Essays on Commodity Prices and their impact on Economic Growth and Financial Markets

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Abstract

This thesis empirically investigates the impact of commodity prices on the macroeconomy and financial markets by drawing explicitly upon their forecasting power for economic growth and stock market returns. Typically, primary commodity trade generates a significant proportion of national income in resource-rich countries and, therefore, any short-run movements in primary commodity prices may have important consequences for economic growth and national financial markets. Supported by a thorough review of the existing literature, the analysis is carried out in three empirical chapters.

Chapter 2 provides an advancement of the index number theory by developing improved index measures of national commodity export prices for a wide range of countries and territories, 217 in total, over the period of January 1980 to April 2017. It proposes a new approach for data collection, which builds upon the past studies by accommodating more precise and accurate data sets. This study demonstrates empirically that the constructed index series outperform those created in past studies.

Chapter 3 looks at the forecasting power of commodity prices for economic growth for a set of 33 commodity-dependent countries between January 1980 and December 2016. Using a mixed-frequency time-varying approach, the empirical results reveal evidence of in-sample causality from commodity prices to economic growth in the case of 31 out of 33 countries. This inference becomes weaker when the estimation horizon becomes longer. Moreover, the commodity-based predictive regressions outperform the benchmark models in 79% of the countries. The substantial evidence found in support of a link between commodity prices and economic growth indicates the long-standing requirement for trade diversification in countries that remain heavily dependent on commodities.

Chapter 4 investigates the relationship between global commodities and national financial markets for 63 countries and territories between January 1951 and March 2018. The study considers five measures of global commodities that are defined as global shocks: world oil prices, world oil demand, world oil supply, world commodity prices (all items) and world metal prices. Using a mixed-frequency time-varying approach, this study provides evidence that commodity prices can predict stock market returns. In the best-case scenario, the world economic activity, denoted as world oil demand, has forecasting power on stock market returns for 54 out of 63 countries. Whereas, in the worst-case scenario, the world oil (metal) prices predict stock market returns for 42 out of 63 countries. This study demonstrates that world commodity prices (all items) exert more influence on stock market returns than oil prices.

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Chapter 1. Introduction

In recent decades, several commodity-dependent countries have experienced remarkable changes in their economies due to a fast economic progress (for example, the rise of emerging powers, like Brazil, Russia, India, China, South Africa, in the world economy), a series of economic and financial events, such as the many crises that have affected the world economy since the 1980s (for example, the 1980s Latin American debt crisis, the 1997 Asian Financial Crisis, the 1998 Russian Financial Crisis, the 2007–2008 Global Financial Crisis) and the ensuing reforms in both the financial and real sectors.

This thesis contributes to the existing literature by providing a comprehensive database of up-to-date country-specific price indexes of commodity exports; it also explores the effects of commodity price dynamics on economic growth and national stock markets.[1]

Recently, as more detailed trade data has become available, the research exploring disaggregated commodity data has been growing. For countries that are recovering after a crisis, one of the main challenges is making their economies and financial markets less dependent on commodity prices. In this direction, a precise measure of the national commodity price movements is required, especially for policymakers to be able to evaluate the dependence of the country's economy and financial markets on commodity dynamics.

At present, up-to-date indexes of national commodity export prices exist only for three countries in the world, namely Australia, Canada and New Zealand. This is clearly not a representative sample of all commodity export-dependent economies.

To help fill this gap, the first empirical chapter of this thesis (i.e. Chapter 2) contributes to the existing literature by constructing country-specific commodity export price indexes using disaggregated trade data for both developed and developing countries. Particularly, Chapter 2 builds on earlier work by (1) constructing a monthly index series for 217 countries and territories, (2) covering the period from January 1980 to April 2017 and (3) providing commodity price sub-indexes for 13 different commodity categories. To be more explicit, the chapter makes use of an index number formula that allows the database to be easily updated and, therefore, to be extended to the most recent period, so that it can serve as a reference point for future studies that focus on commodity-dependent economies.

As highlighted by Deaton and Miller (DM) (1995), the commodity weights used in the construction of the national commodity price indexes should be held fixed over time in order

[1] This thesis uses the definition of commodity as a meaning of a raw or unprocessed material that requires processing before consumption (i.e. primary commodity).

to construct a potentially exogenous variable and, thus, exclude the volume effects of changes in commodity prices. In addition, the abundance of commodity products in the index basket serves as a sufficient and necessary condition for creating a standardised database with a precise index series that represents the true movements of the national commodity price series. Hence, the choice of the index formula and the richness of the index commodity basket are two important aspects in the construction of a world database of country-specific price indexes of commodity exports. Chapter 2 of this thesis considers these aspects with respect to the following research question (RQ):

RQ 1. What is the most appropriate index formula for constructing a world database of country-specific price indexes of commodity exports?

The index number formula of DM is used for the construction of a database of country-specific price (sub-)indexes of commodity exports. The DM formula is considered the most appropriate, as it allows the construction of a database that is rich in terms of the number of countries by using the data available in trade statistics. More precisely, most of the other index formulas require continuous volume data for the construction of the index series, e.g. Fisher and Paasche indexes, while the DM formula does not. However, the volume trade data that is disaggregated to a national level of commodity-specific exports is rarely available, especially for developing countries. Even when such data is available, it is either for a short or discontinuous duration due to various economic and political events such as export bans. All of these arguments indicate that the use of the DM index is the most appropriate for the construction of a comprehensive database of country-specific price indexes of commodity exports. In addition, the DM index is the only existing index number formula that can utilise the available trade data to construct a database that (1) contains a monthly frequency series and (2) spans from January 1980 to April 2017.

The DM index formula reveals the true price movements in the national commodity markets by ignoring the volume effects from the index construction. In fact, the commodity export weights used in the construction of our national commodity price indexes are held fixed over time and, therefore, the index movements are unaffected by the changes in the quantity of commodity exports. The study aims to construct a potentially exogenous variable and, thus, excludes the volume effects of changes in commodity export prices (Cashin et al., 2004).

A unique feature of our new index database of national commodity export prices is the inclusion of the prices of dairy products in the index commodity basket. This significantly improves the accuracy of the index series as compared to those constructed in previous studies, e.g. those of Sahay et al. (2002) and Cashin et al. (2004). As highlighted by Sahay et al. (2002), their constructed index and the official national commodity export price indexes are quite similar for Australia and Canada, while they differ somewhat for New Zealand due to the exclusion of dairy products from their constructed index. In contrast, our constructed index for New Zealand is found to be strongly correlated with the official index that is constructed by the Australia and New Zealand Banking Group (ANZ Bank). A possible reason for this is the inclusion of dairy products within our index commodity basket owing to their large share in the total exports of New Zealand. Given this, our study attempts to contribute to the commodity index literature, since up until now, most of the literature has neglected the prices of dairy products from the construction of the national commodity export index.

In general, this study provides a comprehensive database of country-specific price indexes of commodity exports for 217 countries and territories over the period of January 1980–April 2017. Chapter 2 follows the same framework as that of the United Nations Statistics Division (UNSD), collated via Common Format for Transient Data Exchange and Conference on Trade and Development, when selecting the list of countries for the world database. As some of the countries in our database have experienced changes in their geopolitical borders, country indexes should be selected from our database after careful consideration.

Chapter 3 focuses on quantifying the relationship between commodity prices and economic growth for both commodity-importing and exporting countries. The evidence of such a link existing may be vital to better understand the stages of economic development in developing countries, especially in those that are still heavily dependent on commodities. In fact, a better understanding of the level of commodity dependence is important for a country's ability to design trade policy. This is especially true in times of crisis, such as the 1997 Asian Financial Crisis and the 2007–2008 Global Financial Crisis, when a majority of the commodity prices steeply increased. As such, this study aims to shed light on the commodity-growth nexus.

Ideally, the role of the decision-makers, policymakers in particular, is to use the fluctuations in the commodity prices in order to facilitate sustainable economic growth. The policymakers would probably ignore changes in the world commodity prices if these are known to have no effect on economic growth. However, if the world commodity prices have a direct impact on economic growth, the policymakers have to keep an eye on the short-term changes they bring about. If they fail to do so, it can lead to a serious sacrifice in terms of long-term economic growth. In practical terms, the policymakers in commodity-dependent countries react whenever there is a shock in the world commodity prices. However, any attempt to identify whether the changes in the commodity prices affect economic growth is a real challenge for them. Chapter 3 addresses the following pertinent research question:

RQ 2. Do commodity prices cause economic growth?

In order to answer the above question, this study adopts the recent mixed-frequency approach of Ghysels et al. (2016), which accounts for the data sampled at different frequencies. This feature of the approach is crucial, as data on economic growth is predominantly available at least at a quarterly frequency, while commodity prices are available at a higher frequency. Since standard (single-frequency) VAR literature requires all variables to have the same frequency, high-frequency data is usually aggregated into a lower-frequency, such as that of commodity prices. Such temporal aggregation is known to have an adverse impact on statistical inference (see Marcellino, 1999; McCrorie and Chambers, 2006; Ghysels, 2016; Ghysels et al., 2016 for discussion). Therefore, in addition to the standard VAR models, this study adopts a mixed-frequency vector autoregressive (MF-VAR) model specification proposed by Ghysels et al. (2016). The obtained results align with the statements of Ghysels et al. (2016) that the MF causality tests better recover causal patterns as compared to the traditional low-frequency (LF) approach.

Furthermore, while most studies in the literature focus on commodity exporters, less evidence is available for commodity-importing countries. Chapter 3 builds upon the past literature by considering countries that are commodity-dependent with respect to import and export or both. In fact, the study of Narayan et al. (2014), who consider the impact of oil prices on economic growth, is the most closely related to this study. However, while the authors distinguish between developing and developed economies, they do not provide a clear conclusion in terms of commodity-importing vs. exporting economies.

A further relevant issue is that the forecasting ability of commodity prices for economic growth is attributable to the speed of information transmission (Kang, 2003). In other words, considering prediction in a single-horizon period may fail to reveal a commodity-growth causal pattern that actually exists. Therefore, this thesis considers prediction at different horizon periods and provides evidence of any patterns that are detected with respect to short- and long-horizon periods. Nonetheless, it should be noted that prediction in long-horizon periods may require the consideration of additional control variables in the VAR models. This is because the commodity-growth relationship is potentially exposed to the effect of other external factors (e.g. political events and trade policy reforms as well as macroeconomic variables such as inflation, interest rates, exchange rates, and others) when the estimation horizon increases (see Cavalcanti et al., 2015 for discussion). Future research using the results from this thesis as a motivation should test for predictability using a set of control variables when estimating longer horizon periods.

In addition, Chapter 3 demonstrates that the commodity-growth relationship may be unstable over time. In particular, the results from Andrews' (1993) Quandt likelihood ratio (QLR) tests provide evidence that favours parameter instability in the low-frequency models. To account for this, the full sample MF-VAR models are extended to a time-varying framework that allows for the analysis of the dynamic nature of the commodity-growth relationship. Using a mixed-frequency time-varying approach, the empirical results reveal evidence of short-horizon in-sample predictability from commodity prices to economic growth in the case of 31 out of 33 countries. Meanwhile, the feedback causality is discovered for 23 out of 33 countries. More precisely, Chapter 3 adds to the earlier studies of Ghysels (2016) and Ghysels et al. (2016) by providing concrete evidence in support of the appropriateness of mixed-frequency models for estimating the causal link between commodity prices and economic growth.

Further, Chapter 3 examines the extent to which world commodity prices can help out-of-sample forecasting economic growth. The motivation for extending the analysis to out-of-sample predictability is driven by the claim of Timmermann (2006) that the in-sample predictive ability often fails to translate into out-of-sample success. Although this is a widely documented pattern in the forecasting literature, this study found strong evidence in support of the commodity price out-of-sample predictability for economic growth. It must be highlighted that the commodity prices clearly outperform the random walk benchmarks. That is, the forecast combination results indicate that the commodity-based predictive regression models outperform the benchmark models for 79% of the total number of countries for at least two of the three benchmarks. This inference is valid regardless of the estimation method.

Last but not least, the robustness of the results is confirmed using different proxies of commodity prices. On the one hand, the world commodity price indexes are selected in a manner that allows their diversity in terms of construction and commodity baskets. For example, some indexes weigh commodities equally in the index basket, while others apply different weighing formulas.[2] This in fact ensures that the choice of the index series does not affect the conclusions of the study. On the other hand, this study uses the national commodity export index from Chapter 2 to examine the commodity-growth link in terms of national

[2] For example, the IMF non-fuel commodity price index uses commodity weights derived from their relative export trade values compared to the total world export trade, while the Thomson Reuters Core Commodity Index weighs equal all commodities in the index basket.

commodity prices. The outcomes from the robustness check section confirm the main finding of this study, which is that commodity prices have a causal effect on economic growth. The results from this study provide policymakers with evidence for the predictive content of commodity prices on economic growth for both commodity exporters and importers.

Given the importance of commodity markets, Chapter 4 supplements the third chapter of this thesis by determining the link between commodity and financial markets. Chapter 4 demonstrates that the interaction between world commodity prices and national stock market returns can be essential by shedding new light on the widely debated issues surrounding the commodity-stock relationship. The following fundamental research question is addressed in Chapter 4:

RQ 3. Do commodity prices cause stock market returns?

To answer this research question, Chapter 4 presents a cross-country analysis of the link between global commodities and national financial markets in a set of 63 countries and territories over the period of January 1951–March 2018. The study considers five measures of global commodities that we define as *global shocks* (henceforth): world oil prices, world oil demand, world oil supply, world commodity prices (all items) and world metal prices.

Much of the commodity-stock research has focused on stock markets in developed countries – mainly the US. Little is known about the predictive power of global shocks on national stock market returns beyond the US. Since several developing countries are commodity-dependent, as highlighted by Smith (2004), we examine the effect of global shocks on national financial markets for both developed and developing countries.

Furthermore, the literature on the impact of commodities on stock markets has mostly concentrated on the effects of oil prices, while the evidence of a relationship existing between commodities in general and national stock markets is still limited.[3] In particular, a smaller but recent strand of papers has examined the co-movement between non-fuel commodity prices and stock market returns. For example, authors have looked at metals such as gold (Baur and McDermott, 2010; Hood and Malik, 2013; Arouri et al., 2015; Basher and Sadorsky, 2016; Mensi et al., 2018) and copper (Sadorsky, 2014); in addition, foodstuffs, such as sugar, coffee and cocoa (Creti et al., 2013), have been investigated as well. While a handful of recent studies have been conducted on the co-movement between non-fuel commodity prices and stock market returns, the direction of causality has not yet been fully investigated. Therefore,

[3] For example, see Sadorsky (1999), Park and Ratti (2008), Kilian (2009), Kilian and Park (2009), Narayan and Sharma (2011), Adams and Glück (2015), Chiang and Hughen (2017) and Christoffersen and Pan (2018).

we extend our analysis to include other commodities, such as metals and commodities in general.

Compared to past studies, Chapter 4 further contributes by using the most recent econometric methods, which account for the presence of data sampled at different frequencies (Ghysels, 2016). This is required because the high-frequency data typically used to investigate national stock markets is often unavailable at daily or weekly frequencies for developing countries and for long historical time series – a 65-year period in our case. At the same time, low-frequency stock market data, such as quarterly or annual, may cause a loss of information in empirical models (Orcutt et al., 1968). Our study accounts for this issue by using monthly stock price data in the estimation of the commodity-stock relationship.

However, a series of world commodity prices are available at weekly frequency.[4] To prevent a loss of information from temporal aggregation, as discussed by Ghysels (2016), the MF-VAR modelling approach is adopted. The advantage of employing the MF-VAR model is that it enables the estimation of both weekly and monthly frequency variables together in the same framework. Since classical models require all variables to have the same frequency, variables typically available at a high-frequency, such as commodity prices, are often aggregated at the lowest frequency. However, recent research has shown that the temporal aggregation has an adverse impact on statistical inference (see McCrorie and Chambers, 2006; Andreou et al., 2010; Götz et al., 2014; Eraker et al., 2015; Schorfheide and Song, 2015; Ghysels, 2016; Ghysels et al., 2016; Motegi and Sadahiro, 2018). For example, given that commodity prices are known to be highly volatile (see Deaton and Laroque, 1992; Deaton, 1999), working with a common low-frequency approach is likely to cause the omission of useful information regarding the time series properties of the data (Götz et al., 2016). It is particularly well known that the Granger causality in a VAR framework is not invariant to temporal aggregation (see Granger and Lin, 1995; Marcellino, 1999). Therefore, we use the MF-VAR procedure of Ghysels et al. (2016), which aims to overcome the potential issues that arise from temporal aggregation.

In recent years, the time-varying nature of the commodity-stock relationship has been considered by a number of authors (for example, see Miller and Ratti, 2009; Broadstock and Filis, 2014; Kang et al., 2015). Although numerous studies have been conducted on the commodity-stock relationship, its time-varying nature has not been fully investigated yet. For instance, the past literature has mainly focused on the time-varying oil-stock relationship and

[4] We cannot use daily data (even if available) in combination with monthly data in the same model because this leads to parameter proliferation (see Ghysels, 2016 for discussion).

less is known for other commodities such as metals and commodities in general. This study contributes to the understanding of the significance of the commodity market as a whole, especially for those countries that are still heavily dependent on non-fuel commodities for their main source of income (see Collier and Goderis, 2008 for discussion). The results from this study provide policymakers with evidence regarding whether the connection between commodities and stock markets varies over time.

Therefore, not only does our investigation include a wider set of countries and world prices but also adopts a rich methodological approach, where the MF-VAR and the low-frequency VAR (LF-VAR) models are constructed and a battery of Granger-causality tests are performed to gauge the commodity-stock markets relationship. The analysis is extended to a time-varying framework in order to account for periods during which the world economy and the national stock markets have experienced several large price swings and structural changes.

In a nutshell, Chapter 4 presents a cross-country analysis on the links between global shocks and national financial markets for 63 countries and territories between January 1951 and March 2018. The data modifications and combinations contribute to the field by revealing new empirical evidence that can help policymakers make decisions.

Chapter 2. Towards a New Database of Country-Specific Price Indexes of Commodity Exports

2.1 Introduction

Commodity price indexes have been increasingly used for macroeconomic research. In particular, their utilisation has grown rapidly since the publication of the seminal article of Grilli and Yang (1988). The two economists contribute to the literature by constructing a *commodity-specific* price index database that enables researchers to not only investigate the interaction between commodity prices and, for example, other macroeconomic variables but also analyse the variables on their own – that is, univariate analysis.[5] Their database focuses only on global price trends of a specific commodity, however, not much is known about the movements of *country-specific* commodity prices. Deaton and Miller (1995) made the first major contribution in terms of country-specific commodity price indexes. The authors created a database with annual country-specific price indexes of commodity exports for sub-Saharan African countries. However, commodity export-dependent countries are not restricted to only the sub-Saharan African region. Therefore, other studies, such as those of Sahay et al. (2002) and Cashin et al. (2004), extend the Deaton and Miller (1995) database to include countries and regions from the rest of the world. Another exclusive feature of these two studies is that they create non-fuel country-specific price indexes.[6]

Nonetheless, past studies have created databases that may be impracticable in certain cases due to them (1) restricting the country sample to a specific geographical region and (2) considering a particular group of commodities when constructing the index series, including only non-energy products in the index construction, for example. Therefore, this study aims to provide a worldwide database of country-specific price indexes of commodity exports that can serve a wider array of research objectives.

This study builds on earlier work by (1) constructing a monthly index series for 217 countries and territories, (2) covering the period from January 1980 to April 2017 and (3) providing each country with a commodity price sub-index for 13 commodity categories, if applicable.[7] To be more explicit, we use an index formula that allows our database to be easily updated and, therefore, to be extended to the most recent period in order to establish a convenient starting point for an empirical analysis.

[5] Prior to Grilli and Yang (1988), most studies in the economic literature have excluded commodity prices from their empirical analysis due to the lack of continuous commodity price data.

[6] Both studies construct the database for their analysis using the same (1) index formula, (2) time and country coverage and (3) set of commodities. In other words, they use identical databases.

[7] A sub-index is an index that represents a sector of a larger one, i.e. the all commodities index. The sub-index is not constructed only if the total exports of the given commodity group is equal to zero.

At present, up-to-date indexes of national commodity export prices exist only for three countries in the world – Australia, Canada and New Zealand.[8] This is clearly not a representative sample of all export-dependent economies. Therefore, our analysis is extended to include a monthly index series for 217 countries and territories from around the world.

Furthermore, the construction of the index series for each of the three countries is different, e.g. using non-identical index formulas, which causes inconsistency if cross-country analysis is undertaken. Specifically, the Fisher formula is used for Canada, while the Laspeyres formula is used for Australia and New Zealand.[9] These index formulas are inappropriate for constructing a monthly large-scale commodity price database due to the following shortcomings. First, the Fisher formula requires monthly data on a country's quantity commodity exports. However, the availability of such data is extremely scarce in world trade statistics. Second, the Laspeyres formula uses the arithmetic mean; therefore, any price change from the current period to the base period is not reciprocal to the original price change. In other words, the Laspeyres formula exhibits an upward bias and overestimates the "true" price change (see Boskin et al., 1998; Hill, 2004; IMF, 2009 for discussion). That is to say, the *time reversibility property* is not satisfied (see Diewert, 1998 for discussion). An alternative index formula for the construction of country-specific price indexes of commodity exports is the one proposed by Deaton and Miller (1995). The index formula proposed by Deaton and Miller (1995) uses geometric mean, which is more desirable than the arithmetic average because it satisfies the *time reversibility* condition, as emphasised by Diewert (1998).[10]

In particular, the DM index is a more suitable choice than the above-mentioned formulas for several reasons. First, the DM index does not require quantity data for its completion. This makes it more appropriate in practice than the Fisher index, for example, due to the scarcity of volume trade data in the world trade statistics. Second, the DM index uses the geometric mean for its construction. Therefore, the DM index accounts for the relative price changes of the commodity export prices, whereas the Laspeyres formula does not. Under these circumstances, we might conclude that the DM index is an appropriate index number formula for constructing our world database of country-specific price indexes of commodity exports. A principal reason for this choice is the current trade data availability.

[8] The index databases of Deaton and Miller (1995) and Cashin et al. (2004) are updated up to the years 1992 and 2002 respectively.

[9] The Canadian index is made available by Bank of Canada, whereas Australia and New Zealand's indexes can be obtained from the Reserve Bank of Australia and ANZ Bank respectively.

[10] *Time reversibility* is a property from the axiomatic price index theory that requires the resulting price index to be the reciprocal of the original price index if the prices and quantities in the two periods being compared are interchanged.

In fact, the country-specific price indexes of commodity exports have not been used as much in research. Most previous studies focus either on the world prices of individual primary commodities (Cuddington and Urzua, 1989) or the country terms of trade (Spraos, 1980), or they construct commodity-specific price indexes (Grilli and Yang, 1988). Nonetheless, these measurements suffer from severe limitations when used for tracking the price movements in national commodity markets, as discussed by Cashin et al. (2004). First, only a few exporters of primary commodities are specialised to the extent that the export prices of an individual commodity product can effectively approximate the true price movements in their national commodity markets. Precisely, most economies in the world export more than one commodity, as highlighted by Dehn (2000). Second, the calculation of the terms of trade indexes includes data on both imports and exports. As such, the term of trade index is highly reliant on a country's composition of trade (Deaton and Miller, 1995). Therefore, the commodity terms of trade can be assumed to be an inappropriate measure of the true price movements in national commodity export markets. Third, the commodity-specific price indexes are likely to poorly represent the price movements in national commodity export markets, as emphasised by Sahay et al. (2002). This corresponds to the fact that the commodity weights of the commodity-specific price indexes do not reflect the trade structure of the individual economies, specifically they remain identical for all economies (see Cashin et al., 1999 for discussion). With this in mind, we conclude that the DM index is the most appropriate index formula to represent the price movements in national commodity export markets. Once again, this statement is made on the basis of the current data availability in world trade statistics.

Indeed, this study aims to alter the perception of working with T time series observations of N countries, where either T or N is quite small. It compiles information for a large number of countries (N = 217) without sacrificing information in the time series dimension; precisely, the number of observations (T) in each time series is equal to 448. Therefore, we can link our database with the term *data-rich environment*, where N and T are both large, as defined by Bernanke and Boivin (2003).[11] Therefore, our database can find applications in a broad range of economic fields. Examples of future work can focus on, but do not have to be limited to, the law of one price (see Ardeni, 1989; Parsley and Wei, 1996), resource curse (see Sachs and Warner, 1999; Collier and Goderis, 2008; Frankel, 2010), exchange rate dynamics (see Corden, 1984; Deaton and Miller, 1995; Amano and Van Norden, 1998; Hinkle and Monteil, 1999; Chen and Rogoff, 2003; Cashin et al., 2004; Chen et al., 2010), the foreseeability of

[11] Bernanke and Boivin (2003, p. 15) coined the term "data set is a "rich" one that contains much more information than can be extracted from a relatively small set of macroeconomic time series".

economic activity (see Pindyck and Rotemberg, 1990; Deaton, 1999; Akram, 2009; Narayan et al., 2014), determination of current and future inflation (see Gospodinov and Ng, 2013; Chen et al., 2014) and the reallocation between the tradable and non-tradable sectors (see Goldstein et al., 1980; Grilli and Yang, 1988).

To sum up, the motivation of this study is exemplified by the subsequent appealing features. First, we create a novel framework for the data collection, which relieves researchers from having to manage data changes and revisions. Second, we use a finer level of disaggregation for the different classes of commodity products when constructing the index series. This aims to improve the accuracy of the index series in a manner that more closely represents the true price movements in the national commodity markets. Third, we include as many as 72 commodity products in the process of index construction. On the one hand, this is the largest basket of commodities when compared to all previous studies that have constructed databases of national commodity export price indexes. As a result, our database is able to accommodate an index series for a set of 217 countries for the period between January 1980 and April 2017. This makes it the largest database of country-specific price indexes of commodity exports in terms of (1) number of countries and (2) number of time observations.[12] On the other hand, having a large basket of commodities allows us to construct a more precise measure of national commodity prices. As an illustration, Sahay et al. (2002, p. 53) concludes that their constructed index series for New Zealand "differ somewhat, due to the exclusion of dairy products from the constructed index" in contrast to the official ANZ Bank indexes.[13] Accordingly, we are the first to include data on dairy products in the construction of national commodity price indexes. Based on the results from the Pearson correlation, our New Zealand indexes have a higher accuracy than the index series constructed by Sahay et al. (2002) and Cashin et al. (2004). More precisely, the Pearson correlation coefficient between the index series of Sahay et al. (2002) and the official ANZ Bank index for New Zealand is found to be 0.407, while the Pearson correlation coefficient between the constructed series and the official ANZ Bank series is 0.940. Fourth, we provide sector-specific national commodity export indexes for the following 13 categories: (1) All commodities, (2) Non-energy commodities, (3) Food, (4) Cereals, (5) Vegetable oils and protein meals, (6) Meat, (7) Dairy, (8) Beverages, (9) Agricultural raw materials, (10) Metals, (11) Energy, (12) Fertilizers and (13)

[12] In contrast, some of the most influential studies on national commodity export price indexes are based on the following setting: Deaton and Miller (1995) use 21 commodities for 32 countries, Dehn (2000) uses 57 commodities for 113 countries, Sahay et al. (2002) use 44 commodities for 58 countries, Cashin et al. (2004) use 44 commodities for 58 countries and Bodart et al. (2012) use 42 commodities for 68 countries.

[13] The ANZ Bank is the organisation that publishes the official commodity price index for New Zealand on a regular basis.

Precious Metals.[14] The sector-specific indexes help tracking the performance of specific sectors more precisely (for example, Cashin et al. (2004) use non-fuel national commodity export index).[15] In brief, this study greatly contributes to the literature by (1) providing a data-rich environment for economic analysis, (2) relieving the researchers from having to manage data changes and revisions and (3) creating a consistent in methodology database that facilitates the replication and comparison of results.[16]

The rest of the chapter is organised as follows. Section 2.2 provides an overview of the contribution of this study to the literature and briefly discusses the background of commodity price index databases and methods. Section 2.3 highlights the limitations of the existing national commodity price indexes and provides a timely solution to them. Section 2.4 presents the methodology adopted for construction of country-specific price indexes of commodity exports in the chapter. Further, the data and sources are discussed in Section 2.5. Section 2.6 follows this, where our index series are compared to the national commodity export indexes that are calculated by central and commercial banks as well as to the index series constructed by Sahay et al. (2002) and Cashin et al. (2004). Section 2.7 concludes the chapter.

2.2 Motivation

This section briefly outlines the main contributions of the research. Moreover, it reveals the potential limitations that can be faced when constructing a large world database of country-specific price indexes of commodity exports. In addition, it highlights the shortcomings of past studies and suggests steps for improvement.

First, this study introduces new guidelines for data collection that improves the accuracy of the information obtained from the international trade statistics. We identify differences in the reported trade data through different classification systems of the United Nations Common Format for Transient Data Exchange (UN Comtrade) and United Nations Conference on Trade and Development (UNCTAD) databases. For example, we notice that when exporting countries report trade data information to UN Comtrade and UNCTAD, there may be two different values for the same commodity owing to the different revisions of trade classification systems. Thus, it is useful to reconcile these into a single figure.

[14] The categories have been taken from the International Monetary Fund (IMF, 2018). More information on commodity classification is available at http://www.imf.org/external/np/res/commod/index.aspx. Additional categories are included as well.
[15] If the sum of export values of all commodities included in a category classified above is equal to zero, the country is assumed to be a non-exporter for this group of commodities and, therefore, a sector-specific index is not constructed.
[16] This study embellishes the construction differences of the existing country-specific price indexes provided by central and commercial banks.

Unfortunately, the issue of data quality has been neglected in the current literature, as discussed by Gaulier and Zignago (2010). Such an issue may affect the accuracy of the index series. This is because wrongly recorded data may affect the weighting of commodities in the index basket and, therefore, lead to imprecise index movements. In other words, if one does not take into consideration the current data quality issues in international trade statistics, one may fail to construct an accurate commodity price index. Therefore, this study proposes a new method for data collection with the aim of improving the quality (accuracy) of the index series in our database.

In particular, the method proposed in this study aims to identify incorrectly reported and missing trade data values.[17] While we focus on the data reported by the UN Comtrade and UNCTAD databases, the same procedure can be applied to the other statistical databases. Specifically, the use of a robust data collection procedure aims to obviate the discrepancies in different trade databases and combine all trade data in a common data set. In our study, this data set is used for constructing the index weights. That is to say, our index formula requires data on trade values and, therefore, a failure to construct accurate commodity weights may result in imprecise index movements; for example, a given commodity price may be either under- or over-weighted.

Second, this study emphasises that commodities are heterogeneous goods and assigning them identical prices regardless of their distinct features may cause distortions in the movements of national commodity export price indexes. Indeed, the heterogeneity of commodity products reflects in their pricing at the commodity market, which has been confirmed by Pindyck and Rotemberg (1990, p. 1174), who state, “All commodities are at least somewhat heterogeneous.” Unfortunately, this feature of commodity products is often ignored in the past studies, such as in those conducted by Deaton and Miller (1995), Dehn (2000), Sahay et al. (2002) and Cashin et al. (2004). An example is wool. All of the four studies consider trade data on “wool” instead of using separate data for “coarse wool” and “fine wool”. This is important as international prices for “coarse wool” and “fine wool” are not identical in the world commodity market and do not always tend to move together. Other examples of commodities for which heterogeneity features are neglected by predominant part of the past studies, but considered in ours, include natural gas (liquefied and in gaseous state), timber

[17] For example, the trade data on the product soybean meal is neither available under SITC Rev. 1 nor any other SITC system at UN Comtrade. But this data is available through any HS trade system at the UN Comtrade. Another example is that of the trade data on “Crude petroleum” for Syria, which is missing from each HS system of the UN Comtrade for the period before 2001 (see Section 2.3.1 for a detailed discussion). However, this data is publicly available on UN Comtrade through SITC Rev. 1 and SITC Rev. 2 systems. Thus, solely using a trade classification system to obtain trade data may lead to the incorrect construction of the index weights. This is especially true in the case of Syria, where crude oil is the main exported product.

(hardwood and softwood), milk powder (skim milk powder and whole milk powder). Given these examples, this study concludes that the trade data for each commodity should be collected at a level of disaggregation that best explains the commodity's characteristics. This is important due to the differences in the characteristics of various commodity products. Neglecting these differences may have an adverse impact on the accuracy of index price movements. Therefore, this study uses the exact definition of commodities, where the data allows it, with the aim of constructing more precise national commodity price indexes.

Third, we find that certain past studies were unable to construct a precise national commodity price index due to the insufficient number of commodities in their index basket. Therefore, our study considers a sample of 72 primary commodity products (including dairy products) in the process of index construction. This is the largest index basket of commodities that is used in the compilation of a database of national commodity price indexes among all previous studies in the economic literature.[18] The usage of such a large basket of commodities reduces the possibility of "missing" a commodity product that is part of the country's major primary exports. In that way, this chapter provides rigorous quality assurance to guarantee accuracy and consistency of the national commodity price indexes in our database.

Unfortunately, the issue of "missing" commodity products from the index construction is rather common in the past literature (for example, see Deaton and Miller, 1995; Dehn, 2000; Cashin et al., 2004; Bodart et al., 2012). An example is the earlier study conducted by Deaton and Miller (1995) that explores the relationship between commodity prices and exchange rates in Sub-African countries. The authors do not include precious metals and timber products in the index basket when calculating their national indexes of commodity exports. However, these two groups of commodities are major exports for most African countries, as discussed by Wood and Mayer (2001).[19] Another example is the study of Cashin et al. (2004), who omit some important commodities, such as barley, hardwood sawn, olive oil, poultry and swine meat, from their index calculations. A possible reason for this may be data unavailability. Nonetheless, these commodities are a part of the primary commodity export basket of several countries around the world. Some examples are barley (for Canada, Australia, Russia, Ukraine, Argentina), olive oil (for Tunisia, Morocco), poultry (for New Zealand, Canada), hardwood sawn (for Central African Republic, Canada, Russia, Thailand, Chile, Malaysia, New Zealand) and swine meat (for Canada, Brazil, Mexico, Chile). Since

[18] For example, Deaton and Miller (1995) use 21 commodities, Dehn (2000) uses 57 commodities, Sahay et al. (2002) use 44 commodities, Cashin et al. (2004) use 44 commodities and Bodart et al. (2012) use 42 commodities.

[19] Examples of African countries that are exporters of precious metals are Burkina Faso, Mali, South Africa, Sudan, Suriname and Tanzania, and those of timber are Central African Republic and Gambia.

Cashin et al. (2004) construct a non-fuel index of national commodity export prices, the importance of including these commodities in the construction process of the index series increases. A specific example of this is the Central African Republic, where the export of hardwood sawn accounted for 43% of the total export in 2017 (UN Comtrade, 2018). Therefore, excluding hardwood sawn from the index basket for Central African Republic may lead to an imprecise measurement of the country's commodity export prices; specifically, the index weighting may be incorrect. Given these facts, our study considers the importance of all commodity products and includes these primary commodity products, which are "missing" from the commodity index baskets of past studies, if the data allows it. This aims to improve the performance as well as the accuracy of the national commodity export indexes provided in our database as compared to those constructed in past studies.

Importantly, it must be noted that a large number of commodities in the process of index construction not only have a plausible impact on the accuracy of the index itself but also provide a favourable environment for sector-specific national commodity export indexes.

Fourth, this study makes a key contribution, by constructing a broad set of *sector-specific* national commodity export price indexes for 13 different categories. In particular, we construct *sector-specific* indexes for all countries in our sample; however, if a country is not an exporter of any of the commodity products that are included in the sub-index commodity basket, the sub-index has not been constructed for this country. In fact, we define the sector-specific index (namely sub-index) as an index that covers a particular group of commodities, energy or precious metals for example. This allows researchers to explore the impact of a particular group of commodities on the macroeconomic environment for a given country, for instance.

Some of the most well-known databases of commodity price sub-indexes include Grilli and Yang (1988), the International Monetary Fund (IMF), the World Bank, the United Nations and the Commodity Research Bureau (CRB). Unfortunately, none of these databases provide country-specific sub-indexes of commodity exports. The importance as well as the demand for sub-indexes that focus on a particular commodity group can be seen in the past literature, such as the studies of Sahay et al. (2002) and Cashin et al. (2004), who created indexes of non-fuel national commodity export prices. To put it in other words, our database provides national price sub-indexes of commodity exports (including non-energy indexes) that allow for a more conventional economic analysis of a country.

Consequently, there are various areas where our database can be applicable. Examples include, but are not limited to, resolving economic puzzles, such as the excess comovement hypothesis (see Leybourne et al., 1994; Deb et al., 1996; Ai et al., 2006) and the natural resource curse hypothesis (see Collier and Goderis, 2008; Frankel, 2010). It can also be used in the exploration of the forecasting power of commodity prices with regard to the exchange rates (see Chen et al., 2010), inflation (see Gospodinov and Ng, 2013) or economic growth (see Narayan et al., 2014).

In brief, this study creates a world database of national price sub-indexes of commodity exports with the aim of providing a favourable environment for applied economic work.

There are other key features that are an indivisible part of the database construction process, such as the data frequency and index formula. On the one hand, we create high-frequency index series that accommodate time series analysis and capture the price fluctuations in the commodity markets with greater precision. Particularly, we construct our database at a monthly frequency. We were unable to construct our database with a higher frequency due to the unavailability of data. On the other hand, we use the DM index formula because (1) it does not require quantity data for its completion, which makes it more appropriate in practice than the Fisher index for example, and (2) it does not overestimate the "true" price changes, which is the case when using the Paasche and Laspeyres index formulas for instance (Diewert, 1998). This is to say that our database is consistent with the axiomatic theory. In addition, we provide empirical evidence supporting the reliability of the DM indexes as compared to the index numbers from previous studies. We find the existence of a strong correlation between our constructed index series and the official index series provided by central and commercial banks. A further discussion on this has been presented in the empirical section of this chapter.

Moreover, the constructed database consists of *monthly* national commodity price indexes, which contrasts to the majority of the previous studies in the existing literature that rely mainly on lower-frequency data such as annual frequency data (for example, see Grilli and Yang, 1988; Deaton and Miller, 1995; Bleaney and Greenaway, 2001; Cashin and McDermott, 2002; Cavalcanti et al., 2015) and quarterly frequency data (for example, see Borensztein and Reinhart, 1994; Dehn, 2000; Akram, 2009; Chen et al., 2010; Jacks et al., 2011). Given that commodity prices are known to be highly volatile (see Deaton and Laroque, 1992; Deaton, 1999), working with a lower-frequency data is likely to cause omission of useful information regarding the time series properties of the data (Götz et al., 2016). In addition, the finite-sample power of testing procedures may fall when number of available observations is small (Marcellino, 1999). Moreover, Ferraro et al. (2015, p. 139) highlight, "The most likely explanations for why the existing literature has been unable to find evidence of predictive power in commodity prices are that researchers have focused on low frequencies where the short-lived effects of commodity prices wash away and that the predictive ability in commodity prices is very transitory". Given these arguments, we acknowledge the importance of constructing a relatively higher-frequency (monthly) index series. Unfortunately, we were unable to construct a database with a frequency higher than monthly owing to data unavailability.

To highlight the advantages of our database, the table below provides information regarding the time span and the number of observations for the most commonly used databases of national commodity export price indexes.

Study/Institution	Data frequency	Time span	Total number of observations per time series	Sample of countries
Deaton and Miller (1995)	Annual	1958 – 1992	35	32
Dehn (2000)	Quarterly	1957Q1–1997Q4	164	113
Sahay et al. (2002)	Monthly	1980M1–2002M3	276	58
Cashin et al. (2004)	Monthly	1980M1–2002M3	276	58
Bodart et al. (2012)	Monthly	1980M1–2008M12	348	68
Bank of Canada	Monthly/weekly	1972M1–2017M4	544	1
Reserve Bank of Australia	Monthly	1982M8–2017M4	416	1
ANZ Bank	Monthly	1986M1–2017M4	376	1
Our index	Monthly	1980M1–2017M4	448	217

Table 2.1 Databases of Country-Specific Price Indexes of Commodity Exports

Table 2.1 demonstrates that our study has the second largest number of time observations among all the existing databases of national commodity export price indexes. In terms of the number of observations, only the regularly updated databases constructed by Bank of Canada, Reserve Bank of Australia and ANZ Bank are comparable to this study. These databases provide commodity price indexes for a single country – Canada, Australia and New Zealand respectively. As can be noted, a database with present-day data of national commodity price indexes for other economies apart from these three does not exist. This study contributes to the literature by providing a high-frequency *monthly* environment that allows for a time series analysis of price fluctuations in national commodity markets.

Furthermore, the selection of an appropriate index formula for the construction of national commodity price indexes represents one of the biggest challenges for studies of this kind. The index formula should be both applicable in terms of the database size and consistent with the economic and axiomatic approaches. Therefore, this chapter employs the DM index formula not only because of its desirable properties based on axiomatic theory but also for its suitability with economic theory. Moreover, this study is the first in the existing literature to show empirically the strong correlation between the DM commodity price indexes and the official commodity price indexes created by the central and commercial banks. A further insight for choosing the DM index formula is provided below.

One may choose a chain-link formula because it allows the "index commodity basket" to be updated on a (ir)regular basis, whereas a fixed-base formula holds all the weights constant over time.[20] In other words, the chain index allows for the substitution of the commodities within the index basket (over time), while the fixed-base index does not. However, this substitution comes with the price of a chain-drift bias, as per axiomatic theory (Diewert, 1995). "A chain index is said to drift if it does not return to unity when prices in the current period return to their levels in the base period" (IMF, 2009, p. 607). As such, the chain index is unable to account for the relative price changes of the commodity exports (Malmquist, 1953).[21] For example, if a commodity price index has an upwards (downwards) chain-drift, the index overvalues (undervalues) the commodity prices within the country's export market. This causes the relative price change between two different periods to be inaccurate. Hence, one can conclude that the chain index approach is unsuitable for constructing commodity price indexes, based on the economic theory.

Indeed, the fixed-base DM formula grants an advantage, as the "export shares of the index are time-invariant to ensure that the time series variation of the international export price index is exogenous to changes in the domestic economic environment" (Brückner, 2012, p. 19). Importantly, our study uses time-invariant weights because its purpose is to construct a variable that is exogenous and uncorrelated with the supply responses from the world commodity market, as discussed by Deaton and Miller (1995). Therefore, the volume effect has to be controlled by holding the quantities fixed throughout time, as has been done in this study.

[20] There are two main types of index number formulas – chain and fixed-base (Diewert, 1978). Precisely, the DM formula should be classified as a fixed-base index formula because it holds all commodity weights constant over time (see IMF, 2009 for discussion).
[21] This can be verified by the multi-period identity test of Diewert (1988).

Nonetheless, one may argue that the DM index formula has its limitations, as it is unable to capture natural resource discoveries (due to quantities being fixed), and this does not reflect in the movements of the index series. We agree that this is a limitation of the DM formula and leave this issue to be resolved by a further research work. Meanwhile, the empirical section of this study examines the robustness of our commodity price indexes and finds them to be highly correlated with the official national commodity export price indexes provided by central and commercial banks. Interestingly, we find a strong correlation between the official indexes and the constructed indexes, even though the official indexes use chain-link formulas and, therefore, account for quantity changes. This finding provides some relief in terms of the accuracy of the constructed index series in our database.

Moreover, an alternative source of endogeneity can operate through the individual countries' commodity export prices (Deaton and Miller, 1995). In other words, Chen and Rogoff (2003) note that endogeneity may arise through the market power that certain countries may possess in the world commodity markets. For example, since Chilean copper exports have a large share in the global copper market, the world price of copper may be significantly influenced by the value of the Chilean peso. Another example is Indonesia, which has a vast share in the global palm oil market and, therefore, the world price of palm oil is presumed to be exerted by the value of the Indonesian rupiah.[22] Broda (2004) indicates that only a small number of countries exert such an influence, and they do so on a small share of commodities that they export. The substitution across similar commodity products further mitigates the market power these countries have, even within the specific markets that they appear to dominate (Chen and Rogoff, 2003). Cashin et al. (2004) also conclude that commodity-exporting countries are price-takers in world commodity markets and have negligible long-term market power in terms of their commodity exports (see Mendoza, 1995 for discussion). To address this potential form of endogeneity, our study uses the world commodity prices in the construction of each national commodity price index because they are normally exogenous to the behaviour of individual countries, as highlighted by Deaton and Miller (1995) and Sahay et al. (2002). In addition, the world commodity prices have the advantage of greater accuracy and availability; most importantly, they can be considered exogenous by individual countries that produce a relatively small share of the same commodity (Blattman et al., 2007).

In summary, taking into consideration the above-mentioned concerns, we use the DM index formula to construct our world database of national commodity price indexes. The DM index

[22] Other examples are Brazil's iron ore exports and Cote d'Ivoire's cocoa exports (Broda, 2004).

formula is undoubtedly the most suitable instrument for compiling our database.[23] In addition, the DM index is well-established in the current economic literature as a robust instrument for constructing a national commodity price index (for example, see Dehn, 2000; Sahay et al., 2002; Cashin et al., 2004; Raddatz, 2007; Collier and Goderis, 2008; Brückner and Ciccone, 2010; Bodart et al., 2012; Bodart et al., 2015; Ciccone, 2018). This instils more faith in the reliability of the DM index.

2.3 Data Collection Framework

2.3.1 Trade data synchronisation

If identical trade data values are reported within all the revisions of a given trade classification system, the data collection process would be straightforward. However, this is not always the case with the international trade statistics. That is to say, the reliability of data reported in the international trade statistics is often doubted, especially for developing countries (for discussion, see Balassa and Bauwens, 1987; Yeats, 1990; Yeats, 1999; Fukao et al., 2003; Gaulier and Zignago, 2010).

Internationally, the UN Comtrade database is the main source of trade statistics used by researchers. In line with this, our research obtains bilateral trade data from the UN Comtrade database to construct the commodity index weights. This is because the UN Comtrade database contains a rich data set of country-specific commodity data from all revisions of the Standard International Trade Classification (SITC) and Harmonized System (HS) trade classification systems for all countries in the world. Particularly, the accuracy of our index series would improve only if all the available data is considered in the construction process. Therefore, the data abundance of UN Comtrade is one of the main reasons why it is the preferred data source for our study as well as for several others in the economic literature (for example, see Grilli and Yang, 1988; Deaton and Miller, 1995; Cashin et al., 2004).

In addition, the UN Comtrade database reports the trade data with respect to the classification code provided in the revision of its corresponding trade classification system. More precisely, there are two most commonly used trade classification systems, SITC and HS, which have four and five revisions respectively. Usually, past studies in the literature obtain all the data from only one revision while ignoring the data availability in the others. However, the reported data obtained from different revisions of trade classification systems lacks consistency. In fact, the main difficulty is almost entirely caused due to the changing

[23] The database also includes countries that are not classified as "commodity-dependent" but may have a significant share in the world export of a particular commodity (group). For example, the US has substantial share in the world cereal market, i.e. wheat, soybean, corn (see Chambers and Just, 1981; Mitchell, 2008).

definitions and data availability. These issues may cause the incorrect construction of the index weights if the data is collected from only one revision of the trade classification system. As a result, this may lead to inaccurate national commodity price indexes. Along these lines, two main problems occur in the data collection process. The first one is within the particular trade classification system, while the other is between different systems.

One of the most popular trade classification systems in the world is the SITC. It is used as a source for the selection of trade statistics from UN Comtrade in the earlier studies of Sahay et al. (2002), Cashin et al. (2004), Pfaffenzeller et al. (2007) and Chen and Lee (2014). Unfortunately, the reported export value for the same commodity product, in a given year, may differ with respect to the revision of the SITC.

In other words, the SITC system is not synchronised along its various revisions, which entails that incorrectly reported values are recorded. Hence, the construction of the index weights with inaccurate data may create misleading test results in empirical studies. To highlight this, two examples of incorrect and misreported data in the SITC are presented in the following table:

Year	Rice (Saudi Arabia)									Crude petroleum (Syria)								
	SITC 1	SITC 2	SITC 3	SITC 4	HS1992	HS1996	HS2002	HS2007	HS2012	SITC 1	SITC 2	SITC 3	SITC 4	HS1992	HS1996	HS2002	HS2007	HS2012
1991	4508444	4508444	4508444		4508444													
1992	**576**	**2761465**	**2761465**		**2761465**					1850645376	1850645376							
1993	**492**	**2541966**	**2541966**		**2541966**													
1994		**716843**	**716843**		**716843**													
1995	**3345**	**3449437**	**3449437**		**3449437**					2205121280	2205121280							
1996	5711203	5711203	5711203		5711203					2540826880	2540826880							
1997										2173474304	2173474304							
1998	1825739	1825739	1825739		1825739						1378467200							
1999	2478607	2478607	2478607		2478607	2478607					2179996160							
2000	2370491	2370491	2370491		2370491	2370491				3203686912	3203686912	3203686912						
2001	1990332	1990332	1990332		1990332	1990332				3586002510	3586002510	3586002510		3586002510	3586002510			
2002	2681868	2681868	2681868		2681868	2681868	2681868			4243373385	4243373385	4243373385		4243373385	4243373385			
2003	2531989	2531989	2531989		2531989	2531989	2531989			3583553024	3583553024	3583553024		3583553024	3583553024			
2004	3261257	3261257	3261257		3261257	3261257	3261257			2936056822	2936056822	2936056822		2936056822	2936056822			
2005	7742381	7742381	7742381		7742381	7742381	7742381			3864305830	3864305830	3864305830		3864305830	3864305830	3864305830		
2006	9723464	9723464	9723464		9723464	9723464	9723464			3666558497	3666558497	3666558497		3666558497	3666558497	3666558497		
2007	17674616	17674616	17674616	17674616	17674616	17674616	17674616	17674616		3986018202	3986018202	3986018202		3986018202	3986018202	3986018202		
2008										4708793736	4708793736	4708793736	4708793736	4708793736	4708793736	4708793736	4708793736	
2009										2865468676	2865468676	2865468676	2865468676	2865468676	2865468676	2865468676	2865468676	
2010	7844276	7844276	7844276	7844276	7844276	7844276	7844276	7844276		4325838490	4325838490	4325838490	4325838490	4325838490	4325838490	4325838490	4325838490	
2011	12815216	12815216	12815216	12815216	12815216	12815216	12815216	12815216										
2012	12090415	12090415	12090415	12090415	12090415	12090415	12090415	12090415	12090415									
2013	14472285	14472285	14472285	14472285	14472285	14472285	14472285	14472285	14472285									
2014	10824975	10824975	10824975	10824975	10824975	10824975	10824975	10824975	10824975									
2015	9916279	9916279	9916279	9916279	9916279	9916279	9916279	9916279	9916279									

Note: The table presents the total value of exports reported by different revisions of the trade classification system for rice (exported from Saudi Arabia) and crude petroleum (exported from Syria) for a given year in US $. The source of the data is UN Comtrade (2018).

Table 2.2 Total Value of Exports for a Specific Commodity, for a Given Country, in US$

Table 2.2 presents the total value of exports reported by different revisions of the trade classification system for rice (exported from Saudi Arabia) and crude petroleum (exported from Syria) for a given year in US $. The UN Comtrade provides different exports value of Saudi Arabian rice in years 1992, 1993, 1994 and 1995 with respect to the SITC Rev. 1 and other revisions of the trade classification system. In fact, the total value of the exported rice from Saudi Arabia as per the SITC Rev. 1 is $576 in 1992, whereas the other revisions report a value of $2,761,465 for the same year. This example demonstrates the inconsistency in the reported trade data between different revisions within the same reporting system, namely the SITC.

Further, a similar conclusion can be reached when considering the Syrian export values of crude petroleum. At first, we emphasise that the crude petroleum contributes 88% of the total share of exports in Syria, as per Cashin et al. (2004). As such, crude petroleum is used in the construction of the index series for Syria in the studies of Sahay et al. (2002) and Cashin et al. (2004). More precisely, the two studies use the SITC Rev. 1 UN Comtrade data on Syrian export values of crude petroleum for the period of 1991–1999 in the construction process of the commodity index weights. However, as can be seen in table 2.2, the crude petroleum data reported for 1998 and 1999 in the SITC Rev. 1 is actually missing as compared to the SITC Rev. 2. Therefore, constructing an index series with incorrectly recorded export values for crude petroleum may have a significant adverse impact on the index weighting and, therefore, on the index accuracy. In general, there is no country wherein the trade data is immune to the above-mentioned issues. To emphasise this, we acknowledge that the possible inaccuracy in the previously created index series is not the author’s fault but is due to compilation issues with the international trade statistics.

Furthermore, another well-known trade classification system is the HS. Similar to SITC, the HS exhibits inconsistency in the reported values throughout its various revisions. An example is presented in the table below:

Year	Barley (Kazakhstan)									Wheat (Kazakhstan)								
	SITC 1	SITC 2	SITC 3	SITC 4	HS1992	HS1996	HS2002	HS2007	HS2012	SITC 1	SITC 2	SITC 3	SITC 4	HS1992	HS1996	HS2002	HS2007	HS2012
1995	58106	58106	58106		58106					228472	228472	228472		228472				
1996	93060	93060	93060		93060					311853	311853	311853		311853				
1997	73302	73302	73302		73302					430414	430414	430414		430414				
1998	**24130**	**24130**	**24130**		**24130**	**24612**				**256367**	**256367**	**256367**		**256367**	**258836**			
1999	32951	32951	32951		32951					267084	267084	267084		267084				
2000	40062	40062	40062		40062	40062				449737	449737	449737		449737	449737			
2001	19466	19466	19466		19466	19466				321071	321071	321071		321071	321071			
2002	19177	19177	19177		19177	19177				325139	325139	325139		325139	325139			
2003	37108	37108	37108		37108	37108				522568	522568	522568		522568	522568			
2004	28965	28965	28965		28965	28965	28965			389550	389550	389550		389550	389550	389550		
2005	11321	11321	11321		11321	11321	11321			219727	219727	219727		219727	219727	219727		
2006	39503	39503	39503		39503	39503	39503			522755	522755	522755		522755	522755	522755		
2007	111366	111366	111366		111366	111366	111366			1170507	1170507	1170507		1170507	1170507	1170507		
2008	156642	156642	156642		156642	156642	156642			1458780	1458780	1458780		1458780	1458780	1458780		
2009	39054	39054	39054	39054	39054	39054	39054	39054		632852	632852	632852	632852	632852	632852	632852	632852	
2010	51442	51442	51442	51442	51442	51442	51442	51442		911491	911491	911491	911491	911491	911491	911491	911491	
2011	111017	111017	111017	111017	111017	111017	111017	111017		609419	609419	609419	609419	609419	609419	609419	609419	
2012	76315	76315	76315	76315	76315	76315	76315	76315	76315	1599128	1599128	1599128	1599128	1599128	1599128	1599128	1599128	1599128
2013	60329	60329	60329	60329	60329	60329	60329	60329	60329	1253937	1253937	1253937	1253937	1253937	1253937	1253937	1253937	1253937
2014	142762	142762	142762	142762	142762	142762	142762	142762	142762	960072	960072	960072	960072	960072	960072	960072	960072	960072
2015	103559	103559	103559	103559	103559	103559	103559	103559	103559	1244415	1244415	1244415	1244415	1244415	1244415	1244415	1244415	1244415

Note: The table presents the total value of exports reported by different revisions of the trade classification system for barley (exported from Kazakhstan) and wheat (exported from Kazakhstan) for a given year in US$. Data prior to 1995 has not been reported in either revision. The source of the data is UN Comtrade (2018).

Table 2.3 Total Value of Exports for a Specific Commodity, for a Given Country, in US$ thousands

As can be seen from table 2.3, differences are found in the Kazakh export values of wheat and barley in 1998 with respect to the HS1996 and other revisions of the trade classification system. This provides support for our assumption that the HS is vulnerable to incorrectly reported data.

Nonetheless, another major problem in certain revisions of the above-mentioned trade classification systems is that the trade data is not recorded.[24] This can only be identified if the trade values for identical products are compared across different trade classification revisions (see tables 2.2 and 2.3). The comparison demonstrates that the SITC Rev. 1 is heavily affected by the issue of misreported export values, and other revisions are no exceptions. Therefore, the importance of a single commodity product in the calculation of the index weights may be under/overestimated if the aforementioned shortcomings are not taken into consideration.

Unfortunately, a majority of the current studies in the literature use a single revision of the trade classification systems to collect trade data rather than multiple revisions. This increases the possibility of constructing an inexact index series. Some examples of studies where single trade classification revision is considered are Sahay et al. (2002) – SITC Rev. 1, Cashin et al. (2004) – SITC Rev. 1, Pfaffenzeller et al. (2007) – SITC Rev. 2, Bodart et al. (2012) – HS1992 and Chen and Lee (2014) – SITC Rev. 2. Although all these studies use robust index number formulas for the construction of their commodity price indexes, the inclusion of inaccurate trade data in the process of index calculation may result in incorrectly tracking the price movements in the national commodity market. Yeats (1999, p.34) highlights that "Significant progress in updating the accuracy, and coverage, of trade statistics will require improved procedures for data collection and reporting at the country level". Therefore, our study proposes a new method for trade data collection that aims to reduce the influence of the aforementioned issues and improves the accuracy of the index series.

This study develops a novel approach for data collection that obliterates the dissimilarities in the reported data through various revisions of the HS and SITC trade classification systems in the UN Comtrade and UNCTAD databases.[25] This is required because the international trade statistics suffer from inconsistencies in the reported data, as illustrated in tables 2.2 and 2.3. Therefore, this study provides a rigorous procedure for data collection that aims to reduce the possibility of data inaccuracy affecting the index movements in our database. The procedure is briefly explained below:

[24] The discussion is given when trade actually occurs; however, no trade data is recorded for the particular product. Missing data due to trade barriers or other economic reasons is not suspected as an issue in this study.
[25] UNCTAD produces more than 150 indicators and statistical time series on international trade and commodities for varied groups of countries and territories. It uses the SITC classification system to record its trade data.

First, a common correspondence table for the HS and SITC conversion is constructed. In other words, each commodity in our sample of 72 commodities has been allocated a trade classification code from HS1992, HS1996, HS2002, HS2007, HS2012, SITC1, SITC2, SITC3 and SITC4. It is found that certain series have altered commodity codes; thus, the first task is to allocate the commodities under their new codes. The correlation tables are used, as provided by UN Trade Statistics (UNSTATS, 2017). Then, each commodity is allocated to a code that best describes its characteristics. This is important for cross-checking the precision of the reported trade values among different revisions of trade classification systems. If an inconsistency is identified in the values of reported trade data, the procedure discussed below is applied.

Meanwhile, the primary difficulty is almost entirely caused by changing the definitions (codes) in the revisions of trade classification systems and data availability, as highlighted by McCracken and Ng (2016). As an example, wheat is reported by a 3-digit code in the SITC Rev. 1, i.e. “041”, whereas HS1992 reports wheat using a 4-digit code, i.e. “1001”. Another example is the trade data for soya bean oil. It is recorded under three different codes in the SITC, i.e. SITC Rev. 1 – “4212”, SITC Rev. 2 – “4232”, SITC Rev. 3 – “4211”. In the same manner, the data for soya bean oil is reported in the HS by the code “1507”. As evident, the SITC and HS codes for the same commodity product are not identical across different revisions. Therefore, if one has to collect 40 years of trade data, one cannot avoid splicing the data from different trade classification systems. This makes the data collection process time-consuming.

Second, a common correspondence table, as discussed above, is assigned for each country in our sample. In the case where the code is not available due to the preference of one country to report through a certain revision of trade classification system over the other, the non-zero empty value is assigned. This is important as the non-zero empty value represents no information or missing data, whereas the value of zero entails no export for a given commodity. Neglecting this may influence the structure of commodity weights and lead to incorrect index movements. Therefore, this study distinguishes between the “non-zero empty” and “zero” export values reported through the trade classification systems in both UN Comtrade and UNCTAD databases. Due to a lack of information, we are not aware whether previous studies consider this aspect when constructing their index series.

Third, the commodity codes in each country’s common correspondence table are assigned with relative export trade values from the UN Comtrade and UNCTAD databases. We gather trade export data for each commodity and for all the countries included in our database. The data from each revision of the HS and SITC is recorded separately at first. The separation of the initial trade data is crucial for the identification of any existing inconsistencies in the reported data across different revisions of the trade classification systems.

Fourth, we examine whether the export trade values of the same commodity product are identical among different revisions of trade classification systems. This is done with the aim of checking for data discrepancy. In fact, this study improves upon the past literature, which uses trade data as it appears in a single revision classification trade system in the UN Comtrade (for example, see Grilli and Yang, 1988; Deaton and Miller, 1995; Dehn, 2000; Cashin et al., 2004; Brückner and Ciccone, 2010; Bodart et al., 2012; Ciccone, 2018). In particular, we consider four main assumptions with the objective of reducing the impact of incorrect and misreported data throughout the overall movements of our index series.

Assumption 1: When the trade data for a given commodity is identical throughout all revisions of the HS and SITC, the reported value is accepted as the true value.

This first assumption provides us with a primary piece of information. Analytically, it indicates whether the export trade values reported in the UN Comtrade are the country's true exports. This is the case when they are identical throughout all revisions of the HS and SITC. For instance, table 2.3 shows that the export value of barley for Kazakhstan in 1995 is \$58,106,486 in all trade classification systems. Hence, \$58,106,486 is accepted as the true export value of barley in 1995 for Kazakhstan. Then, this value is used in the construction of the index commodity weights.

Assumption 2: If the trade value for a given commodity is missing in (at least) one revision of the trade classification systems but appears in the others, it is recorded as given in the others.

The second assumption is a consequence of Assumption 1, but it differs in the sense that there exists a revision of the trade classification system where data for a given commodity is not recorded. In other words, a revision of the trade classification system may have a missing data value. For example, table 2.2 displays oil exports for Syria in 1997 in both the SITC Rev. 1 and the SITC Rev. 2, whereas Syrian oil exports in years 1998 and 1999 are available in the SITC Rev. 2 but not in the SITC Rev. 1. As such, one may construct a commodity price index for Syria using the SITC Rev. 1 and, therefore, obtain inaccurate commodity index weights. None of the revisions of trade classification systems are immune against this issue. Hence, Assumption 2 is vital to ensure the quality of the data collection process.

Assumption 3: When the export value reported through one of the trade classification revisions is different from the others, the highest trade value is assumed to be the true representative of the country's export for that specific commodity.

Assumption 3 is imposed in order to ensure that the recorded figure of the country's exports includes both exports and re-exports. This is in line with the definition provided by the United Nations Statistics Division, Department of Economic and Social Affairs, which states that re-exports are to be included in the country exports (UNSD, 2011). In particular, most of the export values reported through the UN Comtrade include re-exports. Here, we say "most, not all" because there are few exceptions where the country's export does not include re-exports. An example of this is the rice exports of Saudi Arabia in the years 1992–1995 as reported in the SITC Rev. 1 (see table 2.2). The reported values exclude the country's re-exports for rice. One can easily verify this if one executes Assertions 1–3 and makes a comparison with the other revisions of trade classification systems. Therefore, if Assumption 3 is ignored, one may end up using exports value data wherein the re-exports are included for some commodities while being excluded for the others. In our case, this may cause distortions in the index weights construction. Therefore, one should check whether Assertion 4 holds. If it does not, then the higher value is incorrectly recorded and the lower value should be accepted.

Assumption 4: In a special case of only two reported values for a given commodity, among all revisions of the trade classification systems, the higher value is considered as the true one.

Assumption 4 is a consequence of Assumption 3 and Assertion 1 (see below Assertion 1). It assures that the re-export trade value is included in the country exports. In addition, it is important to check whether this is in line with Assertion 4 as well. If Assertion 4 fails, the lower value has to be accepted in place of the higher value.

Complimentary to the assumptions above, the following assertions are made:

Assertion 1: The commodity exports value that includes non-zero re-exports should be larger than the one that does not include re-exports.

Assertion 2: The commodity exports value does not include re-exports if it is lower than the value of re-exports.

Assertion 3: The reported commodity exports value includes both the country's exports and re-exports only if the sum of the country's exports plus re-exports is equal to the reported commodity exports value.

Assertion 4: The export data for a given commodity that is provided in a high-digit level of detail, e.g. "1234", should be less than that of one provided in a low-digit level of detail, e.g. "12". For example, the wheat export value should be lower than the export value of cereals for the same year and country.

All of the above assumptions and assertions are made in line with this statement: "Re-exports are to be included in the country exports. They are also recommended to be recorded separately for analytical purposes" (UNSD, 1998, p. 28). The requirement for this statement stems from the fact that certain trade classification systems include the re-exports values in the country export, whereas some exclude them. Further, the commodity trade value series from a single revision may exclude the re-exports from the country exports for one year and include them in another.[26] Therefore, the trade data is not consistent in terms of the *definition* of whether country exports include re-exports across all revisions, which may lead to severe distortions in the index movements.[27]

Therefore, we aim to use a *definition* of trade export values that is consistent throughout products, countries, years and trade classification systems. As such, we follow the *definition* provided by the UNSD (1998) by combining the country's re-exports and exports while constructing our database.[28] For this reason, Assumptions 1–4 and Assertions 1–5 should all hold and prevent any inconsistencies in the data collection process.

Fifth, after the whole procedure is completed, the final step is to "remove the revisions" by combining all trade data in one common data set. In fact, this data set is in the foundation of the index weights construction in our database.

In summary, "It is difficult if not impossible to automate the "data collection" process because judgment is involved" (McCracken and Ng, 2016, p. 4). This study is the first in the literature that emphasises the dissimilarities between the different revisions of the HS and SITC trade classifications systems and, subsequently, their adverse impact on index construction. This issue should not be neglected, as it might lead to fallacious movements in the national commodity index series. Our study questions the accuracy of the data reported through the

[26] The same assumption can be made for the imports and re-imports.
[27] For instance, one country may not be a producer of a particular good (commodity); however, the data obtained from the UN Comtrade may suggest that the country exports it – rice from Saudi Arabia, for example. Further, such data may be included in the data set of empirical studies on the bilateral trade of the researched country(s). For more information, see the exports and re-exports values for rice from Saudi Arabia to the World in the SITC Rev. 1 and other trade classification systems in the UN Comtrade.
[28] One may wish to exclude the re-exports from the country exports when calculating the national commodity price index. If done, one can examine whether there is a change in the movements of the index series. This is out of the scope of this study, so we leave it for future research.

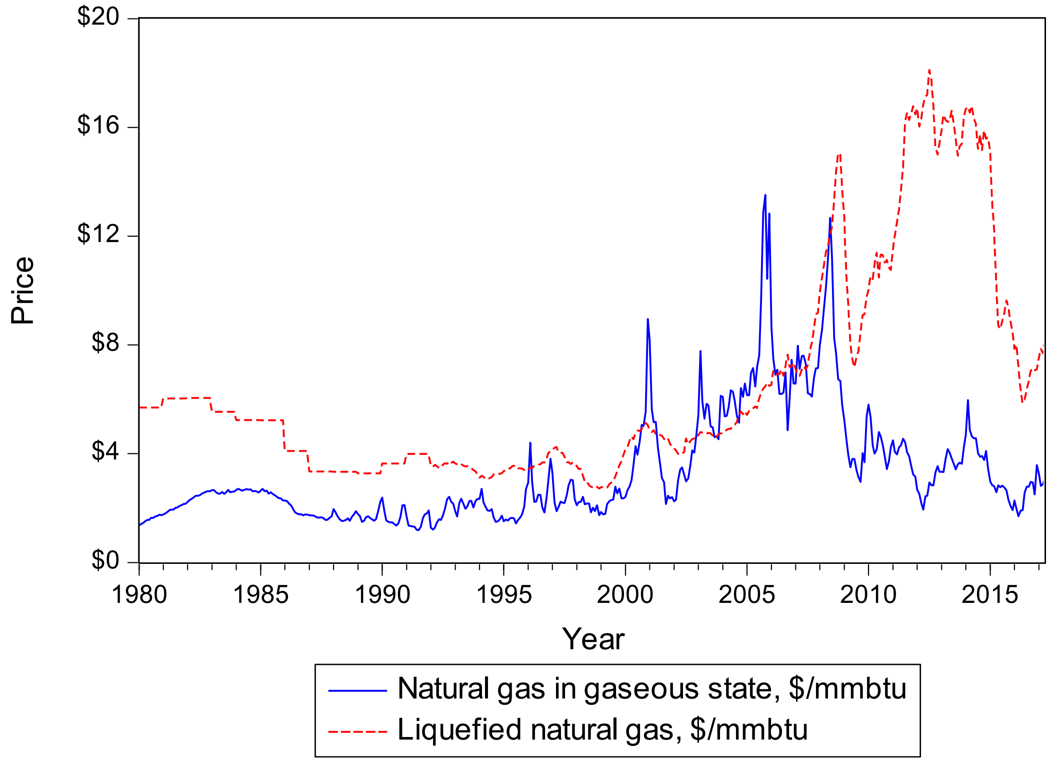
UN Comtrade and UNCTAD databases and proposes a method for mitigating this issue, as illustrated above.

2.3.2 Disaggregation

As clarified in the above section, the data collection process is a vital element in constructing a reliable database of national price indexes of commodity exports. Another important factor for having an index that closely tracks the price movements in the national commodity markets is the use of a finer level of disaggregation data in its construction (Isard, 1977). This is essential because some commodities are heterogeneous goods. In other words, prices of different commodities do not tend to move in parallel and have different end uses as well (Cashin et al., 1999). Although fluctuations in world demand impart common components of several price series, supply conditions differ across goods and relative prices are far from constant (Deaton, 1999). Consequently, the disaggregation of the commodity sector into more detailed products aims to improve the explanatory power of our index series.

Particularly, this study uses further disaggregated levels for data on the following commodities: natural gas, wool, timber and milk powder. These commodities are the main exporting products in the energy, raw material and dairy sectors respectively.

For example, natural gas is an energy product that is primarily traded in two states: gaseous and liquefied. Due to these product specifications, there are two main types of natural gas products: natural gas in a gaseous state and liquefied natural gas (Yergin and Stoppard, 2003). As such, the prices of these two commodities are not identical in the international commodity market, as illustrated by the following diagram:



Source: World Bank (2018)

Figure 2.1 International Prices of Liquefied and Gaseous State Natural Gas

The prices of natural gas in different states, i.e. gaseous and liquefied, significantly diverge from one another in the world commodity market (see figure 2.1). Therefore, neglecting this fact during the process of index calculation may result in an imprecise index series.

Unfortunately, the majority of past studies either use a single price for natural gas when constructing their national commodity price indexes (for example, Bodart et al., 2012) or simply do not include natural gas in the index commodity baskets (for example, Deaton and Miller, 1995; Dehn, 2000). Dehn (2000, p. 36) states, "A few important commodities have not been included in the index due to lack of adequate data. These are natural gas and uranium ore." This study builds on the important contribution of Dehn (2000) by including natural gas and uranium ore as well as numerous other commodities in the process of our database construction.

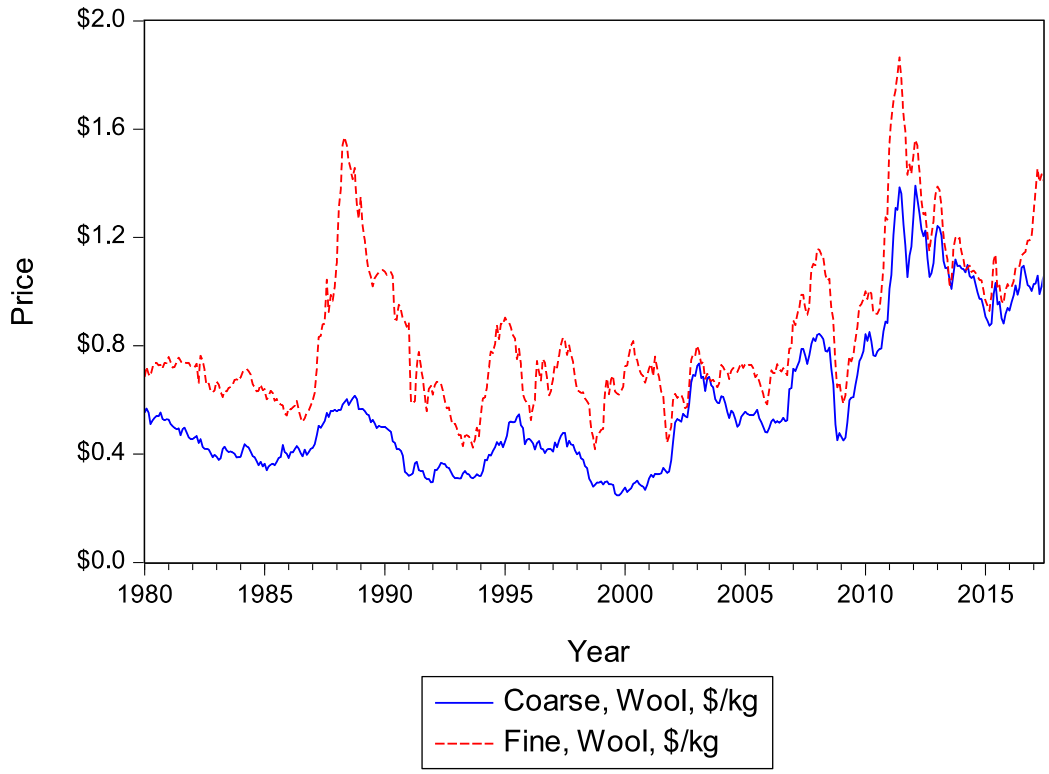
Another key point is that the data for commodity trade export values has been rarely disaggregated to a level that best explains the commodity characteristics. Our study addresses this, which importance is revealed in the following features. On the one hand, if a country is an exporter of liquefied natural gas but trade data for gaseous state natural gas is used in the process of index construction, the impact of the natural gas on the price movements of the national commodity index may be underestimated. On the other hand, if one obtains an export value for liquefied natural gas but uses the price of the natural gas in gaseous state while calculating the commodity price index, one may end up with an inexact national index series of commodity exports. Therefore, our study overcomes this issue by employing disaggregate data for both commodity prices and trade values, if the data allows it.

An example of this can be noted in the study conducted by Sahay et al. (2002), who used natural gas in the construction of national commodity export indexes for four countries, namely Indonesia, Mexico, Norway and Syria. They overlooked the heterogeneous behaviour of the prices of natural gas, liquefied and in the gaseous state, and only used data for natural gas in the gaseous state to construct their index series.[29] However, Norway mostly exports natural gas in the gaseous state, while Indonesia mainly exports liquefied natural gas (Reymond, 2007). Therefore, the national commodity index for Indonesia is likely to be inaccurately constructed because the prices of these two types of natural gas do not tend to move together in the world commodity market (see figure 2.1). This issue may occur for any other commodity product in the commodity market. We mitigate this problem by using disaggregate level data of commodity export values in the construction process of our database. Nonetheless, the usage of disaggregation level data for each commodity product is highly dependent on data availability.

This chapter acknowledges the gaps in the past studies and provides a two-step procedure for index calculation. First, we assure that the data obtained for export values is correct. That is to say, it should be verified whether country exports liquefied or gaseous the natural gas. Then, we collect data only for the specific type of gas but not as a whole. Second, the correct price of natural gas is assigned with respect to whether the country exports either gaseous or liquefied natural gas or both. For consistency reasons, this procedure is applied for all commodity products in our database, if the data allows it.

In addition, we provide another example of a commodity that is commonly used in the construction of national commodity price indexes – wool (see Deaton and Miller, 1995; Dehn, 2000; Sahay et al., 2002; Cashin et al., 2004). In fact, wool has two major sub-products: fine wool and coarse wool (see Angel et al., 1990 for discussion). The prices of fine wool and coarse wool in the international market are:

[29] Our study is unable to comment on whether Sahay et al. (2002) used the correct data on natural gas export values because such information is not provided in their article. However, the authors used the SITC Rev. 1 for obtaining data on commodity export values, where no information is available on the export of liquefied natural gas. Therefore, we assume that the authors do not distinguish between the different states of natural gas, i.e. either gaseous or liquefied, when constructing national commodity export indexes.



Source: IMF (2018)

Figure 2.2 International Prices of Wool, Coarse and Fine

Figure 2.2 shows that the prices of the sub-products of wool diverge from each other in the international market. This can be explained by the differences in the physical attributes of wool, which affect its spinning characteristics and suitability for different end uses (Angel et al., 1990). For example, exporters of fine and coarse wools are Australia and New Zealand respectively. On the one hand, Australia is a leading exporter of fine wools, with around 75% of their wool typically being 23 microns or finer. On the other hand, New Zealand is a leading exporter of coarse wool, with around 75.6% of their wool typically being 33 microns or coarser (Angel et al., 1990). Due quality differences, we can conclude that fine and coarse wools are heterogeneous products, the specific features of which reflect on their prices in the world wool market. Unfortunately, several past studies in the literature neglect this fact (for example, see Deaton and Miller, 1995; Dehn, 2000; Sahay et al., 2002). This may result in the calculation of an imprecise index series of national commodity export.

To resolve this issue, this chapter perceives the heterogeneous behaviour of commodity prices in the world market by allocating to each commodity product a trade classification code that most closely explains the commodity's characteristics. Once selected, the product classification codes are held constant for all countries. In the case of data unavailability, the low-digit trade classification codes are not considered because this may disregard the heterogeneity assumption of the commodity products.[30] With this in mind, the disaggregation of the commodity sector into more detailed products aims to improve the explanatory power of our index series.

Overall, our study uses disaggregate level data for the construction of national price indexes of commodity exports. This has been done by weighing each price series with its relevant trade data wherever data is available. If the relevant data is not available, the low-aggregation export value data is not considered, and the commodity is excluded from the index basket. This prevents possible distortions in the price movements of our national commodity export indexes.

2.3.3 Number of commodities

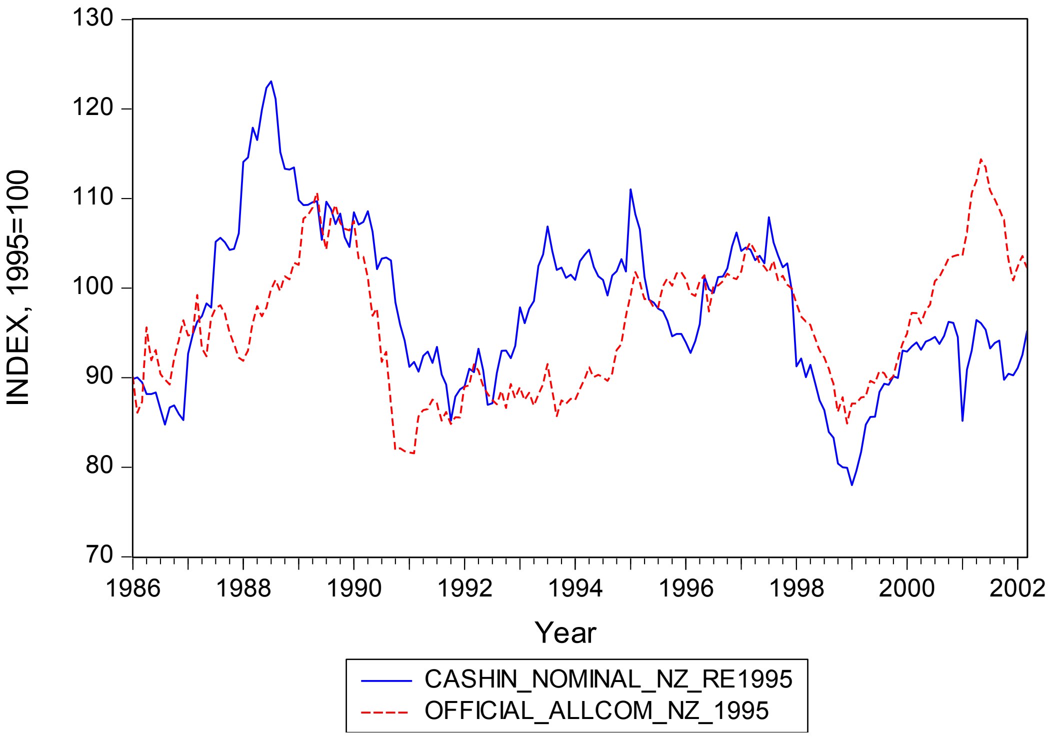
The disaggregation of the commodity data allows more primary products to be included in the index basket of each national commodity export index. In other words, as the number of commodities increases, the index has a more precise explanation of the price movements in the country's commodity sector as well as the possibility of missing a major exporting product from the country's index commodity basket is reduced. Therefore, this study considers a sample of 72 primary commodities when constructing our database.

Until this point, there is no study in the literature that has used such a large sample of commodity products in the calculation of national commodity export indexes. For example, Deaton and Miller (1995) focus on the sub-Saharan African countries and consider only 21 commodities in the construction process of their national commodity export indexes. This is about 70% less than the number of commodities in our study. More precisely, Deaton and Miller (1995) do not consider the following commodities in the index composition: wheat, barley, maize, rice, sorghum, beef, lamb, swine meat, poultry, natural gas, and others. In addition, they do not include timber in the construction of their index series. According to Wood and Mayer (2001), timber has a large share in the export baskets of certain sub-Saharan African countries, such as Cameroon (rough wood and sawn wood: 19% of the total export in 2017), Central African Republic (rough wood and sawn wood: 55% of the total export in 2017) and Gambia (rough wood: 51% of the total export in 2017). Therefore, the exclusion of timber from the commodity index baskets of some sub-Saharan African economies may result

[30] For instance, the HS1992 code for wheat is "1001". This code is used in the process of collection of wheat trade data for all countries in the sample. Let us assume that no trade data has been allocated to a country under HS1992 code "1001", but the data is available under HS1992 code "10", i.e. "cereals". As such, the index has been constructed by weighing the wheat prices with the values from the two-digit code instead of the four-digit one. However, the suggested two-digit code contains trade data on wheat, maize, rice, barley and other cereals. Also, the prices for these products, such as wheat and rice, are not identical in the world market. Therefore, the wheat product is overweighed, and incorrect movements are expected in the price index.

in the imprecise movements of their national commodity export indexes. As such, we consider a large number of commodity products in the construction process of our database (including timber), which aims to increase the accuracy of our index series.

Another example is the study of Cashin et al. (2004), who extended the DM database with respect to both the number of countries and the number of commodities. Particularly, the authors use 44 commodities in the calculation of their non-fuel national commodity export indexes. Further, Cashin et al. (2004) construct a database for 58 countries without imposing any geographical restriction on the sample of countries; for example, the DM database only focuses on sub-Saharan African countries. However, Cashin et al. (2004) disregard some important commodities from the index commodity baskets of their index series. An example is dairy products, which are not included in the calculation process of their indexes. The exclusion of dairy products from the index commodity basket has a large and adverse impact on the accuracy of the index series for dairy exporters such as New Zealand. Indeed, Sahay et al. (2002, p. 53) noted that their "constructed and bank indexes for New Zealand differ somewhat, due to the exclusion of dairy products from the constructed index.".[31] To illustrate that, we present the index series of Sahay et al. (2002) for New Zealand (i.e. Cashin_official_NZ_re1995) and the official index series of ANZ Bank for New Zealand (i.e. Official_all_NZ_1995) on the following graph:



Source: Cashin et al. (2004) and Australia and New Zealand Banking Group (2018)

[31] As noted previously, the data sets of Sahay et al. (2002) and Cashin et al. (2004) are identical.

Figure 2.3 Country-Specific Price Index of Commodity Exports for New Zealand

Figure 2.3 shows a substantial divergence in the index movements between the official series and those constructed by Sahay et al. (2002). This finding highlights the importance of the dairy products for the construction of precise national commodity export indexes. With that in mind, this study contributes to the existing literature by being the first (as far as we know) to include dairy products in the construction process of a world database of country-specific price indexes of commodity exports.

Moreover, the importance of having a large basket of commodity products should not be belittled when one creates a world database of national commodity export indexes.[32] This is because the structure of primary commodity exports is not homogenous across countries, as discussed by Hoekman and Djankov (1997). Therefore, the creation of a world commodity database should consider as many commodities as the data allows.

Although, one may say that the calculation of an accurate national commodity price index is possible even with a small number of commodities. However, the commodities should be pre-selected in advance and should be the top exporting products for each individual country. An example can be given with the commodity price indexes that are constructed by central and commercial banks. In particular, there are only three countries in the world that have regularly updated national commodity export indexes:

- *Canada*: Bank of Canada commodity price index (BCPI) and its relevant sub-indexes; source: Bank of Canada (2018); 26 commodities in the index basket (as of April 2017)
- *Australia*: Reserve Bank of Australia (RBA) Index of Commodity Prices and its relevant sub-indexes; source: Reserve Bank of Australia (2018): 22 commodities in the index basket (as of April 2017)
- *New Zealand*: ANZ Commodity Price Index and its relevant sub-indexes; source: Australia and New Zealand Banking Group (2018): 17 commodities in the index basket (as of April 2017)

As evident, the creation of an accurate national index of commodity exports does not require numerous commodities in the index basket. However, the commodity products for each national index have to be pre-selected. This is a time-consuming process that can be easily overwhelmed by using a large index basket of commodities. Therefore, we use a common

[32] For example, Cashin et al. (2004) point out that inadequate price series for barley, hardwood sawn, olive oil, poultry and swine meat precludes them from including these commodities in the final construction of their index series. We fill this gap in the study of Cashin et al. (2004) by including all these commodities within our index series.

large commodity basket for constructing the index series in our database. Further, our study uses an index number formula that automatically excludes those commodity products that are not a part of the country's export basket. Another advantage of using a large basket of commodity products is that it allows for constructing national commodity export sub-indexes. The sub-indexes are an important part of several studies in the past literature. For example, the study by Cashin et al. (2004) uses non-fuel national commodity export indexes to conduct their analysis.

In summary, one can assume that there is a positive relationship between the number of commodities and the number of countries. In other words, an increase in the number of commodities is a sufficient condition for creating a database with large N countries and numerous T time observations. This is particularly true due to the diversification of the trade exports around the world (Massell, 1970). In brief, this study considers the sample of 72 primary commodity products in order to complete the largest possible database of country-specific price indexes of commodity exports.

2.3.4 Sub-indexes

Some of the most well-known databases of commodity price sub-indexes are those constructed by Grilli and Yang (1988), the IMF, the World Bank, the United Nations and the CRB. However, none of these databases provide country-specific sub-indexes of commodity exports. That is to say, these studies and organisations are likely to provide commodity price sub-indexes that poorly represent the true price movements in the national commodity export markets (Deaton, 1999). Yet, the only countries (as far as we know) for which there are regularly updated databases of national commodity price sub-indexes are Canada, Australia and New Zealand. This is not a truly representative sample of all commodity-dependent economies. Therefore, our study contributes to the literature by constructing 13 categories of sub-indexes for a sample of 217 countries. Notably, we apply an index construction process that is identical for all index series in our database.

To emphasise, the creation of a world database of national commodity price sub-indexes makes an important contribution to the literature. On the one hand, these indexes provide information about the industry (sector) segmentation in a given country. This allows researchers and policymakers to perceive the performance of a given sector or its impact on the country's overall economy[33]. For example, the officials can use the potential predictive ability of national commodity export indexes to undertake on time the necessary reforms or

[33] For example, Gilbert et al. (2013) show that the agricultural exports have mixed effect on economic growth in Cameroon. Coffee and banana exports have a positive impact on economic growth, while cocoa export has a negative impact on economic growth.

actions for developing or protecting a particular sector in the economy of the country. On the other hand, Cashin et al. (2004) note that the inclusion of oil prices in the construction of the national commodity price index can render it an endogenous variable. A possible reason is that oil prices are determined by the oil production, and the oil production is sometimes driven by the OPEC's productions cuts (Filis et al., 2011) or by political events in the OPEC countries (Kilian, 2009).[34] Therefore, oil prices can be seen as a partially endogenous variable, and certain studies may want to exclude them from the empirical analysis – Cashin et al. (2004) for example. The same conclusion may be valid for other commodity products as well. Therefore, one may wish to exclude one or the other commodity product from the country's index basket. Unfortunately, this is a time-consuming process, especially when the number of countries is large. This study overwhelms this process by providing the literature with a world database that contains 13 categories of national commodity export (sub-)indexes for a sample of 217 countries.

In particular, we construct new sub-indexes for each country in our sample for the following categories: (1) All commodities, (2) Non-energy commodities, (3) Food, (4) Cereals, (5) Vegetable oils and protein meals, (6) Meat, (7) Dairy, (8) Beverages, (9) Agricultural raw materials, (10) Metals, (11) Energy, (12) Fertilizers and (13) Precious Metals. We follow the commodity grouping that is used by the IMF when constructing our sub-index series. Any decision to deviate from the IMF's classification is identified as follows:

(1) All commodities (COMPI): Wheat, Maize, Rice, Barley, Sorghum, Soybeans, Soybean meal, Soybean oil, Palm oil, Fish meal, Sunflower Oil, Olive oil, Groundnuts, Rapeseed oil, Coconut oil, Copra, Palm kernel oil, Cottonseed oil, Groundnut (peanut) oil, Linseed oil, Canola, Beef, Lamb, Swine Meat, Poultry, Butter, Cheese, Skim milk powder, Whole milk powder, Salmon, Shrimp, Sugar, Bananas, Orange, Pepper, Coffee, Cocoa Beans, Tea, Sawnwood (Hardwood), Logs (Hardwood), Logs (Softwood), Sawnwood (Softwood), Plywood, Pulp, Cotton, Wool (Fine), Wool (Coarse), Rubber, Hides, Tobacco, Jute, Sisal, Copper, Aluminium, Iron Ore, Tin, Nickel, Zinc, Lead, Uranium, Crude oil, Natural Gas (Gaseous state), Natural Gas (Liquefied), Coal, Phosphate rock, Potash, DAP, TSP, UREA, Gold, Silver, Platinum

(2) Non-energy commodities: Wheat, Maize, Rice, Barley, Sorghum, Soybeans, Soybean meal, Soybean oil, Palm oil, Fish meal, Sunflower Oil, Olive oil, Groundnuts, Rapeseed oil,

[34] The Organization of the Petroleum Exporting Countries (OPEC) is an intergovernmental organisation of 14 nations as of May 2017. As of 2016, the 14 countries accounted for about 44% of the world oil production. Therefore, we can assume that the OPEC has a major influence on global oil prices. See also Golub (1983), Griffin (1985), Backus and Crucini (2000).

Coconut oil, Copra, Palm kernel oil, Cottonseed oil, Groundnut (peanut) oil, Linseed oil, Canola, Beef, Lamb, Swine Meat, Poultry, Butter, Cheese, Skim milk powder, Whole milk powder, Salmon, Shrimp, Sugar, Bananas, Orange, Pepper, Coffee, Cocoa Beans, Tea, Sawnwood (Hardwood), Logs (Hardwood), Logs (Softwood), Sawnwood (Softwood), Plywood, Pulp, Cotton, Wool (Fine), Wool (Coarse), Rubber, Hides, Tobacco, Jute, Sisal, Copper, Aluminium, Iron Ore, Tin, Nickel, Zinc, Lead, Uranium, Phosphate rock, Potash, DAP, TSP, UREA, Gold, Silver, Platinum

(3) Food: Wheat, Maize, Rice, Barley, Sorghum, Soybeans, Soybean meal, Soybean oil, Palm oil, Fish meal, Sunflower Oil, Olive oil, Groundnuts, Rapeseed oil, Coconut oil, Copra, Palm kernel oil, Cottonseed oil, Groundnut (peanut) oil, Linseed oil, Canola, Beef, Lamb, Swine Meat, Poultry, Butter, Cheese, Skim milk powder, Whole milk powder, Salmon, Shrimp, Sugar, Bananas, Orange, Pepper

(5) Vegetable oils and protein meals: Soybeans, Soybean meal, Soybean oil, Palm oil, Fish meal, Sunflower Oil, Olive oil, Groundnuts, Rapeseed oil, Coconut oil, Copra, Palm kernel oil, Cottonseed oil, Groundnut (peanut) oil, Linseed oil, Canola

(6) Meat: Beef, Lamb, Swine Meat, Poultry

(7) Dairy: Butter, Cheese, Skim milk powder, Whole milk powder

(8) Beverages: Coffee, Cocoa Beans, Tea

(9) Agricultural raw materials: Sawnwood (Hardwood), Logs (Hardwood), Logs (Softwood), Sawnwood (Softwood), Plywood, Pulp, Cotton, Wool (Fine), Wool (Coarse), Rubber, Hides, Tobacco, Jute, Sisal

(10) Metals: Copper, Aluminium, Iron Ore, Tin, Nickel, Zinc, Lead, Uranium

(11) Energy: Crude oil, Natural Gas (Gaseous state), Natural Gas (Liquefied), Coal

(12) Fertilizers: Phosphate rock, Potash, DAP, TSP, UREA

(13) Precious Metals: Gold, Silver, Platinum

The above national commodity export price sub-indexes are available for each country in our sample. However, if a country is not an exporter of either of the commodity products in the sub-index commodity basket, the sub-index is not constructed. As such, our study aims to create a *data-rich environment* that will help researchers enhance their understanding of how cross-country and within-country commodity markets operate. In other words, our database has an N-rich environment, where N is the number of the countries. However, what about the T dimension, where T is the number of time series observations in each index series?

2.3.5 Data frequency

The commodity prices are known to be highly volatile, especially in recent years, which increases the importance of using a high-frequency data in the economic analysis (Cavalcanti et al., 2015). Having a high-frequency index series is important as it provides a more precise explanation of the price fluctuations in the countries' commodity markets. To capture these trends, this study constructs monthly national commodity export indexes for all countries in our sample. In fact, the construction of a higher frequency index series is not conceivable due to data limitation.

In particular, long-lasting commodity price data is limited for frequencies higher than monthly. In other words, the availability of daily or weekly commodity price data for all of our 72 commodity products is either restricted to a recent short-time period or unavailable. The former implies that the time span of the index series should be shortened, while the latter suggests that the number of commodities in the database has to be reduced. Based on these facts, the data frequency of our database is chosen to be monthly.

Furthermore, Deaton and Miller (1995) and Dehn (2000) construct their data sets with annual and quarterly frequencies respectively. In contrast, this study improves on them by constructing monthly frequency series that are better at capturing the fluctuations in national commodity markets. In addition, our study is consistent with recent studies from the literature, where the highest frequency national commodity export indexes are monthly, such as Sahay et al. (2002), Cashin et al. (2004) and Bodart et al. (2012).[35] A weakness of these data sets is that they focus only on a particular commodity group (for example, Sahay et al. (2002) and Cashin et al. (2004): *non-fuel commodities*) or country group (for example, Bodart et al. (2012): *developing and emerging countries*). Our database improves on the past literature by providing a monthly world database without imposing restrictions on either the commodity or the country groups.

2.3.6 Index formula

The central and commercial banks have different preferences regarding the selection of the index number formula for the construction of their national commodity price indexes. For example, Bank of Canada uses the chain Fisher index, while Reserve Bank of Australia and ANZ Bank use the chain Laspeyres index. In addition, Bank of Canada updates the weights of

[35] An exception is the national commodity price index constructed by Bank of Canada. The index is available in a weekly/monthly frequency only for Canada (see table 2.1).

their index irregularly, whereas Reserve Bank of Australia does so annually. This methodological dissimilarity in the process of index construction may reflect upon the reliability of the index series for cross-country analysis. In fact, we acknowledge two main differences among the methods implemented by the central and commercial banks: (1) in the index weighting and (2) in the index number formula.

First, the choice of index weighting has always been a topic of debate in the literature. In particular, there are two main statistical methods for the allocation of index weights: chain-linking and fixed-base. These methods can be applied to all the aforementioned index formulas as well as to the one proposed by Deaton and Miller (1995).

In fact, chain-linking requires regular (in most cases, annual) updating of the index weights. This makes it an unsuitable method for compiling large databases of country-specific commodity export price indexes due to the scarcity of regularly updated data on national export values, especially for developing countries. More importantly, the country-specific commodity export values are rarely available for the period before the early 1990s. This may result in a discontinuity of our index series.

The fixed-base method provides a solution for this limitation. In particular, it allows the base period to be arbitrarily selected with respect to the availability of national level data. Further, the fixed-base method allows for an appropriate choice of a base period length. As such, the index series are able to reflect the seasonal behaviour of the prices of certain commodity products (for example, wheat and barley), whereas a month-to-month chain index cannot. For this reason, the fixed-base method has been chosen over the chain-linking for constructing our index weights.

Second, the index formula selected for the construction of our database of national commodity export indexes is the one proposed by Deaton and Miller (1995). In fact, it is chosen over the other index formulas that are currently used in the literature. One of the most commonly used index formulas is the Laspeyres index, which significantly overestimates the "true" price change (Braithwait, 1980).[36] An alternative formula that overcomes this issue is the Fisher formula (see Fisher, 1922; Diewert, 1976; Diewert, 1998; Hill, 2004). However, the Fisher index is impossible to carry out when N and T are both large due to the difficulties in obtaining current-period volume data of the country's commodity exports. In particular,

[36] Similarly, the Paasche index underestimates the "true" price changes (Diewert, 1998). Specifically, the time reversal test is not satisfied by Laspeyres and Paasche index numbers (Samuelson and Swamy, 1974). The test states that the price index number for time period $t(n)$ relative to time period $t(0)$ is the reciprocal of the price index number for time period $t(0)$ relative to time period $t(n)$, where $t(i)$ is a continuous time period such as $i \epsilon \{0,1,2, \dots n \dots\}$ (Balk, 1995).

such data is rarely available in the international trade statistics, especially in monthly frequency, rendering the Fisher index impracticable for the construction of monthly world databases of national commodity export price indexes (see Persons, 1921 for discussion). Thus, our study uses the DM index instead of the Fisher index because the former does not require the availability of volume data for its calculation. Moreover, the DM index is handier in practice than the Fisher index, and it does not overestimate the "true" price change like the Laspeyres index.

Nonetheless, we acknowledge that a potential weakness of the DM index is that it does not allow for updating the commodity index weights. As a result, the national commodity price index does not cope well with shifts in the structure of trade. For example, the index is not able to capture resource discoveries and other quantity shocks, as highlighted by Dehn (2000). With this in mind, the empirical part of our study compares our index series to those constructed by central and commercial banks. Even though the index series provided by central and commercial banks account for the quantity changes, their index movements are found to oscillate together with our constructed indexes (see Section 2.6.1). These results provide some assurance regarding the accuracy of our national commodity export price indexes.

In addition, using fixed-base weights in the construction of commodity price indexes is a common practice for the IMF, The Economist, the World Bank, Grilli and Yang (1988), Deaton and Miller (1995), Deaton (1999), Sahay et al. (2002), Cashin et al. (2004), Bodart et al. (2012), Bodart et al. (2015), and numerous others. This further validates our choice of using the fixed-base DM index for constructing our database.

In summary, the creation of a *data-rich environment* of national commodity price indexes is highly dependent on both the index formula and the data availability. Therefore, we select the DM index formula because it satisfies all the following conditions: (1) produces a robust measure for the movements of the national commodity export prices, (2) fulfils its requirements with the data available from the international trade statistics, (3) is applicable for large set of countries and (4) has strong foundations in the existing literature.

2.4 Methodology

This section outlines the procedure used to construct the country-specific price indexes of commodity exports (COMPI) and their relevant sub-indexes. As mentioned earlier, our study uses the formula proposed by Deaton and Miller (1995) for constructing all the index series. Any decisions to deviate from this are identified below.

2.4.1 Nominal commodity price index

The IMF (2009, p. 33) highlights, "The geometric indexes are likely to be less subject than their arithmetic counterparts to the kinds of index number biases". Further, the arithmetic index fails on the time reversal test, while the geometric mean index satisfies it (Fisher, 1922). Due to these conditions, the arithmetic index likely overestimates or underestimates the "true" price changes (Cuddington and Wei, 1992). Hence, this chapter overcomes this issue by using the weighted geometric mean of price relatives instead of the arithmetic one. The index formula used for the construction of nominal national commodity export price index (NCOMPI) is:

$$NCOMPI = \prod_{k=1}^{K} P_k{}^{W_k} = exp\left\{\sum\nolimits_{k=1}^{K} (W_k(ln\, P_k))\right\} \tag{2.1}$$

where P_k is the US dollar-based world price of commodity k, W_k is the commodity k weighted item and ω is the set of all K commodities included in the corresponding country's basket of commodities. Then, $\forall k \in \omega$. This formula is also used for the construction of national commodity export price sub-indexes, where W_k is revised with respect to the relevant commodity grouping.

2.4.2 Index weights

The index weights for commodity k are calculated using the following formula:

$$W_k = \frac{P_{jk}Q_{jk}}{\sum_{l=1}^{K} P_{jl}Q_{jl}} \tag{2.2}$$

where W_k is the weighted item, which is the value of exports of commodity k in the total export value of all K commodities for the fixed-weight period j, i.e. $k \in \omega$; $P_{jk}Q_{jk}$ is the value of exports of commodity k for the fixed-weight period j.[37] Also, l is a commodity that is a part of the country's commodity basket, i.e. $l \in \omega$. Then, there is a case where $k \equiv l$. For the sub-index calculation, φ is the set of those commodities that are included in the relevant sub-index basket of commodities; hence, $\varphi \subset \omega$. Indeed, the commodity index weights W_k are held fixed over time.

Moreover, the commodity weight W_k is calculated by dividing the 1995–2010 total value of each individual k commodity export by the 1995–2010 total value of all K primary commodity exports. Specifically, the fixed-weight period j covers the years 1995–2010.

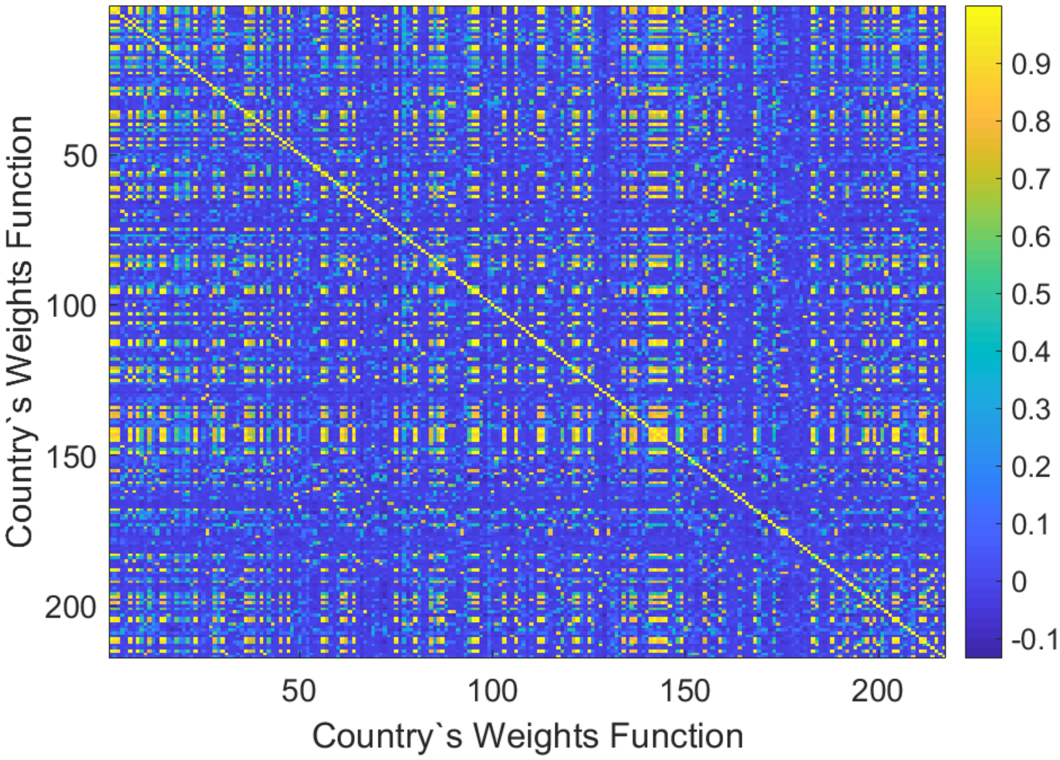
[37] In detail, P_{jk} is the country's export price of commodity k for the fixed-weight period j, whereas Q_{jk} is the country's quantity of export for commodity k for the fixed-weight period j.

Therefore, the formula for calculating the commodity index weights for commodity k can be re-written as follows:

$$W_k = \frac{\sum_{i=1995}^{2010} P_{ik} Q_{ik}}{\sum_{l=1}^{K} \sum_{i=1995}^{2010} P_{il} Q_{il}} \tag{2.3}$$

where i is a calendar year for which $i \in j$.

The country's weight function $\{W_1, W_2 \dots W_K\}$ might possibly suffer from endogeneity bias. Chen and Rogoff (2003) note that endogeneity may arise through the market power that certain countries may possess in the world commodity markets. However, Broda (2004) indicates that only a small number of countries exert such an influence, and they do so on a small share of commodities that they export. The substitution across similar commodity products further mitigates the market power these countries have, even within the specific markets that they appear to dominate (Chen and Rogoff, 2003). Cashin et al. (2004) also conclude that commodity-exporting countries are price-takers in world commodity markets and have negligible long-term market power in terms of their commodity exports (see Mendoza, 1995 for discussion). To address this potential form of endogeneity, our study undertakes a sensitivity analysis of the range of values of the 'weight' function by calculating Pearson correlation coefficients for each pair of countries' weights functions, see figure 2.4. The outcome of the correlation tests clearly shows a principal evidence of negative or no correlation between countries' weights functions. This implies that country-specific W_k is exogenous in nature and, therefore, have location-independent and idiosyncrasy-free estimates.



Note: The plot represents graphically Pearson correlation coefficients between country ķ weights function and country ļ weights function, where $ķ \in \{1,2, \ldots 217\}$ *and* $ļ \in \{1,2, \ldots 217\}$*; there is a case where* $ķ \equiv ļ$*.The colour bar denotes the correlation coefficient. Negative correlation is when the correlation coefficient is lower than 0, while positive is when the correlation coefficient is above 0. No correlation exists when the correlation coefficient is equal to 0.*

Figure 2.4 Heat Map of Correlation Coefficients between Countries' Weights Functions

It is important to emphasise that the choice of the base period is crucial in the index calculation. Our study follows Grilli and Yang (1988), Cashin et al. (2004), IMF (2018) in its way of selecting a non-singular year fixed-weight period. This reduces the impact of the data that has been highly affected by a single event (for example, a trade embargo in a particular year) on the index weights structure. As such, we set the lower bound of j to be 1995, while the upper bound of j is 2010. On the one hand, j's lower bound is selected to be 1995 in order to allow for a sufficient adjustment time for the structural changes at the national commodity markets after the events of the early 1990s, namely after the re-shaping of the physico-geographical borders of numerous countries. On the other hand, j's upper bound is chosen to be 2010 in order to allow enough time for actual export values to be recorded in the international trade statistics. For instance, the primarily data from input-output tables computed by Statistics Canada are only available with a four-year lag. That is to say, the actual trade data in the statistics of certain developing countries might be available with a significantly larger time-lag. Under these circumstances, the selection of j's upper bound to be 2010 is appropriate for reducing the possibility of including *estimated* instead of *actual* export values in the index weights construction.[38] Moreover, all W_k are country specific and, therefore, each commodity price index is inimitable.

Furthermore, the country-specific W_k is matched with the relevant world price of commodity k, i.e. P_k. The usage of international commodity prices instead of the export unit values in the calculation process of national commodity export indexes is consistent with the past academic literature, including but not limited to Deaton and Miller (1995), Dehn (2000), Cashin et al. (2004) as well as non-academic organisations such as ANZ Bank, Bank of Canada and Reserve Bank of Australia.

More precisely, the international commodity prices are useful for two reasons. First, the world commodity prices are typically unaffected by the behaviour of individual countries, as discussed by Deaton and Miller (1995). This reduces the possibility of having endogeneity problems in the index construction process. Second, a country's level trade data in monthly frequency of both prices and quantity is rarely available for a sample of 217 countries. Although data on export unit values is sometimes available, its level of disaggregation might not be sufficient to represent the true behaviour of individual commodity prices (see Silver, 2010 for discussion).[39] Therefore, the export unit values are inconceivable for the construction of a monthly world database of national commodity export indexes.

In brief, "We must admit that there is a large amount of measurement error in these unit values from the Comtrade Database" (Feenstra and Romalis, 2014, p. 496). Under these circumstances, our study uses the international commodity prices in the construction process of our database.

2.4.3 Seasonal adjustment

Following Cashin et al. (2004), we adjust for seasonality each nominal price index of national commodity exports by using the Census X-13ARIMA-SEATS procedure. In addition, we provide a separate database with index series that are not seasonally adjusted. This allows choosing an index series that best fits the research hypotheses' data requirements.

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[38] More precisely, the *estimated* (or *predicted*) values are sometimes reported by the national statistical agencies to provide a proxy for a given economic indicator when the actual data is not available. Nonetheless, the statistical agencies replace the expected values with actual data when the latter becomes available. In other words, the actual values are recorded in the international trade statistics with some time-lag.

[39] The export unit values for commodity k are obtained as the export value is divided by the relevant export volume of the commodity k (see Junz and Rhomberg, 1973; Isard, 1977).

2.4.4 Real commodity price index

This study provides an additional database with real national commodity export price indexes.[40] The real national commodity export index is calculated as the nominal index is deflated by the unit value index of manufactured goods exports (abbreviated as the MUV). The MUV index is used as a deflator in most of the earlier studies in the literature. Some examples include Grilli and Yang (1988), Cuddington and Urzúa (1989), Cuddington (1992), Deaton and Miller (1995), Cuddington and Liang (1998), Deaton (1999), Cashin et al. (2004), Kellard and Wohar (2006), Baffes (2007), Harvey et al. (2010) and Janus and Riera-Crichton (2015). Indeed, the real national commodity price index (RCOMPI) is calculated using the following equation:

$$RCOMPI = \frac{exp\{\sum_{k=1}^{K}(W_k(ln\,P_k))\}}{MUV} * 100 = \frac{NCOMPI}{MUV} * 100 \qquad (2.4)$$

In addition, other deflators that are widely used in the economic literature are the Consumer Price Index (CPI) and the Unit Labour Costs (ULC). While this study follows the prevailing part of the commodity price literature by using the MUV index as a deflator for our national commodity export indexes, we also provide the general formula for calculating the real national commodity export price index, with the use of different deflators:

$$RCOMPI^{*} = \frac{NCOMPI}{Def} * 100 \qquad (2.5)$$

where Def is a selected deflator.

Next, the base $1995 = 100$ is selected for each index number. If necessary, a researcher may replace this index base with another one. Due to the axiomatic properties of the fixed-base DM index, such a change does not have an impact on the preciseness of the national commodity export price indexes. Further, the reliability of the data sources is another important factor for creating a robust index database. The following section provides more information on this.

2.5 Data

This section describes the data sources and the countries coverage that are considered in the construction of our database.

2.5.1 Data sources

The data of export values for each commodity is taken from the UN Comtrade (2018) and UNCTAD (2018) databases. These databases are subject to irregular revision by the United

[40] See also Deaton and Miller (1995), Deaton (1999), Dehn (2000), Cashin et al. (2004) and Bodart et al. (2012).

Nations Statistical Department and, therefore, it is important to point out that our study considers the data collected as of August 2017. As such, more recent data revisions of these databases have not been taken into account in this study. Similarly, if any of the international price data sources have revised their data after August 2017, this has not been taken into consideration in the current version of our index series.

We collected monthly frequency data of international commodity prices for the period from January 1980 to April 2017. The price data is mainly taken from the IMF (2018) and the International Financial Statistics (IFS) (IFS, 2018) databases. In cases when the commodity price data is not accessible through either the IFS or IMF databases, other data sources are used. An example is the price of canola, which is taken from the Canola Council of Canada (CCC) (Canola Council of Canada, 2018). This is due to the unavailability of canola price series in either the IFS or IMF databases. Moreover, the CCC does not provide data on canola prices for the period prior to April 1983. Hence, our study follows Dehn (2000) for holding the pre-April 1983 canola prices constant at the value of the first available observation.

Furthermore, the price series for a few important commodities have been collected from the World Bank (2018), FAO (2018) and UNCTAD (2018).[41] This is done due to the lack of credible data in either the IMF or IFS databases. More details on the price series and their corresponding data sources are provided in Appendix A.1.

2.5.2 Dairy data sources

One of the main strengths of our database is the inclusion of the dairy price series in the index calculation. More precisely, our study considers the following dairy products: butter, cheese, skim milk powder and whole milk powder. The world prices of these commodities are obtained from the FAO (2018) database. Unfortunately, the price series of dairy products in the FAO database are only available starting January 1990. As such, there is a 10-year gap of missing dairy price data for the period between January 1980 and January 1990. There are three different options of dealing with this problem.

First, we may follow the method of Dehn (2000) by holding the pre-January 1990 prices constant at the value of the first available observation. This may cause the imputation of 480 observations in the world prices data set. As such, the national commodity price index for a country where the main exports come from dairy products may provide unreliable index series. One example for such a country is New Zealand.

[41] FAO is an acronym for the Food and Agriculture Organization (of the United Nations).

Second, the price series of dairy products can be excluded from the index calculation either partially or as a whole. Then, the index series for dairy-exporting countries may not be accurate representative of the price movements in the national commodity markets. An obvious example is the index series for New Zealand that are constructed by Sahay et al. (2002).

Third, the missing dairy data is obtained from the *Status Report on the World Market for Dairy Products, International Dairy Arrangement*, published by the United Nations Secretariat (International Dairy Arrangement, 1981; 1982; 1983; 1984; 1985; 1986; 1987; 1988; 1989; 1990). The reported data is of quarterly frequency; therefore, the monthly values can be obtained after data interpolation. More precisely, the mid-points of international price ranges can be obtained and then held constant during the relevant quarter period. We assume that this procedure is an improvement of the Dehn (2000) method because it allows the prices to fluctuate on a quarterly basis instead of holding a constant price for nearly 120 months – particularly for the pre-January 1990 period.

Given these three options, our study selects the third method to fill the 10-year gap of missing price data for the dairy products. This is because keeping the dairy prices constant or excluding them from the index calculation is likely to bring about certain concerns regarding the robustness of the index series (for example, see Sahay et al., 2002). Therefore, our study uses the third method for filling the gaps in the dairy price series and then calculates the national commodity export price indexes.

2.5.3 Conversion units

All international prices are converted to have the same unit values, namely US dollars per metric tonne. First, the international commodity prices that are provided by the World Bank, the IMF, UNCTAD, CCC, FAO and International Dairy Arrangement are all in US dollars. Therefore, there is no need for a currency conversion to be carried out.[42] Second, the conversion to metric tonnes is done by using the conversion factors provided in the *Forest products conversion factors for the UNECE Region* (Fonseca, 2010) and *Monthly Bulletin of Statistics* by United Nations Statistics Division (DESA, 2017). More details for the mathematical conversion factors and formulas are reported in Appendices A.2 and A.3.

2.5.4 The choice of data deflator

The nominal commodity price indexes are deflated with the MUV index that is reported by UNCTAD (2018). This is consistent with the earlier studies of Grilli and Yang (1988) and

[42] If prices are given in cents, then we use the conversion ratio of 1 US$ is equal to 100 US cents.

Harvey et al. (2010). Particularly, we do not use the MUV index from the IMF and the World Bank, as of Deaton and Miller (1995) and Cashin et al. (2004), because the MUV index of UNCTAD is likely to provide a better reflection of the price volatility of manufactured goods exports from developed countries.[43] This is because the MUV index of UNCTAD is of quarterly frequency, while the MUV index provided by the IMF and the World Bank is of annual frequency. This implies that the former is likely to capture more precisely the changes in the manufactured goods exports as compared to the latter. Therefore, we use the MUV index provided by UNCTAD to deflate the nominal index series in our database.

2.5.5 Country coverage

Our database covers 217 countries, of which seven are Northern African countries, 17 are from Eastern Africa, nine are from Middle Africa, five are from Southern Africa, 17 are from Western Africa, 20 are from Caribbean, eight are from Central America, 13 are from South America, five are from Northern America, five are from Central Asia, seven are from Eastern Asia, 11 are from South-eastern Asia, nine are from Southern Asia, 18 are from Western Asia, ten are from Eastern Europe, 11 are from Northern Europe, 15 are from Southern Europe, seven are from Western Europe and 23 are from Oceania. The certain regional split of the countries is taken from the United Nations classification "Standard Country or Area Codes for Statistical Use", which is originally published as Series M, No. 49 (UNSTATS, 2018). For basic descriptive statistics on each country's structure of trade and regional affiliation, see Appendix A.4. It should be noted that our database includes some countries that either no longer exist after 2017 or have altered their borders during the period of 1980–2017. Therefore, it is crucial that the researcher does not neglect this fact in the country selection process. Nonetheless, the accessibility to data for dissolved countries is predisposed for an economic history analysis to be undertaken.

2.5.6 Official benchmark indexes of national commodity export prices

Three *official* benchmark indexes are used for checking the robustness of our database. The first one is the Bank of Canada commodity price index (BCPI), which is a proxy for the national commodity export prices for Canada; the index is obtained from the Bank of Canada (2018). The second one is the ANZ Commodity Price Index, which is a proxy for the national commodity export prices for New Zealand; the index is obtained from the Australia and New Zealand Banking Group (2018). The third one is the Reserve Bank of Australia Index of Commodity Prices, which is a proxy for the national commodity export prices for Australia; the index is obtained from the Reserve Bank of Australia (2018).

[43] The IMF and the World Bank use the same method to construct the MUV index.

2.6 An Empirical Comparison of Indexes

In order to ensure the robustness of our database, we compare empirically the performance of our commodity price indexes with those constructed by the central and commercial banks.[44] To reiterate, there are only three institutions in the world that provide regularly updated national commodity export price indexes, namely the Bank of Canada (for Canada), the Reserve Bank of Australia (for Australia) and the ANZ Bank (for New Zealand). We denote the index series that are calculated by these three institutions as "*official*" and use them as a "*benchmark*" for our index series.[45]

To emphasise, Sahay et al. (2002) also use the commodity price indexes from the aforementioned data sources as a benchmark for the robustness of their data series. However, the authors find that the official index is highly uncorrelated with their constructed one for New Zealand. The authors presume the exclusion of dairy products from their index commodity basket as a possible reason for this result. In fact, our study fills this gap by including the world dairy prices in the construction process of our index series. In addition, the nominal commodity price indexes from the database of Sahay et al. (2002), later used by Cashin et al. (2004), are also included in the analysis below.[46] Indeed, the index series that are considered in this analysis and their relevant sources are listed in Appendix A.5.

Similar to Melser (2018), the performance of the indexes is evaluated on the basis of three main factors. First is the correlation between the indexes. This measures the strength of the linear relationship between each pair of indexes. In other words, a correlation coefficient with a high positive value (close to one) entails that there is a strong relationship between the two index series. In other words, the series move in the same direction. Otherwise, a correlation coefficient with a high negative value (close to minus one) entails that one variable increases as the other decreases, and vice versa. In the case when a correlation coefficient is equal to zero, there is no relationship between the two index series. The second factor is the differences between the indexes. This is measured by the difference in the average annual percentage change between each of the indexes and all the other indexes. In fact, this provides a measure of deviation for each pair of indexes. The resemblance between the indexes is the third factor. This is determined by comparing the absolute difference in the annual percentage change between each index. Otherwise speaking, the similarity of the indexes and the extent to which they record different measures of price change are identified (Melser, 2018). With

[44] As far as we know, there is no a national statistical agency that provides regularly updated data on national price index of commodity exports.
[45] These institutions report the commodity price indexes in nominal terms.
[46] The data is available from January 1980 to March 2002 and was kindly provided by Dr Paul Cashin.

respect to the last two factors, it should be noted that the percentage change is considered instead of the reported index value. This aims to provide a relative measure that accounts for the size of the index number.[47]

2.6.1 *Correlation between the indexes*

To illustrate the robustness of our database, the correlation coefficient between the index series is calculated.[48] Due to the exclusion of fuel primary products from the calculation of Cashin et al. (2004) indexes, their series are compared with the relevant national non-energy commodity price indexes. In fact, the correlation for each pair of indexes is estimated by the following methods: Pearson's linear and Spearman's rank-order correlation tests. Importantly, the results should be interpreted with caution due to possible dissimilarities between the index commodity baskets. The results from the Pearson's correlation tests are presented in table 2.4:

Index pairs		**Correlation Coefficient**
Official_all_Aus	NCOMPI_all_Aus	0.987***
Official_all_Can	NCOMPI_all_Can	0.989***
Official_all_NZ	NCOMPI_all_NZ	0.940***
Official_ne_Can	NCOMPI_ne_Can	0.985***
Official_ne_Can	Cashin_official_Can	0.881***
NCOMPI_ne_Can	Cashin_official_Can	0.883***
Official_all_Aus	Cashin_official_Aus	0.843***
Official_all_NZ	NCOMPI_ne_NZ	0.939***
Official_all_NZ	Cashin_official_NZ	0.407***
Official_dairy_NZ	NCOMPI_dairy_NZ	0.988***
Official_energy_Can	NCOMPI_energy_Can	0.986***
Official_food_Can	NCOMPI_food_Can	0.966***
Official_m_Aus	NCOMPI_m_Aus	0.776***
Official_m_Can	NCOMPI_m_Can	0.974***
Official_m_NZ	NCOMPI_m_NZ	0.975***

*Note: *, ** and *** denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table 2.4 Pearson Correlation Test

[47] For example, the absolute difference of five, between one and six, is more significant than the same absolute difference between 105 and 100.

[48] Correlation coefficients are used to measure the strength of the relationship between two series. They can vary numerically between -1.000 and 1.000. The closer the correlation is to 1.000, the stronger the relationship between the index series. In other words, a positive correlation coefficient means that as series one increases, series two increases, and conversely, as series one decreases, series two decreases. Thus, positive correlation is a relationship between two series which move in tandem – that is, in the same direction. Whereas, negative correlation is a relationship between two series that move in opposite directions. Further, quantile correlation coefficients are reported in order to provide evidence that a high correlation is irrespective to a general rise or decrease in the price trend of commodities.

The results in table 2.4 indicate that the NCOMPI_all_Aus, NCOMPI_all_Can and NCOMPI_all_NZ indexes are highly correlated with their corresponding benchmarks, namely Official_all_Aus, Official_all_Can, Official_all_NZ respectively. In fact, the correlation coefficients for those comparisons are 0.987, 0.989 and 0.940 in terms of Australia, Canada and New Zealand respectively. Therefore, we conclude that our index series have a strong positive relationship with their corresponding benchmarks.

Importantly, the correlation results from the non-energy indexes provide further support for the reliability of our index series. First, the correlation coefficient between NCOMPI_ne_Can and Official_ne_Can is 0.985, whereas the correlation coefficient between Official_ne_Can and Cashin_official_Can is down to 0.881. This implies that our non-energy index for Canada has a stronger linear relationship with the benchmark index than the one constructed by Cashin et al. (2004). Second, the correlation coefficient between Official_all_NZ and NCOMPI_ne_NZ is 0.939, whereas the coefficient of correlation between Official_all_NZ and Cashin_official_NZ is 0.407.[49] The latter result is consistent with the statement of Sahay et al. (2002, p. 53) that their "constructed and bank indexes for New Zealand differ somewhat". This statement is also valid for Cashin et al. (2004) because the two studies employ the same data set. In brief, the findings of high positive correlation demonstrate that our index series track the price movement in the national commodity markets more accurately than the index series constructed by past studies such as that constructed by Cashin et al. (2004). This conclusion is based on the results given in table 2.4.

In addition, the correlation coefficients between our indexes and their corresponding benchmarks are found to be at least 0.939, as shown in table 2.4. The only exception is the coefficient of correlation between NCOMPI_m_Aus and Official_m_Aus, which is equal to 0.776. A possible reason for reaching this outcome is the exclusion of iron ore from the index basket of Official_m_Aus, whereas this commodity is included in the calculation of NCOMPI_m_Aus index. Nonetheless, iron ore is included in the index baskets of both NCOMPI_all_Aus and Official_all_Aus, as it has a major share in the Australian export.[50] To iterate, the correlation coefficient between NCOMPI_all_Aus and Official_all_Aus is 0.987. As such, we can assume that the correlation coefficient of 0.776 is more due to the differences in the index commodity grouping rather than a weakness of our database.

[49] The official index of New Zealand that is constructed by ANZ Bank does not include any energy commodity products. Thus, Official_all_NZ can also be classified as a non-energy commodity index.

[50] The iron ore comprises of almost one-fifth of the commodity index basket of Official_all_Aus as of April 2017 (Reserve Bank of Australia, 2018).

Further, the calculation of Pearson's correlation coefficient is sensitive to skewed distributions and outliers. Hence, we examine the reliability of our outcomes from the Pearson's correlation test by using the Spearman's correlation test. The results from Spearman's rank correlation are presented in table 2.5:

Index pairs		Correlation Coefficient
Official_all_Aus	NCOMPI_all_Aus	0.975***
Official_all_Can	NCOMPI_all_Can	0.971***
Official_all_NZ	NCOMPI_all_NZ	0.924***
Official_ne_Can	NCOMPI_ne_Can	0.972***
Official_ne_Can	Cashin_official_Can	0.855***
NCOMPI_ne_Can	Cashin_official_Can	0.857***
Official_all_Aus	Cashin_official_Aus	0.840***
Official_all_NZ	NCOMPI_ne_NZ	0.920***
Official_all_NZ	Cashin_official_NZ	0.457***
Official_dairy_NZ	NCOMPI_dairy_NZ	0.972***
Official_energy_Can	NCOMPI_energy_Can	0.976***
Official_food_Can	NCOMPI_food_Can	0.934***
Official_m_Aus	NCOMPI_m_Aus	0.911***
Official_m_Can	NCOMPI_m_Can	0.990***
Official_m_NZ	NCOMPI_m_NZ	0.985***

*Note: *, ** and *** denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table 2.5 Spearman Rank-Order Correlation Test

As can be noted, there is a negligible difference between the results obtained by Pearson's linear and Spearman's rank-order correlation tests (see tables 2.4 and 2.5). Moreover, all correlation coefficients are significant at 1% significance level. That is to say, the correlation tests confirm the superiority of our index series over those constructed in the past literature. This provides certain reliability for the robustness of our national commodity export price indexes.

Index pairs		Correlation Coefficient				
		q10	q25	q50	q75	q90
Official_all_Aus	NCOMPI_all_Aus	0.985***	0.986***	0.987***	0.988***	0.989***
Official_all_Can	NCOMPI_all_Can	0.987***	0.988***	0.989***	0.990***	0.990***
Official_all_NZ	NCOMPI_all_NZ	0.933***	0.937***	0.940***	0.944***	0.947***
Official_ne_Can	NCOMPI_ne_Can	0.983***	0.984***	0.985***	0.986***	0.987***
Official_ne_Can	Cashin_official_Can	0.865***	0.873***	0.881***	0.889***	0.895***
NCOMPI_ne_Can	Cashin_official_Can	0.866***	0.874***	0.884***	0.891***	0.899***
Official_all_Aus	Cashin_official_Aus	0.819***	0.832***	0.843***	0.855***	0.865***

Official_all_NZ	NCOMPI_ne_NZ	0.932***	0.936***	0.939***	0.942***	0.945***
Official_all_NZ	Cashin_official_NZ	0.339***	0.375***	0.408***	0.441***	0.474***
Official_dairy_NZ	NCOMPI_dairy_NZ	0.986***	0.987***	0.988***	0.989***	0.990***
Official_energy_Can	NCOMPI_energy_Can	0.984***	0.985***	0.986***	0.987***	0.988***
Official_food_Can	NCOMPI_food_Can	0.962***	0.964***	0.966***	0.968***	0.970***
Official_m_Aus	NCOMPI_m_Aus	0.753***	0.763***	0.776***	0.792***	0.803***
Official_m_Can	NCOMPI_m_Can	0.971***	0.973***	0.974***	0.976***	0.977***
Official_m_NZ	NCOMPI_m_NZ	0.969***	0.972***	0.976***	0.979***	0.982***

*Note: *, ** and *** denote the statistical significance at the 10%, 5% and 1% level respectively.q10, q25, q50, q75, q90 denotes the 10^{th}, 25^{th},50^{th}, 75^{th}, and 90^{th} quantiles respectively*

Table 2.6 Spearman Rank-Order Correlation Test across five quantiles, from the 10th through 90th percentile

Table 2.6 provides the results from Spearman's rank correlation across five quantiles, from the 10th through 90th. One advantage of quantile correlation is that the quantile estimates are more robust against outliers in the response measurements, especially when the data is skewed. Therefore, we estimate the correlation coefficients at various quantiles of the distribution, in order to capture heterogeneity and bias. The empirical results reported in table 2.6 show evidence of strong positive correlation between our constructed index series and the official benchmark indexes at various quantiles of the distribution. This demonstrates that our constructed index series and the official benchmark indexes move in the same direction with the same magnitude.

As shown, the results in table 2.6 are rather similar to those reported in tables 2.4 and 2.5. More importantly, the findings from quantile correlation coefficients demonstrate a high positive correlation between our index series and the benchmark indexes. This implies a strong interdependence between our indexes and the benchmarks. The evidence of high correlation may be due to the fact that commodity prices respond quickly to general economic shocks such as increases in demand (Furlong and Ingenito, 1996). Alternatively, the evidence of high correlation may perhaps be due to the general rise in the prices (the trend) irrespective of the type of indexes constructed. This is apparent from the empirical findings obtained at high quantiles of the probability distribution, see table 2.6. Precisely, table 2.6 demonstrates that the strongest correlation between the index series is achieved at the highest level of the probability distribution.

2.6.2 Differences between the indexes

According to Melser (2018), this is one of the most important features of the empirical analysis, which is likely to provide some reassurance to statistical agencies, i.e. when comparing two indexes, say A and B. Thus, the difference in annual average percentage change between A and B is as follows:

$$\bar{d}^{AB} = \frac{\sum_{t=1}^{T} d_t^{AB}}{T} = \frac{\sum_{t=1}^{T} \Delta\alpha_t^A - \Delta\alpha_t^B}{T} \tag{2.6}$$

where α_t^A is the index A in time period t and, therefore, $\Delta\alpha_t^A = 100 * [\frac{\alpha_t^A - \alpha_{t-12}^A}{\alpha_{t-12}^A}]$ ($\Delta\alpha_t^B$ is defined analogously). Then, $d_t^{AB} = \Delta\alpha_t^A - \Delta\alpha_t^B$ is the difference between indexes A and B in time period t. If there exists a time period t where either A or B, or both, do not have an index value reported, this period is removed from the index calculation. As such, T is the total number of all calculated d_t^{AB} differences between indexes A and B. The results for each pair of indexes are shown in table 2.7:

Index pairs		Coefficient of difference in %
Official_all_Aus	NCOMPI_all_Aus	-0.202
Official_all_Can	NCOMPI_all_Can	-0.447
Official_all_NZ	NCOMPI_all_NZ	-0.607
Official_ne_Can	NCOMPI_ne_Can	-0.287
Official_ne_Can	Cashin_official_Can	-0.241
NCOMPI_ne_Can	Cashin_official_Can	0.294
Official_all_Aus	Cashin_official_Aus	0.460
Official_all_NZ	NCOMPI_ne_NZ	-0.481
Official_all_NZ	Cashin_official_NZ	-0.470
Official_dairy_NZ	NCOMPI_dairy_NZ	-2.213
Official_energy_Can	NCOMPI_energy_Can	0.073
Official_food_Can	NCOMPI_food_Can	-0.005
Official_m_Aus	NCOMPI_m_Aus	-0.121
Official_m_Can	NCOMPI_m_Can	-0.859
Official_m_NZ	NCOMPI_m_NZ	-0.142

Table 2.7 Differences in Annual % Change

None of the index pairs presented in table 2.7 provide obviously erroneous results. All differences have values less than 1% in the annual percentage change. The only exception is the index pair of Official_dairy_NZ and NCOMPI_dairy_NZ. The difference in the annual percentage change for this particular index pair is 2.213%, which is a negligible number with respect to the time length that both index series cover. The particular index series cover a period of more than 37 years including periods of highly volatile commodity prices, such as the 2007–2008 Global Financial Crisis. Overall, the results for the pairs of *all commodities* indexes, as shown in table 2.7, reveal a negligible difference in the average annual percentage change between our index series and the official indexes. This outcome provides further support for the reliability of our index database.

2.6.3 Resemblance between the indexes

This section presents the third criterion, which is the resemblance between the index pairs. This study follows the method suggested by Melser (2018) to calculate the absolute difference in the annual average percentage change for each index pair. In other words, this is our measurement of the index similarity. Specifically, when comparing indexes, say A and B, the absolute difference in the annual average percentage change between A and B, i.e. $\bar{d}_{abs}^{AB}$, is derived from Equation (2.6). Then, the $\bar{d}_{abs}^{AB}$ is calculated using the following formula:

$$\bar{d}_{abs}^{AB} = \frac{\sum_{t=1}^{T} |d_t^{AB}|}{T} = \frac{\sum_{t=1}^{T} |\Delta\alpha_t^{A} - \Delta\alpha_t^{B}|}{T} \qquad (2.7)$$

where the $\left|\Delta\alpha_t^{A} - \Delta\alpha_t^{B}\right|$ is the absolute difference between $\Delta\alpha_t^{A}$ and $\Delta\alpha_t^{B}$ at time period t, as defined above. The estimation results from the resemblance between the index pairs are provided in table 2.8:

Index pairs		**Coefficient of absolute difference in %**
Official_all_Aus	NCOMPI_all_Aus	6.454
Official_all_Can	NCOMPI_all_Can	4.233
Official_all_NZ	NCOMPI_all_NZ	5.749
Official_ne_Can	NCOMPI_ne_Can	4.203
Official_ne_Can	Cashin_official_Can	5.359
NCOMPI_ne_Can	Cashin_official_Can	5.311
Official_all_Aus	Cashin_official_Aus	4.153
Official_all_NZ	NCOMPI_ne_NZ	5.634
Official_all_NZ	Cashin_official_NZ	8.792
Official_dairy_NZ	NCOMPI_dairy_NZ	6.796
Official_energy_Can	NCOMPI_energy_Can	6.193
Official_food_Can	NCOMPI_food_Can	5.104
Official_m_Aus	NCOMPI_m_Aus	17.131
Official_m_Can	NCOMPI_m_Can	6.201
Official_m_NZ	NCOMPI_m_NZ	2.965

Table 2.8 Absolute Differences in Annual % Change

Notably, none of these index pairs appear to show large deviations (see table 2.8). However, there is one exception – the absolute difference in annual average percentage change between NCOMPI_m_Aus and Official_m_Aus, which is 17.131%. This result indicates dissimilarity between our metal index series and the official metal index for Australia. In fact, the possible reason for this dissimilarity has already been discussed in the correlation results section (see Section 2.6.1). Therefore, the result for the NCOMPI_m_Aus and Official_m_Aus index pair should not be taken into a serious consideration for the overall performance of our index database. In general, table 2.8 shows that all index pairs, apart from the aforementioned one, have an absolute difference in the annual average percentage change of less than nine percentage points. This finding lends support for the reliability of our index database.

2.6.4 Plots of national commodity prices and national aggregates

This section provides plots of national aggregates, such as employment and real GDP per capita, which may closely proximate national commodity prices of respective commodity-dependent countries. In absence of suitable comparison of the national commodity prices, as a benchmark, these aggregates can capture broader trends. Below, this thesis provides some plots that illustrate the movements of employment, real GDP per capita and RCOMPI for representative countries, subject to data accessibility. Data for employment and real GDP per capita is obtained from IMF (2018) and Datastream (2018), respectively, for Bahrain, Bolivia, Canada, Chile, Colombia, Ecuador, Kazakhstan, Kenya, Norway, Peru, Russian Federation, Tanzania, Venezuela, and Viet Nam. For commodity-dependent economies, national commodity prices can capture broader trends in employment and real GDP per capita for numerous countries, as shown in figures 2.5 – 2.18.

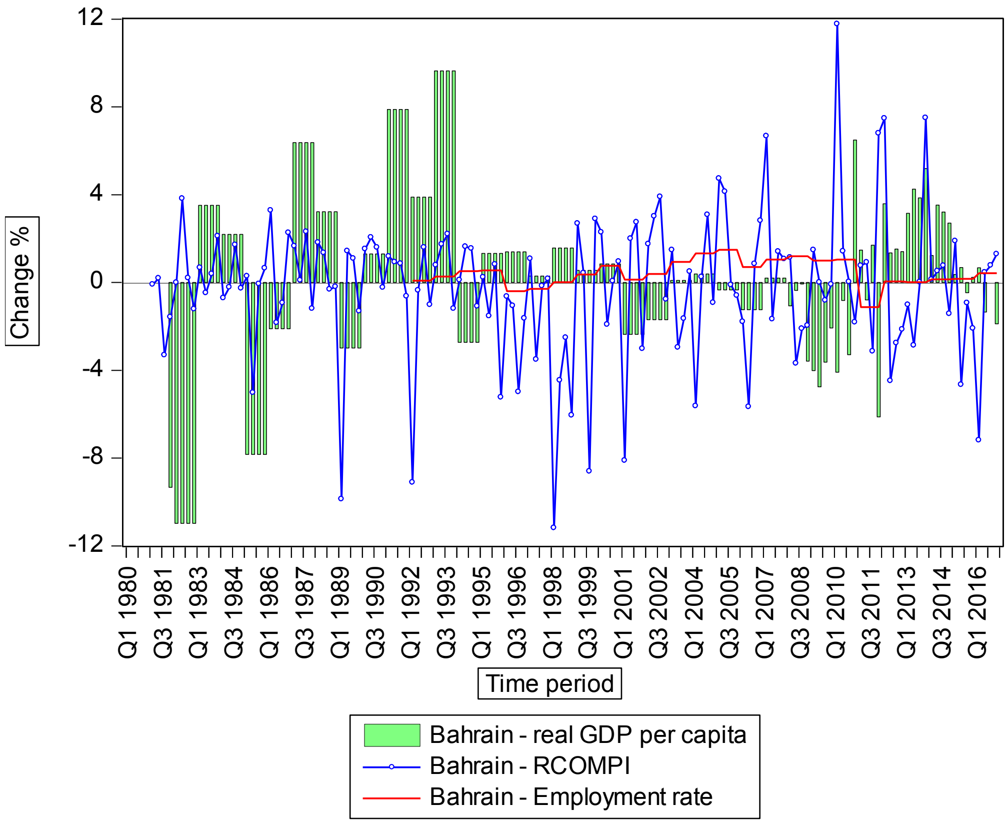

Figure 2.5 National Aggregates and RCOMPI – Bahrain

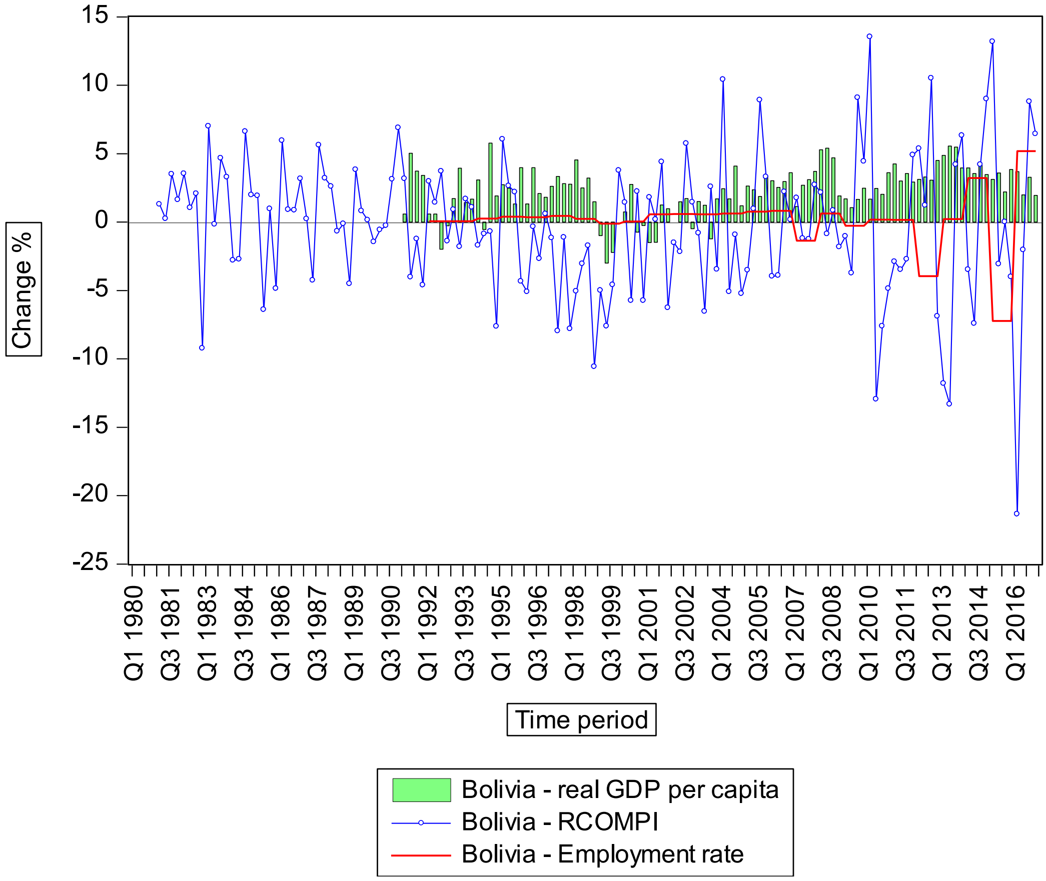

Figure 2.6 National Aggregates and RCOMPI – Bolivia

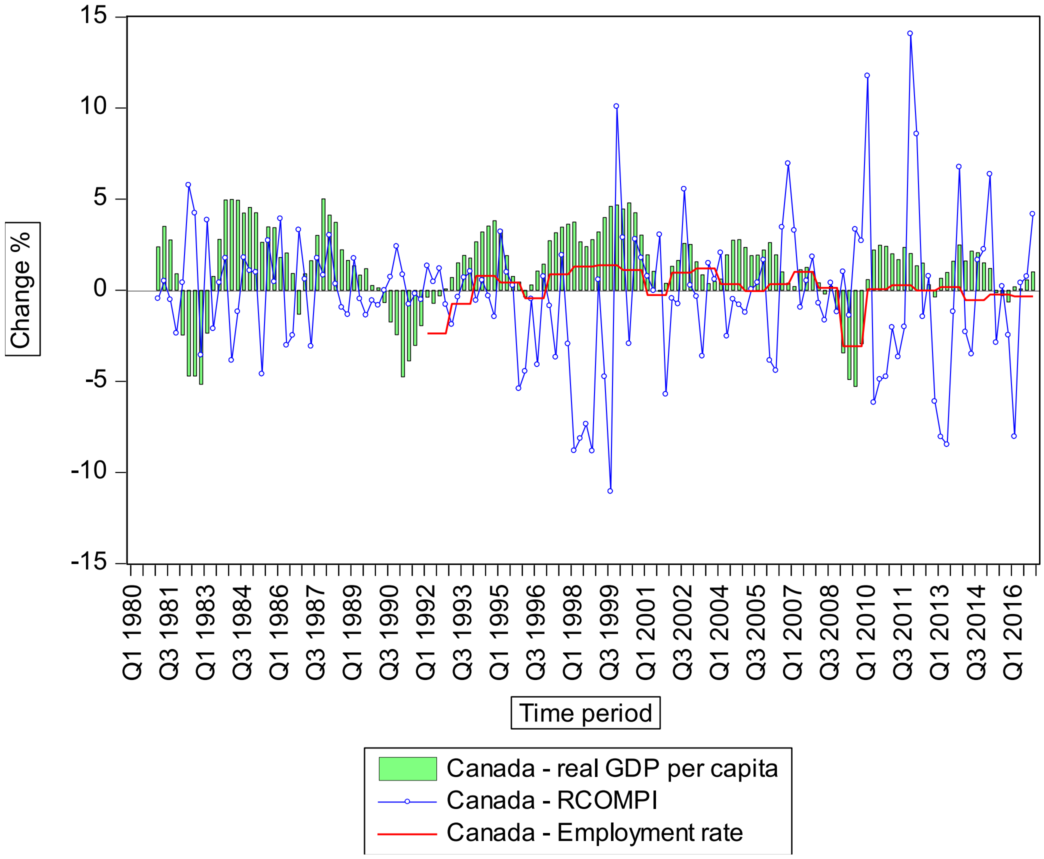

Figure 2.7 National Aggregates and RCOMPI – Canada

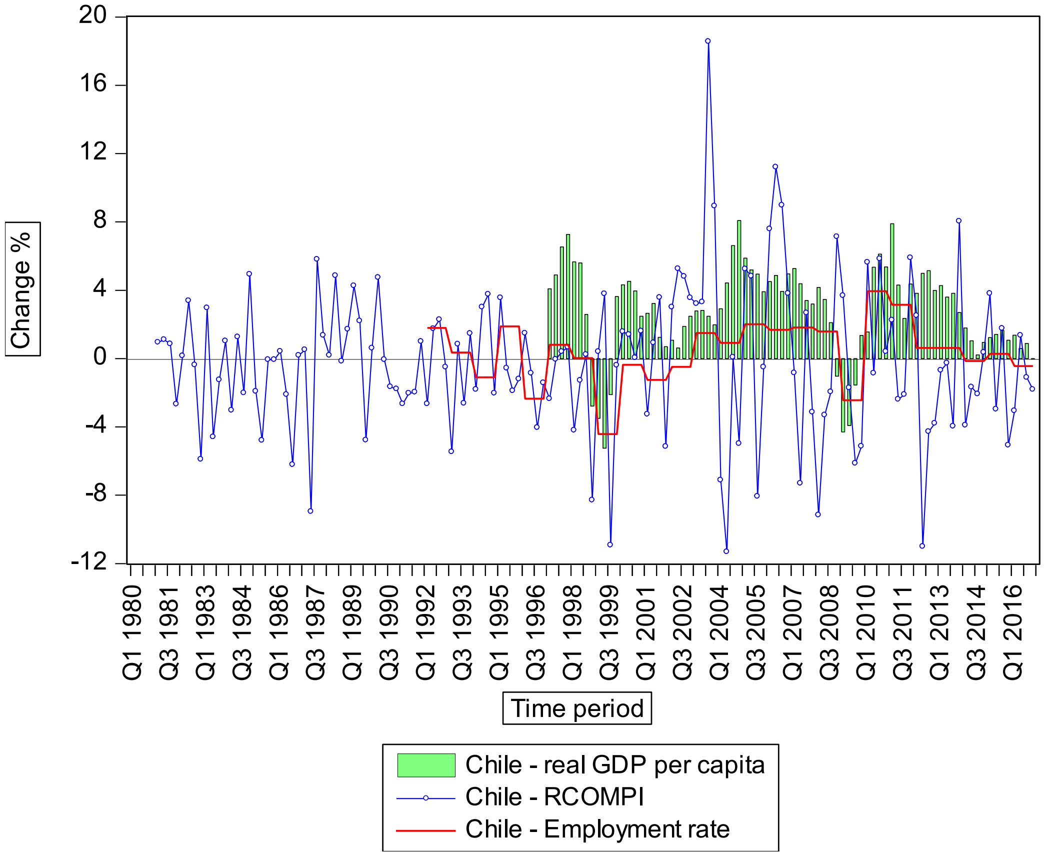

Figure 2.8 National Aggregates and RCOMPI – Chile

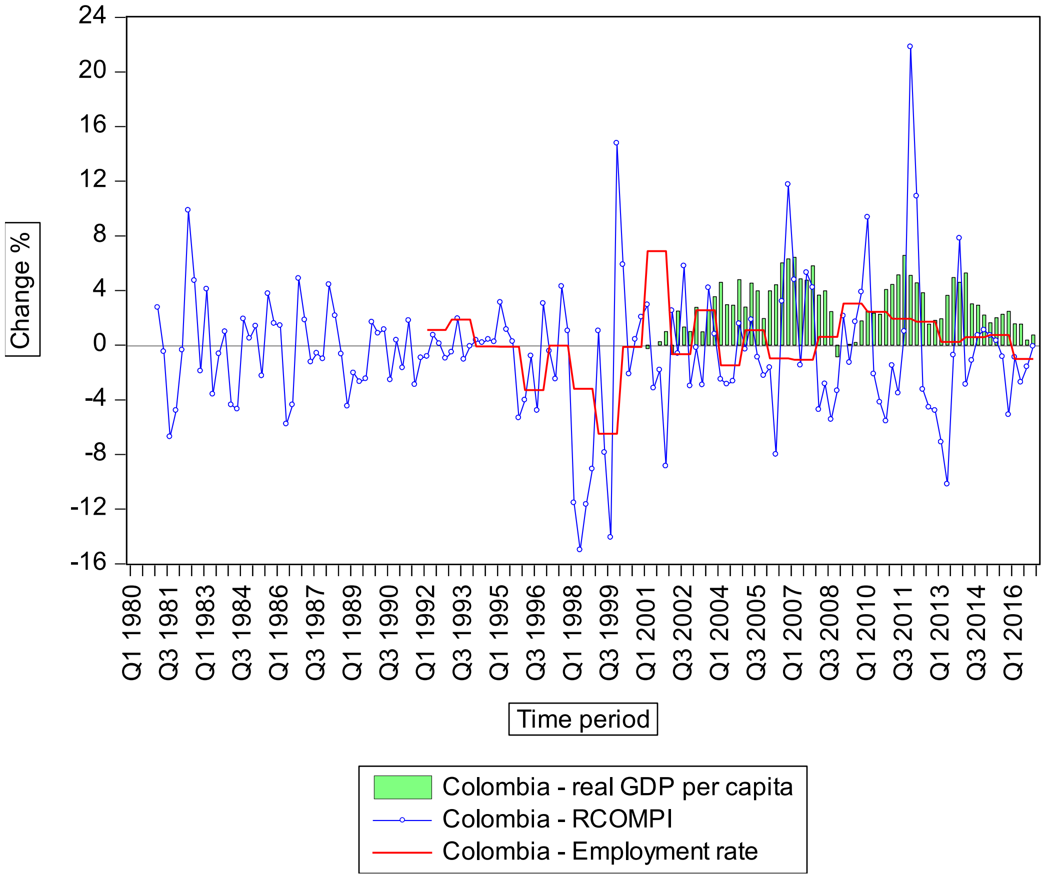

Figure 2.9 National Aggregates and RCOMPI – Colombia

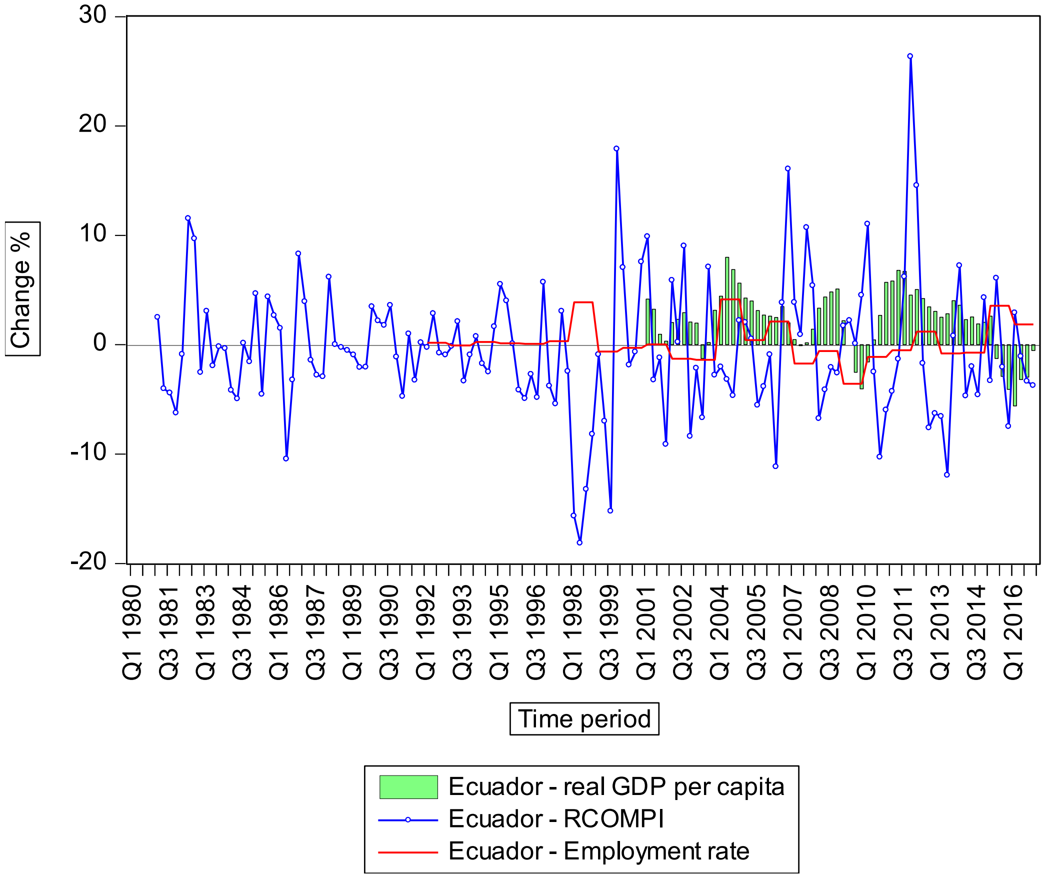

Figure 2.10 National Aggregates and RCOMPI – Ecuador

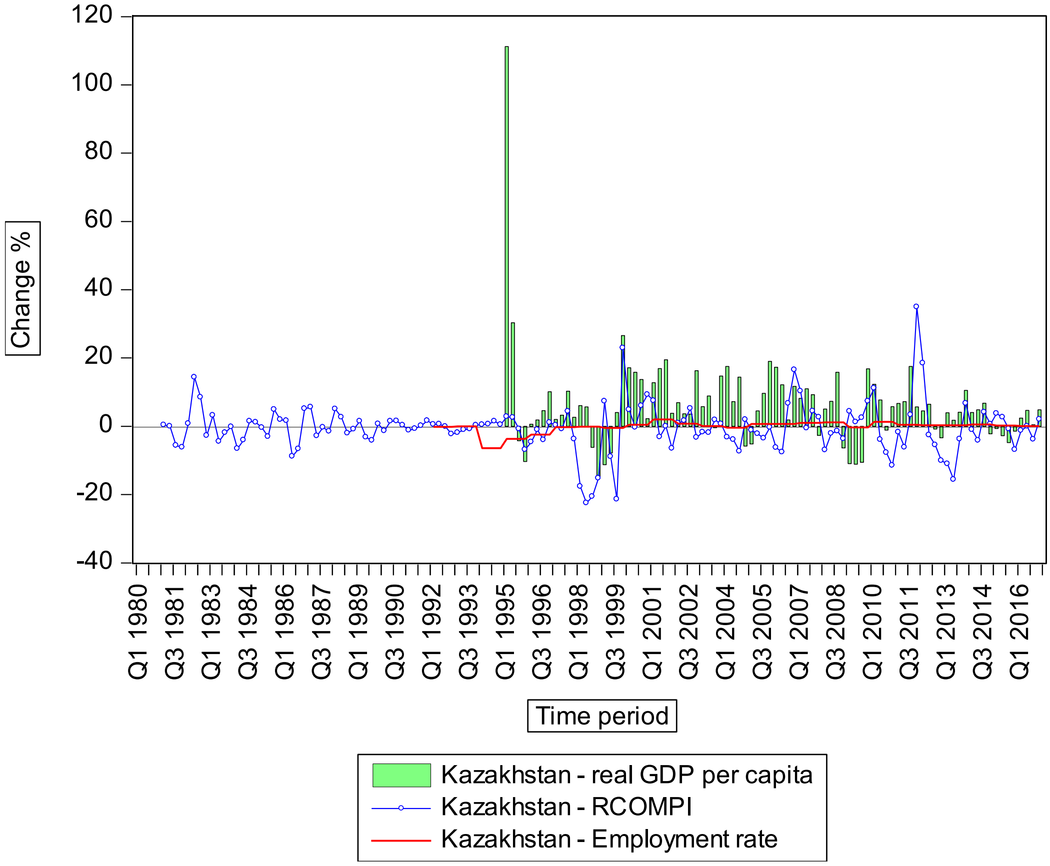

Figure 2.11 National Aggregates and RCOMPI – Kazakhstan

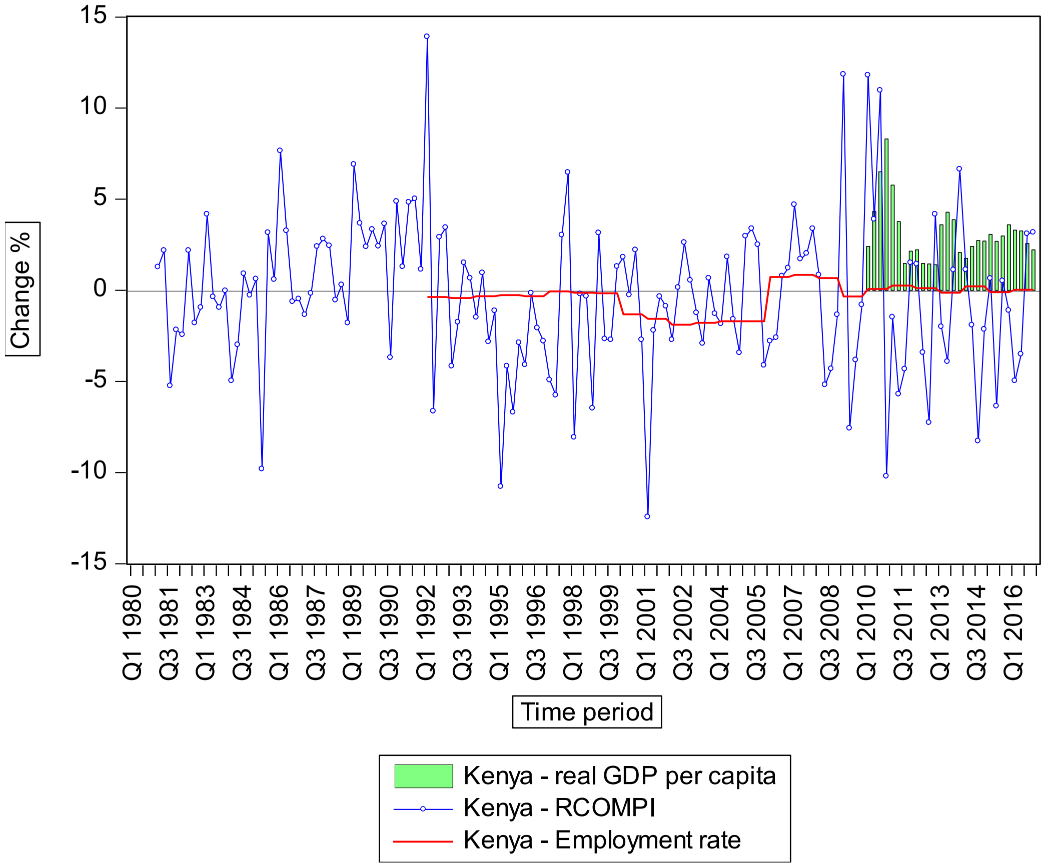

Figure 2.12 National Aggregates and RCOMPI – Kenya

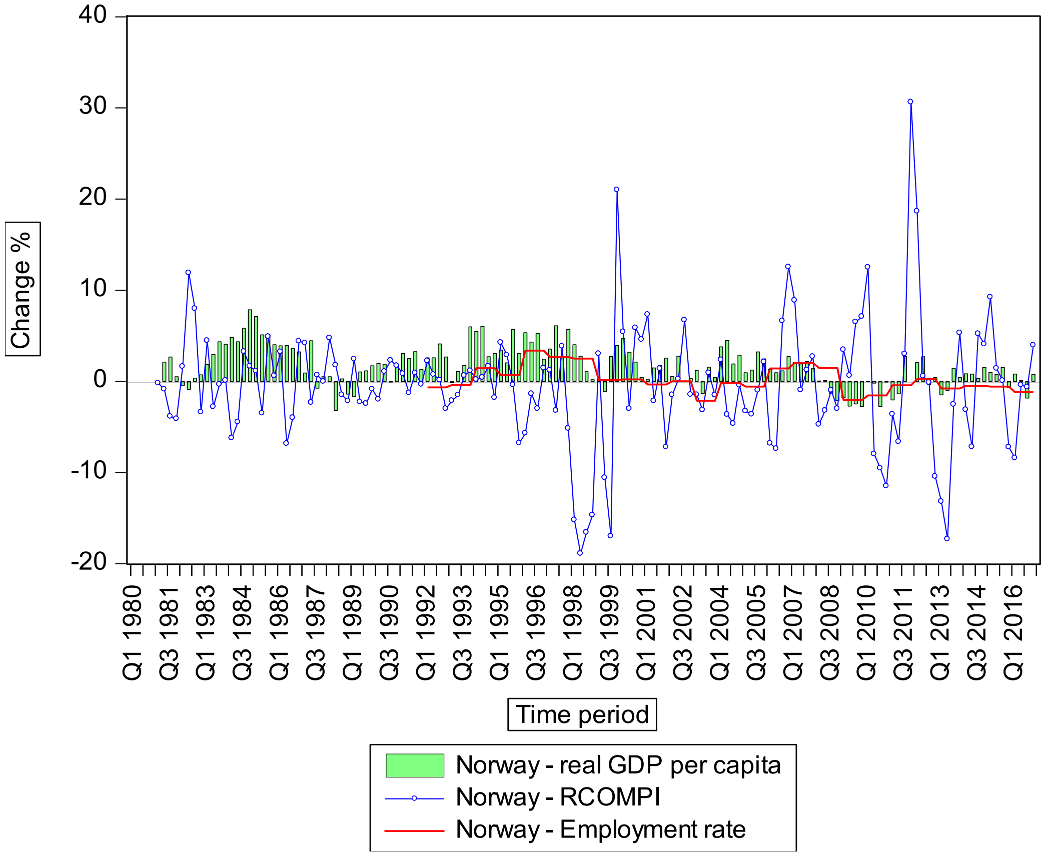

Figure 2.13 National Aggregates and RCOMPI – Norway

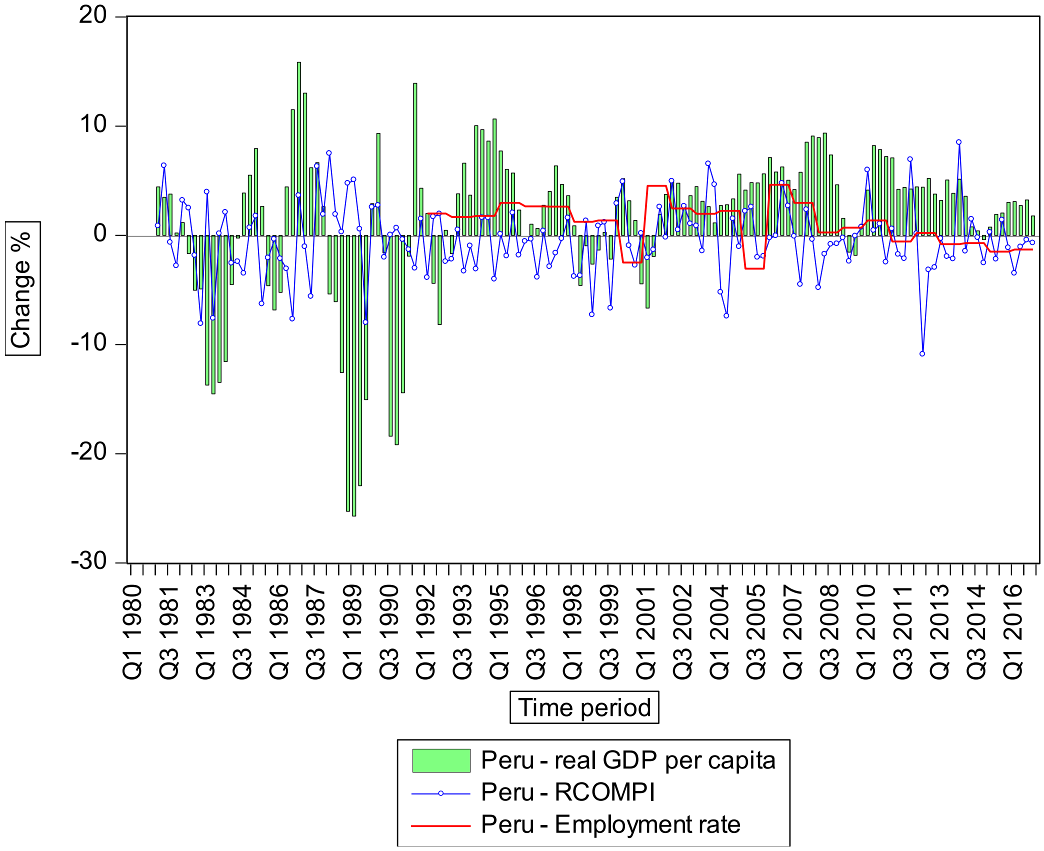

Figure 2.14 National Aggregates and RCOMPI – Peru

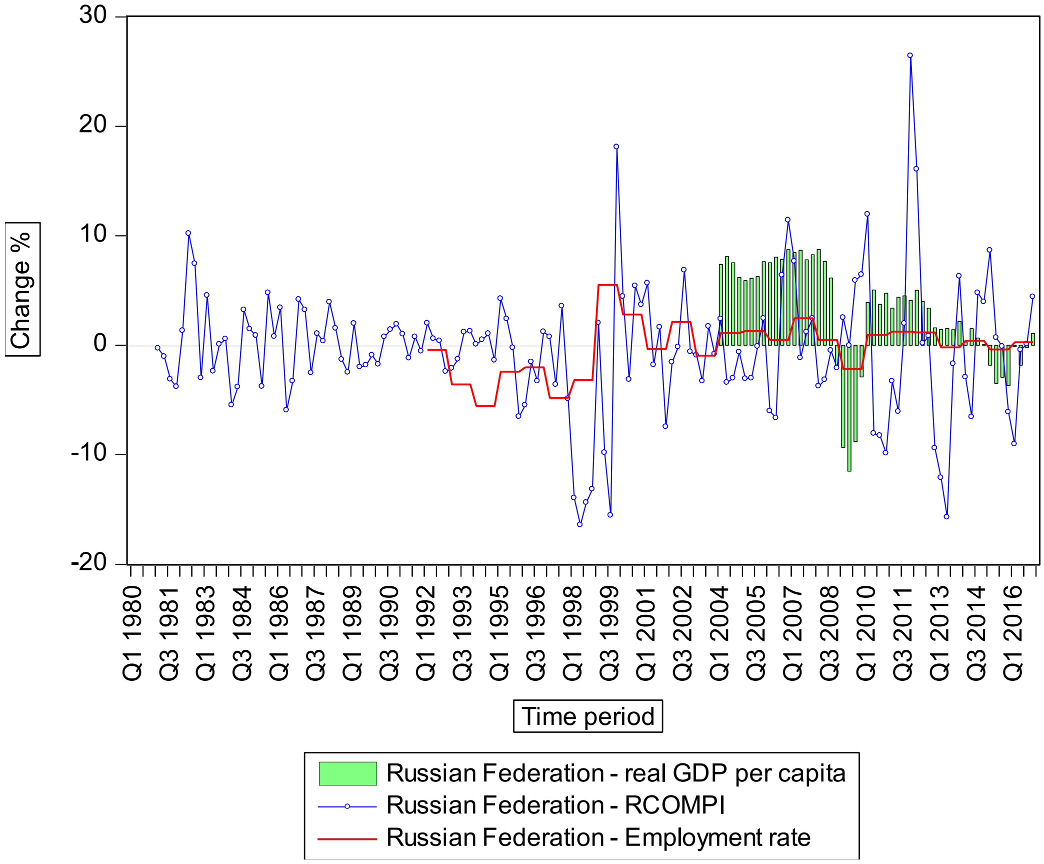

Figure 2.15 National Aggregates and RCOMPI – Russian Federation

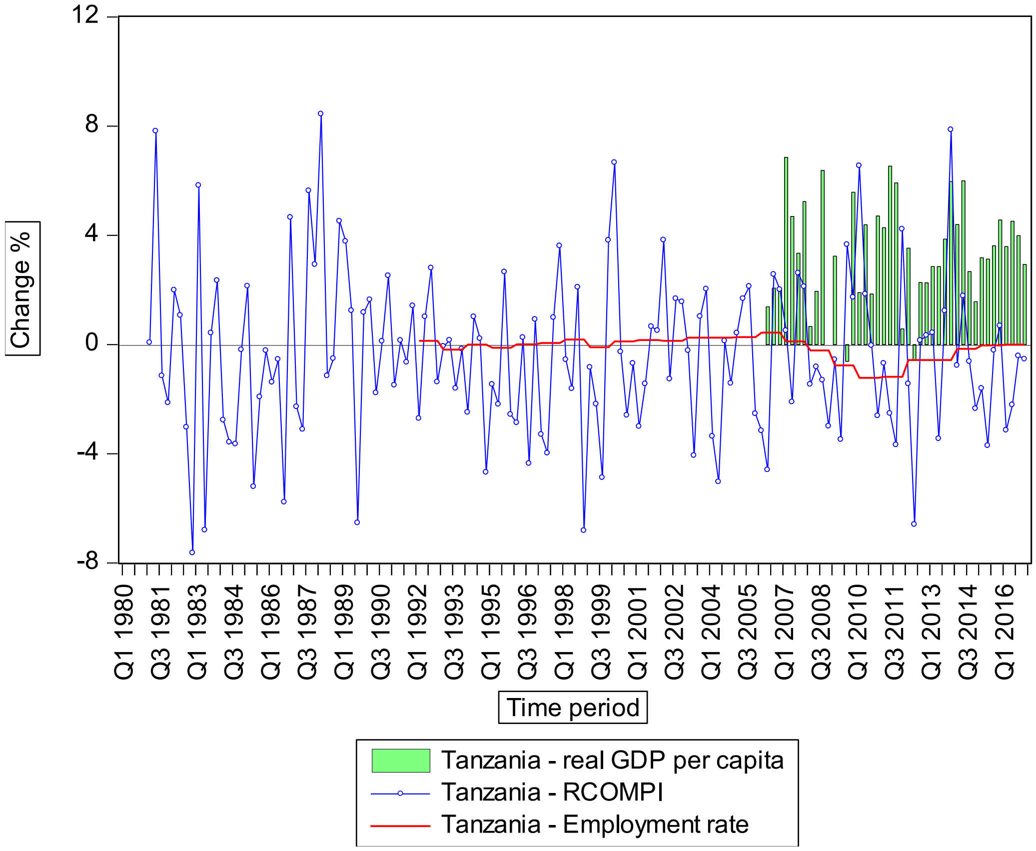

Figure 2.16 National Aggregates and RCOMPI – Tanzania

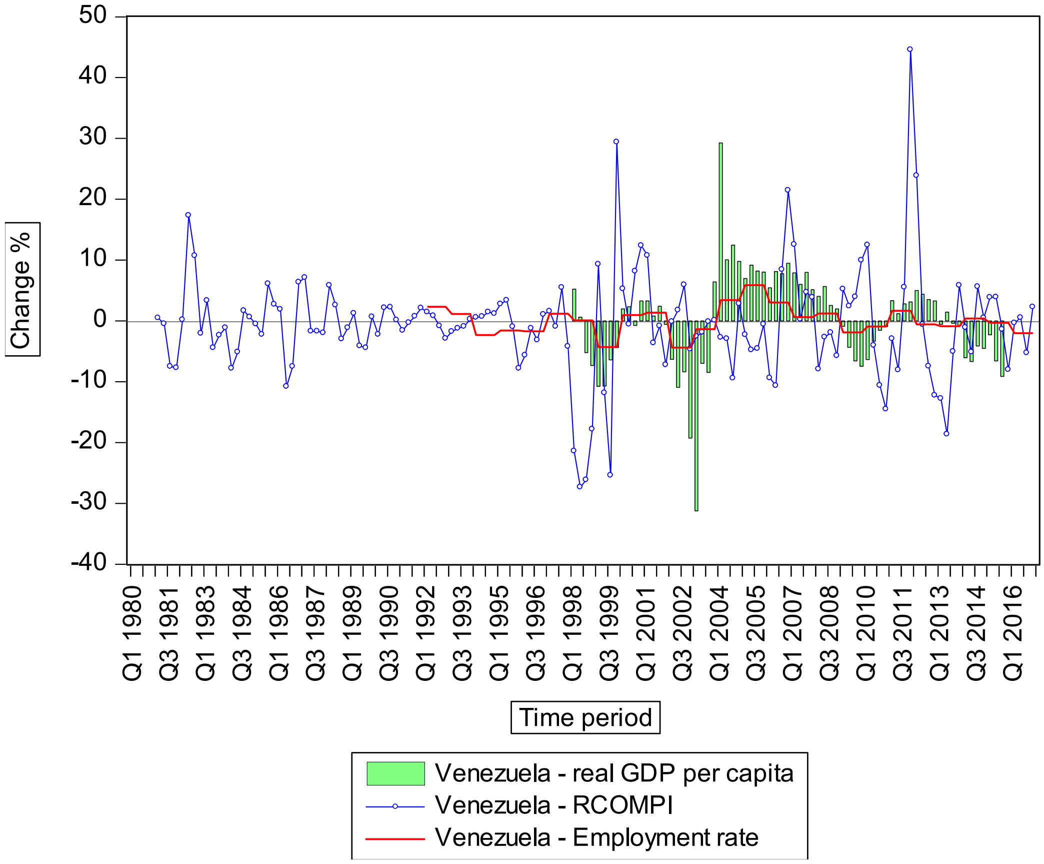

Figure 2.17 National Aggregates and RCOMPI – Venezuela

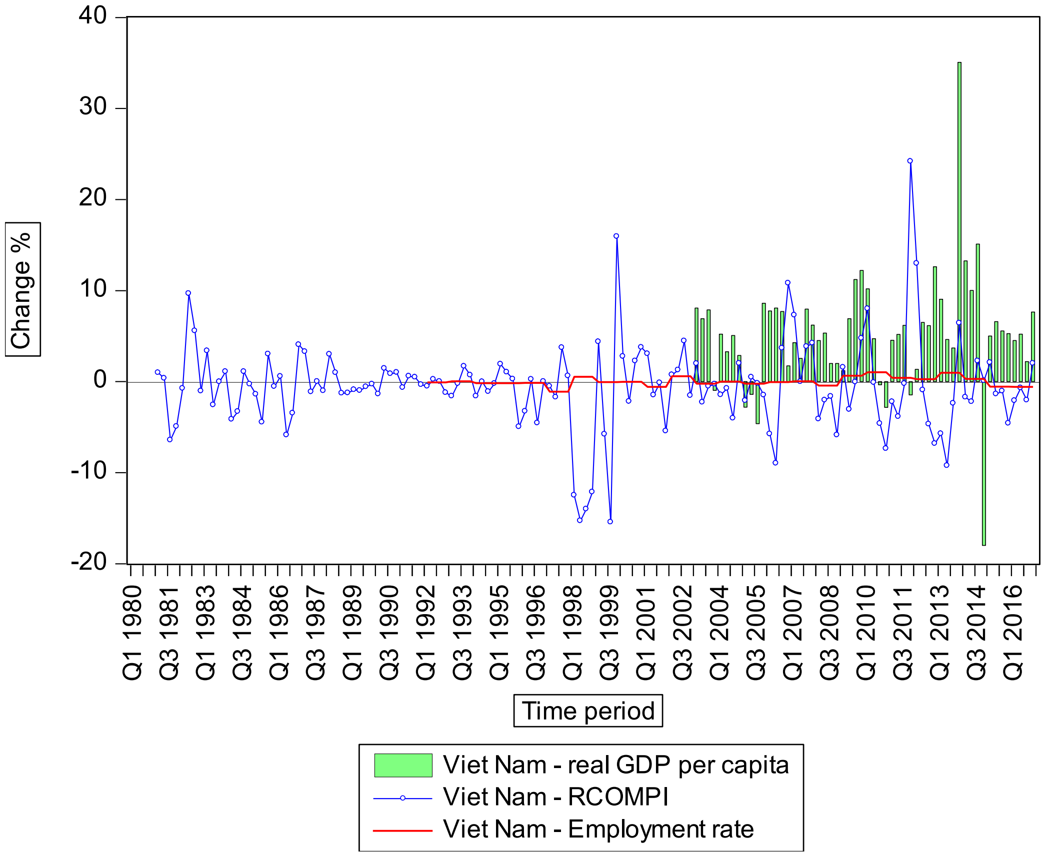

Figure 2.18 National Aggregates and RCOMPI – Viet Nam

To sum up, this study provides a large world database of national commodity export price indexes, which reliability we demonstrate empirically throughout a rigorous testing procedure.

2.7 Conclusions

This study presents the first attempt in the existing literature to create a monthly world database of national commodity export price indexes where the number of time series observations and the number of countries are both large. Specifically, our database has a total sample of 217 countries, where each index series contains 448 observations. Thus, our research provides a substantial contribution to economic literature by establishing a convenient starting point for an empirical analysis in a *data-rich environment*.

More broadly, this study contributes to the literature by creating a new robust framework for index construction. On the one hand, our study identifies the lack of consistency between export values in different revisions of trade classification systems. We show that this can create imprecise movements in the national commodity export indexes. To correct for this, we suggest a solution by providing a new data collection process that synchronises information from various revisions of the trade classification systems. On the other hand, our study uses a large basket of primary commodity products, including dairy, which improves the accuracy of the index series. The large commodity index basket also predisposes the creation of national commodity export price sub-indexes. Therefore, further contribution of our study is the construction of national commodity export price sub-indexes for each country in our database.

More precisely, our database provides monthly national commodity export indexes for the following 13 categories: (1) All commodities, (2) Non-energy commodities, (3) Food, (4) Cereals, (5) Vegetable oils and protein meals, (6) Meat, (7) Dairy, (8) Beverages, (9) Agricultural raw materials, (10) Metals, (11) Energy, (12) Fertilizers and (13) Precious Metals. The index series cover the period from January 1980 to April 2017. This reduces the overhead of macro-econometric modelling by providing the researcher with a favourable working environment. In addition, we use a consistent methodology that facilitates the replication and comparison of the results.

Last but not least, this study emphasises the plausibility of the index formula proposed by Deaton and Miller (1995) for constructing a world database of national commodity export indexes. Employing numerous empirical tests, we have demonstrated that our index series are highly correlated with the official index series provided by the central and commercial banks. This provides some relief in terms of the accuracy of our index series. After all, we can conclude that the data availability and the choice of index formula are both are inimitable ingredients for the construction of a *data-rich* database of national commodity export price indexes.

Chapter 3. Do Commodity Prices Predict Economic Growth? A Mixed-Frequency Time-Varying Investigation

3.1 Introduction

Commodity markets play an important role over the course of the economic cycle (Sachs and Warner, 1999; 2001). Stock and Watson (2003) highlight that commodity prices are forward-looking economic variables, which make them a class of potentially useful predictors of future economic growth. The interest in commodity prices as leading indicators of output dates back to the 1970s, which is a period of growing dependence on imported oil and poor macroeconomic performance in industrial countries, such as Australia, Canada, Japan, New Zealand and the US (Barsky and Kilian, 2004). Therefore, separating international macroeconomics and trade can result in a failure to account for the influence of commodity price dynamics on economic development as well as the aggregate feedback effects of commodity price patterns (see Deaton and Miller, 1995 for discussion). This chapter contributes to the existing body of literature by shedding new light on the causal relationship between commodity prices and economic growth. In particular, we explore the causal patterns between world commodity prices and economic growth for a sample of 33 commodity-dependent countries between January 1980 and December 2016.

In order to investigate the role of commodity prices on economic growth, this study presents a cross-country analysis by classifying countries into the following groups: commodity exporters, commodity importers and “hybrid” economies.[51] This allows for an investigation of the effects of changes in commodity prices on economic growth in relation to the country’s dependency on commodity exports, imports or both. This has important implications for the design of the country’s trade policy. For example, countries that adopt free-market trade policies may also adopt free-market domestic policies and stable fiscal and monetary policies (Frankel and Romer, 1999). However, free-market trade policies, such as lower tariff and non-tariff trade barriers, are significantly associated with economic growth, as highlighted by Rodriguez and Rodrik (2000). Subsequently, countries’ trade policies are likely to be correlated with factors that are often omitted from the income equation, such as commodity prices (Barro et al., 1991). Thus, one needs to identify whether world commodity prices have an impact on economic growth. Providing insights on this question is a contribution of this study.

A large body of literature is devoted to the study of commodity prices and its effects on macro-economic variables such as inflation (see Beckerman and Jenkinson, 1986; Boughton

[51] For a more detailed explanation of countries’ taxonomy, see Section 3.5.

and Branson, 1992; Cecchetti and Moessner, 2008; Gospodinov and Ng, 2013; Chen et al., 2014), exchange rates (see Chen and Rogoff, 2003; Cashin et al., 2004; Chen et al., 2010; Gopinath et al., 2010; Ferraro et al., 2015) and interest rates (see Sargent, 1969; Roll, 1972; Akram, 2009). Attention has also been given to the effects of commodity prices, mainly of oil prices, on economic growth.[52] For example, Hamilton (1983) concludes that escalations in oil prices are responsible for declines in the US Gross National Product, while Hamilton (1996) finds that oil prices Granger-cause the US Gross Domestic Product (GDP). Deaton (1999) asserts that economic growth in African economies remains heavily dependent on exports of primary commodities; as such, these economies do better when the prices of commodities are rising rather than when they are falling. A recent study by Collier and Goderis (2008) argues that half of the current growth of Africa's commodity-exporting economies is attributable to the short-term effects of the commodity price boom. While past studies have shown a substantial interest in this topic, the following gaps have been identified.

To begin with, the existing literature has mainly focused on the oil market; however, not all commodity-dependent countries are reliant on oil, as discussed by Deaton and Miller (1995). Cashin et al. (2004, p. 245) state, "Few exporters of nonfuel commodities are so specialized that the export prices of a single commodity can well approximate movements in an index of commodity export prices based on the export baskets of individual commodity-exporting countries." The same conclusion validity also pertains to the international commodity market where, in 2011, fuels accounted for 52% of all commodity exports, minerals, ores and metals for 20% in total commodity exports and the share of agricultural commodity exports is 28% (UNCTAD, 2018). As can be seen, the role that non-fuel commodities play in the international commodity market is substantial and, therefore, their inclusion in the forecasting models may improve the predictive ability of commodity prices in general. Therefore, this study considers an index of world commodity prices that contains commodities from all commodity groups; specifically, both fuel and non-fuel commodities have been included. Particularly, this analysis aims to provide new evidence on whether world commodity prices as a whole predict economic growth in commodity-dependent countries.

A further contribution of this study is its use of a novel econometric method that overcomes the problems associated with temporal aggregation (Ghysels, 2016). In particular, a problem with analysing the link between commodity prices and economic growth is the data sampled

[52] See, among others, Deaton and Miller (1995), Deaton (1999), Dehn (2000), Blattman et al. (2007), Hausmann et al. (2007), Collier and Goderis (2008), Alexeev and Conrad (2009), Beny and Cook (2009), Torvik (2009), Berument et al. (2010), Frankel (2010), Collier and Goderis (2012), Narayan et al. (2014), Cavalcanti et al. (2015), Mohaddes and Raissi (2017).

at different frequencies. Since classical models require all variables to have a single-frequency, the variables typically available at a high-frequency, such as commodity prices, are often aggregated at the lowest frequency.[53] However, recent research has shown that temporal aggregation has an adverse impact on statistical inference (see Orcutt et al., 1968; McCrorie and Chambers, 2006; Andreou et al., 2010; Eraker et al., 2015; Schorfheide and Song, 2015; Chambers, 2016; Ghysels, 2016; Ghysels et al., 2016). In fact, given that commodity price data is known to be highly volatile (see Deaton, 1999; Deaton and Laroque, 1992), working with a common low-frequency approach is likely to omit useful information about the variables (Götz et al., 2016). In other words, choosing low-frequency (e.g. quarterly) commodity price data may lead to the loss of information in the empirical models (Adams et al., 1979; Garner, 1989). Therefore, forecasting performance might improve by making use of the extra information contained in the high-frequency observations, as highlighted by Götz et al. (2014). Additionally, the finite-sample power of testing procedures may fall when temporally aggregated data are used due to the smaller number of available observations (Marcellino, 1999). Therefore, this study adopts the mixed-frequency vector autoregressive (MF-VAR) approach of Ghysels et al. (2016) to prevent the loss of information from temporal aggregation. The MF-VAR model allows both monthly and quarterly frequency variables to be estimated together in the same model.

To emphasise, a key point is that Granger causality in a VAR framework is not invariant to temporal aggregation (see Granger and Lin, 1995; Marcellino, 1999). To put it another way, McCrorie and Chambers (2006) claim that spurious Granger causality relationships can arise from temporal aggregation. Similarly, Rossana and Seater (1995, p. 441) state, "The observed time series behaviour of temporal aggregated data is not a reliable guide to the true cyclical properties of the underlying economy." Therefore, one can assume that the causal patterns discovered by the low-frequency (LF) model, and not the mixed-frequency (MF) model, are largely a result of temporal aggregation bias – not due to the true properties of the underlying data.[54] This study overcomes the issues from temporal aggregation by using the MF-VAR procedure of Ghysels et al. (2016).

Another aspect of the commodity-growth literature that has been left mostly unattended is parameter stability. Hansen (1992) and Lin and Teräsvirta (1994) highlight that the parameter non-constancy may have severe consequences on statistical inference if left unattended.

[53] The economic growth data is available at frequency equal to or lower than quarterly, especially for developing countries and for long historical time series (37 years in our case).

[54] The problem of measurement errors is a classic issue in statistical theory. For more details, see Fuller (1987), Wansbeek and Meijer (2000) and Carroll et al. (2006).

Further, Lee and Chang (2005) claim that, when there are structural breaks, the various test statistics are biased towards the non-rejection of a null hypothesis. Whereas the past studies have mainly used a standard linear framework to examine the linkage between commodity prices and economic growth, several studies have identified structural breaks in the commodity-growth relationship, such as those of Hamilton (1983), Hamilton (1996), Hooker (2002), Cuñado and De Gracia (2003; 2005) Narayan et al. (2014) and Cavalcanti et al. (2015). Therefore, this study contributes to the literature by investigating the stability of the commodity-growth relationship. Precisely, the Andrews' (1993) QLR structural break test is applied, which confirms the presence of parameter instability in the commodity-growth relationship.

Given these conditions, we apply the MF-VAR and LF-VAR models and conduct a battery of Granger causality tests using a time-varying framework in order to analyse the dynamic nature of the commodity-growth relationship. This approach yields interesting results. For instance, the world commodity prices are found to predict economic growth for 21 out of total 33 countries by the time-varying LF method, while the time-varying MF method finds predictability for 31 out of 33 countries. This highlights that there are around 30% fewer cases when causality is determined by the LF method as compared to the MF method. This result is consistent with the study of Ghysels et al. (2016) who point out that the MF test achieves higher power than its LF counterpart.

Moreover, the in-sample predictive ability often fails to translate into out-of-sample success, as highlighted by Timmermann (2006). Although this is a widely documented pattern in the forecasting literature, this study provides solid evidence in support of the commodity price out-of-sample predictive power for economic growth. In particular, we use three benchmark forecasting models to evaluate the out-of-sample forecasting performance of our commodity price-based models, namely an autoregressive, a random walk and a random walk with drift. The forecasting combination results show that the commodity-based models outperform the benchmark models for 79% of the total number of countries for at least two of the three benchmarks, according to both the LF and MF methods. This result builds upon the study of Narayan et al. (2014), who find evidence of out-of-sample predictability of oil prices on economic growth for around 70% of the countries. This finding confirms our assumption that the commodity market, as a whole, does not play a less significant role in economic growth than the oil market.

Last but not least, the robustness of the results is confirmed in several ways. First, the study presents the outcomes under different lag-order scenarios. Second, a measure of national commodity export prices is used to examine the impact of national commodity prices on economic growth for a selected number of countries. The national index series are obtained from Chapter 2. Third, different synthetic measures of world commodity prices are used to confirm that our findings are immune against the selection of proxy for world commodity prices. The five different proxies of world commodity prices are Reuters/Jeffries, Goldman Sachs, Moody's, Thompson Reuters Core Commodity Equal Weighted and IMF non-fuel commodity price indexes. The index series are selected so that they differ in terms of construction and commodity basket. In particular, the robustness check aims to confirm the main findings of this study in the context of the role that commodity prices have in forecasting economic growth.

The rest of this chapter is organised as follows. The review of the literature is presented in Section 3.2, while a conceptual framework is provided in Section 3.3. The econometric methodology in terms of both in-sample and out-of-sample methods is presented in Section 3.4. Data information is provided in Section 3.5. Next, Section 3.6 explores the empirical results in terms of world commodity prices. A brief discussion about the robustness of the results is presented in Section 3.7. Finally, Section 3.8 concludes the chapter.

3.2 Literature Review

Commodity prices have been found to be one of the most prominent determinants of economic growth, especially in developing countries (see Mendoza, 1997; Collier and Gunning, 1999; Deaton, 1999; Temple, 1999; Frankel, 2010 for discussion). As such, the commodity-growth relationship has been examined by a myriad of published research. As the main objective of this study is to determine whether there exists a causal relationship between commodity prices and economic growth, it is essential to first explore the literature that investigates the existence of such a relationship in general.

The effect of natural resources on economic growth has long been a debated topic. Certain studies observe the effect of commodities on economic growth by considering the volatility of commodity prices (see Bleaney and Greenaway, 2001; Blattman et al., 2007; Cavalcanti et al., 2015; Mohaddes and Raissi, 2017), while others focus on the impact of growth rates in commodity prices (see Deaton and Miller, 1995; Dehn, 2000; Raddatz, 2007; Collier and Goderis, 2008; 2012; Addison et al., 2016). However, the findings remain inconclusive as demonstrated by the discussion of the literature to follow.

Bleaney and Greenaway (2001) use fixed effects panel regressions to estimate the impact of level and volatility of terms of trade on economic growth for the period between 1980 and

1995. The sample of countries includes only 14 developing sub-Saharan African economies, while a larger sample is not considered due to data limitations.[55] Using annual data, they find that volatility in the terms of trade has a negative impact on growth, while growth depends positively on the current level of the terms of trade and negatively on the lagged change.[56] In addition, the choice of countries is not arbitrary. The focus on the sub-Saharan African region reflects the intuition that if the terms of trade volatility matters, it should do so in countries most dependent on primary commodities. This is because the high volatility of the world commodity prices causes severe volatility in the output per capita growth in countries that depend heavily on primary commodities, as discussed by van der Ploeg and Poelhekke (2009).

The negative effect of terms of trade volatility on economic growth found by Bleaney and Greenaway (2001) is consistent with findings obtained in several other studies. Blattman et al. (2007) investigate the impact of terms of trade volatility on economic growth during the period between 1870 and 1939.[57] The authors use a larger sample of countries than Bleaney and Greenaway (2001) – 35 developing and developed economies from around the world. Using annual data in a standard ordinary least squared (OLS) regression framework, Blattman et al. (2007) find no statistically significant relationship between the terms of trade growth and income growth in the commodity-specialised Periphery.[58] Nonetheless, the authors conclude that the terms of trade volatility matters for the larger and diversified industrial nations. Particularly, the volatility has a negative impact on growth in these economies. In contrast, the Core results show evidence that income growth is positively correlated with the terms of trade growth. Therefore, the impact of terms of trade on economic growth is not alike for all countries.

A recent work conducted by Cavalcanti et al. (2015) examines the impact of growth and volatility of commodity terms of trade (CToT) on long-run economic growth between 1970 and 2007.[59] They use annual data for a panel data set of 118 countries that is split into two sets: (a) 62 primary commodity exporters and (b) 56 other countries that have more

[55] The sample includes Botswana, Burkina Faso, Cameroon, Côte d'Ivoire, The Gambia, Ghana, Kenya, Malawi, Mauritius, Niger, Senegal, Tanzania, Togo and Zimbabwe.
[56] The volatility of the terms of trade is estimated from a GARCH (1, 1) model.
[57] The terms of trade growth measure is defined as the percentage change in the trend in the terms of trade over the decade, while volatility is measured by the standard deviation of variations from this trend.
[58] The definition of "Periphery" nations, as stated by the authors, includes Denmark, Greece, Norway, Portugal, Russia, Sweden, Serbia, Spain, Australia, Canada, New Zealand, Argentina, Brazil, Chile, Colombia, Cuba, Mexico, Peru, Uruguay, Burma, Ceylon, China, Egypt, India, Indonesia, Japan, Philippines, Thailand and Turkey, while the "Core" nations includes Austria, France, Germany, Italy, the United Kingdom and the US.
[59] The CToT index is based on the index formula of Spatafora and Tytell (2009). It includes the prices of 32 primary commodities. The CToT volatility is a standard deviation of CToT growth in a five-year interval, while the CToT growth is measured as the growth rate of CToT index.

diversified export basket. To estimate the commodity-growth relationship, the authors use a standard system generalised methods of moments approach and a dynamic common correlated effects pooled mean group methodology that account for cross-country heterogeneity, cross-sectional dependence and feedback effects.[60] The main finding is that, while the CToT growth enhances real GDP per capita, the CToT volatility has a negative impact on economic growth in commodity-exporting countries. These results hold for the subsample of the 62 primary commodity exporters but not for the remaining 56 countries, which have a more diversified export basket. For the latter, the authors find that changes in commodity prices (or their volatility) do not have any major impact on their economies. This finding is somewhat consistent with the conclusions of Blattman et al. (2007) that the impact of terms of trade on economic growth is not alike for all countries.

A further study by Mohaddes and Raissi (2017) analyses the impact of CToT volatility on long-run economic growth for a sample of 69 commodity-dependent countries between 1981 and 2014.[61] The authors create an annual panel data set (by averaging monthly data) and use a panel Cross-Sectionally augmented Autoregressive Distributive Lag approach to account for the endogeneity, cross-country heterogeneity and cross-sectional dependence that arise from unobserved common factors. Similar to Cavalcanti et al. (2015), the authors find that the CToT growth enhances the real output per capita, while the CToT volatility exerts a negative impact on economic growth.

Despite the accumulation of evidence for the effect of terms of trade volatility on economic growth, the validity of these outcomes may be challenged by other researchers. On the one hand, the above-mentioned studies use annual data to calculate volatility. However, commodity prices are known to be highly volatile (see Deaton, 1999; Deaton and Laroque, 1992) and, therefore, averaging itself induces a loss of information. Further, Cavalcanti et al. (2015) highlight that using year averages could underestimate the importance of volatility. On the other hand, terms of trade may be an imprecise measure of commodity price movements (see Deaton and Miller, 1995; Cashin et al., 2004) because the terms of trade indexes are typically calculated using export and unit values, which are affected by the composition of exports and, consequently, by the composition of GDP (Deaton and Miller, 1995). Another branch of studies overcomes these issues by analysing the level (mean) relationship between commodity prices and economic activity.

[60] The size-adjusted power gain from using the generalised least squares (GLS) test statistic over the OLS test statistic in small sample sizes is estimated to be around 20% (Westerlund and Narayan, 2012).
[61] The CToT index is based on the index formula of Spatafora and Tytell (2009). It includes the prices of 45 primary commodities. The CToT volatility is constructed as the standard deviation of the year-on-year growth rates of CToT.

Empirical studies by Deaton and Miller (1995) for Africa and Raddatz (2007) for low-income countries use panel vector autoregressive (VAR) models and reveal that higher commodity prices significantly raise income in the short-run.

In an influential paper, Deaton and Miller (1995) study the relationship between national commodity export prices (as discussed in Chapter 2) and economic growth in 32 sub-Saharan African countries in the period of 1958–1992.[62] Using annual data, the authors find zero or a negative correlation between commodity prices and economic growth for 14 countries, namely Benin, Botswana, Burkina Faso, Kenya, Liberia, Malawi, Mali, Mauritania, Mauritius, Niger, Rwanda, Senegal, Sudan and Uganda. In addition, they estimate a VAR for each of the 32 countries, including the lagged rate of growth of GDP and the lagged rate of growth of commodity prices. The VAR results suggest evidence for a commodity-growth relationship for only four countries (for the Central African Republic, Ghana, Liberia and Mauritania).

A further study by Raddatz (2007) constructs an annual DM index to fit a panel VAR model and examines the effect of commodity price shocks on economic growth for 40 low-income countries between 1965 and 1997.[63] The author finds that one standard deviation shock to the commodity prices corresponds to a 14% increase in the commodity prices with respect to their baseline level and results in a significant 0.9% increase in the GDP after four years. However, the author does not attempt to disentangle the different sources of shocks and their relative importance – positive vs. negative shocks, for example.

In contrast to the above studies, Dehn (2000) uses a pooled OLS to examine the effects of both positive and negative shocks on economic growth in a large sample of 113 countries for the period from 1957Q1 to 1997Q4. Using quarterly data, the author finds that per capita economic growth is significantly reduced by the negative commodity price shocks, while positive commodity price shocks do not exert influence on economic growth. As such, the positive commodity price shocks have no long-run impact on growth. This result is supportive

[62] The sample of countries includes Benin, Botswana, Burkina Faso, Burundi, Cameroon, Central African Republic, Congo, Ethiopia, Gabon, Gambia, Ghana, Ivory Coast, Kenya, Lesotho, Liberia, Madagascar, Malawi, Mali, Mauritania, Mauritius, Niger, Nigeria, Rwanda, Senegal, Sierra Leone, Sudan, Tanzania, Togo, Uganda, Zaire, Zambia and Zimbabwe.

[63] The sample of countries includes Angola, Bangladesh, Benin, Burkina Faso, Burundi, Cameroon, Central African Republic, Chad, Democratic Republic of Congo, Republic of Congo, Cote d'Ivoire, Ethiopia, Gambia, Ghana, Guinea, Guinea-Bissau, Haiti, Honduras, India, Indonesia, Kenya, Lesotho, Madagascar, Malawi, Mali, Mauritania, Mozambique, Nepal, Nicaragua, Niger, Nigeria, Pakistan, Rwanda, Senegal, Sierra Leone, Tanzania, Togo, Uganda, Zambia and Zimbabwe. The region with highest presence in the country sample is sub-Saharan Africa, with 32 countries, followed by South Asia with four, Latin America and the Caribbean with three and East Asia and Pacific with one.

to the studies of Deaton and Miller (1995) and Raddatz (2007), who find evidence of a positive effect of commodity prices on income in the short-run.

In a more sophisticated study, Addison et al. (2016) apply an impulse response analysis, due to Kilian and Vigfusson (2011), to uncover whether positive or negative agricultural price shocks evoke a different response from economic growth. As such, Addison et al. (2016) build upon the study of Dehn (2000) by determining whether a positive commodity price shock has a larger effect on economic growth than a negative one. The authors conclude that there is negligible evidence to suggest that a positive price shock leads to a significantly different response in per capita income as opposed to a negative price shock and, therefore, the evidence of asymmetry is minimal. The authors conducted their analysis for 9 sub-Saharan African countries in the period of 1960–2010 using five different annual commodity prices from the Grilli and Yang (1988) database, namely cocoa, coffee, cotton, tea and tobacco.[64]

Further, the long-run relationship between national commodity prices and economic growth is confirmed by other studies, such as those of Collier and Goderis (2008; 2012).[65]

Collier and Goderis (2008) adopt a panel cointegration methodology to study the long-run effects of national commodity prices on economic growth. The study considers annual frequency data and covers the period from 1963 to 2003 for a sample of 129 countries. Notably, the authors decompose the annual commodity price index into two sub-indexes: agricultural and non-agricultural. For general commodity prices, the authors find that high commodity export prices reduce the long-run economic growth in commodity-exporting countries. The same conclusion is obtained for non-agricultural commodity export prices. In contrast, the agricultural commodity prices are found to have a positive (and insignificant) effect on long-run economic growth. This finding emphasises that the choice of proxy for commodity prices matters.

Further, Collier and Goderis (2012) indicate that an increase in the commodity prices has positive effects on output in the short-run; however, the effect on the output largely depends on the type of commodity and the quality of governance in the long-run. They find that an increase in non-agricultural commodity prices has a negative effect on the long-run economic growth in countries with poor governance, whereas countries with sufficiently good governance do not suffer from this adverse effect. They instead may even benefit from higher

[64] These countries are Benin, Burkina Faso, Burundi, Cameroon, Cote d'Ivoire, Ghana, Malawi, Kenya and Rwanda.
[65] The commodity export price indexes are constructed using the methodology of Deaton and Miller (1995).

commodity prices both in the short- and long-run. The analysis includes 37 commodity-exporting countries over the period of 1963–2008.

A major drawback of these two studies is that the dynamics are common across cross-sectional units and the country specific dynamics cannot be estimated due to the limited time series data available, as discussed by Addison et al. (2016).

Thus far, the literature has shown that although the commodity-growth relationship has been heavily researched, it is not enough known whether commodity prices have any predictive power for economic growth. This embodies another strand of literature that owes much to the earlier work of Hamilton (1983). Typically, these studies fit a predictive regression model of economic growth wherein commodity prices appear as a predictor variable. Sublime studies that investigate the causal commodity-growth relationship include that of Stock and Watson (2003; 2004).

Stock and Watson (2003) adopt an h-step ahead forecast approach to examine the predictive performance of commodity prices for inflation and real output growth. They undertake an empirical analysis of quarterly data on 38 indicators (mainly commodity prices) for seven developed OECD countries (namely Canada, France, Germany, Italy, Japan, the United Kingdom and the US) for the period of 1959Q1–1999Q4. At first, the authors use QLR statistics to examine the structural stability of the estimated models and find instability in the output forecasts for all countries. Then, the authors fit projection regressions and find that commodity prices have small predictive content for output at the second, fourth and eighth quarter horizon.

More recent study by Stock and Watson (2004) uses forecast combination methods to forecast output growth for seven developed OECD countries (namely Canada, France, Germany, Italy, Japan, the United Kingdom and the US). The authors employ a quarterly economic data set that covers the period of 1959Q1–1999Q4, with up to 73 predictors per country (i.e. unbalanced panel). They find that the forecasts based on individual predictors are unstable over time and across countries and, on average, perform worse than the AR benchmark. However, the combination forecasts often improve upon the AR forecasts. This finding is consistent with the past study of Stock and Watson (2003).

Cuñado and De Gracia (2003) study the impact of oil price changes on both inflation and industrial production growth in a quarterly data set of 15 countries that covers the period from

1960 to 1999.[66] The authors estimate a VAR model for each of the 15 developed countries and find that growth rates of oil prices possess in-sample predictive power on industrial production growth for only seven countries, namely Belgium, Greece, Ireland, Luxembourg, Netherlands, Sweden and the United Kingdom. However, no evidence of a long-run relationship between oil prices and economic activity is found for either of the countries in the sample. This suggests that the impact of oil prices on economic activity is limited only to the short-run.

In a further study, Cuñado and De Gracia (2005) use quarterly data to investigate the causal relationship between oil prices, economic activity and consumer price indexes for six Asian developing countries (namely Japan, Malaysia, Philippines, Singapore, South Korea and Thailand) from 1975Q1 to 2002Q2.[67] The authors find that oil prices Granger-cause economic activity in Japan, South Korea and Thailand, while no evidence of causality is discovered for the other countries in the sample. Similar to their previous study, the authors are unable to provide evidence for a long-run relationship between oil prices and economic activity, which suggests that the impact of oil prices is limited to the short-run.

In a work conducted by Pradhan et al. (2015), a panel VAR model is used to identify the possible causality between economic growth, oil prices, stock market depth and three other macroeconomic variables (exchange rate, inflation and interest rate). The study considers an annual frequency series and covers the period from 1961 to 2012 for a sample of all G-20 countries.[68] The authors do not find any evidence to support the short-run Granger-causal relationship between economic growth and oil prices.

Most closely related in motivation to this chapter is Narayan et al. (2014) who examine the causal relationship between oil prices and economic growth for a sample of 45 countries between 1983Q2 and 2010Q4.[69] The authors use quarterly data for the estimation of bias-adjusted OLS and GLS regressions. The main difference between the two approaches is that the former estimator accounts only for persistency and endogeneity, while the latter estimator is flexible enough to cater for persistency, endogeneity and heteroskedasticity. In fact, the use

[66] The countries included in the study are Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden and the United Kingdom.
[67] In fact, the economic activity is proxied by the industrial production index for Japan and South Korea, by the Manufacturing Production index for Singapore and by the real GDP for Malaysia, Philippines and Thailand.
[68] As of 2015, there are 20 members of the G-20 group: Argentina, Australia, Brazil, Canada, China, the European Union, France, Germany, India, Indonesia, Italy, Japan, Mexico, Russia, Saudi Arabia, South Africa, South Korea, Turkey, the United Kingdom and the US.
[69] The sample of countries include (1) developed countries – Australia, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Hong Kong, Hungary, Iceland, Ireland, Israel, Italy, Japan, Malta, Netherlands, New Zealand, Norway, Portugal, Singapore, South Korea, Spain, Sweden, Switzerland, the United Kingdom and the US and (2) developing countries – Argentina, Brazil, Chile, China, Colombia, Croatia, India, Indonesia, Malaysia, Mexico, Philippines, Poland, Russia, South Africa, Sri Lanka, Thailand and Turkey.

of methods that account for heteroskedasticity is important because commodity price data are indeed heteroskedastic (see Bollerslev, 1987; Bernard et al., 2008). The existing heteroskedasticity leads to biased standard errors and, thus, biased inference, so results of hypothesis tests are possibly wrong. Therefore, we adopt Gonçalves and Kilian's (2004) recursive design parametric wild bootstrap that is robust to conditional heteroskedasticity of an unknown form.

Moreover, Narayan et al. (2014) find in-sample prediction from nominal oil prices to economic growth for 37 countries, while the results from real oil prices suggest in-sample predictability for 36 countries.[70] In contrast, the evidence for out-of-sample predictability is much weaker. Particularly, two of the three out-of-sample evaluation techniques agree that nominal oil prices have out-of-sample predictability for economic growth for 26 countries at forecasting horizon of one quarter. Whereas, this evidence is found for only 16 countries when real oil prices are considered.[71] Regardless the forecasting horizon, the authors find that nominal oil prices have out-of-sample predictability for economic growth for 34 countries, while real oil prices for only 22 countries.

Recent studies on predictability of commodity prices on economic activity (such as Cuñado and De Gracia, 2003; 2005; Narayan et al., 2014; Pradhan et al., 2015) have mainly focused on the crude oil market. A major drawback of these studies is the assumption that all countries are dependent on a single commodity – in this case, crude oil. On the one hand, the economy of several commodity export-dependent countries relies on non-fuel primary commodity products, as discussed by Cashin et al. (2004). For example, the main export is agricultural commodities for Argentina (soybean meal – 15%, corn – 6.8%, soybean oil – 6.6%, soybean – 4.8%, wheat – 4.3% of total exports in 2017), is copper for Chile (almost 50% of total exports in 2017), is gold for Hong Kong (19% of total exports in 2017), while the major imports for these countries is manufacturing products. In addition, Deaton (1999) highlights that many sub-Saharan African countries are export-dependent on metals and agricultural commodities. On the other hand, numerous countries are exporters or importers of crude oil but fuel products do not have large share in their export/import baskets. Examples include newly industrialised economies such as Australia and New Zealand (see Chen and Rogoff, 2003 for

[70] The oil prices are not found to in-sample predict economic growth in the case of (1) nominal oil prices – Greece, Hong Kong, Iceland, Italy, Malaysia, Norway, Sweden and the US and (2) real oil prices – Brazil, China, Mexico, Norway, Poland, Russia, South Africa, Turkey and the US.

[71] The oil prices are not found to have the out-of-sample predictive ability on economic growth in the case of (1) nominal oil prices – Canada, China, Croatia, Czech Republic, Denmark, Finland, Greece, Hong Kong, Indonesia, Ireland, Israel, Malaysia, Mexico, Poland, Russia, Singapore, South Africa, South Korea and Sri Lanka and (2) real oil prices – Argentina, Australia, Belgium, Brazil, Canada, Croatia, Greece, Hong Kong, Hungary, India, Indonesia, Ireland, Italy, Malaysia, Mexico, Philippines, Poland, Portugal, Russia, Singapore, South Africa, South Korea, Spain, Sri Lanka, Sweden, Switzerland, Turkey, the United Kingdom and the US.

discussion). For this reason, this study builds on Narayan et al. (2014) by not focusing on a single commodity, e.g. oil, because this will probably result in the underestimation of the impact of commodity prices on economic growth (see Deaton and Miller, 1995; Deaton, 1999). Therefore, we follow Stock and Watson (2003; 2004) by using the general commodity price index as a proxy for world commodity prices.[72]

As a case in point, the past commodity-growth studies have been found to aggregate the commodity price time series to a frequency that suits the requirements of their empirical models. In fact, the earlier studies have mainly used annual commodity data (see Cavalcanti et al., 2015; Pradhan et al., 2015; Mohaddes and Raissi, 2017), while there are some exceptions such as Dehn (2000) and Narayan et al. (2014) – they use quarterly commodity prices to analyse the relationship between commodity prices and economic growth. Nonetheless, commodity price data is known to be highly volatile (see Deaton, 1999; Deaton and Laroque, 1992) and, therefore, aggregating commodity price series is likely to omit useful information about the forecasting ability of the variables (Götz et al., 2016). Unfortunately, the issues that may arise from temporal data aggregation are often misjudged. As such, a possible reason behind the failure of past studies to find causality may be the aggregation of the high-frequency data to low-frequency, as highlighted by Ghysels et al. (2016). Particularly, we allow data of different frequency to be estimated in the same framework. As such, we overcome the issues that may arise from temporal aggregation. The next section provides more information about the conceptual approach employed in this chapter.

3.3 Conceptual Framework

Numerous theories have been developed to describe the commodity-growth relationship. Among them are the commodity wealth and the cost-of-production theories, which we consider to explain whether a change in commodity prices affects the amount of goods and services produced by an economy over time (see also Hamilton, 1996; 2003; 2009; Cuñado and De Gracia, 2003; 2005; Kilian, 2008; 2009). This section illustrates the two aforementioned theories, which are the main theoretical frameworks of this study, and discusses how they link to the cases of commodity-importing and exporting countries. The concept of these two channels is presented in figure 3.1.

[72] We acknowledge the potential importance of the commodity price volatility for predicting economic growth; however, in this study, we focus on the growth rates in commodity prices (consistent with Narayan et al., 2014). As such, we leave the commodity price volatility for future research work.

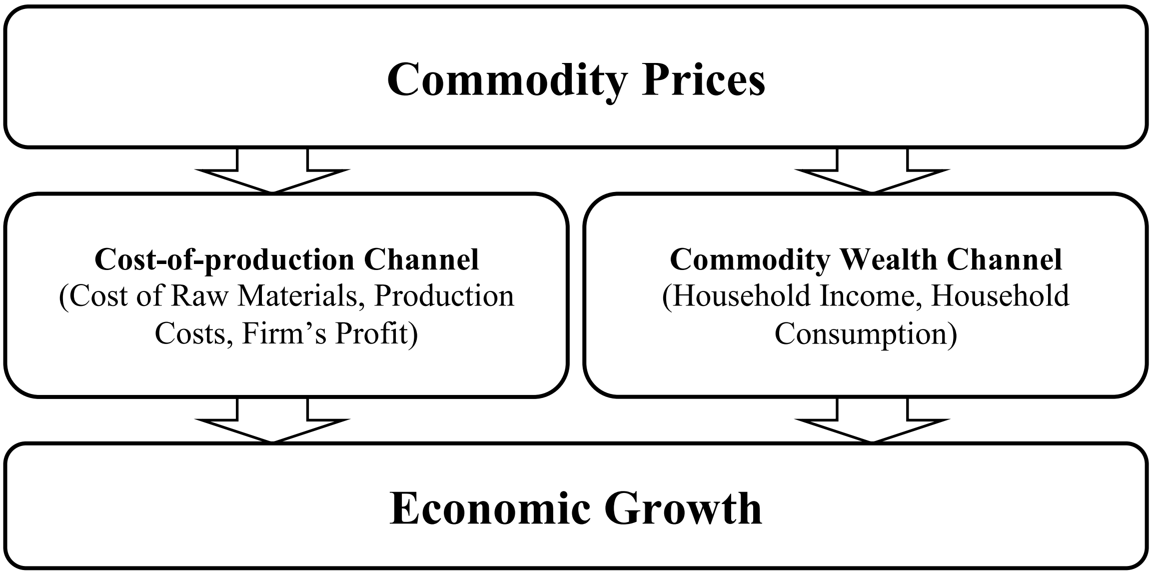

Figure 3.1 Commodity Prices and Economic Growth

Under this section, the thesis considers a small open economy model as of Ferraro and Peretto (2018) to theoretically describe the link between commodity prices and economic growth. The time is viewed as a continuous variable, which is indexed by subscript t, where $t \geq 0$. The specified model provides finer details for the link between commodity prices and economic growth.

3.3.1 Commodity wealth channel

The commodity wealth channel represents the notion that changes in economic growth may be a result of fluctuations in commodity prices in international markets (Arezki and Brückner, 2012). Thus, if disposable income decreases due to changes in commodity prices, households have less money to either save or spend, which naturally leads to a decline in consumption and, thus, it is very likely that economic growth will also slow (Svensson, 2005). Therefore, commodity prices directly affect household income by changing the value of the commodity endowment and, therefore, inducing income/wealth effects (Ferraro and Peretto, 2018).

Actually, the commodity wealth channel does not influence in the same way all economies but depends on whether an economy is an exporter or importer of commodities. Let us assume that there is an increase in the prices of commodities. On the one hand, the discretionary income of households in a commodity-importing country tends to decline due to the change in retail prices (Edelstein and Kilian, 2009). The higher expenditure on home consumption goods is determined by the increase in the price of primary commodity products; for example, fuels and agricultural products. This causes a decline in the household spending, which is the essential driving force of economic growth – it represents more than half of GDP in most developed economies (Chai, 2018). Thus, the economic growth is expected to slow down. On the other hand, the retail prices in a commodity-exporting country upsurge due to higher production costs. However, the household sector experiences a positive income/wealth effect, which is generally larger than the cost-of-production effect and, therefore, the economy grows faster (Degiannakis et al., 2018). A detailed theoretical setting is presented below.

Based on the theoretical framework of Ferraro and Peretto (2018), a small open economy is populated by a representative household that allocates disposable income on consumption of home and foreign goods and savings by borrowing and lending freely in a competitive market for financial assets at the current market interest rate (Nerlove, 1974). The household income comprises of labour incomes (WL), returns on investments (rA), profits (Π) and commodity income ($p\Omega$). The commodity income is specified as the commodity endowment (Ω) valued at the world commodity price (p). According to Ferraro and Peretto (2018), economies with a larger commodity endowment are commodity exporters for a larger range of commodity prices, whereas economies with no commodity endowment are constrained to be commodity importers. Therefore, the commodity wealth channel makes up the initial key transmission mechanism of the changes in the prices of primary commodities.

A simple model of commodity wealth channel is constructed in Ferraro and Peretto (2018) and is summarised here to embed the link between commodity prices and economic growth. Based on the theoretical model of Ferraro and Peretto (2018), we state that representative household chooses (Y_H, Y_F) to maximise the lifetime utility function $U(t)$, which is

$$\max_{\{Y_H, Y_F\}} U(t) = \int_t^\infty e^{-\rho(s-t)} \log u(s) ds \tag{3.1}$$

with

$$\log u = \varphi \log\left(\frac{Y_H}{P_H L}\right) + (1-\varphi)\log\left(\frac{Y_F}{P_F L}\right) \tag{3.2}$$

subject to the budget constraint

$$\dot{A} = rA + WL + \Pi_H + \Pi_M + p\Omega - Y_H - Y_F \tag{3.3}$$

where ρ is the discount rate, $\rho > 0$; φ controls the degree of home bias in preferences, $0 < \varphi < 1$; A is assets holding; r is the rate of return on financial assets, e.g. investments; W is the wage rate; L is population size; Y_H is expenditure on home consumption goods at price P_H; Y_F is expenditure on foreign consumption goods at price P_F; Π_H denotes the dividends paid by the producers of the home consumption goods; Π_M is the dividends paid by firms in the materials sector; Ω is the domestic commodity endowment. Based on the Ferraro and Peretto (2018) model, there is not preference for leisure, the population size (L) equals labour supply and, therefore, the labour income is WL.

Consequently, an increase in commodity prices (p) has twofold effect. On the one side, a household in a commodity-importing economy ($O > \Omega$) faces a reduction in the profit due to an increase in the cost of wages and, hence, the rate of return on financial assets declines, where O denotes the home use of the commodity. Simultaneously, higher commodity prices are likely to reduce the household consumption due to less labour income (WL). This leads to a reduction in aggregate consumption and, thus, the aggregate output decreases. On the other side, a representative household in a commodity-exporting economy ($O < \Omega$) experiences an increase in the cost of wages, but the income effect is stronger than the cost effect and, therefore, the profit increases together with the rate of return on investment. At the same time, the labour income (WL) raises due to the increase in the demand for labour. This leads to higher disposable income and faster economic growth (income effect).

Specifically, the economy can either be an importer or exporter of the commodity. On the one hand, a commodity importing economy sells the home consumption good (Y_H) to buy the commodity in the world market. On the other hand, a commodity exporting economy accepts the foreign consumption good (Y_F) as payment for its commodity exports. Based on Ferraro and Peretto (2018) framework, the foreign consumption good is imported at the exogenous and constant price P_F. Only final goods and the commodity are tradable. The balanced trade condition, which is also the market clearing condition for the consumption good market, is

$$Y = Y_H + Y_F + p(O - \Omega) \tag{3.4}$$

where Y is the aggregate value of consumption and the revenues from sales of the domestic commodity endowment are $p(O - \Omega)$. In the case of commodity-importing country ($O > \Omega$), higher commodity prices (p) lead to less demand for O, Y declines and, thus, aggregate output decreases. In the case of commodity-exporting country ($O < \Omega$), an increase in p leads to higher demand for O and Y increases. Hence, changes in commodity prices (p) affect the household consumption and, therefore, aggregate output increases.

Accordingly, an increase in the commodity prices tend to lower the discretionary income of households, due to the changes in retail prices (as a result of increased production costs), but also due to an increase in the prices of primary commodities (Edelstein and Kilian, 2009). Then, lower income leads to lower consumption and, thus, the aggregate output diminishes. Indeed, an increase in commodity prices will worsen the terms-of-trade for a commodity-importing economy, which results in lower income and a negative wealth effect on consumption and, in turn, lowers aggregate demand (Svensson, 2005; 2006). Therefore, economic growth is directly affected by commodity prices through changing the value of the commodity endowment and, therefore, inducing income/wealth effects (Ferraro and Peretto, 2018).

3.3.2 Cost-of-production channel

An alternative transmission mechanism by which commodity prices affect directly aggregate output is the cost-of-production channel (Ferraro and Peretto, 2018). Based on the theoretical model of Basu (1995), the production side of a small open economy comprises of competitive firms that employ labour to make and then use intermediate goods in the manufacturing of final products. Upon entry (horizontal innovation), manufacturing firms combine labour services and materials to produce intermediate goods, while they also aim to decrease the unit production costs (vertical innovation). The supply of materials is reliant on an upstream competitive sector, which uses labour services and primary commodities as inputs. As such, a change in commodity prices will spread through the entire vertical cost structure of the economy, affecting the unit production costs and correspondingly the aggregate output (Bernanke et al., 1997). This represents the cost-of-production channel, which is a key transmission mechanism that determines the link between commodity prices and economic growth.

Importantly, the impact of cost-of-production channel is not expected to be alike for all economies but it depends on whether an economy is commodity-importing or commodity-exporting. Let us assume that there is an increase in the prices of commodities. On the one hand, the firms in a commodity-importing country that purchase commodity-based products (for example, fuels, agricultural products and metals) are likely to have less net income to spend on other goods and services (Edelstein and Kilian, 2009). In particular, higher prices of commodities tend to increase the cost of raw materials and, hence, the productions costs upsurge. This causes a reduction in the firm profit and, thus, aggregate output falls (Bohi, 1991). On the other hand, the firms in a commodity-exporting country also experience an increase in the cost of raw materials, which leads to higher production costs. However, the firms gain from the commodity products they sell, which induces a positive commodity wealth effect. Therefore, the net effect depends on a series of factors, as discussed below.

Based on the theoretical framework of Ferraro and Peretto (2018), the representative firm produces materials along an infinitely elastic supply curve such that the price of materials, P_M, equals the marginal cost of materials production

$$P_M = C_M(W, p) \tag{3.5}$$

where C_M is an unit-cost function, W in the wage rate and p is the commodity price, while the production technology of materials, M, is

$$M = f(L_M, O) \tag{3.6}$$

where L_M represents the labour service, O denotes the commodity product, both of which are inputs to produce material, M, that is purchases by the manufacturing sector at price P_M. Therefore, the total production costs (TPC) is

$$TPC = C_M(W, p)M \tag{3.7}$$

Given the vertical structure of production, a commodity price change directly affects production costs and so the price, P_M, in the upstream materials sector through the unit-cost function $C_M(W, p)$. Earlier study by Garner (1989) highlights that primary commodities are important inputs into the production of manufactured goods and, therefore, changes in commodity prices directly affect production costs (TPC), and then aggregate output.

Even so, the net effect of the two aforementioned channels remains uncertain. We will investigate further whether there is a direct causal link between commodity prices and economic growth in the empirical analysis below.

3.4 Methodology

This study adopts a MF-VAR model that was proposed by Ghysels et al. (2016) to examine the relationship between monthly commodity prices and quarterly economic growth. This procedure builds upon the shortcomings of the standard VAR models that are designed for single-frequency data and often suffer from temporal aggregation bias (see Ghysels, 2016 for discussion).[73] Using the MF-VAR model, we fit Granger causality tests based on Wald statistics and test the null hypothesis of non-causality. Then, we aggregate the simulated MF data into LF and fit causality tests again. This allows for a direct comparison between the MF and the traditional LF methods.

3.4.1 Mixed-frequency VAR

The MF-VAR model is an observation-driven model that directly relates to standard VAR model settings and is suitable for exploiting Granger causality tests (Ghysels, 2016). Following the notation of Ghysels et al. (2016), we denote m to be the *ratio of sampling frequencies*, i.e. the number of high-frequency time periods in each low-frequency time period τ, where $\tau \epsilon \{1, 2, \ldots, T_L\}$ is a time sequence. The variables are at monthly and quarterly

[73] Time series are often sampled at different frequencies, and it is well known that temporal aggregation adversely affects Granger's (1969) notion of causality.

frequencies, so that $m = 3$. Therefore, the MF-VAR has the dimension $m + 1$ in the case of a bivariate MF setting. We emphasise that parameter proliferation occurs only if m is large and becomes precipitously worse as the VAR lag order increases, as discussed by Ghysels et al. (2017). However, parameter proliferation is unlikely to occur in our case due to the relatively small ratio of sampling frequencies (Ghysels et al., 2016). Thus, let $CP_{(\tau,j)}$ denote the series of commodity prices at the j-th month of quarter τ with $j\epsilon\{1,2,3\}$, while $EG_{(\tau)}$ denotes the series of economic growth at quarter τ. Section 5 provides more details on data construction. Assume that covariance stationarity is satisfied for each series. Then, the MF-VAR (p) model is specified as follows:

$$\underbrace{\begin{bmatrix} CP_{(\tau,1)} \\ CP_{(\tau,2)} \\ CP_{(\tau,3)} \\ EG_{(\tau)} \end{bmatrix}}_{\equiv X_{(\tau)}} = \sum_{k=1}^{p} \underbrace{\begin{bmatrix} a_{11,k} & a_{12,k} & a_{13,k} & a_{14,k} \\ a_{21,k} & a_{22,k} & a_{23,k} & a_{24,k} \\ a_{31,k} & a_{32,k} & a_{33,k} & a_{34,k} \\ a_{41,k} & a_{42,k} & a_{43,k} & a_{44,k} \end{bmatrix}}_{\equiv A_k} \underbrace{\begin{bmatrix} CP_{(\tau-k,1)} \\ CP_{(\tau-k,2)} \\ CP_{(\tau-k,3)} \\ EG_{(\tau-k)} \end{bmatrix}}_{\equiv X_{(\tau-k)}} + \underbrace{\begin{bmatrix} \varepsilon_{(\tau,1)} \\ \varepsilon_{(\tau,2)} \\ \varepsilon_{(\tau,3)} \\ \varepsilon_{(\tau,4)} \end{bmatrix}}_{\equiv \varepsilon_{(\tau)}} \tag{3.8}$$

where A_k is a coefficient square matrix for $k = 1, \ldots, p$, p is the lag length, and $\varepsilon_{(\tau)}$is the vector of residuals. Rather than working on aggregate quarterly data, all of the monthly observations are stacked in each quarter period τ to

obtain $X_{(\tau)} = \left[CP_{(\tau,1)}{}', CP_{(\tau,2)}{}', CP_{(\tau,3)}{}', EG_{(\tau)}{}'\right]'$. Following Ghysels et al. (2016), the constant term is not included in Equation (3.8). This notation is consistent with Kuzin et al. (2011), Ghysels (2016) and Ghysels et al. (2016). Therefore, $X_{(\tau)}$ should be thought as of a de-meaned process. The MF-VAR (p) model in Equation (3.8) can then be written as:

$$X_{(\tau)} = \sum_{k=1}^{p} A_k X_{(\tau-k)} + \varepsilon_{(\tau)} \tag{3.9}$$

$\varepsilon_{(\tau)} \sim (0, \sigma^2)$, $\sigma^2 > 0$.

To investigate the long-horizon Granger causality between commodity prices and economic growth, we iterate Equation (3.9) over the desired test horizon h and lag order p, obtaining the following MF-VAR (p, h) model:

$$X_{(\tau+h)} = \sum_{k=1}^{p} A_k^{(h)} X_{(\tau+1-k)} + u_{(\tau)}^{(h)} \tag{3.10}$$

where $A_k^{(1)} = A_k$, $A_k^{(i)} = A_{k+i-1} + \sum_{l=1}^{i-1} A_{i-l}\, A_k^{(l)}$ for $i \geq 2$, $u_{(\tau)}^{(h)} = \sum_{k=0}^{h-1} \Psi_k\, \varepsilon_{(\tau-k)}$ and, by convention, $A_k^{(1)} = 0_{k\times k}$ whenever $k > p$.

Following Ghysels et al. (2016), we make the following assumptions.[74] First, all roots of the polynomial $det(I_4 - \sum_{k=1}^{p} A_k\, z^k) = 0$ lie outside the unit circle, where $det(.)$ is the determinant. This ensures that the MF-VAR is state stationary. Second, $\varepsilon_{(\tau)}$ is a strictly stationary martingale difference sequence with finite second moment. Third, $\{X_{(\tau)}, \varepsilon_{(\tau)}\}$ obey α-mixing that satisfies $\sum_{h=0}^{\infty} \alpha_{2^h} < \infty$. This is a standard assumption to ensure the validity of the bootstrap for VAR models (for example, see Paparoditis, 1996; Kilian, 1998; Cavaliere et al., 2012; Cavaliere et al., 2014; Götz et al., 2016). In fact, these assumptions ensure the consistency and asymptotic normality of the least squares estimator $\widehat{A_k}$.[75]

Next, we exploit Wald statistics based on the coefficients of MF-VAR (p, h), $B(h) = [A_1^{(h)}, \ldots, A_p^{(h)}]'$. For example, CP do not Granger-cause EG given a MF information set equal to $a_{41,1} = \cdots = a_{42,1} = \cdots = a_{43,p} = 0_{1\times m}$, whereas EG does not Granger-cause CP given a MF information set equal to $a_{14,1} = \cdots = a_{24,1} = \cdots = a_{34,p} = 0_{m\times 1}$. Therefore, the null hypothesis of non-causality is a linear restriction defined as follows:

$$H_0 \colon Rvec[B(h)] = r \tag{3.11}$$

where R is a $q \times pK^2$selection matrix of full row rank q. $K = mK_H + K_L$, where K_H is the number of high frequency variables and $K_L$is the number of low-frequency variables. Here, $K_H = 1$ and $K_L = 1$. The complete details of the construction of R can be found in the study of Ghysels et al. (2016). r is a restricted vector, and zeros are always chosen when performing Granger causality tests. Thus, the null hypothesis of the MF Granger causality test is expressed via the following Wald statistic:

$$W_{T_L^*}[H_0(h)] \equiv T_L^*(Rvec[\hat{B}(h)] - r)' \times \left(R\hat{\sum}_p(h)R'\right)^{-1} \times \left(Rvec[\hat{B}(h)] - r\right) \tag{3.12}$$

where $T_L^* \equiv T_L - h + 1$ is the effective sample size of the MF-VAR (p, h) model, $\hat{B}(h)$ is the least square estimator of the MF-VAR (p, h) model, $\hat{\sum}_p\ (h)$ is positive semi-definite for any $T_L^* \geq 1$, and $\hat{\sum}_p\ (h) \xrightarrow{p} \sum_p(h)$ where $\sum_p(h)$ is positive definite (Ghysels et al., 2016).[76]

Next, we adopt a time-varying approach for the MF-VAR models in order to account for (1) time-varying relationships between commodity prices and economic growth and (2) structural changes in the time series. In particular, a substantial body of literature discovers structural shifts, time-varying volatility and nonlinearity over time in the commodity-growth

[74] In terms of the asymptotic theory, MF-VAR can be treated in the same way as a classical VAR. Therefore, all standard regularity conditions carry over.
[75] See Ghysels et al. (2016) for technical details.
[76] Following Ghysels et al. (2016), this study uses Newey and West's (1987) Bartlett kernel-based HAC covariance estimator, which ensures positive semi-definiteness for any $T_L^* \geq 1$, with Newey and West's (1994) automatic bandwidth selection.

relationship (for example, see Hooker, 1996; Hamilton, 1996; 2003). Therefore, we extend the full sample MF models of Ghysels et al. (2016) to a time-varying setting that allows for capturing structural changes in the commodity-growth relationship over time. This study follows Chen et al. (2010) by using a rolling rather than recursive window estimation, as it adapts more quickly to possible structural changes. The rolling procedure is relatively robust against the presence of time-varying parameters and requires no explicit assumption regarding the nature of the time variation in the data. We use a rolling window with a size of 50 quarters to estimate the model parameters. This choice of a rolling window size is determined by the following factors: (1) the availability of economic growth data and (2) the power properties of the MF-VAR, as discussed by Ghysels et al. (2016).

Last, this study follows Ghysels et al. (2016) by using parametric bootstraps by Gonçalves and Kilian (2004) in order to circumvent size distortions for small samples $\tau \epsilon \{50, 100\}$. Gonçalves and Kilian's (2004) recursive design parametric wild bootstrap does not require knowledge of the true error distribution and is robust to conditional heteroskedasticity of an unknown form. The bootstrap method is employed to improve the empirical size in small samples (see Davidson and MacKinnon, 2006 for discussion). The Wald statistic p-values are computed based on the non-robust covariance matrix and bootstraps with $N = 999$ replications. The selected number of bootstrap replications is consistent with the past studies, such as Clark and West (2006), Kilian (2009), Brüggemann et al. (2016), Götz et al. (2016), Ghysels et al. (2016) and Ghysels et al. (2017). Hence, we compute the resulting p-value of Equation (3.12), defined as follows:

$$\hat{p}_N(W_{T_L^*}[H_0(h)]) = \frac{1}{N+1} \times \left(1 + \sum_{i=1}^{N} I(W_i[H_0(h)] \geq W_{T_L^*}[H_0(h)])\right) \quad (3.13)$$

where $W_i[H_0(h)]$ is the Wald test statistic based on the ith simulation sample and the null hypothesis $\mathrm{H_0(h)}$ is rejected at level α if $\hat{\mathrm{p}}_{\mathrm{N}}\left(\mathrm{W}_{\mathrm{T_L^*}}[\mathrm{H_0(h)}]\right) \leq \alpha$.[77]

The next section presents the setting of the LF-VAR model. In particular, this study uses both methods (LF-VAR and MF-VAR) in order to exemplify that the choice of sampling frequency has an impact on the empirical outcomes.

3.4.2 Low-frequency VAR

This section formulates the LF-VAR model to examine the relationship between quarterly commodity prices and quarterly economic growth. The LF-VAR is a standard single-frequency VAR model. The notation $CP_{(\tau)}^{Q}$ denote the commodity prices at quarter τ, while the

[77] See Ghysels et al. (2016) for technical details.

notation $EG_{(\tau)}$ denotes the economic growth at quarter τ. Superscript "Q" is placed in order to explicitly distinguish between a quarterly level and a monthly level data. Since the commodity price changes are a flow variable, the aggregated commodity price variable is defined as follows:

$$CP^{Q}_{(\tau)} = \frac{1}{3}\sum_{j=1}^{3} CP_{(\tau,j)} \tag{3.14}$$

To avoid notational confusion, this study hereafter distinguishes between monthly commodity prices $\{CP_{(\tau,1)}, CP_{(\tau,2)}, CP_{(\tau,3)}\}$, quarterly commodity prices $CP^{Q}_{(\tau)}$ and a general notion of commodity prices CP. Then, the specification of the LF-VAR (p) model is given as follows:

$$\begin{bmatrix} CP^{\ Q}_{\ (\tau)} \\ EG_{(\tau)} \end{bmatrix} = \sum_{k=1}^{p} \begin{bmatrix} a_{11,k} & a_{12,k} \\ a_{21,k} & a_{22,k} \end{bmatrix} \begin{bmatrix} CP^{\ Q}_{\ (\tau-k)} \\ EG_{(\tau-k)} \end{bmatrix} + \begin{bmatrix} \varepsilon_{(\tau,1)} \\ \varepsilon_{(\tau,2)} \end{bmatrix} \tag{3.15}$$

Following Ghysels et al. (2016), the constant term is omitted and each series is de-meaned before fitting the model. In line with the empirical study of Collier and Goderis (2012), we consider first-differenced log of level series (see Section 3.5). If a time series does not satisfy the covariance stationarity after the first differencing, this series (i.e. country) is excluded from the study sample. This is because a further differencing of the level series, i.e. a second or higher differencing, does not make sense economically. All series are normalised by their full sample mean and standard deviation. The assumptions made for the MF-VAR models are valid for the LF-VAR models as well. For a detailed discussion on data handling, see Section 3.5.

Last but not least, a time-varying analysis is performed in terms of the LF-VAR models. The LF-VAR setting is identical to the one considered for the MF-VAR models in order to allow for direct comparison of the results.

3.4.3 Selection of lag length

The MF-VAR models are sensitive to the choice of lag length, such as the standard VAR models. The current MF-VAR literature discusses various methods for lag length selection but a parsimonious criterion is not suggested. On the one hand, certain past studies simply add an autoregressive lag to capture potential seasonality, such as those of Ghysels et al. (2007), Clements and Galvão (2008) and Motegi and Sadahiro (2018). On the other hand, studies such as Andreou et al. (2013) and Kuzin et al. (2011) recommend using Information Criterion (IC). In particular, Andreou et al. (2013) suggest using the Akaike Information Criterion (AIC) or Bayesian Information Criterion (BIC), while Kuzin et al. (2011) propose using BIC.

The authors of the latter study suggest that the maximum lag should be set to four when there is a mixture between quarterly and monthly data. This is the case in our study as well. Similarly, Bai et al. (2013) select the lag length by using the BIC with a maximum lag of four in a setting of quarterly and monthly data. In addition, Ghysels et al. (2016) choose lag one for a tri-variable MF-VAR model and argue that including redundant lags would have a large adverse effect on power.[78] Last but not least, Engle et al. (2013) suggest profiling the likelihood function to decide upon the lag structure.

All things considered, we follow Lütkepohl (2005) and choose the lag length by comparing the results for the selection of ICs. Following Harvey et al. (2017), we use the Bayesian, Akaike and Hannan-Quinn ICs. When the three ICs agree, that lag length is selected. When they disagree, the IC that shows the most evidence of Granger causality is displayed. The maximum lag length is set to be four, which is consistent with the works of Kuzin et al. (2011) and Bai et al. (2013). For comparison purposes, we fit the same lag orders to both the LF-VAR and MF-VAR models. In addition, we do a robustness check of the main findings of this study by using different lag-order scenarios, where the lag order $p \epsilon \{1,2,3,4\}$. This aims to provide some confidence regarding the reliability of our empirical results.

3.4.4 Out-of-sample forecasting models

This section discusses the procedure that we follow when evaluating the extent to which commodity prices can help forecast economic growth out-of-sample. Specifically, we use a rolling window, which is half the size of the total sample size, to estimate the model parameters (what we call "commodity-based forecasts").[79] This setting is consistent with the one considered by Chen et al. (2010). To reiterate, the rolling forecast procedure is often chosen in the empirical literature because it is relatively robust to the presence of time-varying parameters and requires no explicit assumption regarding the nature of the time variation in the data (for example, see Alquist and Kilian, 2010; Baumeister et al., 2015; Ball and Ghysels, 2017).

Further, we use three benchmark models to evaluate the out-of-sample forecasting performance of the commodity price models. Consistent with Chen et al. (2010), we consider the following benchmark models: an autoregressive (AR), a random walk (RW) and a random walk with drift (RWWD) (see Section 3.4.4.1). These benchmark forecasts are used to

[78] Since parameter proliferation is less of an issue in the LF-VAR, Ghysels et al. (2016) let the lag order be four in order to take into account a potential seasonality.
[79] The estimations are heteroskedasticity and serial-correlation consistent, and the results are based on the Newey and West (1987) procedure with bandwidth $T^{1/3}$, where T is the sample size (Chen et al., 2010).

evaluate the performance of our mixed data sampling (MIDAS) forecast models, which have been presented in Section 3.4.4.2.

3.4.4.1 Benchmark models

The first benchmark forecast is a basic AR time series regression model that has been used in a number of past studies (for example, see Stock and Watson, 2003; Bradshaw et al., 2012; Gospodinov and Ng, 2013; Ball and Ghysels, 2017). It is given below:

$$EG_{(\tau)} = \gamma + \sum_{i=1}^{I^*} \gamma_i \, EG_{(\tau-i)} + \varepsilon_{(\tau)} \tag{3.16}$$

where $EG_{(\tau)}$ is the economic growth in the current quarter, τ; γ and γ_i are model parameters; $\varepsilon_{(\tau)}$ is the error term; I^* is the number of lag quarters of $EG_{(\tau)}$ included in the model and selected using the BIC. This is consistent with the works of Chen et al. (2010) and Ball and Ghysels (2017).[80]

The second benchmark is based on a RW model, which is commonly used in the economic literature (for example, see Stock and Watson, 2004; Alquist and Kilian, 2010). Following Chen et al. (2010), the model is a specific case of a basic AR process, where the model parameters, as given in Equation (3.16), are assigned as $\gamma = 0$ and $\sum_{i=1}^{I^*} \gamma_i = 0$. In fact, the model is estimated without the lagged dependent variable because we consider the first differences of the variables involved in the Equation (3.16) (see Granger and Newbold, 1974 for discussion).[81] Therefore, the RW model is defined as follows:

$$EG_{(\tau)} = \varepsilon_{(\tau)} \tag{3.17}$$

The third benchmark model is based on a RWWD model (for example, see Alquist and Kilian, 2010; Chen et al., 2010). Similar to the RW model, the RWWD is estimated without the lagged dependent variable because we take the first differences of the variables present in Equation (3.16). Consequently, the RWWD model is specified as follows:

$$EG_{(\tau)} = \gamma + \varepsilon_{(\tau)} \tag{3.18}$$

Following Chen et al. (2010), the forecast comparisons between the commodity-based forecast and the benchmark are based on the following information. First, we report the number of the differences between the mean square forecast errors (MSFE) of the commodity-based forecasts and the MSFE of the benchmark, where both are re-scaled by a

[80] To elaborate, $EG_{(\tau)}$ is a weakly stationary variable due to (1) $E(EG_{(\tau)}) = EG_0$, since errors have zero expectation, and (2) $Var(EG_{(\tau)}) = \tau\sigma^2$.

[81] Another study in the literature that excludes the lagged dependent variable when constructing the random walk (with drift) model is that of Moosa and Vaz (2016).

measure of their variability. A negative number indicates that the commodity-based model outperforms the benchmark.

In addition, we use Clark and McCracken's (2001) test of equal MSFEs to compare the models. A rejection of the null hypothesis implies that the additional regressor contains out-of-sample forecasting power for the dependent variable. As discussed by Chen et al. (2010), Clark and McCracken's (2001) test of equal MSFEs is a more powerful statistical test, which corrects for finite-sample bias in the MSFE comparison between the models. The bias correction considers the possibility for the model to outperform the benchmark even when the computed MSFE differences are positive. In fact, Clark-McCracken's (2001) correction considers the case wherein two nested models with unequal size are estimated. Hence, the smaller model has an unfair advantage relative to the larger one because it imposes, rather than estimates, some of the parameters. Particularly, under the null hypothesis that the smaller model is the true specification, both models should have the same MSFE in population. However, the sample MSFE of the larger models is expected to be greater. Therefore, without correcting the test statistic, the researcher may ineptly conclude that the smaller model is better, resulting in size distortions because the larger model is rejected more frequently. The Clark and McCracken (2001) test makes a correction that addresses this finite-sample bias.

In brief, for each unique observation (country n and quarter τ), we estimate benchmark model parameters via OLS using a rolling window that is half the size of the total sample size. This is consistent with Chen et al. (2010). The AR benchmark forecast, $\widehat{EG}_{n,\tau}^{AR}$, is equal to the out-of-sample predicted value in quarter τ from the estimated model (Ball and Ghysels, 2017). The same assertion holds for the RW benchmark forecast, $\widehat{EG}_{n,\tau}^{RW}$, and the RWWD benchmark forecast, $\widehat{EG}_{n,\tau}^{RWWD}$. The three benchmark models are used to evaluate the performance of our commodity-based forecasts. The next section provides more details regarding the procedure of constructing the commodity-based forecasts.

3.4.4.2 MIDAS-forecast models

This section describes the procedure that we use to appraise the out-of-sample forecasting performance of our commodity-based models. In particular, we consider two approaches: regression-based forecast and forecast combinations.

First, we estimate regression-based forecasts based on whether the predictor variables are available at a relatively low quarterly frequency or a relatively high monthly frequency. Second, we use a forecast combination technique to aggregate the individual time series model forecasts into one composite forecast, which we have termed the MIDAS-combination forecast. This estimation setting is consistent with the earlier study of Ball and Ghysels (2017).

To begin with, we use the temporally aggregated commodity price series, $CP_{(\tau)}^{Q}$ (as described in Section 3.4.2), to construct LF regression-based forecast models. Specifically, we augment the AR model with one low-frequency predictor variable, $CP_{(\tau)}^{Q}$, as follows:

$$EG_{(\tau)} = \gamma + \sum_{i=1}^{I^*} \gamma_i \, EG_{(\tau-i)} + \sum_{f=1}^{F^*} \delta_f \, CP_{(\tau-f)}^{Q} + \varepsilon_{(\tau)} \tag{3.19}$$

where $EG_{(\tau)}$ is the economic growth in the current quarter, τ; $CP_{(\tau)}^{Q}$ is the aggregated commodity prices variable in the current quarter τ, as specified in Section 3.4.2; γ, γ_i and δ_f are model parameters; ε_τ is the error term; $F^*$and I^* are the number of lag quarters of $CP_{(\tau)}^{Q}$ and $EG_{(\tau)}$ respectively that are included in the model and are selected using the BIC. This is consistent with the works of Chen et al. (2010) and Ball and Ghysels (2017).

Analogously, we augment the RW model with one low-frequency predictor variable, $CP_{(\tau)}^{Q}$, as follows:

$$EG_{(\tau)} = \sum_{f=1}^{F^*} \delta_f \, CP_{(\tau-f)}^{Q} + \varepsilon_{(\tau)} \tag{3.20}$$

Further, the RWWD model with one low-frequency predictor variable, $CP_{(\tau)}^{Q}$, is augmented as follows:

$$EG_{(\tau)} = \gamma + \sum_{f=1}^{F^*} \delta_f \, CP_{(\tau-f)}^{Q} + \varepsilon_{(\tau)} \tag{3.21}$$

Meanwhile, we separately estimate the commodity-based forecast parameters via OLS, where we use a rolling window that is half the size of the total sample size for each unique observation (country n and quarter τ), which is consistent with Chen et al. (2010). The resulting forecast models are equal to the out-of-sample predicted value in quarter τ from the estimated model.

Particularly, the estimation of the above regression models requires the temporal aggregation of the monthly values to quarterly observations, which coincides with the frequency of the national accounts data (i.e. GDP per capita). However, the LF method has two notable downsides. First, the temporal aggregation limits the ability of the time series model to optimally use the real-time flow of information during the quarter, as highlighted by Ball and

Ghysels (2017). In other words, the information from the high-frequency observations is useful for providing updated real-time forecasts at short-horizons within the quarter period (Ghysels, 2016). Second, the use of quarterly regressors, based on aggregated high-frequency data, implicitly restricts the regression parameters, δ_f, to be temporally constant during the quarter period. Therefore, if certain months contain more relevant forecasting information than others do, that information will be lost in the process of aggregating the high-frequency data (Ball and Easton, 2013).

Next, we consider the regression-based forecast models in a MIDAS framework that are designed to exploit high-frequency information embedded in the commodity price predictor variables. Explicitly, we augment the AR model with a MIDAS specification that uses a commodity price predictor variable, $CP_{(\tau,j)}$, as shown below:

$$EG_{(\tau)} = \gamma + \sum_{i=1}^{I^*} \gamma_i EG_{(\tau-i)} + \sum_{f=0}^{F^*}\sum_{j=1}^{3} \theta_{f,j} CP_{(\tau-f,j)} + \varepsilon_{(\tau)} \tag{3.22}$$

where $EG_{(\tau)}$ is the economic growth in the current quarter, τ; $CP_{(\tau,j)}$ is the high-frequency commodity prices variable at month j of current quarter τ (as described in Section 3.4.1); γ, $\gamma_i$and $\theta_{f,j}$ are model parameters; $\varepsilon_{(\tau)}$ is the error term; $F^*$and I^* are the number of lags of $CP_{(\tau,j)}$ and $EG_{(\tau)}$ respectively that are included in the model and are selected using the BIC. This is consistent with the works of Chen et al. (2010) and Ball and Ghysels (2017).

Likewise, we augment the RW model with a single quarter, τ, high-frequency predictor variables, $CP_{(\tau,j)}$, as follows:

$$EG_{(\tau)} = \sum_{f=0}^{F^*}\sum_{j=1}^{3} \theta_{f,j} CP_{(\tau-f,j)} + \varepsilon_{(\tau)} \tag{3.23}$$

Similarly, the RWWD model is augmented with a single quarter, τ, high-frequency predictor variables, $CP_{(\tau,j)}$, as given below:

$$EG_{(\tau)} = \gamma + \sum_{f=0}^{F^*}\sum_{j=1}^{3} \theta_{f,j} CP_{(\tau-f,j)} + \varepsilon_{(\tau)} \tag{3.24}$$

The commodity-based forecast parameters are estimated via OLS, where we use a rolling window that is half the size of the total sample size for each unique observation (country n and quarter τ), which is consistent with the study of Chen et al. (2010). The resulting forecast models are equal to the out-of-sample predicted value in quarter τ from the estimated model.

Afterwards, we consider forecast combinations. This is an alternative way of exploiting information contained in the commodity prices. According to Baumeister and Kilian (2015), the forecast combinations have the following advantages. First, the forecast combinations provide assurance against forecast failures. Baumeister and Kilian (2012) highlight that forecast combinations are more robust against model misspecification than estimating an individual model-based forecast with all predictors included. Second, previous research has shown that certain forecasting models are more accurate at short-horizons, while others at longer horizons (Baumeister and Kilian, 2015). For instance, forecasting models based on the third month from a given quarter, τ, may expose superior accuracy than those based on the first month from a given quarter, τ. In other words, the aggregation limits the ability of the time series model to optimally use the real-time information flow during the quarter period (Ball and Ghysels, 2017). Third, the forecasting model with the lowest MSFE may potentially be improved by incorporating information from other models with a higher MSFE (Baumeister and Kilian, 2014). Baumeister and Kilian (2014) conclude that the equally weighted averages of quarterly forecasts are more accurate at short-horizons than the model itself. Therefore, we employ the forecast combination method because it offers an effective way to summarise a large amount of information provided by high-frequency predictors.

Moreover, Timmermann (2006) provides an excellent survey of the forecast combination methods. The author points out that estimating a separate regression for each (high-frequency) predictor, and then using the forecast combination method, is more robust against model misspecification and measurement error than estimating a single forecasting model with all predictors included. Another key point is that forecast combinations can deal with model instability and structural breaks under certain conditions and, therefore, simple strategies such as equally weighting (mean) schemes can produce more stable forecasts than the individual forecasts, as discussed by Andreou et al. (2013). As has been noted, the current literature strongly supports the fact that combination methods have better out-of-sample forecasting performance than the best performing individual model (for example, see Stock and Watson, 2003; Hendry and Clements, 2004; Aiolfi and Timmermann, 2006; Chen et al., 2010; Andreou et al., 2013; Baumeister and Kilian, 2015). Therefore, following Chen et al. (2010), we use a rolling window with size equal to half the total sample size to estimate the model parameters and generate one-quarter-ahead economic growth forecasts, using the rolling procedure:

$$EG_{(\tau)}^{j} = \gamma_j + \omega_j \sum_{f=0}^{F^*} CP_{(\tau-f,j)} + \varepsilon_{(\tau)} \tag{3.25}$$

where $EG_{(\tau)}^{j}$ is the low-frequency economic growth variable that is held fixed for all months $j\epsilon\{1,2,3\}$ in the current quarter, τ; $CP_{(\tau,j)}$ is the high-frequency commodity prices variable at month j of current quarter τ; $\gamma_j$and ω_j are model parameters; $\varepsilon_{(\tau)}$ is the error term; F^* is the number of lag quarters of $CP_{(\tau,j)}$ included in the model and selected using the BIC.

Following Baumeister and Kilian (2014), we consider the equal-weighting scheme and compare the out-of-sample forecasts of economic growth, $\frac{\widehat{EG}_{(\tau)}^{1}+\widehat{EG}_{(\tau)}^{2}+\widehat{EG}_{(\tau)}^{3}}{3}$, with the benchmark models, as described in Section 3.4.4.1. More precisely, we observe whether the MSFE differences are negative, indicating that the economic growth forecasts constructed by combining individual monthly commodity-based forecasts outperform the random walk and the autoregressive forecasts. To judge the significance of the forecast combinations, we use critical values obtained from Diebold and Mariano (1995).[82]

3.5 Data

We use annual log-differences of (1) monthly world commodity prices from January 1980 to December 2016 and (2) quarterly real GDP per capita from 1980:Q1 to 2016:Q4 for the MF-VAR modelling.[83] The main reason for differencing the commodity data is that commodity prices at levels do not affect the real GDP per capita growth; however, the rate of growth of commodity prices does enter the portfolio behaviour as one of the constituents of national income (Tobin, 1969). This choice is consistent with the previous works of Brückner and Ciccone (2010), Chen et al. (2014) and Ciccone (2018). Also, we take the year-to-year difference in order to eliminate the potential seasonality from the commodity price data, as discussed by Ghysels et al. (2016). Following Chen et al. (2014) and Gargano and Timmermann (2014), the world commodity prices are represented by the CRB Commodity Price Index, which is obtained from the Datastream (2018) database.

Next, the value of the real economic output is derived from the quarterly real GDP per capita data at a constant 2010 US$, which is downloaded from the Datastream (2018) database. Following Collier and Goderis (2012), economic growth is computed as the annual log-differences of real GDP per capita quoted at a quarterly rate. The real GDP per capita data is

[82] For further details, see Alquist and Kilian (2010), Chen et al. (2010), Chen et al. (2014) and Diebold (2015).

[83] Specifically, we use an unbalanced panel of data and, therefore, the time period for each individual country depends on the data availability. Likewise, the choice of countries to be included in our sample is mainly determined by the availability of economic growth data (see Appendix B.1).

collected for both commodity-exporting and importing countries. Specifically, commodity exporters (importers) are defined as countries with a ratio of primary commodity exports (imports) to GDP that exceeds 7%, resulting in highly commodity-dependent subsamples that contain 12 exporters and eight importers (see Appendix B.1).[84] The countries that are both commodity exporters and importers are classified as "hybrid economies" in this study. The number of those countries is 13.

Due to the inclusion of energy-related products into the index basket of CRB Commodity Price Index, we use the IMF non-fuel index as an alternative measure of the world commodity prices. In this manner, we aim to control for the effect of prices of energy-related products over the overall index movements (see Cashin et al., 2004 for discussion). Previous studies that use the IMF non-fuel index are those of Chen and Rogoff (2003), De Broeck and Sløk (2006) and Chen et al. (2010). Moreover, we use alternative proxies of the world commodity prices as a robustness check for the main results of our analysis, namely Reuters/Jeffries, Goldman Sachs, Moody's, Thompson Reuters Core Commodity Equal Weighted and IMF non-fuel commodity indexes. The selection of alternative indexes is made in line with the study of Chen et al. (2010). The data for all alternative proxies of world commodity prices is obtained from the Global Financial Database (2018).

Furthermore, we investigate the impact of national commodity export prices on economic growth by using monthly country-specific commodity export price indexes. The data of national commodity export price indexes spans from January 1980 to December 2016 and is obtained from the newly constructed database in Chapter 2. The country sample covers Bahrain, Bolivia, Chile, Ecuador, Peru, Seychelles and Venezuela. The selection of the countries is made according to (1) the commodity export-dependent status of the country and (2) the data availability. As a source of information on commodity export dependence, we use the State of Commodity Dependence 2016 report, which was published by the United Nations (UNCTAD, 2017).

Lastly, the data for all nominal commodity price indexes is transformed in real terms, as the nominal series are deflated by the MUV index from the UNCTAD (2018) database. The choice of deflator is consistent with the past studies of Grilli and Yang (1988) and Cashin et al. (2004).

[84] We follow Makhlouf et al. (2017) in the manner in which we select our sample of countries.

3.6 Economic Growth and Commodity Prices: Empirical Results

The information contained in world commodity prices can be helpful in predicting economic growth (see Stock and Watson, 2003; 2004). However, not much is known regarding the predictive content of world commodity prices for economic growth in terms of high-frequency, time-varying, in-sample and out-of-sample frameworks. This section builds upon the past literature by providing evidence for the (non-)causal patterns between world commodity prices and economic growth in a set of 33 commodity-dependent economies.

3.6.1 Preliminary statistics

We begin by considering the main features of the data, namely stationarity, which matters for the accuracy of the estimated predictive regression model. Kormendi and Meguire (1990) emphasise that a failure to account for possible unit roots by differencing may have serious statistical consequences, such as regressions estimated from data with unit roots can have non-stationary residuals, leading to spurious regression results. A proper understanding of the extent of stationarity in the data handling is therefore essential when interpreting the results. For testing stationarity, we perform Augmented Dickey-Fuller (ADF) (1979) and Phillips and Perron (PP) (1988) unit root tests for all the time series, as specified in Section 3.5. For both tests, we specify the null hypothesis of a unit root against the alternative of stationarity.

The results that we obtain from both tests denote that the null hypothesis is rejected in favour of stationarity for all economic growth and commodity price series. The unit root results as well as other preliminary statistics are discussed in more details in Appendix B.2 of this thesis.

3.6.2 Full sample Granger causality approach

We choose $h \in \{1,2,3,4,6\}$ to model the short- and longer-horizon causal relationship between world commodity prices and economic growth in order to investigate the influence of the time horizon over the empirical findings (see Dufour and Taamouti, 2010 for discussion).[85] Particularly, we use identical lag length and forecasting horizon for both the LF-VAR and MF-VAR models in order to allow direct comparison of the test results. The optimal lag length for each of the three ICs is provided in Appendix B.3. The results presented below are for the selected lag order, as discussed in Section 3.4.3.

3.6.2.1 Short-horizon investigation

The short-horizon Granger causality test results for both the MF-VAR and the LF-VAR models are reported in table 3.1. The null hypothesis of non-causality is tested against the

[85] We define *short-horizon* as of horizon of one quarter ($h = 1$), while *longer-horizon* is any horizon longer than one quarter ($h > 1$).

alternative of causality. The rejection of the null hypothesis $H_0: CP \nRightarrow EG$ implies that commodity prices Granger-cause economic growth, against the alternative hypothesis that commodity prices do not Granger-cause economic growth. Analogously, we test the null hypothesis $H_0: EG \nRightarrow CP$.

	Panel A: Mixed-frequency model		**Panel B: Low-frequency model**	
	EG $\nRightarrow$ CP	**CP $\nRightarrow$ EG**	**EG $\nRightarrow$ CP**	**CP $\nRightarrow$ EG**
Exporters				
Australia	0.532	0.585	0.410	0.509
Bolivia	0.406	0.101	0.513	0.052
Canada	0.866	0.003	0.516	0.010
Chile	0.076	0.327	0.363	0.246
Denmark	0.513	0.084	0.779	0.040
Ecuador	0.556	0.010	0.103	0.034
Kazakhstan	0.412	0.022	0.049	0.836
New Zealand	0.854	0.578	0.728	0.150
Norway	0.609	0.329	0.198	0.952
Peru	0.795	0.244	0.969	0.637
South Africa	0.608	0.010	0.560	0.004
Venezuela	0.111	0.800	0.011	0.304
Importers				
Czech Republic	0.923	0.004	0.355	0.034
Dominican Republic	0.751	0.031	0.148	0.011
Hungary	0.469	0.029	0.300	0.121
Luxembourg	0.215	0.137	0.137	0.025
Malta	0.055	0.021	0.484	0.581
Philippines	0.010	0.010	0.007	0.640
Slovakia	0.492	0.002	0.388	0.042
Slovenia	0.259	0.018	0.052	0.002
Both (Hybrid)				
Bahrain	0.440	0.063	0.170	0.340
Belgium	0.444	0.445	0.011	0.111
Hong Kong	0.854	0.002	0.126	0.122
Estonia	0.111	0.001	0.340	0.156
Iceland	0.840	0.832	0.337	0.437
Israel	0.432	0.208	0.592	0.333
Latvia	0.027	0.001	0.019	0.015
Netherlands	0.009	0.006	0.366	0.006
Serbia	0.140	0.410	0.775	0.045
Seychelles	0.072	0.044	0.147	0.939
Singapore	0.079	0.004	0.003	0.491
Thailand	0.411	0.025	0.437	0.019
Viet Nam	0.054	0.629	0.289	0.805

Note: Table 3.1 reports the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizon of one quarter, i.e. short-horizon. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0: CP \nRightarrow EG$ ($\nRightarrow$ means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ replications. All variables are mean-centred and annual log-differenced, as specified in Section 3.5.

Table 3.1 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices, Short-horizon

Table 3.1 reports the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizon of one quarter, i.e. $h = 1$. For example, the p-value for testing the null hypothesis $H_0: EG \nRightarrow CP$ ($\nRightarrow$ means "does not Granger-cause") that economic growth does not Granger-cause world commodity prices in the case of the full sample MF model for Australia is 0.532. This implies that the null hypothesis cannot be rejected at a 10% significance level and, thereby, Australian economic growth does not have predictive power on world commodity prices. The same conclusion is reached when using the full sample LF model.

In contrast, the p-value for testing the null hypothesis that Chilean economic growth does not Granger-cause world commodity prices in the case of the full sample MF model is 0.076. This infers that the null hypothesis can be rejected at a 10% level of significance. Based on this result, we can conclude that Chilean economic growth possesses an in-sample predictability for world commodity prices. This is not a surprise, as Chile is the top copper producing and exporting country and, as such, Chilean economic growth has a major role to play in the world copper market. On the contrary, the full sample LF method fails to discover a causal relationship between Chilean economic growth and world commodity prices. More precisely, the p-value for testing the null hypothesis that economic growth does not Granger-cause world commodity prices in the case of the full sample LF model for Chile is 0.363. Hence, the null hypothesis cannot be rejected, which infers that Chilean economic growth has no predictive power on world commodity prices. A potential reason for the LF outcome is the loss of information from temporal data aggregation, as discussed by Ghysels (2016).

To begin with, we find that the impact of world commodity prices on economic growth is considerably large (see table 3.1). In particular, world commodity prices are found to predict economic growth for 14 out of 33 countries by the LF method, while the MF method finds predictability for 20 out of 33 countries. This signifies the vital role of world commodity prices in forecasting economic growth for both commodity-importing and exporting economies. Moreover, the results from the MF method suggest that economic growth of eight countries Granger-causes world commodity prices, while the LF method provides evidence for only seven countries. That is to say, economic growth of only few countries has predictive power on world commodity prices. Most of these countries are hybrid economies. Therefore, we can conclude that commodity prices are a good predictor for economic growth in many countries, while the feedback hypothesis is valid only for few of them.

Furthermore, the results in table 3.1 show that the MF approach has greater power than the LF approach in terms of capturing causal patterns. This finding confirms the overall superiority of the MF method over its LF counterpart. As such, our study offers further empirical evidence to earlier works of Granger and Lin (1995), Marcellino (1999) and Ghysels et al. (2016), who advocate the appropriability of the MF method for better capturing causality in an underlying high-frequency process as compared to the traditional LF approach.

Certainly, this study primarily aims to provide evidence for the existence (or absence) of the commodity-growth relationship rather than comparing two methodological approaches, i.e. the LF-VAR and MF-VAR methods. Therefore, we investigate whether the causal relationship between economic growth and world commodity prices is more common for commodity-dependent countries that are exporter, importers or both.

Particularly, the MF approach suggests that world commodity prices Granger-cause economic growth mainly for commodity-importing countries, namely Czech Republic, Dominican Republic, Hungary, Malta, Philippines, Slovakia and Slovenia (see table 3.1). In fact, the full sample MF tests are unable to detect causality in the case of only one commodity-importing country – Luxembourg. It is interesting to say that Luxembourg is a small European economy with a well-diversified trade sector. The broad diversification of trade might be the reason why world commodity prices are not found to cause economic growth in the country. At the same time, the results from the full sample LF approach suggest that world commodity prices have an impact on economic growth for only five out of eight commodity-importing countries, namely Czech Republic, Dominican Republic, Luxembourg, Slovakia and Slovenia. All but Dominican Republic are small European economies that are dependent on commodity imports. This result implies that world commodity prices have an important role for the economic growth of small European countries. In addition, this finding builds on the study of Pradhan et al. (2015), who do not find any evidence of causality between oil prices and economic growth in the case of European countries.

Meanwhile, the impact of world commodity prices on economic growth is found to be less influential in hybrid economies. Nonetheless, the MF approach finds that world commodity prices Granger-cause economic growth for eight out of 13 hybrid economies, namely Bahrain, Hong Kong, Estonia, Latvia, Netherlands, Seychelles, Singapore and Thailand. This finding suggests that world commodity prices have predictability on economic growth not only for European countries but for Asian countries as well.

Furthermore, the LF approach discovers causality from world commodity prices to economic growth for only four out of 13 hybrid economies, namely Latvia, Netherlands, Serbia and Thailand. On contrary, the MF approach finds causal patterns in twice as many hybrid economies than its LF counterpart (see table 3.1). Put differently, the LF approach might be less powerful to capture causal patterns due to a loss of information caused by temporal aggregation of the commodity price data. This finding signifies the appropriability of using the MF-VAR approach within this study.

Surprisingly, the economic growth in commodity-exporting countries is found to be relatively unaffected by the movements of world commodity prices. Specifically, world commodity prices Granger-cause economic growth in the case of five out of 12 commodity-exporting countries, according to both the LF and MF methods. The countries found by the MF approach are Canada, Denmark, Ecuador, Kazakhstan and South Africa, while the LF method discovers causal patterns for Bolivia, Canada, Denmark, Ecuador and South Africa. Although the results differ to some extent, a straightforward conclusion can be reached from these outcomes. That is to say, world commodity prices have only a negligible ability to predict economic growth in commodity-exporting economies. This result adds to the study of Deaton and Miller (1995) to a certain degree. In particular, the authors find evidence for the existence of a long-run relationship between national commodity prices and economic growth for only four out of 32 commodity-exporting economies (namely the Central African Republic, Ghana, Liberia and Mauritania).

Briefly, while a predominant part of the literature focuses on the oil market (for example, see Stock and Watson, 2003; 2004; Cuñado and De Gracia, 2003; 2005; Narayan et al., 2014; Pradhan et al., 2015), the importance of other commodities for economic growth is often disregarded. The aforementioned findings indicate that the impact of the world commodity prices in general on economic growth is sometimes even stronger than oil prices, especially for commodity-importing countries.

Moreover, this study is consistent with earlier works of Deaton and Miller (1995), Deaton (1999), Dehn (2000), Collier and Goderis (2008; 2012) and Ciccone (2018), who consider the prices of all commodities rather than focusing on a single commodity when observing the interaction between economic growth and commodity prices. Nonetheless, while these studies focus on the long-run analysis of a commodity-growth relationship, our study compliments to them by providing evidence for the existence of a short-run relationship between economic growth and commodity prices.

This section provides a piece of evidence for a short-run relationship between world commodity prices and economic growth. However, Gargano and Timmermann (2014) emphasise that the commodity price predictability may vary with the forecast horizon. We consider the two-quarter, three-quarter, four-quarter and six-quarter horizons separately in order to address this question. The leading reason for why we expect the outcome to change is that the demand and supply of commodities can be important in the short-run, but we would expect them to be resolved in the long-run. Therefore, the next section of this study considers the longer-horizon predictability as an additional robustness check.

3.6.2.2 Longer-horizon investigation

The longer-horizon Granger causality test results for both the MF-VAR and LF-VAR models are reported in table 3.2. The null hypothesis of non-causality is tested against the alternative of causality. The rejection of the null hypothesis $H_0: CP \nRightarrow EG$ means that commodity prices Granger-cause economic growth, against the alternative hypothesis that commodity prices do not Granger-cause economic growth. Analogously, we test the null hypothesis $H_0: EG \nRightarrow CP$.

	Panel A: Mixed-frequency model								Panel B: Low-frequency model							
	horizon = 2		horizon = 3		horizon = 4		horizon = 6		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Exporters																
Australia	0.843	0.499	0.312	0.287	0.620	0.530	0.446	0.347	0.576	0.687	0.191	0.347	0.295	0.583	0.325	0.221
Bolivia	0.246	0.430	0.204	0.387	0.144	0.020	0.881	0.187	0.056	0.234	0.165	0.216	0.224	0.005	0.488	0.028
Canada	0.615	0.003	0.341	0.468	0.484	0.097	0.102	0.159	0.229	0.011	0.330	0.456	0.244	0.047	0.239	0.081
Chile	0.405	0.182	0.685	0.433	0.717	0.162	0.288	0.820	0.184	0.123	0.211	0.132	0.424	0.191	0.712	0.950
Denmark	0.819	0.034	0.394	0.272	0.323	0.054	0.578	0.282	0.292	0.082	0.137	0.041	0.251	0.072	0.838	0.302
Ecuador	0.341	0.012	0.313	0.049	0.305	0.026	0.347	0.804	0.115	0.020	0.048	0.047	0.122	0.053	0.059	0.678
Kazakhstan	0.312	0.121	0.126	0.380	0.049	0.455	0.551	0.269	0.103	0.705	0.333	0.180	0.675	0.133	0.432	0.215
New Zealand	0.825	0.056	0.492	0.012	0.665	0.140	0.885	0.041	0.373	0.051	0.126	0.175	0.312	0.344	0.700	0.063
Norway	0.145	0.057	0.224	0.124	0.226	0.069	0.238	0.121	0.133	0.866	0.093	0.652	0.136	0.635	0.143	0.977
Peru	0.827	0.151	0.314	0.594	0.533	0.890	0.062	0.676	0.662	0.734	0.199	0.947	0.344	0.775	0.183	0.250
South Africa	0.336	0.038	0.485	0.476	0.555	0.037	0.238	0.013	0.349	0.012	0.544	0.490	0.901	0.008	0.422	0.034
Venezuela	0.035	0.528	0.023	0.540	0.963	0.388	0.095	0.841	0.001	0.146	0.088	0.287	0.519	0.112	0.218	0.593
Importers																
Czech Republic	0.490	0.491	0.747	0.263	0.698	0.008	0.473	0.031	0.489	0.519	0.646	0.003	0.706	0.001	0.960	0.004
Dominican Republic	0.136	0.719	0.052	0.571	0.037	0.904	0.341	0.837	0.688	0.602	0.852	0.717	0.952	0.356	0.958	0.751
Hungary	0.594	0.198	0.355	0.345	0.342	0.277	0.313	0.261	0.223	0.281	0.166	0.122	0.285	0.088	0.381	0.110
Luxembourg	0.525	0.008	0.589	0.277	0.149	0.141	0.171	0.021	0.126	0.025	0.205	0.036	0.209	0.015	0.308	0.005
Malta	0.110	0.012	0.227	0.159	0.418	0.350	0.139	0.387	0.325	0.079	0.188	0.107	0.081	0.093	0.126	0.148
Philippines	0.003	0.004	0.008	0.010	0.036	0.128	0.006	0.310	0.017	0.005	0.011	0.008	0.010	0.145	0.012	0.239
Slovakia	0.266	0.015	0.178	0.288	0.592	0.643	0.383	0.993	0.651	0.018	0.118	0.821	0.674	0.360	0.233	0.602
Slovenia	0.061	0.160	0.132	0.109	0.044	0.047	0.348	0.361	0.313	0.017	0.094	0.059	0.041	0.027	0.042	0.062
Both (Hybrid)																
Bahrain	0.182	0.007	0.472	0.004	0.968	0.016	0.960	0.009	0.129	0.019	0.442	0.011	0.676	0.085	0.589	0.012
Belgium	0.506	0.515	0.710	0.154	0.597	0.043	0.222	0.024	0.449	0.485	0.035	0.183	0.288	0.117	0.297	0.050
Hong Kong	0.912	0.004	0.897	0.003	0.617	0.005	0.278	0.015	0.331	0.003	0.666	0.010	0.542	0.004	0.283	0.023
Estonia	0.542	0.025	0.408	0.313	0.320	0.573	0.257	0.707	0.523	0.302	0.115	0.508	0.313	0.248	0.160	0.202
Iceland	0.654	0.629	0.071	0.689	0.129	0.639	0.198	0.264	0.298	0.389	0.087	0.861	0.079	0.374	0.370	0.264
Israel	0.906	0.349	0.393	0.218	0.437	0.071	0.335	0.022	0.582	0.129	0.132	0.161	0.424	0.251	0.415	0.211
Latvia	0.853	0.038	0.438	0.029	0.462	0.032	0.889	0.205	0.064	0.001	0.265	0.008	0.126	0.081	0.096	0.021
Netherlands	0.009	0.014	0.015	0.007	0.559	0.001	0.435	0.044	0.244	0.002	0.151	0.006	0.204	0.001	0.422	0.004
Serbia	0.284	0.323	0.233	0.270	0.144	0.758	0.645	0.785	0.719	0.196	0.193	0.285	0.650	0.629	0.587	0.622
Seychelles	0.135	0.003	0.491	0.484	0.572	0.409	0.069	0.867	0.185	0.687	0.491	0.490	0.080	0.791	0.061	0.949
Singapore	0.078	0.004	0.228	0.003	0.014	0.002	0.679	0.008	0.017	0.005	0.028	0.006	0.048	0.001	0.809	0.002
Thailand	0.201	0.043	0.485	0.737	0.565	0.677	0.733	0.466	0.367	0.122	0.114	0.119	0.219	0.576	0.195	0.290
Viet Nam	0.127	0.781	0.120	0.244	0.234	0.394	0.579	0.153	0.070	0.036	0.263	0.702	0.295	0.240	0.633	0.020

Note: Table 3.2 reports the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizon larger than one quarter, i.e. longer-horizon. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0\colon CP \nRightarrow EG$ (⇏ means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ replications. All variables are mean-centred and annual log-differenced, as specified in Section 3.5.

Table 3.2 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices, Longer-horizon

Table 3.2 reports the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizons longer than one quarter, i.e. $h \geq 2$. The interpretation of the p-values is in line with the one made for the short-horizon. We consider horizons $h \in \{2,3,4,6\}$ and find that the results of the two sets of tests lead to rather similar conclusions. To emphasise, world commodity prices cause economic growth in 20 (23) economies, according to the LF (MF) approach at $h \geq 2$ (see table 3.2). This finding can be used in support of the so-called commodity-led growth hypothesis.[86] In particular, the outcomes from the full sample MF approach suggest that world commodity prices possess a predictive ability on economic growth mostly in commodity-importing and hybrid economies. A similar conclusion can be inferred from the LF approach. In contrast, when $h \geq 2$, world commodity prices are found to forecast economic growth for only seven out of 12 commodity-exporting countries, namely Bolivia, Canada, Denmark, Ecuador, New Zealand, Norway and South Africa. This result is consistent for both the MF and LF approaches. Surprisingly, the two approaches are unable to detect causality from world commodity prices to economic growth in the following commodity exporters: Australia, Chile, Kazakhstan, Peru and Venezuela. All except Australia are developing economies with less diversified export basket. More precisely, the predominant export of Kazakhstan and Venezuela is crude petroleum, while Chile and Peru are exporters of copper. Arouri et al. (2012) highlight that episodes of world geo-political tensions, the Gulf wars, the Asian crisis, the Global Financial Crisis and the current global economic weaknesses affect metal prices, which can cause sudden breaks in precious metal prices. Furthermore, Narayan et al. (2014) highlight that a potential instability in the coefficient of oil prices may alter the possibility of the full sample tests to detect a true causal relationship between commodity prices and economic growth. Therefore, we acknowledge that the relationship between commodity prices and economic growth may vary over time and the next section investigates this by adopting time-varying models.

3.6.3 Time-varying Granger causality approach

The Granger causality analysis reported so far is conducted on the full sample with the assumption that the relationship between commodity prices and economic growth remains stable over time. However, several studies have acknowledged that the commodity-growth relationship may be affected by structural breaks (for example, see Hamilton, 1983; Hamilton, 1996; Hooker, 2002; Cuñado and De Gracia, 2003; 2005; Cavalcanti et al., 2015). Given that commodity price data is known to be highly volatile (see Deaton and Laroque, 1992; Deaton,

[86] The commodity-led growth hypothesis postulates that commodity growth is one of the key determinants of economic growth.

1999), ignoring the structural changes may be crucial in detecting a potential Granger-causal relationship. Therefore, this study investigates the existence of parameter instability in the commodity-growth relationship at first before proceeding with the time-varying analysis.

3.6.3.1 Accounting for structural changes

This study addresses the issue of parameter instability by implementing Andrews' (1993) QLR test, which is a prominent test in the economic literature (see Rossi, 2005; Chen et al., 2010; Ben-Rephael et al., 2012; Bekaert et al., 2013).[87] The Andrews' (1993) QLR test is used to indicate the parameter instability when the number and location of structural breaks are unknown.[88] The null hypothesis of structural stability is specified. The testing outcomes are reported in table 3.3.

	EG $\nRightarrow$ CP	Break Dates	CP $\nRightarrow$ EG	Break Dates
Exporters				
Australia	0.410	–	0.509	–
Bolivia	0.513	–	0.052	Q2-2009
Canada	0.516	–	0.010	Q4-2005
Chile	0.363	–	0.246	–
Denmark	0.779	–	0.040	Q3-2006
Ecuador	0.103	–	0.034	Q2-2007
Kazakhstan	0.049	Q1-2013	0.836	–
New Zealand	0.728	–	0.150	–
Norway	0.198	–	0.952	–
Peru	0.969	–	0.637	–
South Africa	0.560	–	0.004	Q3-1992
Venezuela	0.011	Q2-2008	0.304	–
Importers				
Czech Republic	0.355	–	0.034	Q4-2011
Dominican Republic	0.148	–	0.011	Q2-2004
Hungary	0.300	–	0.121	–
Luxembourg	0.137	–	0.025	Q1-2003
Malta	0.484	–	0.581	–
Philippines	0.007	Q2-1987	0.640	–
Slovakia	0.388	–	0.042	Q3-2007

[87] The test against a one-time reversal is implemented with trimming values 0.15 and 0.85. Such trimming values are a conventional choice for the implementation of Andrews' (1993) test, as discussed by Stock and Watson (2003).

[88] The Andrews' (1993) QLR test is designed for same-frequency data models. Therefore, we test only the parameters in our LF models for structural instability. We are unable to perform tests for structural breaks for our MF models due to the unavailability in the current literature of an appropriate method to do so. Therefore, we leave the performance of structural breaks in a mixed-frequency environment for future research. Meanwhile, we use the results from the LF approach as evidence for the existence, or absence, of structural breaks in the commodity-growth relationship.

Slovenia	0.052	Q3-2013	0.002	Q1-2008
Both (Hybrid)				
Bahrain	0.170	–	0.340	–
Belgium	0.011	Q4-2009	0.111	–
Hong Kong	0.126	–	0.122	–
Estonia	0.340	–	0.156	–
Iceland	0.337	–	0.437	–
Israel	0.592	–	0.333	–
Latvia	0.019	Q2-2009	0.015	Q4-2006
Netherlands	0.366	–	0.006	Q1-2000
Serbia	0.775	–	0.045	Q3-2009
Seychelles	0.147	–	0.939	–
Singapore	0.003	Q3-1998	0.491	–
Thailand	0.437	–	0.019	Q3-2012
Viet Nam	0.289	–	0.805	–

Note: The table reports the p-values of Andrews' QLR (1993) test for instabilities.

Table 3.3 P-values of Andrews' QLR (1993) Test for Instabilities

Table 3.3 reports the p-values and the estimated break dates for Andrews' (1993) QLR test of parameter stability in the case of LF models. When the test rejects the null hypothesis of parameter stability, the estimated break dates are reported. For example, the p-value for testing the null hypothesis of parameter stability in the case of $CP \nRightarrow EG$ (where EG is dependent variable) model for Bolivia is 0.052. Therefore, the null hypothesis of parameter stability is rejected within a 90% confidence level. This provides evidence for the existence of a structural break in the commodity-growth relationship in the case of Bolivia, with the estimated break date being the second quarter of 2009 (i.e. Q2-2009).

In particular, the results from Andrews' (1993) QLR tests indicate the existence of evidence in favour of parameter instability in both commodity-importing and exporting countries. Specifically, the null hypothesis of structural stability is rejected for 14 out of 33 countries at a 10% level of significance. This requires the use of models that account for structural changes in the commodity-growth relationship. Consequently, this study adopts time-varying MF Granger causality tests to address the issue of parameter instability. For comparison purposes, we replicate the analysis in the LF framework by using a specification setting that is identical to that of the MF models.

3.6.3.2 Mixed-frequency versus low-frequency time-varying Granger causality

This section presents the outcomes of the time-varying mixed- and low-frequency Granger causality tests at quarterly horizons $h \in \{1,2,3,4,6\}$. Similar to Ghysels et al. (2016), the lag number is selected to be one because the inclusion of redundant lags has a substantial adverse impact on power. For consistency reasons, we employ the time-varying Granger causality tests for all countries in the sample regardless of whether evidence for structural breaks is identified by the Andrews' QLR (1993) test for instabilities.

To begin with, the short-horizon time-varying Granger causality test results of both the MF-VAR and LF-VAR models are reported in table 3.4.

	Panel A: Mixed-frequency model				Panel B: Low-frequency model			
	EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG	
	5%	10%	5%	10%	5%	10%	5%	10%
Exporters								
Australia	0	0.074	0.411	0.568	0.095	0.105	0.032	0.095
Bolivia	0.036	0.091	0.055	0.273	0.436	0.509	0.491	0.564
Canada	0	0	0.389	0.505	0	0	0.126	0.179
Chile	0.065	0.129	0.032	0.129	0	0	0.097	0.161
Denmark	0	0.039	0.588	0.647	0	0.020	0	0
Ecuador	0.067	0.800	1.000	1.000	0.800	0.933	1.000	1.000
Kazakhstan	0	0	0.081	0.568	0	0	0	0
New Zealand	0.015	0.121	0.076	0.076	0.470	0.576	0.076	0.091
Norway	0	0	0.495	0.611	0	0	0.053	0.105
Peru	0.084	0.179	0.326	0.453	0.063	0.084	0.126	0.253
South Africa	0	0	0.337	0.337	0	0.021	0.042	0.084
Venezuela	0.136	0.182	0.773	0.955	0	0	1.000	1.000
Importers								
Czech Republic	0.097	0.129	0.968	1.000	0.065	0.065	0	0
Dominican Republic	0	0	0.135	0.189	0	0	0	0.027
Hungary	0.029	0.057	0.457	0.771	0	0	0	0
Luxembourg	0	0.067	0.533	1.000	0	0	0	0
Malta	0.333	0.533	0.467	0.667	0	0	0	0
Philippines	0.066	0.154	0.484	0.560	0.209	0.363	0.033	0.110
Slovakia	0.143	0.200	0.771	0.857	0	0	0.257	0.371
Slovenia	0	0.029	0.829	0.857	0.029	0.057	0	0
Both (Hybrid)								
Bahrain	0.022	0.098	0.076	0.141	0.130	0.185	0.011	0.120
Belgium	0.029	0.057	0.829	0.857	0	0	0	0
Hong Kong	0.105	0.263	0.516	0.642	0.621	0.758	0.168	0.305
Estonia	0	0	0.829	0.829	0	0	0	0
Iceland	0	0	0	0	0	0	0	0.037
Israel	0	0	0.743	0.800	0	0.086	0	0
Latvia	0.029	0.057	0.400	0.886	0.029	0.057	0.029	0.314
Netherlands	0	0	0.903	0.968	0	0	0	0
Serbia	0	0.032	0.194	0.452	0	0	0.613	0.710
Seychelles	0.011	0.126	0.232	0.411	0.179	0.347	0.316	0.368
Singapore	0.105	0.337	0.316	0.368	0.632	0.842	0.011	0.021
Thailand	0.581	0.605	0.767	0.860	0.628	0.628	0.163	0.349
Viet Nam	0	0	0	0	0	0	0	0

Note: Table 3.4 reports the rejection frequencies at different significance levels for rolling window mixed- and low-frequency Granger causality tests with a window size of 50 quarters for the forecasting horizon of one quarter, i.e. short-horizon. For each rolling window, the Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ *replications. "CP" denotes the commodity prices, while "EG" denotes the* *economic growth variables.* H_0: $CP \nRightarrow EG$ *(*$\nRightarrow$ *means "does not Granger-cause"). All variables are mean-centred and annual log-differenced, as specified in Section 3.5.*

Table 3.4 Rejection Frequencies at Different Significant Levels for Rolling Window Granger Causality Tests, Short-horizon

Table 3.4 reports the rejection frequencies at the 5% and 10% significance levels for time-varying mixed- and low-frequency Granger causality tests at the horizon of one quarter, i.e. $h = 1$. The rejection frequency for a single country is calculated as the total number of p-values within a 5% (or 10%) significance level is divided by the total number of rolling window tests. For example, the rejection frequency for testing the null hypothesis H_0: $CP \nRightarrow EG$ ($\nRightarrow$ means "does not Granger-cause") that world commodity prices do not Granger-cause economic growth in the case of the time-varying MF model for Australia is found to be 0.568 at a 10% level of significance. This implies that the null hypothesis can be rejected for 56.8% of all cases at a 10% significance level; therefore, we can conclude that world commodity prices have predictive power on Australian economic growth. The same conclusion is reached when using the time-varying LF approach.

Notably, we find that the impact of world commodity prices on economic growth is massive. More precisely, world commodity prices are found to predict economic growth for 21 out of 33 countries using the time-varying LF method, while the time-varying MF method finds predictability for 31 out of 33 countries (see table 3.4). According to the time-varying MF approach, the causality from world commodity prices to economic growth is found for all countries apart from Iceland and Viet Nam. To emphasise, world commodity prices are found to predict economic growth in all commodity-exporting and importing economies. This finding adds to the predominant literature that focuses on the long-run analysis of the commodity-growth relationship (see Deaton and Miller, 1995; Deaton, 1999; Collier and Goderis, 2008; 2012), while far less attention has been given to the predictive ability of world commodity prices for economic growth in the short-run.

Indeed, the time-varying analysis provide much stronger support in terms of the vital role of world commodity prices in forecasting economic growth than the full sample tests (see table 3.1). Particularly, this chapter supports somehow the assertion of Lee and Chang (2005) that the various test statistics are biased towards the non-rejection of a null hypothesis when there are structural breaks. Therefore, we can conclude that accounting for structural breaks has significantly improved the power of both LF and MF approaches in detecting causal patterns (see Ghysels et al., 2017).

Moreover, the time-varying approach also discovers more cases of causality than the full sample approach when causality runs from economic growth to world commodity prices (see table 3.4). Particularly, the time-varying MF Granger causality tests demonstrate that economic growth Granger-causes world commodity prices for 70% of all cases at a 10% level of significance. In contrast, the cases of causality found by the full sample MF method are only 24% of all cases at a 10% level of significance. That is to say, the time-varying MF results provide substantial support in favour of the predictive power of economic growth for world commodity prices. In fact, we find that economic growth of commodity-importing countries has the strongest influence on world commodity prices, while less evidence is found for the other countries. This suggests that the economic growth of commodity-importing countries is a good predictor for future changes in world commodity prices.

Put differently, the time-varying LF Granger causality tests discover that economic growth Granger-causes world commodity prices for 52% of all cases at a 10% level of significance, whereas the full sample LF approach suggests the same for only 21% of all cases at a 10% level of significance. This finding implies that the time-varying method is better in capturing causality than the full sample one. In addition, this study concludes that the results from the LF tests are less supportive than their MF counterparts. Therefore, our study contributes to the existing literature by providing evidence supporting the existence of a short-run causal relationship between commodity prices and economic growth in both directions. It also provides evidence for the appropriability of the MF method for better capturing causality in an underlying high-frequency process as compared to the traditional LF approach. Additionally, the time-varying MF approach outperforms its time-varying LF counterpart regardless of the direction of causality considered (see table 3.4). It is still to be studied whether this conclusion changes if the predictability for a longer time horizon is considered.

Next, this chapter investigates the existence of a relationship between commodity prices and economic growth in a longer-horizon. The longer-horizon time-varying Granger causality test results of both the MF-VAR and LF-VAR models are reported in tables 3.5 and 3.6 respectively.

	Panel A: Mixed-frequency model															
	horizon = 2				horizon = 3				horizon = 4				horizon = 6			
	EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																
Australia	0.074	0.084	0.116	0.147	0	0	0.021	0.053	0	0	0.032	0.116	0	0.042	0.032	0.084
Bolivia	0	0.145	0	0	0	0.073	0	0	0.145	0.218	0.127	0.182	0	0	0	0.036
Canada	0	0.053	0.284	0.421	0	0.011	0.126	0.274	0	0	0.032	0.095	0	0.011	0.021	0.158
Chile	0.129	0.226	0.355	0.710	0.194	0.194	0.032	0.065	0	0	0.065	0.065	0.129	0.194	0	0
Denmark	0	0	0.157	0.275	0	0	0.176	0.373	0.039	0.078	0.255	0.333	0	0.039	0.098	0.118
Ecuador	0.533	0.800	1.000	1.000	0.600	0.867	0.800	0.933	0	0.600	0.733	1.000	0	0	0	0
Kazakhstan	0	0	0.135	0.216	0.027	0.135	0	0.054	0	0.135	0	0	0	0	0.432	0.514
New Zealand	0	0.045	0.515	0.652	0	0	0.500	0.591	0.030	0.045	0.015	0.106	0	0	0.242	0.409
Norway	0.032	0.042	0.253	0.263	0	0.063	0	0.063	0	0.074	0	0.011	0.084	0.105	0	0.021
Peru	0.011	0.074	0.263	0.284	0	0	0.253	0.274	0	0	0	0	0.105	0.147	0	0
South Africa	0	0	0.295	0.305	0	0	0.316	0.358	0.032	0.074	0.032	0.263	0	0.011	0.189	0.253
Venezuela	0	0	1.000	1.000	0	0	0.591	0.864	0	0	0.045	0.227	0	0	0	0
Importers																
Czech Republic	0.097	0.161	0.903	0.968	0.097	0.129	0.355	0.710	0.032	0.097	0.129	0.290	0	0.097	0	0.065
Dominican Republic	0	0	0	0.027	0	0.027	0	0.081	0	0.081	0	0	0	0	0	0
Hungary	0	0.029	0.200	0.486	0	0	0	0	0	0	0	0	0	0	0	0
Luxembourg	0	0.067	1.000	1.000	0	0	0.133	0.800	0	0	0	0.600	0	0	0.200	0.867
Malta	0.333	0.733	0.067	0.267	0	0	0	0.133	0	0	0	0	0	0	0	0
Philippines	0.121	0.220	0.341	0.385	0.121	0.275	0.319	0.319	0.154	0.319	0.077	0.165	0.088	0.110	0	0.033
Slovakia	0	0.086	0.886	0.886	0	0.057	0.400	0.714	0	0.086	0	0.086	0	0	0	0
Slovenia	0	0	0.686	0.743	0	0	0.029	0.029	0	0	0	0.057	0	0	0	0
Both (Hybrid)																
Bahrain	0.207	0.304	0.076	0.109	0.054	0.087	0.098	0.185	0.043	0.076	0.326	0.478	0.120	0.239	0.022	0.141
Belgium	0	0	0.714	0.829	0	0	0.029	0.114	0	0	0.171	0.200	0	0	0.343	0.343
Hong Kong	0.063	0.126	0.326	0.421	0.063	0.074	0.263	0.274	0.053	0.063	0.232	0.263	0.021	0.053	0.232	0.274
Estonia	0	0.057	0.743	0.829	0	0	0.114	0.286	0	0	0.057	0.600	0	0.400	0	0
Iceland	0	0	0	0	0	0	0	0	0	0.185	0	0	0	0.074	0	0.037
Israel	0	0	0.429	0.686	0	0	0.057	0.171	0	0	0	0	0	0	0	0.086
Latvia	0	0	0.171	0.286	0	0.029	0.200	0.429	0	0	0.114	0.143	0	0	0	0.057
Netherlands	0	0.097	0.581	0.710	0	0.097	0	0.065	0.226	0.258	0.065	0.129	0	0.032	0	0
Serbia	0.032	0.065	0.161	0.161	0.258	0.290	0.065	0.097	0.323	0.323	0.097	0.194	0	0	0	0
Seychelles	0.042	0.116	0.063	0.147	0.032	0.095	0	0.063	0.074	0.147	0	0.116	0.074	0.179	0.021	0.137
Singapore	0.158	0.295	0.274	0.411	0.042	0.063	0.042	0.126	0	0	0.105	0.253	0.032	0.116	0.011	0.095
Thailand	0.488	0.628	0.814	0.930	0.442	0.512	0.163	0.326	0.372	0.419	0.047	0.070	0.116	0.233	0	0
Viet Nam	0	0	0	0	0	0	0	0	0	0.909	0	0	0	0	0.273	0.727

Note: The table reports the rejection frequencies at different significance levels for rolling window MF Granger causality tests with a window size of 50 quarters for the forecasting horizon larger than one quarter, i.e. longer-horizon. For each rolling window, the Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ *replications. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables.* $H_0: CP \nRightarrow EG$ *(⇏ means "does not Granger-cause"). All variables are mean-centred and annual log-differenced, as specified in Section 3.5.*

Table 3.5 Rejection Frequencies at Different Significant Levels for Rolling Window MF Granger Causality Tests, Longer-horizon

	Panel B: Low-frequency model															
	horizon = 2				horizon = 3				horizon = 4				horizon = 6			
	EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																
Australia	0.074	0.084	0.011	0.021	0.011	0.074	0	0	0	0.042	0	0	0	0.042	0.074	0.084
Bolivia	0.036	0.400	0.018	0.145	0.055	0.273	0.018	0.200	0.055	0.091	0.218	0.455	0	0	0.218	0.455
Canada	0	0	0	0.011	0	0	0.032	0.063	0	0	0.074	0.084	0	0.053	0.032	0.105
Chile	0	0	0	0.065	0	0	0	0	0	0	0	0	0	0	0	0
Denmark	0	0.020	0.157	0.275	0	0	0.137	0.216	0	0	0.020	0.255	0	0	0	0.059
Ecuador	0.867	1.000	1.000	1.000	1.000	1.000	1.000	1.000	0.867	1.000	0.733	1.000	0	0	0	0
Kazakhstan	0	0	0	0	0	0	0	0	0	0	0	0	0.027	0.162	0.270	0.378
New Zealand	0.106	0.182	0.152	0.212	0.030	0.212	0.500	0.606	0.121	0.197	0.091	0.182	0	0	0.015	0.030
Norway	0	0	0.021	0.053	0	0	0.021	0.021	0	0.053	0.011	0.032	0.021	0.095	0.095	0.168
Peru	0.011	0.011	0.042	0.105	0	0	0.032	0.147	0	0	0	0.011	0.021	0.063	0	0.021
South Africa	0	0	0	0	0	0.074	0.011	0.011	0.095	0.116	0.021	0.084	0	0	0.137	0.253
Venezuela	0	0	1.000	1.000	0	0	0.909	0.955	0	0	0.682	0.864	0	0	0	0
Importers																
Czech Republic	0.097	0.097	0	0.032	0.097	0.097	0	0.032	0.097	0.097	0	0.161	0.032	0.097	0.194	0.258
Dominican Republic	0	0	0.081	0.297	0	0	0.054	0.135	0	0	0	0	0	0	0	0
Hungary	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Luxembourg	0	0	0	0	0	0	0	0	0	0	0	0.067	0	0	1.000	1.000
Malta	0	0	0	0	0	0	0	0.133	0	0.067	0.067	0.067	0	0.067	0	0.133
Philippines	0.176	0.286	0.011	0.022	0.055	0.132	0.022	0.066	0.209	0.396	0.055	0.132	0.088	0.110	0	0.011
Slovakia	0	0	0.171	0.286	0	0	0	0.029	0	0	0.114	0.171	0	0	0	0.057
Slovenia	0	0	0	0	0	0	0	0.057	0	0	0.029	0.114	0	0	0.029	0.200
Both (Hybrid)																
Bahrain	0.076	0.217	0.196	0.228	0	0.043	0.239	0.250	0	0.043	0.217	0.228	0.011	0.054	0.207	0.217
Belgium	0	0	0	0	0	0	0	0.086	0	0	0.114	0.229	0	0	0.143	0.257
Hong Kong	0.189	0.400	0.221	0.368	0.126	0.189	0.337	0.495	0.063	0.105	0.316	0.463	0.042	0.053	0.232	0.326
Estonia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Iceland	0	0	0	0	0	0	0	0	0	0.185	0	0	0	0	0	0
Israel	0	0	0	0	0	0	0	0	0	0	0	0.057	0	0	0.171	0.229
Latvia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Netherlands	0	0	0.032	0.032	0	0	0.097	0.226	0	0	0.129	0.194	0	0	0	0.129
Serbia	0	0.194	0.161	0.419	0	0	0.194	0.194	0	0	0.161	0.226	0	0	0	0
Seychelles	0.084	0.263	0.211	0.263	0.021	0.168	0.137	0.221	0.084	0.242	0.116	0.189	0.147	0.168	0.105	0.200
Singapore	0.137	0.368	0.053	0.095	0.021	0.147	0.116	0.158	0	0	0.084	0.168	0	0	0.305	0.400
Thailand	0.558	0.628	0.349	0.651	0.512	0.651	0.605	0.884	0.419	0.419	0.023	0.140	0.093	0.302	0	0
Viet Nam	0	0	0	0	0	0	0	0	0	0.182	0	0	0	0	0.727	1.000

Note: The table reports the rejection frequencies at different significance levels for rolling window LF Granger causality tests with a window size of 50 quarters for the forecasting horizon larger than one quarter, i.e. longer-horizon. For each rolling window, the Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ *replications. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables.* $H_0: CP \nRightarrow EG$ *(⇏ means "does not Granger-cause"). All variables are mean-centred and annual log-differenced, as specified in Section 3.5.*

Table 3.6 Rejection Frequencies at Different Significant Levels for Rolling Window LF Granger Causality Tests, Longer-horizon

Tables 3.5 and 3.6 reports the rejection frequencies at the 5% and 10% significance levels for the time-varying mixed- and low-frequency Granger causality tests at the horizons longer than one quarter, i.e. $h \geq 2$. Specifically, we consider time horizons $h \in \{2,3,4,6\}$. As such, we find that world commodity prices cause economic growth in 29 (33) economies, according to the LF (MF) approach at $h \geq 2$ (see tables 3.5 and 3.6). Remarkably, the time-varying MF tests detect causality is the case of all countries in our sample, while the time-varying LF approach is able to discover causality in all but four economies, namely Hungary, Estonia, Iceland and Latvia. All of these countries are small European economies with well-developed financial and economic sectors. Interestingly, two out of the three developed countries (Estonia and Latvia) are East European economies that have recently been promoted as advanced economies by the IMF. Therefore, this study concludes that world commodity prices have a crucial role to play in forecasting economic growth regardless of the country's status in terms of economic development.

In addition, the time-varying MF approach has greater power in detecting causality than the time-varying LF method among all time horizons. The only exception is when the time horizon $h = 6$ is considered. In that case, we find that world commodity prices cause economic growth in 23 (20) economies, according to the LF (MF) approach. This may suggest that the power of the MF approach becomes weaker at time horizons longer than one year, while it is superior to the LF one in detecting causality for time horizons within one year. This finding provides empirical support to the study of Gargano and Timmermann (2014) who emphasise that the commodity price predictability may vary with the forecast horizon.

Furthermore, economic growth is found to have a substantial influence on world commodity prices at $h \geq 2$. Particularly, the time-varying MF approach finds evidence for causality from economic growth to world commodity prices for 29 out of 33 economies when $h \geq 2$. The only economies for which causality from economic growth to world commodity prices is not discovered are Belgium, Israel, Slovenia and Venezuela. Notably, all countries apart from Venezuela are characterised by being small oil-importing economies with negligible influence on the world commodity market. On contrary, Venezuela is one of the largest exporters of oil. We find that its growth has immediate (short-run) effect on world commodity market, while this effect disappears in the long-run. Moreover, the time-varying LF approach indicates that economic growth possesses in-sample predictive power on world commodity prices for only 21 out of 33 countries (see table 3.6). In fact, the causality is not detected in the following economies: Belgium, Chile, Dominican Republic, Estonia, Hungary, Israel, Latvia, Luxembourg, Netherlands, Slovakia, Slovenia and Venezuela. This sample of countries includes all countries for which the time-varying MF approach does not discover causality from economic growth to world commodity prices, namely Venezuela, Slovenia, Belgium and Israel. Therefore, we can conclude that the results from the time-varying LF tests are less supportive of the existence of a causal link from economic growth to world commodity prices than their MF counterparts. This assertion is valid for all time horizons under investigation.

Overall, the above results highlight the importance of world commodity prices in predicting economic growth in commodity-dependent economies. Also, we find that economic growth in commodity-dependent economies is a robust predictor for world commodity prices. These findings add to the earlier studies such as those of Narayan et al. (2014) and Pradhan et al. (2015) that focus solely on the oil market, and particularly on the predictability of world oil prices for economic growth. Moreover, this study examines the existence of predictability in different horizons. We find a relatively high number of causal patterns at shorter time horizons, while there is a tendency of this number to decrease when the time horizon becomes longer. Hooker (1996) highlights that a potential explanation for why commodity prices no longer Granger-cause macroeconomic indicator variables is that they are endogenous. Although, the evidence in Hooker's (1996) study does not support the endogeneity hypothesis, we leave the answer of this question for a future investigation.

Furthermore, Timmermann (2006) claims that the in-sample predictive ability often fails to translate into out-of-sample success. Therefore, the subsequent section is going to provide evidence of whether the in-sample predictability of world commodity prices for economic growth translates to out-of-sample success.

3.6.4 Out-of-sample forecasts

The analysis reported so far concerns the in-sample relationship between commodity prices and economic growth. This section further examines the extent to which world commodity prices can help forecast economic growth out-of-sample. We compare the performance of the commodity-based model forecasts against the alternative benchmarks: an AR, a RW and a RWWD. All of these benchmarks are widely adopted in the literature (for example, see Stock and Watson, 2003; Chen et al., 2010; Baumeister and Kilian, 2012; Baumeister and Peersman, 2013; Gospodinov and Ng, 2013; Ball and Ghysels, 2017). Following Chen et al. (2010), we use a rolling forecast procedure to address the parameter instability issue.

3.6.4.1 *Mixed-frequency versus low-frequency out-of-sample forecast*

Here, we enhance the significance of using the MF time series over the LF one, as combining the information from all high-frequency commodity prices may improve the model's out-of-sample predictive ability. The MF case is similar to the multivariate prediction model, as explained by Chen et al. (2010).

To begin with, the differences between the MSFE of the commodity-based and the benchmark models via regression-based forecasts are reported in table 3.7.

	Panel A: Mixed-frequency model			**Panel B: Low-frequency model**		
	AR Benchmark	**RW Benchmark**	**RWWD Benchmark**	**AR Benchmark**	**RW Benchmark**	**RWWD Benchmark**
Exporters						
Australia	14.851***	-0.627	-0.954	0.466	2.282*	1.432*
Bolivia	5.994***	2.746*	8.411***	5.342***	3.155**	8.736***
Canada	11.603***	-6.665	-4.373	-0.411	-2.510	0.266
Chile	1.835	4.029**	7.426***	3.915***	9.841***	14.518***
Denmark	6.060***	1.243	-0.199	-0.493	5.324***	2.954**
Ecuador	2.357*	1.042	2.296*	1.398*	0.778	2.010**
Kazakhstan	8.567***	14.776***	16.716***	-0.232	4.901***	6.313***
New Zealand	-0.211	-0.117	0.594	1.843*	-0.574	0.066
Norway	5.689***	20.173***	2.411*	-1.266	17.93***	-1.153
Peru	-1.325	4.698**	-1.923	-0.526	4.997***	-1.169
South Africa	19.156***	32.977***	27.938***	0.592	36.913***	32.506***
Venezuela	4.123**	-0.275	1.477	3.656**	-0.351	1.584*
Importers						
Czech Republic	16.438***	1.100	1.954	-0.351	3.662**	4.991***
Dominican Republic	0.255	-2.586	-2.185	-0.735	-3.043	-2.625
Hungary	10.021***	-3.129	-1.715	-0.727	-1.452	0.484
Luxembourg	-0.602	0.313	0.407	-0.365	0.953	1.052
Malta	0.597	-1.073	-0.366	-0.195	-1.088	-0.376
Philippines	16.871***	45.847***	24.282***	-0.463	43.150***	20.858***
Slovakia	1.255	2.777*	5.052**	4.379***	13.377***	18.554***
Slovenia	26.121***	-0.241	1.211	-0.350	2.541*	4.604***
Both (Hybrid)						
Bahrain	-1.072	-5.414	-5.413	-3.120	-3.699	-3.266
Belgium	18.243***	4.096**	4.36**	-0.121	4.873***	5.244***
Hong Kong	13.934***	23.976***	32.639***	3.697***	4.714**	10.903***
Estonia	6.711***	2.014	4.038**	0.544	5.878***	8.344***
Iceland	-0.059	-1.110	-1.339	0.260	-1.225	-1.438
Israel	12.743***	13.678***	14.666***	-0.885	12.101***	12.524***
Latvia	15.076***	-2.528	-0.309	5.953***	0.160	3.179**
Netherlands	19.734***	5.548**	1.016	-0.487	12.925***	6.649***
Serbia	1.976	-1.588	0.305	2.512**	-2.030	-0.085
Seychelles	3.203*	-3.531	-1.933	0.298	-2.124	-0.715
Singapore	4.163**	25.082***	26.508***	-0.024	16.606***	17.497***
Thailand	11.004***	19.513***	20.488***	0.471	6.554***	6.686***
Viet Nam	0.978	0.538	1.085	-0.283	-0.889	-0.581

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Clark and McCracken's (2001) critical values. All variables are mean-centred and annual log-differenced, as specified in Section 3.5.

Table 3.7 Tests for Out-of-Sample Forecasting Ability – Regression Based Forecast Models

Table 3.7 presents three sets of information on the forecast comparisons. This helps us evaluate the model performance in the following ways. First, negative values indicate that the commodity-based model forecasts outperform the benchmark. Second, we use Clark and McCracken's (2001) test of equal MSFEs, for which we specify the null hypothesis that the benchmark model forecasts better than the commodity-based model against the alternative hypothesis that the commodity-based model forecasts better than the benchmark model. For example, the MSFE difference between the commodity-based model and the AR benchmark forecast in the case of Australia is found to be 14.851. On the one hand, the value is positive, implying that the benchmark outperforms the commodity-based forecast. On the other hand, the null hypothesis of Clark and McCracken's (2001) test of equal MSFEs, which states that the benchmark is better, is rejected at a 1% level of significance. This suggests that the commodity-based forecast outperforms the AR benchmark. Therefore, both methods lead to different conclusions. However, Clark and McCracken's (2001) show that their test of equal MSFEs has higher local asymptotic power than that achieved by using the differences of the MSFEs. Moreover, Clark and McCracken's (2001) test is the preferred method for assessing the model's out-of-sample performance for a wide range of past studies, such as those of Lettau and Lidvigson (2001), Clarida et al. (2003), Stock and Watson (2003; 2004), Welch and Goyal (2007), Ludvigson and Ng (2009), Molodtsova and Papell (2009), Chen et al. (2010) and Rapach et al. (2010). Hence, we follow the previous literature by using Clark and McCracken's (2001) test of equal MSFEs as the primary method for assessing the commodity-based model's out-of-sample performance.

Indeed, the results in table 3.7 suggest that world commodity prices have hardly been found to have out-of-sample predictive ability on economic growth, as compared to the benchmark models.

According to the MF approach, the commodity-based forecast outperforms the AR benchmark for 22 out of 33 countries at a 10% level of significance. At the same time, the MF approach demonstrates that both random walk benchmarks, a RW and a RWWD, are better than the commodity-based forecast for 19 out of 33 countries, respectively, at a 10% level of significance. More precisely, the commodity-based forecasts beat all the benchmark models for Belgium, Bolivia, Hong Kong, Israel, Kazakhstan, Norway, Philippines, Singapore, South Africa and Thailand. Most of these are Asian countries, while there are only two European country (Belgium and Norway), one African country (South Africa) and one Latin American country (Bolivia). This implies that world commodity prices have greater out-of-sample predictability on economic growth in Asian economies than the economies from the other continents. Notably, the commodity-based forecasts outperform the benchmark models mainly in the commodity-exporting and hybrid economies. However, a lack of evidence for out-of-sample predictability of world commodity prices on economic growth is mainly noticed in commodity-importing economies. This conclusion is valid regardless of the benchmark model.

On contrary, the LF approach finds that the commodity-based forecasts outperform the AR benchmark for nine out of 33 countries at a 10% level of significance. Likewise, the LF method discovers that both random walk benchmarks, a RW and a RWWD, are better at forecasting economic growth than the commodity-based models in 14 and 13 countries, respectively, at a 10% level of significance. Consistent with the MF approach, the LF method finds that the commodity-based forecasts outperform the benchmark models mainly in the commodity-exporting and hybrid economies. As has been noted, the evidence for out-of-sample predictability of world commodity prices on economic growth is limited in terms of the MF approach. Therefore, we can conclude that the results from the LF regression-based forecasts are more supportive for the existence of out-of-sample predictability from world commodity prices to economic growth than their MF counterparts.

Overall, the aforementioned findings challenge the widely documented pattern in the forecasting literature that the in-sample predictability often fails to translate into out-of-sample success. Unfortunately, we cannot confirm or deny this perception due to the relatively mixed results from both the LF and MF models. The next section considers forecast combinations to provide further support for the out-of-sample predictability of world commodity prices on economic growth. This is an alternative way for successfully capturing the information content in the high-frequency commodity prices, as discussed by Timmermann (2006).

3.6.4.2 Forecast combinations

This section considers the forecast combinations for examining the out-of-sample forecast performance of the commodity-based models against the alternative benchmarks. This method is consistent with the past literature, such as the studies of Stock and Watson (2003), Aiolfi and Timmermann (2006), Chen et al. (2010) and Baumeister and Kilian (2014). The advantage of using forecast combinations lies in their ability to deal with model instability and structural breaks, as highlighted by Andreou et al. (2013).

To begin with, the differences between the MSFE of the commodity-based forecast and the benchmark models via forecast combinations are reported in Table 3.8.

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	0.474	-4.871***	-4.979***	0.627	-4.831***	-4.948***
Bolivia	-0.272	-2.385**	-3.217***	-0.131	-2.321**	-3.154***
Canada	0.844	-3.623***	-3.686***	1.438	-3.579***	-3.651***
Chile	-0.626	-2.894***	-3.173***	-0.700	-2.883***	-3.169***
Denmark	0.633	-2.281**	-2.265**	1.224	-2.270**	-2.254**
Ecuador	-1.347	-2.547**	-2.586***	-1.265	-2.543**	-2.583***
Kazakhstan	-0.314	-0.346	-0.806	0.743	-0.156	-0.625
New Zealand	-0.134	-2.760***	-2.981***	0.018	-2.723***	-2.949***
Norway	1.492	-3.525***	-2.546**	1.733*	-3.462***	-2.468**
Peru	1.776*	-5.686***	-4.224***	1.823*	-5.627***	-4.196***
South Africa	-0.923	-6.486***	-7.936***	0.774	-6.382***	-7.865***
Venezuela	-0.897	-3.018***	-2.749***	-0.752	-2.950***	-2.698***
Importers						
Czech Republic	0.188	-2.556**	-3.035***	0.927	-2.280**	-2.774***
Dominican Republic	1.271	-0.981	-1.004	1.463	-0.962	-0.986
Hungary	0.355	-2.270**	-2.337**	1.801*	-2.244**	-2.315**
Luxembourg	1.075	0.706	0.414	1.183	0.702	0.418
Malta	0.623	-5.084***	-4.874***	0.993	-4.937***	-4.813***
Philippines	0.037	-6.867***	-6.014***	0.905	-6.607***	-5.672***
Slovakia	-0.582	-1.737*	-1.936*	-0.368	-1.687*	-1.886*
Slovenia	-1.355	-2.687***	-2.619***	1.105	-2.641***	-2.585***
Both (Hybrid)						
Bahrain	3.615***	-0.301	-2.312**	3.510***	-0.248	-2.241**
Belgium	-0.115	-2.103**	-1.882*	0.919	-2.062**	-1.851*
Hong Kong	0.013	-3.808***	-4.210***	0.338	-3.761***	-4.170***
Estonia	-2.535**	-1.927*	-1.888*	-2.271**	-1.917*	-1.879*
Iceland	0.595	-0.110	-0.567	0.647	-0.084	-0.547
Israel	0.269	-2.618***	-3.008***	1.920*	-2.544**	-2.928***
Latvia	-3.055***	-2.563**	-2.469**	-2.560**	-2.570**	-2.475**
Netherlands	-0.190	-1.947*	-2.179**	1.753*	-1.754*	-1.897*
Serbia	1.908*	0.535	-0.876	1.966**	0.868	-0.390
Seychelles	0.782	-6.706***	-7.934***	0.946	-6.668***	-7.890***
Singapore	0.076	-3.540***	-3.625***	0.674	-3.452***	-3.557***
Thailand	-0.278	-1.273	-1.436	0.216	-1.015	-1.158
Viet Nam	0.919	1.472	1.174	1.055	1.532	1.258

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Diebold and Mariano's (1995) critical values. All variables are mean-centred and annual log-differenced, as specified in Section 3.5.

Table 3.8 Tests for Out-of-Sample Forecasting Ability – Combination Forecast Models

The negative numbers reported in table 3.8 indicate that the commodity-based forecast outperforms the benchmark model. Similar to the previous section, we use the Clark and McCracken's (2001) test to evaluate the model's forecast performance relative to the three benchmark forecasts, namely an AR, a RW and a RWWD. The null hypothesis is that the benchmark model forecasts better than the commodity-based model against the alternative hypothesis that the commodity-based model is better. Following Chen et al. (2010), we use the Diebold and Mariano's (1995) critical values to judge the significance in terms of forecast combinations. The asterisks denote the rejection of the null hypothesis.

Considering the results from the forecast combinations, as shown in table 3.8, we find that the AR benchmark outperforms the commodity-based forecasts in 24 (28) economies, according to the LF (MF) approach. This implies that world commodity prices are poor predictors of economic growth as compared to the AR benchmark. Nevertheless, this assertion is only partially valid for hybrid economies where the commodity-based forecasts outperform economic growth for about half of the cases.

Moreover, the commodity-based forecasts outperform the RW (RWWD) benchmark for 25 (26) out of 33 countries, according to both the LF and MF methods. This finding strongly emphasizes the importance of commodity prices in forecasting economic growth. First, the commodity-based forecasts are unable to outperform the RW benchmark for only eight countries, namely Bahrain, Iceland, Kazakhstan, Luxembourg, Serbia, Singapore, Thailand and Viet Nam. Particularly, world commodity prices are found to be a decent out-of-sample predictor for the economic growth in all African and Latin American countries. Second, the commodity-based models are unable to beat the RWWD benchmark for only seven countries, namely Iceland, Kazakhstan, Luxembourg, Serbia, Singapore, Thailand and Viet Nam. This demonstrates that the results from the RWWD benchmark provide even stronger support than the RW results for the out-of-sample predictability power of world commodity prices on economic growth. Consequently, based on the results from the two random walk benchmarks we conclude that economic growth in Africa and Latin America is largely dependent on the behaviour of world commodity prices.

Other key findings from the aforementioned results are as follows. First, we find that the movements of world commodity prices have affected the economic growth of all large developed economies. This builds on the previous studies such as those of Pradhan et al. (2015) who were unable to provide any evidence of a short-run relationship existing between economic growth and oil prices in the case of the G-20 countries. Second, we also find that the commodity-based forecasts outperform the two random walk benchmarks in all commodity-exporting countries, except for Kazakhstan, and all commodity-importing countries apart from Luxembourg. While rising commodities prices are beneficial for commodity-exporters such as Australia, Canada, New Zealand and Norway, they increase the risk of a downturn in commodity-importing countries such as Czech Republic, Slovakia and Slovenia. Therefore, policymakers can use the commodity prices to help them forecast economic risks and take appropriate actions. They can use so-called fiscal policy to use of government spending and taxation to influence the level of aggregate demand and, therefore, economic activity. Such measures include requiring governments to lower taxes or increase government spending in attempt to increase economic growth during a recession for example.

Overall, the results of this analysis contribute to the existing literature as follows.

First, we demonstrate that the in-sample predictability transforms to out-of-sample success. Particularly, the results from the forecast combination models signify that the commodity-based regressions outperform the benchmark models in 79% of the total number of countries for at least two of the three benchmarks. This finding may be of potential interest to policymakers to assess the current state of the economy and its expected developments in real time, explicitly using commodity prices as a predictor variable. Notably, the data of GDP is released quarterly (and typically with a substantial temporal delay), while commodity prices are timely available at a monthly or even higher frequency. Hence, decision-makers may want to construct a forecast of the GDP growth based on the available higher frequency information at commodity prices.

Second, the forecast combination analysis reveals the most robust evidence for the out-of-sample forecasting power of world commodity prices on economic growth. This finding provides empirical evidence to the study of Timmermann (2006), who claims that the forecasting combination models perform better than the alternatives based on forecasts from a single model. Therefore, this study recommends highly the use of forecast combination models in analysing the out-of-sample predictability of world commodity prices on economic growth.

Finally, the substantial evidence that we provide for a link existing between commodity prices and economic growth indicates the long-standing need for trade diversification in countries that remain heavily dependent on a few basic commodities. Without diversification, the commodity-dependent countries are significantly more vulnerable to external commodity price shocks, which potentially affect their capacity for sustainable growth. Therefore, we advise those countries for which we find economic growth to be caused by world commodity prices to expand less volatile sectors of their economy. This has to be accompanied by fiscal reforms to restructure and broaden the revenue base in order to reduce fiscal dependency on short-term commodity revenue (UN, 2018). The massive economic costs related to recent commodity price realignments prove this point (UN, 2018).

3.7 Robustness Check

3.7.1 Lag selection

This section provides evidence regarding whether causal patterns exist between commodity prices and economic growth regardless of the choice of lag order. In fact, this aims to provide some robustness for the outcomes of our Granger causality analysis from Section 3.6. Therefore, this study reports all bootstrapped p-values for both mixed- and low-frequency Granger causality tests of lag orders $p\epsilon\{1,2,3,4\}$ in Appendix B.4. In general, we conclude that the main conclusions from our analysis in Section 3.6 remain valid in different lag-order scenarios.

3.7.2 National commodity export prices

This section provides evidence of whether commodity prices possess predictive ability for economic growth when the national commodity export prices are considered. More precisely, we explore the predictive ability of national commodity export prices (taken from Chapter 2) on economic growth in terms of several commodity-exporting economies. We replicate the above analysis by using the national commodity export prices instead of world commodity prices. We select only countries that are classified as commodity export-dependent based on the State of Commodity Dependence 2016 report published by UNCTAD (2017), and for which continuous quarterly data on real GDP per capita is available. This restricts our sample to the following countries: Bahrain, Bolivia, Chile, Ecuador, Peru, Seychelles and Venezuela.

The innovation of using national commodity prices, instead of world commodity prices, in examining the commodity-growth link is hidden behind the fact that they reflect more closely the trade structure of a given country. Particularly, synthetic measures of world commodity prices are good proxy for price fluctuations in world commodity market, while national commodity price indexes consider the trade structure of the corresponding country in their construction. This makes them to proximate national aggregates of the respective countries more closely. In other words, the ground for using national commodity prices lies on the fact that economic activity of a commodity-dependent country responds quicker to changes in national commodity prices compared to their world counterparts (Deaton and Miller, 1995). As such, national commodity export prices have been used within some of the most recognised and popular economic texts such as Deaton and Miller (1995), Dehn (2000), Cashin et al. (2004), Brückner and Ciccone (2010), Collier and Goderis (2012), Bodart et al. (2015), Caselli and Tesei (2016) and Ciccone (2018). Following that, Chapter 3 considers the importance of national commodity prices as part of the world commodity market and, as such, includes them in the estimation of the relationship between commodity prices and economic growth.

On the basis of that, the existing literature uses different ways to proxy the movements of national commodity export prices. One of the most common methods is the usage of terms of trade as a measure of national commodity prices (for example, see Cavalcanti et al., 2015; Mohaddes and Raissi, 2017). However, the terms of trade indexes are typically calculated using export and unit values; therefore, these indexes are affected by the composition of exports and by the composition of the GDP, as discussed by Deaton and Miller (1995). Hence, this thesis develops an improved index of national commodity export prices to overcome this problem (see Chapter 2 for further discussion). The index series of national commodity export prices that are constructed within Chapter 2 are used in the robustness check section of this chapter.

In particular, we find solid evidence for both in-sample and out-of-sample forecasting ability of the national commodity export prices on economic growth (for the empirical results see Appendix B.5). Using a MF full sample approach, the empirical results reveal evidence of in-sample causality from national commodity prices to economic growth in the case of all countries. In contrast, the LF full sample method is able to detect causality for only four out of seven countries (i.e. Bahrain, Chile, Ecuador and Venezuela). Furthermore, the feedback effect is found for all countries but Ecuador, according to the MF full sample method. Whereas, the LF full sample method identifies the existence of feedback effect for only two countries (i.e. Bahrain and Venezuela). These findings once again justify the adverse effect of temporally aggregated data on statistical inference (as discussed by Marcellino, 1999). More importantly, these findings prove that our constructed index series contain an important predictive power for countries' economic growth.

Moreover, the predictive content inherited in our developed commodity price indexes is confirmed by the MF and LF time-varying estimation results. Specifically, both methods provide evidence of in-sample causality from national commodity prices to economic growth in the case of all countries. Further to that, this thesis finds solid evidence in support of the out-of-sample predictive ability of national commodity prices on economic growth. The forecast combination results provide evidence that the commodity-based models outperform the benchmark models in six (five) out of seven countries for at least two of the three benchmarks, according to the MF (LF) approach. The substantial evidence found in support of a link between national commodity prices and economic growth indicates the long-standing requirement for trade diversification in countries that remain heavily dependent on commodities.

Overall, the outcomes of this robustness check are consistent with the conclusions we made in terms of world commodity prices. It also highlights that our constructed index series contain an important predictive power for countries' economic growth. Therefore, our findings can be generalised to state that commodity prices are a robust predictor for economic growth.

3.7.3 Alternative proxies for world commodity prices

Here, we use different proxies for world commodity prices in order to provide evidence regarding whether our findings are immune to the choice of index proxy. This is required because the indexes of world commodity prices differ in terms of (1) the composition of the index commodity basket and (2) the index construction – for example, the index weighting and the index formula (see Diewert, 1976). Therefore, one may argue that the findings from our analysis are largely due to the choice of index measure.

In light of this debate, our results seem incomplete. Therefore, for the sake of completeness, we replicate the above analysis using five different proxies of world commodity prices, namely Reuters/Jeffries, Goldman Sachs, Moody's, Thompson Reuters Core Commodity Equal Weighted and IMF non-fuel commodity indexes. The selection of the indexes is made in line with the study of Chen et al. (2010), and it represents an inimitable part of the robustness check. In particular, the selected indexes differ in terms of (1) index basket composition and (2) index weighting.

Particularly, we use the IMF non-fuel index to isolate the effect world energy prices have on the overall index movements (see Cashin et al., 2004 for discussion). The IMF non-fuel index provides rather similar results to those achieved by the CRB Commodity Price Index. To reiterate, we find the CRB world commodity prices to have in-sample predictive power on economic growth for all economies in our sample, according to the time-varying MF approach. Whereas, the IMF world commodity prices have in-sample forecasting power on economic growth for all countries apart from Dominican Republic when $h \geq 2$, according to the time-varying MF approach. The possible reason for the difference in the results is that the imports of Dominican Republic are nearly twice as much as its exports ($16.7 billion vs. $8.73 billion in 2017 respectively), with energy products comprising 18% of the total imports in 2017 (UN Comtrade, 2018). Therefore, the exclusion of the energy products from the index basket is a possible reason for not finding evidence of causality when using the IMF non-fuel index. This signifies the importance of our robustness check analysis, which aims to remove the presumption that our findings are due to the choice of index proxy rather than the actual predictive power the commodity prices have on economic growth of commodity-dependent countries.

Furthermore, the results for the reverse causality are supportive when considering the IMF non-fuel index. To reiterate, the time-varying MF approach finds causality from economic growth to the CRB world commodity prices for 29 out of 33 economies when $h \geq 2$. The only economies where causality is unrevealed are Belgium, Israel, Slovenia and Venezuela. However, the economic growth has in-sample predictive power for the IMF world commodity prices for all of these economies apart from Venezuela when $h \geq 2$, according to the time-varying MF approach. The inability of the IMF non-fuel index to provide evidence of causality in the case of Venezuela is consistent with the economic endowments of the country. In particular, Venezuela is a large producer and exporter of oil products; therefore, the energy commodities play an essential role in the economic growth of the country. However, the non-fuel commodity products are excluded from the IMF non-fuel index basket, which may be the reason why causality is not detected. Overall, we can conclude that the IMF non-fuel index provides even further support for the existence of a causal relationship between commodity prices and economic growth.

Similarly, we use an index with equal weights in order to check whether the index weighting is a factor that may alter our conclusions. For this purpose, we consider the Thompson Reuters Core Commodity Equal Weighted index, which equally weighs all commodity products. In fact, the replication of our analysis with an index with equal weights is crucial, as these types of indexes are usually immune to the volume effect and reflect only the prices movements, as discussed by Deaton and Miller (1995). Unsurprisingly, the test results obtained with the Thompson Reuters Core Commodity Equal Weighted index are relatively similar to those of the CRB Commodity Price Index. This implies that the forecasting power of world commodity prices on economic growth is unaffected by the index weighting.

Overall, we can conclude that the choice of a proxy does not affect the main findings of this study. The test results for all five additional proxies are provided in Appendix B.6.

3.8 Conclusions

This chapter examines whether there exists a causal relationship between commodity prices and economic growth. A sample of 33 commodity import and export-dependent countries is considered with an investigation period of January 1980–December 2016. Since the data for commodity prices is normally available at a high-frequency, while economic growth data is usually available at a lower-frequency, the use of standard same-frequency models requires temporal aggregation of the high-frequency series. However, the temporal aggregation can often generate spurious and hidden effects. Therefore, this study addresses this limitation by adopting the mixed-frequency approach of Ghysels et al. (2016).

The full sample mixed-frequency tests provide evidence in support of the forecasting ability that world commodity prices have in predicting economic growth. In particular, we find that world commodity prices have forecasting ability for economic growth in the commodity-exporting countries, while several more cases of causality are revealed in terms of the commodity-importing and hybrid economies. The findings from this study build upon the important contribution of Deaton and Miller (1995) and Deaton (1999), who examine only the long-run relationship between economic growth and commodity prices, while we provide evidence of a relationship existing in the short-run.

We also acknowledge that the relationship between commodity prices and economic growth may vary over time. This is confirmed empirically by using the Andrews' (1993) QLR structural break tests, which provide evidence in favour of the parameter instability. This study finds evidence for short-horizon in-sample causal patterns from commodity prices to economic growth for 31 out of 33 countries by using a MF time-varying approach. Meanwhile, the feedback causality is revealed for 23 out of 33 countries.

Further, we test for the existence of causality in the long-run. Notably, we find evidence of causality for certain countries when the estimation horizon is longer, while causal patterns are not detected for these countries in the short-horizon. Such an example is Viet Nam, where causal patterns are detected only when the estimation horizon is at least four quarters, i.e. at the longer-horizon. This finding suggests a possible lagged effect of commodity prices when predicting economic growth. Another explanation is that the investigation of the commodity-growth relationship in terms of a longer time horizon is redundant, especially when no control variables are included in the VAR models. Further research is required in this area.

Moreover, we find concrete evidence in support of the out-of-sample predictive ability of commodity prices for economic growth. The forecast combination results suggest that the commodity-based models outperform the benchmark models for 79% of the total number of countries according to both the LF and MF methods. This finding adds to the past study of Narayan et al. (2014), who find evidence for out-of-sample predictability of oil prices on economic growth for only 70% of all countries, while we emphasise that the commodity prices, in general, have a somewhat higher power than the oil prices in forecasting economic growth.

To confirm our main findings, we test the robustness of our results by considering different proxies of commodity prices. The robustness check indicates that the conclusions made in the main section of this analysis remain valid regardless of the choice of proxy for the commodity prices. The outcomes from the robustness check further support the in-sample predictive ability of commodity prices on economic growth, which successfully translates to out-of-sample success. This finding provides empirical evidence to the statement of Timmermann (2006) that the in-sample predictability often fails to translate into out-of-sample success.

Finally, the results from this study signify the crucial role of commodity markets in economic development of commodity-dependent economies. As such, we suggest countries that remain heavily dependent on a few basic commodities should heed to the adoption of trade diversification policies. Without diversification, the commodity-dependent countries are much more vulnerable to external commodity price shocks, which affect their capacity for sustainable growth. Future research, using our results as motivation, may explore the interaction between commodity prices and economic growth by considering other channels, such as inflation, interest rates, real effective exchange rates and so on.

Chapter 4. Global Commodity Markets and National Financial Markets: A Mixed-Frequency Time-Varying Investigation

4.1 Introduction

Given the importance of commodity markets in international trade, their impact on national stock markets has been subject to extensive academic analysis (for example, see Sadorsky, 1999; Papapetrou, 2001; Basher and Sadorsky, 2006; Kilian and Park, 2009; Basher et al., 2018; Smyth and Narayan, 2018). The extent of the reliance of national financial markets on commodity prices may hinder their development and influence the ability of national firms to find resources for investment through local financial markets (Aghion et al., 2010). As documented by Basher et al. (2018), the impact that oil market shocks have on stock prices in oil-exporting countries has implications for both domestic and international investors. In particular, there is extensive evidence suggesting that oil prices have a substantial impact on the stock markets.[89] A key question is whether, besides oil prices, the commodities in general (both energy and non-energy commodities) and metals (for example, copper, steel, lead, and others) have an impact on the national stock markets.

The vicious circle of lower commodity prices, metals in particular, caused the spending on certain types of capital goods to plunge starting mid-2015. Spending on agricultural machinery in 2016 fell by 38% since 2014, while the number for petroleum and natural gas structures, such as oil drilling rigs, was down massively – by 60% (The New York Times, 2018). With the fall in domestic capital investment in these industries, the earnings of the companies in associated industries shrank. For example, Caterpillar, a maker of heavy equipment, had 30% lower revenue in 2016 as compared to 2014 (The New York Times, 2018). The stock prices of the company felt down by 26% over this two-year period.[90] The reduction in stock liquidity (due to lower commodity prices) results in a reduction in stock prices and an increase in expected stock returns (Amihud et al., 2006). This makes it vital to look into the causal relationship between commodity and stock markets and determine the direction of causality. In sharp contrast to the extensive investigation of the oil-stock

[89] For example, see Basher and Sadorsky (2006), Kilian and Park (2009), Narayan and Narayan (2010), Filis et al. (2011), Narayan and Sharma (2011), Basher et al. (2012), Wang et al. (2013), Broadstock and Filis (2014), Adams and Glück (2015), Kang et al. (2015), Chiang and Hughen (2017), Christoffersen and Pan (2018).
[90] The stock's 52-week average closing price for Caterpillar was $100.8 in 2014 and $80.8 in 2016.

relationship, little research has been devoted to investigating the repercussions of the commodity market as a whole, and metals in particular, on financial markets.

A smaller but recent strand of papers has examined the relationship between non-fuel commodities and stock market returns. For example, authors have investigated precious metals (Baur and McDermott, 2010; Hood and Malik, 2013; Arouri et al., 2015; Basher and Sadorsky, 2016; Mensi et al., 2018), copper (Sadorsky, 2014) and foodstuffs such as sugar, coffee and cocoa (Creti et al., 2013). While these studies evaluate the role of metals and other non-fuel commodities relative to stock market returns in terms of *a hedge and safe haven* hypothesis, the evidence of a causal relationship existing between global commodity prices, especially metals, and stock market returns is limited. Moreover, much of the commodity-stock research has focused on stock markets in developed countries, mainly on the US and the UK stock markets, while less evidence exists for the other national financial markets.

This study tries to fill this gap by contributing to the existing literature through the examination of the impact of global commodity prices on the national stock markets for 63 countries and territories between January 1951 and March 2018. Several of its appealing features are distinct. First, five different measures of global commodities are considered: world oil prices, world oil demand, world oil supply, world commodity prices (all items) and world metal prices. Each of these measures is henceforth defined as a global shock variable.[91] The usage of different measures of global commodities aims to determine which one exercises the greatest influence over national stock markets. Second, by sampling stock markets from around the world, this study aims to identify the region with the highest influence of global commodity measures on its national stock markets. Third, a long historical period is considered because the commodity dependence of stock markets could have increased or decreased over time. If a country's stock market has reduced its dependence on commodities, it can be an indication that the rest of the economy is becoming stronger and that financial markets are less vulnerable to fluctuations in the commodity markets. As such, an extensive cross-country analysis is performed to examine and compare the impact of an extensive set of prices, including those of fuel, metals and all commodities.

A further contribution of the chapter is to make use of novel econometric methods that account for data sampled at different frequencies (Ghysels, 2016). This is required because high-frequency (daily or weekly) continuous data for stock markets is seldom available, especially (1) for developing countries and (2) for long historical time series – 65 years in our

[91] The definition of *global shocks* is used only for simplicity and corresponds to the five proxies considered in this study.

case. However, continuous data for world commodity prices is available in weekly frequency for a lengthy period.[92] The mixed-frequency vector autoregressive (MF-VAR) modelling approach is adopted to prevent loss of information from temporal aggregation, as discussed by Ghysels (2016). The MF-VAR allows both weekly and monthly frequency variables to be estimated together in the same framework.

There are several advantages of employing the MF-VAR method. Variables typically available at high-frequency, such as commodity prices, are often aggregated at the lower-frequency because classical models require all variables to have the same frequency. However, recent research has demonstrated that the temporal aggregation is known to have an adverse impact on the statistical inference (see McCrorie and Chambers, 2006; Andreou et al., 2010; Götz et al., 2014; Eraker et al., 2015; Schorfheide and Song, 2015; Ghysels, 2016; Ghysels et al., 2016; Motegi and Sadahiro, 2018). For example, given that the commodity price data is known to be highly volatile (see Deaton, 1999; Deaton and Laroque, 1992), working with a common low-frequency approach is likely to omit useful information about the variables (Götz et al., 2016). Moreover, the Granger causality tests in a VAR framework are not invariant to temporal aggregation (see Granger and Lin, 1995; Marcellino, 1999). Therefore, we use the MF-VAR procedure of Ghysels et al. (2016) to overcome the possible issues that arise from temporal aggregation.

Further, few studies have identified that the relationship between commodity and stock markets changes over time (see Filis et al., 2011; Chang and Yu, 2013; Broadstock and Filis, 2014; Kang et al., 2015). While a handful of studies have been carried out recently, the time-varying relationship has not been fully exploited yet, especially in the case of non-oil commodities and the stock markets of developing countries. Therefore, we apply both the MF-VAR and LF-VAR models and conduct a battery of Granger causality tests using a time-varying framework in order to analyse the dynamic nature of the commodity-stock markets relationship. This approach yields interesting results. For instance, the time-varying MF tests suggest that world oil supply shocks Granger-cause stock market returns for 78% of all cases at a 10% level of significance. However, the time-varying LF Granger causality test discovers that world oil supply shocks Granger-cause stock market returns for only 54% of all cases at a 10% level of significance. This signifies that there are around 24% fewer cases of causality found by the LF method as compared to the MF method. This result provides empirical support to the study of Ghysels (2016) by highlighting the advantages of the MF data analysis

[92] Our estimation framework cannot employ daily data (even if it is available) in combination with monthly data, as this may lead to parameter proliferation (see Ghysels, 2016 for discussion). A further research should aim to fill this gap.

and allowing us to obtain a better understanding of the causal relationships as compared to the LF models.

The rest of the chapter is organised as follows. A review of the literature on the relationship between global shocks and stock market returns is presented in Section 4.2. Next, a conceptual framework is provided in Section 4.3. Section 4.4 illustrates the methodological approach in the form of mixed and low-frequency VAR models. Section 4.5 and Section 4.6 describe the data and discuss the empirical results respectively. Section 4.7 concludes the chapter.

4.2 Literature Review

Investigations of the relationship between the commodity prices, oil in particular, and stock returns are not new. In the *Energy Economics* journal, there have been almost 70 articles published since 2008 on the relationship between oil markets and stock markets, as discussed by Smyth and Narayan (2018). Overall, they identified that the leading finance journals have published over 100 studies that were concerned with the various aspects of how oil prices influence stock returns. However, the links between non-fuel commodities and financial markets have not yet been fully understood.

Moreover, most studies focus on individual countries, the US in particular. Smyth and Narayan (2018) highlight that there are relatively few studies that examine a large number of countries – exceptions include Driesprong et al. (2008) (48 countries) and Cuñado and De Gracia (2014) (12 countries). Park and Ratti (2008), who use a sample that includes the US and 13 European countries, suggest that "It is important to consider the effects of oil prices on stock prices in a number of countries in order to better identify effects that may be systematic across countries rather than country specific" (Park and Ratti, 2008, p. 2588).

The main objective of this study is to address some of the limitations of the past literature in regards to the subjects mentioned above and examine the causal impact of commodity prices on stock market returns. Therefore, the literature review section starts with a discussion of the preceding research on the relationship between oil prices and stock market returns before proceeding to studies that focus on non-fuel commodities.

Beginning with Jones and Kaul (1996), who found that oil prices have a negative impact on the stock returns in Canada, Japan, the United Kingdom and the US, a subset of the existing literature has examined how changes in oil prices influence stock returns.[93] Papapetrou (2001) uses a multivariate VAR framework and finds that monthly oil price shocks have a negative impact on monthly Greek stock returns in the period between January 1989 and June 1999. A more recent study by Cong et al. (2008) fails to find evidence of a relationship existing between oil prices and real stock returns in China using a multivariate VAR framework and monthly data between January 1996 and December 2007. Therefore, one may conclude that the oil prices exercise a negative effect on stock markets in developed economies, while no relationship exists between oil and stock markets in developing ones.

Generally speaking, this conclusion may be incorrect if it is drawn for either a single country or a small sample of countries (e.g. sampling bias). Along these lines, one of the few studies that provide global evidence on the oil-stock relationship is conducted by Driesprong et al. (2008). Using monthly stock market data for 48 countries, Driesprong et al. (2008) find that the changes in oil prices predict stock market returns worldwide and an increase in oil prices drastically lowers the future stock returns.[94] For developed markets, they find evidence that the changes in oil prices do not predict future market returns for only three out of 18 developed markets that are considered, namely Hong Kong, Japan and Singapore. At the same time, the results for emerging markets are less pronounced. In most cases, the sign of the oil return coefficient is found to be negative (consistent with the results for developed economies), while oil prices are found to predict future market returns for only 11 out of 30 emerging markets, namely Brazil, Finland, Ireland, Jordan, New Zealand, Portugal, South Korea, Taiwan, Thailand, India and Israel. In particular, for the shorter emerging markets series starting in 1993 or after (14 countries in total), the authors find predictability for only two countries (India and Israel). They conclude, “This does not necessarily indicate that there is no significant predictability. These countries might exhibit a significant oil effect, but we simply do not have enough data to confirm this” (Driesprong et al., 2008, p. 314). Consequently, the investigation of the oil-stock relationship in terms of developing countries appears to be incomplete. While it is fair to acknowledge the strong support of the literature

[93] Although they find evidence for significant Granger causality from oil futures to stocks of individual oil companies, they detect no impact on a broad-based index such as the S&P 500. The authors use quarterly data between 1947 and 1991, and the world oil prices were measured by the US producer price index for oil.
[94] The authors consider stock returns of both (1) 18 developed markets – Australia, Austria, Belgium, Canada, Denmark, France, Germany, Hong Kong, Italy, Japan, the Netherlands, Norway, Singapore, Spain, Sweden, Switzerland, the United Kingdom and the US and (2) 30 emerging markets – Argentina, Brazil, Chile, China, Columbia, Czech Republic, Egypt, Finland, Hungary, India, Indonesia, Ireland, Israel, Jordan, Malaysia, Mexico, Morocco, New Zealand, Pakistan, Peru, the Philippines, Poland, Portugal, Russia, South Africa, South Korea, Taiwan, Thailand, Turkey and Venezuela. The authors use weekly data (only for developed markets) and monthly data with coverage from October 1973 to April 2003. They fit standard OLS regressions and use the t-values, based on White standard errors, to provide evidence for predictability.

for the negative effect of the oil prices on stock markets, the idea that increases/decreases in oil prices have the same effect on financial markets is not entirely plausible.

Several studies have found that increases/decreases in oil prices have asymmetric effects on macroeconomic variables (Hamilton, 1983) and stock prices (Sadorsky, 1999). Wan (2005) provides a theoretical justification for why oil prices may have asymmetric effects on stock returns. The author suggests that the optimal decision for listed companies is to make dividend payments only when their expected present value is above a certain threshold. In a period of oil price rises, the expected present value is likely to be below the threshold and, therefore, the firm will choose not to pay dividends and face a decline in stock prices; whereas, if the oil prices fall, the firm will pay a higher dividend and the stock price are likely to increase. In particular, the mechanism of the oil-stock link can also be applied to the commodity market in general. Analogous example of this is the share prices of the automobile manufacturers and the price of steel.

Accordingly, a strand of papers has investigated the relationship between oil prices and stock market returns in terms of asymmetry. Sadorsky (1999), who uses monthly data over the period of January1947–April 1996 and identifies that oil price shocks and its volatility play an important part in explaining the US real stock returns, conducted one of the earliest studies on the topic. In particular, the author split the time period into two samples, i.e. before and after 1986, and found that the impact of oil prices on stock returns is significantly stronger after 1986, when turbulence increased in the oil market.

A more recent study by Park and Ratti (2008) builds upon the earlier study of Sadorsky (1999) by providing evidence for the asymmetric behaviour of the oil prices on stock returns. Using linear and nonlinear multivariate VAR specifications, the authors estimate the effects of oil price shocks and oil price volatility on the real stock returns for a sample of 14 developed countries (namely Austria, Belgium, Denmark, Finland, France, Germany, Greece, Italy, the Netherlands, Norway, Spain, Sweden, the United Kingdom and the US). The estimation period is between January 1986 and December 2005, and the data frequency is monthly. They find some evidence of asymmetric effects for the US and Norway and little evidence for the oil-importing European countries.[95] Additionally, the authors find that oil price shocks have a statistically significant contemporaneous or one-month-lag impact on the real stock returns.

In an influential paper, Kilian and Park (2009) use a structural VAR (SVAR) model to estimate the impact of demand- and supply-driven oil price shocks on the US stock market

[95] The only exception is Greece.

returns over the period of January 1973–December 2006. Using monthly data, they find that the response of the US real stock returns to oil prices can be positive or negative depending on the nature of the shock. For example, the demand shocks that result from the uncertainty of future oil supply shortfalls create a negative relationship between oil prices and stock returns. Whereas, higher oil prices that result from an unanticipated global expansion have a positive effect on stock returns. The authors argue that, at the beginning of the business cycle, there will be a positive correlation between oil prices and stock returns, reflecting a strong demand for industrial commodities that drives up both oil prices and stock returns.

Apergis and Miller (2009) use monthly data to fit a SVAR approach that examines whether oil price changes affect stock-market returns in a sample of eight developed countries throughout 1981–2007.[96] Their results suggest that real oil price shocks temporally cause stock market returns in Germany, Italy, the United Kingdom and the US. In the case of Australia, only oil supply shocks temporally cause stock market returns; whereas, in the case of France, only global oil demand shocks temporarily lead stock market returns. For Canada and Japan, no such causal linkage is found.

For an extended sample of countries, Cuñado and De Gracia (2014) use a multivariate VECM to verify whether oil price changes are able to predict stock market returns for 12 oil-importing European countries using monthly data between February 1973 and December 2011.[97] The authors find that oil demand shocks have a significant negative effect on stock returns only in Germany, Italy, Luxembourg and the United Kingdom, while they have a significant positive effect on stock returns in France. However, oil supply shocks have a significant negative effect on stock returns in Belgium, Finland, France, Germany, Italy, Netherlands, Portugal, Spain and the United Kingdom. As such, they conclude that oil supply shocks exert more negative effects on European stock returns than oil demand shocks. This result is in line with the findings of Kilian and Park (2009) for the US stock returns and those of Apergis and Miller (2009) for developed economies, who also identify different effects caused by oil supply and oil demand shocks.

While all aforementioned studies focus on developed economies, no clear evidence is provided for the emerging countries. Wang et al. (2013) use monthly data in a SVAR model to observe the effect of oil price shocks on stock market returns for nine oil-importing and

[96] The data set includes Australia, Canada, France, Germany, Italy, Japan, the United Kingdom and the US.
[97] The sample of countries includes Austria, Belgium, Denmark, Finland, France, Germany, Italy, Luxembourg, Netherlands, Portugal, Spain and the United Kingdom.

seven oil-exporting countries during the period of January 1999–December 2011.[98] The empirical tests suggest that the null hypothesis of nonlinearity cannot be rejected for most of the countries in their sample. The only exception where the null hypothesis is rejected (at a 10% significance level) is that of Korea. As such, the authors conclude that there is no evidence for asymmetric effects from oil price shocks to stock market returns for all importing and exporting countries in their sample. Moreover, their empirical analysis suggests that there is no evidence for nonlinear causality from oil price changes to stock market returns for most countries. The evidence of causality is found only for two countries – Russia and Norway. A possible reason for this finding may be the short period of investigation that was considered by the authors.

Another study that explores the oil-stock relationship in terms of developing countries is Fang and You (2014). The authors follow the procedure of Kilian and Park (2009) and investigate the manner in which oil price shocks affect monthly stock market returns in China, India and Russia between January 2001 and May 2012. They find that the oil prices in India always have a negative impact on the country's economy. In Russia, there is a significant positive impact on stock returns only when the Russian oil-specific supply shocks drive oil price changes. In China, the oil-specific demand-driven shocks have a significant negative effect on stock returns during the third to sixth month, whereas oil price shocks driven by global oil demand have no significant effect. Overall, these findings provide some evidence for the asymmetric effects of oil price shocks on stock market returns in developing countries, which have not been captured by the study of Wang et al. (2013).

Together with asymmetry, another branch of papers has examined the time-varying relationship between oil prices and stock market returns (among others, see Ciner, 2001; Miller and Ratti, 2009; Kang et al., 2015). In particular, those studies have mainly focused on developed economies. Ciner (2001) uses a nonlinear Granger causality approach to examine the dynamic linkages between daily future oil prices and the US stock returns. The author uses two data samples for the following periods: (1) from 9th October 1979 to 16th March 1990 and (2) from 20th March 1990 to 2nd March 2000. He finds significant nonlinear Granger causality from crude oil future returns to S&P 500 index returns in both samples. There is also evidence that stock index returns affect crude oil futures, suggesting a feedback effect. Kang et al. (2015) combined the Kilian and Park (2009) SVAR model with a time-varying parameter VAR to examine the impact of structural oil price shocks on the US stock

[98] The data set includes nine oil-importing countries (China, France, Germany, India, Italy, Japan, South Korea, the United Kingdom and the US) and seven oil-exporting countries (Canada, Kuwait, Mexico, Norway, Russia, Saudi Arabia and Venezuela).

market returns between January 1968 and December 2012 at a monthly frequency. They find that oil price shocks contain useful information for forecasting US real stock returns, while the coefficients and the nature of shocks have varied over time. This finding adds to the study of Ciner (2001) by suggesting that the relationship between oil prices and the US stock returns changes with time (i.e. there are periods of non-causality).

While most studies focus on the US market, Miller and Ratti (2009) consider a larger sample of developed countries. The authors use a vector error correction model (VECM) to investigate the long-run relationship between monthly world oil prices and monthly stock returns for six OECD markets (Canada, France, Germany, Italy, the United Kingdom and the US) between January 1971 and March 2008.[99] Considering the full sample period, they find no evidence for either a short- or long-run relationship between oil prices and stock returns for either of the countries. Then, the authors use the testing procedure of Hansen and Johansen (1993) to identify the possible structural breaks in the oil-stock relationship and, subsequently, split the full sample period into sub-periods relative to the identified breakpoints. As a result, the study shows that the long-run relationships exist between oil and stock markets, particularly during the periods of January 1971–May 1980 (for Germany, Italy, the United Kingdom and the US) and February 1988–September 1999 (for Canada, France, Germany, Italy, the United Kingdom and the US). Further, no evidence for either a short- or long-run relationship is found for the period of June 1980–January 1988. Moreover, evidence for a short- and long-run relationship appears only in the case of Canada for the period after September 1999. This finding is in line with the conclusions of some past studies, such as Kang et al. (2015), that the oil-stock relationship changes over time. While such evidence is provided for the oil market, there is limited evidence for other commodity markets.

As can be seen, the literature testing the commodity-stock relationship has mainly focused on the oil market; however, there are only few studies that examine the effect of non-fuel commodities on the stock market returns (among others, see Baur and McDermott, 2010; Creti et al., 2013; Hood and Malik, 2013; Sadorsky, 2014; Arouri et al., 2015; Basher and Sadorsky, 2016; Mensi et al., 2018). Actually, most if not all studies in the literature examine the time-varying correlation behaviour of non-fuel commodities with respect to stock markets in developed countries and that of the US stock market in particular (for example, Creti et al., 2013; Hood and Malik, 2013). Creti et al. (2013) investigate the dynamic correlation between daily price returns for 25 commodities and the US stock returns for the period of January

[99] The study divides the full sample period into four sub-periods with break dates: May 1980, January 1988 and September 1999. This aims to account for potential asymmetries in the oil-stock relationship.

2001–November 2011. They find that the correlations between commodity and stock markets evolve with time and are highly volatile, particularly since the 2007–2008 financial crisis. In addition, some commodities, such as oil, coffee and cocoa, are found to have a stronger correlation with the S&P 500 returns when stock prices are increasing, while the correlation becomes weaker in times of bearish financial markets. For precious metals such as gold, they find that its correlation with stock returns is mostly negative and diminishes in times of declining stock prices. Similarly, Hood and Malik (2013) look at the correlation between precious metals (such as gold, silver and platinum) and the US stock market returns. Using daily data from November 1995 to November 2010, they find that gold, unlike other precious metals, serves as a hedge and a weak safe haven for the US stock market. This result confirms the finding of Creti et al. (2013) regarding the role of gold as a hedge against investment risks.

Expanding the sample to other developed countries, Baur and McDermott (2010) discover that the relationship between commodity and stock markets vary across markets. Baur and McDermott (2010) examine the daily conditional volatility of the individual country stock indexes and gold returns between 2nd March 1979 and 2nd March 2009.[100] They analyse the time-varying behaviour of gold with respect to global stock markets by using a rolling window regression with window length set to 250 daily observations, which approximately represents one calendar year. The authors find that gold is both a hedge and a safe haven for major European stock markets (namely France, Germany, Italy, Switzerland and the United Kingdom) and the US but not for Australia, Canada, Japan and large emerging markets such as the BRIC countries.[101] Therefore, one may say that the link between gold prices and emerging stock markets is stronger than that found in developed markets.

In particular, less evidence has been provided in the case of developing stock markets. A possible reason is stock data scarcity (or discontinuity) for developing markets. To overcome this problem, certain studies, such as those of Sadorsky (2014) and Basher and Sadorsky (2016), use aggregate stock market indexes. Sadorsky (2014) examines the correlations between the MSCI Emerging Markets Index and the prices of copper, oil and wheat between 3rd January 2000 and 29th June 2012. The author shows that the dynamic conditional correlations between the stock market index and the commodity prices increased between 2008 and 2009 and that oil provides the cheapest hedge for emerging markets' stock prices.

[100] The sub-set comprises the seven largest developed countries (G7), the largest emerging markets (BRIC countries), Australia and Switzerland. They include Australia to represent a small developed country with a large commodity market and Switzerland as a small European and non-Euro market with a strong and potentially important currency.

[101] BRIC is an acronym that refers to four important emerging stock markets: Brazil, Russia, India and China. The acronym BRICS includes the aforementioned countries and South Africa.

This result is consistent with the findings of Creti et al. (2013) for the US market, who discover that volatility between different asset classes increased or changed after the most recent financial crisis. Another study by Basher and Sadorsky (2016) uses a fixed rolling window multivariate GARCH to model conditional correlations between MSCI Emerging Markets Index, oil prices and gold prices between 4th January 2000 and 31st July 2014. Consistent with the study of Sadorsky (2014), the authors conclude that oil is the best hedge for emerging market stock prices because the stock/oil hedge has the highest hedging effectiveness in most cases. At the same time, gold has been found to be less effective for hedging emerging market stock prices. This finding is in line with the conclusions of Baur and McDermott (2010).

While these studies use aggregate indexes of emerging market stock prices, the characteristics of the individual stock markets may be wiped away. A study by Arouri et al. (2015) explores the return and volatility spillovers between daily world gold futures and daily stock market prices in China between 22nd March 2004 and 31st March 2011. They discover that the gold asset serves as a safe haven for stocks in the Chinese stock market. A similar result was found by Mensi et al. (2018) who use daily data to examine the co-movements between commodity prices (gold and oil) and BRICS stock market returns between 29th September 1997 and 4th March 2016. Their results indicate that BRICS stock returns co-move with the WTI crude oil price at long horizons. Moreover, the authors find stronger co-movement during the onset of the Global Financial Crisis. No evidence of co-movement is detected between the BRICS stock market returns and gold prices over time and across frequencies (horizons). The latter implies that gold can act as a hedge or a safe haven asset for BRICS economies against extreme market movements. This is in sharp contrast with the findings of Sadorsky (2014) and Basher and Sadorsky (2016), who claim that gold has a strong link with emerging market stock prices.

Therefore, there is a lack of a definite conclusion regarding the link between non-fuel commodity prices (metals in particular) and stock returns for most stock markets. Additionally, the past studies that investigate the commodity-stock relationship in terms of non-fuel commodities have mainly looked at the process of correlation, while the direction of causality remains unknown to a large extent.

While the aforementioned studies clearly make a significant contribution to the topic, the literature still possesses some substantial gaps that we try to fill in this chapter.

First, the past literature has mainly focused on the oil market, while little or no importance has been given to the non-fuel commodities. In this study, we address this gap by looking not only at fuel but also at non-fuel commodities, such as metals, and considering a general price index of all commodities.

Second, it is well-known that commodity prices alternate between periods of relative tranquillity and periods of turbulence. This is particularly true for long time periods where the link between commodity prices and stock markets is more likely to change. In this study, we consider a data set that, compared to past studies, includes a larger combination of countries and time periods. This allows us to investigate the time-varying nature between world commodity prices and financial markets over a 65-year period.

Third, we address an important data issue that has generally been overlooked. Since daily or weekly stock price data is often unavailable for long historical time periods, especially for developing countries, we use monthly stock market returns, which is in line with most studies in the existing literature. However, since commodity price data is available at a weekly frequency for a longer time period, the temporal aggregation of this data may lead to a loss of information (see Ghysels, 2016). We overcome this issue by using the mixed-frequency approach developed by Ghysels et al. (2016). Combining time series at different sampling frequencies, this approach allows us to improve upon the past studies by examining the causal patterns between commodity prices and national stock market returns without the loss of data characterisation and properties.

4.3 Conceptual Framework

In describing the theoretical link between commodity price changes and stock market returns, we consider the equity pricing channel, which is the direct channel by which commodity prices influence stock markets (Degiannakis et al., 2018). In an equity pricing model, the price of equity, at any point in time, is equal to the expected present value of discounted future cash flow (Huang et al., 1996).

Economic theory suggests that stock prices are determined by company's expected discounted cash flows (Williams, 1938). Consequently, any factor that could alter the expected discounted cash flows has a significant effect on stock prices. Commodities, along with capital, labour and materials represent important components into the production of most goods and services and changes in the prices of these inputs affect future cash flows (Basher and Sadorsky, 2006; Keun Yoo, 2006). In other words, any commodity price increase is likely to be accompanied by a decrease in stock prices.

Following the theoretical justification proffered by Jones and Kaul (1996), asset values are determined by expected discounted cash flow, which embodies the cash flow hypothesis (Williams, 1938; Sadorsky, 1999). Then, higher (lower) commodity prices tend to decrease (increase) the future cash flow and, therefore, stock prices decrease (increase). More explicitly, the following two equations summarise the theoretical fundamentals of the cash flow hypothesis.

First, we define the stock returns as the first log-difference:

$$SP_{i,t}^{R} = \ln(SP_{i,t}) - \ln(SP_{i,t-1}) \tag{4.1}$$

where $SP_{\mathrm{i},t}$ denotes the stock price of firm i at time t.

Second, economic theory suggests that the current stock prices reflect the discounted future cash flow of a particular stock (Huang et al., 1996). This can be expressed as follows:

$$SP_{i,t} = \sum_{n=t+1}^{N} \left(\frac{E(CF_n)}{\left(1 + E(r)\right)^n} \right) \tag{4.2}$$

where CF_n is the cash flow at time n, r is the discount rate and $E(\cdot)$ denotes the expectation operator.

The above two equations, Equations (4.1) and (4.2), illustrate that the stock returns are impacted by factors that can alter the expected cash flow and/or the discount rate, including commodity prices. Therefore, any changes in the commodity prices can alter a firm's future cash flow either positively or negatively, based on whether the firm is a commodity-consumer or a commodity-producer (Oberndorfer, 2009; Mohanty et al., 2011).

In particular, commodities such as fuels and metals, along with capital, labour and materials, represent essential components in the production of most goods and services; therefore, any changes in the prices of these inputs affect the cash flow. For instance, in the absence of the complete substitution effects between the factors of production, an increase in the international price of the primary commodity will increase the production costs (Basher and Sadorsky, 2006). Consequently, higher production costs dampen cash flow and reduce stock prices. The overall impact of rising commodity prices on stock prices depends on whether a company is a consumer or a producer of commodities and commodity-related products.

4.3.1 Commodity prices and stock price fluctuations in commodity-importing countries

For commodity-importing countries, higher commodity prices affect stock markets through the production cost function by reducing the net amount of the commodity used in production (Kim and Loungani, 1992; Backus and Crucini, 2000; Hooker, 2002; Bjørnland, 2009;

Mohanty et al., 2011). Consequently, an increase in commodity prices leads to a rise in production costs, inducing firms to lower cash flows. This reduction in cash flows and income induces rational firms in commodity-importing countries to hold back on investment spending and, therefore, value growth falls. In a nutshell, higher production costs dampen cash flows and reduce stock prices (Basher and Sadorsky, 2006).

Based on the theoretical concept of Degiannakis et al. (2018), higher commodity prices lead to negative income effect and shift the short-run aggregate supply curve to the left. This shift is determined by the upsurge in the production costs, mainly due to increases in the cost of wages and raw materials. Therefore, the leftward shift of the aggregates supply curve leads to cost-push inflation (Hooker, 2002). Specifically, Basher and Sadorsky (2006) highlight that rising commodity prices are often indicative of inflationary pressures which central banks can control by raising interest rates. Higher interest rates make bonds look more attractive than stocks leading to a fall in stock prices.

At the same time, higher commodity prices lead to lower income by pushing the aggregate demand curve to the left (Hamilton, 1996; 2009). This is due to the higher expenditure on home consumption goods whose prices are determined by the price of primary commodity products; for example, fuels and agricultural products (e.g. wheat, barley, maize). In addition, the aggregated demand curve shifts to the left due to production cost effects as some portions of these are passed onto consumers via non-tradables, such as retail and wholesale trade (Benguria et al., 2018). Hence, there is an increase in the retail prices, which lowers the home consumption. Those factors influence the risk in cash flow and required rate of return, which then impact cost of capital and, therefore, stock prices decrease.

In case when the monetary authority tries to counteract potential increases in inflation by reducing the money supply, the short-run interest rates will be higher. This forces firms to reduce their investment activity (Bernanke et al., 1997). This is because higher interest rates increase the cost of borrowing, which discourages businesses to increase investment spending. Based on the theoretical concept of Smyth and Narayan (2018), higher commodity prices can lead to an overestimation of expected inflation and higher nominal interest rates. Because interest rates are used to discount expected future cash flows, this will depress earnings, dividends and, hence, stock returns. Therefore, this requires investment to have a higher rate of return to be profitable (Jones and Kaul, 1996).

Overall, tightening the money supply discourages business expansion and consumer spending and negatively impacts stock markets (Smyth and Narayan, 2018).

4.3.2 Commodity prices and stock price fluctuations in commodity-exporting countries

For commodity-exporting countries, higher commodity prices may affect stock markets in two ways: (1) through profit margins channel and (2) through production costs channel (Bjørnland, 2009).

On the subject of the first channel, higher commodity prices represent an immediate transfer of wealth from commodity importers to commodity exporters and, hence, there is an increase in the profit margins in commodity-producing firms. The potential for profitable output from the commodity-producing sector can also provide huge investment and business opportunities in the overall economy, with increased demand for labour and capital (Bjørnland, 2009). This could reflect stronger business performance and the concomitant impact on stock markets, as discussed by Kollias et al. (2013). However, the high rate of economic activity may put upward pressures on inflation and on the domestic currency, which often appreciates in commodity-exporting countries and, hence, have an adverse impact on stock markets (Haldane, 1997).

Regarding the second channel, higher commodity prices lead to a commodity induced recession in the commodity-importing countries and, therefore, they will demand less export of traditional goods and services from the commodity-exporting countries. Consequently, this channel provides a negative stimulus to the commodity-exporting countries, especially if the commodity-exporting country has a large sector of commodity-consuming firms (Bjørnland, 2009; Mohanty et al., 2011). On the top of that, higher commodity prices lead to an increase in the production costs of commodity-consumer firms. This has a negative impact on the cash flow and, hence, stock prices decrease (Degiannakis et al., 2018).

Furthermore, an increase in the commodity prices lead to higher government spending. Then if the monetary authority of the commodity-exporting economy responds with contractionary monetary policy, the interest rates will rise, which makes lending more expensive (Bjørnland, 2009). At the same time, an increase in the commodity prices forces the commodity-producing firms to increase their investment activity and, therefore, borrow money from the bank. The revenue effect is generally larger than the production cost effect in a commodity-exporting economy and, therefore, the stock prices are expected to increase (Oberndorfer, 2009).

Overall, the theoretical framework of this study is based on the concept of equity pricing model. On the one hand, the primary commodities are one of the major production factors for a commodity-consuming firm; therefore, a raise in commodity prices increases the production costs. As such, profit levels are expected to decline and, thus, the future cash flow decreases (Mork et al., 1994; Filis et al., 2011). On the other hand, for a commodity-producer, the commodity price increase results in an increase of profit margins and, thus, the expected cash flow rises. As there are more companies in the world that consume commodities than the ones that produce them, the overall impact of rising commodity prices on stock markets is expected to be negative (Degiannakis et al., 2018).

In summary, the equity pricing model suggests that there could be a relationship between commodity prices and stock returns, due to the following reasons:

- First, because some commodities, such as fuels and metals, are primary inputs for most firms, an increase in their prices raises the cost of production, reduces the future cash flow, earnings and dividends and, therefore, the stock returns decline.
- Second, an upsurge in the prices of metals and fuels can lead to an overestimation of the expected inflation and higher nominal interest rates. As interest rates are used to discount the expected future cash flow, this will depress earnings, dividends and, thus, stock returns (Smyth and Narayan, 2018).

The diagram below visualises the transmission channels from commodity prices to firms' stock prices.

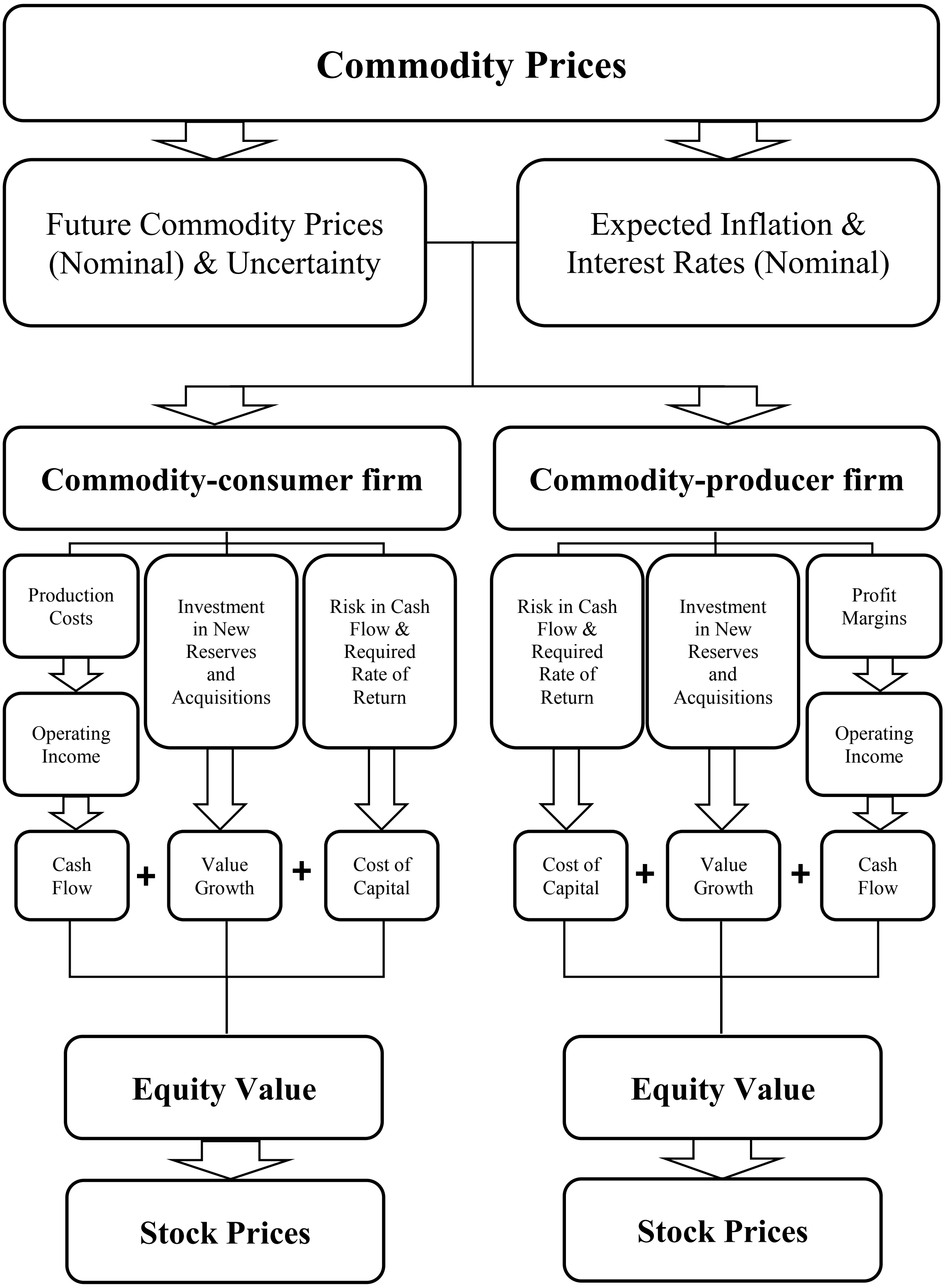

Figure 4.1 Commodity Prices and Stock Markets

4.4 Methodology

In this study, we investigate the relationship between world commodity prices and stock markets by using VAR models and Granger causality tests in a time-varying setting. We particularly exploit the MF-VAR approach proposed by Ghysels et al. (2016), a procedure that, as aforementioned, overcomes some of the shortcomings of standard single-frequency VAR models and, more specifically, the temporal aggregation bias (Ghysels, 2016). This allows us to exploit a richer data set and perform a more insightful analysis. Using the MF-VAR model, we conduct Granger causality tests based on the Wald statistics to test the null hypothesis of non-causality. Then, the simulated MF data is aggregated into LF, and causality tests are conducted based on both approaches. We allow for a direct comparison between the MF and the traditional LF methods.

The subsequent sections provide a brief description of the methods and specifications used in the empirical analysis.

4.4.1 Mixed-frequency VAR

The MF-VAR model is an observation-driven model that directly relates to standard VAR model settings and is suitable for exploiting Granger causality tests (Ghysels, 2016). Following the notation of Ghysels et al. (2016), we denote m to be the *ratio of sampling frequencies*. In other words, m represents the number of high-frequency time periods in each low-frequency time period τ , where $\tau \epsilon\{1,2,\dots,T_L\}$ is a time sequence. Thus, let $CP_{(\tau,j)}$ denote the series of commodity prices at the j-th week of month τ with $j\epsilon\{1,2,3,4\}$, while $SP_{(\tau,j)}$ denote the series of stock prices at month τ. Section 4.5 provides more details about the data construction. Assume that covariance stationarity is satisfied for each series. The variables are at weekly and monthly frequencies, so that we fix $m = 4$ according to Ghysels et al. (2017). Then, the MF-VAR (p) model is specified as follows:

$$\underbrace{\begin{bmatrix} CP_{(\tau,1)} \\ CP_{(\tau,2)} \\ CP_{(\tau,3)} \\ CP_{(\tau,4)} \\ SP_{(\tau)} \end{bmatrix}}_{\equiv X_{(\tau)}} = \sum_{k=1}^{p} \underbrace{\begin{bmatrix} a_{11,k} & a_{12,k} & a_{13,k} & a_{14,k} & a_{15,k} \\ a_{21,k} & a_{22,k} & a_{23,k} & a_{24,k} & a_{25,k} \\ a_{31,k} & a_{32,k} & a_{33,k} & a_{34,k} & a_{35,k} \\ a_{41,k} & a_{42,k} & a_{43,k} & a_{44,k} & a_{45,k} \\ a_{51,k} & a_{52,k} & a_{53,k} & a_{54,k} & a_{55,k} \end{bmatrix}}_{\equiv A_k} \underbrace{\begin{bmatrix} CP_{(\tau-k,1)} \\ CP_{(\tau-k,2)} \\ CP_{(\tau-k,3)} \\ CP_{(\tau-k,4)} \\ SP_{(\tau-k)} \end{bmatrix}}_{\equiv X_{(\tau-k)}} + \underbrace{\begin{bmatrix} \varepsilon_{(\tau,1)} \\ \varepsilon_{(\tau,2)} \\ \varepsilon_{(\tau,3)} \\ \varepsilon_{(\tau,4)} \\ \varepsilon_{(\tau,5)} \end{bmatrix}}_{\equiv \varepsilon_{(\tau)}} \tag{4.3}$$

where A_k is a coefficient square matrix for $k = 1,\dots,p$, where p is the lag length and $\varepsilon_{(\tau)}$ is the vector of residuals. Rather than working on aggregate monthly data, all of the weekly observations are stacked in each month period τ to obtain $X_{(\tau)} = \left[CP_{(\tau,1)}{}', CP_{(\tau,2)}{}', CP_{(\tau,3)}{}', CP_{(\tau,4)}{}', SP_{(\tau)}{}'\right]'$. Following Ghysels et al. (2016), the constant term is not included in Equation (4.3). The lag length is selected by using the BIC, which is consistent with the studies of Kuzin et al. (2011) and Bai et al. (2013). Therefore, $X_{(\tau)}$ should be thought of as a de-meaned process. The MF-VAR (p) model in Equation (4.3) can be written as follows:

$$X_{(\tau)} = \sum_{k=1}^{p} A_k X_{(\tau-k)} + \varepsilon_{(\tau)} \tag{4.4}$$

$\varepsilon_{(\tau)} \sim (0, \sigma^2)$, $\sigma^2 > 0$.

To investigate the long-run Granger causality between commodity and stock prices, we iterate Equation (4.4) over the desired test horizon h and lag order p, obtaining the following MF-VAR (p, h) model:

$$X_{(\tau+h)} = \sum_{k=1}^{p} A_k^{(h)} X_{(\tau+1-k)} + u_{(\tau)}^{(h)} \tag{4.5}$$

where $A_k^{(1)} = A_k$, $A_k^{(i)} = A_{k+i-1} + \sum_{l=1}^{i-1} A_{i-l}\, A_k^{(l)}$ for $i \geq 2$, $u_{(\tau)}^{(h)} = \sum_{k=0}^{h-1} \Psi_k\, \varepsilon_{(\tau-k)}$ and by convention, $A_k^{(1)} = 0_{k\times k}$ whenever $k > p$.

Following Ghysels et al. (2016), we make the following assumptions.[102] First, all roots of the polynomial $det(I_5 - \sum_{k=1}^{p} A_k\, z^k) = 0$ lie outside the unit circle, where $det(.)$ is the determinant. This ensures that the MF-VAR is state stationary. Second, $\varepsilon_{(\tau)}$ is a strictly stationary martingale difference sequence with a finite second moment. Third, $\{X_{(\tau)}, \varepsilon_{(\tau)}\}$ obey α-mixing that satisfies $\sum_{h=0}^{\infty} \alpha_{2^h} < \infty$. This is a standard assumption to ensure the validity of the bootstrap for VAR models (for example, see Paparoditis, 1996; Kilian, 1998; Cavaliere et al., 2012; Cavaliere et al., 2014). In fact, these assumptions ensure the consistency and asymptotic normality of the least squares estimator $\widehat{A_k}$.[103]

Next, we exploit Wald statistics based on the coefficients of MF-VAR (p, h), $B(h) = [A_1^{(h)}, \ldots, A_p^{(h)}]'$. For example, CP do not Granger-cause SP given a MF information set equal to $a_{51,1} = \cdots = a_{52,1} = \cdots = a_{53,1} = \cdots = a_{54,p} = 0_{1\times m}$, whereas SP do not Granger-cause CP given a MF information set equal to $a_{15,1} = \cdots = a_{25,1} = \cdots = a_{35,1} = \cdots = a_{45,p} = 0_{m\times 1}$. Therefore, the null hypothesis of non-causality is a linear restriction defined as follows:

$$H_0\colon Rvec[B(h)] = r \tag{4.6}$$

where R is a $q \times pK^2$selection matrix of full row rank q. $K = mK_H + K_L$, where K_H is the number of high frequency variables and $K_L$is the number of low-frequency variables. Here, $K_H = 1$ and $K_L = 1$. The complete details of the construction of R can be found in Ghysels et al. (2016). r is a restricted vector, and zeros are always chosen when performing Granger

[102] In terms of the asymptotic theory, the MF-VAR can be treated in the same manner as a classical VAR. Therefore, all standard regularity conditions carry over.
[103] See Ghysels et al. (2016) for technical details.

causality tests. Thus, the null hypothesis of the MF Granger causality test can be expressed via the following Wald statistic:

$$W_{T_L^*}[H_0(h)] \equiv T_L^*(Rvec[\hat{B}(h)] - r)' \times \left(R\hat{\Sigma}_p(h)R'\right)^{-1} \times \left(Rvec[\hat{B}(h)] - r\right) \quad (4.7)$$

where $T_L^* \equiv T_L - h + 1$ is the effective sample size of the MF-VAR (p, h) model, $\hat{B}(h)$ is the least square estimator of the MF-VAR (p, h) model, $\hat{\Sigma}_p\ (h)$ is positive semi-definite for any $T_L^* \geq 1$, and $\hat{\Sigma}_p\ (h) \xrightarrow{p} \Sigma_p(h)$ where $\Sigma_p(h)$ is positive definite (Ghysels et al., 2016).[104]

Last, following Ghysels et al. (2016) to circumvent size distortions for small samples $\tau\epsilon\{50, 100\}$, parametric bootstraps by Gonçalves and Kilian (2004) are employed. Gonçalves and Kilian's (2004) recursive design parametric wild bootstrap does not require knowledge of the true error distribution and is robust to conditional heteroskedasticity of an unknown form. The bootstrap method is employed to improve the empirical size in small samples (Davidson and MacKinnon, 2006). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 499$ replications, as suggested by Ghysels et al. (2016). Hence, we compute the resulting p-value of Equation (4.7), which is defined as follows:

$$\hat{p}_N(W_{T_L^*}[H_0(h)]) = \frac{1}{N+1} \times \left(1 + \sum_{i=1}^{N} I(W_i[H_0(h)] \geq W_{T_L^*}[H_0(h)])\right) \quad (4.8)$$

where $W_i[H_0(h)]$ is the Wald test statistic based on the ith simulation sample and the null hypothesis $\mathrm{H_0(h)}$ is rejected at level α if $\hat{\mathrm{p}}_{\mathrm{N}}\left(\mathrm{W}_{\mathrm{T_L^*}}[\mathrm{H_0(h)}]\right) \leq \alpha$.[105]

The next section presents the setting of the LF-VAR model. In particular, we use both of the methods, i.e. LF-VAR and MF-VAR, in order to provide evidence that the choice of sampling frequency can alter the empirical results considerably.

4.4.2 Low-frequency VAR

This section formulates the LF-VAR model, which is used to examine the relationship between monthly commodity prices and monthly stock prices. The LF-VAR is a standard single-frequency VAR model. The notation $CP_{(\tau)}^M$ is the commodity price at month τ; $SP_{(\tau)}$ is the stock price at month τ. Superscript "M" is put in order to explicitly distinguish a monthly level from a weekly level data. Since the number of weeks contained in each month τ is not

[104] Following Ghysels et al. (2016), this study uses Newey and West's (1987) Bartlett kernel-based HAC covariance estimator, which ensures positive semi-definiteness for any $T_L^* \geq 1$, with Newey and West's (1994) automatic bandwidth selection.
[105] See Ghysels et al. (2016) for technical details.

constant, this study follows Ghysels et al. (2017) in terms of simplifying the analysis by taking a sample average for each τ.

This chapter hereafter distinguishes between weekly commodity prices $\{CP_{(\tau,1)}, CP_{(\tau,2)}, CP_{(\tau,3)}, CP_{(\tau,4)}\}$, monthly commodity prices $CP_{(\tau)}^M$ and a general notion of commodity prices CP in order to avoid notational confusion. Then, the specification of the LF-VAR (p) model is given as follows:

$$\begin{bmatrix} CP_{(\tau)}^M \\ SP_{(\tau)} \end{bmatrix} = \sum_{k=1}^{p} \begin{bmatrix} a_{11,k} & a_{12,k} \\ a_{21,k} & a_{22,k} \end{bmatrix} \begin{bmatrix} CP_{(\tau-k)}^M \\ SP_{(\tau-k)} \end{bmatrix} + \begin{bmatrix} \varepsilon_{(\tau,1)} \\ \varepsilon_{(\tau,2)} \end{bmatrix} \tag{4.9}$$

Following Ghysels et al. (2016), the constant term is omitted and each series is de-meaned before fitting the model. In line with the empirical study of Kilian and Park (2009), we consider the period-to-period log-difference of the level series (see Section 4.5 for further details). If a time series does not satisfy the covariance stationarity after the first differencing, that series (i.e. country) is excluded from the study sample. This is because a further differencing of the level series, i.e. a second or higher differencing, does not make economic sense. All series are normalised by their full sample mean and standard deviation. The assumptions that are made for the MF-VAR (p) models apply to the LF-VAR models as well. For a detailed discussion on the data handling, see Section 4.5.

4.4.3 Time-varying estimation

We adopt a time-varying approach for both the MF-VAR and LF-VAR models in order to account for the structural changes that occurred during the time period that is considered, especially for the case of emerging markets (see Smyth and Narayan, 2018 for further discussion). We follow Chen et al. (2010) by using a rolling window estimation than a recursive one, as it adapts more quickly to possible structural changes. The rolling procedure is relatively robust against the presence of time-varying parameters and requires no explicit assumption regarding the nature of time variation in the data. Following Chen et al. (2010), we estimate the model parameters using a rolling window that is half the size of the total sample size.

4.5 Data

To conduct the empirical analysis mixed monthly and weekly data for financial and commodity markets, respectively, of 63 countries and territories from January 1951 to March 2018 is used. The start date is influenced by the availability of data (see Appendix C.1). It should be noted that continuous stock market data is not available at a daily or weekly frequency so far into the past, especially for developing countries. Therefore, consistent with previous studies, we use monthly stock market indexes.[106] This allows us to observe a larger number of countries than otherwise possible.

Specifically, as a proxy for the national stock market returns, we use the main stock index for each country (see Appendix C.1). Following Kilian and Park (2009), the stock market returns are calculated as log returns, i.e. $SP_t^R = \ln(SP_t) - \ln(SP_{t-1})$.

Further, the study considers five measures of global commodities that are defined as global shocks: world oil price, world oil production, world economic activity, world commodity prices in general (which include commodity prices from all sectors) and world metal prices.[107] The data for global shocks is available at a weekly frequency, thereby requiring temporal aggregation to fit the traditional single-frequency methods. As has been argued extensively in the MF literature (see Clements and Galvão, 2008; Marcellino and Schumacher, 2010; Ghysels and Miller, 2015; Ghysels et al., 2016 for discussion), working in a single LF setting has certain disadvantages due to the potential loss of information (Götz et al., 2016). Therefore, we combine the weekly data on global shocks with the monthly data on stock markets in a MF setting. All data is sourced from the Datastream (2018) database.

As a proxy for world oil prices, we use the weekly price data of West Texas Intermediate (WTI) crude oil. The WTI oil price is widely used as the benchmark for oil pricing and is highly correlated with the prices of the two other major categories of crude oil – Brent and Dubai crude oils (see Borenstein et al., 1997; Kilian, 2009; Kilian and Park, 2009; Phan et al., 2015). The WTI price data is denominated in US dollars and is obtained by averaging daily data.

In addition, we use two different proxies for global oil supply and demand (Kilian, 2009). As a proxy for supply, we use the weekly global oil production data (consistent with Kilian, 2009) and, as a proxy for demand, we use the weekly Baltic Exchange Dry Index (BDI) to estimate the scale of global economic activity (consistent with Conrad et al., 2018). The BDI index is highly correlated with the global oil demand proxy that was constructed by Kilian (2009). The BDI index is used in nominal terms in order to retain consistency with the stock index series.

Last but not least, we use the weekly CRB Commodity Price Index as a proxy for the world commodity price. This index is broadly used in the literature as a proxy for global commodity

[106] For example, see Sadorsky (1999), Park and Ratti (2008) and Kilian and Park (2009).
[107] Unless otherwise stated, the term “world commodity prices (all items)” refers hereafter to the “world commodity prices in general”.

prices (see Creti et al., 2013; Silvennoinen and Thorp, 2013; Chen et al., 2014). Moreover, it has the advantage of being available since January 1951. In addition, we consider the weekly CRB Metals Sub-Index as a proxy for global metal prices. Selecting the CRB Metals Sub-Index as a representative of the world metal prices is consistent with the most recent commodity-stock literature (for example, see Beckmann et al., 2014; Lu and Jacobsen, 2016), examining the link between individual metal prices (predominantly gold) and stock market prices (see Hood and Malik, 2013; Sadorsky, 2014; Arouri et al., 2015; O'Connor et al., 2015; Basher and Sadorsky, 2016). Finally, it should be noted that all the above series are denominated in US dollars to ensure comparability and are calculated as one-period log returns. All data is found to be stationary using the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests.[108]

4.6 Empirical Results

This section presents and discusses the results of the empirical testing. First, the relationship between commodity and financial markets is examined by using the full sample Granger causality tests for both the LF-VAR and MF-VAR models. Second, we use time-varying Granger causality tests to explore the short-run dynamic relationship between commodity and stock markets. Similar to the full sample approach, the results for both the LF-VAR and the MF-VAR are presented.

4.6.1 Full sample Granger causality tests

The full sample Granger causality test results for both the MF-VAR and LF-VAR models are reported in tables 4.1 and 4.2 respectively. The null hypothesis of non-causality is tested against the alternative of causality. The rejection of the null hypothesis $H_0: SP \nRightarrow CP$ means that stock market returns Granger-cause global shocks, against the alternative hypothesis that stock market returns do not Granger-cause global shocks. Analogously, we test the null hypothesis $H_0: CP \nRightarrow SP$.

[108] Details of the unit root tests and other preliminary statistics are available in Appendix C.2.

	World oil prices		Oil supply shocks		Oil demand shocks		World commodity prices (all items)		World metal prices	
	SP ⇏ CP	CP ⇏ SP	SP ⇏ CP	CP ⇏ SP	SP ⇏ CP	CP ⇏ SP	SP ⇏ CP	CP ⇏ SP	SP ⇏ CP	CP ⇏ SP
Africa										
Panel A: Northern Africa										
Egypt	0.570	0.384	0.604	0.018	0.004	0.400	0.440	0.280	0.236	0.286
Morocco	0.888	0.466	0.994	0.056	0.148	0.254	0.948	0.002	0.132	0.002
Tunisia	0.516	0.408	0.814	0.840	0.200	0.698	0.078	0.876	0.384	0.310
Panel B: Sub-Saharan Africa										
Kenya	0.612	0.566	0.766	0.154	0.650	0.160	0.566	0.464	0.010	0.810
Malawi	0.084	0.862	0.402	0.862	0.382	0.376	0.360	0.486	0.198	0.812
Mauritius	0.186	0.310	0.384	0.574	0.016	0.558	0.364	0.394	0.082	0.078
Uganda	0.952	0.282	0.580	0.558	0.330	0.876	0.208	0.068	0.114	0.684
Tanzania	0.168	0.870	0.628	0.610	0.882	0.786	0.526	0.284	0.612	0.166
Zambia	0.726	0.882	0.770	0.200	0.230	0.202	0.250	0.050	0.076	0.016
Namibia	0.026	0.254	0.066	0.426	0.816	0.760	0.332	0.248	0.104	0.202
South Africa	0.382	0.302	0.036	0.002	0.154	0.336	0.220	0.016	0.354	0.070
Ghana	0.698	0.642	0.286	0.668	0.774	0.746	0.848	0.334	0.478	0.012
Nigeria	0.780	0.926	0.508	0.544	0.800	0.854	0.812	0.206	0.936	0.240
West African Economic and Monetary Union	0.278	0.452	0.062	0.010	0.446	0.802	0.008	0.002	0.032	0.010
Americas										
Panel C: Latin America and the Caribbean										
Mexico	0.378	0.980	0.318	0.292	0.676	0.916	0.860	0.502	0.508	0.802
Argentina	0.198	0.758	0.060	0.690	0.122	0.346	0.924	0.124	0.356	0.052
Brazil	0.240	0.226	0.974	0.074	0.054	0.442	0.276	0.330	0.250	0.164
Chile	0.828	0.372	0.828	0.102	0.156	0.406	0.508	0.106	0.032	0.082
Colombia	0.914	0.974	0.700	0.014	0.416	0.566	0.504	0.904	0.360	0.748
Panel D: Northern America										
Canada	0.636	0.978	0.574	0.630	0.348	0.710	0.110	0.066	0.062	0.360
United States of America (the US)	0.352	0.906	0.542	0.618	0.478	0.312	0.484	0.338	0.036	0.664
Asia										
Panel E: Central Asia										
Kazakhstan	0.168	0.142	0.220	0.242	0.134	0.054	0.030	0.452	0.028	0.336
Panel F: Eastern Asia										
China (Mainland)	0.792	0.562	0.188	0.908	0.292	0.280	0.496	0.958	0.572	0.490
Hong Kong	0.540	0.776	0.130	0.966	0.104	0.496	0.632	0.100	0.576	0.214
Japan	0.728	0.974	0.360	0.988	0.092	0.994	0.286	0.368	0.036	0.856
South Korea	0.814	0.376	0.910	0.018	0.186	0.450	0.916	0.672	0.472	0.968
Taiwan	0.614	0.996	0.660	0.882	0.164	0.304	0.140	0.458	0.062	0.154
Panel G: South-eastern Asia										
Indonesia	0.760	0.842	0.302	0.482	0.474	0.630	0.200	0.128	0.060	0.204
Malaysia	0.502	0.466	0.972	0.760	0.332	0.676	0.112	0.492	0.368	0.250
Philippines	0.638	0.482	0.676	0.410	0.428	0.566	0.168	0.330	0.240	0.280
Thailand	0.404	0.478	0.396	0.236	0.302	0.936	0.072	0.746	0.220	0.646
Panel H: Southern Asia										
Bangladesh	0.876	0.134	0.910	0.044	0.154	0.964	0.926	0.050	0.640	0.414

India	0.428	0.876	0.246	0.542	0.486	0.384	0.446	0.076	0.086	0.032
Iran	0.010	0.178	0.106	0.250	0.170	0.408	0.188	0.460	0.300	0.530
Sri Lanka	0.626	0.938	0.142	0.098	0.162	0.854	0.352	0.318	0.942	0.154
Panel I: Western Asia										
Israel	0.840	0.194	0.266	0.590	0.852	0.086	0.840	0.142	0.926	0.098
Saudi Arabia	0.270	0.210	0.574	0.060	0.458	0.120	0.162	0.930	0.282	0.240
Turkey	0.776	0.724	0.452	0.072	0.186	0.226	0.510	0.570	0.728	0.752
						Europe				
Panel J: Eastern Europe										
Czech Republic	0.608	0.974	0.058	0.012	0.384	0.310	0.376	0.184	0.228	0.058
Hungary	0.760	0.834	0.184	0.008	0.642	0.716	0.314	0.104	0.240	0.142
Poland	0.912	0.804	0.018	0.364	0.596	0.652	0.136	0.266	0.156	0.086
Russia	0.392	0.402	0.628	0.540	0.732	0.548	0.562	0.144	0.036	0.008
Slovakia	0.772	0.942	0.168	0.174	0.282	0.990	0.300	0.432	0.020	0.768
Ukraine	0.082	0.048	0.240	0.008	0.242	0.120	0.232	0.444	0.042	0.546
Panel K: Northern Europe										
Finland	0.656	0.948	0.112	0.546	0.228	0.924	0.830	0.410	0.028	0.052
Iceland	0.006	0.224	0.374	0.130	0.154	0.742	0.158	0.720	0.344	0.230
Ireland	0.776	0.754	0.622	0.726	0.388	0.852	0.950	0.222	0.764	0.034
Lithuania	0.060	0.980	0.450	0.602	0.104	0.874	0.274	0.316	0.076	0.320
Norway	0.138	0.004	0.370	0.002	0.066	0.740	0.168	0.014	0.280	0.002
United Kingdom	0.168	0.634	0.416	0.996	0.504	0.614	1.000	0.230	0.326	0.690
Panel L: Southern Europe										
Croatia	0.502	0.560	0.416	0.746	0.538	0.578	0.244	0.190	0.096	0.174
Greece	0.524	0.368	0.244	0.462	0.074	0.576	0.502	0.744	0.124	0.406
Italy	0.656	0.030	0.814	0.388	0.264	0.744	0.678	0.044	0.358	0.310
Portugal	0.626	0.572	0.080	0.014	0.062	0.862	0.652	0.386	0.376	0.338
Spain	0.180	0.900	0.904	0.268	0.078	0.184	0.828	0.062	0.506	0.040
Panel M: Western Europe										
Belgium	0.336	0.686	0.680	0.526	0.700	0.750	0.390	0.448	0.776	0.704
France	0.210	0.568	0.592	0.044	0.032	0.878	0.316	0.054	0.506	0.068
Germany	0.526	0.468	0.680	0.860	0.052	0.532	0.912	0.010	0.320	0.670
Netherlands	0.394	0.448	0.476	0.880	0.378	0.740	0.770	0.150	0.332	0.576
Switzerland	0.410	0.952	0.056	0.300	0.892	0.934	0.914	0.044	0.174	0.162
Panel N: Europe										
Euro Zone	0.264	0.892	0.376	0.002	0.002	0.862	0.512	0.278	0.872	0.136
						Oceania				
Panel O: Australia and New Zealand										
Australia	0.536	0.948	0.260	0.332	0.290	0.772	0.412	0.040	0.064	0.096
New Zealand	0.810	0.828	0.354	0.444	0.822	0.116	0.296	0.048	0.372	0.342

Note: The table contains bootstrapped p-values for the full sample MF Granger causality tests. The MF approach uses weekly measures of global variables and monthly stock returns. "SP" denotes the stock market returns, while "CP" denotes the global variables. $H_0\!: SP \nRightarrow CP$ *($\nRightarrow$ means "does not Granger-cause"). We follow Ghysels et al. (2016) and use bootstrapped p-values with N = 499 replications (Gonçalves and Kilian, 2004). All variables are mean-centred and log-differenced.*

Table 4.1 P-values for Mixed-Frequency Tests of Non-Causality

	World oil prices		Oil supply shocks		Oil demand shocks		World commodity prices (all items)		World metal prices	
	SP $\nRightarrow$ CP	CP $\nRightarrow$ SP	SP $\nRightarrow$ CP	CP $\nRightarrow$ SP	SP $\nRightarrow$ CP	CP $\nRightarrow$ SP	SP $\nRightarrow$ CP	CP $\nRightarrow$ SP	SP $\nRightarrow$ CP	CP $\nRightarrow$ SP
Africa										
Panel A: Northern Africa										
Egypt	0.118	0.622	0.198	0.736	0.002	0.498	0.134	0.746	0.120	0.118
Morocco	0.346	0.310	0.684	0.368	0.032	0.554	0.524	0.002	0.066	0.002
Tunisia	0.102	0.472	0.896	0.842	0.034	0.626	0.044	0.434	0.104	0.048
Panel B: Sub-Saharan Africa										
Kenya	0.910	0.378	0.734	0.134	0.328	0.836	0.098	0.772	0.026	0.314
Malawi	0.904	0.952	0.614	0.954	0.178	0.216	0.148	0.608	0.664	0.166
Mauritius	0.040	0.460	0.516	0.158	0.006	0.136	0.052	0.080	0.048	0.022
Uganda	0.698	0.008	0.368	0.172	0.986	0.904	0.478	0.734	0.580	0.606
Tanzania	0.888	0.988	0.714	0.578	0.338	0.160	0.482	0.330	0.928	0.170
Zambia	0.084	0.386	0.546	0.040	0.510	0.168	0.082	0.006	0.034	0.002
Namibia	0.764	0.048	0.174	0.080	0.244	0.456	0.654	0.092	0.578	0.012
South Africa	0.092	0.294	0.422	0.002	0.192	0.358	0.082	0.068	0.076	0.044
Ghana	0.214	1.000	0.056	0.342	0.306	0.530	0.780	0.394	0.910	0.014
Nigeria	0.600	0.322	0.676	0.274	0.698	0.422	0.574	0.066	0.552	0.686
West African Economic and Monetary Union	0.210	0.186	0.858	0.038	0.046	0.462	0.304	0.002	0.184	0.002
Americas										
Panel C: Latin America and the Caribbean										
Mexico	0.578	0.846	0.332	0.094	0.406	0.534	0.488	0.102	0.286	0.298
Argentina	0.490	0.530	0.960	0.322	0.016	0.084	0.508	0.016	0.142	0.024
Brazil	0.036	0.286	0.782	0.554	0.016	0.344	0.158	0.884	0.174	0.020
Chile	0.304	0.618	0.778	0.162	0.040	0.778	0.048	0.024	0.006	0.008
Colombia	0.878	0.468	0.716	0.222	0.300	0.472	0.284	0.320	0.166	0.244
Panel D: Northern America										
Canada	0.292	0.764	0.988	0.206	0.022	0.492	0.236	0.614	0.116	0.186
United States of America (the US)	0.908	0.646	0.966	0.848	0.710	0.240	0.430	0.468	0.278	0.740
Asia										
Panel E: Central Asia										
Kazakhstan	0.050	0.598	0.030	0.604	0.186	0.390	0.006	0.882	0.002	0.406
Panel F: Eastern Asia										
China (Mainland)	0.986	0.486	0.390	0.840	0.038	0.812	0.900	0.722	0.176	0.234
Hong Kong	0.222	0.768	0.320	0.882	0.074	0.366	0.366	0.176	0.150	0.708
Japan	0.416	0.986	0.226	0.764	0.074	0.828	0.060	0.814	0.020	0.424
South Korea	0.574	0.656	0.954	0.314	0.038	0.226	0.682	0.352	0.404	0.458
Taiwan	0.348	0.998	0.960	0.612	0.306	0.116	0.306	0.276	0.046	0.310
Panel G: South-eastern Asia										
Indonesia	0.772	0.656	0.474	0.214	0.666	0.306	0.078	0.188	0.026	0.108
Malaysia	0.042	0.698	0.800	0.668	0.354	0.946	0.006	0.876	0.162	0.168
Philippines	0.554	0.818	0.750	0.976	0.926	0.214	0.226	0.086	0.956	0.050
Thailand	0.992	0.324	0.190	0.580	0.160	0.478	0.378	0.910	0.164	0.474
Panel H: Southern Asia										
Bangladesh	0.708	0.042	0.316	0.060	0.294	0.952	0.884	0.136	0.784	0.086

India	0.388	0.252	0.848	0.788	0.306	0.648	0.096	0.918	0.028	0.008
Iran	0.148	0.106	0.134	0.020	0.436	0.318	0.222	0.002	0.656	0.038
Sri Lanka	0.408	0.902	0.412	0.106	0.024	0.612	0.088	0.018	0.360	0.022
Panel I: Western Asia										
Israel	0.656	0.704	0.988	0.198	0.290	0.044	0.574	0.088	0.744	0.006
Saudi Arabia	0.062	0.040	0.094	0.030	0.634	0.940	0.234	0.784	0.224	0.174
Turkey	0.702	0.774	0.486	0.320	0.070	0.490	0.244	0.576	0.572	0.312
						Europe				
Panel J: Eastern Europe										
Czech Republic	0.860	0.414	0.964	0.744	0.156	0.322	0.050	0.992	0.208	0.016
Hungary	0.634	0.278	0.534	0.040	0.312	0.906	0.158	0.234	0.442	0.012
Poland	0.780	0.538	0.818	0.718	0.332	0.444	0.124	0.808	0.362	0.076
Russia	0.274	0.226	0.592	0.226	0.380	0.452	0.196	0.088	0.090	0.002
Slovakia	0.988	0.540	0.856	0.844	0.068	0.854	0.324	0.664	0.676	0.304
Ukraine	0.104	0.034	0.640	0.026	0.070	0.586	0.048	0.140	0.016	0.386
Panel K: Northern Europe										
Finland	0.244	0.412	0.948	0.506	0.076	0.488	0.232	0.816	0.312	0.252
Iceland	0.030	0.200	0.514	0.964	0.064	0.312	0.570	0.616	0.916	0.280
Ireland	0.328	0.968	0.406	0.546	0.102	0.578	0.640	0.424	0.918	0.120
Lithuania	0.220	0.604	0.562	0.566	0.074	0.374	0.302	0.490	0.158	0.038
Norway	0.010	0.484	0.404	0.002	0.002	0.328	0.060	0.002	0.054	0.002
United Kingdom	0.198	0.232	0.210	0.678	0.124	0.268	0.968	0.808	0.832	0.832
Panel L: Southern Europe										
Croatia	0.472	0.458	0.526	0.468	0.204	0.468	0.346	0.252	0.098	0.022
Greece	0.170	0.150	0.930	0.846	0.006	0.634	0.166	0.360	0.064	0.080
Italy	0.506	0.084	0.580	0.382	0.036	0.360	0.936	0.258	0.534	0.052
Portugal	0.474	0.192	0.720	0.850	0.016	0.912	0.204	0.436	0.176	0.040
Spain	0.688	0.490	0.718	0.068	0.080	0.220	0.338	0.076	0.848	0.046
Panel M: Western Europe										
Belgium	0.296	0.564	0.438	0.822	0.174	0.276	0.136	0.294	0.488	0.266
France	0.158	0.470	0.588	0.002	0.006	0.678	0.222	0.154	0.260	0.124
Germany	0.146	0.334	0.568	0.838	0.266	0.210	0.612	0.498	0.254	0.416
Netherlands	0.180	0.098	0.680	0.504	0.272	0.242	0.550	0.898	0.500	0.438
Switzerland	0.652	0.956	0.810	0.052	0.406	0.580	0.670	0.008	0.696	0.038
Panel N: Europe										
Euro Zone	0.248	0.738	0.154	0.002	0.030	0.940	0.810	0.096	0.910	0.234
						Oceania				
Panel O: Australia and New Zealand										
Australia	0.504	0.818	0.272	0.054	0.282	0.274	0.236	0.710	0.334	0.464
New Zealand	0.570	0.610	0.674	0.526	0.720	0.344	0.428	0.306	0.500	0.100

Note: The table contains bootstrapped p-values for the full sample LF Granger causality tests. The LF approach uses monthly measures of global variables and monthly stock returns. "SP" denotes the stock market returns, while "CP" denotes the global variables. $H_0: SP \nRightarrow CP$ *(*$\nRightarrow$ *means "does not Granger-cause"). We follow Ghysels et al. (2016) and use bootstrapped p-values with N = 499 replications (Gonçalves and Kilian 2004). All variables are mean-centred and log-differenced.*

Table 4.2 P-values for Low-Frequency Tests of Non-Causality

Tables 4.1 and 4.2 display the bootstrapped p-values for the full sample MF and LF Granger causality tests respectively. For example, the p-value for testing the null hypothesis $H_0: SP \nRightarrow CP$ ($\nRightarrow$ means "does not Granger-cause") that stock market returns do not Granger-cause world oil prices in the case of the full sample MF model for Egypt is 0.570. Therefore, the null hypothesis cannot be rejected at a 10% significance level, which implies that the Egyptian stock market returns do not have predictive power on the world oil prices. The same conclusion is reached when using the LF method. Another example is Kazakhstan. The p-value for testing the null hypothesis that oil demand shocks do not Granger-cause Kazakh stock market returns in terms of the MF model is 0.054. This concludes that the oil demand shocks possess a predictive power over Kazakh stock market returns at a 10% significance level. This is not surprising, as Kazakhstan can be classified as a large oil-exporting country. However, the LF method fails to identify a causal relationship between oil demand shocks and Kazakh stock market returns. This discrepancy between the two methods can possibly be attributed to the loss of information from temporal data aggregation, as discussed by Ghysels (2016).

As a whole, the full sample tests suggest that a causal relationship between world oil prices and national stock market returns is hardly detected. Using the MF approach, causality from world oil prices to national stock returns is found only for the European countries: Italy, Norway and Ukraine. As Norway is a major world exporter of oil and gas and Italy is a major importer, it is not surprising that their stock markets are influenced by oil prices. These results are consistent with earlier works of Driesprong et al. (2008) and Wang et al. (2013) in the context of Norway and that of Apergis and Miller (2009) in the context of Italy.

As discussed in the introduction, while the literature has concentrated on developed countries, less is known about developing economies. The results of the LF approach, presented in table 4.2, suggest that the world oil prices have impact on stock market returns in five developing countries (namely Uganda, Namibia, Bangladesh, Saudi Arabia and Ukraine) and in only two developed countries (namely Italy and Netherlands). Notably, all these countries, apart from Saudi Arabia, are oil-importers. Therefore, we can conclude that world oil prices have a greater impact on the oil-importing rather than oil-exporting economies.

At the same time, the impact of oil supply shocks on national stock markets is found to be much stronger than that of the world oil prices. In particular, the MF approach suggests that oil supply shocks Granger-cause stock market returns in 18 countries and regions. Most of these countries (and regions) are oil-importers, i.e. Bangladesh, Czech Republic, the Euro Zone, France, Hungary, Morocco, Portugal, South Africa, South Korea, Sri Lanka, Ukraine and West African Economic and Monetary Union, suggesting a higher sensitivity of their national stock markets to oil supply shocks as compared to the exporting countries, i.e. Egypt, Colombia, Brazil, Norway, Saudi Arabia and Turkey. The LF results are similar and confirm this conclusion.

For the specific case of France, both the MF and LF approaches suggest that the French stock market is affected by oil supply shocks. This adds to the study of Apergis and Miller (2009) who find that only global oil demand shocks temporarily lead the French stock market returns. The longer time span of our sample can contribute to explaining the difference between their finding and ours.

Furthermore, the impact of oil demand shocks is found to be less influential on national stock markets (see tables 4.1 and 4.2). Both the MF and LF approaches find that oil demand shocks lead stock market returns for only two out of 63 countries. According to the MF approach, these countries are Kazakhstan and Israel, however, the LF method discovers causal patterns for Mexico and Israel. Despite the difference in the findings, the overall conclusion from these results is that world economic activity has a negligible influence on stock market returns.

While the literature primarily focuses on the oil-stock relationship, less is known about the association between commodity markets as a whole and national stock markets. As highlighted by Deaton and Miller (1995), a single commodity price (e.g. oil price) may not well approximate the entire commodity market. This is true, especially for the newly industrialised economies that have more diversified trade exports and imports but are still dependent on primary commodities – for example, Australia, Canada and New Zealand (see also Chen and Rogoff, 2003; Cashin et al., 2004 for discussion). Therefore, we build upon the important contribution of Apergis and Miller (2009) and other studies that entirely focus on the oil market by including non-oil commodities as part of our analysis. Since the main purpose of this study is to investigate the global commodity-national stock relationship, we first look into the world commodity prices (all items) and, thereafter, at the world metal prices.

According to our test results, world commodity prices (all items) have a predictive power on national stock market returns comparable to that of oil supply shocks (in terms of the causal patterns discovered). In particular, world commodity prices (all items) Granger-cause national stock returns in 18 (17) economies, according to the LF (MF) approach.

The results of the full sample LF approach suggest that world commodity prices (all items) have an impact on stock market returns for 13 developing countries (and regions), i.e. Argentina, Chile, Iran, Mauritius, Morocco, Namibia, Nigeria, Philippines, Russia, South Africa, Sri Lanka, West African Economic and Monetary Union and Zambia, and for five developed countries (and regions), i.e. the Euro Zone, Israel, Norway, Spain and Switzerland. Consequently, we can conclude that world commodity prices (all items) Granger-cause stock market returns predominantly in developing economies, especially in Africa, where the economies are still heavily dependent on commodities – Nigeria and Zambia for example.

Additionally, this finding highlights that oil is not the only commodity playing a role in the national stock markets. The LF approach discovers causal patterns from world commodity prices (all items) to stock market returns for seven countries (namely Chile, Mauritius, Morocco, Nigeria, Philippines, Russia and Sri Lanka), for which no evidence of causality is found in terms of the oil shock proxies (see table 4.2).

Moreover, this outcome contributes to earlier works of Driesprong et al. (2008), Apergis and Miller (2009) and Wang et al. (2013), all of whom have focussed on the oil market, neglecting the effect of world commodity prices in general. Consistent with Apergis and Miller (2009), the MF test results do not conclude in favour of causality between oil and stock markets in the case of Canada. However, we identify that world commodity prices (all items) Granger-cause stock market returns for Canada. This finding is somewhat consistent with the study of Baur and McDermott (2010) who find evidence for a relationship existing between gold and stock markets in the case of Canada.

Similar to Wang et al. (2013), we are unable to identify causal patterns from oil to stock markets in the cases of India and Canada. However, we find that the world commodity prices (all items) lead stock market returns in Canada and India – a result obtained from both the MF and the LF tests. Additionally, in line with Driesprong et al. (2008), we are unable to find causality from world oil prices to national stock returns in the case of Hong Kong. In fact, our study cannot discover causality irrespective of whichever proxies for global oil shocks is considered. However, the full sample MF approach reveals that causality exists from world commodity prices (all items) to stock market returns in terms of Hong Kong. Based on these findings, we can conclude that the world commodity prices (all items) have a predictive power on stock market returns in cases where the oil prices fail irrespective of the empirical approach that is used (the LF or the MF).

Furthermore, the MF approach is able to capture causal patterns that are missed by the LF method. According to the MF approach, the world commodity prices (all items) have a predictive power on stock market returns in ten countries, for which causality remains unrevealed by the LF tests. These countries are Australia, Bangladesh, Canada, France, Germany, Hong Kong, India, Italy, New Zealand and Uganda. Therefore, our finding provides further empirical support to the study of Ghysels (2016) who claims that the MF approach is able to capture causal links that are missed by the standard LF methods.

Last but not least, we examine whether the information contained in metal prices can be used to predict national stock market returns. We find that world metal prices cause national stock returns in 30 (19) countries and regions, according to the LF (MF) approach. This finding leads to the following conclusions. First, the stock markets in developing economies are the most affected by world metal prices, as two-thirds of the causality cases from world metal prices to stock market returns are found there. Second, the stock markets in Europe and Africa are those that are primarily affected by the movements of world metal prices, according to both the LF and the MF approaches. Therefore, this finding suggests that world metal prices have a strong impact on European stock markets regardless of how developed the financial sectors of these countries are. Third, both methods (the LF and the MF) fail to identify causal patterns from world metal prices to stock market returns in terms of North American and East Asian stock markets. This might be an indication that the stock markets in these regions are well protected from the fluctuations of world metal prices. However, Hood and Malik (2013) highlight that a potential instability in the coefficient of commodity prices may alter the possibility of the full sample tests to detect a true causal relationship between global commodity markets and national financial markets. Therefore, we acknowledge that the relationship between global shocks and stock market returns may vary over time and the next section investigates this by adopting time-varying models.

4.6.2 Time-varying Granger causality tests

This section implements the time-varying MF approach together with the standard time-varying LF method. These two methods are used to discover relationships between commodity and stock markets in the presence of parameter instability. We extend our study to subsample causality analysis because national stock prices are known to be sensitive to political and economic shocks, especially in emerging economies (Bekaert and Harvey, 1995). In our time-varying analysis, the lag number is selected using the BIC, as done by Kuzin et al. (2011) and Bai et al. (2013). The BIC is a preferred method for lag selection in the case of the time-varying approach, as the inclusion of redundant lags has a large adverse impact on the asymptotic power, as highlighted by Ghysels et al. (2016).

Tables 4.3–4.7 report the rejection frequencies (at different significant levels) for rolling window MF and LF Granger causality tests.[109] The null hypothesis of non-causality is specified for each rolling window. The rejection frequency for a single country is calculated as the total number of p-values within a 5% (or 10%) significance level is divided by the total number of rolling window tests. For example, the rejection frequency for testing the null hypothesis $H_0: CP \nRightarrow SP$ ($\nRightarrow$ means "does not Granger-cause") that world oil supply shocks do not Granger-cause stock market returns in terms of the time-varying MF model for Egypt is found to be 0.083 at a 10% level of significance. This implies that the null hypothesis is rejected for 8.3% of all rolling window MF Granger causality tests for Egypt at a 10% level of significance. Therefore, we can conclude that world oil prices have predictive power on Egyptian stock market returns, which is subject to time-variability.

[109] The rolling window size is equal to half of the sample size. This is consistent with the study of Chen et al. (2010).

	World oil prices		Oil supply shocks		Oil demand shocks		World commodity prices (all items)		World metal prices	
	Significance Level		Significance Level		Significance Level		Significance Level		Significance Level	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Africa										
Panel A: Northern Africa										
Egypt	0.008	0.083	0.909	1.000	0.430	0.562	0.050	0.116	0.066	0.107
Morocco	0	0	0.355	0.432	0	0	0.672	0.749	0	0
Tunisia	0	0.081	0	0.057	0.024	0.057	0.439	0.553	0.057	0.130
Panel B: Sub-Saharan Africa										
Kenya	0.271	0.329	0.041	0.141	0.276	0.359	0	0	0	0
Malawi	0	0.034	0	0	0	0	0	0.017	0	0
Mauritius	0.012	0.029	0.179	0.231	0.046	0.116	0	0	0	0
Uganda	0.072	0.217	0	0	0	0	0.530	0.627	0	0
Tanzania	0	0	0	0	0.188	0.319	0	0	0.014	0.058
Zambia	0.047	0.063	0	0.016	0.039	0.156	0.375	0.578	0.281	0.313
Namibia	0.045	0.182	0.170	0.273	0.193	0.330	0	0.011	0	0
South Africa	0.267	0.405	0.996	1.000	0.005	0.025	0	0	0.033	0.063
Ghana	0	0	0	0	0	0	0	0	0	0.023
Nigeria	0	0.020	0.059	0.098	0	0	0.039	0.137	0	0.059
West African Economic and Monetary Union	0.008	0.127	0.907	0.932	0.136	0.212	0.008	0.008	0.051	0.085
Americas										
Panel C: Latin America and the Caribbean										
Mexico	0	0	0.004	0.031	0.085	0.171	0.054	0.147	0.004	0.089
Argentina	0	0	0.007	0.007	0	0.013	0	0	0	0
Brazil	0.154	0.256	0.280	0.436	0.101	0.156	0	0	0	0.014
Chile	0	0	0.216	0.304	0.014	0.047	0.412	0.480	0.520	0.642
Colombia	0	0	0.631	0.667	0	0.046	0	0	0	0
Panel D: Northern America										
Canada	0	0.026	0	0	0	0	0	0.006	0.125	0.243
United States of America (the US)	0	0.041	0.004	0.037	0.241	0.317	0	0.002	0.012	0.101
Asia										
Panel E: Central Asia										
Kazakhstan	0.103	0.495	0	0.037	0.065	0.150	0	0.037	0.318	0.449
Panel F: Eastern Asia										
China (Mainland)	0	0	0.031	0.070	0.086	0.188	0.016	0.039	0	0.063
Hong Kong	0	0.021	0	0	0.055	0.136	0.031	0.090	0	0.003
Japan	0	0.010	0	0	0	0	0.201	0.323	0.003	0.041
South Korea	0	0	0.559	0.693	0.131	0.246	0.011	0.057	0	0
Taiwan	0	0.010	0.010	0.075	0.005	0.050	0	0	0.090	0.149
Panel G: South-eastern Asia										
Indonesia	0	0	0.023	0.146	0.006	0.029	0.070	0.146	0	0
Malaysia	0.005	0.021	0.017	0.083	0.060	0.206	0.139	0.400	0.004	0.078
Philippines	0	0	0	0.013	0	0	0	0.013	0.063	0.113
Thailand	0	0	0.116	0.212	0	0.005	0.004	0.097	0.012	0.031
Panel H: Southern Asia										
Bangladesh	0.207	0.396	0.426	0.538	0	0.065	0	0	0	0.006

India	0	0	0.005	0.022	0.082	0.158	0	0	0	0.005
Iran	0.239	0.269	0.254	0.313	0	0.015	0.134	0.358	0.194	0.299
Sri Lanka	0	0	0.425	0.525	0.085	0.141	0.070	0.180	0.070	0.170
Panel I: Western Asia										
Israel	0.229	0.357	0.013	0.076	0.076	0.229	0	0	0.032	0.064
Saudi Arabia	0.008	0.074	0.172	0.402	0.008	0.049	0	0.025	0.008	0.090
Turkey	0	0	0.104	0.352	0.209	0.401	0.132	0.143	0	0
						Europe				
Panel J: Eastern Europe										
Czech Republic	0	0.007	0.041	0.189	0.088	0.209	0	0.014	0	0
Hungary	0	0	0.049	0.268	0.140	0.274	0.037	0.171	0.177	0.329
Poland	0.092	0.123	0	0.074	0.270	0.374	0.061	0.080	0	0
Russia	0.185	0.250	0	0.016	0.040	0.250	0	0	0	0
Slovakia	0.107	0.349	0	0.040	0	0	0	0	0	0
Ukraine	0.460	0.548	0.008	0.032	0.105	0.444	0	0	0	0
Panel K: Northern Europe										
Finland	0.087	0.210	0	0	0.090	0.171	0	0.019	0	0
Iceland	0.013	0.111	0	0	0	0.007	0.320	0.523	0	0
Ireland	0.005	0.041	0	0.004	0.005	0.040	0.033	0.094	0	0
Lithuania	0	0	0	0	0.018	0.018	0.054	0.243	0.297	0.369
Norway	0.141	0.200	1.000	1.000	0.015	0.052	0.311	0.481	0.207	0.415
United Kingdom	0.072	0.159	0	0	0.111	0.196	0	0.003	0	0.006
Panel L: Southern Europe										
Croatia	0.375	0.414	0.016	0.148	0.039	0.047	0.688	0.813	0.555	0.633
Greece	0.149	0.314	0.005	0.020	0	0.056	0.085	0.291	0.025	0.101
Italy	0.082	0.210	0.007	0.059	0.136	0.201	0.017	0.101	0.007	0.132
Portugal	0	0	0.126	0.264	0.033	0.049	0.005	0.038	0	0.005
Spain	0.015	0.118	0	0	0.131	0.231	0.003	0.030	0	0.008
Panel M: Western Europe										
Belgium	0	0	0	0	0.025	0.055	0.004	0.048	0	0
France	0.077	0.159	0.304	0.392	0	0.020	0	0	0	0.017
Germany	0	0.005	0	0	0.181	0.191	0	0	0	0.020
Netherlands	0	0.015	0	0.004	0.070	0.101	0	0.006	0	0.031
Switzerland	0	0.006	0.006	0.066	0.138	0.254	0.260	0.425	0.050	0.166
Panel N: Europe										
Euro Zone	0	0	0.993	1.000	0.010	0.070	0	0.007	0.445	0.739
						Oceania				
Panel O: Australia and New Zealand										
Australia	0	0.026	0.139	0.187	0.055	0.126	0.004	0.032	0	0
New Zealand	0	0	0.083	0.255	0.214	0.302	0.208	0.380	0.469	0.688

Note: The table shows the rejection frequencies at different significant levels for rolling window MF Granger causality tests of non-causality from global shocks to stock market returns. "SP" denotes the stock market returns, while "CP" denotes the global variables. $H_0: CP \nRightarrow SP$ ($\nRightarrow$ means "does not Granger-cause"). We follow Ghysels et al. (2016) and use bootstrapped p-values with N = 499 replications (Gonçalves and Kilian 2004). All variables are mean-centred and log-differenced, as specified in Section 4.5.

Table 4.3 Rejection Frequencies at Different Significant Levels for Rolling Window Mixed-Frequency Granger Causality Tests of Non-Causality from Global Shocks to Stock Market Returns

	World oil prices		Oil supply shocks		Oil demand shocks		World commodity prices (all items)		World metal prices	
	Significance Level		Significance Level		Significance Level		Significance Level		Significance Level	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Africa										
Panel A: Northern Africa										
Egypt	0.058	0.099	0.058	0.174	0	0.008	0	0	0.008	0.058
Morocco	0	0	0.022	0.142	0	0	0.044	0.262	0	0
Tunisia	0	0	0.114	0.195	0.057	0.244	0.008	0.008	0.163	0.252
Panel B: Sub-Saharan Africa										
Kenya	0.218	0.276	0.041	0.241	0	0.035	0	0	0	0.006
Malawi	0	0	0	0	0.034	0.034	0.034	0.052	0.034	0.034
Mauritius	0.023	0.075	0	0.006	0.254	0.439	0	0	0	0
Uganda	0.193	0.422	0	0.012	0	0.012	0.072	0.145	0	0
Tanzania	0	0	0	0	0	0	0	0.130	0	0.058
Zambia	0.047	0.234	0.039	0.234	0.070	0.375	0.922	0.953	0.945	1.000
Namibia	0.375	0.648	0.375	0.500	0.102	0.239	0	0.068	0	0.011
South Africa	0.138	0.482	1.000	1.000	0	0.025	0	0	0.129	0.158
Ghana	0	0	0	0	0	0.114	0	0.023	0.045	0.091
Nigeria	0	0	0.020	0.314	0	0.078	0.118	0.216	0.137	0.255
West African Economic and Monetary Union	0.297	0.763	0.941	0.941	0.017	0.034	0.017	0.161	0.059	0.093
Americas										
Panel C: Latin America and the Caribbean										
Mexico	0	0.005	0.004	0.036	0.317	0.492	0	0.009	0	0.004
Argentina	0	0	0	0	0.060	0.309	0	0	0.007	0.013
Brazil	0	0.005	0	0	0.241	0.437	0	0	0.014	0.124
Chile	0	0.041	0	0	0	0	0.047	0.081	0.034	0.230
Colombia	0	0	0.215	0.508	0	0	0.010	0.062	0	0.015
Panel D: Northern America										
Canada	0	0	0	0	0.010	0.050	0	0	0.150	0.274
United States of America (the US)	0.021	0.103	0	0	0.312	0.397	0.136	0.358	0.452	0.652
Asia										
Panel E: Central Asia										
Kazakhstan	0	0.019	0.009	0.037	0	0.009	0	0	0	0.037
Panel F: Eastern Asia										
China (Mainland)	0	0.008	0	0	0	0	0	0.023	0.148	0.578
Hong Kong	0	0	0	0	0.060	0.201	0	0	0.012	0.150
Japan	0	0	0	0	0	0	0	0	0	0
South Korea	0	0	0	0	0.307	0.508	0.077	0.134	0.077	0.172
Taiwan	0	0.010	0.005	0.050	0.065	0.156	0	0.005	0	0.070
Panel G: South-eastern Asia										
Indonesia	0.006	0.070	0.006	0.070	0.053	0.181	0	0.012	0	0
Malaysia	0	0.031	0	0	0	0	0	0	0	0.009
Philippines	0	0	0	0	0	0.075	0	0	0.225	0.350
Thailand	0	0	0	0	0	0.065	0.564	0.764	0	0
Panel H: Southern Asia										
Bangladesh	0.414	0.527	0.420	0.456	0.036	0.071	0	0	0.053	0.349

India	0.087	0.268	0	0	0.158	0.268	0	0	0.005	0.027
Iran	0.328	0.537	0.254	0.254	0	0	0.463	0.582	0.343	0.567
Sri Lanka	0	0	0.260	0.450	0.035	0.095	0.265	0.430	0.105	0.300
Panel I: Western Asia										
Israel	0.102	0.166	0.159	0.204	0.414	0.694	0.019	0.064	0	0.064
Saudi Arabia	0.426	0.680	0.213	0.508	0	0	0.049	0.057	0.123	0.287
Turkey	0	0	0	0	0.033	0.121	0.005	0.033	0	0
						Europe				
Panel J: Eastern Europe										
Czech Republic	0.054	0.149	0.007	0.027	0.020	0.108	0	0.007	0	0
Hungary	0	0.012	0.024	0.159	0	0	0.012	0.189	0.396	0.500
Poland	0.006	0.135	0	0	0.006	0.092	0.043	0.080	0.012	0.037
Russia	0.008	0.032	0	0	0.065	0.089	0.065	0.089	0	0.024
Slovakia	0.007	0.047	0	0	0	0	0	0	0	0
Ukraine	0	0	0.065	0.089	0	0	0.024	0.073	0	0
Panel K: Northern Europe										
Finland	0.128	0.262	0.011	0.062	0.281	0.417	0	0.096	0.190	0.273
Iceland	0.039	0.105	0	0	0	0	0.150	0.601	0	0
Ireland	0.169	0.297	0	0	0.010	0.060	0	0.022	0	0
Lithuania	0	0	0	0	0.009	0.099	0.018	0.207	0	0
Norway	0	0	1.000	1.000	0.052	0.089	0.422	0.630	0.593	0.770
United Kingdom	0.021	0.138	0	0	0.161	0.261	0.003	0.053	0.178	0.445
Panel L: Southern Europe										
Croatia	0.023	0.102	0	0.172	0.047	0.133	0.219	0.398	0.273	0.375
Greece	0.124	0.232	0	0	0	0	0.186	0.307	0.126	0.417
Italy	0.379	0.426	0	0.011	0.121	0.216	0.270	0.382	0.003	0.074
Portugal	0	0	0	0	0	0.016	0.187	0.258	0	0.060
Spain	0.246	0.333	0.051	0.161	0.201	0.216	0	0.058	0	0
Panel M: Western Europe										
Belgium	0	0.036	0	0	0.090	0.186	0	0	0.043	0.139
France	0	0.005	0.286	0.366	0.005	0.015	0.003	0.050	0	0
Germany	0.010	0.072	0	0.011	0.236	0.281	0	0.031	0.151	0.196
Netherlands	0.087	0.287	0	0	0.181	0.302	0	0	0.283	0.336
Switzerland	0.149	0.188	0.392	0.586	0.249	0.381	0	0	0	0
Panel N: Europe										
Euro Zone	0	0	0.445	0.596	0.010	0.040	0	0.004	0.011	0.029
						Oceania				
Panel O: Australia and New Zealand										
Australia	0.005	0.092	0.004	0.070	0.231	0.332	0	0	0	0
New Zealand	0	0	0	0	0.240	0.344	0.250	0.401	0.583	0.599

Note: The table shows the rejection frequencies at different significant levels for rolling window LF Granger causality tests of non-causality from global shocks to stock market returns. "SP" denotes the stock market returns, while "CP" denotes the global variables. $H_0: CP \nRightarrow SP$ ($\nRightarrow$ means "does not Granger-cause"). We follow Ghysels et al. (2016) and use bootstrapped p-values with N = 499 replications (Gonçalves and Kilian 2004). All variables are mean-centred and log-differenced, as specified in Section 4.5.

Table 4.4 Rejection Frequencies at Different Significant Levels for Rolling Window Low-Frequency Granger Causality Tests of Non-Causality from Global Shocks to Stock Market Returns

	World oil prices		Oil supply shocks		Oil demand shocks		World commodity prices (all items)		World metal prices	
	Significance Level		Significance Level		Significance Level		Significance Level		Significance Level	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Africa										
Panel A: Northern Africa										
Egypt	0	0	0	0	0.240	0.364	0	0	0.008	0.190
Morocco	0	0	0	0	0.022	0.131	0.191	0.377	0	0
Tunisia	0	0.008	0	0	0.065	0.081	0	0	0	0
Panel B: Sub-Saharan Africa										
Kenya	0	0	0	0	0	0	0	0	0	0
Malawi	0.672	0.793	0.017	0.086	0.224	0.224	0	0	0.190	0.517
Mauritius	0.081	0.116	0	0	0.422	0.578	0	0	0	0
Uganda	0	0	0	0.012	0.241	0.373	0.169	0.530	0.181	0.289
Tanzania	0	0	0.058	0.101	0	0	0.029	0.101	0.159	0.275
Zambia	0.164	0.336	0.039	0.141	0	0	0.055	0.078	0.016	0.047
Namibia	0.034	0.068	0.011	0.011	0	0	0	0.034	0	0
South Africa	0	0.026	0.029	0.158	0.065	0.176	0	0	0	0
Ghana	0	0	0	0	0	0	0	0.045	0	0
Nigeria	0	0	0	0.020	0	0	0.039	0.157	0	0.020
West African Economic and Monetary Union	0.017	0.169	0.585	0.720	0	0.085	0	0	0	0
Americas										
Panel C: Latin America and the Caribbean										
Mexico	0	0	0	0.036	0	0.005	0	0.067	0	0.004
Argentina	0.087	0.228	0.027	0.027	0.309	0.463	0.013	0.013	0.013	0.067
Brazil	0	0	0	0	0.171	0.327	0.289	0.514	0	0.037
Chile	0.027	0.088	0	0	0	0.054	0	0	0	0.007
Colombia	0	0	0	0	0	0	0.005	0.021	0	0
Panel D: Northern America										
Canada	0	0	0.059	0.201	0.015	0.035	0.377	0.570	0	0.003
United States of America (the US)	0	0.005	0	0.007	0	0.015	0	0.010	0.121	0.205
Asia										
Panel E: Central Asia										
Kazakhstan	0	0.019	0.140	0.280	0.355	0.355	0.009	0.037	0	0.009
Panel F: Eastern Asia										
China (Mainland)	0	0	0	0	0	0.023	0	0.008	0	0
Hong Kong	0	0	0	0.007	0.010	0.060	0	0.003	0	0
Japan	0	0	0	0	0.020	0.050	0	0.003	0.019	0.188
South Korea	0	0	0.015	0.061	0.236	0.357	0	0.004	0.027	0.103
Taiwan	0	0	0	0	0	0.050	0	0	0.035	0.149
Panel G: South-eastern Asia										
Indonesia	0	0	0	0.012	0	0.006	0.175	0.345	0.193	0.485
Malaysia	0	0	0	0	0.080	0.151	0	0.026	0	0
Philippines	0	0	0	0	0	0	0	0	0	0.038
Thailand	0	0	0	0.023	0.040	0.085	0	0	0	0
Panel H: Southern Asia										
Bangladesh	0	0	0	0	0.107	0.361	0.107	0.189	0.296	0.414

India	0.005	0.011	0.005	0.022	0	0.060	0.016	0.077	0.071	0.169
Iran	0.552	0.612	0	0.090	0	0	0.015	0.134	0	0.015
Sri Lanka	0	0	0	0.020	0.060	0.256	0	0.030	0.310	0.600
Panel I: Western Asia										
Israel	0	0	0	0	0	0	0.096	0.369	0.019	0.217
Saudi Arabia	0.082	0.123	0	0	0	0.025	0	0	0	0.074
Turkey	0	0	0	0	0.044	0.170	0.005	0.049	0	0
						Europe				
Panel J: Eastern Europe										
Czech Republic	0	0	0.223	0.318	0.250	0.358	0	0	0.007	0.115
Hungary	0	0.055	0	0.043	0.055	0.226	0	0.024	0	0
Poland	0.018	0.061	0.098	0.166	0.209	0.405	0	0	0	0.025
Russia	0.016	0.065	0.008	0.032	0.008	0.056	0	0.016	0	0.032
Slovakia	0.148	0.383	0.054	0.342	0.181	0.336	0	0.013	0	0.007
Ukraine	0.073	0.226	0	0.089	0.210	0.702	0	0	0.105	0.210
Panel K: Northern Europe										
Finland	0.005	0.036	0.267	0.293	0.010	0.116	0	0	0	0
Iceland	0.739	0.745	0	0	0.007	0.033	0.039	0.092	0.516	0.784
Ireland	0	0	0	0	0.010	0.181	0	0.011	0.749	0.901
Lithuania	0.514	0.784	0	0.009	0	0	0.009	0.027	0	0
Norway	0.252	0.444	0	0	0.044	0.104	0	0	0	0
United Kingdom	0	0	0	0.007	0.050	0.126	0	0	0	0
Panel L: Southern Europe										
Croatia	0	0.039	0	0	0	0	0	0	0	0
Greece	0	0	0.070	0.141	0.040	0.061	0	0.020	0	0.020
Italy	0	0.072	0	0.022	0.045	0.176	0.030	0.159	0	0
Portugal	0	0	0	0.033	0.390	0.665	0	0.011	0.038	0.099
Spain	0	0.010	0	0	0.211	0.427	0	0.017	0.666	0.738
Panel M: Western Europe										
Belgium	0	0.056	0	0	0.337	0.377	0	0.022	0	0
France	0	0.021	0	0.011	0.271	0.377	0	0.006	0.028	0.204
Germany	0	0	0.110	0.201	0.005	0.085	0	0	0	0.003
Netherlands	0	0	0	0	0.075	0.126	0.564	0.607	0.044	0.212
Switzerland	0	0.011	0.204	0.276	0.088	0.155	0	0.017	0.022	0.210
Panel N: Europe										
Euro Zone	0.108	0.205	0.195	0.434	0.417	0.688	0	0	0.004	0.040
						Oceania				
Panel O: Australia and New Zealand										
Australia	0	0	0.004	0.062	0.030	0.070	0	0.004	0	0
New Zealand	0	0	0.005	0.026	0	0	0.068	0.130	0	0

Note: The table shows the rejection frequencies at different significant levels for rolling window MF Granger causality tests of non-causality from stock market returns to global shocks. "SP" denotes the stock market returns, while "CP" denotes the global variables. $H_0: SP \nRightarrow CP$ ($\nRightarrow$ means "does not Granger-cause"). We follow Ghysels et al. (2016) and use bootstrapped p-values with N = 499 replications (Gonçalves and Kilian 2004). All variables are mean-centred and log-differenced, as specified in Section 4.5.

Table 4.5 Rejection Frequencies at Different Significant Levels for Rolling Window Mixed-Frequency Granger Causality Tests of Non-Causality from Stock Market Returns to Global Shocks

	World oil prices		Oil supply shocks		Oil demand shocks		World commodity prices (all items)		World metal prices	
	Significance Level		Significance Level		Significance Level		Significance Level		Significance Level	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Africa										
Panel A: Northern Africa										
Egypt	0.066	0.331	0.446	0.545	0.752	0.860	0.008	0.149	0.041	0.058
Morocco	0	0.011	0	0.011	0.011	0.131	0.060	0.279	0.022	0.257
Tunisia	0.081	0.195	0	0	0.016	0.089	0	0	0	0
Panel B: Sub-Saharan Africa										
Kenya	0	0	0	0	0	0	0	0	0	0
Malawi	0	0.103	0.103	0.190	0.224	0.224	0	0.034	0.241	0.310
Mauritius	0.272	0.497	0	0	0.757	0.861	0	0	0	0
Uganda	0.012	0.048	0	0	0	0	0	0	0	0
Tanzania	0	0	0	0.029	0.014	0.130	0.174	0.232	0	0.014
Zambia	0.680	0.727	0.031	0.055	0	0	0.117	0.188	0.070	0.320
Namibia	0	0	0	0.011	0	0.023	0	0	0	0
South Africa	0.133	0.205	0.022	0.099	0.347	0.462	0	0	0	0
Ghana	0	0	0.023	0.023	0.227	0.250	0.023	0.068	0	0
Nigeria	0	0	0	0	0	0	0.020	0.098	0	0.039
West African Economic and Monetary Union	0.127	0.483	0	0	0.093	0.364	0	0	0	0.008
Americas										
Panel C: Latin America and the Caribbean										
Mexico	0.005	0.021	0	0.004	0	0.075	0.031	0.058	0.152	0.379
Argentina	0.054	0.235	0.020	0.020	0.597	0.638	0.007	0.020	0.181	0.208
Brazil	0.026	0.113	0	0.005	0.201	0.337	0.890	0.940	0.018	0.174
Chile	0.324	0.534	0	0	0.588	0.608	0.088	0.101	0.115	0.243
Colombia	0	0.051	0.128	0.236	0.226	0.415	0	0	0.062	0.272
Panel D: Northern America										
Canada	0.005	0.046	0	0	0.286	0.467	0.520	0.642	0.072	0.221
United States of America (the US)	0	0	0	0.015	0	0	0	0	0.005	0.027
Asia										
Panel E: Central Asia										
Kazakhstan	0.346	0.598	0.364	0.794	0.047	0.056	0.009	0.019	0	0.047
Panel F: Eastern Asia										
China (Mainland)	0	0	0	0	0.375	0.477	0	0	0.008	0.016
Hong Kong	0	0	0	0.015	0.050	0.327	0.037	0.081	0	0
Japan	0	0	0	0.040	0.186	0.387	0	0	0	0
South Korea	0	0.031	0.015	0.027	0.251	0.452	0	0.015	0.015	0.119
Taiwan	0.051	0.256	0.005	0.075	0.010	0.035	0	0	0.478	0.622
Panel G: South-eastern Asia										
Indonesia	0.035	0.152	0.006	0.076	0	0.023	0.304	0.503	0.702	0.784
Malaysia	0.200	0.559	0	0	0	0	0	0	0	0.013
Philippines	0	0	0	0	0	0	0	0	0.013	0.050
Thailand	0	0	0	0.023	0.151	0.276	0.004	0.008	0	0
Panel H: Southern Asia										
Bangladesh	0	0	0.124	0.207	0.249	0.320	0.154	0.207	0.148	0.183

India	0	0	0	0.044	0.208	0.306	0.005	0.005	0	0.005
Iran	0	0.030	0.164	0.418	0.075	0.194	0	0.015	0	0
Sri Lanka	0	0	0	0	0.487	0.543	0.420	0.450	0.305	0.435
Panel I: Western Asia										
Israel	0.006	0.102	0.013	0.051	0	0.032	0.121	0.376	0	0
Saudi Arabia	0.123	0.180	0.008	0.025	0.008	0.025	0	0	0	0
Turkey	0.005	0.126	0.038	0.121	0.187	0.462	0	0	0.071	0.192
						Europe				
Panel J: Eastern Europe										
Czech Republic	0	0	0.088	0.264	0.074	0.149	0	0	0.074	0.203
Hungary	0	0.006	0	0.024	0.128	0.348	0	0	0	0
Poland	0.153	0.184	0	0.006	0.178	0.331	0	0	0	0.025
Russia	0.613	0.815	0.145	0.250	0.266	0.573	0.137	0.258	0.040	0.194
Slovakia	0	0	0	0	0.094	0.376	0	0	0	0
Ukraine	0.145	0.774	0.008	0.065	0.685	0.831	0.508	0.847	0.387	0.927
Panel K: Northern Europe										
Finland	0	0	0.165	0.271	0.085	0.427	0	0.008	0	0
Iceland	0.386	0.523	0.007	0.013	0.209	0.261	0.255	0.425	0.255	0.575
Ireland	0	0.041	0	0	0.427	0.503	0	0	0.854	0.917
Lithuania	0	0.009	0	0	0.045	0.189	0	0.018	0	0
Norway	0.837	0.852	0.156	0.548	0.600	0.711	0.022	0.052	0.037	0.067
United Kingdom	0	0	0	0.007	0.015	0.246	0	0	0	0
Panel L: Southern Europe										
Croatia	0.016	0.086	0	0	0.055	0.258	0	0	0.008	0.023
Greece	0.077	0.196	0	0.015	0.369	0.475	0.020	0.085	0	0.020
Italy	0.021	0.113	0	0	0.427	0.623	0	0	0.017	0.118
Portugal	0.165	0.253	0	0.104	0.412	0.577	0	0	0	0
Spain	0	0.026	0	0	0.372	0.462	0.072	0.141	0.746	0.785
Panel M: Western Europe										
Belgium	0.031	0.128	0	0	0.010	0.131	0	0	0.048	0.199
France	0	0.015	0.026	0.136	0.412	0.472	0.003	0.050	0.309	0.627
Germany	0.113	0.190	0.187	0.242	0	0.005	0.281	0.594	0	0.054
Netherlands	0	0.010	0	0	0	0	0.445	0.558	0.106	0.252
Switzerland	0	0	0.022	0.127	0.099	0.337	0	0.011	0	0.022
Panel N: Europe										
Euro Zone	0	0	0.301	0.404	0.508	0.548	0	0.011	0	0
						Oceania				
Panel O: Australia and New Zealand										
Australia	0	0.015	0	0.004	0.040	0.116	0	0.007	0	0
New Zealand	0	0	0.005	0.063	0	0	0	0	0	0

Note: The table shows the rejection frequencies at different significant levels for rolling window LF Granger causality tests of non-causality from stock market returns to global shocks. "SP" denotes the stock market returns, while "CP" denotes the global variables. $H_0: SP \nRightarrow CP$ ($\nRightarrow$ means "does not Granger-cause"). We follow Ghysels et al. (2016) and use bootstrapped p-values with N = 499 replications (Gonçalves and Kilian 2004). All variables are mean-centred and log-differenced, as specified in Section 4.5.

Table 4.6 Rejection Frequencies at Different Significant Levels for Rolling Window Low-Frequency Granger Causality Tests of Non-Causality from Stock Market Returns to Global Shocks

Tables 4.3 and 4.4 show that, compared to the full sample results, the rolling Granger causality tests provide more evidence in favour of causality from global shocks to stock market returns for both the MF and LF models. We consider rejections that occur at the 5% and 10% levels and find that the results of the two sets of tests suggest rather different conclusions.

The MF results in table 4.3 reveal that world oil prices Granger-cause stock market returns for 42 out of 63 countries (and regions) at a 10% level of significance. This represents 67% of all countries (and regions) in our sample. In particular, the time-varying MF Granger causality test cannot reject the null hypothesis that world oil prices do not Granger-cause stock market returns for all Latin American and the Caribbean and South-eastern Asian countries apart from Brazil and Malaysia. In comparison, the rolling window LF Granger causality tests discover that world oil prices Granger-cause stock market returns for 65% of all countries at a 10% level of significance (see table 4.4). Therefore, the cases of causality that are identified by the time-varying LF method are 2% less than those found by the time-varying MF method. In addition, this study adds to the important contribution of Apergis and Miller (2009), who claim that there is no causality from world oil prices to national stock markets for Canada and Japan. This result is consistent with the findings of our time-varying LF approach. However, the results from the time-varying MF approach reveal evidence of temporal causality from world oil prices to stock market returns for Canada and Japan. This outcome suggests that the MF approach can capture causal patterns where the LF approach fails – a finding that provides further empirical support to the recent studies of Ghysels (2016) and Ghysels et al. (2016).

Moreover, the full sample MF tests find that world oil prices Granger-cause stock market returns for only three out of 63 countries (and regions). The weak evidence of causality may lead to misleading conclusions regarding the crucial role that commodity markets play in national stock markets. A possible reason for the difference between the full sample and time-varying methods is the inability of the former to account for potential structural breaks in the commodity-stock relationship. This once again confirms the advantages of using the time-varying models when analysing the commodity prices, which are known to be highly volatile, especially during financial crises (e.g. the Global Financial Crisis).

Furthermore, we acknowledge that the reaction of the stock market returns to an oil price shock differs greatly depending on whether the change in the price of oil is driven by demand or supply shocks in the oil market, as highlighted by Kilian and Park (2009). In particular, the MF results in table 4.3 reveal substantial causality from oil supply shocks to stock market returns. Particularly, the results from the time-varying MF tests suggest that world oil supply shocks Granger-cause stock market returns for 78% of all countries at a 10% level of significance. Another way of interpretation is that the time-varying MF Granger causality tests reject the null $H_0: CP \nRightarrow SP$ for all countries in Central Asia, Eastern Europe, the Euro Zone, Latin America and the Caribbean, Northern Africa, Oceania, South-eastern Asia, Southern Asia and Western Asia. Interestingly, the world oil supply shocks have a "persistent" impact on stock markets in Egypt, the Euro Zone, Norway and South Africa, regardless of the time period. Except for Norway, all other countries are classified as oil-importers.[110] Therefore, we conclude that oil supply shocks have a greater influence on oil-importing countries than oil-exporting ones. This evidence is consistent with the study of Park and Ratti (2008).

Nonetheless, the world oil demand shocks are found to be a slightly better predictor for the movements of national stock returns than the world oil supply shocks. The results from the time-varying MF tests suggest that world oil demand shocks Granger-cause stock market returns for 86% of all countries at a 10% level of significance. In particular, the world economic activity has a causal effect on national stock market returns for all Oceania countries, all European countries apart from Slovakia, all Americas countries apart from Canada, and all Asian countries apart from Japan and Philippines. In contrast, the world economic activity has a less pronounced effect on national stock markets in Africa. In fact, world economic activity does not Granger-cause stock market returns in the following African countries: Ghana, Malawi, Morocco, Nigeria and Uganda. All of these countries are classified as developing economies that are commodity export-dependent. The only exception is Morocco, which is an importer of primary commodity products such as oil. Equally important is the finding of the time-varying MF tests that world oil supply shocks Granger-cause stock market returns for only 78% of all countries at a 10% level of significance. This indicates that the world oil demand shocks have a greater effect on stock market returns than the world oil supply shock.

It is important to emphasise the fact that the time-varying MF approach reveals more cases of causality than the time-varying LF approach when causality runs from oil supply (demand) shocks to stock market returns. In particular, the rolling window LF Granger causality tests identify that world oil supply shocks Granger-cause stock market returns for only 54% of all countries at a 10% level of significance (see table 4.4). In fact, the number of countries (and

[110] Most of the countries in the Euro Zone are oil-importers; therefore, we assume that the Euro Zone is an oil-importing region.

regions) found by the time-varying LF method are about 24% less than those found by the time-varying MF method. Similarly, the rolling window LF Granger causality tests find that world oil demand shocks Granger-cause stock market returns for only 78% of all countries at a 10% level of significance, whereas the time-varying MF approach suggests this for 86% of all countries at a 10% level of significance (see tables 4.3 and 4.4). Thus, we can conclude that the results from the time-varying LF tests provide less support for the existence of a link between oil and stock markets than their MF counterparts.

Another key point is that some stock markets might be less dependent on oil than on other commodities (Hood and Malik, 2013). Therefore, this study investigates the impact of both world metal prices and world commodity prices (all items) on national stock markets. The results from the MF methods in table 4.3 show substantial causality from world metal prices to stock market returns. In fact, the time-varying MF results suggest that world metal prices Granger-cause stock market returns for 67% of all countries at a 10% level of significance (see table 4.3). In a similar manner, the results from the rolling window LF Granger causality tests indicate that world metal prices Granger-cause stock market returns for 73% of all countries at a 10% level of significance (see table 4.4). Hence, the overall conclusions that are drawn from the time-varying LF and MF methods are rather similar. In other words, both methods identify a substantial number of causal patterns from world metal prices to national stock market returns for all countries in Central Asia, the Euro Zone, Northern America and Southern Asia. In addition, the MF method finds a link between world metal prices and stock market returns for all countries in Southern Europe, whereas the LF method finds such a link for all countries in Southern Europe apart from Spain. Therefore, the differences in the overall results between the time-varying LF and MF methods, as given in tables 4.3 and 4.4, are almost negligible, at least in the case of world metal prices.

Another import proxy of global shocks that we consider in our analysis is the world commodity prices (all items). The empirical outcomes of the MF tests show that world commodity prices (all items) Granger-cause stock market returns for 73% of all countries at a 10% level of significance (see table 4.3). In fact, the time-varying MF Granger causality tests reject the null hypothesis that world commodity prices (all items) do not Granger-cause stock market returns for all countries in Central Asia, the Euro Zone, Northern Africa, Northern America, Northern Europe, Oceania and Southern Europe. Similarly, the results from the time-varying LF method confirm the existence of a solid link between commodity and financial markets (see table 4.4). In other words, the time-varying LF method reveals that world commodity prices (all items) Granger-cause stock market returns for 70% of all countries at a 10% level of significance. More precisely, the time-varying LF method finds causality for all countries in the Euro Zone, Northern Europe, Southern Europe and Western Asia. Based on the aforementioned results, we conclude that the world commodity prices (all items) exert more influence on stock market returns as compared to the oil prices. Thus, one should avoid focussing on a single commodity, such as oil or gold, especially when investigating the existence of a causal link between commodity and stock markets in terms of well-developed and functioning financial markets.

Last but not least, we also discover a solid evidence of causality from stock market returns to global shocks. In particular, the highest number of causal patterns is found in the case of world oil demand shocks. The results from the time-varying MF tests suggest that the stock market returns Granger-cause world oil demand shocks for 79% of all countries (and regions) at a 10% level of significance (see table 4.5). More precisely, the stock market returns of all countries in Central Asia, Eastern Asia, Eastern Europe, the Euro Zone, Northern Africa, Northern America and Western Europe are found to have impact on world economic activity. At the same time, less evidence is found with respect to the financial markets in the sub-Saharan African region. This result is not surprising as only few companies trade their shares on national stock markets in sub-Saharan Africa as compared to Europe (e.g. the United Kingdom), Asia (e.g. China) and North America (e.g. the US). Furthermore, a handful of causal patterns from stock market returns to global shocks are identified in the case of four other proxies for global shocks, namely world oil prices, world oil supply, world metal prices and world commodity prices (all items). The evidence for the existence of causality is less pronounced as compared to that found by world oil demand. All things considered, we conclude that stock markets in developed economies play an important role in world economic activity.

In summary, this chapter provides evidence of a relationship existing between commodity and financial markets for both developing and developed countries. It shows that oil is not the only relevant commodity that plays a role in the national financial markets. Indeed, the world metal prices and the world commodity prices (all items) also play an important role in the development of stock markets. Furthermore, this study extends the full sample MF models of Ghysels et al. (2016) to a time-varying setting that accounts for potential structural breaks in the commodity-stock relationship. Rossi (2005) highlights that the presence of structural breaks may reduce the power of full sample tests and, therefore, an existing causal link may remain hidden. The adoption of the time-varying method along with the long historical time series led to identifying a substantial number of causal patterns that may otherwise stay uncovered. Finally, the past studies have mainly looked at the relationship between non-oil commodities and stock markets in terms of correlation, while less evidence has been provided to the case of causality.[111] This study contributes to the existing literature by revealing the important role of fuel and non-fuel commodities in national stock markets irrespective of the stage of development of the financial markets.[112]

4.7 Robustness Check

This section looks at a different perspective on the connection between commodity prices and stock market returns. Here, we use a national index of commodity prices instead of a global proxy. Particularly, we use the index series of national commodity export prices that we construct in Chapter 2 to examine the relationship between (national) commodity prices and stock market returns. Whereas, national commodity prices have been frequently used in the existing commodity-exchange rate literature (see Chen and Rogoff, 2003; Chen et al., 2010; Bodart et al., 2012; 2015; Ferraro et al., 2015), evidence for a relationship between national commodity prices and stock market returns is limited.

This section aims to tackle this limitation by using the newly constructed database of national commodity price indexes (see Chapter 2). Important to highlight is that we use only same-frequency (LF) model in this section as data frequency of both variables, i.e. stock and commodity prices, is monthly. Another caveat is that national commodity price indexes are not constructed for Euro Zone and Taiwan (see Appendix A.4); therefore, they are excluded from our analysis. This is due to the unavailability of 1995-2010 period trade data for construction of index weights for Euro Zone and Taiwan and, thus, their corresponding national commodity price index series from UN Comtrade (2018). Further to that, we use national commodity price indexes for each country part of West African Economic and Monetary Union, i.e. Benin, Burkina Faso, Guinea-Bissau, Ivory Coast/Cote d'Ivoire, Mali, Niger, Senegal and Togo, as of November 2019, to examine the relationship between national commodity prices and stock market returns. Therefore, the total number of countries under investigation increases to 68.

We begin by considering the main features of the data, namely stationarity, which matters for the accuracy of the VAR model. A failure to account for possible unit roots by differencing may have serious statistical consequences, such as regressions estimated from data with unit roots can have non-stationary residuals, leading to spurious regression results (Kormendi and

[111] For example, see Apergis and Miller (2009), Kilian and Park (2009), Park and Ratti (2009), Wang et al. (2013) and Fang and You (2014) among others.
[112] We provide an additional analysis on the relationship between financial markets and national commodity prices (taken from Chapter 2) in Appendix C.3.

Meguire, 1990). A proper understanding of the extent of stationarity in the data handling is therefore essential when interpreting the results. For testing stationarity, we perform Augmented Dickey-Fuller (ADF) (1979) and Phillips and Perron (PP) (1988) unit root tests for national commodity export price series, as discussed by Narayan et al. (2014). For both tests, we specify the null hypothesis of a unit root against the alternative of stationarity. The lag length is selected by using the BIC. The results that we obtain from both tests denote that the null hypothesis is rejected in favour of stationarity for all national commodity price series. The results from the unit root tests are reported in table C.7 in Appendix C.3.

Next, we proceed with discussing the results from the empirical testing. First, the relationship between national commodity prices and financial markets is examined by using full sample LF-VAR Granger causality tests. Second, we use time-varying LF-VAR Granger causality tests to explore the dynamic relationship between national commodity prices and stock market returns. The latter considers the time-varying nature of the relationship between commodity and stock prices, as discussed by Miller and Ratti (2009).

The bootstrapped p-values from full sample Granger causality tests are reported in table C.8 in Appendix C.3. The null hypothesis of non-causality is tested against the alternative of causality. The rejection of the null hypothesis $H_0: SP \nRightarrow CP$ ($\nRightarrow$ means "does not Granger-cause") means that stock market returns Granger-cause national commodity prices, against the alternative hypothesis that stock market returns do not Granger-cause national commodity prices. Analogously, we test the null hypothesis $H_0: CP \nRightarrow SP$.

The full sample tests show a substantial number of causal patterns from stock market returns to national commodity prices. In fact, we find that stock market returns have a causal impact on national commodity prices in 49 out of 68 countries. In contrast, stock prices have weak impact on global shock variables (see tables 4.1 and 4.2). Particularly, stock markets exhibit the largest predictive power for global oil demand shocks among all global shock variables. In other words, stock market returns have causal impact on global oil demand shocks in the case of 27 out of 63 countries, according to the full sample MF tests. This finding implies that stock markets have greater influence over national commodity prices, while global commodities are less affected by movements in the stock prices.

At the same time, we find that national commodity prices have large impact on stock market returns in the countries from West African Economic and Monetary Union. The full sample tests show causality from national commodity prices to stock market returns in the case of five out of the eight countries part of the West African Economic and Monetary Union. This finding suggests that our constructed index of national commodity prices is a reliable predictor for stock market returns in developing economies that are heavily reliant on commodities. Otherwise, the full sample LF tests find minor evidence of causality from national commodity prices to stock market returns for the rest of the countries in our sample. The weak evidence for causality may be due to the structural changes that affect national commodity markets (see Chen et al., 2010) and, as such, the relationship between national commodity prices and stock market returns is not likely to remain stable over long time period.

As highlighted by Hood and Malik (2013) potential instability in the coefficient of commodity prices may alter the possibility of the full sample tests to detect causal patterns between commodity prices and stock market returns. Further to that, national stock prices are known to be sensitive to political and economic shocks, especially in emerging economies (Bekaert and Harvey, 1995). Therefore, we investigate the relationship between national commodity prices and stock market returns by adopting time-varying models.

Table C.9 reports the rejection frequencies (at different significant levels) for rolling window LF Granger causality tests, see Appendix C.3.[113] The null hypothesis of non-causality is specified for each rolling window. The rejection frequency for a single country is calculated as the total number of p-values within a 5% (or 10%) significance level is divided by the total number of rolling window tests. For example, the rejection frequency for testing the null hypothesis $H_0: CP \nRightarrow SP$ ($\nRightarrow$ means "does not Granger-cause") that national commodity prices do not Granger-cause stock market returns in Morocco is found to be 0.520 at a 10% level of significance. This implies that the null hypothesis is rejected at 52.0% of all rolling window tests for Morocco at a 10% level of significance. Therefore, we can conclude that national commodity prices have predictive power on Moroccan stock market returns, which is subject to time-variability.

Moreover, the LF tests show evidence of causality from stock market returns to national commodity prices for 62 out of 68 countries. This is a remarkable result suggesting that financial markets are still an important factor for national commodity markets around the world. In fact, financial markets of only five Sub-Saharan Africa economies (i.e. Ghana, Kenya, Mauritius, Nigeria and Uganda) and China are not found to have impact on national commodity markets. A possible reason for non-existence of causal patterns from stock to national commodity prices in the five aforementioned Sub-Saharan Africa countries is that

[113] The rolling window size is equal to half of the sample size. This choice is consistent to the main analysis provided in Chapter 4, as well as, to the study of Chen et al. (2010).

financial markets in smaller economies are much less developed than in other countries (IMF, 2008). Further to that, only 40% of smaller economies have a stock exchange, and trading in many of them is so low that their economic impact is minimal (IMF, 2008). In the case of China, our result is consistent to the findings of Cong et al. (2008).

At the same time, we find that national commodity prices have temporal impact on stock market returns in 50 out of 68 countries. This finding leads to two important conclusions. First, our constructed index series contain an important predictive power for numerous stock markets around the world. More explicit evidence is found for African countries, which are known to be heavily dependent on commodity trade (see Deaton and Miller, 1995; Deaton, 1999). Second, the time-varying LF tests show evidence of causality from commodity to stock prices in the case of five times more countries than their full sample counterparts; i.e. the full sample method finds evidence for only nine economies. This implies that the relationship between national commodity prices and stock market returns in exposed to time-variability. Overall, our constructed index series in Chapter 2 show certain predictive power over stock prices.

In a nutshell, this thesis contributes to the existing literature by revealing the important role of national commodity export prices for predicting stock market returns. This section adds to Chapter 3 that focuses primarily on economic growth. In that way, we show that both financial markets and economic growth are tightly interlinked with commodity prices.

4.8 Conclusions

Chapter 4 presents a cross-country analysis of the connection between global commodities and national financial markets for 63 countries and territories between January 1951 and March 2018. This study considers five measures of global commodities that we define as global shocks: world oil prices, world oil demand, world oil supply, world commodity prices (all items) and world metal prices.

Using a full sample MF-VAR approach proposed by Ghysels et al. (2016), we find that the oil market has an impact on stock market returns mainly through oil supply shocks. The MF approach suggests predictability from world oil supply shocks to stock market returns in the case of 18 out of 63 countries, while the LF approach finds such predictability for 16 out of 63 countries. Most of these countries (and regions) are oil-importers, suggesting a higher sensitivity of their national stock markets to oil supply shocks as compared to the exporting countries. Along these lines, the world oil prices and world oil demand have negligible influence on stock returns, especially in Africa, America and Oceania.

Moreover, this study considers the vital influence of world commodity prices (all items) over national stock markets. In terms of the MF approach, world commodity prices (all items) are found to predict stock market returns for 17 out of 63 countries, predominantly in developing economies, especially in Africa, where the economies are still heavily dependent on commodities. In a similar manner, this study provides a solid evidence for in-sample predictability from world metal prices to stock market returns, of which world metal prices have the strongest impact on national stock markets in Africa and Europe. This finding signifies the important role of the other commodities, apart from oil, on national stock markets.

Nonetheless, it should be noted that the full sample tests provide rather weak evidence in support of a causal link between global shocks and national stock returns. A potential reason for this is that the commodity-stock relationship has been changing over time. Therefore, we extend the full sample approach of Ghysels et al. (2016) to a time-varying framework.

Using a MF time-varying approach, this study reveals abundant evidence that commodity prices predict stock market returns. In the best-case scenario, we find that the world economic activity, denoted as world oil demand shocks, Granger-causes stock market returns in the case of 54 out of 63 countries. In the worst-case scenario, the world oil (metal) prices predict stock market returns in the case of 42 out of 63 countries. The results from world oil prices and world metal prices disclose an identical number of countries. Given these points, we confidently conclude that the commodity market plays an important role in the development of stock markets.

Is it possible to draw useful policy lessons from these arguments and results? Given the sustained and sharp increase in the variance of commodity prices, global stock markets remain heavily dependent on primary commodities. As one might have expected, financial development may affect growth by affecting saving rates, the allocation of saving, the profit margins and the business costs involved in the production of goods (Devereux and Smith, 1994). Equally important, we find that stock markets are more and more prone to events in global commodity market, which favours the presence of automated trading strategies operated by computers on multiple assets. This evolution in commodity and stock linkages reduces their potential substitutability in portfolios, as highlighted by Creti et al. (2013).

Future development of this study may consider extending the mixed-frequency analysis to a framework that investigates the reaction of national stock markets to a structural shock in the global commodities – a concept that is consistent with the works of Apergis and Miller (2009) and Kilian and Park (2009).

Chapter 5. Conclusions

5.1 Overview of the Research

This thesis is mainly concerned with exploring the unidentified effects of commodity price dynamics on the economic and financial sectors of commodity-exporting and importing economies.

After an introductory chapter, this thesis addresses the following aspects concerning commodity prices and their impact on the macroeconomy and financial markets. Based on theoretical foundations, a database with national commodity export price indexes has been constructed in Chapter 2. These index measures are used in the robustness check section of Chapter 3 to identify the forecasting ability of national commodity export prices on economic growth. With this intention, the chief contribution of Chapter 3 relates to investigating the forecasting ability of world commodity prices on economic growth for commodity-dependent exporting and importing countries. The next chapter, i.e. Chapter 4, analyses the causal links between commodity and stock markets.

Particularly, the objectives of this study include the following: **(1)** creating a world database with precise index measures of national commodity export prices (Chapter 2), **(2)** testing the in-sample and out-of-sample predictability of commodity prices for economic growth in commodity-dependent economies (Chapter 3), **(3)** verifying whether the potential commodity-growth relationship changes over time (Chapter 3), **(4)** investigating the historical link between world commodity prices and national stock market returns (Chapter 4) and **(5)** examining which global commodity measure exercises the greatest influence over national stock markets (Chapter 4). In essence, this study has successfully enhanced the understanding of the influences of commodity prices as an important factor for investors and policymakers who wish to extract macroeconomic expectations from the commodity price fluctuations and take action regarding the link between commodity prices with economic growth and national stock markets.

During the process of achieving the above objectives, certain choices and trade-offs are inevitably made in the context of the study sample and methodology. For example, the investigation is predominantly conducted in commodity-dependent economies due to their high economic and financial reliance on commodity trade. Moreover, the study employs a recent mixed-frequency vector autoregressive (MF-VAR) approach as the main statistical method to derive the final findings due to a number of limitations that are inherited in the same-frequency VAR approach. The mixed-frequency approach allows data of different frequencies to be estimated in the same empirical framework. The results from the mixed-frequency approach provide strong evidence of a causal link between commodity prices and economic growth (Chapter 3) and for the existence of causality from commodity prices to national stock market returns (Chapter 4). To provide a better insight into the main findings of the three empirical studies of this thesis, more information has been provided in the subsequent section.

5.2 Main Findings of the Empirical Analysis

The central findings of this thesis have been illustrated through the following appealing features.

5.2.1 Main findings of Chapter 2: Towards a new database of country-specific price indexes of commodity exports

To begin with, Chapter 2 contributes to the commodity literature by constructing a new world database of country-specific price indexes of commodity exports for a set of 217 countries. All index series are represented in a monthly frequency. The coverage of the database spans from January 1980 to April 2017.

The construction of the database is based on the index number formula of Deaton and Miller (1995). The main reasons for selecting the DM index as the most appropriate formula have been listed as follows. First, it allows constructing a database that is rich in terms of the number of countries it includes by using the existing data in the world trade statistics. Second, it allows constructing lengthy index series that predispose undertaking a (historical) time series analysis.

Particularly, most index formulas require continuous volume data for the construction of index series, e.g. Fisher and Paasche indexes, while the DM formula does not. More precisely, the volume trade data that is disaggregated to a national level of commodity-specific exports is rarely available, especially for developing countries. Although such data may be available, it is commonly discontinued due to external and internal factors, such as civil war and trade barriers. The DM index overcomes these limitations, which is why we select it as the most appropriate index formula to construct our world database of country-specific price indexes of commodity exports.

Another important feature of the DM index is that it allows the database to be easily updated. Therefore, the researcher is able to explore the most recent price fluctuations in the national commodity markets.

Furthermore, Chapter 2 improves upon past studies, such as those of Deaton and Miller (1995), Dehn (2000) and Cashin et al. (2004), by providing a robust framework for index construction. A drawback of the previous studies is that they focus more on the choice of index formula instead of the appropriateness (quality) of the data. For example, past studies obtain their data for commodity index weights from a single revision trade classification system. However, Chapter 2 identifies that using a single revision trade classification system leads to an inaccurate construction of commodity index weights and, therefore, the constructed indexes may fail to represent the true price movements in national commodity markets. Thus, this study suggests a new approach for data collection that synchronises information from various revisions of trade classification systems, thereby overcoming the issues arising from data collection with the aim of accommodating more precise and accurate data sets.

Another key contribution of this study is the inclusion of 72 commodities in the index basket of the country-specific commodity price indexes in our database. The advantages of including such a large set of commodities are illustrated in the following appealing features.

First, it reduces the possibility of omitting a major exporting commodity from the individual index basket, thereby improving the accuracy of the constructed indexes. Such an example is the exclusion of dairy products from the index basket of the constructed index series by Sahay et al. (2002) for New Zealand (see Section 2.2.3).[114]

It is empirically confirmed that the accuracy of the index series improves by including dairy products in the index commodity basket. The Pearson correlation coefficient between the index series of Sahay et al. (2002) and the official benchmark for New Zealand is found to be 0.407, whereas the correlation coefficient between the constructed index series and the official benchmark for New Zealand is found to be 0.940. Thus, the inclusion of dairy products in the index basket leads to improvement of the accuracy of the constructed index series.

Second, it allows constructing country-specific price sub-indexes of commodity exports, thereby making it a convenient tool for commodity-specific research. For example, Cashin et al. (2004) focus only on non-fuel commodities, while future studies may be interested in other commodity markets such as metals or dairy. For this reason, our database includes country-

[114] The reason for the exclusion of dairy products from the index series of Sahay et al. (2002) is not due to a measurement error. It is a consequence of data unavailability, as highlighted by the authors.

specific price *sub-indexes* of commodity exports for 13 different commodity categories (see Section 2.3.4).[115]

In particular, this study discovers a strong positive correlation between the constructed sub-index series and the official benchmark sub-indexes. More precisely, the Pearson correlation coefficient between the constructed non-energy index and the official non-energy index for Canada is found to be 0.989, and the correlation coefficient between the constructed dairy index series and the official dairy benchmark index for New Zealand is found to be 0.988. These findings provide evidence supporting the high accuracy of the sub-index series in our database.

Consequently, Chapter 2 concludes that there is a statistically significant correlation between the constructed index series and the official benchmark indexes. This is to say, the two correlation tests (i.e. Pearson and Spearman tests) provide solid support for a strong positive correlation existing between the constructed index series and the official benchmark indexes. Moreover, all correlations are significant at 1% significance level. This provides further evidence of the reliability of our constructed national commodity export price indexes.

As academic researchers are increasingly paying more attention to whether a correlation exists between the chain-linking and the fixed-based indexes, this study provides evidence of a correlation in terms of commodity price indexes.

Chapter 2 empirically demonstrates that the index series constructed using the DM formula are highly correlated with the official index series of central and commercial banks, which are created through chain-linking. This provides more confidence with respect to the appropriateness of the DM index formula for constructing a world database of country-specific price indexes of commodity exports, as well as for the accuracy of our database by making it (possibly) a key factor that underpins evidence for future research on commodity markets.

5.2.2 Main findings of Chapter 3: Do commodity prices predict economic growth? A mixed-frequency time-varying investigation

The next contribution of this thesis is made in Chapter 3, which looks at the forecasting power of commodity prices for economic growth for a set of 33 commodity-dependent countries between January 1980 and December 2016.

[115] The construction of a sub-index is dependent on the country's export basket. If a certain group of commodities are a part of the country's exports, the sub-index for this group is not constructed.

As economic growth data is mainly available in at least a quarterly frequency, while commodity prices are available at a higher frequency, the use of standard same-frequency models requires temporal aggregation of the high-frequency series. However, the temporal aggregation generates spurious and hidden effects, as discussed by Marcellino (1999). On the one hand, the finite-sample power of the testing procedure may decline when temporally aggregated data is used. This attributes to the small number of available observations (Haug, 2002). On the other hand, the temporal aggregation commonly leads to high mean squared forecast errors (Marcellino et al., 2016). Therefore, Chapter 3 addresses this limitation by adopting the mixed-frequency approach proposed by Ghysels et al. (2016), which allows data in different frequencies to be analysed in the same framework.

The full sample mixed-frequency tests provide evidence that commodity prices predict economic growth. In particular, world commodity prices are found to have forecasting power on economic growth in commodity-exporting countries, whereas many more cases of causality are revealed for commodity-importing and hybrid economies.

Moreover, we acknowledge that the relationship between commodity prices and economic growth may change over time. This is confirmed empirically by using the Andrews' (1993) QLR structural break tests, which led to discovering evidence of instability for nearly half of the countries in the sample.

Hence, by using a mixed-frequency time-varying approach, this study provides evidence of short-horizon in-sample causality from commodity prices to economic growth in the case of 31 out of 33 countries. In the meantime, the feedback causality is found for 23 out of 33 countries.

Furthermore, this study finds stronger evidence for causality from commodity prices to economic growth in a short-horizon than a longer one. Nonetheless, there are some countries where causal patterns are not detected in the short-horizon, while evidence appears when the estimation horizon is longer. Such an example is Viet Nam, where causal patterns are detected only when the estimation horizon is at least four quarters, i.e. at a longer-horizon. This finding suggests a possible lagged effect of commodity prices for predicting economic growth.

In addition, this thesis finds solid evidence in support of the out-of-sample predictive ability of commodity prices on economic growth. The forecast combination results provide evidence that the commodity-based models outperform the benchmark models in 79% of the total number of countries for at least two of the three benchmarks. In other words, the commodity-based predictive regressions outperform the benchmark models in the majority of the countries. This finding is consistent along both the LF and MF methods. It also provides empirical evidence that the in-sample predictability translates into out-of-sample success. Despite the potential for natural resource wealth to promote economic growth and development, excessive commodity dependence, for the most part, remains a barrier to economic development (UN, 2019). The thesis findings emphasise that countries that are heavily dependent on commodities for their main source of income are in need of diversification. Therefore, this study suggests that greater economic diversification may help protect economies against commodity price volatility. This is closely associated with developing linkages between the commodity sector and the rest of the economy.

Furthermore, it is necessary for policy makers to use the predictive content of commodity prices to establish such strategies that reduce the commodity harmfulness on economic growth. Achieving this requires a comprehensive approach to commodity management embedded within a broad sustainable development strategy, careful management of resource revenues and firm policy commitments. The UN (2019) suggests three key policy objectives that commodity-dependent countries need to address – building resilience against volatility, expanding linkages from the commodity sector to the rest of the economy and developing necessary human and physical capital. Moreover, the governments of commodity-dependent economies should adopt countercyclical fiscal policies, accumulating savings during times of price booms, and raising government spending when commodity prices are low to compensate for the economic slowdown. In this aspect, some countries have established revenue stabilisation funds as buffers against commodity price fluctuations. The advisory boards of the revenue stabilisation funds can use the findings reported in this thesis to enhanced understanding of the inter-linkages between commodity prices and economic growth. Particularly, it is vital to ensure that the returns from natural resources are widely shared across society and directed towards promoting development objectives and productive investment (UN, 2019).

5.2.3 Main findings of Chapter 4: Global commodity markets and national financial markets: a mixed-frequency time-varying investigation

Chapter 4 presents a cross-country analysis of the causal relationship between global commodities and national financial markets in a set of 63 countries and territories between January 1951 and March 2018 using the MF-VAR approach proposed by Ghysels et al. (2016). The main reason for employing this method is that temporal aggregation in Granger causality tests is an important, and yet often overlooked, problem that can generate spurious and hidden effects. The MF approach addresses this limitation and exploits all available data irrespective of their sampling frequency.

Particularly, this study exploits recent econometric methods that account for data sampled at different frequencies, which is unfortunately the case of stock markets and commodity prices. This study considers five measures of global commodities where each is defined as a global shock variable: world oil prices, world oil demand, world oil supply, world commodity prices (all items) and world metal prices.

The full sample tests provide weak evidence in support of the causal relationship between global shock variables and national stock market returns. A possible reason for this is that the commodity-stock relationship is affected by structural breaks. Therefore, the full sample approach of Ghysels et al. (2016) is extended to a time-varying framework.

Using a MF time-varying approach, the empirical results of this study present evidence that commodity prices predict stock market returns. In the best-case scenario, it is found that the world economic activity, proxied by world oil demand shocks, Granger-causes stock market returns in the case of 54 out of 63 countries. In the worst-case scenario, it is identified that the stock market returns are Granger-caused by world oil (metal) prices in the case of 42 out of 63 countries. The results from the world oil prices and the world metal prices provide an identical total number of countries. Given these points, this study confidently concludes that the commodity prices play an important role in forecasting stock market returns. In particular, this study finds world oil demand to be a better predictor of stock market returns in comparison to world oil (metal) prices.

Based on the aforementioned findings, Chapter 4 concludes that none of our measures for global commodities has a dominant influence (i.e. beat others) over all the stock markets around the world. A summary of the time-varying MF results is provided in the following table:

Continent	World oil prices	World oil supply shock	World oil demand shock	World commodity prices	World metal prices
Africa	11(14)	10(14)	9(14)	9(14)	8(14)
Americas	3(7)	6(7)	6(7)	4(7)	5(7)
Asia	9(17)	15(17)	15(17)	13(17)	14(17)
Europe	18(23)	16(23)	22(23)	18(23)	14(23)
Oceania	1(2)	2(2)	2(2)	2(2)	1(2)

Note: The table shows the number of countries for which causality has been detected from global shocks to national stock returns. The number in brackets denotes the total number of countries considered in the analysis, per each continent. The results are based on the time-varying MF approach, as reported in table 4.3.

Table 5.1 Number of Countries for which Causality has been Detected from Global Shocks to National Stock Returns

In brief, Chapter 4 draws conclusions that might have important implications for policy design, financial advisers, investment managers and so on.

Financial advisers and investment managers can use the findings reported in this thesis as an indicator of the individual stock market dependency on commodities. This provides invaluable information for stock investors who seek diversification benefits from commodity investments, as they can design optimal portfolios and hedging strategies based on the detected interrelations. Equally important, this study finds that stock markets are more and more prone to events in global commodity market, which favours the presence of automated trading strategies operated by computers on multiple assets. This evolution in commodity and stock linkages reduces their potential substitutability in portfolios, as highlighted by Creti et al. (2013). Therefore, investors need to observe carefully the economic situation and the stock market exposure to the commodity risks.

Furthermore, the findings of this thesis suggest policy makers to attempt in understanding the exposure of individual stock markets to commodity risks. Particularly, it is necessary for policy makers to organise such strategies that help to reduce the commodity shocks harmfulness on financial market. Given the results of this research, world oil demand shock is the main mechanism causing stock market returns in European stock markets, while African stock markets are mainly driven by the changes in world oil prices. Therefore, it is critical for policy makers to adopt an integrated risk management approach. For instance, employing alternative energy resources can decrease the dependence on oil for production purposes. Specifically, this thesis recommends governments to adopt an explicitly integrated approach to assessing the key risks of stock market dependency on commodities and to develop options on how best to mitigate the identified risks, including through developing a national risk management strategy and supporting institutions.

5.3 Suggestions for Future Research

With detailed discussions about the study's findings and contributions, this final section aims to identify future directions and research opportunities in the area.

- By employing export trade data, this research cannot explicitly draw conclusions about the changes in national commodity prices for import commodity-dependent economies. Therefore, future research should consider using import trade data to construct a new database of national commodity import price indexes.
- One potential issue of the commodity price index measurement is that the difference in index fluctuations across periods may instead result from different structural changes in (national) commodity markets. Therefore, future research should attempt to develop a new index measure of (national) commodity prices that can tackle this issue.
- The database of national commodity export prices has a limitation such that, without the presence of the diamonds, the quality of the index series for diamond exporting countries (e.g. Botswana) cannot be ensured despite numerous quality-check methods being utilised. Therefore, future research should employ additional commodity trade data, including diamonds, to confirm the findings obtained in this thesis.
- Utilising the database of national commodity export prices from this thesis, further research should be conducted to obtain a general conclusion regarding how these indexes are associated with other economic and financial variables such as exchange rate, inflation, unemployment rate, interest rate, industrial production, economic growth, stock market returns and so on.
- Using aggregate data of commodity prices may lead the empirical results to be subject to temporal aggregation bias. In particular, the world commodity price data is commonly available at a frequency higher than monthly. Future research may consider collecting daily or weekly data of commodity prices and fitting it within a commodity-growth model. With such data, a new mixed-frequency method that addresses the issue of parameter proliferation should be adopted.
- Potential controlling and moderating factors that are related to the commodity-growth relationship should be taken into consideration in future research, for example, inflation. In essence, the high inflation tends to drive up the prices of commodities

(Gospodinov and Ng, 2013), leading to a decline in the purchasing power of money which, in turn, reduces consumption and, therefore, the GDP decreases.

- ❖ Although a myriad of research work has been conducted to study the effects of commodity prices on economic growth and stock returns, minimal research has been conducted in terms of impulse response analysis in a mixed-frequency (multivariate) setting, considering putting restrictions within mixed-frequency VAR and providing clear identification strategy to explore the reaction of economic and financial variables to structural price shocks in commodity markets. This thesis provides room for future research to fill this gap.
- ❖ Another possible extension of this thesis includes the estimation of the long-run relationship between variables and determining the speed of disequilibrium error correction (Enders, 2014). In fact, the speed of adjustment determines the rate at which the dependent variable corrects short-run deviations. Then, when the variables are out of long-run equilibrium, there are economic forces, captured by the adjustment coefficients, that push the model back to long-run equilibrium. The speed of adjustment toward equilibrium is determined by the magnitude of γ_s. For example, γ_s=0.5 implies that roughly one half of the disequilibrium errors corrected in one time period. If γ_s= 1 then the entire disequilibrium is corrected in one period. If γ_s=1.5 then the correction overshoots the long-run equilibrium (Zivot and Wang, 2006). The impulse response function may be used to determine the speed of adjustment to long-run equilibrium, and the half-life of a shock can be estimated (see Rossi, 2005; Zivot and Wang, 2006). This entails new macro-econometric techniques to be introduced for the analysis of mixed-frequency datasets.
- ❖ Future research can potentially consider doing rolling window estimation of VAR effects and collect the coefficients of error correction parameters, and roll the same.
- ❖ Future investigation should attempt to identify the underlying reason behind the variation in the commodity-stock relationship. For example, the current thesis adopts a time-varying approach that can be extended to a graphical analysis that illustrates the exact periods of shifts from non-causality to causality (and vice versa) and looks at the findings more theoretically or identifies alternative explanations.
- ❖ Last but not least, we acknowledge the importance of commodity price volatility; however, in this study, we ignore it and leave it for a further research.

Appendices

Appendix A

A.1 Data specifications and sources

Primary Commodity	Price Specifications	Unit	Sources
Aluminium	London Metal Exchange, standard grade, spot price, minimum purity 99.5 percent, CIF U.K. ports (Wall Street Journal, New York and Metals Week, New York). Prior to 1979, U.K. producer price, minimum purity 99%	US$/mt	IMF (2018)
Bananas	Central American and Ecuador, first class quality tropical pack, Chiquita Dole and Del Monte, US importer's price FOB US ports (Sopisco News, Guayaquil)	US$/mt	IMF (2018)
Barley	Through May 2012: Canadian No. 1 Western Barley, spot price (Winnipeg Commodity Exchange). From June 2012 onwards: US No. 2 feed barley, Minneapolis delivery spot price (USDA) (Datastream)	US$/mt	IMF (2018)
Beef	Australian and New Zealand, frozen boneless, 85% visible lean cow meat, US import price FOB port of entry	US Cts/lb	IMF (2018)
Butter	Butter (European & Oceania average indicative export prices, FOB) Average of mid-point of price ranges reported bi-weekly by Dairy Market News (USDA). Prior to January 1990, International Dairy Arrangement	US$/mt	International Dairy Arrangement (1981; 1982; 1983; 1984; 1985; 1986; 1987; 1988; 1989; 1990) / FAO (2018)
Canola	August 2012 to Present – Canadian International Merchandise Trade Database, Statistics Canada Note: April 1983–July 2000 (Crude Degummed Oil) FOB PLANTS; August 2000 to July 2012 (Crude Degummed Oil) FOB. Vancouver – Source: Cereals & Oilseeds Review, Statistics Canada	US$/mt	Canola Council of Canada (2018)
Cheese	Cheddar Cheese (European & Oceania average indicative export prices, FOB) Average of mid-point of price ranges reported bi-weekly by Dairy Market News (USDA). Prior to January 1990, International Dairy Arrangement	US$/mt	International Dairy Arrangement (1981; 1982; 1983; 1984; 1985; 1986; 1987; 1988; 1989; 1990) / FAO (2018)
Coal	Australian thermal coal, 12000 btu/pound, less than 1% sulfur, 14% ash, fob prices, Newcastle/Port Kembla (Argus Media Group)	US$/mt	IMF (2018)
Coarse, Wool	23 micron (AWEX, Australian Wool Exchange) Sydney, Australia	US Cts/kg	IMF (2018)
Cocoa Beans	International Cocoa Organization cash price. Average of the three nearest active futures trading months in the New York Cocoa Exchange at noon and the London Terminal market at closing time, CIF US and European ports (The Financial Times, London).	US$/mt	IMF (2018)

Coconut oil	Crude, in bulk, Philippines/Indonesia, CIF Rotterdam. Prior to 1973: Sri Lanka, 1% bulk, CIF European ports. (Oil World, Hamburg, Germany)	US$/mt	UNCTAD (2018)
Coffee	ICO Composite indicator price, weighted as follows: - 10% Colombian milds (54% the US and 46% the EU) - 23% Other mild (41% the US and 59% the EU) Arabicas - 30% Brazilian naturals (26% the US and 74% the EU) - 37% Robustas (17% the US and 83% the EU) For previous weights of I.C.A., please refer to the International Coffee Organization's (ICO) website. (International Coffee Organization (ICO), London, United Kingdom)	US Cts/lb	UNCTAD (2018)
Copper	London Metal Exchange, grade A cathodes, spot price, CIF European ports (Wall Street Journal, New York and Metals Week, New York) Prior to July 1986, higher grade wirebars or cathodes	US$/mt	IMF (2018)
Copra	Copra (Philippines/Indonesia), bulk, CIF N.W. Europe	US$/mt	IMF (2018)
Cotton	Middling 1-3/32 inch staple, Liverpool Index "A", average of the cheapest five of fourteen styles, CIF Liverpool (Cotton Outlook, Liverpool). From January 1968 to May 1981 strict middling 1-1/16 inch staple. Prior to 1968, Mexican 1-1/16	US Cts/lb	IMF (2018)
Cottonseed oil	The US, crude cottonseed oil, FOB Mississippi Valley. (The Public Ledger, London, the United Kingdom) Prior to January 2014, Crude, in bulk, the US, Prime Bleachable Summer Yellow (PBSY), FOB Gulf. Prior to October 1994: the US, PBSY, CIF Rotterdam (Oil World, Hamburg, Germany)	US$/mt	UNCTAD (2018)
Crude oil	Average of U.K. Brent (light), Dubai (medium) and West Texas Intermediate (heavy), equally weighted	US$/bbl	IMF (2018)
DAP	DAP (diammonium phosphate), standard size, bulk, spot, FOB US Gulf	US$/mt	IMF (2018)
Fine, Wool	19 micron (AWEX, Australian Wool Exchange) Sydney, Australia	US Cts/kg	IMF (2018)
Fish meal	Peru Fish meal/pellets 65% protein, CIF Germany (Datastream)	US$/mt	IMF (2018)
Gaseous state, Natural Gas	Natural Gas (US), spot price at Henry Hub, Louisiana	US$/mm btu	World Bank (2018)
Gold	Gold (UK), 99.5% fine, London afternoon fixing, average of daily rates	US$/troy oz	World Bank (2018)
Groundnut /peanut/ oil	Groundnut oil (any origin), CIF Rotterdam.	US$/mt	IMF (2018)
Groundnuts	40/50 (40 to 50 count per ounce), in-shell, cif Argentina (Datastream)	US$/mt	IMF (2018)

Hides	US, Chicago packer's heavy native steers, over 53 lbs.,wholesale dealer's price, (formerly over 58 lbs.), FOB shipping point (Wall Street Journal, New York). Prior to November 1985, US Bureau of Labor Statistics, Washington, D. C.	US Cts/lb	IMF (2018)
Iron Ore	Iron Ore (FE63.5%) in CIF China (Datastream)	US$/mt	IMF (2018)
Jute	Bangladesh, BWD (Bangladesh White D), FOB Mongla (Food and Agricultural Organisation, Rome, Italy) (Prior to 2004: The Public Ledger, London, United Kingdom) Prior to March 1980: Chittagong-Chalna, minimum export price (Ministry of Jute, Bangladesh)	US$/mt	IFS (2018)
Lamb	New Zealand, PL, frozen, wholesale price at Smithfield Market, London (National Business Review, Auckland, New Zealand)	US Cts/lb	IMF (2018)
Lead	London Metal Exchange, 99.97% pure, spot price, CIF European ports (Wall Street Journal, New York and Metals Week, New York)	US$/mt	IMF (2018)
Linseed oil	Crude, in bulk, any origin, ex-tank Rotterdam. Prior to January 1977: any origin, CIF London/Hull. Prior to 15 September 1969: Argentina, in bulk, CIF United Kingdom (Oil World, Hamburg, Germany)	US$/mt	IFS (2018)
Liquefied, Natural Gas	Natural gas LNG (Japan), import price, CIF, recent two months' averages are estimates	US$/mm btu	World Bank (2018)
Logs, Hardwood	Malaysian, meranti, Sarawak best quality, sale price charged by importers, Japan (World Bank, Washington, D.C.). From January 1988 to February 1993, average of Sabah and Sarawak in Tokyo weighted by their respective import volumes in Japan. From February 1993 to present, Sarawak only	US$/Cm	IMF (2018)
Logs, Softwood	Oregon Logs: (free alongside ship (FAS) Value)/(First Unit of Quantity) by FAS Value for US domestic exports, exported through Oregon ports	US$/Cm	IMF (2018)
Maize	US No. 2 yellow, prompt shipment, FOB Gulf of Mexico ports (USDA, Grain and Feed Market News, Washington, D.C.)	US$/mt	IMF (2018)
Nickel	London Metal Exchange, melting grade, spot price, CIF Northern European ports (Wall Street Journal, New York and Metals Week, New York). Prior to 1980 INCO, melting grade, CIF Far East and American ports (Metal Bulletin, London)	US$/mt	IMF (2018)
Olive oil	United Kingdom ex-tanker prices, extra virgin olive oil, 1%> ffa (free fatty acid) (Datastream). From December 2011 onwards: Olive Oil Eu / Extra Virgin, Italy CIF (Cost, Insurance, Freight) (Bloomberg)	US$/mt	IMF (2018)

Orange	Miscellaneous Oranges, French import price (FruiTROP and World Bank)	US$/mt	IMF (2018)
Palm kernel oil	Crude, in bulk, Malaysia/Indonesia, CIF Rotterdam. Prior to September 1980: Dutch, FOB Ex-Mill (Oil World, Hamburg, Germany)	US$/mt	UNCTAD (2018)
Palm oil	Palm Oil Futures (first contract forward) 4–5 percent FFA Bursa Malaysian Derivatives Berhad	US$/mt	IMF (2018)
Pepper	Indonesian Muntok, white, FAQ, EXW Rotterdam. Prior to July 2012: White Muntok, FAQ, spot. (The Public Ledger, London, United Kingdom). Prior to June 2003: white Sarawak, closing quotations, Singapore (Market News Service, ITC, UNCTAD/WTO, Geneva, Switzerland)	US Cts/lb	IFS (2018)
Phosphate rock	Phosphate rock (Morocco), 70% BPL, contract, f.a.s. Casablanca	US$/mt	IFS (2018)
Platinum	Platinum (UK), 99.9% refined, London afternoon fixing	US$/troy oz	World Bank (2018)
Plywood	Africa and Southeast Asia, Lauan, 3-ply, extra, 91 cm x 182 cm x 4 mm, wholesale price, spot Tokyo (World Bank, Washington D.C., the US) Prior to January 2002: Ministry of Agriculture, Forestry and Fisheries, Tokyo, Japan	US$/Cm	UNCTAD (2018)
Potash	Potassium chloride (muriate of potash), standard grade, spot, FOB Vancouver	US$/mt	IFS (2018)
Poultry	Georgia docks, ready to eat whole body chicken, packed in ice, spot price (USDA).	US Cts/lb	IMF (2018)
Pulp	Woodpulp (Sweden), softwood, sulphate, bleached, air-dry weight, CIF North Sea ports	US$/mt	IFS (2018)
Rapeseed oil	Rapeseed Oil European Union Ex-Mill Free on Board Rotterdam Current Month (Datastream)	US$/mt	IMF (2018)
Rice	Export Prices (FOB) of Thailand 5% Grade Parboiled Rice (USDA, Rice Market News, Little Rock, Arkansas)	US$/mt	IMF (2018)
Rubber	SGX Ribbed Smoked Sheet 3 (RSS3) Futures (Bloomberg, RG1 Comdty)	US Cts/lb	IMF (2018)
Salmon	Norwegian Salmon, fresh or chilled, fish-farm bred, export price (Statistics Norway)	US$/kg	IMF (2018)
Sawnwood, Hardwood	Malaysian sawnwood, dark red meranti, select and better quality, standard density, C&F U.K. Port (Tropical Timbers, Surrey, England).	US$/Cm	IMF (2018)
Sawnwood, Softwood	Oregon Lumber: (free alongside ship (FAS) Value)/(First Unit of Quantity) by FAS Value for US domestic exports, exported through Oregon ports	US$/Cm	IMF (2018)
Shrimp	Mexican, west coast, white, No. 1, shell-on, headless, 26 to 30 count per pound, wholesale price at New York	US$/kg	IMF (2018)
Silver	Silver (UK), 99.9% refined, London afternoon fixing; prior to July 1976 Handy & Harman. Grade prior to 1962	US$/troy oz	World Bank (2018)

	unrefined silver		
Sisal	Tanzania/Kenya No. 3 UG, FOB. (Food and Agricultural Organization, Rome, Italy) Prior to 2007: CIF main European ports (Prior to 2004: The Public Ledger, London, United Kingdom)	US$/mt	IFS (2018)
Skim milk powder	Skim Milk Powder (European & Oceania average indicative export prices, FOB) Average of mid-point of price ranges reported bi-weekly by Dairy Market News (USDA). Prior to January 1990, International Dairy Arrangement	US$/mt	International Dairy Arrangement (1981; 1982; 1983; 1984; 1985; 1986; 1987; 1988; 1989; 1990) / FAO (2018)
Sorghum	Sorghum (US), no. 2 milo yellow, FOB Gulf ports.	US$/mt	IFS (2018)
Soybean meal	Soybean Meal Futures (first contract forward) Minimum 48 percent protein (Chicago Board of Trade)	US$/mt	IMF (2018)
Soybean oil	Soybean Oil Futures (first contract forward) exchange approved grades (Chicago Board of Trade)	US$/mt	IMF (2018)
Soybeans	Soybean futures contract (first contract forward) No. 2 yellow and par (Chicago Board of Trade)	US$/mt	IMF (2018)
Sugar	CSCE contract No. 11, nearest future position Cts/lb (Coffee, Sugar and Cocoa Exchange, New York Board of Trade)	US Cts/lb	IMF (2018)
Sunflower Oil	Sunflower Oil, crude, US export price from Gulf of Mexico (Datastream)	US$/mt	IMF (2018)
Swine Meat	51-52% (.8 - .99 inches of back fat at measuring point) lean Hogs, USDA average base cost price of back fat measured at the tenth rib (USDA)	US Cts/lb	IMF (2018)
Tea	Tea , average three auctions, arithmetic average of quotations at Kolkata, Colombo and Mombasa/Nairobi	US$/kg	World Bank (2018)
Tin	London Metal Exchange, standard grade, spot price, CIF European ports (Wall Street Journal, New York, New York). From Dec. 1985 to June 1989 Malaysian, straits, minimum 99.85 percent purity, Kuala Lumpur Tin Market settlement price. Prior to November 1985, London Metal Exchange (Wall Street Journal, New York and Metals Week, New York)	US$/mt	IMF (2018)
Tobacco	Tobacco (any origin), unmanufactured, general import , CIF, US	US$/mt	IFS (2018)
TSP	TSP (triple superphosphate), bulk, spot, beginning October 2006, Tunisian origin, granular, fob; previously US origin, FOB US Gulf	US$/mt	IFS (2018)
Uranium	Metal Bulletin Nuexco Exchange Uranium (U3O8 restricted) price.	US$/lb	IMF (2018)
UREA	Urea, (Black Sea), bulk, spot, FOB Black Sea (primarily Yuzhnyy) beginning July 1991; for 1985–91 (June) FOB Eastern Europe	US$/mt	IFS (2018)

Wheat	US No. 1 hard red winter, ordinary protein, prompt shipment, Kansas City (USDA, Grain and Feed Market News, Washington, DC)	US$/mt	IMF (2018)
Whole milk powder	Whole Milk Powder (European & Oceania average indicative export prices, FOB) Average of mid-point of price ranges reported bi-weekly by Dairy Market News (USDA). Prior to January 1990, International Dairy Arrangement	US$/mt	International Dairy Arrangement (1981; 1982; 1983; 1984; 1985; 1986; 1987; 1988; 1989; 1990) / FAO (2018)
Zinc	London Metal Exchange, high grade 98 percent pure, spot price, CIF U.K. ports (Wall Street Journal and Metals Week, New York). Prior to January 1987, standard grade	US$/mt	IMF (2018)

Table A.1 Data Specifications and Sources

A.2 Products conversion formulas

Pound-mass (lb) to metric tonne (mt)

When $\boldsymbol{P_k}$ is the world price of commodity k that is reported in US dollars per pound-mass, it can be converted to US dollars per metric tonnes by dividing it by a standard conversion factor of $\mathbf{0.000453592}$. The converted world price is then obtained as follows:

$$\boldsymbol{P_{k,US\$/mt}} = \frac{\boldsymbol{P_{k,US\$/lb}}}{\mathbf{0.000453592}} \tag{A.1}$$

When $\boldsymbol{P_k}$ is the world price of commodity k that is reported in US cents per pound-mass, it can be converted to US cents per metric tonnes by dividing it by a standard conversion factor of $\mathbf{0.000453592}$. Then, prices are converted from cents to US dollars by multiplying by a standard conversion factor of $\mathbf{0.01}$. The converted world price is then obtained as described below:

$$\boldsymbol{P_{k,US\$/mt}} = \left(\frac{\boldsymbol{P_{k,US\,Cts/lb}}}{\mathbf{0.000453592}}\right) * \mathbf{0.01} \tag{A.2}$$

Kilogram (kg) to metric tonne (mt)

When $\boldsymbol{P_k}$ is the world price of commodity k that is reported in US dollars per kilogram, it can be converted to US dollars per metric tonnes by dividing it by a standard conversion factor of $\mathbf{0.001}$. The converted world price is then obtained as shown below:

$$\boldsymbol{P_{k,US\$/mt}} = \frac{\boldsymbol{P_{k,US\$/kg}}}{\mathbf{0.001}} \tag{A.3}$$

Troy ounce (troy oz) to metric tonne (mt)

When $\boldsymbol{P_k}$ is the world price of commodity k that is reported in US dollars per troy ounce, it can be converted to US dollars per metric tonnes by dividing it by a standard conversion factor of $\mathbf{0.000031103477}$. The converted world price is then obtained as follows:

$$\boldsymbol{P_{k,US\$/mt}} = \frac{\boldsymbol{P_{k,US\$/troy\,oz}}}{\mathbf{0.000031103477}} \tag{A.4}$$

Cubic meter (m^3) to metric tonne (mt)

When $\boldsymbol{P_k}$ is the world price of commodity k that is reported in US dollars per troy ounce, it can be converted to US dollars per metric tonnes by dividing it by a standard conversion factors of $\mathbf{1.43}$ (for Sawnwood, Hardwood), $\mathbf{1.25}$ (for Logs, Hardwood), $\mathbf{1.43}$ (for Logs, Softwood), $\mathbf{1.82}$ (for Sawnwood, Softwood) and $\mathbf{1.54}$ (for Plywood). The converted world price is then obtained as follows:

$$P_{k,US\$/mt} = \frac{P_{k,US\$/m^3}}{conversion\ factor\ for\ the\ specific\ timber's\ product} \quad \text{(A.5)}$$

Barrels (bbl) to metric tonne (mt)

When $\boldsymbol{P_k}$ is the world price of commodity k that is reported in US dollars per bbl, it can be converted to US dollars per metric tonnes by dividing it by a standard conversion factor of $\mathbf{7.352941176}$. The converted world price is then obtained as follows:

$$P_{k,US\$/mt} = \frac{P_{k,US\$/bbl}}{7.352941176} \quad \text{(A.6)}$$

One million British Thermal Units (mmbtu) to metric tonne (mt)

When $\boldsymbol{P_k}$ is the world price of commodity k that is reported in US dollars per mmbtu, it can be converted to US dollars per metric tonnes by dividing it by a standard conversion factor of $\mathbf{40.2}$. The converted world price is then obtained as shown below:

$$P_{k,US\$/mt} = \frac{P_{k,US\$/mmbtu}}{40.2} \quad \text{(A.7)}$$

US cents (US¢) to US dollars ($)

When $\boldsymbol{P_k}$ is the world price of commodity k that is reported in US cents, it can be converted from US cents (US¢) to US dollars (US$) by multiplying it by a standard conversion factor of $\mathbf{0.01}$. The converted world price is then obtained as described below:

$$P_{k,US\$} = P_{k,US\ Cts} * 0.01 \quad \text{(A.8)}$$

A.3 Products conversion factors

Product	Conversion unit factor
	m^3 / tonne
Sawnwood, Hardwood	1.43
Logs, Hardwood	1.25
Logs, Softwood	1.43
Sawnwood, Softwood	1.82
Plywood	1.54
	lb / tonne
Beef	0.000453592
Lamb	0.000453592
Swine Meat	0.000453592
Poultry	0.000453592
Coffee	0.000453592
Cotton	0.000453592
Rubber	0.000453592
Hides	0.000453592
Pepper	0.000453592
Sugar	0.000453592
Uranium	0.000453592
	troy oz / tonne
Gold	0.000031103477
Silver	0.000031103477
Platinum	0.000031103477
	kg / tonne
Salmon	0.001
Shrimp	0.001
Tea	0.001
Fine, Wool	0.001
Coarse, Wool	0.001
	bbl / tonne
Crude oil	7.352941176
	mmbtu / tonne
Gaseous state, Natural Gas	40.2
Liquefied, Natural Gas	40.2
Currency	**Conversion currency factor**
US cents / US dollars	0.01

Table A.2 Products Conversion Factors

A.4 Country's trade structure and regional affiliation

M49 code	Country	ISO-alpha 3 code	Sub-region name	Sub-region code	1995–2010 value of indexed commodities (US$m)	1995–2010 value of total exports (US$m)	1995–2010 indexed commodities as a share of total exports	1995–2010 total GDP (US$m)	1995–2010 total exports as a share of GDP
4	Afghanistan	AF/AFG	Southern Asia	34	267.840	4104.450	0.070	78250.390	0.050
8	Albania	AL/ALB	Southern Europe	39	854.780	9577.470	0.090	104115.600	0.090
12	Algeria	DZ/DZA	Northern Africa	15	406617.400	517926.500	0.790	1377817.000	0.380
16	American Samoa	AS/ASM	Polynesia	9	159.420	6009.300	0.030	4889.000	1.230
20	Andorra	AD/AND	Southern Europe	39	52.220	1408.310	0.040	37722.640	0.040
24	Angola	AO/AGO	Middle Africa	17	306778.000	322656.500	0.950	469818.600	0.690
660	Anguilla	AI/AIA	Caribbean	29	6.610	115.830	0.060	No data	No data
28	Antigua and Barbuda	AG/ATG	Caribbean	29	38.800	1048.110	0.040	14772.540	0.070
32	Argentina	AR/ARG	South America	5	218425.700	600441.900	0.360	4185368.000	0.140
51	Armenia	AM/ARM	Western Asia	145	996.420	9592.050	0.100	71295.520	0.130
533	Aruba	AW/ABW	Caribbean	29	533.300	41291.450	0.010	32739.760	1.260
36	Australia	AU/AUS	Australia and New Zealand	9	865183.700	1558556.000	0.560	9684881.000	0.160
40	Austria	AT/AUT	Western Europe	155	74361.510	1552599.000	0.050	4541670.000	0.340
31	Azerbaijan	AZ/AZE	Western Asia	145	125728.300	153280.600	0.820	265717.900	0.580
44	Bahamas	BS/BHS	Caribbean	29	101.580	8003.550	0.010	105186.400	0.080
48	Bahrain	BH/BHR	Western Asia	145	12665.130	133331.300	0.090	213567.700	0.620
50	Bangladesh	BD/BGD	Southern Asia	34	9048.390	140798.600	0.060	1051492.000	0.130
52	Barbados	BB/BRB	Caribbean	29	724.390	5240.450	0.140	55696.800	0.090
112	Belarus	BY/BLR	Eastern Europe	151	31401.060	216384.700	0.150	430512.400	0.500
56	Belgium	BE/BEL	Western Europe	155	228809.700	4430448.000	0.050	5526768.000	0.800
84	Belize	BZ/BLZ	Central America	13	1789.850	4613.670	0.390	15747.230	0.290
204	Benin	BJ/BEN	Western Africa	11	3518.300	10674.260	0.330	65790.530	0.160
60	Bermuda	BM/BMU	Northern America	21	16.070	709.210	0.020	67721.510	0.010
64	Bhutan	BT/BTN	Southern Asia	34	142.100	4188.750	0.030	11579.960	0.360
68	Bolivia (Plurinational	BO/BOL	South America	5	25187.200	45346.480	0.560	167906.300	0.270

	State of)								
70	Bosnia Herzegovina	BA/BIH	Southern Europe	39	4853.060	32577.990	0.150	147160.400	0.220
72	Botswana	BW/BWA	Southern Africa	18	1842.340	54684.350	0.030	123055.100	0.440
92	British Virgin Islands	VG/VGB	Caribbean	29	9.730	481.700	0.020	No data	No data
76	Brazil	BR/BRA	South America	5	591077.600	1559929.000	0.380	15900844.000	0.100
96	Brunei Darussalam	BN/BRN	South-eastern Asia	35	53930.830	80674.610	0.670	127656.000	0.630
100	Bulgaria	BG/BGR	Eastern Europe	151	24739.730	161568.200	0.150	419259.700	0.390
854	Burkina Faso	BF/BFA	Western Africa	11	4323.660	7625.040	0.570	74701.430	0.100
108	Burundi	BI/BDI	Eastern Africa	14	1333.860	1638.330	0.810	17941.590	0.090
132	Cabo Verde	CV/CPV	Western Africa	11	3.670	1093.970	0.000	14814.840	0.070
116	Cambodia	KH/KHM	South-eastern Asia	35	1006.920	39416.250	0.030	93150.040	0.420
120	Cameroon	CM/CMR	Middle Africa	17	30108.770	41981.860	0.720	232918.200	0.180
124	Canada	CA/CAN	Northern America	21	1282382.000	4769257.000	0.270	15828999.000	0.300
136	Cayman Islands	KY/CYM	Caribbean	29	2.660	343.800	0.010	4219.480	0.080
140	Central African Republic	CF/CAF	Middle Africa	17	569.580	2363.170	0.240	20737.270	0.110
148	Chad	TD/TCD	Middle Africa	17	19836.950	25241.680	0.790	73081.650	0.350
152	Chile	CL/CHL	South America	5	246465.800	548491.500	0.450	1808546.000	0.300
156	China	CN/CHN	Eastern Asia	30	241404.200	9894861.000	0.020	36718925.000	0.270
344	China, Hong Kong SAR	HK/HKG	Eastern Asia	30	106256.400	4046800.000	0.030	2902950.000	1.390
446	China, Macao SAR	MO/MAC	Eastern Asia	30	251.970	34729.220	0.010	189936.800	0.180
170	Colombia	CO/COL	South America	5	156247.200	307886.300	0.510	2270135.000	0.140
174	Comoros	KM/COM	Eastern Africa	14	2.210	210.580	0.010	5295.570	0.040
178	Congo	CG/COG	Middle Africa	17	51894.250	64465.710	0.800	84134.860	0.770
184	Cook Islands	CK/COK	Polynesia	9	0.080	80.190	0.000	No data	No data
188	Costa Rica	CR/CRI	Central America	13	17688.220	101371.700	0.170	314330.600	0.320
384	Côte d'Ivoire	CI/CIV	Western Africa	11	47527.510	98860.250	0.480	254584.100	0.390
191	Croatia	HR/HRV	Southern Europe	39	9670.940	118453.800	0.080	615694.600	0.190
192	Cuba	CU/CUB	Caribbean	29	3491.440	37572.950	0.090	646061.600	0.060
196	Cyprus	CY/CYP	Western Asia	145	1215.610	19857.840	0.060	257185.500	0.080
203	Czech Republic	CZ/CZE	Eastern Europe	151	36206.720	1028042.000	0.040	1875821.000	0.550

408	Democratic People's Republic of Korea	KP/PRK	Eastern Asia	30	2362.650	20238.000	0.120	No data	No data
180	Democratic Republic of the Congo	CD/COD	Middle Africa	17	3821.490	34311.500	0.110	183539.100	0.190
208	Denmark	DK/DNK	Northern Europe	154	151068.400	1118018.000	0.140	3741562.000	0.300
262	Djibouti	DJ/DJI	Eastern Africa	14	91.880	966.880	0.100	11049.890	0.090
212	Dominica	DM/DMA	Caribbean	29	173.080	711.810	0.240	5597.170	0.130
214	Dominican Republic	DO/DOM	Caribbean	29	4738.180	88159.320	0.050	480488.700	0.180
218	Ecuador	EC/ECU	South America	5	99623.600	137832.100	0.720	598972.700	0.230
818	Egypt	EG/EGY	Northern Africa	15	54530.820	175643.300	0.310	1727194.000	0.100
222	El Salvador	SV/SLV	Central America	13	5231.610	50239.130	0.100	246782.000	0.200
226	Equatorial Guinea	GQ/GNQ	Middle Africa	17	60768.690	74150.350	0.820	95469.700	0.780
232	Eritrea	ER/ERI	Eastern Africa	14	57.810	477.380	0.120	16541.340	0.030
233	Estonia	EE/EST	Northern Europe	154	10903.520	104430.700	0.100	183220.300	0.570
231	Ethiopia	ET/ETH	Eastern Africa	14	6462.070	13742.470	0.470	220227.300	0.060
234	Faroe Islands	FO/FRO	Northern Europe	154	1837.790	9267.840	0.200	21838.700	0.420
238	Falkland Islands (Malvinas)	FK/FLK	South America	5	1.150	1809.800	0.000	No data	No data
242	Fiji	FJ/FJI	Melanesia	9	1880.580	10629.810	0.180	39071.750	0.270
246	Finland	FI/FIN	Northern Europe	154	81504.250	916835.400	0.090	2883891.000	0.320
251	France	FR/FRA	Western Europe	155	306149.400	6132932.000	0.050	31375045.000	0.200
258	French Polynesia	PF/PYF	Polynesia	9	49.510	3240.610	0.020	22523.850	0.140
583	FS Micronesia	FM/FSM	Micronesia	9	71.200	355.810	0.200	3879.230	0.090
266	Gabon	GA/GAB	Middle Africa	17	60504.230	69202.010	0.870	128761.000	0.540
270	Gambia	GM/GMB	Western Africa	11	128.140	338.050	0.380	12104.210	0.030
268	Georgia	GE/GEO	Western Asia	145	1492.330	10463.030	0.140	94031.690	0.110
276	Germany	DE/DEU	Western Europe	155	349760.800	13337008.000	0.030	42953150.000	0.310
288	Ghana	GH/GHA	Western Africa	11	28578.660	49322.090	0.580	211049.000	0.230
292	Gibraltar	GI/GIB	Southern Europe	39	48.150	2842.560	0.020	No data	No data
300	Greece	GR/GRC	Southern Europe	39	29865.070	261214.100	0.110	3398889.000	0.080
304	Greenland	GL/GRL	Northern America	21	1730.460	5616.230	0.310	24579.250	0.230
308	Grenada	GD/GRD	Caribbean	29	29.460	557.800	0.050	9320.780	0.060

316	Guam	GU/GUM	Micronesia	9	1.150	1087.380	0.000	37884.000	0.030
320	Guatemala	GT/GTM	Central America	13	23402.230	71829.960	0.330	400251.000	0.180
324	Guinea	GN/GIN	Western Africa	11	3349.870	14001.130	0.240	58152.040	0.240
624	Guinea-Bissau	GW/GNB	Western Africa	11	154.460	1142.030	0.140	7820.160	0.150
328	Guyana	GY/GUY	South America	5	5829.120	9422.030	0.620	17378.830	0.540
332	Haiti	HT/HTI	Caribbean	29	26.950	5835.210	0.000	69165.520	0.080
340	Honduras	HN/HND	Central America	13	11286.840	64117.850	0.180	139545.800	0.460
348	Hungary	HU/HUN	Eastern Europe	151	29520.610	800049.500	0.040	1379054.000	0.580
352	Iceland	IS/ISL	Northern Europe	154	14345.350	46240.770	0.310	189790.500	0.240
699	India	IN/IND	Southern Asia	34	147203.300	1388173.000	0.110	11868186.000	0.120
360	Indonesia	ID/IDN	South-eastern Asia	35	543433.500	1266505.000	0.430	4780994.000	0.260
364	Iran	IR/IRN	Southern Asia	34	600866.500	770847.500	0.780	3337131.000	0.230
368	Iraq	IQ/IRQ	Western Asia	145	341570.800	353486.000	0.970	622354.400	0.570
372	Ireland	IE/IRL	Northern Europe	154	53757.210	1426851.000	0.040	2557779.000	0.560
376	Israel	IL/ISR	Western Asia	145	6596.750	582926.600	0.010	2339244.000	0.250
381	Italy	IT/ITA	Southern Europe	39	119018.800	5274698.000	0.020	25871490.000	0.200
388	Jamaica	JM/JAM	Caribbean	29	2385.940	23796.910	0.100	159353.800	0.150
392	Japan	JP/JPN	Eastern Asia	30	81162.870	8505698.000	0.010	75629579.000	0.110
400	Jordan	JO/JOR	Western Asia	145	6029.530	59292.530	0.100	202562.800	0.290
398	Kazakhstan	KZ/KAZ	Central Asia	143	287590.500	374138.000	0.770	881423.300	0.420
404	Kenya	KE/KEN	Eastern Africa	14	14120.900	45846.530	0.310	320474.900	0.140
296	Kiribati	KI/KIR	Micronesia	9	34.820	88.290	0.390	1494.210	0.060
414	Kuwait	KW/KWT	Western Asia	145	351231.300	529476.300	0.660	1028865.000	0.510
417	Kyrgyzstan	KG/KGZ	Central Asia	143	5152.660	13303.480	0.390	40503.830	0.330
418	Lao People's Democratic Republic	LA/LAO	South-eastern Asia	35	191.880	9571.090	0.020	46582.570	0.210
428	Latvia	LV/LVA	Northern Europe	154	13719.260	65771.950	0.210	239742.600	0.270
422	Lebanon	LB/LBN	Western Asia	145	5887.740	32816.260	0.180	340540.100	0.100
426	Lesotho	LS/LSO	Southern Africa	18	140.830	7938.820	0.020	21375.020	0.370
430	Liberia	LR/LBR	Western Africa	11	184.210	4722.230	0.040	9066.210	0.520
434	Libya	LY/LBY	Northern Africa	15	311446.700	374532.500	0.830	694358.500	0.540
440	Lithuania	LT/LTU	Northern Europe	154	14814.020	151420.400	0.100	346603.500	0.440

442	Luxembourg	LU/LUX	Western Europe	155	7001.700	223870.800	0.030	527019.200	0.420
450	Madagascar	MG/MDG	Eastern Africa	14	2241.790	14114.060	0.160	85390.800	0.170
454	Malawi	MW/MWI	Eastern Africa	14	8419.280	9824.340	0.860	54066.500	0.180
458	Malaysia	MY/MYS	South-eastern Asia	35	355170.000	1933171.000	0.180	2151114.000	0.900
462	Maldives	MV/MDV	Southern Asia	34	27.510	2422.720	0.010	18047.250	0.130
466	Mali	ML/MLI	Western Africa	11	11923.030	16299.900	0.730	86902.070	0.190
470	Malta	MT/MLT	Southern Europe	39	161.310	39230.320	0.000	90892.680	0.430
584	Marshall Islands	MH/MHL	Micronesia	9	3.840	277.200	0.010	2065.410	0.130
478	Mauritania	MR/MRT	Western Africa	11	8961.770	12245.930	0.730	34925.290	0.350
480	Mauritius	MU/MUS	Eastern Africa	14	5148.420	29868.190	0.170	97578.050	0.310
484	Mexico	MX/MEX	Central America	13	403994.600	2932426.000	0.140	11859135.000	0.250
496	Mongolia	MN/MNG	Eastern Asia	30	3276.500	17222.660	0.190	41118.560	0.420
499	Montenegro	ME/MNE	Southern Europe	39	1392.080	2623.480	0.530	28631.880	0.090
500	Montserrat	MS/MSR	Caribbean	29	4.710	37.280	0.130	No data	No data
504	Morocco	MA/MAR	Northern Africa	15	27361.240	167041.400	0.160	925668.100	0.180
508	Mozambique	MZ/MOZ	Eastern Africa	14	12461.090	19908.760	0.630	105656.200	0.190
104	Myanmar	MM/MMR	South-eastern Asia	35	5667.430	54014.470	0.100	208176.500	0.260
580	Northern Mariana Islands	MP/MNP	Micronesia	9	19.370	10080.150	0.000	9255.000	1.090
516	Namibia	NA/NAM	Southern Africa	18	5801.010	41084.470	0.140	94616.860	0.430
520	Nauru	NR/NRU	Micronesia	9	8.800	427.080	0.020	153.310	2.790
524	Nepal	NP/NPL	Southern Asia	34	160.750	11181.660	0.010	123790.600	0.090
530	Netherlands Antilles	AN/ANT	Caribbean	29	777.240	19418.660	0.040	No data	No data
528	Netherlands	NL/NLD	Western Europe	155	378555.400	5454727.000	0.070	9581183.000	0.570
540	New Caledonia	NC/NCL	Melanesia	9	651.560	15581.620	0.040	19425.050	0.800
554	New Zealand	NZ/NZL	Australia and New Zealand	9	140851.500	302538.600	0.470	1445243.000	0.210
558	Nicaragua	NI/NIC	Central America	13	8908.930	21964.140	0.410	95165.150	0.230
562	Niger	NE/NER	Western Africa	11	1237.320	7965.710	0.160	49369.910	0.160
566	Nigeria	NG/NGA	Western Africa	11	542026.000	592863.900	0.910	1643113.000	0.360
570	Niue	NU/NIU	Polynesia	9	0.010	5.890	0.000	No data	No data
807	North Macedonia	MK/MKD	Southern Europe	39	2350.910	30514.830	0.080	92155.800	0.330

579	Norway	NO/NOR	Northern Europe	154	855954.200	1335085.000	0.640	4159086.000	0.320
512	Oman	OM/OMN	Western Asia	145	198302.000	259051.200	0.770	458531.300	0.560
586	Pakistan	PK/PAK	Southern Asia	34	20426.310	206803.100	0.100	1626136.000	0.130
585	Palau	PW/PLW	Micronesia	9	0.520	183.600	0.000	2507.760	0.070
591	Panama	PA/PAN	Central America	13	6993.240	62902.240	0.110	255842.100	0.250
598	Papua New Guinea	PG/PNG	Melanesia	9	14876.930	50695.530	0.290	81616.200	0.620
600	Paraguay	PY/PRY	South America	5	24069.220	52590.050	0.460	170678.900	0.310
604	Peru	PE/PER	South America	5	113093.400	235545.600	0.480	1212231.000	0.190
608	Philippines	PH/PHL	South-eastern Asia	35	36461.130	587240.300	0.060	1725203.000	0.340
616	Poland	PL/POL	Eastern Europe	151	80495.810	1164949.000	0.070	4372156.000	0.270
620	Portugal	PT/PRT	Southern Europe	39	24796.820	556058.900	0.040	2727639.000	0.200
634	Qatar	QA/QAT	Western Asia	145	329506.200	383256.400	0.860	684389.100	0.560
410	Republic of Korea	KR/KOR	Eastern Asia	30	69411.900	3833002.000	0.020	11751596.000	0.330
498	Republic of Moldova	MD/MDA	Eastern Europe	151	2233.860	14871.070	0.150	45302.090	0.330
642	Romania	RO/ROU	Eastern Europe	151	20295.690	357863.600	0.060	1387627.000	0.260
643	Russian Federation	RU/RUS	Eastern Europe	151	1654293.000	3086484.000	0.540	11053880.000	0.280
646	Rwanda	RW/RWA	Eastern Africa	14	977.000	2215.190	0.440	42928.920	0.050
654	Saint Helena	SH/SHN	Western Africa	11	8.150	278.050	0.030	No data	No data
659	Saint Kitts and Nevis	KN/KNA	Caribbean	29	86.980	549.940	0.160	8084.330	0.070
662	Saint Lucia	LC/LCA	Caribbean	29	414.980	1548.860	0.270	14260.610	0.110
666	Saint Pierre and Miquelon	PM/SPM	Northern America	21	0.100	103.580	0.000	No data	No data
670	Saint Vincent and the Grenadines	VC/VCT	Caribbean	29	307.210	732.300	0.420	7916.650	0.090
882	Samoa	WS/WSM	Polynesia	9	30.680	878.490	0.030	6248.200	0.140
678	Sao Tome and Principe	ST/STP	Middle Africa	17	47.960	95.510	0.500	1334.230	0.070
682	Saudi Arabia	SA/SAU	Western Asia	145	1595659.000	2071440.000	0.770	4412696.000	0.470
686	Senegal	SN/SEN	Western Africa	11	2945.450	21318.920	0.140	123098.000	0.170
688	Serbia	RS/SRB	Southern Europe	39	4577.650	48846.050	0.090	408000.800	0.120
891	Serbia and Montenegro	CS/SCG	Southern Europe	39	3967.000	44287.310	0.090	No data	No data
690	Seychelles	SC/SYC	Eastern Africa	14	104.350	4105.900	0.030	12038.910	0.340

694	Sierra Leone	SL/SLE	Western Africa	11	61.170	1863.540	0.030	23045.010	0.080
702	Singapore	SG/SGP	South-eastern Asia	35	59790.350	3096813.000	0.020	2020952.000	1.530
703	Slovakia	SK/SVK	Eastern Europe	151	19858.500	456367.500	0.040	838702.700	0.540
705	Slovenia	SI/SVN	Southern Europe	39	7941.550	258063.600	0.030	515164.100	0.500
90	Solomon Islands	SB/SLB	Melanesia	9	956.080	2088.670	0.460	7755.220	0.270
706	Somalia	SO/SOM	Eastern Africa	14	161.540	4295.000	0.040	No data	No data
710	South Africa	ZA/ZAF	Southern Africa	18	161246.800	662503.700	0.240	3294154.000	0.200
724	Spain	ES/ESP	Southern Europe	39	153194.500	2621532.000	0.060	15462620.000	0.170
144	Sri Lanka	LK/LKA	Southern Asia	34	13498.260	91088.630	0.150	386170.600	0.240
275	State of Palestine	PS/PSE	Western Asia	145	248.840	6275.050	0.040	No data	No data
729	Sudan	SD/SDN	Northern Africa	15	47895.850	65572.250	0.730	417366.200	0.160
740	Suriname	SR/SUR	South America	5	715.810	13857.170	0.050	28937.500	0.480
748	Swaziland	SZ/SWZ	Southern Africa	18	2185.340	22027.940	0.100	39075.870	0.560
752	Sweden	SE/SWE	Northern Europe	154	119688.500	1795022.000	0.070	5559633.000	0.320
757	Switzerland	CH/CHE	Western Europe	155	57030.650	1899256.000	0.030	6126207.000	0.310
760	Syria	SY/SYR	Western Asia	145	60283.120	198549.500	0.300	282286.100	0.700
762	Tajikistan	TJ/TJK	Central Asia	143	631.850	14648.760	0.040	37041.720	0.400
764	Thailand	TH/THA	South-eastern Asia	35	185734.700	1578802.000	0.120	3037222.000	0.520
626	Timor-Leste	TL/TLS	South-eastern Asia	35	33.810	211.800	0.160	6160.000	0.030
768	Togo	TG/TGO	Western Africa	11	2316.050	9016.660	0.260	31489.990	0.290
772	Tokelau	TK/TKL	Polynesia	9	0.320	9.450	0.030	No data	No data
776	Tonga	TO/TON	Polynesia	9	1.010	170.980	0.010	3911.410	0.040
780	Trinidad and Tobago	TT/TTO	Caribbean	29	43536.640	113458.400	0.380	205449.500	0.550
788	Tunisia	TN/TUN	Northern Africa	15	30338.470	152663.700	0.200	466351.400	0.330
792	Turkey	TR/TUR	Western Asia	145	34483.610	943650.500	0.040	6411247.000	0.150
795	Turkmenistan	TM/TKM	Central Asia	143	5129.890	66375.190	0.080	129200.400	0.510
796	Turks and Caicos Islands	TC/TCA	Caribbean	29	4.830	184.700	0.030	No data	No data
798	Tuvalu	TV/TUV	Polynesia	9	0.030	1.980	0.010	305.650	0.010
800	Uganda	UG/UGA	Eastern Africa	14	6640.130	13388.620	0.500	150013.900	0.090
804	Ukraine	UA/UKR	Eastern Europe	151	74748.690	450619.600	0.170	1212913.000	0.370
784	United Arab Emirates	AE/ARE	Western Asia	145	775086.400	1530031.000	0.510	2487462.000	0.620

826	United Kingdom	GB/GBR	Northern Europe	154	448306.500	5399046.000	0.080	32871255.000	0.160
834	United Republic of Tanzania	TZ/TZA	Eastern Africa	14	11399.550	24247.740	0.470	238730.400	0.100
842	United States of America	US/USA	Northern America	21	1077078.000	13751191.000	0.080	184000000.000	0.070
858	Uruguay	UY/URY	South America	5	23730.07	53598.81	0.44	358877.7	0.15
860	Uzbekistan	UZ/UZB	Central Asia	143	535.120	85060.010	0.010	287640.400	0.300
548	Vanuatu	VU/VUT	Melanesia	9	96.350	574.010	0.170	6016.870	0.100
862	Venezuela (Bolivarian Republic of)	VE/VEN	South America	5	451595.200	650393.800	0.690	2545618.000	0.260
704	Viet Nam	VN/VNM	South-eastern Asia	35	144108.000	448496.700	0.320	849647.100	0.530
876	Wallis and Futuna Islands	WF/WLF	Polynesia	9	0.000	5.900	0.000	No data	No data
732	Western Sahara	EH/ESH	Northern Africa	15	0.690	124.010	0.010	No data	No data
887	Yemen	YE/YEM	Western Asia	145	60311.670	70126.010	0.860	227124.700	0.310
894	Zambia	ZM/ZMB	Eastern Africa	14	22198.030	37400.740	0.590	130311.900	0.290
716	Zimbabwe	ZW/ZWE	Eastern Africa	14	14488.870	39176.940	0.370	108122.200	0.360

Source: Datastream (2018), UN Comtrade (2018) and UNCTAD (2018)

Table A.3 Country's Trade Structure and Regional Affiliation

A.5 Description of the index series and their relevant sources

Index	Description	Source
NCOMPI_all_Aus	This is a nominal COMPI index, i.e. NCOMPI_all, for Australia taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The base year is 1995.	Author's calculations
NCOMPI_all_Can	This is a nominal COMPI index, i.e. NCOMPI_all, for Canada taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017.	Author's calculations
NCOMPI_all_NZ	This is a nominal COMPI index, i.e. NCOMPI_all, for New Zealand taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The base year is 1995.	Author's calculations
NCOMPI_m_Aus	This is a nominal metals index, i.e. NCOMPI_m, for Australia taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The base year is 1995.	Author's calculations
NCOMPI_m_Can	This is a nominal metals index, i.e. NCOMPI_m, for Canada taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The base year is 1995.	Author's calculations
NCOMPI_m_NZ	This is a nominal metals index, i.e. NCOMPI_m, for New Zealand taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The base year is 1995.	Author's calculations
NCOMPI_ne_Aus	This is a nominal non-energy index, i.e. NCOMPI_ne, for Australia taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The base year is 1995.	Author's calculations
NCOMPI_ne_Can	This is a nominal non-energy index, i.e. NCOMPI_ne, for Canada taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The	Author's calculations

	base year is 1995.	
NCOMPI_ne_NZ	This is a nominal non-energy index, i.e. NCOMPI_ne, for New Zealand taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The base year is 1995.	Author's calculations
NCOMPI_energy_Can	This is a nominal energy index, i.e. NCOMPI_energy, for Canada taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The base year is 1995.	Author's calculations
NCOMPI_food_Can	This is a nominal food index, i.e. NCOMPI_food, for Canada taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The base year is 1995.	Author's calculations
NCOMPI_dairy_NZ	This is a nominal energy index, i.e. NCOMPI_dairy, for Canada taken from my constructed database. The structure of its commodity basket is given above. The index starts at January 1980 and ends at April 2017. The base year is 1995.	Author's calculations
Official_all_Aus	This is an "Index of commodity prices; All items; US$", i.e. Official_all, for Australia as reported by Reserve Bank of Australia. It is in nominal terms. The index starts at July 1982 and ends at April 2017. The index is rebased to 1995=100 for a consistency reasons.	Reserve Bank of Australia (2018)
Official_all_Can	This is a "Monthly Bank of Canada commodity price index – Total", i.e. Official_all, for Canada as reported by Bank of Canada. It is in nominal terms. The index starts at January 1980 and ends at April 2017. The index is rebased to 1995=100 for a consistency reasons.	Bank of Canada (2018)
Official_all_NZ	This is a "World Price Index Expressed in US$", i.e. Official_all, for New Zealand as reported by ANZ Bank. It is in nominal terms. The index starts at January 1986 and ends at April 2017. The index is rebased to 1995=100 for a consistency reasons.	Australian and New Zealand Banking Group (2018)
Official_m_Aus	This is an "Index of commodity prices; Non-rural component – Base metals; US$", i.e. Official_m, for Australia as reported by Reserve Bank of Australia. It is in nominal terms. The index starts at July 1982 and ends at April 2017. The index is rebased to 1995=100 for a	Reserve Bank of Australia (2018)

	consistency reasons.	
Official_m_Can	This is a “Monthly Bank of Canada commodity price index - Metals and Minerals”, i.e. Official_m, for Canada as reported by Bank of Canada. It is in nominal terms. The index starts at January 1980 and ends at April 2017. The index is rebased to 1995=100 for a consistency reasons.	Bank of Canada (2018)
Official_m_NZ	This is an “Aluminium” index, i.e. Official_m, for New Zealand as reported by ANZ Bank. It is in nominal terms. The index starts at January 1986 and ends at April 2017. The index is rebased to 1995=100 for a consistency reasons.	Australian and New Zealand Banking Group (2018)
Official_ne_Can	This is a “Monthly Bank of Canada commodity price index - Excluding Energy”, i.e. Official_ne, for Canada as reported by Bank of Canada. It is in nominal terms. The index starts at January 1980 and ends at April 2017. The index is rebased to 1995=100 for a consistency reasons.	Bank of Canada (2018)
Official_energy_Can	This is a “Monthly Bank of Canada commodity price index – Energy”, i.e. Official_energy, for Canada as reported by Bank of Canada. It is in nominal terms. The index starts at January 1980 and ends at April 2017. The index is rebased to 1995=100 for a consistency reasons.	Bank of Canada (2018)
Official_food_Can	This is a “Monthly Bank of Canada commodity price index – Agriculture”, i.e. Official_food, for Canada as reported by Bank of Canada. It is in nominal terms. The index starts at January 1980 and ends at April 2017. The index is rebased to 1995=100 for a consistency reasons.	Bank of Canada (2018)
Official_dairy_NZ	This is a “Dairy Products” index, i.e. Official_dairy, for New Zealand as reported by ANZ Bank. It is in nominal terms. The index starts at January 1986 and ends at April 2017. The index is rebased to 1995=100 for a consistency reasons.	Australian and New Zealand Banking Group (2018)
Cashin_official_Aus	This is a nominal non-fuel commodities index, i.e. Cashin_official, for Australia as constructed in Cashin et al. (2004). The data is initially provided by Dr Paul Cashin in real terms; hence, the MUV deflator (base 1995=100) is taken from the IFS for transforming the series in nominal terms. The same deflator is used by Cashin et al. (2004) for obtaining the real values of the index series. The index starts at January 1980 and ends at March 2002. The index is rebased to 1995=100 for a	Cashin et al. (2004)

	consistency reasons. A brief description of the individual commodity prices is available in a longer working paper version of Cashin et al. (2004), which is Sahay et al. (2002). The data is kindly provided by Dr Paul Cashin.	
Cashin_official_Can	This is a nominal non-fuel commodities index, i.e. Cashin_official, for Canada as constructed in Cashin et al. (2004). The data is initially provided by Dr Paul Cashin in real terms; hence, the MUV deflator (base 1995=100) is taken from the IFS for transforming the series in nominal terms. The same deflator is used by Cashin et al. (2004) for obtaining the real values of the index series. The index starts at January 1980 and ends at March 2002. The index is rebased to 1995=100 for a consistency reasons. A brief description of the individual commodity prices is available in a longer working paper version of Cashin et al. (2004), which is Sahay et al. (2002). The data is kindly provided by Dr Paul Cashin.	Cashin et al. (2004)
Cashin_official_NZ	This is a nominal non-fuel commodities index, i.e. Cashin_official, for New Zealand as constructed in Cashin et al. (2004). The data is initially provided by Dr Paul Cashin in real terms; hence, the MUV deflator (base 1995=100) is taken from the IFS for transforming the series in nominal terms. The same deflator is used by Cashin et al. (2004) for obtaining the real values of the index series. The index starts at January 1980 and ends at March 2002. The index is rebased to 1995=100 for a consistency reasons. A brief description of the individual commodity prices is available in a longer working paper version of Cashin et al. (2004), which is Sahay et al. (2002). The data is kindly provided by Dr Paul Cashin.	Cashin et al. (2004)

Table A.4 Description of the Index Series and Their Relevant Sources

Appendix B

B.1 Ranking of the countries by export/import as a share of GDP

M49 code	Country	ISO-alpha3 code	Region (M49 code)	Primary Export in $m of 1995–2010 total of primary commodities, precious stones and non-monetary gold (SITC 0 + 1 + 2 + 3 + 4 + 68 + 667+ 971)	Primary import in $m of 1995–2010 total of primary commodities, precious stones and non-monetary gold (SITC 0 + 1 + 2 + 3 + 4 + 68 + 667+ 971)	1995–2010 total GDP (US$m)	Primary exports as a share of GDP	Primary import as a share of GDP	Data Period (From/To)	
36	Australia	AU/AUS	9	73310	22515	1047457	0.07	0.02	Q1 1980	Q4 2016
48	Bahrain	BH/BHR	145	6058	1873	17834	0.34	0.11	Q4 1981	Q4 2016
56	Belgium	BE/BEL	155	66585	76722	426218	0.16	0.18	Q1 1996	Q4 2016
68	Bolivia	BO/BOL	5	2463	510	14925	0.17	0.03	Q1 1991	Q4 2016
124	Canada	CA/CAN	21	114240	50190	1399178	0.08	0.04	Q1 1980	Q4 2016
152	Chile	CL/CHL	5	29369	8763	166062	0.18	0.05	Q1 1997	Q4 2016
203	Czech Republic	CZ/CZE	151	6941	12563	170436	0.04	0.07	Q1 1997	Q4 2016
208	Denmark	DK/DNK	154	21251	13733	301956	0.07	0.05	Q1 1992	Q4 2016
214	Dominican Republic	DO/DOM	29	1094	2783	36912	0.03	0.08	Q1 1993	Q3 2014
218	Ecuador	EC/ECU	5	7880	2319	54319	0.15	0.04	Q1 2001	Q4 2016
233	Estonia	EE/EST	154	1961	2402	16638	0.12	0.14	Q1 1996	Q4 2016
344	Hong Kong	HK/HKG	30	19317	35339	173381	0.11	0.20	Q1 1980	Q4 2016
348	Hungary	HU/HUN	151	6406	8295	117067	0.06	0.07	Q1 1996	Q4 2016
352	Iceland	IS/ISL	154	2384	953	11395	0.21	0.08	Q1 1998	Q4 2016
376	Israel	IL/ISR	145	13585	13429	179260	0.08	0.08	Q1 1996	Q4 2016
398	Kazakhstan	KZ/KAZ	143	19848	2947	93850	0.21	0.03	Q3 1995	Q4 2016
428	Latvia	LV/LVA	154	1568	1988	20299	0.08	0.10	Q1 1996	Q4 2016
442	Luxembourg	LU/LUX	155	1909	4265	43126	0.04	0.10	Q1 2001	Q4 2016
470	Malta	MT/MLT	39	441	909	7349	0.06	0.12	Q1 2001	Q4 2016
528	Netherlands	NL/NLD	155	106132	94684	746501	0.14	0.13	Q1 1997	Q4 2016
554	New Zealand	NZ/NZL	9	13010	4949	123462	0.11	0.04	Q2 1988	Q4 2016
579	Norway	NO/NOR	154	63337	10357	382844	0.17	0.03	Q1 1980	Q4 2016
604	Peru	PE/PER	5	12738	3901	101282	0.13	0.04	Q1 1981	Q4 2016
608	Philippines	PH/PHL	35	4621	10892	144665	0.03	0.08	Q1 1982	Q4 2016
688	Serbia	RS/SRB	39	3418	5684	45368	0.08	0.13	Q1 1997	Q4 2016
690	Seychelles	SC/SYC	14	228	237	792	0.29	0.30	Q1 1980	Q4 2016
702	Singapore	SG/SGP	35	32357	43219	155548	0.21	0.28	Q1 1980	Q4 2016

703	Slovakia	SK/SVK	151	3918	7010	66215	0.06	0.11	Q1 1996	Q4 2016
705	Slovenia	SI/SVN	39	1761	4232	40468	0.04	0.11	Q1 1996	Q4 2016
710	South Africa	ZA/ZAF	18	21424	12781	299152	0.07	0.04	Q1 1980	Q4 2016
764	Thailand	TH/THA	35	25631	28310	258525	0.10	0.11	Q1 1994	Q4 2016
862	Venezuela	VE/VEN	5	36058	4192	320459	0.11	0.01	Q1 1998	Q4 2016
704	Viet Nam	VN/VNM	35	12632	9392	74965	0.17	0.13	Q1 2002	Q4 2016

Source: UN Comtrade (2018) and Datastream (2018)

Table B.1 Ranking of the Countries by Export/Import as a Share of GDP

B.2 Results for unit root tests and other preliminary sample statistics

The stationarity is a requirement that needs to be fulfilled. Therefore, we perform ADF and PP unit root tests for all the time series, as described in Section 3.5. We consider unit root tests with an intercept and without trend, similar to Narayan et al. (2014). The lags number is selected using BIC for the ADF test. In terms of the PP test, the lag number is selected by Bartlett kernel with the Newey-West automatic bandwidth selection. The results from the ADF and PP unit root tests are reported in Appendices B.2–B.6.

Commodity prices: When the ADF (PP) test equation has an intercept and no trend, the null hypothesis of unit root is rejected at the 10% level for all high-frequency commodity price series. In the case of low-frequency series, when the ADF test equation has an intercept and no trend, the null hypothesis is not rejected at the 10% level for Thailand while being rejected at the 10% level for all other low-frequency commodity price series. Otherwise, for the PP test equation with an intercept and no trend, the null hypothesis is rejected at the 10% level for all low-frequency commodity price series. Therefore, we conclude that the condition of stationarity is satisfied for all commodity price series based on the results from ADF and PP tests.

Economic growth: When the ADF test equation has an intercept and no trend, the null hypothesis of unit root cannot be rejected at the 10% level for nine economic growth series, namely Belgium, Bolivia, Denmark, Ecuador, Estonia, Hungary, Kazakhstan and Philippines. In contrast, when the PP test equation has an intercept and no trend, the null hypothesis of unit root cannot be rejected at the 10% level for only two economic growth series, namely Czech Republic and Netherlands. Based on the results from both ADF and PP tests, as shown in Appendices B.2–B.6, we conclude that stationarity is satisfied for all the economic growth series.

	Economic Growth	
	ADF with intercept, no trend	PP with intercept, no trend
Australia	-4.360***	-3.117**
Bahrain	-2.897**	-4.522***
Belgium	-2.402	-3.116**
Bolivia	-1.882	-5.473***
Canada	-3.56***	-3.750***
Chile	-4.089***	-3.271**
Czech Republic	-3.948***	-2.442
Denmark	-2.192	-3.438**
Dominican Republic	-4.014***	-4.132***
Ecuador	-2.264	-2.739*
Estonia	-2.574	-2.861*
Hong Kong	-2.719*	-4.690***
Hungary	-2.277	-2.670*
Iceland	-3.326**	-3.571***
Israel	-3.676***	-3.432**
Kazakhstan	-1.898	-5.390***
Latvia	-4.225***	-2.683*
Luxembourg	-2.707*	-3.164**
Malta	-3.680***	-3.615***
Netherlands	-3.117**	-2.565
New Zealand	-2.690*	-4.296***
Norway	-2.593*	-5.202***
Peru	-3.108**	-3.455**
Philippines	-2.441	-2.880*
Serbia	-2.658*	-4.609***
Seychelles	-3.546***	-2.963**
Singapore	-3.317**	-3.191**
Slovakia	-3.176**	-3.339**
Slovenia	-2.771*	-2.636*
South Africa	-2.847*	-3.061**
Thailand	-3.516***	-3.826***
Venezuela	-4.298***	-3.311**
Viet Nam	-6.464***	-6.461***

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.2 Results for Unit Root Tests for Economic Growth

	Commodity Prices: Low-Frequency	
	ADF with intercept, no trend	PP with intercept, no trend
Australia	-3.668***	-4.387***
Bahrain	-3.678***	-4.297***
Belgium	-4.221***	-3.719***
Bolivia	-4.627***	-3.127**
Canada	-3.668***	-4.387***
Chile	-3.865***	-3.602***
Czech Republic	-3.865***	-3.602***
Denmark	-4.501***	-4.123***
Dominican Republic	-4.197***	-3.797***
Ecuador	-4.184***	-3.196**
Estonia	-4.221***	-3.719***
Hong Kong	-3.668***	-4.387***
Hungary	-4.221***	-3.719***
Iceland	-3.944***	-3.028**
Israel	-4.221***	-3.719***
Kazakhstan	-4.162***	-3.766***
Latvia	-4.221***	-3.719***
Luxembourg	-4.184***	-3.196**
Malta	-4.184***	-3.196**
Netherlands	-3.865***	-3.602***
New Zealand	-4.686***	-3.259**
Norway	-3.668***	-4.387***
Peru	-3.668***	-4.387***
Philippines	-3.691***	-4.279***
Serbia	-3.865***	-3.602***
Seychelles	-3.668***	-4.387***
Singapore	-3.668***	-4.387***
Slovakia	-4.221***	-3.719***
Slovenia	-4.221***	-3.719***
South Africa	-3.668***	-4.387***
Thailand	-2.561	-3.918***
Venezuela	-3.799***	-3.166**
Viet Nam	-4.336***	-3.053**

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.3 Results for Unit Root Tests for Commodity Prices at Low-Frequency

	Commodity Prices: Mixed-Frequency					
	ADF with intercept, no trend			PP with intercept, no trend		
	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	-6.256***	-5.209***	-5.253***	-4.303***	-4.495***	-4.199***
Bahrain	-6.305***	-5.193***	-5.250***	-4.213***	-4.410***	-3.955***
Belgium	-4.233***	-3.970***	-4.182***	-3.460**	-4.001***	-4.162***
Bolivia	-5.312***	-4.414***	-4.585***	-3.186**	-3.597***	-3.542***
Canada	-6.256***	-5.209***	-5.253***	-4.303***	-4.495***	-4.199***
Chile	-4.163***	-4.249***	-4.092***	-3.416**	-3.916***	-4.094***
Czech Republic	-4.163***	-4.249***	-4.092***	-3.416**	-3.916***	-4.094***
Denmark	-5.106***	-4.280***	-4.488***	-4.197***	-4.285***	-4.404***
Dominican Republic	-4.441***	-3.978***	-4.161***	-3.856***	-3.963***	-4.278***
Ecuador	-4.076***	-4.551***	-3.917***	-3.623***	-3.152**	-3.613***
Estonia	-4.233***	-3.970***	-4.182***	-3.460**	-4.001***	-4.162***
Hong Kong	-6.256***	-5.209***	-5.253***	-4.303***	-4.495***	-4.199***
Hungary	-4.233***	-3.970***	-4.182***	-3.460**	-4.001***	-4.162***
Iceland	-4.979***	-4.421***	-4.416***	-3.634***	-3.472**	-3.662***
Israel	-4.233***	-3.970***	-4.182***	-3.460**	-4.001***	-4.162***
Kazakhstan	-4.183***	-3.938***	-4.117***	-3.546***	-4.047***	-4.210***
Latvia	-4.233***	-3.970***	-4.182***	-3.460**	-4.001***	-4.162***
Luxembourg	-4.076***	-4.551***	-3.917***	-3.623***	-3.152**	-3.613***
Malta	-4.076***	-4.551***	-3.917***	-3.623***	-3.152**	-3.613***
Netherlands	-4.163***	-4.249***	-4.092***	-3.416**	-3.916***	-4.094***
New Zealand	-5.509***	-4.539***	-4.716***	-3.422**	-3.741***	-3.653***
Norway	-6.256***	-5.209***	-5.253***	-4.303***	-4.495***	-4.199***
Peru	-6.256***	-5.209***	-5.253***	-4.303***	-4.495***	-4.199***
Philippines	-6.282***	-5.159***	-5.245***	-4.251***	-4.335***	-3.922***
Serbia	-4.163***	-4.249***	-4.092***	-3.416**	-3.916***	-4.094***
Seychelles	-6.256***	-5.209***	-5.253***	-4.303***	-4.495***	-4.199***
Singapore	-6.256***	-5.209***	-5.253***	-4.303***	-4.495***	-4.199***
Slovakia	-4.233***	-3.970***	-4.182***	-3.460**	-4.001***	-4.162***
Slovenia	-4.233***	-3.970***	-4.182***	-3.460**	-4.001***	-4.162***
South Africa	-6.256***	-5.209***	-5.253***	-4.303***	-4.495***	-4.199***
Thailand	-4.877***	-3.997***	-2.941**	-4.001***	-4.208***	-4.366***
Venezuela	-4.966***	-4.224***	-4.219***	-3.449**	-3.340**	-3.553***
Viet Nam	-3.949***	-3.722***	-3.810***	-3.430**	-2.908*	-3.283**

Note: The table reports the test statistics obtained from the unit root tests for monthly commodity price series at a given month of each quarter period τ.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.4 Results for Unit-Root Tests for Commodity Prices at Mixed-Frequency

The above unit root tests are supplemented by a test of (trend) stationarity. The test of stationarity is added because some macroeconomic aggregates appear to display long-memory as discussed by Barkoulas (1998) and Lee and Schmidt (1996). In particular, long memory is a phenomenon that may arise in the analysis of time series data. This phenomenon is considered to have long memory if the statistical dependence decays more slowly than an exponential decay. Therefore, we perform a KPSS (1992), hereafter KPSS, test which is a test of trend stationarity in order to underpin our results of non-stationarity/stationarity of the variables before performing MF-VAR. In the KPSS test trend stationarity is the null hypothesis to be tested against the alternative of a unit root. We consider stationarity test with an intercept and without trend, consistent to Narayan et al. (2014), and to allow for direct comparison between different preliminary tests. Similar to PP test, the lag number is selected by Bartlett kernel with the Newey-West automatic bandwidth selection. The results from the KPSS test of trend stationarity are reported in Appendices B.2–B.6.

Commodity prices: When KPSS test equation has an intercept and no trend, the null hypothesis of trend stationarity cannot be rejected at the 10% level for all high-frequency commodity price series. Similarly, in the case of KPSS test equation with an intercept and no trend, the null hypothesis of trend stationarity cannot be rejected at the 10% level for all low-frequency commodity price series. Therefore, we conclude that the condition of stationarity is satisfied for all commodity price series based on the results from KPSS tests, in addition to ADF and PP tests.

Economic growth: When the KPSS test equation has an intercept and no trend, the null hypothesis of trend stationarity is rejected at the 1% level only for the economic growth series of Philippines. For all other countries, economic growth series are found to be stationary within 10% level of significance. Importantly, the PP test suggests stationarity for economic growth series of Philippines at 10% level of significance. Our study consents with the results of the PP test in terms of stationary for economic growth series of Philippines due to the claim of Caner and Kilian (2001) that asymptotic critical values of KPSS test make no distinction between a process that is white noise and a highly persistent stationary process. Moreover, KPSS test has high rate of Type I errors (Das, 2019). Type I error is the rejection of a true null hypothesis which leads to false finding and conclusion (Caner and Kilian, 2001). This may be the case with Philippines. Therefore, we conclude that the condition of stationarity is satisfied for all economic growth series based on the combined results from KPSS, ADF and PP tests (see Perron, 2019).

	KPSS with intercept, no trend				
	Economic Growth (EG)	CP(τ)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	0.128	0.135	0.124	0.140	0.134
Bahrain	0.131	0.153	0.148	0.159	0.144
Belgium	0.525**	0.090	0.093	0.091	0.083
Bolivia	0.509**	0.091	0.096	0.090	0.083
Canada	0.087	0.135	0.124	0.140	0.134
Chile	0.078	0.122	0.127	0.121	0.113
Czech Republic	0.125	0.122	0.127	0.121	0.113
Denmark	0.387*	0.069	0.074	0.070	0.063
Dominican Republic	0.056	0.070	0.073	0.070	0.065
Ecuador	0.215	0.133	0.134	0.131	0.121
Estonia	0.280	0.090	0.093	0.091	0.083
Hong Kong	0.230	0.135	0.124	0.140	0.134
Hungary	0.277	0.090	0.093	0.091	0.083
Iceland	0.153	0.167	0.171	0.168	0.155
Israel	0.048	0.090	0.093	0.091	0.083
Kazakhstan	0.167	0.085	0.084	0.087	0.083
Latvia	0.229	0.090	0.093	0.091	0.083
Luxembourg	0.130	0.133	0.134	0.131	0.121
Malta	0.545**	0.133	0.134	0.131	0.121
Netherlands	0.381*	0.122	0.127	0.121	0.113
New Zealand	0.138	0.068	0.067	0.069	0.066
Norway	0.562**	0.135	0.124	0.140	0.134
Peru	0.564**	0.135	0.124	0.140	0.134
Philippines	0.742***	0.143	0.142	0.146	0.132
Serbia	0.206	0.122	0.127	0.121	0.113
Seychelles	0.094	0.135	0.124	0.140	0.134
Singapore	0.269	0.135	0.124	0.140	0.134
Slovakia	0.155	0.090	0.093	0.091	0.083
Slovenia	0.441*	0.090	0.093	0.091	0.083
South Africa	0.536**	0.135	0.124	0.140	0.134
Thailand	0.067	0.069	0.071	0.070	0.064
Venezuela	0.125	0.191	0.209	0.189	0.173
Viet Nam	0.117	0.138	0.143	0.137	0.124

Note: The table reports the test statistics obtained from the KPSS unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.5 Results for KPSS Unit-Root Test

Tables B.6 to B.10 presents the summary statistics for the commodity prices and economic growth series. Specifically, $CP_{(\tau,1)}$, $CP_{(\tau,2)}$ and $CP_{(\tau,3)}$ have some interesting differences. First, with respect to the median, the commodity prices from the last month in each quarter perform worse than those in the first month in each quarter in most cases. Second, $CP_{(\tau,3)}$ has weaker asymmetry than $CP_{(\tau,1)}$ and $CP_{(\tau,2)}$ in terms of skewness. The heterogeneous characteristics of $CP_{(\tau,1)}$, $CP_{(\tau,2)}$ and $CP_{(\tau,3)}$ suggest a potential benefit of the MF-VAR. In particular, a major advantage of MF-VAR relative to single-frequency VAR is that high-frequency variables are allowed to have heterogeneous impacts on a low-frequency variable within each quarterly time period (see Ghysels et al., 2016 for discussion).

In addition, the Kolmogorov-Smirnov test rejects the null hypothesis of normality for economic growth at the 1% level for all countries in our sample. Further, the Anderson-Darling test rejects the null hypothesis of normality for economic growth for 29 out of a total of 33 countries at a 10% significance level. Similar results are confirmed by the Jarque-Bera test statistics. This implies that the economic growth series are likely to have non-normal distributions. While most standard causality models require the assumption of normality to be fulfilled, as demonstrated in the study of Ghysels et al. (2016), the asymptotic theory of MF-VAR models does not require the normality assumption. This is another benefit of using MF-VAR modelling within our empirical framework.

	CP(τ,1)										
	Obs.	Mean	Median	Minimum	Maximum	Std. Dev.	Skewness	Kurtosis	p-KS	p-AD	p-JB
Australia	144	-0.378	0.955	-35.370	29.780	11.059	-0.462	3.660	0.000	0.005	0.027
Bahrain	141	-0.504	0.920	-35.370	29.780	11.117	-0.440	3.631	0.000	0.007	0.033
Belgium	84	0.643	1.800	-26.870	29.780	10.813	-0.240	3.373	0.000	0.112	0.435
Bolivia	104	0.401	1.330	-26.870	29.780	10.058	-0.199	3.650	0.000	0.077	0.198
Canada	144	-0.378	0.955	-35.370	29.780	11.059	-0.462	3.660	0.000	0.005	0.027
Chile	80	0.385	1.330	-26.870	29.780	11.001	-0.177	3.284	0.000	0.259	0.500
Czech Republic	80	0.385	1.330	-26.870	29.780	11.001	-0.177	3.284	0.000	0.259	0.500
Denmark	100	0.713	1.830	-26.870	29.780	10.128	-0.274	3.704	0.000	0.022	0.120
Dominican Republic	86	1.290	2.725	-26.870	29.780	10.512	-0.399	3.723	0.000	0.003	0.078
Ecuador	64	1.808	1.800	-26.870	29.780	10.578	-0.092	3.590	0.000	0.592	0.500
Estonia	84	0.643	1.800	-26.870	29.780	10.813	-0.240	3.373	0.000	0.112	0.435
Hong Kong	144	-0.378	0.955	-35.370	29.780	11.059	-0.462	3.660	0.000	0.005	0.027
Hungary	84	0.643	1.800	-26.870	29.780	10.813	-0.240	3.373	0.000	0.112	0.435
Iceland	76	0.122	0.840	-26.870	29.780	11.214	-0.114	3.186	0.000	0.516	0.500
Israel	84	0.643	1.800	-26.870	29.780	10.813	-0.240	3.373	0.000	0.112	0.435
Kazakhstan	86	0.743	1.940	-26.870	29.780	10.714	-0.265	3.428	0.000	0.069	0.340
Latvia	84	0.643	1.800	-26.870	29.780	10.813	-0.240	3.373	0.000	0.112	0.435
Luxembourg	64	1.808	1.800	-26.870	29.780	10.578	-0.092	3.590	0.000	0.592	0.500
Malta	64	1.808	1.800	-26.870	29.780	10.578	-0.092	3.590	0.000	0.592	0.500
Netherlands	80	0.385	1.330	-26.870	29.780	11.001	-0.177	3.284	0.000	0.259	0.500
New Zealand	115	0.599	1.660	-26.870	29.780	9.731	-0.233	3.808	0.000	0.069	0.083
Norway	144	-0.378	0.955	-35.370	29.780	11.059	-0.462	3.660	0.000	0.005	0.027
Peru	144	-0.378	0.955	-35.370	29.780	11.059	-0.462	3.660	0.000	0.005	0.027
Philippines	140	-0.462	0.955	-35.370	29.780	11.146	-0.450	3.627	0.000	0.005	0.032
Serbia	80	0.385	1.330	-26.870	29.780	11.001	-0.177	3.284	0.000	0.259	0.500
Seychelles	144	-0.378	0.955	-35.370	29.780	11.059	-0.462	3.660	0.000	0.005	0.027
Singapore	144	-0.378	0.955	-35.370	29.780	11.059	-0.462	3.660	0.000	0.005	0.027
Slovakia	84	0.643	1.800	-26.870	29.780	10.813	-0.240	3.373	0.000	0.112	0.435
Slovenia	84	0.643	1.800	-26.870	29.780	10.813	-0.240	3.373	0.000	0.112	0.435
South Korea	144	-0.378	0.955	-35.370	29.780	11.059	-0.462	3.660	0.000	0.005	0.027
Thailand	92	0.942	2.025	-26.870	29.780	10.421	-0.320	3.602	0.000	0.017	0.144
Venezuela	71	0.498	1.660	-26.870	29.780	11.445	-0.187	3.141	0.000	0.368	0.500
Viet Nam	60	1.812	2.370	-26.870	29.780	10.904	-0.091	3.398	0.000	0.788	0.500

Note: The table reports the primary statistics obtained for monthly commodity price series at the first month of each quarter period τ. CP(τ,1) series are log-differenced, as specified in Section 3.5. "p-KS" signifies a p-value of the Kolmogorov-Smirnov test for normality. "p-AD" signifies a p-value of the Anderson-Darling test for normality. "p-JB" signifies a p-value of the Jarque-Bera test for normality.

Table B.6 Sample Statistics for Commodity Prices CP(τ,1), Mixed-Frequency

	CP(τ,2)										
	Obs.	Mean	Median	Minimum	Maximum	Std. Dev.	Skewness	Kurtosis	p-KS	p-AD	p-JB
Australia	144	-0.401	0.250	-35.000	26.120	10.920	-0.579	3.577	0.000	0.015	0.016
Bahrain	141	-0.492	0.210	-35.000	26.120	10.985	-0.563	3.543	0.000	0.022	0.019
Belgium	84	0.688	2.150	-30.320	26.120	10.846	-0.352	3.042	0.000	0.160	0.320
Bolivia	104	0.480	1.615	-30.320	26.120	10.116	-0.319	3.266	0.000	0.206	0.269
Canada	144	-0.401	0.250	-35.000	26.120	10.920	-0.579	3.577	0.000	0.015	0.016
Chile	80	0.459	2.150	-30.320	26.120	11.019	-0.305	2.961	0.000	0.248	0.448
Czech Republic	80	0.459	2.150	-30.320	26.120	11.019	-0.305	2.961	0.000	0.248	0.448
Denmark	100	0.776	2.150	-30.320	26.120	10.190	-0.390	3.320	0.000	0.099	0.148
Dominican Republic	86	1.306	2.995	-30.320	26.120	10.682	-0.504	3.251	0.000	0.023	0.088
Ecuador	64	1.970	3.605	-23.170	26.120	10.333	-0.131	2.792	0.000	0.191	0.500
Estonia	84	0.688	2.150	-30.320	26.120	10.846	-0.352	3.042	0.000	0.160	0.320
Hong Kong	144	-0.401	0.250	-35.000	26.120	10.920	-0.579	3.577	0.000	0.015	0.016
Hungary	84	0.688	2.150	-30.320	26.120	10.846	-0.352	3.042	0.000	0.160	0.320
Iceland	76	0.151	0.755	-30.320	26.120	11.188	-0.241	2.889	0.000	0.397	0.500
Israel	84	0.688	2.150	-30.320	26.120	10.846	-0.352	3.042	0.000	0.160	0.320
Kazakhstan	86	0.719	2.150	-30.320	26.120	10.720	-0.365	3.116	0.000	0.156	0.278
Latvia	84	0.688	2.150	-30.320	26.120	10.846	-0.352	3.042	0.000	0.160	0.320
Luxembourg	64	1.970	3.605	-23.170	26.120	10.333	-0.131	2.792	0.000	0.191	0.500
Malta	64	1.970	3.605	-23.170	26.120	10.333	-0.131	2.792	0.000	0.191	0.500
Netherlands	80	0.459	2.150	-30.320	26.120	11.019	-0.305	2.961	0.000	0.248	0.448
New Zealand	115	0.597	1.610	-30.320	26.120	9.755	-0.342	3.447	0.000	0.183	0.134
Norway	144	-0.401	0.250	-35.000	26.120	10.920	-0.579	3.577	0.000	0.015	0.016
Peru	144	-0.401	0.250	-35.000	26.120	10.920	-0.579	3.577	0.000	0.015	0.016
Philippines	140	-0.414	0.250	-35.000	26.120	10.985	-0.581	3.578	0.000	0.016	0.016
Serbia	80	0.459	2.150	-30.320	26.120	11.019	-0.305	2.961	0.000	0.248	0.448
Seychelles	144	-0.401	0.250	-35.000	26.120	10.920	-0.579	3.577	0.000	0.015	0.016
Singapore	144	-0.401	0.250	-35.000	26.120	10.920	-0.579	3.577	0.000	0.015	0.016
Slovakia	84	0.688	2.150	-30.320	26.120	10.846	-0.352	3.042	0.000	0.160	0.320
Slovenia	84	0.688	2.150	-30.320	26.120	10.846	-0.352	3.042	0.000	0.160	0.320
South Korea	144	-0.401	0.250	-35.000	26.120	10.920	-0.579	3.577	0.000	0.015	0.016
Thailand	92	0.971	2.390	-30.320	26.120	10.451	-0.428	3.267	0.000	0.063	0.134
Venezuela	71	0.346	2.250	-30.320	26.120	11.503	-0.281	2.785	0.000	0.234	0.500
Viet Nam	60	1.930	3.950	-23.170	26.120	10.569	-0.121	2.712	0.000	0.183	0.500

Note: The table reports the primary statistics obtained for monthly commodity price series at the second month of each quarter period τ. CP(τ,2) series are log-differenced, as specified in Section 3.5. "p-KS" signifies a p-value of the Kolmogorov-Smirnov test for normality. "p-AD" signifies a p-value of the Anderson-Darling test for normality. "p-JB" signifies a p-value of the Jarque-Bera test for normality.

Table B.7 Sample Statistics for Commodity Prices CP(τ,2), Mixed-Frequency

	CP(τ,3)										
	Obs.	Mean	Median	Minimum	Maximum	Std. Dev.	Skewness	Kurtosis	p-KS	p-AD	p-JB
Australia	144	-0.388	0.765	-33.230	27.760	11.385	-0.621	3.824	0.000	0.003	0.008
Bahrain	141	-0.434	0.730	-33.230	27.760	11.466	-0.611	3.783	0.000	0.003	0.010
Belgium	84	0.720	1.645	-33.230	27.760	11.489	-0.574	3.914	0.000	0.043	0.027
Bolivia	104	0.507	1.060	-33.230	27.760	10.732	-0.531	4.150	0.000	0.049	0.014
Canada	144	-0.388	0.765	-33.230	27.760	11.385	-0.621	3.824	0.000	0.003	0.008
Chile	80	0.471	1.600	-33.230	27.760	11.665	-0.528	3.808	0.000	0.079	0.042
Czech Republic	80	0.471	1.600	-33.230	27.760	11.665	-0.528	3.808	0.000	0.079	0.042
Denmark	100	0.846	1.215	-33.230	27.760	10.783	-0.610	4.277	0.000	0.017	0.009
Dominican Republic	86	1.393	2.190	-33.230	27.760	11.347	-0.729	4.168	0.000	0.003	0.010
Ecuador	64	1.942	2.050	-33.230	27.760	10.797	-0.333	4.143	0.000	0.487	0.059
Estonia	84	0.720	1.645	-33.230	27.760	11.489	-0.574	3.914	0.000	0.043	0.027
Hong Kong	144	-0.388	0.765	-33.230	27.760	11.385	-0.621	3.824	0.000	0.003	0.008
Hungary	84	0.720	1.645	-33.230	27.760	11.489	-0.574	3.914	0.000	0.043	0.027
Iceland	76	0.186	1.130	-33.230	27.760	11.859	-0.472	3.695	0.000	0.155	0.069
Israel	84	0.720	1.645	-33.230	27.760	11.489	-0.574	3.914	0.000	0.043	0.027
Kazakhstan	86	0.699	1.215	-33.230	27.760	11.354	-0.575	4.000	0.000	0.035	0.023
Latvia	84	0.720	1.645	-33.230	27.760	11.489	-0.574	3.914	0.000	0.043	0.027
Luxembourg	64	1.942	2.050	-33.230	27.760	10.797	-0.333	4.143	0.000	0.487	0.059
Malta	64	1.942	2.050	-33.230	27.760	10.797	-0.333	4.143	0.000	0.487	0.059
Netherlands	80	0.471	1.600	-33.230	27.760	11.665	-0.528	3.808	0.000	0.079	0.042
New Zealand	115	0.599	1.090	-33.230	27.760	10.306	-0.560	4.435	0.000	0.021	0.006
Norway	144	-0.388	0.765	-33.230	27.760	11.385	-0.621	3.824	0.000	0.003	0.008
Peru	144	-0.388	0.765	-33.230	27.760	11.385	-0.621	3.824	0.000	0.003	0.008
Philippines	140	-0.354	0.765	-33.230	27.760	11.468	-0.629	3.819	0.000	0.002	0.008
Serbia	80	0.471	1.600	-33.230	27.760	11.665	-0.528	3.808	0.000	0.079	0.042
Seychelles	144	-0.388	0.765	-33.230	27.760	11.385	-0.621	3.824	0.000	0.003	0.008
Singapore	144	-0.388	0.765	-33.230	27.760	11.385	-0.621	3.824	0.000	0.003	0.008
Slovakia	84	0.720	1.645	-33.230	27.760	11.489	-0.574	3.914	0.000	0.043	0.027
Slovenia	84	0.720	1.645	-33.230	27.760	11.489	-0.574	3.914	0.000	0.043	0.027
South Korea	144	-0.388	0.765	-33.230	27.760	11.385	-0.621	3.824	0.000	0.003	0.008
Thailand	92	0.998	1.645	-33.230	27.760	11.088	-0.642	4.184	0.000	0.011	0.012
Venezuela	71	0.323	2.030	-33.230	27.760	12.146	-0.501	3.613	0.000	0.120	0.076
Viet Nam	60	1.879	1.620	-33.230	27.760	11.104	-0.312	3.948	0.000	0.585	0.106

Note: The table reports the primary statistics obtained for monthly commodity price series at the third month of each quarter period τ. CP(τ,3) series are log-differenced, as specified in Section 3.5. "p-KS" signifies a p-value of the Kolmogorov-Smirnov test for normality. "p-AD" signifies a p-value of the Anderson-Darling test for normality. "p-JB" signifies a p-value of the Jarque-Bera test for normality.

Table B.8 Sample Statistics for Commodity Prices CP(τ,3), Mixed-Frequency

	CP(τ)										
	Obs.	Mean	Median	Minimum	Maximum	Std. Dev.	Skewness	Kurtosis	p-KS	p-AD	p-JB
Australia	144	-0.389	0.963	-34.033	23.693	10.742	-0.625	3.607	0.000	0.003	0.011
Bahrain	141	-0.477	0.750	-34.033	23.693	10.810	-0.607	3.569	0.000	0.005	0.014
Belgium	84	0.683	2.702	-29.990	23.693	10.504	-0.438	3.200	0.000	0.086	0.151
Bolivia	104	0.463	2.273	-29.990	23.693	9.812	-0.398	3.404	0.000	0.085	0.114
Canada	144	-0.389	0.963	-34.033	23.693	10.742	-0.625	3.607	0.000	0.003	0.011
Chile	80	0.438	2.512	-29.990	23.693	10.673	-0.383	3.107	0.000	0.184	0.265
Czech Republic	80	0.438	2.512	-29.990	23.693	10.673	-0.383	3.107	0.000	0.184	0.265
Denmark	100	0.778	2.702	-29.990	23.693	9.863	-0.477	3.497	0.000	0.031	0.062
Dominican Republic	86	1.330	3.265	-29.990	23.693	10.314	-0.603	3.480	0.000	0.005	0.041
Ecuador	64	1.907	3.245	-23.383	23.693	9.885	-0.163	2.987	0.000	0.544	0.500
Estonia	84	0.683	2.702	-29.990	23.693	10.504	-0.438	3.200	0.000	0.086	0.151
Hong Kong	144	-0.389	0.963	-34.033	23.693	10.742	-0.625	3.607	0.000	0.003	0.011
Hungary	84	0.683	2.702	-29.990	23.693	10.504	-0.438	3.200	0.000	0.086	0.151
Iceland	76	0.153	1.675	-29.990	23.693	10.858	-0.316	3.010	0.000	0.442	0.437
Israel	84	0.683	2.702	-29.990	23.693	10.504	-0.438	3.200	0.000	0.086	0.151
Kazakhstan	86	0.721	2.702	-29.990	23.693	10.385	-0.453	3.276	0.000	0.067	0.121
Latvia	84	0.683	2.702	-29.990	23.693	10.504	-0.438	3.200	0.000	0.086	0.151
Luxembourg	64	1.907	3.245	-23.383	23.693	9.885	-0.163	2.987	0.000	0.544	0.500
Malta	64	1.907	3.245	-23.383	23.693	9.885	-0.163	2.987	0.000	0.544	0.500
Netherlands	80	0.438	2.512	-29.990	23.693	10.673	-0.383	3.107	0.000	0.184	0.265
New Zealand	115	0.598	2.173	-29.990	23.693	9.459	-0.428	3.599	0.000	0.062	0.055
Norway	144	-0.389	0.963	-34.033	23.693	10.742	-0.625	3.607	0.000	0.003	0.011
Peru	144	-0.389	0.963	-34.033	23.693	10.742	-0.625	3.607	0.000	0.003	0.011
Philippines	140	-0.410	0.963	-34.033	23.693	10.819	-0.624	3.593	0.000	0.003	0.012
Serbia	80	0.438	2.512	-29.990	23.693	10.673	-0.383	3.107	0.000	0.184	0.265
Seychelles	144	-0.389	0.963	-34.033	23.693	10.742	-0.625	3.607	0.000	0.003	0.011
Singapore	144	-0.389	0.963	-34.033	23.693	10.742	-0.625	3.607	0.000	0.003	0.011
Slovakia	84	0.683	2.702	-29.990	23.693	10.504	-0.438	3.200	0.000	0.086	0.151
Slovenia	84	0.683	2.702	-29.990	23.693	10.504	-0.438	3.200	0.000	0.086	0.151
South Korea	144	-0.389	0.963	-34.033	23.693	10.742	-0.625	3.607	0.000	0.003	0.011
Thailand	92	0.970	2.938	-29.990	23.693	10.127	-0.518	3.439	0.000	0.020	0.060
Venezuela	71	0.389	2.650	-29.990	23.693	11.125	-0.365	2.946	0.000	0.258	0.345
Viet Nam	60	1.874	3.442	-23.383	23.693	10.154	-0.153	2.861	0.000	0.505	0.500

Note: The table reports the primary statistics obtained for quarterly commodity price series at each quarter period τ. CP(τ) series are log-differenced, as specified in Section 3.5. "p-KS" signifies a p-value of the Kolmogorov-Smirnov test for normality. "p-AD" signifies a p-value of the Anderson-Darling test for normality. "p-JB" signifies a p-value of the Jarque-Bera test for normality.

Table B.9 Sample Statistics for Commodity Prices CP(τ), Low-Frequency

	EG(τ)										
	Obs.	Mean	Median	Minimum	Maximum	Std. Dev.	Skewness	Kurtosis	p-KS	p-AD	p-JB
Australia	144	1.720	1.680	-4.890	6.503	1.823	-0.742	4.369	0.000	0.002	0.002
Bahrain	141	0.343	0.381	-10.974	9.638	3.985	-0.373	4.273	0.000	0.001	0.009
Belgium	84	1.230	1.233	-4.636	4.825	1.685	-0.893	5.176	0.000	0.005	0.002
Bolivia	104	2.285	2.486	-2.993	5.779	1.794	-0.580	3.252	0.000	0.028	0.041
Canada	144	1.273	1.537	-5.283	5.014	2.218	-0.913	4.024	0.000	0.001	0.001
Chile	80	2.806	3.211	-5.245	8.077	2.702	-0.798	3.777	0.000	0.007	0.015
Czech Republic	80	2.216	2.302	-6.334	6.701	2.899	-0.728	3.482	0.000	0.079	0.025
Denmark	100	1.200	1.457	-6.891	5.853	2.140	-1.255	6.130	0.000	0.001	0.001
Dominican Republic	86	3.789	3.783	-3.289	10.658	3.115	-0.179	2.786	0.000	0.350	0.500
Ecuador	64	2.153	2.575	-5.599	7.996	2.880	-0.584	3.108	0.000	0.018	0.087
Estonia	84	4.305	5.146	-21.270	13.137	6.416	-1.742	6.930	0.000	0.001	0.001
Hong Kong	144	3.353	3.626	-10.412	14.785	4.235	-0.360	4.108	0.000	0.012	0.014
Hungary	84	2.486	3.452	-7.522	5.143	2.741	-1.868	6.809	0.000	0.001	0.001
Iceland	76	2.210	2.498	-9.897	10.082	4.467	-0.548	3.243	0.000	0.025	0.080
Israel	84	1.671	2.126	-5.172	8.118	2.284	-0.470	4.219	0.000	0.002	0.023
Kazakhstan	86	4.946	4.721	-15.103	26.563	8.277	-0.077	2.825	0.000	0.363	0.500
Latvia	84	5.006	5.491	-16.567	14.232	6.364	-1.414	5.322	0.000	0.001	0.001
Luxembourg	64	1.003	1.749	-9.535	7.536	3.543	-1.031	4.501	0.000	0.001	0.005
Malta	64	2.558	2.467	-3.922	9.334	2.765	-0.177	2.969	0.000	0.592	0.500
Netherlands	80	1.409	1.573	-4.975	4.935	2.122	-0.808	3.565	0.000	0.023	0.017
New Zealand	115	1.538	1.856	-4.615	6.752	2.210	-0.593	3.130	0.000	0.001	0.034
Norway	144	1.712	1.538	-3.213	7.848	2.219	0.186	2.759	0.000	0.636	0.500
Peru	144	1.371	3.174	-25.730	15.841	7.137	-1.583	6.193	0.000	0.001	0.001
Philippines	140	1.381	2.217	-14.328	8.487	3.903	-1.855	7.301	0.000	0.001	0.001
Serbia	80	3.146	2.770	-22.730	19.640	6.351	-0.428	6.015	0.000	0.005	0.001
Seychelles	144	2.393	3.593	-10.055	12.413	4.764	-0.460	2.571	0.000	0.001	0.041
Singapore	144	3.809	4.424	-11.791	14.909	4.418	-0.766	3.986	0.000	0.001	0.003
Slovakia	84	3.779	3.771	-6.394	12.913	3.499	-0.762	4.633	0.000	0.001	0.005
Slovenia	84	2.267	2.934	-10.534	7.125	3.478	-1.725	6.366	0.000	0.001	0.001
South Korea	144	5.377	5.524	-8.607	14.191	3.708	-0.681	4.845	0.000	0.012	0.001
Thailand	92	2.828	3.244	-14.208	13.661	4.263	-1.338	6.434	0.000	0.001	0.001
Venezuela	71	0.140	1.167	-31.260	29.255	8.163	-0.351	6.548	0.000	0.019	0.001
Viet Nam	60	5.175	5.132	-18.007	35.057	6.331	0.939	12.162	0.000	0.001	0.001

Note: The table reports the primary statistics obtained for economic growth series at each quarter period τ. EG(τ) series are log-differenced, as specified in Section 3.5. "p-KS" signifies a p-value of the Kolmogorov-Smirnov test for normality. "p-AD" signifies a p-value of the Anderson-Darling test for normality. "p-JB" signifies a p-value of the Jarque-Bera test for normality.

Table B.10 Sample Statistics for Economic Growth

B.3 Optimal VAR order selection criteria

	AIC	SC	HQ
Australia	4	2	3
Bahrain	4	1	2
Belgium	2	2	2
Bolivia	4	1	2
Canada	4	2	4
Chile	2	1	2
Czech Republic	2	2	2
Denmark	4	1	2
Dominican Republic	3	1	3
Ecuador	2	1	2
Estonia	4	1	4
Hong Kong	4	2	4
Hungary	4	2	2
Iceland	2	2	2
Israel	3	2	2
Kazakhstan	4	1	1
Latvia	4	4	4
Luxembourg	2	2	2
Malta	2	2	2
Netherlands	4	2	2
New Zealand	4	1	1
Norway	4	2	4
Peru	4	2	2
Philippines	4	1	4
Serbia	4	1	1
Seychelles	4	2	4
Singapore	4	2	2
Slovakia	4	1	4
Slovenia	4	2	3
South Africa	4	2	2
Thailand	4	1	4
Venezuela	4	1	1
Viet Nam	2	2	2

Note: The table shows the lag order that is selected by different ICs. The maximum lag of 4 is specified in each model.

Table B.11 Optimal VAR Order Selection Criteria

B.4 Results of full sample Granger causality tests for economic growth and world commodity prices for different lag orders and horizons

	Panel A: Mixed-frequency model								Panel B: Low-frequency model							
	lag order = 1		lag order = 2		lag order = 3		lag order = 4		lag order = 1		lag order = 2		lag order = 3		lag order = 4	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Exporters																
Australia	0.287	0.250	0.366	0.656	0.444	0.431	0.532	0.585	0.099	0.207	0.073	0.365	0.141	0.456	0.410	0.509
Bolivia	0.300	0.260	0.240	0.139	0.585	0.017	0.406	0.101	0.100	0.037	0.033	0.015	0.217	0.080	0.513	0.052
Canada	0.557	0.001	0.287	0.001	0.774	0.005	0.866	0.003	0.239	0.108	0.004	0.010	0.165	0.007	0.516	0.010
Chile	0.627	0.124	0.076	0.327	0.758	0.193	0.926	0.080	0.289	0.058	0.363	0.246	0.462	0.344	0.956	0.643
Denmark	0.726	0.025	0.148	0.017	0.508	0.045	0.513	0.084	0.798	0.410	0.364	0.017	0.571	0.085	0.779	0.040
Ecuador	0.052	0.001	0.556	0.010	0.220	0.001	0.296	0.019	0.033	0.001	0.103	0.034	0.692	0.289	0.421	0.279
Kazakhstan	0.546	0.049	0.346	0.214	0.299	0.002	0.412	0.022	0.588	0.383	0.295	0.200	0.296	0.638	0.049	0.836
New Zealand	0.279	0.396	0.355	0.174	0.569	0.118	0.854	0.578	0.060	0.404	0.401	0.063	0.356	0.332	0.728	0.150
Norway	0.670	0.001	0.291	0.016	0.287	0.005	0.609	0.329	0.302	0.461	0.232	0.405	0.323	0.434	0.198	0.952
Peru	0.904	0.909	0.768	0.176	0.694	0.093	0.795	0.244	0.887	0.895	0.974	0.905	0.980	0.877	0.969	0.637
South Africa	0.470	0.001	0.340	0.019	0.523	0.002	0.608	0.010	0.917	0.167	0.035	0.004	0.255	0.007	0.560	0.004
Venezuela	0.326	0.007	0.441	0.112	0.027	0.369	0.111	0.800	0.602	0.001	0.404	0.008	0.010	0.041	0.011	0.304
Importers																
Czech Republic	0.477	0.011	0.923	0.004	0.744	0.004	0.678	0.028	0.931	0.555	0.355	0.034	0.919	0.092	0.968	0.135
Dominican Republic	0.693	0.184	0.808	0.102	0.751	0.031	0.107	0.420	0.531	0.733	0.562	0.015	0.148	0.011	0.364	0.097
Hungary	0.233	0.032	0.645	0.039	0.123	0.189	0.469	0.029	0.318	0.460	0.344	0.367	0.625	0.293	0.300	0.121
Luxembourg	0.138	0.063	0.215	0.137	0.673	0.348	0.610	0.340	0.322	0.998	0.137	0.025	0.257	0.161	0.277	0.304
Malta	0.082	0.069	0.055	0.021	0.300	0.016	0.066	0.192	0.503	0.672	0.484	0.581	0.646	0.460	0.385	0.358
Philippines	0.206	0.013	0.332	0.014	0.189	0.027	0.010	0.010	0.005	0.119	0.114	0.091	0.044	0.151	0.007	0.640
Slovakia	0.180	0.039	0.420	0.002	0.326	0.003	0.492	0.002	0.866	0.007	0.051	0.059	0.754	0.143	0.388	0.042
Slovenia	0.276	0.027	0.170	0.008	0.175	0.002	0.259	0.018	0.739	0.461	0.022	0.001	0.063	0.001	0.052	0.002
Both (Hybrid)																
Bahrain	0.219	0.186	0.579	0.126	0.101	0.222	0.440	0.063	0.086	0.355	0.083	0.228	0.036	0.321	0.170	0.340
Belgium	0.916	0.017	0.444	0.445	0.532	0.239	0.786	0.322	0.960	0.520	0.011	0.111	0.023	0.321	0.139	0.450
Hong Kong	0.081	0.009	0.441	0.022	0.459	0.009	0.854	0.002	0.003	0.231	0.024	0.006	0.072	0.019	0.126	0.122
Estonia	0.921	0.001	0.192	0.004	0.276	0.001	0.111	0.001	0.468	0.044	0.012	0.001	0.162	0.005	0.340	0.156
Iceland	0.897	0.713	0.840	0.832	0.725	0.865	0.507	0.352	0.619	0.587	0.337	0.437	0.208	0.699	0.397	0.610
Israel	0.517	0.003	0.596	0.516	0.432	0.208	0.643	0.441	0.147	0.313	0.178	0.130	0.592	0.333	0.954	0.336
Latvia	0.799	0.007	0.378	0.046	0.003	0.001	0.027	0.001	0.541	0.013	0.011	0.010	0.037	0.001	0.019	0.015
Netherlands	0.582	0.001	0.454	0.016	0.744	0.010	0.009	0.006	0.313	0.721	0.070	0.013	0.354	0.009	0.366	0.006
Serbia	0.518	0.012	0.729	0.384	0.432	0.040	0.140	0.410	0.147	0.001	0.274	0.138	0.740	0.339	0.775	0.045
Seychelles	0.675	0.200	0.397	0.185	0.154	0.027	0.072	0.044	0.622	0.990	0.118	0.806	0.206	0.858	0.147	0.939
Singapore	0.005	0.004	0.030	0.084	0.029	0.079	0.079	0.004	0.007	0.095	0.005	0.101	0.003	0.075	0.003	0.491
Thailand	0.062	0.001	0.100	0.010	0.034	0.010	0.411	0.025	0.039	0.158	0.113	0.034	0.196	0.089	0.437	0.019
Viet Nam	0.238	0.497	0.054	0.629	0.006	0.445	0.005	0.198	0.333	0.858	0.289	0.805	0.062	0.655	0.032	0.956

Note: The table shows the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizon of one quarter, i.e. short-horizon, and the lag orders p∊{1,2,3,4}. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0\colon CP \nRightarrow EG$ *(⇏ means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with* $N = 999$ *replications. All variables are mean-centred and annual log-differenced.*

Table B.12 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices at the Horizon 1

	Panel A: Mixed-frequency model								Panel B: Low-frequency model							
	lag order = 1		lag order = 2		lag order = 3		lag order = 4		lag order = 1		lag order = 2		lag order = 3		lag order = 4	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Exporters																
Australia	0.067	0.499	0.379	0.507	0.202	0.712	0.843	0.499	0.067	0.299	0.403	0.490	0.183	0.968	0.576	0.687
Bolivia	0.044	0.395	0.101	0.113	0.166	0.312	0.246	0.430	0.005	0.186	0.062	0.247	0.195	0.199	0.056	0.234
Canada	0.252	0.001	0.061	0.074	0.674	0.007	0.615	0.003	0.100	0.516	0.174	0.493	0.265	0.004	0.229	0.011
Chile	0.132	0.142	0.405	0.182	0.814	0.131	0.814	0.374	0.212	0.128	0.184	0.123	0.502	0.254	0.549	0.253
Denmark	0.649	0.076	0.705	0.007	0.827	0.047	0.819	0.034	0.599	0.866	0.276	0.217	0.374	0.138	0.292	0.082
Ecuador	0.185	0.005	0.341	0.012	0.437	0.028	0.210	0.038	0.020	0.001	0.115	0.020	0.238	0.087	0.181	0.089
Kazakhstan	0.381	0.156	0.189	0.097	0.143	0.049	0.312	0.121	0.701	0.675	0.141	0.304	0.068	0.792	0.103	0.705
New Zealand	0.381	0.009	0.894	0.057	0.754	0.147	0.825	0.056	0.166	0.140	0.368	0.021	0.718	0.050	0.373	0.051
Norway	0.130	0.419	0.080	0.665	0.154	0.933	0.145	0.057	0.142	0.453	0.187	0.583	0.122	0.893	0.133	0.866
Peru	0.849	0.613	0.465	0.485	0.361	0.474	0.827	0.151	0.856	0.932	0.506	0.473	0.510	0.498	0.662	0.734
South Africa	0.265	0.006	0.056	0.531	0.004	0.549	0.336	0.038	0.940	0.333	0.102	0.509	0.011	0.515	0.349	0.012
Venezuela	0.387	0.002	0.088	0.099	0.032	0.266	0.035	0.528	0.454	0.002	0.017	0.001	0.006	0.053	0.001	0.146
Importers																
Czech Republic	0.529	0.027	0.490	0.491	0.447	0.498	0.419	0.407	0.968	0.958	0.489	0.519	0.485	0.508	0.496	0.483
Dominican Republic	0.184	0.091	0.655	0.202	0.136	0.719	0.083	0.450	0.257	0.016	0.278	0.088	0.688	0.602	0.626	0.249
Hungary	0.167	0.020	0.169	0.537	0.495	0.109	0.594	0.198	0.197	0.750	0.207	0.082	0.412	0.200	0.223	0.281
Luxembourg	0.656	0.001	0.525	0.008	0.309	0.168	0.210	0.281	0.519	0.396	0.126	0.025	0.306	0.117	0.431	0.084
Malta	0.049	0.115	0.110	0.012	0.062	0.052	0.019	0.044	0.514	0.393	0.325	0.079	0.123	0.212	0.181	0.056
Philippines	0.040	0.182	0.080	0.194	0.014	0.152	0.003	0.004	0.035	0.081	0.041	0.026	0.001	0.180	0.017	0.005
Slovakia	0.295	0.003	0.719	0.001	0.455	0.001	0.266	0.015	0.574	0.111	0.195	0.002	0.722	0.018	0.651	0.018
Slovenia	0.377	0.132	0.258	0.478	0.134	0.045	0.061	0.160	0.518	0.676	0.247	0.502	0.111	0.073	0.313	0.017
Both (Hybrid)																
Bahrain	0.270	0.294	0.098	0.418	0.279	0.022	0.182	0.007	0.135	0.093	0.016	0.163	0.174	0.189	0.129	0.019
Belgium	0.500	0.125	0.506	0.515	0.510	0.516	0.517	0.507	0.640	0.338	0.449	0.485	0.217	0.500	0.501	0.527
Hong Kong	0.197	0.004	0.386	0.010	0.880	0.004	0.912	0.004	0.005	0.018	0.065	0.003	0.247	0.016	0.331	0.003
Estonia	0.408	0.020	0.384	0.068	0.248	0.035	0.542	0.025	0.474	0.520	0.266	0.067	0.576	0.194	0.523	0.302
Iceland	0.572	0.836	0.654	0.629	0.289	0.747	0.558	0.631	0.318	0.718	0.298	0.389	0.252	0.812	0.411	0.359
Israel	0.532	0.029	0.504	0.249	0.906	0.349	0.803	0.217	0.459	0.571	0.033	0.004	0.582	0.129	0.422	0.142
Latvia	0.833	0.041	0.350	0.012	0.555	0.006	0.853	0.038	0.404	0.039	0.120	0.004	0.035	0.003	0.064	0.001
Netherlands	0.451	0.117	0.474	0.531	0.062	0.010	0.009	0.014	0.271	0.371	0.241	0.444	0.226	0.001	0.244	0.002
Serbia	0.190	0.078	0.249	0.134	0.148	0.153	0.284	0.323	0.074	0.067	0.252	0.030	0.604	0.101	0.719	0.196
Seychelles	0.364	0.635	0.080	0.001	0.337	0.062	0.135	0.003	0.465	0.874	0.142	0.510	0.438	0.490	0.185	0.687
Singapore	0.023	0.014	0.041	0.050	0.136	0.077	0.078	0.004	0.009	0.040	0.007	0.005	0.025	0.112	0.017	0.005
Thailand	0.001	0.001	0.011	0.009	0.195	0.033	0.201	0.043	0.004	0.006	0.097	0.070	0.295	0.012	0.367	0.122
Viet Nam	0.382	0.957	0.127	0.781	0.097	0.200	0.020	0.342	0.100	0.892	0.036	0.856	0.112	0.889	0.169	0.260

Note: The table shows the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizon of two quarters, i.e. longer-horizon, and lag orders p∊{1,2,3,4}. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0: CP \nRightarrow EG$ (⇏ means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ replications. All variables are mean-centred and annual log-differenced.

Table B.13 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices at the Horizon 2

	Panel A: Mixed-frequency model								Panel B: Low-frequency model							
	lag order = 1		lag order = 2		lag order = 3		lag order = 4		lag order = 1		lag order = 2		lag order = 3		lag order = 4	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Exporters																
Australia	0.261	0.475	0.311	0.939	0.390	0.791	0.312	0.287	0.033	0.272	0.045	0.408	0.052	0.650	0.191	0.347
Bolivia	0.045	0.297	0.113	0.537	0.086	0.569	0.204	0.387	0.022	0.238	0.100	0.335	0.043	0.332	0.165	0.216
Canada	0.287	0.023	0.479	0.413	0.464	0.491	0.341	0.468	0.117	0.785	0.004	0.055	0.493	0.479	0.330	0.456
Chile	0.177	0.210	0.685	0.433	0.818	0.461	0.526	0.474	0.130	0.153	0.211	0.132	0.259	0.277	0.522	0.275
Denmark	0.343	0.089	0.563	0.021	0.266	0.028	0.394	0.272	0.374	0.439	0.017	0.065	0.231	0.037	0.137	0.041
Ecuador	0.267	0.019	0.313	0.049	0.486	0.111	0.482	0.099	0.028	0.001	0.048	0.047	0.072	0.046	0.147	0.047
Kazakhstan	0.174	0.968	0.118	0.334	0.296	0.732	0.126	0.380	0.274	0.843	0.301	0.417	0.478	0.349	0.333	0.180
New Zealand	0.487	0.245	0.225	0.196	0.713	0.024	0.492	0.012	0.045	0.056	0.240	0.121	0.371	0.152	0.126	0.175
Norway	0.116	0.725	0.345	0.253	0.229	0.205	0.224	0.124	0.153	0.517	0.059	0.704	0.065	0.574	0.093	0.652
Peru	0.979	0.539	0.809	0.433	0.337	0.518	0.314	0.594	0.758	0.944	0.800	0.973	0.362	0.776	0.199	0.947
South Africa	0.787	0.005	0.579	0.120	0.101	0.499	0.485	0.476	0.857	0.637	0.091	0.001	0.318	0.499	0.544	0.490
Venezuela	0.403	0.017	0.006	0.293	0.120	0.222	0.023	0.540	0.197	0.006	0.074	0.032	0.198	0.105	0.088	0.287
Importers																
Czech Republic	0.722	0.161	0.747	0.263	0.135	0.094	0.078	0.210	0.939	0.385	0.646	0.003	0.575	0.023	0.609	0.053
Dominican Republic	0.476	0.134	0.102	0.296	0.052	0.571	0.058	0.629	0.667	0.204	0.836	0.349	0.852	0.717	0.944	0.877
Hungary	0.236	0.383	0.257	0.287	0.628	0.656	0.355	0.345	0.138	0.298	0.052	0.070	0.087	0.109	0.166	0.122
Luxembourg	0.703	0.096	0.589	0.277	0.309	0.149	0.593	0.164	0.689	0.237	0.205	0.036	0.146	0.022	0.166	0.025
Malta	0.400	0.087	0.227	0.159	0.220	0.354	0.033	0.134	0.322	0.100	0.188	0.107	0.051	0.018	0.077	0.027
Philippines	0.013	0.055	0.024	0.006	0.001	0.007	0.008	0.010	0.034	0.035	0.001	0.001	0.005	0.003	0.011	0.008
Slovakia	0.756	0.074	0.495	0.037	0.357	0.200	0.178	0.288	0.513	0.477	0.100	0.528	0.198	0.850	0.118	0.821
Slovenia	0.497	0.261	0.586	0.043	0.146	0.168	0.132	0.109	0.414	0.066	0.030	0.011	0.104	0.029	0.094	0.059
Both (Hybrid)																
Bahrain	0.629	0.124	0.353	0.062	0.509	0.005	0.472	0.004	0.476	0.094	0.202	0.009	0.465	0.003	0.442	0.011
Belgium	0.532	0.258	0.710	0.154	0.301	0.117	0.302	0.456	0.401	0.202	0.035	0.183	0.142	0.160	0.202	0.107
Hong Kong	0.225	0.010	0.685	0.001	0.794	0.001	0.897	0.003	0.188	0.005	0.073	0.002	0.504	0.001	0.666	0.010
Estonia	0.743	0.075	0.153	0.731	0.230	0.472	0.408	0.313	0.449	0.950	0.008	0.479	0.058	0.526	0.115	0.508
Iceland	0.317	0.714	0.071	0.689	0.397	0.891	0.589	0.038	0.207	0.777	0.087	0.861	0.095	0.560	0.214	0.679
Israel	0.644	0.279	0.710	0.241	0.393	0.218	0.478	0.032	0.875	0.305	0.038	0.033	0.132	0.161	0.362	0.226
Latvia	0.574	0.112	0.653	0.041	0.712	0.033	0.438	0.029	0.316	0.292	0.112	0.004	0.221	0.031	0.265	0.008
Netherlands	0.457	0.104	0.092	0.029	0.070	0.035	0.015	0.007	0.134	0.156	0.023	0.011	0.213	0.008	0.151	0.006
Serbia	0.178	0.417	0.069	0.353	0.155	0.408	0.233	0.270	0.100	0.087	0.709	0.215	0.674	0.314	0.193	0.285
Seychelles	0.276	0.736	0.616	0.495	0.508	0.496	0.491	0.484	0.366	0.958	0.061	0.274	0.412	0.464	0.491	0.490
Singapore	0.081	0.017	0.038	0.021	0.109	0.005	0.228	0.003	0.076	0.020	0.001	0.004	0.004	0.005	0.028	0.006
Thailand	0.028	0.090	0.166	0.221	0.159	0.475	0.485	0.737	0.022	0.040	0.097	0.043	0.056	0.056	0.114	0.119
Viet Nam	0.600	0.467	0.120	0.244	0.064	0.561	0.135	0.089	0.157	0.995	0.263	0.702	0.440	0.463	0.653	0.104

Note: The table shows the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizon of three quarters, i.e. longer-horizon, and lag orders pϵ{1,2,3,4}. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. H_0: $CP \nRightarrow EG$ (⇏ means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ replications. All variables are mean-centred and annual log-differenced.

Table B.14 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices at the Horizon 3

	Panel A: Mixed-frequency model								Panel B: Low-frequency model							
	lag order = 1		lag order = 2		lag order = 3		lag order = 4		lag order = 1		lag order = 2		lag order = 3		lag order = 4	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Exporters																
Australia	0.157	0.585	0.424	0.864	0.516	0.743	0.620	0.530	0.071	0.457	0.080	0.487	0.307	0.634	0.295	0.583
Bolivia	0.035	0.320	0.016	0.264	0.092	0.082	0.144	0.020	0.030	0.135	0.054	0.148	0.143	0.039	0.224	0.005
Canada	0.165	0.034	0.469	0.192	0.214	0.064	0.484	0.097	0.089	0.402	0.057	0.054	0.180	0.052	0.244	0.047
Chile	0.431	0.390	0.717	0.162	0.581	0.248	0.910	0.318	0.176	0.278	0.424	0.191	0.629	0.258	0.932	0.150
Denmark	0.180	0.041	0.278	0.013	0.194	0.033	0.323	0.054	0.183	0.181	0.105	0.170	0.333	0.040	0.251	0.072
Ecuador	0.164	0.005	0.305	0.026	0.554	0.190	0.594	0.369	0.030	0.017	0.122	0.053	0.147	0.054	0.163	0.092
Kazakhstan	0.278	0.626	0.243	0.763	0.014	0.364	0.049	0.455	0.880	0.856	0.932	0.691	0.551	0.138	0.675	0.133
New Zealand	0.246	0.160	0.420	0.140	0.683	0.095	0.665	0.140	0.118	0.097	0.221	0.148	0.153	0.205	0.312	0.344
Norway	0.130	0.929	0.124	0.033	0.144	0.060	0.226	0.069	0.043	0.588	0.051	0.629	0.062	0.624	0.136	0.635
Peru	0.864	0.469	0.914	0.712	0.457	0.872	0.533	0.890	0.629	0.972	0.823	0.887	0.341	0.830	0.344	0.775
South Africa	0.672	0.017	0.838	0.021	0.433	0.185	0.555	0.037	0.970	0.949	0.177	0.006	0.423	0.014	0.901	0.008
Venezuela	0.453	0.030	0.511	0.236	0.454	0.406	0.963	0.388	0.280	0.008	0.585	0.019	0.535	0.107	0.519	0.112
Importers																
Czech Republic	0.834	0.100	0.698	0.008	0.785	0.057	0.798	0.211	0.860	0.168	0.706	0.001	0.986	0.014	0.901	0.012
Dominican Republic	0.976	0.589	0.188	0.747	0.037	0.904	0.020	0.856	0.782	0.244	0.832	0.467	0.952	0.356	0.961	0.349
Hungary	0.223	0.257	0.486	0.370	0.663	0.451	0.342	0.277	0.182	0.139	0.205	0.083	0.224	0.088	0.285	0.088
Luxembourg	0.502	0.118	0.149	0.141	0.456	0.101	0.515	0.134	0.964	0.148	0.209	0.015	0.214	0.008	0.236	0.013
Malta	0.215	0.332	0.418	0.350	0.279	0.360	0.115	0.646	0.148	0.074	0.081	0.093	0.062	0.170	0.061	0.205
Philippines	0.042	0.111	0.009	0.095	0.072	0.041	0.036	0.128	0.017	0.044	0.018	0.088	0.021	0.121	0.010	0.145
Slovakia	0.477	0.237	0.610	0.302	0.322	0.433	0.592	0.643	0.486	0.423	0.320	0.163	0.506	0.224	0.674	0.360
Slovenia	0.554	0.139	0.067	0.035	0.023	0.052	0.044	0.047	0.366	0.098	0.038	0.021	0.039	0.090	0.041	0.027
Both (Hybrid)																
Bahrain	0.577	0.007	0.476	0.015	0.552	0.106	0.968	0.016	0.737	0.038	0.409	0.011	0.526	0.017	0.676	0.085
Belgium	0.327	0.158	0.597	0.043	0.573	0.079	0.635	0.175	0.259	0.168	0.288	0.117	0.329	0.060	0.195	0.145
Hong Kong	0.669	0.193	0.502	0.005	0.673	0.012	0.617	0.005	0.407	0.007	0.222	0.002	0.502	0.005	0.542	0.004
Estonia	0.304	0.088	0.042	0.633	0.322	0.839	0.320	0.573	0.319	0.854	0.110	0.131	0.122	0.319	0.313	0.248
Iceland	0.092	0.893	0.129	0.639	0.296	0.531	0.145	0.609	0.081	0.710	0.079	0.374	0.156	0.499	0.232	0.209
Israel	0.244	0.331	0.250	0.170	0.437	0.071	0.571	0.013	0.834	0.296	0.122	0.117	0.424	0.251	0.265	0.168
Latvia	0.364	0.209	0.675	0.384	0.514	0.086	0.462	0.032	0.306	0.610	0.019	0.085	0.082	0.116	0.126	0.081
Netherlands	0.139	0.159	0.660	0.029	0.515	0.005	0.559	0.001	0.114	0.082	0.025	0.007	0.132	0.013	0.204	0.001
Serbia	0.036	0.411	0.102	0.470	0.168	0.474	0.144	0.758	0.460	0.179	0.757	0.146	0.830	0.429	0.650	0.629
Seychelles	0.513	0.760	0.704	0.190	0.590	0.260	0.572	0.409	0.250	0.971	0.060	0.832	0.067	0.745	0.080	0.791
Singapore	0.316	0.003	0.041	0.003	0.106	0.001	0.014	0.002	0.279	0.018	0.014	0.001	0.051	0.001	0.048	0.001
Thailand	0.080	0.376	0.067	0.484	0.312	0.616	0.565	0.677	0.062	0.109	0.116	0.146	0.139	0.235	0.219	0.576
Viet Nam	0.059	0.401	0.234	0.394	0.548	0.188	0.392	0.228	0.126	0.888	0.295	0.240	0.553	0.016	0.503	0.025

Note: The table shows the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizon of four quarters, i.e. longer-horizon, and lag orders pϵ{1,2,3,4}. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0: CP \nRightarrow EG$ *(⇏ means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with* $N = 999$ *replications. All variables are mean-centred and annual log-differenced.*

Table B.15 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices at the Horizon 4

	Panel A: Mixed-frequency model								Panel B: Low-frequency model							
	lag order = 1		lag order = 2		lag order = 3		lag order = 4		lag order = 1		lag order = 2		lag order = 3		lag order = 4	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Exporters																
Australia	0.147	0.922	0.387	0.636	0.205	0.604	0.446	0.347	0.141	0.852	0.214	0.310	0.272	0.389	0.325	0.221
Bolivia	0.369	0.496	0.933	0.475	0.969	0.415	0.881	0.187	0.127	0.196	0.331	0.408	0.649	0.588	0.488	0.028
Canada	0.243	0.303	0.021	0.295	0.099	0.107	0.102	0.159	0.074	0.115	0.178	0.055	0.215	0.083	0.239	0.081
Chile	0.366	0.626	0.288	0.820	0.334	0.364	0.583	0.602	0.291	0.765	0.712	0.950	0.809	0.460	0.877	0.458
Denmark	0.062	0.135	0.406	0.017	0.574	0.112	0.578	0.282	0.368	0.073	0.637	0.107	0.746	0.105	0.838	0.302
Ecuador	0.403	0.564	0.347	0.804	0.519	0.867	0.129	0.875	0.793	0.634	0.059	0.678	0.142	0.287	0.256	0.458
Kazakhstan	0.722	0.155	0.990	0.096	0.843	0.354	0.551	0.269	0.353	0.168	0.257	0.140	0.338	0.324	0.432	0.215
New Zealand	0.868	0.019	0.580	0.039	0.772	0.058	0.885	0.041	0.954	0.493	0.260	0.510	0.423	0.104	0.700	0.063
Norway	0.008	0.801	0.055	0.593	0.134	0.414	0.238	0.121	0.041	0.869	0.045	0.942	0.066	0.945	0.143	0.977
Peru	0.977	0.998	0.275	0.980	0.011	0.841	0.062	0.676	0.757	0.905	0.210	0.886	0.081	0.278	0.183	0.250
South Africa	0.453	0.247	0.121	0.128	0.122	0.022	0.238	0.013	0.954	0.142	0.915	0.045	0.809	0.024	0.422	0.034
Venezuela	0.512	0.787	0.579	0.988	0.456	0.671	0.095	0.841	0.677	0.390	0.188	0.743	0.229	0.446	0.218	0.593
Importers																
Czech Republic	0.047	0.040	0.473	0.031	0.310	0.221	0.085	0.211	0.886	0.022	0.960	0.004	0.700	0.009	0.492	0.034
Dominican Republic	0.809	0.784	0.431	0.671	0.341	0.837	0.373	0.775	0.861	0.758	0.867	0.664	0.958	0.751	0.935	0.548
Hungary	0.133	0.196	0.439	0.086	0.466	0.137	0.313	0.261	0.600	0.056	0.339	0.039	0.372	0.050	0.381	0.110
Luxembourg	0.361	0.014	0.171	0.021	0.473	0.111	0.741	0.072	0.152	0.005	0.308	0.005	0.253	0.006	0.331	0.016
Malta	0.189	0.154	0.139	0.387	0.096	0.287	0.116	0.070	0.143	0.060	0.126	0.148	0.119	0.276	0.141	0.124
Philippines	0.021	0.452	0.018	0.235	0.003	0.320	0.006	0.310	0.041	0.293	0.012	0.441	0.026	0.221	0.012	0.239
Slovakia	0.524	0.922	0.849	0.944	0.645	0.974	0.383	0.993	0.248	0.909	0.463	0.943	0.307	0.501	0.233	0.602
Slovenia	0.450	0.075	0.898	0.118	0.891	0.319	0.348	0.361	0.216	0.052	0.378	0.030	0.352	0.025	0.042	0.062
Both (Hybrid)																
Bahrain	0.857	0.037	0.885	0.079	0.843	0.001	0.960	0.009	0.676	0.011	0.376	0.021	0.522	0.039	0.589	0.012
Belgium	0.381	0.114	0.222	0.024	0.198	0.046	0.554	0.064	0.160	0.003	0.297	0.050	0.365	0.075	0.502	0.107
Hong Kong	0.929	0.004	0.653	0.001	0.538	0.003	0.278	0.015	0.954	0.002	0.907	0.002	0.766	0.019	0.283	0.023
Estonia	0.016	0.482	0.086	0.572	0.098	0.446	0.257	0.707	0.132	0.428	0.062	0.187	0.052	0.153	0.160	0.202
Iceland	0.463	0.386	0.198	0.264	0.383	0.383	0.456	0.214	0.210	0.911	0.370	0.264	0.599	0.423	0.664	0.579
Israel	0.232	0.464	0.372	0.032	0.335	0.022	0.653	0.019	0.235	0.178	0.243	0.434	0.415	0.211	0.483	0.255
Latvia	0.135	0.426	0.806	0.111	0.957	0.057	0.889	0.205	0.153	0.749	0.350	0.027	0.607	0.038	0.096	0.021
Netherlands	0.240	0.015	0.595	0.004	0.411	0.016	0.435	0.044	0.124	0.013	0.125	0.001	0.235	0.002	0.422	0.004
Serbia	0.757	0.946	0.317	0.912	0.714	0.631	0.645	0.785	0.783	0.654	0.826	0.772	0.372	0.894	0.587	0.622
Seychelles	0.213	0.999	0.075	0.512	0.035	0.720	0.069	0.867	0.084	0.962	0.039	0.972	0.059	0.976	0.061	0.949
Singapore	0.497	0.002	0.210	0.003	0.474	0.001	0.679	0.008	0.969	0.001	0.920	0.001	0.878	0.003	0.809	0.002
Thailand	0.201	0.905	0.292	0.295	0.457	0.395	0.733	0.466	0.067	0.670	0.132	0.039	0.119	0.317	0.195	0.290
Viet Nam	0.722	0.052	0.579	0.153	0.360	0.263	0.372	0.673	0.510	0.046	0.633	0.020	0.791	0.068	0.902	0.127

Note: The table shows the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizon of six quarters, i.e. longer-horizon, and lag orders p∈{1,2,3,4}. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0: CP \nRightarrow EG$ (⇏ means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ replications. All variables are mean-centred and annual log-differenced.

Table B.16 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices at the Horizon 6

B.5 Results from national commodity export prices

This section reports the empirical results when using national commodity export prices as a proxy for commodity prices. The national commodity export prices are obtained from Chapter 2.

	Economic Growth	
	ADF with intercept, no trend	**PP with intercept, no trend**
Bahrain	-2.897**	-4.522***
Bolivia	-1.882	-5.473***
Chile	-4.089***	-3.271**
Ecuador	-2.264	-2.739*
Peru	-3.108**	-3.455**
Seychelles	-3.546***	-2.963**
Venezuela	-4.298***	-3.311**

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.17 Results for Unit Root Tests for Economic Growth, National Commodity Export Prices

	Commodity Price Growth: Low-Frequency	
	ADF with intercept, no trend	**PP with intercept, no trend**
Bahrain	-3.002**	-3.245**
Bolivia	-5.558***	-2.885*
Chile	-4.697***	-3.418**
Ecuador	-4.469***	-3.404**
Peru	-3.033**	-3.235**
Seychelles	-4.373***	-3.903***
Venezuela	-4.341***	-2.654*

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.18 Results for Unit Root Tests for Commodity Prices at Low-Frequency, National Commodity Export Prices

	Commodity Prices: Mixed-Frequency					
	ADF with intercept, no trend			PP with intercept, no trend		
	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Bahrain	-3.073**	-3.224**	-2.358	-3.783***	-3.529***	-3.655***
Bolivia	-5.661***	-4.936***	-5.664***	-3.599***	-3.781***	-3.020**
Chile	-4.133***	-4.436***	-4.616***	-3.407**	-3.522***	-3.541***
Ecuador	-4.369***	-2.196	-4.075***	-3.542***	-3.438**	-3.545***
Peru	-3.400**	-2.965**	-2.783*	-3.781***	-3.791***	-3.351**
Seychelles	-4.684***	-4.642***	-4.422***	-3.840***	-4.091***	-3.886***
Venezuela	-3.319**	-2.868*	-4.024***	-3.068**	-2.883*	-3.262**

Note: The table reports the test statistics obtained from the unit root tests for monthly commodity price series at a given month of each quarter period τ.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.19 Results for Unit Root Tests for Commodity Prices at Mixed-Frequency, National Commodity Export Prices

	KPSS with intercept, no trend				
	Economic Growth (EG)	CP(τ)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Bahrain	0.131	0.194	0.189	0.194	0.197
Bolivia	0.509**	0.152	0.14	0.172	0.117
Chile	0.078	0.138	0.144	0.138	0.131
Ecuador	0.215	0.307	0.326	0.304	0.268
Peru	0.564**	0.475**	0.478**	0.471**	0.466**
Seychelles	0.094	0.162	0.158	0.168	0.154
Venezuela	0.125	0.164	0.154	0.15	0.184

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.20 Results for KPSS Unit Root Tests, National Commodity Export Prices

KPSS test results suggest stationarity for all economic growth and commodity price series at 1% level of significance. In the case of Peru, both ADF and PP tests provide evidence of stationarity for economic growth and commodity price series. Similar results are found for Bolivia, where PP test reject the null hypothesis of unit root at 1% level of significance for economic growth series. Therefore, we conclude that the condition of stationarity is satisfied for all economic growth and commodity price series based on the combined results from KPSS, ADF and PP tests.

	AIC	SC	HQ
Bahrain	4	1	1
Bolivia	4	2	2
Chile	2	2	2
Ecuador	2	2	2
Peru	3	2	2
Seychelles	4	2	3
Venezuela	4	2	4

Note: The table shows the lag order that is selected by different ICs. The maximum lag of 4 is specified in each model.

Table B.21 Optimal VAR Order Selection Criteria, National Commodity Export Prices

	Panel A: Mixed-frequency model									
	horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Bahrain	0.062	0.182	0.635	0.227	0.905	0.408	0.606	0.111	0.417	0.012
Bolivia	0.050	0.544	0.299	0.268	0.087	0.024	0.015	0.037	0.083	0.338
Chile	0.203	0.005	0.192	0.007	0.040	0.016	0.135	0.070	0.564	0.314
Ecuador	0.309	0.038	0.653	0.016	0.796	0.085	0.293	0.189	0.820	0.756
Peru	0.713	0.043	0.511	0.499	0.283	0.503	0.612	0.148	0.021	0.337
Seychelles	0.095	0.036	0.514	0.485	0.413	0.498	0.048	0.457	0.155	0.685
Venezuela	0.003	0.071	0.004	0.307	0.088	0.670	0.241	0.479	0.888	0.597
	Panel B: Low-frequency model									
	horizon = 1		**horizon = 2**		**horizon = 3**		**horizon = 4**		**horizon = 6**	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Bahrain	0.867	0.331	0.891	0.518	0.749	0.392	0.583	0.181	0.444	0.053
Bolivia	0.288	0.761	0.143	0.627	0.053	0.847	0.013	0.727	0.070	0.981
Chile	0.311	0.020	0.228	0.001	0.101	0.008	0.112	0.031	0.171	0.610
Ecuador	0.746	0.500	0.451	0.032	0.520	0.094	0.979	0.193	0.754	0.299
Peru	0.263	0.478	0.485	0.501	0.216	0.321	0.131	0.109	0.039	0.252
Seychelles	0.216	0.909	0.504	0.483	0.383	0.397	0.291	0.770	0.274	0.813
Venezuela	0.003	0.006	0.005	0.005	0.101	0.021	0.352	0.073	0.932	0.692

Note: The table reports the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizons $h \epsilon \{1,2,3,4,6\}$*. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables.* $H_0: CP \nRightarrow EG$ *(⇏ means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with* $N = 999$ *replications. All variables are mean-centred and annual log-differenced.*

Table B.22 Results of Full Sample Granger Causality Tests for Economic Growth and National Commodity Export Prices

Panel A: Mixed-frequency model

	horizon = 1				horizon = 2				horizon = 3				horizon = 4				horizon = 6			
	EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Bahrain	0.185	0.239	0.272	0.315	0	0.021	0	0.021	0	0.011	0	0.011	0	0.021	0	0	0.200	0.253	0	0
Bolivia	0	0	0	0	0	0	0	0.164	0.018	0.036	0	0	0	0.018	0	0	0.345	1	0	0.236
Chile	0.129	0.161	0.774	1.000	0.032	0.097	0.806	0.839	0.097	0.129	0.806	0.968	0	0	0.548	0.871	0	0	0	0
Ecuador	0	0	1	1	0	0	0.467	0.533	0	0	0.200	0.333	0	0	0.200	0.333	0	0	0	0
Peru	0.179	0.274	0.126	0.189	0.147	0.253	0.032	0.137	0	0.053	0.200	0.221	0	0	0.137	0.211	0.147	0.379	0.095	0.137
Seychelles	0.011	0.074	0.021	0.032	0	0	0.864	1	0.045	0.227	0	0.682	0	0	0	0.091	0	0	0	0
Venezuela	0	0	0.955	1.000	0.163	0.207	0.065	0.163	0.022	0.065	0.033	0.076	0.130	0.141	0.043	0.065	0	0	0.109	0.196

Panel B: Low-frequency model

	horizon = 1				horizon = 2				horizon = 3				horizon = 4				horizon = 6			
	EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Bahrain	0	0.011	0.087	0.283	0.116	0.137	0.053	0.105	0.063	0.105	0	0.042	0.074	0.105	0	0.032	0.295	0.347	0	0
Bolivia	0	0	0	0	0	0	0	0	0	0.018	0	0	0.018	0.055	0	0	0.418	0.491	0.073	0.109
Chile	0.355	0.452	1	1	0.226	0.323	1	1	0.065	0.097	1	1	0	0	1	1	0	0	0	0
Ecuador	0.067	0.133	0.600	0.600	0	0.067	0.467	0.467	0	0	0.267	0.333	0	0	0	0.133	0	0	0	0
Peru	0.284	0.358	0.263	0.347	0.274	0.347	0.095	0.221	0.137	0.189	0.263	0.347	0	0	0.295	0.389	0.042	0.074	0.063	0.105
Seychelles	0.137	0.137	0.105	0.168	0	0.045	1	1	0	0.045	0.955	1	0	0	0.455	0.773	0	0	0	0
Venezuela	0	0.045	1	1	0.011	0.076	0.087	0.196	0	0.011	0.076	0.163	0.011	0.109	0	0.141	0	0.011	0	0.141

Note: The table reports the rejection frequencies at different significance levels for rolling window mixed- and low-frequency Granger causality tests with a window size of 50 quarters for the horizons $h\epsilon\{1,2,3,4,6\}$*. For each rolling window, the Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with* $N = 999$ *replications. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables.* $H_0: CP \nRightarrow EG$ *(⇏ means "does not Granger-cause"). All variables are mean-centred and annual log-differenced.*

Table B.23 Rejection Frequencies at Different Significant Levels for Rolling Window Granger Causality Tests, National Commodity Export Prices

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Bahrain	0.385	-1.452	2.471*	1.405*	1.242	5.155***
Bolivia	-1.684	-3.152	-1.154	0.030	-2.377	-0.072
Chile	9.713***	9.329***	14.534***	9.082***	8.159***	13.790***
Ecuador	3.553**	17.374***	22.070***	5.565***	13.118***	17.125***
Peru	-2.077	3.633*	-3.987	-4.009	1.408	-6.548
Seychelles	-0.370	1.785	5.780***	1.168	2.438*	6.745***
Viet Nam	1.634	2.431	5.000**	7.080***	3.374**	7.383***

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Clark and McCracken's (2001) critical value. All variables are mean-centred and annual log-differenced, as specified in Section 3.5.

Table B.24 Tests for Out-of-Sample Forecasting Ability – Regression Based Forecast Models, National Commodity Export Prices

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	**RW Benchmark**	**RWWD Benchmark**	**AR Benchmark**	**RW Benchmark**	**RWWD Benchmark**
Bahrain	0.322	-0.603**	-2.538***	0.302	-0.604**	-2.534***
Bolivia	0.845*	-2.264***	-3.408***	0.686*	-2.275***	-3.418***
Chile	-0.602	-3.502*	-3.617***	-0.522	-3.430	-3.552***
Ecuador	-2.128**	-2.740***	-2.723***	-2.113**	-2.761***	-2.737***
Peru	2.301	-5.127***	-4.076***	2.368	-4.879***	-3.967***
Seychelles	0.196*	-6.836***	-8.062***	0.249*	-6.836***	-8.059***
Viet Nam	-1.846	-2.970	-2.709	-1.718	-2.927	-2.684

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Diebold and Mariano's (1995) critical values. All variables are mean-centred and annual log-differenced, as specified in Section 3.5.

Table B.25 Tests for Out-of-Sample Forecasting Ability – Combination Forecast Models, National Commodity Export Prices

B.6 Alternative proxies for world commodity prices

This section presents the results from the empirical tests when a proxy for world commodity prices is any used of these indexes: Reuters/Jeffries (RJ), Goldman Sachs (GS), Moody's (MD), Thompson Reuters Core Commodity Equal Weighted (TR) and IMF non-fuel commodity price index (IMF non-fuel), as discussed in Section 3.5.

Particularly, this study identifies that the outcomes from all additional proxies are relatively similar to those obtained from the CRB index. Due to this reason, we only provide summary tables with the outcomes from the empirical tests, the interpretation of the results is consistent with the one provided in the main part of the Chapter 3.

	Panel A: Mixed-frequency model										Panel B: Low-frequency model									
	horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6		horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG
CRB																				
Exporters	1(12)	5(12)	1(12)	6(12)	1(12)	2(12)	1(12)	6(12)	2(12)	2(12)	2(12)	5(12)	2(12)	5(12)	3(12)	2(12)	0(12)	5(12)	1(12)	4(12)
Importers	2(8)	7(8)	2(8)	4(8)	2(8)	1(8)	3(8)	2(8)	1(8)	2(8)	2(8)	5(8)	1(8)	5(8)	2(8)	4(8)	3(8)	5(8)	2(8)	3(8)
Both (Hybrid)	5(13)	8(13)	2(13)	8(13)	2(13)	5(13)	1(13)	7(13)	1(13)	6(13)	3(13)	4(13)	3(13)	6(13)	3(13)	5(13)	3(13)	5(13)	2(13)	7(13)
Total	**8(33)**	**20(33)**	**5(33)**	**18(33)**	**5(33)**	**8(33)**	**5(33)**	**15(33)**	**4(33)**	**10(33)**	**7(33)**	**14(33)**	**6(33)**	**16(33)**	**8(33)**	**11(33)**	**6(33)**	**15(33)**	**5(33)**	**14(33)**
RJ																				
Exporters	1(12)	5(12)	5(12)	3(12)	1(12)	3(12)	1(12)	3(12)	2(12)	2(12)	1(12)	8(12)	5(12)	5(12)	6(12)	6(12)	2(12)	4(12)	1(12)	1(12)
Importers	0(8)	3(8)	2(8)	1(8)	2(8)	1(8)	1(8)	2(8)	2(8)	0(8)	1(8)	3(8)	2(8)	3(8)	3(8)	3(8)	1(8)	1(8)	2(8)	0(8)
Both (Hybrid)	5(13)	10(13)	5(13)	7(13)	2(13)	3(13)	1(13)	1(13)	3(13)	2(13)	5(13)	8(13)	2(13)	6(13)	3(13)	8(13)	4(13)	5(13)	2(13)	3(13)
Total	**6(33)**	**18(33)**	**12(33)**	**11(33)**	**5(33)**	**7(33)**	**3(33)**	**6(33)**	**7(33)**	**4(33)**	**7(33)**	**19(33)**	**9(33)**	**14(33)**	**12(33)**	**17(33)**	**7(33)**	**10(33)**	**5(33)**	**4(33)**
GS																				
Exporters	2(12)	6(12)	5(12)	3(12)	1(12)	3(12)	0(12)	4(12)	1(12)	2(12)	3(12)	5(12)	4(12)	5(12)	3(12)	5(12)	2(12)	3(12)	2(12)	1(12)
Importers	1(8)	1(8)	3(8)	0(8)	2(8)	0(8)	3(8)	1(8)	1(8)	0(8)	1(8)	4(8)	1(8)	3(8)	3(8)	3(8)	2(8)	3(8)	1(8)	0(8)
Both (Hybrid)	4(13)	8(13)	3(13)	6(13)	0(13)	1(13)	2(13)	0(13)	1(13)	1(13)	5(13)	8(13)	3(13)	6(13)	3(13)	3(13)	0(13)	1(13)	0(13)	1(13)
Total	**7(33)**	**15(33)**	**11(33)**	**9(33)**	**3(33)**	**4(33)**	**5(33)**	**5(33)**	**3(33)**	**3(33)**	**9(33)**	**17(33)**	**8(33)**	**14(33)**	**9(33)**	**11(33)**	**4(33)**	**7(33)**	**3(33)**	**2(33)**
Moody																				
Exporters	4(12)	7(12)	6(12)	7(12)	4(12)	5(12)	3(12)	2(12)	1(12)	0(12)	3(12)	8(12)	6(12)	7(12)	5(12)	4(12)	5(12)	5(12)	3(12)	1(12)
Importers	2(8)	4(8)	2(8)	3(8)	1(8)	2(8)	3(8)	2(8)	2(8)	2(8)	3(8)	3(8)	1(8)	4(8)	2(8)	3(8)	2(8)	3(8)	0(8)	3(8)
Both (Hybrid)	5(13)	7(13)	4(13)	5(13)	3(13)	5(13)	2(13)	1(13)	3(13)	1(13)	1(13)	8(13)	3(13)	8(13)	3(13)	5(13)	4(13)	4(13)	2(13)	4(13)
Total	**11(33)**	**18(33)**	**12(33)**	**15(33)**	**8(33)**	**12(33)**	**8(33)**	**5(33)**	**6(33)**	**3(33)**	**7(33)**	**19(33)**	**10(33)**	**19(33)**	**10(33)**	**12(33)**	**11(33)**	**12(33)**	**5(33)**	**8(33)**
TR																				
Exporters	0(12)	5(12)	8(12)	3(12)	4(12)	2(12)	2(12)	2(12)	4(12)	1(12)	2(12)	7(12)	4(12)	6(12)	5(12)	3(12)	3(12)	4(12)	2(12)	2(12)
Importers	1(8)	4(8)	3(8)	2(8)	1(8)	3(8)	2(8)	2(8)	3(8)	1(8)	1(8)	4(8)	2(8)	3(8)	1(8)	4(8)	2(8)	2(8)	0(8)	2(8)
Both (Hybrid)	5(13)	9(13)	5(13)	8(13)	4(13)	6(13)	5(13)	4(13)	3(13)	3(13)	5(13)	6(13)	5(13)	6(13)	4(13)	5(13)	4(13)	5(13)	4(13)	6(13)
Total	**6(33)**	**18(33)**	**16(33)**	**13(33)**	**9(33)**	**11(33)**	**9(33)**	**8(33)**	**10(33)**	**5(33)**	**8(33)**	**17(33)**	**11(33)**	**15(33)**	**10(33)**	**12(33)**	**9(33)**	**11(33)**	**6(33)**	**10(33)**
IMF non-fuel																				
Exporters	2(12)	10(12)	3(12)	5(12)	4(12)	2(12)	1(12)	1(12)	4(12)	1(12)	3(12)	7(12)	6(12)	6(12)	4(12)	3(12)	3(12)	3(12)	3(12)	1(12)
Importers	4(8)	5(8)	2(8)	5(8)	3(8)	4(8)	2(8)	2(8)	0(8)	1(8)	3(8)	5(8)	2(8)	3(8)	1(8)	4(8)	3(8)	2(8)	2(8)	2(8)
Both (Hybrid)	6(13)	9(13)	5(13)	7(13)	2(13)	5(13)	1(13)	6(13)	1(13)	3(13)	3(13)	8(13)	3(13)	7(13)	3(13)	7(13)	3(13)	6(13)	3(13)	6(13)
Total	**12(33)**	**24(33)**	**10(33)**	**17(33)**	**9(33)**	**11(33)**	**4(33)**	**9(33)**	**5(33)**	**5(33)**	**9(33)**	**20(33)**	**11(33)**	**16(33)**	**8(33)**	**14(33)**	**9(33)**	**11(33)**	**8(33)**	**9(33)**

Note: The table presents the number of countries for which full sample causality has been detected for horizons $h \epsilon \{1,2,3,4,6\}$. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0: CP \nRightarrow EG$ ($\nRightarrow$ means "does not Granger-cause"). The number in brackets denotes the total number of countries considered in the analysis, for each group of countries.

Table B.26 Summary of the Full Sample Granger Causality Test Results for Economic Growth and World Commodity Prices

	Panel A: Mixed-frequency model										Panel B: Low-frequency model									
	horizon = 1		**horizon = 2**		**horizon = 3**		**horizon = 4**		**horizon = 6**		**horizon = 1**		**horizon = 2**		**horizon = 3**		**horizon = 4**		**horizon = 6**	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
CRB																				
Exporters	8(12)	12(12)	8(12)	11(12)	6(12)	11(12)	7(12)	10(12)	7(12)	8(12)	7(12)	10(12)	6(12)	10(12)	5(12)	9(12)	6(12)	9(12)	5(12)	9(12)
Importers	7(8)	8(8)	6(8)	8(8)	4(8)	7(8)	4(8)	5(8)	2(8)	3(8)	3(8)	3(8)	2(8)	4(8)	2(8)	6(8)	3(8)	6(8)	3(8)	6(8)
Both (Hybrid)	8(13)	11(13)	8(13)	11(13)	8(13)	11(13)	8(13)	10(13)	8(13)	9(13)	7(13)	8(13)	6(13)	7(13)	5(13)	8(13)	6(13)	9(13)	4(13)	8(13)
Total	**23(33)**	**31(33)**	**22(33)**	**30(33)**	**18(33)**	**29(33)**	**19(33)**	**25(33)**	**17(33)**	**20(33)**	**17(33)**	**21(33)**	**14(33)**	**21(33)**	**12(33)**	**23(33)**	**15(33)**	**24(33)**	**12(33)**	**23(33)**
RJ																				
Exporters	9(12)	10(12)	11(12)	10(12)	10(12)	8(12)	9(12)	7(12)	8(12)	6(12)	8(12)	10(12)	8(12)	8(12)	7(12)	7(12)	5(12)	6(12)	5(12)	5(12)
Importers	3(8)	4(8)	3(8)	3(8)	2(8)	3(8)	4(8)	2(8)	6(8)	1(8)	4(8)	5(8)	2(8)	3(8)	2(8)	3(8)	1(8)	2(8)	5(8)	1(8)
Both (Hybrid)	12(13)	10(13)	8(13)	11(13)	5(13)	8(13)	6(13)	6(13)	7(13)	4(13)	7(13)	9(13)	6(13)	10(13)	5(13)	10(13)	5(13)	8(13)	6(13)	5(13)
Total	**24(33)**	**24(33)**	**22(33)**	**24(33)**	**17(33)**	**19(33)**	**19(33)**	**15(33)**	**21(33)**	**11(33)**	**19(33)**	**24(33)**	**16(33)**	**21(33)**	**14(33)**	**20(33)**	**11(33)**	**16(33)**	**16(33)**	**11(33)**
GS																				
Exporters	6(12)	12(12)	6(12)	10(12)	6(12)	9(12)	5(12)	8(12)	7(12)	5(12)	6(12)	11(12)	6(12)	7(12)	4(12)	8(12)	5(12)	7(12)	5(12)	5(12)
Importers	5(8)	3(8)	5(8)	1(8)	3(8)	2(8)	5(8)	1(8)	6(8)	2(8)	5(8)	4(8)	4(8)	2(8)	3(8)	3(8)	3(8)	2(8)	5(8)	2(8)
Both (Hybrid)	8(13)	11(13)	8(13)	6(13)	7(13)	5(13)	5(13)	5(13)	8(13)	6(13)	5(13)	9(13)	3(13)	9(13)	5(13)	6(13)	5(13)	7(13)	6(13)	4(13)
Total	**19(33)**	**26(33)**	**19(33)**	**17(33)**	**16(33)**	**16(33)**	**15(33)**	**14(33)**	**21(33)**	**13(33)**	**16(33)**	**24(33)**	**13(33)**	**18(33)**	**12(33)**	**17(33)**	**13(33)**	**16(33)**	**16(33)**	**11(33)**
Moody																				
Exporters	8(12)	12(12)	7(12)	11(12)	8(12)	11(12)	10(12)	7(12)	6(12)	5(12)	7(12)	11(12)	8(12)	12(12)	8(12)	10(12)	8(12)	9(12)	4(12)	8(12)
Importers	4(8)	6(8)	5(8)	7(8)	3(8)	5(8)	4(8)	5(8)	1(8)	4(8)	5(8)	5(8)	3(8)	4(8)	3(8)	3(8)	2(8)	4(8)	3(8)	4(8)
Both (Hybrid)	8(13)	11(13)	9(13)	11(13)	7(13)	10(13)	9(13)	4(13)	6(13)	2(13)	9(13)	9(13)	10(13)	8(13)	8(13)	7(13)	7(13)	7(13)	7(13)	6(13)
Total	**20(33)**	**29(33)**	**21(33)**	**29(33)**	**18(33)**	**26(33)**	**23(33)**	**16(33)**	**13(33)**	**11(33)**	**21(33)**	**25(33)**	**21(33)**	**24(33)**	**19(33)**	**20(33)**	**17(33)**	**20(33)**	**14(33)**	**18(33)**
TR																				
Exporters	8(12)	11(12)	9(12)	9(12)	10(12)	8(12)	6(12)	6(12)	6(12)	6(12)	9(12)	10(12)	9(12)	10(12)	6(12)	8(12)	4(12)	5(12)	3(12)	4(12)
Importers	5(8)	5(8)	5(8)	5(8)	3(8)	0(8)	3(8)	2(8)	4(8)	3(8)	3(8)	4(8)	4(8)	2(8)	4(8)	3(8)	3(8)	3(8)	5(8)	3(8)
Both (Hybrid)	7(13)	10(13)	7(13)	9(13)	5(13)	7(13)	7(13)	4(13)	7(13)	5(13)	7(13)	8(13)	6(13)	8(13)	4(13)	7(13)	4(13)	5(13)	5(13)	5(13)
Total	**20(33)**	**26(33)**	**21(33)**	**23(33)**	**18(33)**	**15(33)**	**16(33)**	**12(33)**	**17(33)**	**14(33)**	**19(33)**	**22(33)**	**19(33)**	**20(33)**	**14(33)**	**18(33)**	**11(33)**	**13(33)**	**13(33)**	**12(33)**
IMF non-fuel																				
Exporters	7(12)	12(12)	6(12)	12(12)	3(12)	10(12)	3(12)	11(12)	7(12)	6(12)	9(12)	10(12)	6(12)	11(12)	3(12)	10(12)	3(12)	9(12)	4(12)	7(12)
Importers	6(8)	7(8)	3(8)	6(8)	3(8)	5(8)	3(8)	4(8)	2(8)	5(8)	3(8)	4(8)	3(8)	2(8)	2(8)	2(8)	3(8)	3(8)	4(8)	3(8)
Both (Hybrid)	7(13)	12(13)	8(13)	12(13)	4(13)	10(13)	4(13)	9(13)	7(13)	7(13)	6(13)	9(13)	6(13)	9(13)	5(13)	7(13)	4(13)	6(13)	4(13)	5(13)
Total	**20(33)**	**31(33)**	**17(33)**	**30(33)**	**10(33)**	**25(33)**	**10(33)**	**24(33)**	**16(33)**	**18(33)**	**18(33)**	**23(33)**	**15(33)**	**22(33)**	**10(33)**	**19(33)**	**10(33)**	**18(33)**	**12(33)**	**15(33)**

Note: The table presents the number of countries for which time-varying causality has been detected for horizons $h\epsilon\{1,2,3,4,6\}$. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0: CP \nRightarrow EG$ (⇏ means "does not Granger-cause"). The number in brackets denotes the total number of countries considered in the analysis, for each group of countries.

Table B.27 Summary of the Time-varying Granger Causality Test Results for Economic Growth and World Commodity Prices

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
CRB						
Exporters	9(12)	6(12)	6(12)	5(12)	8(12)	8(12)
Importers	4(8)	2(8)	2(8)	1(8)	4(8)	4(8)
Both (Hybrid)	9(13)	6(13)	6(13)	3(13)	7(13)	8(13)
Total	**22(33)**	**14(33)**	**14(33)**	**9(33)**	**19(33)**	**20(33)**
RJ						
Exporters	7(12)	8(12)	8(12)	7(12)	8(12)	10(12)
Importers	5(8)	1(8)	4(8)	2(8)	1(8)	5(8)
Both (Hybrid)	9(13)	7(13)	9(13)	3(13)	7(13)	10(13)
Total	**21(33)**	**16(33)**	**21(33)**	**12(33)**	**16(33)**	**25(33)**
GS						
Exporters	7(12)	8(12)	9(12)	8(12)	8(12)	8(12)
Importers	3(8)	2(8)	3(8)	1(8)	1(8)	4(8)
Both (Hybrid)	8(13)	7(13)	8(13)	4(13)	7(13)	9(13)
Total	**18(33)**	**17(33)**	**20(33)**	**13(33)**	**16(33)**	**21(33)**
Moody						
Exporters	9(12)	4(12)	5(12)	7(12)	5(12)	8(12)
Importers	4(8)	2(8)	2(8)	3(8)	2(8)	3(8)
Both (Hybrid)	9(13)	6(13)	6(13)	3(13)	7(13)	7(13)
Total	**22(33)**	**12(33)**	**13(33)**	**13(33)**	**14(33)**	**18(33)**
TR						
Exporters	5(12)	7(12)	8(12)	8(12)	7(12)	9(12)
Importers	4(8)	2(8)	3(8)	1(8)	3(8)	5(8)
Both (Hybrid)	6(13)	7(13)	7(13)	3(13)	7(13)	8(13)
Total	**15(33)**	**16(33)**	**18(33)**	**12(33)**	**17(33)**	**22(33)**
IMF non-fuel						
Exporters	8(12)	7(12)	8(12)	7(12)	6(12)	9(12)
Importers	4(8)	2(8)	3(8)	2(8)	3(8)	4(8)
Both (Hybrid)	8(13)	7(13)	7(13)	4(13)	7(13)	8(13)
Total	**20(33)**	**16(33)**	**18(33)**	**13(33)**	**16(33)**	**21(33)**

Note: The table shows the number of countries for which the commodity-based model outperforms the benchmark one, based on regression-based forecasts. The number in brackets denotes the total number of countries considered in the analysis, for each group of countries.

Table B.28 Summary of the Out-of-Sample Forecasting Ability Test Results – Regression Based Forecast Models

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
CRB						
Exporters	1(12)	11(12)	11(12)	2(12)	11(12)	11(12)
Importers	0(8)	6(8)	6(8)	1(8)	6(8)	6(8)
Both (Hybrid)	4(13)	8(13)	9(13)	6(13)	8(13)	9(13)
Total	**5(33)**	**25(33)**	**26(33)**	**9(33)**	**25(33)**	**26(33)**
RJ						
Exporters	2(12)	11(12)	11(12)	2(12)	11(12)	11(12)
Importers	1(8)	5(8)	6(8)	3(8)	5(8)	6(8)
Both (Hybrid)	4(13)	9(13)	10(13)	4(13)	8(13)	10(13)
Total	**7(33)**	**25(33)**	**27(33)**	**9(33)**	**24(33)**	**27(33)**
GS						
Exporters	4(12)	11(12)	11(12)	4(12)	11(12)	11(12)
Importers	3(8)	5(8)	6(8)	3(8)	5(8)	6(8)
Both (Hybrid)	3(13)	6(13)	8(13)	4(13)	7(13)	9(13)
Total	**10(33)**	**22(33)**	**25(33)**	**11(33)**	**23(33)**	**26(33)**
Moody						
Exporters	1(12)	11(12)	11(12)	1(12)	11(12)	11(12)
Importers	1(8)	6(8)	6(8)	1(8)	6(8)	6(8)
Both (Hybrid)	4(13)	8(13)	9(13)	4(13)	7(13)	9(13)
Total	**6(33)**	**25(33)**	**26(33)**	**6(33)**	**24(33)**	**26(33)**
TR						
Exporters	3(12)	11(12)	11(12)	2(12)	11(12)	11(12)
Importers	0(8)	6(8)	6(8)	0(8)	6(8)	6(8)
Both (Hybrid)	4(13)	8(13)	10(13)	4(13)	8(13)	10(13)
Total	**7(33)**	**25(33)**	**27(33)**	**6(33)**	**25(33)**	**27(33)**
IMF non-fuel						
Exporters	3(12)	11(12)	11(12)	3(12)	11(12)	11(12)
Importers	1(8)	6(8)	6(8)	1(8)	6(8)	6(8)
Both (Hybrid)	4(13)	8(13)	9(13)	5(13)	8(13)	9(13)
Total	**8(33)**	**25(33)**	**26(33)**	**9(33)**	**25(33)**	**26(33)**

Note: The table shows the number of countries for which the commodity-based model outperforms the benchmark one, based on combination forecasts. The number in brackets denotes the total number of countries considered in the analysis, for each group of countries.

Table B.29 Summary of the Out-of-Sample Forecasting Ability Test Results – Combination Forecast Models

	Commodity Prices: Low-Frequency, Reuters/Jeffries	
	ADF with intercept, no trend	PP with intercept, no trend
Australia	-2.952**	-3.112**
Bahrain	-2.916**	-3.137**
Belgium	-5.390***	-3.296**
Bolivia	-5.676***	-2.672*
Canada	-2.952**	-3.112**
Chile	-5.346***	-3.477**
Czech Republic	-5.346***	-3.477**
Denmark	-5.577***	-3.871***
Dominican Republic	-4.750***	-3.6720***
Ecuador	-4.968***	-3.410**
Estonia	-5.390***	-3.296**
Hong Kong	-2.952**	-3.112**
Hungary	-5.390***	-3.296**
Iceland	-5.219***	-3.211**
Israel	-5.390***	-3.296**
Kazakhstan	-5.363***	-3.586***
Latvia	-5.390***	-3.296**
Luxembourg	-4.968***	-3.410**
Malta	-4.968***	-3.410**
Netherlands	-5.346***	-3.477**
New Zealand	-4.861***	-2.557
Norway	-2.952**	-3.112**
Peru	-2.952**	-3.112**
Philippines	-2.900**	-3.087**
Serbia	-5.346***	-3.477**
Seychelles	-2.952**	-3.112**
Singapore	-2.952**	-3.112**
Slovakia	-5.390***	-3.296**
Slovenia	-5.390***	-3.296**
South Africa	-2.952**	-3.112**
Thailand	-5.644***	-3.725***
Venezuela	-4.989***	-3.034**
Viet Nam	-4.654***	-3.221**

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.30 Results for Unit Root Tests for Commodity Prices, Reuters/Jeffries, Low-frequency

	Commodity Prices: Mixed-Frequency, Reuters/Jeffries					
	ADF with intercept, no trend			PP with intercept, no trend		
	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	-2.934**	-5.183**	-3.089***	-3.133**	-3.427**	-3.168**
Bahrain	-2.885**	-5.082**	-3.032***	-3.058**	-3.316**	-3.239**
Belgium	-5.339	-4.823***	-2.523***	-3.557***	-3.530***	-3.637***
Bolivia	-2.428*	-4.588	-2.624***	-2.634*	-2.913*	-2.710**
Canada	-2.934**	-5.183**	-3.089***	-3.133**	-3.427**	-3.168**
Chile	-5.539	-4.763***	-2.430***	-3.778***	-3.667***	-3.781***
Czech Republic	-5.539	-4.763***	-2.430***	-3.778***	-3.667***	-3.781***
Denmark	-2.289	-4.497	-2.500***	-3.932***	-4.080***	-4.076***
Dominican Republic	-2.374	-4.598	-2.547***	-3.753***	-3.858***	-3.803***
Ecuador	-4.643***	-4.429***	-4.198***	-3.527***	-3.573**	-3.734***
Estonia	-5.339	-4.823***	-2.523***	-3.557***	-3.530***	-3.637***
Hong Kong	-2.934**	-5.183**	-3.089***	-3.133**	-3.427**	-3.168**
Hungary	-5.339	-4.823***	-2.523***	-3.557***	-3.530***	-3.637***
Iceland	-4.610**	-4.636***	-3.437***	-2.843**	-3.465*	-3.052**
Israel	-5.339	-4.823***	-2.523***	-3.557***	-3.530***	-3.637***
Kazakhstan	-2.631*	-4.746*	-2.856***	-3.664***	-3.561***	-3.698***
Latvia	-5.339	-4.823***	-2.523***	-3.557***	-3.530***	-3.637***
Luxembourg	-4.643***	-4.429***	-4.198***	-3.527***	-3.573**	-3.734***
Malta	-4.643***	-4.429***	-4.198***	-3.527***	-3.573**	-3.734***
Netherlands	-5.539	-4.763***	-2.430***	-3.778***	-3.667***	-3.781***
New Zealand	-2.534*	-4.757	-2.730***	-2.715*	-3.069*	-2.797**
Norway	-2.934**	-5.183**	-3.089***	-3.133**	-3.427**	-3.168**
Peru	-2.934**	-5.183**	-3.089***	-3.133**	-3.427**	-3.168**
Philippines	-2.867**	-5.036*	-3.013***	-2.991**	-3.341**	-3.142**
Serbia	-5.539	-4.763***	-2.43***	-3.778***	-3.667***	-3.781***
Seychelles	-2.934**	-5.183**	-3.089***	-3.133**	-3.427**	-3.168**
Singapore	-2.934**	-5.183**	-3.089***	-3.133**	-3.427**	-3.168**
Slovakia	-5.339	-4.823***	-2.523***	-3.557***	-3.530***	-3.637***
Slovenia	-5.339	-4.823***	-2.523***	-3.557***	-3.530***	-3.637***
South Africa	-2.934**	-5.183**	-3.089***	-3.133**	-3.427**	-3.168**
Thailand	-2.325	-5.024	-2.547***	-3.794***	-3.730***	-3.848***
Venezuela	-5.225**	-4.408***	-3.330***	-2.668*	-3.278*	-2.884**
Viet Nam	-4.16***	-4.227***	-3.939***	-3.396**	-3.315**	-3.502**

Note: The table reports the test statistics obtained from the unit root tests for monthly commodity price series at a given month of each quarter period τ.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.31 Results for Unit Root Tests for Commodity Prices, Reuters/Jeffries, Mixed-frequency

	KPSS with intercept, no trend			
	CP(τ)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	0.132	0.133	0.128	0.129
Bahrain	0.139	0.143	0.134	0.134
Belgium	0.303	0.307	0.284	0.304
Bolivia	0.249	0.248	0.237	0.251
Canada	0.132	0.133	0.128	0.129
Chile	0.228	0.239	0.212	0.214
Czech Republic	0.228	0.239	0.212	0.214
Denmark	0.305	0.300	0.286	0.312
Dominican Republic	0.190	0.190	0.173	0.200
Ecuador	0.353*	0.368*	0.336	0.321
Estonia	0.303	0.307	0.284	0.304
Hong Kong	0.132	0.133	0.128	0.129
Hungary	0.303	0.307	0.284	0.304
Iceland	0.177	0.178	0.167	0.180
Israel	0.303	0.307	0.284	0.304
Kazakhstan	0.290	0.285	0.274	0.296
Latvia	0.303	0.307	0.284	0.304
Luxembourg	0.353*	0.368*	0.336	0.321
Malta	0.353*	0.368*	0.336	0.321
Netherlands	0.228	0.239	0.212	0.214
New Zealand	0.224	0.227	0.212	0.222
Norway	0.132	0.133	0.128	0.129
Peru	0.132	0.133	0.128	0.129
Philippines	0.137	0.141	0.129	0.132
Serbia	0.228	0.239	0.212	0.214
Seychelles	0.132	0.133	0.128	0.129
Singapore	0.132	0.133	0.128	0.129
Slovakia	0.303	0.307	0.284	0.304
Slovenia	0.303	0.307	0.284	0.304
South Africa	0.132	0.133	0.128	0.129
Thailand	0.342	0.344	0.323	0.342
Venezuela	0.129	0.118	0.116	0.151
Viet Nam	0.407*	0.361*	0.393*	0.436*

Note: The table reports the test statistics obtained from the KPSS unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.32 Results for KPSS Unit-Root Test, Reuters/Jeffries

	Commodity Prices: Low-Frequency, Goldman Sachs	
	ADF with intercept, no trend	PP with intercept, no trend
Australia	-4.056***	-3.443**
Bahrain	-4.013***	-3.564***
Belgium	-4.997***	-3.120**
Bolivia	-5.288***	-2.850*
Canada	-4.056***	-3.443**
Chile	-4.845***	-3.034**
Czech Republic	-4.845***	-3.034**
Denmark	-5.300***	-3.928***
Dominican Republic	-4.378***	-3.397**
Ecuador	-4.570***	-3.367**
Estonia	-4.997***	-3.120**
Hong Kong	-4.056***	-3.443**
Hungary	-4.997***	-3.120**
Iceland	-4.819***	-3.415**
Israel	-4.997***	-3.120**
Kazakhstan	-5.053***	-3.625***
Latvia	-4.997***	-3.120**
Luxembourg	-4.570***	-3.367**
Malta	-4.570***	-3.367**
Netherlands	-4.845***	-3.034**
New Zealand	-3.788***	-3.066**
Norway	-4.056***	-3.443**
Peru	-4.056***	-3.443**
Philippines	-4.010***	-3.408**
Serbia	-4.845***	-3.034**
Seychelles	-4.056***	-3.443**
Singapore	-4.056***	-3.443**
Slovakia	-4.997***	-3.120**
Slovenia	-4.997***	-3.120**
South Africa	-4.056***	-3.443**
Thailand	-5.151***	-3.752***
Venezuela	-4.541***	-3.163**
Viet Nam	-4.437***	-3.276**

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.33 Results for Unit Root Tests for Commodity Prices, Goldman Sachs, Low-frequency

	Commodity Prices: Mixed-Frequency, Goldman Sachs					
	ADF with intercept, no trend			PP with intercept, no trend		
	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	-3.244**	-4.459**	-3.363***	-3.934***	-4.162***	-3.809***
Bahrain	-3.179**	-4.386**	-3.319***	-3.952***	-4.129***	-3.914***
Belgium	-2.997**	-4.592**	-3.258***	-3.705***	-3.864***	-3.759***
Bolivia	-2.575***	-3.916	-4.153***	-3.197**	-3.296**	-3.153**
Canada	-3.244**	-4.459**	-3.363***	-3.934***	-4.162***	-3.809***
Chile	-2.855***	-4.414*	-4.352***	-3.640***	-3.761***	-3.663***
Czech Republic	-2.855***	-4.414*	-4.352***	-3.640***	-3.761***	-3.663***
Denmark	-2.449***	-3.694	-3.573***	-4.071***	-4.298***	-4.290***
Dominican Republic	-2.965***	-3.758**	-4.411***	-3.665***	-4.236***	-4.059***
Ecuador	-2.162***	-4.185	-4.153***	-3.534***	-3.587**	-3.689***
Estonia	-2.997**	-4.592**	-3.258***	-3.705***	-3.864***	-3.759***
Hong Kong	-3.244**	-4.459**	-3.363***	-3.934***	-4.162***	-3.809***
Hungary	-2.997**	-4.592**	-3.258***	-3.705***	-3.864***	-3.759***
Iceland	-3.029**	-4.425**	-3.307***	-3.522**	-3.429***	-3.325**
Israel	-2.997**	-4.592**	-3.258***	-3.705***	-3.864***	-3.759***
Kazakhstan	-3.035**	-4.616**	-3.311***	-3.792***	-3.912***	-3.810***
Latvia	-2.997**	-4.592**	-3.258***	-3.705***	-3.864***	-3.759***
Luxembourg	-2.162***	-4.185	-4.153***	-3.534***	-3.587**	-3.689***
Malta	-2.162***	-4.185	-4.153***	-3.534***	-3.587**	-3.689***
Netherlands	-2.855***	-4.414*	-4.352***	-3.640***	-3.761***	-3.663***
New Zealand	-2.817**	-4.219*	-3.018***	-3.505**	-3.669***	-3.361***
Norway	-3.244**	-4.459**	-3.363***	-3.934***	-4.162***	-3.809***
Peru	-3.244**	-4.459**	-3.363***	-3.934***	-4.162***	-3.809***
Philippines	-3.128**	-4.385**	-3.261***	-3.869***	-4.012***	-3.767***
Serbia	-2.855***	-4.414*	-4.352***	-3.640***	-3.761***	-3.663***
Seychelles	-3.244**	-4.459**	-3.363***	-3.934***	-4.162***	-3.809***
Singapore	-3.244**	-4.459**	-3.363***	-3.934***	-4.162***	-3.809***
Slovakia	-2.997**	-4.592**	-3.258***	-3.705***	-3.864***	-3.759***
Slovenia	-2.997**	-4.592**	-3.258***	-3.705***	-3.864***	-3.759***
South Africa	-3.244**	-4.459**	-3.363***	-3.934***	-4.162***	-3.809***
Thailand	-2.355**	-4.704	-3.478***	-3.893***	-4.098***	-3.953***
Venezuela	-3.007**	-4.149**	-3.123***	-3.295**	-3.188**	-3.093**
Viet Nam	-3.865***	-4.036***	-4.059***	-3.382***	-3.432**	-3.624**

Note: The table reports the test statistics obtained from the unit root tests for monthly commodity price series at a given month of each quarter period τ.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.34 Results for Unit Root Tests for Commodity Prices, Goldman Sachs, Mixed-frequency

	KPSS with intercept, no trend			
	CP(τ)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	0.168	0.169	0.162	0.163
Bahrain	0.175	0.183	0.167	0.166
Belgium	0.218	0.234	0.203	0.200
Bolivia	0.146	0.155	0.136	0.142
Canada	0.168	0.169	0.162	0.163
Chile	0.202	0.220	0.189	0.185
Czech Republic	0.202	0.220	0.189	0.185
Denmark	0.151	0.161	0.147	0.145
Dominican Republic	0.073	0.075	0.067	0.073
Ecuador	0.311	0.346	0.297	0.267
Estonia	0.218	0.234	0.203	0.200
Hong Kong	0.168	0.169	0.162	0.163
Hungary	0.218	0.234	0.203	0.200
Iceland	0.213	0.218	0.199	0.207
Israel	0.218	0.234	0.203	0.200
Kazakhstan	0.207	0.217	0.195	0.194
Latvia	0.218	0.234	0.203	0.200
Luxembourg	0.311	0.346	0.297	0.267
Malta	0.311	0.346	0.297	0.267
Netherlands	0.202	0.220	0.189	0.185
New Zealand	0.130	0.139	0.122	0.122
Norway	0.168	0.169	0.162	0.163
Peru	0.168	0.169	0.162	0.163
Philippines	0.160	0.169	0.152	0.151
Serbia	0.202	0.220	0.189	0.185
Seychelles	0.168	0.169	0.162	0.163
Singapore	0.168	0.169	0.162	0.163
Slovakia	0.218	0.234	0.203	0.200
Slovenia	0.218	0.234	0.203	0.200
South Africa	0.168	0.169	0.162	0.163
Thailand	0.172	0.183	0.161	0.162
Venezuela	0.156	0.142	0.139	0.185
Viet Nam	0.422*	0.398*	0.424*	0.414*

Note: The table reports the test statistics obtained from the KPSS unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.35 Results for KPSS Unit-Root Test, Goldman Sachs

	Commodity Prices: Low-Frequency, Moody’s	
	ADF with intercept, no trend	**PP with intercept, no trend**
Australia	-4.228***	-3.367**
Bahrain	-4.187***	-3.702***
Belgium	-4.063***	-3.190**
Bolivia	-2.430	-2.959**
Canada	-4.228***	-3.367**
Chile	-3.981***	-3.097**
Czech Republic	-3.981***	-3.097**
Denmark	-4.395***	-3.717***
Dominican Republic	-4.071***	-3.450**
Ecuador	-3.705***	-2.901*
Estonia	-4.063***	-3.190**
Hong Kong	-4.228***	-3.367**
Hungary	-4.063***	-3.190**
Iceland	-4.005***	-2.978**
Israel	-4.063***	-3.190**
Kazakhstan	-4.166***	-3.230**
Latvia	-4.063***	-3.190**
Luxembourg	-3.705***	-2.901*
Malta	-3.705***	-2.901*
Netherlands	-3.981***	-3.097**
New Zealand	-2.555	-2.882*
Norway	-4.228***	-3.367**
Peru	-4.228***	-3.367**
Philippines	-4.230***	-3.369**
Serbia	-3.981***	-3.097**
Seychelles	-4.228***	-3.367**
Singapore	-4.228***	-3.367**
Slovakia	-4.063***	-3.190**
Slovenia	-4.063***	-3.190**
South Africa	-4.228***	-3.367**
Thailand	-4.399***	-3.351**
Venezuela	-3.812***	-2.843*
Viet Nam	-3.593***	-2.807*

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.36 Results for Unit Root Tests for Commodity Prices, Moody’s, Low-frequency

	Commodity Prices: Mixed-Frequency, Moody's					
	ADF with intercept, no trend			PP with intercept, no trend		
	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	-2.687**	-2.785*	-3.074*	-3.957***	-3.550***	-4.05***
Bahrain	-2.621**	-2.749*	-2.969*	-3.922***	-3.812***	-4.019***
Belgium	-3.893**	-4.242***	-3.278***	-3.349***	-3.440**	-3.674**
Bolivia	-2.327*	-2.507	-2.663	-2.993**	-3.307**	-3.314**
Canada	-2.687**	-2.785*	-3.074*	-3.957***	-3.550***	-4.050***
Chile	-3.820**	-4.141***	-3.371***	-3.233***	-3.359**	-3.570**
Czech Republic	-3.820**	-4.141***	-3.371***	-3.233***	-3.359**	-3.570**
Denmark	-2.170	-2.370	-2.520	-3.648***	-4.058***	-4.028***
Dominican Republic	-1.966	-3.736	-2.316***	-3.359***	-3.722**	-3.754***
Ecuador	-3.551**	-3.912***	-3.198***	-2.556**	-3.181	-3.198**
Estonia	-3.893**	-4.242***	-3.278***	-3.349***	-3.440**	-3.674**
Hong Kong	-2.687**	-2.785*	-3.074*	-3.957***	-3.550***	-4.050***
Hungary	-3.893**	-4.242***	-3.278***	-3.349***	-3.440**	-3.674**
Iceland	-3.800**	-4.144***	-3.165***	-2.923**	-3.340**	-3.165**
Israel	-3.893**	-4.242***	-3.278***	-3.349***	-3.440**	-3.674**
Kazakhstan	-1.880	-3.416	-2.178**	-3.371***	-3.497**	-3.712**
Latvia	-3.893**	-4.242***	-3.278***	-3.349***	-3.440**	-3.674**
Luxembourg	-3.551**	-3.912***	-3.198***	-2.556**	-3.181	-3.198**
Malta	-3.551**	-3.912***	-3.198***	-2.556**	-3.181	-3.198**
Netherlands	-3.820**	-4.141***	-3.371***	-3.233***	-3.359**	-3.570**
New Zealand	-2.438*	-2.593	-2.790*	-3.071**	-3.350**	-3.390**
Norway	-2.687**	-2.785*	-3.074*	-3.957***	-3.550***	-4.050***
Peru	-2.687**	-2.785*	-3.074*	-3.957***	-3.550***	-4.050***
Philippines	-2.593**	-2.734*	-2.945*	-3.524***	-3.697***	-3.743***
Serbia	-3.820**	-4.141***	-3.371***	-3.233***	-3.359**	-3.570**
Seychelles	-2.687**	-2.785*	-3.074*	-3.957***	-3.550***	-4.050***
Singapore	-2.687**	-2.785*	-3.074*	-3.957***	-3.550***	-4.050***
Slovakia	-3.893**	-4.242***	-3.278***	-3.349***	-3.440**	-3.674**
Slovenia	-3.893**	-4.242***	-3.278***	-3.349***	-3.440**	-3.674**
South Africa	-2.687**	-2.785*	-3.074*	-3.957***	-3.550***	-4.050***
Thailand	-1.960	-2.173	-2.291	-3.506***	-3.625***	-3.835***
Venezuela	-1.623**	-4.016	-3.028***	-2.773**	-3.228*	-3.028**
Viet Nam	-3.474**	-3.751**	-3.108***	-2.510**	-3.050	-3.108**

Note: The table reports the test statistics obtained from the unit root tests for monthly commodity price series at a given month of each quarter period τ.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.37 Results for Unit Root Tests for Commodity Prices, Moody's, Mixed-frequency

	KPSS with intercept, no trend			
	CP(τ)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	0.379*	0.391*	0.376*	0.361*
Bahrain	0.149	0.361*	0.336	0.328
Belgium	0.218	0.148	0.143	0.138
Bolivia	0.135	0.125	0.120	0.120
Canada	0.145	0.391*	0.376*	0.361*
Chile	0.145	0.152	0.148	0.142
Czech Republic	0.110	0.152	0.148	0.142
Denmark	0.218	0.104	0.101	0.099
Dominican Republic	0.149	0.103	0.099	0.098
Ecuador	0.145	0.211	0.236	0.215
Estonia	0.149	0.148	0.143	0.138
Hong Kong	0.098	0.391*	0.376*	0.361*
Hungary	0.188	0.148	0.143	0.138
Iceland	0.277	0.176	0.177	0.167
Israel	0.218	0.148	0.143	0.138
Kazakhstan	0.145	0.151	0.149	0.151
Latvia	0.173	0.148	0.143	0.138
Luxembourg	0.345	0.211	0.236	0.215
Malta	0.276	0.211	0.236	0.215
Netherlands	0.283	0.152	0.148	0.142
New Zealand	0.218	0.134	0.131	0.135
Norway	0.122	0.391*	0.376*	0.361*
Peru	0.311	0.391*	0.376*	0.361*
Philippines	0.379*	0.325	0.300	0.301
Serbia	0.110	0.152	0.148	0.142
Seychelles	0.145	0.391*	0.376*	0.361*
Singapore	0.145	0.391*	0.376*	0.361*
Slovakia	0.152	0.148	0.143	0.138
Slovenia	0.379*	0.148	0.143	0.138
South Africa	0.101	0.391*	0.376*	0.361*
Thailand	0.149	0.107	0.109	0.106
Venezuela	0.379*	0.188	0.191	0.179
Viet Nam	0.145	0.262	0.292	0.287

Note: The table reports the test statistics obtained from the KPSS unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.38 Results for KPSS Unit-Root Test, Moody's

	Commodity Prices: Low-Frequency, Thompson Reuters	
	ADF with intercept, no trend	PP with intercept, no trend
Australia	-2.989**	-3.075**
Bahrain	-2.962**	-3.136**
Belgium	-5.364***	-3.445**
Bolivia	-4.504***	-2.893**
Canada	-2.989**	-3.075**
Chile	-5.209***	-3.354**
Czech Republic	-5.209***	-3.354**
Denmark	-5.618***	-3.329**
Dominican Republic	-4.705***	-3.111**
Ecuador	-4.772***	-2.994**
Estonia	-5.364***	-3.445**
Hong Kong	-2.989**	-3.075**
Hungary	-5.364***	-3.445**
Iceland	-5.197***	-3.346**
Israel	-5.364***	-3.445**
Kazakhstan	-5.449***	-3.529***
Latvia	-5.364***	-3.445**
Luxembourg	-4.772***	-2.994**
Malta	-4.772***	-2.994**
Netherlands	-5.209***	-3.354**
New Zealand	-4.651***	-2.699*
Norway	-2.989**	-3.075**
Peru	-2.989**	-3.075**
Philippines	-2.947**	-3.055**
Serbia	-5.209***	-3.354**
Seychelles	-2.989**	-3.075**
Singapore	-2.989**	-3.075**
Slovakia	-5.364***	-3.445**
Slovenia	-5.364***	-3.445**
South Africa	-2.989**	-3.075**
Thailand	-5.609***	-3.686***
Venezuela	-4.958***	-3.161**
Viet Nam	-4.584***	-2.927**

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.39 Results for Unit Root Tests for Commodity Prices, Thompson Reuters, Low-frequency

	Commodity Prices: Mixed-Frequency, Thompson Reuters					
	ADF with intercept, no trend			PP with intercept, no trend		
	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	-3.026**	-3.241**	-3.097**	-3.320**	-3.803**	-3.369***
Bahrain	-2.986**	-3.198**	-3.038**	-3.294**	-3.616**	-3.447***
Belgium	-5.101***	-4.698***	-4.029***	-3.145***	-3.454**	-3.841**
Bolivia	-2.597*	-4.171*	-2.663***	-3.002**	-3.297**	-3.008**
Canada	-3.026**	-3.241**	-3.097**	-3.320**	-3.803**	-3.369***
Chile	-5.034***	-4.546***	-3.974***	-3.046***	-3.355**	-3.751**
Czech Republic	-5.034***	-4.546***	-3.974***	-3.046***	-3.355**	-3.751**
Denmark	-2.386	-4.085	-2.46***	-3.499***	-4.257***	-4.067***
Dominican Republic	-4.100***	-4.146***	-4.821***	-3.673***	-3.641***	-3.845***
Ecuador	-4.087*	-4.234***	-2.836***	-3.250**	-3.510**	-3.271**
Estonia	-5.101***	-4.698***	-4.029***	-3.145***	-3.454**	-3.841**
Hong Kong	-3.026**	-3.241**	-3.097**	-3.320**	-3.803**	-3.369***
Hungary	-5.101***	-4.698***	-4.029***	-3.145***	-3.454**	-3.841**
Iceland	-4.381***	-4.500***	-4.121***	-2.940***	-3.696**	-3.532***
Israel	-5.101***	-4.698***	-4.029***	-3.145***	-3.454**	-3.841**
Kazakhstan	-2.910***	-3.060**	-4.071**	-3.240***	-3.519**	-3.918***
Latvia	-5.101***	-4.698***	-4.029***	-3.145***	-3.454**	-3.841**
Luxembourg	-4.087*	-4.234***	-2.836***	-3.250**	-3.510**	-3.271**
Malta	-4.087*	-4.234***	-2.836***	-3.250**	-3.510**	-3.271**
Netherlands	-5.034***	-4.546***	-3.974***	-3.046***	-3.355**	-3.751**
New Zealand	-2.647*	-4.232*	-2.724***	-2.929**	-3.398**	-2.994**
Norway	-3.026**	-3.241**	-3.097**	-3.320**	-3.803**	-3.369***
Peru	-3.026**	-3.241**	-3.097**	-3.320**	-3.803**	-3.369***
Philippines	-2.970**	-3.187**	-3.021**	-3.195**	-3.663**	-3.300***
Serbia	-5.034***	-4.546***	-3.974***	-3.046***	-3.355**	-3.751**
Seychelles	-3.026**	-3.241**	-3.097**	-3.320**	-3.803**	-3.369***
Singapore	-3.026**	-3.241**	-3.097**	-3.320**	-3.803**	-3.369***
Slovakia	-5.101***	-4.698***	-4.029***	-3.145***	-3.454**	-3.841**
Slovenia	-5.101***	-4.698***	-4.029***	-3.145***	-3.454**	-3.841**
South Africa	-3.026**	-3.241**	-3.097**	-3.320**	-3.803**	-3.369***
Thailand	-2.271	-3.821	-2.343***	-3.387***	-3.680**	-3.904***
Venezuela	-4.171***	-4.286***	-3.967***	-3.083**	-3.536**	-3.360***
Viet Nam	-3.895***	-4.064***	-3.984***	-3.223**	-3.372**	-3.216**

Note: The table reports the test statistics obtained from the unit root tests for monthly commodity price series at a given month of each quarter period τ.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.40 Results for Unit Root Tests for Commodity Prices, Thompson Reuters, Mixed-frequency

	KPSS with intercept, no trend			
	CP(τ)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	0.234	0.231	0.225	0.239
Bahrain	0.262	0.274	0.249	0.255
Belgium	0.126	0.127	0.123	0.121
Bolivia	0.115	0.117	0.112	0.117
Canada	0.234	0.231	0.225	0.239
Chile	0.136	0.137	0.133	0.131
Czech Republic	0.136	0.137	0.133	0.131
Denmark	0.102	0.104	0.102	0.112
Dominican Republic	0.078	0.079	0.072	0.077
Ecuador	0.247	0.247	0.248	0.227
Estonia	0.126	0.127	0.123	0.121
Hong Kong	0.234	0.231	0.225	0.239
Hungary	0.126	0.127	0.123	0.121
Iceland	0.146	0.148	0.144	0.141
Israel	0.126	0.127	0.123	0.121
Kazakhstan	0.123	0.124	0.120	0.119
Latvia	0.126	0.127	0.123	0.121
Luxembourg	0.247	0.247	0.248	0.227
Malta	0.247	0.247	0.248	0.227
Netherlands	0.136	0.137	0.133	0.131
New Zealand	0.108	0.107	0.105	0.112
Norway	0.234	0.231	0.225	0.239
Peru	0.234	0.231	0.225	0.239
Philippines	0.241	0.255	0.226	0.237
Serbia	0.136	0.137	0.133	0.131
Seychelles	0.234	0.231	0.225	0.239
Singapore	0.234	0.231	0.225	0.239
Slovakia	0.126	0.127	0.123	0.121
Slovenia	0.126	0.127	0.123	0.121
South Africa	0.234	0.231	0.225	0.239
Thailand	0.114	0.115	0.111	0.113
Venezuela	0.148	0.145	0.139	0.153
Viet Nam	0.341	0.305	0.339	0.355*

Note: The table reports the test statistics obtained from the KPSS unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.41 Results for KPSS Unit-Root Test, Thompson Reuters

	Commodity Prices: Low-Frequency, IMF	
	ADF with intercept, no trend	PP with intercept, no trend
Australia	-3.920***	-3.233**
Bahrain	-3.879***	-3.407**
Belgium	-3.095**	-3.353**
Bolivia	-3.333**	-2.705*
Canada	-3.920***	-3.233**
Chile	-3.401**	-3.263**
Czech Republic	-3.401**	-3.263**
Denmark	-3.132**	-3.895***
Dominican Republic	-2.986**	-3.642***
Ecuador	-4.251***	-3.168**
Estonia	-3.095**	-3.353**
Hong Kong	-3.920***	-3.233**
Hungary	-3.095**	-3.353**
Iceland	-4.376***	-3.362**
Israel	-3.095**	-3.353**
Kazakhstan	-2.944**	-3.402**
Latvia	-3.095**	-3.353**
Luxembourg	-4.251***	-3.168**
Malta	-4.251***	-3.168**
Netherlands	-3.401**	-3.263**
New Zealand	-3.408**	-3.184**
Norway	-3.920***	-3.233**
Peru	-3.920***	-3.233**
Philippines	-3.863***	-3.243**
Serbia	-3.401**	-3.263**
Seychelles	-3.920***	-3.233**
Singapore	-3.920***	-3.233**
Slovakia	-3.095**	-3.353**
Slovenia	-3.095**	-3.353**
South Africa	-3.920***	-3.233**
Thailand	-3.091**	-3.543***
Venezuela	-4.172***	-3.233**
Viet Nam	-4.258***	-3.065**

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.42 Results for Unit Root Tests for Commodity Prices, IMF, Low-frequency

	Commodity Prices: Mixed-Frequency, IMF					
	ADF with intercept, no trend			PP with intercept, no trend		
	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	-3.661***	-3.761***	-3.893***	-3.607**	-3.240***	-3.314**
Bahrain	-3.647***	-3.707***	-3.842***	-3.746***	-3.361***	-3.570**
Belgium	-3.693***	-3.931***	-3.830***	-3.751**	-3.528***	-3.399***
Bolivia	-4.188**	-3.168***	-3.323**	-3.101*	-2.861**	-2.687*
Canada	-3.661***	-3.761***	-3.893***	-3.607**	-3.240***	-3.314**
Chile	-4.097***	-4.044***	-4.099***	-3.536**	-3.449***	-2.928**
Czech Republic	-4.097***	-4.044***	-4.099***	-3.536**	-3.449***	-2.928**
Denmark	-2.426**	-2.943	-3.147**	-4.140***	-3.874***	-3.769***
Dominican Republic	-2.690**	-2.754*	-3.045*	-3.848***	-3.660***	-3.549***
Ecuador	-4.056**	-3.936***	-3.508***	-3.348**	-3.344**	-3.167**
Estonia	-3.693***	-3.931***	-3.830***	-3.751**	-3.528***	-3.399***
Hong Kong	-3.661***	-3.761***	-3.893***	-3.607**	-3.240***	-3.314**
Hungary	-3.693***	-3.931***	-3.830***	-3.751**	-3.528***	-3.399***
Iceland	-4.189***	-4.185***	-4.225***	-3.463**	-3.541**	-3.368***
Israel	-3.693***	-3.931***	-3.830***	-3.751**	-3.528***	-3.399***
Kazakhstan	-3.736**	-3.961***	-2.967***	-3.783**	-3.587***	-3.454***
Latvia	-3.693***	-3.931***	-3.830***	-3.751**	-3.528***	-3.399***
Luxembourg	-4.056**	-3.936***	-3.508***	-3.348**	-3.344**	-3.167**
Malta	-4.056**	-3.936***	-3.508***	-3.348**	-3.344**	-3.167**
Netherlands	-4.097***	-4.044***	-4.099***	-3.536**	-3.449***	-2.928**
New Zealand	-2.710**	-3.275*	-3.393**	-3.487**	-3.547**	-3.133***
Norway	-3.661***	-3.761***	-3.893***	-3.607**	-3.240***	-3.314**
Peru	-3.661***	-3.761***	-3.893***	-3.607**	-3.240***	-3.314**
Philippines	-3.632***	-3.690***	-3.820***	-3.624**	-3.279***	-3.227**
Serbia	-4.097***	-4.044***	-4.099***	-3.536**	-3.449***	-2.928**
Seychelles	-3.661***	-3.761***	-3.893***	-3.607**	-3.240***	-3.314**
Singapore	-3.661***	-3.761***	-3.893***	-3.607**	-3.240***	-3.314**
Slovakia	-3.693***	-3.931***	-3.830***	-3.751**	-3.528***	-3.399***
Slovenia	-3.693***	-3.931***	-3.830***	-3.751**	-3.528***	-3.399***
South Africa	-3.661***	-3.761***	-3.893***	-3.607**	-3.240***	-3.314**
Thailand	-2.771**	-2.813*	-3.135*	-3.931***	-3.718***	-3.611***
Venezuela	-4.040***	-3.996***	-4.002***	-3.395**	-3.392**	-3.189**
Viet Nam	-4.022***	-3.867***	-4.027***	-3.244**	-3.222**	-3.062**

Note: The table reports the test statistics obtained from the unit root tests for monthly commodity price series at a given month of each quarter period τ.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.43 Results for Unit Root Tests for Commodity Prices, IMF, Mixed-frequency

	KPSS with intercept, no trend			
	CP(τ)	CP(τ,1)	CP(τ,2)	CP(τ,3)
Australia	0.160	0.156	0.165	0.154
Bahrain	0.153	0.160	0.152	0.143
Belgium	0.117	0.124	0.115	0.110
Bolivia	0.103	0.107	0.101	0.100
Canada	0.160	0.156	0.165	0.154
Chile	0.128	0.133	0.131	0.118
Czech Republic	0.128	0.133	0.131	0.118
Denmark	0.089	0.094	0.088	0.083
Dominican Republic	0.092	0.096	0.091	0.087
Ecuador	0.183	0.193	0.188	0.162
Estonia	0.117	0.124	0.115	0.110
Hong Kong	0.160	0.156	0.165	0.154
Hungary	0.117	0.124	0.115	0.110
Iceland	0.149	0.157	0.151	0.137
Israel	0.117	0.124	0.115	0.110
Kazakhstan	0.124	0.127	0.120	0.122
Latvia	0.117	0.124	0.115	0.110
Luxembourg	0.183	0.193	0.188	0.162
Malta	0.183	0.193	0.188	0.162
Netherlands	0.128	0.133	0.131	0.118
New Zealand	0.094	0.093	0.091	0.096
Norway	0.160	0.156	0.165	0.154
Peru	0.160	0.156	0.165	0.154
Philippines	0.137	0.142	0.133	0.130
Serbia	0.128	0.133	0.131	0.118
Seychelles	0.160	0.156	0.165	0.154
Singapore	0.160	0.156	0.165	0.154
Slovakia	0.117	0.124	0.115	0.110
Slovenia	0.117	0.124	0.115	0.110
South Africa	0.160	0.156	0.165	0.154
Thailand	0.097	0.102	0.097	0.090
Venezuela	0.161	0.168	0.164	0.149
Viet Nam	0.220	0.226	0.221	0.202

Note: The table reports the test statistics obtained from the KPSS unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table B.44 Results for KPSS Unit-Root Test, IMF

	Panel A: Mixed-frequency model										Panel B: Low-frequency model									
	horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6		horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Exporters																				
Australia	0.152	0.656	0.245	0.354	0.728	0.563	0.938	0.833	0.811	0.197	0.612	0.378	0.564	0.297	0.269	0.654	0.960	0.682	0.592	0.608
Bolivia	0.489	0.152	0.079	0.254	0.100	0.978	0.018	0.904	0.017	0.974	0.134	0.012	0.009	0.358	0.016	0.868	0.014	0.674	0.004	0.932
Canada	0.504	0.176	0.014	0.486	0.570	0.416	0.236	0.450	0.103	0.880	0.419	0.010	0.071	0.031	0.061	0.001	0.164	0.371	0.650	0.584
Chile	0.020	0.448	0.204	0.340	0.454	0.636	0.634	0.880	0.179	0.914	0.143	0.094	0.298	0.033	0.537	0.209	0.457	0.599	0.437	0.905
Denmark	0.783	0.044	0.268	0.380	0.478	0.573	0.652	0.457	0.041	0.649	0.917	0.037	0.534	0.352	0.776	0.063	0.958	0.328	0.655	0.517
Ecuador	0.597	0.009	0.619	0.016	0.724	0.068	0.383	0.041	0.929	0.925	0.658	0.020	0.272	0.003	0.771	0.009	0.732	0.038	0.406	0.716
Kazakhstan	0.104	0.013	0.205	0.025	0.279	0.203	0.291	0.304	0.871	0.674	0.187	0.004	0.393	0.284	0.425	0.507	0.994	0.681	0.781	0.551
New Zealand	0.163	0.251	0.370	0.170	0.020	0.347	0.317	0.050	0.315	0.078	0.138	0.134	0.406	0.421	0.062	0.020	0.027	0.074	0.404	0.012
Norway	0.166	0.583	0.065	0.434	0.163	0.317	0.115	0.220	0.450	0.053	0.190	0.134	0.005	0.157	0.010	0.763	0.126	0.187	0.377	0.729
Peru	0.294	0.447	0.406	0.507	0.153	0.720	0.758	0.903	0.311	0.760	0.531	0.552	0.166	0.269	0.055	0.951	0.541	0.976	0.257	0.897
South Africa	0.848	0.045	0.052	0.503	0.477	0.047	0.402	0.174	0.313	0.600	0.292	0.004	0.036	0.030	0.002	0.001	0.152	0.052	0.482	0.694
Venezuela	0.456	0.001	0.017	0.005	0.133	0.036	0.107	0.066	0.648	0.933	0.077	0.002	0.021	0.005	0.256	0.003	0.435	0.021	0.525	0.739
Importers																				
Czech Republic	0.691	0.419	0.525	0.529	0.793	0.410	0.857	0.530	0.773	0.574	0.730	0.016	0.494	0.503	0.313	0.008	0.349	0.212	0.597	0.445
Dominican Republic	0.944	0.049	0.506	0.592	0.102	0.309	0.186	0.163	0.139	0.573	0.967	0.103	0.675	0.202	0.282	0.155	0.190	0.315	0.069	0.105
Hungary	0.847	0.330	0.703	0.202	0.044	0.378	0.746	0.738	0.103	0.778	0.787	0.067	0.401	0.217	0.032	0.028	0.644	0.530	0.481	0.838
Luxembourg	0.601	0.881	0.791	0.845	0.187	0.871	0.795	0.732	0.194	0.994	0.441	0.659	0.295	0.506	0.285	0.842	0.457	0.913	0.168	0.534
Malta	0.170	0.346	0.013	0.300	0.003	0.487	0.006	0.041	0.042	0.125	0.097	0.186	0.096	0.058	0.091	0.259	0.060	0.164	0.015	0.154
Philippines	0.697	0.036	0.291	0.078	0.199	0.018	0.129	0.058	0.122	0.430	0.453	0.013	0.037	0.081	0.001	0.001	0.160	0.062	0.160	0.207
Slovakia	0.399	0.748	0.010	0.271	0.314	0.789	0.610	0.840	0.058	0.922	0.773	0.109	0.848	0.075	0.846	0.538	0.958	0.290	0.894	0.565
Slovenia	0.338	0.057	0.244	0.687	0.255	0.632	0.478	0.861	0.544	0.931	0.439	0.102	0.537	0.182	0.900	0.136	0.583	0.442	0.990	0.482
Both (Hybrid)																				
Bahrain	0.796	0.099	0.326	0.785	0.842	0.484	0.135	0.554	0.312	0.093	0.095	0.244	0.039	0.151	0.407	0.177	0.363	0.044	0.837	0.104
Belgium	0.340	0.123	0.541	0.484	0.599	0.341	0.771	0.524	0.595	0.690	0.082	0.581	0.488	0.523	0.674	0.162	0.787	0.616	0.964	0.714
Hong Kong	0.696	0.001	0.850	0.005	0.707	0.065	0.816	0.020	0.486	0.374	0.116	0.011	0.496	0.480	0.002	0.001	0.246	0.070	0.292	0.118
Estonia	0.109	0.003	0.039	0.063	0.799	0.297	0.811	0.311	0.021	0.669	0.950	0.002	0.923	0.058	0.311	0.025	0.393	0.356	0.743	0.971
Iceland	0.609	0.645	0.617	0.851	0.972	0.254	0.899	0.554	0.222	0.531	0.599	0.970	0.793	0.799	0.650	0.586	0.716	0.598	0.768	0.715
Israel	0.705	0.021	0.570	0.023	0.141	0.135	0.400	0.144	0.568	0.079	0.227	0.040	0.166	0.077	0.203	0.051	0.193	0.082	0.216	0.088
Latvia	0.011	0.001	0.030	0.001	0.221	0.055	0.439	0.211	0.514	0.894	0.461	0.012	0.196	0.001	0.235	0.009	0.053	0.079	0.086	0.393
Netherlands	0.391	0.083	0.307	0.073	0.277	0.200	0.508	0.227	0.902	0.271	0.669	0.016	0.380	0.095	0.473	0.001	0.293	0.038	0.907	0.081
Serbia	0.053	0.009	0.091	0.095	0.408	0.087	0.059	0.585	0.484	0.729	0.096	0.004	0.161	0.044	0.573	0.098	0.618	0.215	0.404	0.914
Seychelles	0.164	0.001	0.523	0.516	0.493	0.512	0.109	0.223	0.005	0.929	0.477	0.003	0.436	0.408	0.376	0.432	0.002	0.317	0.027	0.792
Singapore	0.003	0.006	0.042	0.019	0.072	0.410	0.106	0.273	0.877	0.244	0.007	0.140	0.308	0.421	0.001	0.080	0.095	0.123	0.768	0.234
Thailand	0.007	0.004	0.002	0.133	0.019	0.106	0.378	0.156	0.066	0.135	0.026	0.002	0.022	0.058	0.035	0.013	0.034	0.128	0.216	0.069
Viet Nam	0.028	0.788	0.501	0.506	0.459	0.789	0.297	0.283	0.699	0.354	0.439	0.511	0.316	0.279	0.519	0.888	0.314	0.990	0.747	0.501

Note: The table reports the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizons $h \epsilon \{1,2,3,4,6\}$. *"CP" denotes the commodity prices, while "EG" denotes the economic growth variables.* $H_0: CP \nRightarrow EG$ *(⇏ means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with* $N = 999$ *replications. All variables are mean-centred and annual log-differenced.*

Table B.45 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices, Reuters/Jeffries

	Panel A: Mixed-frequency model										Panel B: Low-frequency model									
	horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6		horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Exporters																				
Australia	0.602	0.744	0.390	0.495	0.651	0.898	0.913	0.983	0.947	0.962	0.885	0.452	0.479	0.509	0.62	0.661	0.998	0.905	0.980	0.942
Bolivia	0.553	0.110	0.221	0.246	0.294	0.954	0.136	0.453	0.176	0.995	0.329	0.024	0.115	0.374	0.045	0.947	0.036	0.295	0.078	0.900
Canada	0.493	0.065	0.065	0.206	0.498	0.493	0.179	0.502	0.172	0.387	0.941	0.200	0.645	0.122	0.528	0.484	0.797	0.210	0.605	0.199
Chile	0.065	0.549	0.074	0.175	0.590	0.069	0.693	0.595	0.654	0.838	0.196	0.224	0.405	0.022	0.723	0.225	0.763	0.560	0.769	0.690
Denmark	0.995	0.064	0.816	0.321	0.745	0.403	0.470	0.077	0.469	0.101	0.965	0.194	0.597	0.244	0.571	0.103	0.881	0.167	0.976	0.033
Ecuador	0.334	0.012	0.701	0.013	0.893	0.068	0.549	0.055	0.774	0.832	0.579	0.023	0.372	0.007	0.920	0.022	0.953	0.041	0.581	0.410
Kazakhstan	0.026	0.012	0.083	0.774	0.167	0.669	0.700	0.487	0.203	0.762	0.047	0.007	0.185	0.030	0.479	0.095	0.884	0.852	0.278	0.704
New Zealand	0.468	0.006	0.090	0.002	0.329	0.005	0.858	0.030	0.701	0.069	0.206	0.001	0.176	0.001	0.137	0.001	0.375	0.002	0.848	0.419
Norway	0.458	0.636	0.154	0.874	0.082	0.143	0.151	0.050	0.074	0.087	0.256	0.630	0.050	0.744	0.037	0.546	0.038	0.748	0.026	0.925
Peru	0.190	0.280	0.183	0.478	0.397	0.421	0.566	0.729	0.428	0.492	0.076	0.930	0.049	0.972	0.248	0.729	0.598	0.718	0.139	0.468
South Africa	0.680	0.259	0.078	0.499	0.790	0.206	0.298	0.431	0.684	0.800	0.130	0.118	0.001	0.474	0.135	0.040	0.109	0.318	0.435	0.787
Venezuela	0.140	0.002	0.139	0.004	0.378	0.107	0.471	0.132	0.125	0.889	0.098	0.001	0.004	0.001	0.057	0.001	0.123	0.008	0.843	0.536
Importers																				
Czech Republic	0.883	0.324	0.493	0.516	0.973	0.207	0.985	0.429	0.750	0.772	0.860	0.037	0.496	0.474	0.770	0.016	0.599	0.049	0.871	0.283
Dominican Republic	0.847	0.037	0.775	0.842	0.159	0.622	0.427	0.571	0.487	0.776	0.512	0.036	0.399	0.204	0.244	0.371	0.399	0.622	0.574	0.244
Hungary	0.657	0.432	0.572	0.385	0.134	0.510	0.622	0.674	0.135	0.617	0.690	0.189	0.472	0.070	0.003	0.020	0.463	0.430	0.646	0.377
Luxembourg	0.801	0.729	0.032	0.553	0.091	0.901	0.052	0.999	0.248	0.923	0.538	0.644	0.445	0.314	0.292	0.842	0.162	0.864	0.240	0.617
Malta	0.383	0.503	0.005	0.257	0.011	0.240	0.002	0.056	0.030	0.210	0.039	0.153	0.040	0.076	0.032	0.174	0.021	0.097	0.026	0.109
Philippines	0.184	0.309	0.346	0.251	0.632	0.464	0.078	0.261	0.439	0.347	0.422	0.049	0.253	0.067	0.013	0.004	0.086	0.091	0.403	0.218
Slovakia	0.165	0.413	0.043	0.446	0.304	0.830	0.942	0.894	0.373	0.807	0.982	0.310	0.702	0.309	0.788	0.907	0.919	0.915	0.786	0.343
Slovenia	0.019	0.152	0.322	0.486	0.670	0.688	0.496	0.800	0.606	0.951	0.132	0.021	0.403	0.280	0.993	0.114	0.729	0.324	0.998	0.572
Both (Hybrid)																				
Bahrain	0.103	0.973	0.456	0.860	0.182	0.323	0.056	0.281	0.143	0.043	0.071	0.969	0.458	0.377	0.839	0.076	0.625	0.048	0.489	0.042
Belgium	0.005	0.329	0.462	0.506	0.714	0.324	0.809	0.659	0.903	0.762	0.034	0.068	0.482	0.390	0.775	0.199	0.807	0.343	0.996	0.624
Hong Kong	0.185	0.055	0.182	0.969	0.128	0.762	0.460	0.750	0.179	0.652	0.252	0.228	0.592	0.532	0.230	0.396	0.716	0.755	0.575	0.205
Estonia	0.099	0.011	0.520	0.116	0.795	0.192	0.208	0.695	0.651	0.786	0.674	0.014	0.693	0.128	0.760	0.381	0.798	0.148	0.133	0.472
Iceland	0.644	0.796	0.138	0.976	0.139	0.404	0.661	0.838	0.808	0.295	0.544	0.455	0.791	0.521	0.821	0.198	0.957	0.989	0.829	0.432
Israel	0.423	0.001	0.312	0.028	0.199	0.234	0.645	0.229	0.375	0.321	0.040	0.183	0.081	0.025	0.006	0.220	0.140	0.206	0.321	0.352
Latvia	0.157	0.003	0.110	0.016	0.504	0.120	0.780	0.451	0.708	0.900	0.638	0.003	0.311	0.001	0.634	0.040	0.235	0.127	0.216	0.902
Netherlands	0.787	0.101	0.794	0.113	0.685	0.532	0.767	0.594	0.660	0.750	0.831	0.068	0.909	0.112	0.966	0.353	0.730	0.295	0.961	0.321
Serbia	0.162	0.050	0.042	0.081	0.393	0.083	0.142	0.526	0.214	0.659	0.137	0.015	0.340	0.012	0.530	0.212	0.580	0.276	0.302	0.521
Seychelles	0.306	0.009	0.108	0.064	0.500	0.494	0.571	0.272	0.080	0.876	0.487	0.001	0.465	0.033	0.512	0.495	0.149	0.112	0.165	0.832
Singapore	0.002	0.009	0.007	0.023	0.125	0.411	0.090	0.635	0.738	0.589	0.008	0.216	0.002	0.350	0.004	0.341	0.225	0.397	0.800	0.152
Thailand	0.222	0.018	0.017	0.012	0.150	0.141	0.692	0.258	0.185	0.112	0.048	0.005	0.027	0.004	0.068	0.060	0.174	0.162	0.175	0.147
Viet Nam	0.019	0.225	0.601	0.144	0.237	0.324	0.742	0.766	0.839	0.361	0.302	0.015	0.252	0.009	0.697	0.685	0.835	0.484	0.847	0.600

Note: The table reports the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizons h$\epsilon\{1,2,3,4,6\}$*. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables.* $H_0: CP \nRightarrow EG$ *(*$\nRightarrow$ *means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with* $N = 999$ *replications. All variables are mean-centred and annual log-differenced.*

Table B.46 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices, Goldman Sachs

	Panel A: Mixed-frequency model										Panel B: Low-frequency model									
	horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6		horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG
Exporters																				
Australia	0.727	0.188	0.728	0.545	0.745	0.870	0.218	0.967	0.565	0.498	0.808	0.399	0.450	0.290	0.192	0.239	0.462	0.718	0.484	0.690
Bolivia	0.113	0.162	0.032	0.097	0.093	0.098	0.073	0.419	0.206	0.248	0.005	0.039	0.005	0.098	0.023	0.392	0.005	0.279	0.098	0.094
Canada	0.267	0.040	0.181	0.043	0.497	0.518	0.351	0.147	0.527	0.557	0.308	0.015	0.370	0.092	0.503	0.487	0.257	0.138	0.373	0.159
Chile	0.149	0.085	0.136	0.071	0.426	0.250	0.562	0.241	0.702	0.684	0.211	0.041	0.217	0.004	0.158	0.083	0.176	0.097	0.490	0.912
Denmark	0.346	0.046	0.031	0.306	0.037	0.586	0.064	0.071	0.290	0.250	0.424	0.004	0.009	0.052	0.004	0.175	0.048	0.068	0.343	0.145
Ecuador	0.146	0.019	0.112	0.042	0.103	0.068	0.313	0.113	0.114	0.288	0.027	0.001	0.026	0.007	0.018	0.011	0.066	0.023	0.878	0.130
Kazakhstan	0.075	0.001	0.052	0.023	0.449	0.126	0.150	0.449	0.772	0.913	0.186	0.044	0.316	0.098	0.647	0.181	0.913	0.368	0.503	0.543
New Zealand	0.008	0.068	0.030	0.201	0.007	0.037	0.042	0.192	0.129	0.198	0.166	0.101	0.005	0.181	0.007	0.177	0.039	0.268	0.050	0.193
Norway	0.016	0.235	0.054	0.757	0.057	0.261	0.198	0.462	0.099	0.395	0.341	0.425	0.144	0.479	0.037	0.604	0.025	0.257	0.040	0.964
Peru	0.292	0.475	0.374	0.504	0.922	0.301	0.982	0.703	0.604	0.288	0.493	0.504	0.465	0.452	0.832	0.251	0.981	0.361	0.313	0.480
South Africa	0.957	0.158	0.110	0.001	0.673	0.053	0.249	0.216	0.945	0.622	0.926	0.022	0.010	0.494	0.629	0.013	0.638	0.081	0.886	0.683
Venezuela	0.097	0.049	0.032	0.013	0.530	0.040	0.796	0.044	0.558	0.461	0.015	0.001	0.014	0.004	0.197	0.009	0.604	0.007	0.287	0.189
Importers																				
Czech Republic	0.950	0.400	0.514	0.498	0.647	0.002	0.865	0.004	0.088	0.011	0.993	0.010	0.490	0.510	0.957	0.001	0.912	0.001	0.939	0.002
Dominican Republic	0.072	0.013	0.056	0.305	0.148	0.570	0.744	0.257	0.422	0.912	0.025	0.009	0.471	0.414	0.591	0.379	0.863	0.610	0.913	0.840
Hungary	0.502	0.126	0.057	0.134	0.193	0.025	0.491	0.162	0.176	0.180	0.392	0.323	0.081	0.084	0.023	0.007	0.351	0.039	0.448	0.028
Luxembourg	0.132	0.163	0.519	0.298	0.076	0.416	0.043	0.458	0.212	0.095	0.045	0.241	0.172	0.074	0.075	0.311	0.082	0.164	0.201	0.008
Malta	0.006	0.122	0.169	0.068	0.135	0.195	0.230	0.061	0.500	0.357	0.085	0.219	0.243	0.216	0.197	0.136	0.075	0.077	0.119	0.145
Philippines	0.132	0.004	0.104	0.335	0.664	0.200	0.016	0.481	0.009	0.692	0.448	0.279	0.298	0.219	0.185	0.138	0.278	0.395	0.480	0.530
Slovakia	0.697	0.043	0.116	0.081	0.440	0.296	0.656	0.330	0.735	0.973	0.439	0.153	0.474	0.009	0.613	0.128	0.631	0.174	0.806	0.854
Slovenia	0.775	0.001	0.505	0.043	0.324	0.137	0.089	0.144	0.525	0.554	0.749	0.004	0.626	0.050	0.479	0.052	0.583	0.163	0.153	0.125
Both (Hybrid)																				
Bahrain	0.713	0.377	0.367	0.913	0.423	0.032	0.870	0.022	0.364	0.014	0.982	0.625	0.524	0.348	0.818	0.272	0.839	0.028	0.837	0.129
Belgium	0.087	0.318	0.494	0.510	0.611	0.244	0.181	0.326	0.402	0.105	0.162	0.059	0.471	0.440	0.050	0.089	0.276	0.105	0.365	0.061
Hong Kong	0.438	0.029	0.505	0.021	0.315	0.119	0.632	0.349	0.961	0.423	0.437	0.013	0.264	0.003	0.467	0.070	0.659	0.196	0.731	0.072
Estonia	0.012	0.001	0.067	0.001	0.041	0.435	0.066	0.179	0.257	0.514	0.221	0.041	0.133	0.046	0.127	0.219	0.108	0.167	0.023	0.206
Iceland	0.878	0.638	0.281	0.491	0.288	0.034	0.281	0.718	0.718	0.568	0.360	0.863	0.148	0.763	0.160	0.862	0.092	0.636	0.627	0.469
Israel	0.336	0.055	0.262	0.054	0.321	0.027	0.286	0.111	0.478	0.477	0.567	0.022	0.067	0.031	0.025	0.070	0.170	0.200	0.258	0.449
Latvia	0.980	0.010	0.935	0.059	0.227	0.049	0.384	0.196	0.037	0.375	0.711	0.002	0.432	0.005	0.211	0.004	0.076	0.037	0.012	0.040
Netherlands	0.802	0.015	0.094	0.067	0.037	0.110	0.581	0.149	0.558	0.108	0.202	0.014	0.234	0.002	0.302	0.021	0.168	0.024	0.460	0.035
Serbia	0.295	0.085	0.480	0.423	0.813	0.093	0.752	0.655	0.069	0.700	0.252	0.184	0.379	0.040	0.595	0.185	0.521	0.368	0.238	0.601
Seychelles	0.090	0.002	0.321	0.524	0.495	0.506	0.193	0.489	0.105	0.789	0.591	0.141	0.448	0.387	0.368	0.406	0.031	0.762	0.152	0.866
Singapore	0.041	0.572	0.047	0.325	0.009	0.528	0.101	0.143	0.528	0.434	0.035	0.073	0.018	0.027	0.003	0.232	0.041	0.024	0.299	0.172
Thailand	0.346	0.139	0.329	0.124	0.552	0.496	0.043	0.457	0.201	0.463	0.431	0.016	0.416	0.065	0.548	0.379	0.497	0.546	0.454	0.130
Viet Nam	0.045	0.984	0.030	0.426	0.601	0.641	0.922	0.970	0.046	0.484	0.104	0.920	0.083	0.591	0.585	0.663	0.448	0.789	0.659	0.172

Note: The table reports the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizons $h\epsilon\{1,2,3,4,6\}$*. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables.* $H_0: CP \nRightarrow EG$ *(*$\nRightarrow$ *means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with* $N = 999$ *replications. All variables are mean-centred and annual log-differenced.*

Table B.47 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices, Moody's

	Panel A: Mixed-frequency model										Panel B: Low-frequency model									
	horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6		horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Exporters																				
Australia	0.262	0.772	0.262	0.504	0.570	0.941	0.783	0.975	0.331	0.203	0.702	0.798	0.502	0.534	0.157	0.729	0.336	0.706	0.389	0.597
Bolivia	0.342	0.334	0.062	0.221	0.038	0.632	0.013	0.756	0.095	0.798	0.023	0.040	0.002	0.196	0.002	0.534	0.008	0.706	0.016	0.541
Canada	0.257	0.039	0.051	0.495	0.244	0.442	0.044	0.243	0.039	0.473	0.194	0.039	0.002	0.059	0.030	0.406	0.019	0.083	0.197	0.088
Chile	0.152	0.357	0.079	0.486	0.629	0.804	0.578	0.714	0.251	0.549	0.573	0.279	0.445	0.070	0.492	0.341	0.410	0.644	0.333	0.651
Denmark	0.266	0.219	0.093	0.624	0.435	0.559	0.349	0.104	0.535	0.034	0.762	0.066	0.296	0.206	0.149	0.113	0.368	0.087	0.850	0.014
Ecuador	0.227	0.099	0.126	0.087	0.297	0.456	0.316	0.342	0.088	0.942	0.338	0.064	0.190	0.071	0.035	0.134	0.260	0.124	0.289	0.770
Kazakhstan	0.655	0.027	0.452	0.028	0.732	0.212	0.575	0.980	0.696	0.718	0.668	0.004	0.675	0.064	0.638	0.162	0.962	0.622	0.450	0.564
New Zealand	0.153	0.214	0.090	0.203	0.076	0.270	0.700	0.240	0.434	0.116	0.253	0.183	0.271	0.115	0.356	0.096	0.223	0.103	0.120	0.116
Norway	0.502	0.777	0.019	0.433	0.055	0.193	0.291	0.648	0.034	0.173	0.506	0.394	0.453	0.458	0.031	0.412	0.032	0.454	0.031	0.958
Peru	0.524	0.152	0.500	0.493	0.002	0.576	0.632	0.744	0.446	0.849	0.581	0.313	0.500	0.504	0.103	0.114	0.454	0.635	0.146	0.705
South Africa	0.786	0.005	0.021	0.493	0.358	0.013	0.200	0.067	0.554	0.549	0.253	0.002	0.002	0.013	0.011	0.001	0.365	0.019	0.927	0.413
Venezuela	0.101	0.003	0.006	0.009	0.419	0.008	0.560	0.060	0.984	0.816	0.001	0.005	0.008	0.042	0.131	0.001	0.531	0.007	0.757	0.315
Importers																				
Czech Republic	0.850	0.372	0.507	0.466	0.367	0.013	0.937	0.019	0.706	0.024	0.480	0.011	0.482	0.501	0.540	0.005	0.193	0.007	0.983	0.019
Dominican Republic	0.766	0.003	0.380	0.193	0.415	0.441	0.521	0.183	0.055	0.789	0.188	0.004	0.071	0.095	0.303	0.140	0.175	0.262	0.488	0.412
Hungary	0.864	0.318	0.804	0.112	0.194	0.150	0.712	0.304	0.213	0.109	0.406	0.484	0.175	0.191	0.161	0.030	0.599	0.125	0.643	0.036
Luxembourg	0.338	0.993	0.776	0.749	0.661	0.963	0.253	0.893	0.642	0.994	0.401	0.589	0.488	0.286	0.222	0.709	0.289	0.807	0.166	0.904
Malta	0.151	0.027	0.022	0.035	0.084	0.069	0.062	0.057	0.144	0.479	0.287	0.047	0.231	0.019	0.110	0.077	0.061	0.065	0.123	0.115
Philippines	0.168	0.005	0.044	0.005	0.126	0.050	0.040	0.229	0.072	0.644	0.094	0.420	0.060	0.127	0.077	0.022	0.054	0.154	0.124	0.571
Slovakia	0.001	0.161	0.003	0.244	0.326	0.926	0.650	0.899	0.077	0.990	0.741	0.106	0.622	0.051	0.530	0.397	0.951	0.502	0.911	0.905
Slovenia	0.217	0.015	0.387	0.485	0.531	0.503	0.830	0.766	0.972	0.500	0.191	0.005	0.289	0.208	0.533	0.167	0.345	0.208	0.782	0.339
Both (Hybrid)																				
Bahrain	0.023	0.481	0.022	0.083	0.253	0.021	0.009	0.013	0.729	0.082	0.005	0.245	0.007	0.227	0.288	0.006	0.313	0.061	0.419	0.061
Belgium	0.584	0.837	0.500	0.484	0.997	0.462	0.791	0.367	0.323	0.538	0.405	0.177	0.496	0.462	0.080	0.295	0.433	0.367	0.557	0.629
Hong Kong	0.839	0.001	0.890	0.003	0.373	0.015	0.337	0.026	0.728	0.109	0.644	0.025	0.481	0.213	0.273	0.001	0.369	0.020	0.523	0.047
Estonia	0.421	0.005	0.002	0.006	0.164	0.382	0.096	0.202	0.033	0.569	0.900	0.002	0.837	0.010	0.002	0.266	0.076	0.209	0.007	0.358
Iceland	0.881	0.301	0.895	0.314	0.782	0.099	0.832	0.618	0.377	0.674	0.585	0.970	0.605	0.794	0.478	0.918	0.605	0.818	0.676	0.974
Israel	0.448	0.030	0.280	0.036	0.086	0.167	0.031	0.309	0.485	0.063	0.179	0.346	0.142	0.015	0.001	0.266	0.100	0.210	0.266	0.217
Latvia	0.215	0.004	0.158	0.013	0.139	0.011	0.452	0.026	0.116	0.362	0.541	0.002	0.247	0.006	0.130	0.007	0.042	0.006	0.010	0.093
Netherlands	0.783	0.025	0.297	0.029	0.068	0.024	0.740	0.052	0.605	0.069	0.182	0.005	0.127	0.036	0.148	0.021	0.135	0.019	0.696	0.074
Serbia	0.002	0.028	0.025	0.196	0.004	0.333	0.003	0.779	0.015	0.886	0.092	0.375	0.097	0.201	0.279	0.209	0.203	0.462	0.007	0.847
Seychelles	0.902	0.048	0.054	0.001	0.365	0.481	0.689	0.153	0.033	0.903	0.779	0.442	0.001	0.167	0.193	0.461	0.001	0.189	0.014	0.968
Singapore	0.001	0.013	0.002	0.006	0.007	0.331	0.066	0.210	0.795	0.147	0.001	0.028	0.003	0.038	0.001	0.245	0.017	0.072	0.263	0.095
Thailand	0.037	0.045	0.120	0.307	0.158	0.076	0.477	0.229	0.261	0.147	0.084	0.002	0.049	0.009	0.103	0.073	0.155	0.130	0.329	0.037
Viet Nam	0.002	0.921	0.715	0.956	0.657	0.454	0.542	0.361	0.907	0.516	0.032	0.644	0.449	0.836	0.620	0.754	0.443	0.957	0.569	0.408

Note: The table reports the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizons $h\epsilon\{1,2,3,4,6\}$*. “CP” denotes the commodity prices, while “EG” denotes the economic growth variables.* $H_0: CP \nRightarrow EG$ *(⇏ means “does not Granger-cause”). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian’s (2004) bootstrap with* $N = 999$ *replications. All variables are mean-centred and annual log-differenced.*

Table B.48 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices, Thompson Reuters

	Panel A: Mixed-frequency model										Panel B: Low-frequency model									
	horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6		horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG
Exporters																				
Australia	0.408	0.725	0.474	0.483	0.473	0.301	0.633	0.551	0.609	0.740	0.424	0.882	0.531	0.485	0.022	0.095	0.169	0.418	0.236	0.464
Bolivia	0.352	0.099	0.126	0.109	0.069	0.628	0.012	0.343	0.287	0.281	0.089	0.214	0.012	0.081	0.036	0.499	0.014	0.340	0.128	0.360
Canada	0.050	0.074	0.045	0.495	0.583	0.484	0.168	0.112	0.016	0.452	0.056	0.096	0.002	0.494	0.108	0.492	0.220	0.333	0.104	0.212
Chile	0.008	0.038	0.363	0.028	0.344	0.060	0.125	0.247	0.486	0.593	0.559	0.027	0.290	0.031	0.439	0.125	0.523	0.144	0.727	0.973
Denmark	0.504	0.007	0.142	0.027	0.034	0.330	0.133	0.193	0.037	0.014	0.262	0.001	0.017	0.046	0.002	0.115	0.120	0.035	0.346	0.044
Ecuador	0.118	0.058	0.296	0.376	0.084	0.569	0.200	0.452	0.131	0.646	0.059	0.085	0.059	0.251	0.006	0.383	0.070	0.306	0.063	0.501
Kazakhstan	0.531	0.001	0.448	0.005	0.656	0.446	0.997	0.598	0.517	0.331	0.297	0.011	0.293	0.025	0.286	0.111	0.830	0.580	0.768	0.123
New Zealand	0.673	0.040	0.415	0.075	0.034	0.236	0.276	0.356	0.022	0.446	0.394	0.144	0.427	0.156	0.168	0.275	0.137	0.407	0.091	0.701
Norway	0.369	0.142	0.160	0.402	0.444	0.441	0.301	0.240	0.141	0.826	0.526	0.634	0.351	0.388	0.114	0.526	0.087	0.611	0.058	0.638
Peru	0.433	0.043	0.523	0.512	0.290	0.543	0.692	0.980	0.090	0.848	0.862	0.496	0.523	0.530	0.383	0.569	0.645	0.903	0.124	0.293
South Africa	0.396	0.021	0.003	0.475	0.954	0.125	0.533	0.111	0.876	0.460	0.141	0.001	0.002	0.041	0.646	0.001	0.665	0.038	0.921	0.403
Venezuela	0.207	0.015	0.076	0.003	0.802	0.039	0.966	0.050	0.214	0.715	0.175	0.012	0.018	0.001	0.551	0.001	0.929	0.003	0.181	0.181
Importers																				
Czech Republic	0.305	0.138	0.504	0.492	0.778	0.092	0.786	0.011	0.275	0.004	0.565	0.043	0.501	0.475	0.821	0.005	0.812	0.001	0.878	0.002
Dominican Republic	0.025	0.055	0.382	0.746	0.254	0.931	0.340	0.904	0.388	0.804	0.034	0.010	0.341	0.522	0.557	0.780	0.685	0.819	0.721	0.936
Hungary	0.045	0.014	0.161	0.048	0.133	0.062	0.096	0.056	0.101	0.141	0.237	0.084	0.299	0.104	0.350	0.056	0.190	0.036	0.179	0.050
Luxembourg	0.106	0.022	0.196	0.084	0.087	0.450	0.289	0.416	0.409	0.357	0.085	0.171	0.053	0.164	0.180	0.610	0.064	0.435	0.030	0.492
Malta	0.515	0.313	0.110	0.334	0.033	0.203	0.158	0.173	0.228	0.273	0.625	0.193	0.260	0.237	0.161	0.232	0.081	0.177	0.027	0.215
Philippines	0.042	0.356	0.003	0.017	0.018	0.202	0.014	0.296	0.191	0.554	0.056	0.496	0.033	0.014	0.047	0.020	0.030	0.343	0.139	0.485
Slovakia	0.223	0.084	0.675	0.012	0.294	0.081	0.708	0.486	0.504	0.635	0.243	0.079	0.216	0.002	0.367	0.257	0.535	0.313	0.682	0.754
Slovenia	0.021	0.001	0.027	0.007	0.282	0.070	0.132	0.100	0.542	0.118	0.514	0.004	0.504	0.033	0.271	0.029	0.262	0.148	0.108	0.124
Both (Hybrid)																				
Bahrain	0.111	0.622	0.159	0.022	0.339	0.060	0.109	0.021	0.960	0.134	0.165	0.362	0.323	0.407	0.652	0.009	0.563	0.030	0.603	0.048
Belgium	0.028	0.037	0.537	0.492	0.705	0.072	0.549	0.084	0.446	0.047	0.129	0.076	0.470	0.529	0.279	0.324	0.422	0.302	0.529	0.055
Hong Kong	0.682	0.021	0.568	0.005	0.820	0.013	0.587	0.043	0.569	0.016	0.346	0.001	0.252	0.008	0.452	0.022	0.367	0.054	0.978	0.002
Estonia	0.490	0.012	0.064	0.014	0.246	0.174	0.657	0.153	0.286	0.722	0.644	0.006	0.379	0.104	0.226	0.216	0.227	0.086	0.092	0.370
Iceland	0.778	0.156	0.601	0.107	0.884	0.354	0.882	0.686	0.858	0.393	0.076	0.636	0.494	0.418	0.458	0.808	0.485	0.547	0.839	0.149
Israel	0.463	0.040	0.074	0.365	0.615	0.541	0.200	0.172	0.804	0.489	0.115	0.058	0.069	0.012	0.028	0.052	0.313	0.117	0.634	0.573
Latvia	0.066	0.001	0.190	0.001	0.928	0.005	0.931	0.002	0.467	0.292	0.191	0.001	0.176	0.001	0.518	0.006	0.178	0.004	0.018	0.032
Netherlands	0.568	0.083	0.410	0.471	0.381	0.096	0.612	0.021	0.477	0.021	0.259	0.045	0.183	0.062	0.131	0.033	0.093	0.040	0.340	0.036
Serbia	0.095	0.009	0.177	0.035	0.319	0.159	0.408	0.401	0.176	0.970	0.060	0.001	0.037	0.018	0.080	0.056	0.368	0.155	0.909	0.863
Seychelles	0.657	0.093	0.255	0.040	0.407	0.516	0.167	0.132	0.312	0.881	0.619	0.302	0.446	0.377	0.411	0.434	0.013	0.618	0.068	0.930
Singapore	0.033	0.407	0.023	0.391	0.026	0.178	0.047	0.110	0.543	0.256	0.006	0.325	0.009	0.095	0.001	0.470	0.038	0.065	0.478	0.096
Thailand	0.010	0.112	0.048	0.084	0.033	0.546	0.208	0.463	0.025	0.706	0.518	0.019	0.344	0.009	0.281	0.051	0.375	0.253	0.702	0.201
Viet Nam	0.042	0.064	0.017	0.282	0.606	0.179	0.442	0.088	0.216	0.318	0.161	0.811	0.331	0.301	0.838	0.815	0.887	0.659	0.681	0.203

Note: The table reports the bootstrapped p-values for the full sample mixed- and low-frequency Granger causality tests at the horizons $h \epsilon \{1,2,3,4,6\}$. *"CP" denotes the commodity prices, while "EG" denotes the economic growth variables.* $H_0: CP \nRightarrow EG$ *(⇏ means "does not Granger-cause"). The Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with* $N = 999$ *replications. All variables are mean-centred and annual log-differenced.*

Table B.49 Results of Full Sample Granger Causality Tests for Economic Growth and World Commodity Prices, IMF

Panel A: Mixed-frequency model																				
	horizon = 1				horizon = 2				horizon = 3				horizon = 4				horizon = 6			
	EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																				
Australia	0.305	0.432	0	0.021	0.021	0.063	0.063	0.105	0.053	0.126	0.053	0.116	0.011	0.053	0.242	0.316	0.021	0.032	0.011	0.063
Bolivia	0.055	0.164	0	0	0.036	0.182	0.018	0.018	0.055	0.073	0	0	0.055	0.182	0	0	0.345	0.382	0	0
Canada	0.021	0.042	0.189	0.379	0.011	0.021	0.095	0.168	0.011	0.042	0.032	0.084	0.074	0.126	0.011	0.053	0.011	0.063	0.021	0.021
Chile	0.613	0.839	0	0.032	0.097	0.129	0	0.032	0.129	0.161	0	0	0	0.032	0	0	0	0.032	0	0
Denmark	0	0	0	0.02	0	0	0	0	0	0	0	0	0	0	0	0	0.02	0.039	0	0
Ecuador	0	0	1	1	0	0.067	0.867	0.867	0	0	0.933	1	0	0	0.6	0.933	0	0	0	0
Kazakhstan	0.135	0.351	0.595	0.811	0	0.081	0.054	0.135	0	0.027	0	0	0.216	0.297	0	0	0	0	0	0
New Zealand	0.288	0.561	0.288	0.439	0.439	0.5	0.167	0.379	0.455	0.697	0.242	0.288	0.136	0.288	0.136	0.152	0	0.015	0.061	0.136
Norway	0.011	0.032	0.147	0.263	0.137	0.179	0.021	0.116	0.053	0.084	0.053	0.074	0	0	0	0.084	0.084	0.095	0	0.042
Peru	0	0.032	0	0	0	0.011	0	0	0.042	0.053	0.011	0.105	0	0.011	0.095	0.137	0	0	0.021	0.053
South Africa	0	0	0.253	0.442	0.042	0.053	0.253	0.411	0.053	0.074	0.189	0.389	0.063	0.074	0.074	0.147	0.063	0.095	0.084	0.095
Venezuela	0.045	0.091	1	1	0	0.182	0.455	0.636	0.818	0.864	0	0.091	0.045	0.091	0	0	0	0	0	0
Importers																				
Czech Republic	0	0	0.032	0.032	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Dominican Republic	0	0	0.135	0.297	0	0	0	0.027	0.054	0.108	0	0.189	0.297	0.73	0.486	0.622	0.135	0.324	0.027	0.027
Hungary	0	0.029	0	0	0.057	0.143	0	0.029	0	0	0	0.029	0	0.029	0	0	0.086	0.086	0	0
Luxembourg	0.067	0.2	0	0.133	0	0	0	0	0	0	0	0	0	0	0	0	0.133	0.333	0	0
Malta	0	0	0	0	0.267	0.4	0	0	0	0	0	0	0.4	0.533	0	0	0.4	0.6	0	0
Philippines	0.044	0.066	0.033	0.088	0.011	0.121	0.011	0.033	0.099	0.132	0	0.033	0.088	0.121	0	0.033	0.088	0.143	0	0
Slovakia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.029	0	0
Slovenia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Both (Hybrid)																				
Bahrain	0.141	0.163	0.011	0.043	0.098	0.141	0.043	0.13	0	0.011	0	0.033	0	0	0.054	0.109	0	0.011	0	0
Belgium	0	0.171	0.343	0.571	0.086	0.114	0.143	0.2	0	0	0	0	0	0.029	0	0	0.057	0.343	0	0.029
Hong Kong	0.074	0.158	0.653	0.705	0.084	0.116	0.316	0.547	0	0.074	0.084	0.253	0	0	0.168	0.232	0	0.042	0.274	0.326
Estonia	0.029	0.057	0.4	0.571	0.114	0.343	0.029	0.029	0	0	0.029	0.057	0	0	0.057	0.057	0	0	0	0
Iceland	0.037	0.037	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Israel	0.029	0.057	0.857	0.914	0	0	0.429	0.714	0	0	0	0.029	0	0.229	0	0	0.314	0.429	0	0
Latvia	0	0.029	0.029	0.029	0	0	0.029	0.086	0	0	0	0.029	0	0	0	0.029	0	0	0	0
Netherlands	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Serbia	0.226	0.387	0.194	0.355	0	0	0.097	0.355	0	0	0	0	0.129	0.226	0	0	0	0	0	0
Seychelles	0	0.011	0.095	0.189	0.032	0.095	0.158	0.242	0.137	0.263	0.053	0.242	0.326	0.347	0.021	0.053	0.368	0.526	0.011	0.021
Singapore	0.232	0.326	0.6	0.779	0.168	0.274	0.253	0.495	0.063	0.116	0.095	0.105	0.011	0.053	0.021	0.105	0.032	0.179	0.168	0.284
Thailand	0.581	0.837	0.326	0.395	0.605	0.698	0.047	0.163	0.512	0.535	0	0.023	0.209	0.512	0	0	0.14	0.279	0	0
Viet Nam	0	0.364	0	0	0.364	0.545	0	0.182	0	0	0	0	0	0	0	0	0	0	0	0

Panel B: Low-frequency model										
	horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG

	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																				
Australia	0.042	0.105	0.021	0.084	0.021	0.074	0	0	0.032	0.053	0	0	0.032	0.053	0	0	0	0	0	0.021
Bolivia	0	0.109	0	0	0.036	0.109	0	0	0.073	0.236	0	0	0.073	0.255	0	0	0.527	0.836	0	0
Canada	0.011	0.053	0.074	0.137	0	0	0	0.032	0	0	0.053	0.063	0	0	0.063	0.084	0.116	0.2	0	0.011
Chile	0.097	0.097	0.161	0.226	0.097	0.097	0.032	0.129	0.032	0.032	0	0	0	0	0	0	0	0	0	0
Denmark	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Ecuador	0.267	0.667	0.933	0.933	0	0.067	1	1	0	0	1	1	0	0	0.933	1	0	0	0	0
Kazakhstan	0	0	0.838	0.973	0	0	0.216	0.297	0	0.027	0	0.162	0	0	0	0	0	0	0	0
New Zealand	0.485	0.606	0.167	0.273	0.47	0.712	0.212	0.273	0.576	0.818	0.212	0.53	0.561	0.848	0.045	0.121	0	0.015	0.091	0.106
Norway	0	0	0.221	0.326	0	0.011	0.105	0.158	0	0.032	0.074	0.158	0	0.063	0.084	0.116	0.011	0.095	0	0.042
Peru	0.084	0.221	0.042	0.105	0	0.021	0	0	0	0	0	0	0	0	0	0	0	0	0	0
South Africa	0	0.021	0.021	0.105	0.011	0.021	0	0.053	0.032	0.053	0.021	0.074	0.053	0.063	0.084	0.105	0.032	0.063	0.095	0.116
Venezuela	0	0	0.955	0.955	0	0	0.909	0.955	0	0	0.909	1	0	0	0.318	0.955	0	0	0	0
Importers																				
Czech Republic	0	0.032	0.226	0.355	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Dominican Republic	0	0	0	0.054	0	0	0.054	0.27	0	0	0	0.27	0	0	0	0	0.216	0.865	0	0
Hungary	0	0	0.143	0.429	0	0	0.029	0.057	0	0	0	0.057	0	0	0	0	0	0.086	0	0
Luxembourg	0.133	0.133	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.333	0.6	0	0
Malta	0	0	0	0	0.2	0.467	0	0	0	0.133	0	0	0	0	0	0.067	0.067	0.333	0	0
Philippines	0.143	0.143	0.033	0.165	0.011	0.121	0	0.011	0.011	0.121	0	0.011	0.044	0.143	0	0.011	0	0.066	0	0
Slovakia	0	0	0	0.057	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.057
Slovenia	0	0.029	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Both (Hybrid)																				
Bahrain	0.185	0.337	0.033	0.174	0.217	0.326	0.065	0.217	0.087	0.174	0.076	0.185	0	0.054	0.098	0.174	0	0.054	0.054	0.12
Belgium	0.029	0.143	0.029	0.057	0	0	0	0.057	0	0	0	0.057	0	0	0	0	0	0	0	0
Hong Kong	0.147	0.242	0	0	0.042	0.158	0.021	0.042	0	0.053	0.053	0.158	0.032	0.042	0.158	0.253	0	0.053	0.274	0.432
Estonia	0	0	0.314	0.457	0	0	0.057	0.057	0	0	0.057	0.057	0	0	0.086	0.086	0	0	0	0
Iceland	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.037	0	0
Israel	0	0	0	0	0	0	0	0	0	0	0	0.057	0	0	0	0.057	0.543	0.686	0	0
Latvia	0	0	0.057	0.143	0	0	0.057	0.114	0	0	0.057	0.086	0	0	0	0.086	0	0	0	0
Netherlands	0	0	0	0	0	0	0	0	0	0	0	0.065	0	0	0	0.097	0	0	0	0
Serbia	0.065	0.065	0.613	0.742	0	0.032	0.258	0.355	0	0	0	0.129	0	0	0	0	0	0	0	0
Seychelles	0.032	0.084	0.053	0.179	0	0.074	0.032	0.158	0.032	0.147	0	0	0.168	0.242	0	0	0.442	0.579	0	0.011
Singapore	0.432	0.653	0.042	0.126	0.221	0.284	0.095	0.116	0.126	0.211	0.105	0.116	0.042	0.053	0.053	0.116	0.042	0.211	0.347	0.411
Thailand	0.465	0.581	0.488	0.558	0.488	0.651	0.395	0.628	0.651	0.674	0.419	0.744	0.651	0.674	0	0.093	0	0	0.023	0.209
Viet Nam	0	0	0.273	0.273	0	0	0.091	0.182	0	0	0	0	0	0	0	0	0	0	0	0

Note: The table reports the rejection frequencies at different significance levels for rolling window mixed- and low-frequency Granger causality tests with a window size of 50 quarters for the horizons $h\epsilon\{1,2,3,4,6\}$. For each rolling window, the Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ replications. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0: CP \nRightarrow EG$ ($\nRightarrow$ means "does not Granger-cause"). All variables are mean-centred and annual log-differenced.

Table B.50 Rejection Frequencies at Different Significant Levels for Rolling Window Granger Causality Tests, Reuters/Jeffries

	Panel A: Mixed-frequency model																			
	horizon = 1				horizon = 2				horizon = 3				horizon = 4				horizon = 6			
	EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																				
Australia	0.011	0.042	0.095	0.137	0	0	0.295	0.474	0	0.021	0.2	0.379	0	0.011	0.347	0.421	0	0.063	0.021	0.053
Bolivia	0.091	0.182	0	0.018	0.018	0.036	0	0	0	0	0	0	0	0	0	0	0.073	0.182	0	0
Canada	0	0.095	0.105	0.2	0	0	0.084	0.105	0	0	0.042	0.063	0.032	0.189	0.074	0.084	0.021	0.095	0.063	0.116
Chile	0.097	0.387	0.032	0.161	0.161	0.161	0	0.032	0.194	0.194	0	0	0.032	0.194	0	0	0	0	0	0
Denmark	0	0	0	0.078	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Ecuador	0	0	1	1	0	0.133	0.8	1	0	0	0.467	0.667	0	0	0.4	0.933	0	0.067	0	0
Kazakhstan	0.568	0.595	0.405	0.649	0.27	0.486	0	0.162	0	0.027	0	0.081	0	0	0	0.054	0	0	0	0
New Zealand	0.212	0.242	0.318	0.636	0.303	0.439	0.152	0.364	0.061	0.167	0.197	0.394	0	0	0.197	0.227	0.015	0.045	0.045	0.061
Norway	0	0	0.084	0.084	0.032	0.105	0.053	0.053	0.021	0.084	0.2	0.263	0.053	0.168	0	0	0.126	0.158	0.011	0.105
Peru	0	0	0.179	0.274	0	0	0.032	0.042	0	0.011	0.063	0.116	0	0	0.063	0.095	0	0.011	0	0
South Africa	0	0	0.295	0.432	0	0	0.011	0.053	0	0	0.032	0.053	0	0.011	0.074	0.116	0	0	0.084	0.158
Venezuela	0	0	0.955	1	0	0	1	1	0	0	0.409	0.909	0	0	0	0.045	0	0	0	0
Importers																				
Czech Republic	0	0	0	0.065	0.032	0.097	0	0	0.161	0.161	0	0	0	0.097	0	0	0	0	0	0
Dominican Republic	0.027	0.027	0.135	0.189	0	0	0	0	0	0	0	0.054	0.081	0.216	0	0	0.216	0.243	0	0.027
Hungary	0	0	0	0	0	0.029	0	0	0	0	0	0	0	0	0	0	0.086	0.086	0	0
Luxembourg	0.2	0.267	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.067	0.2	0	0
Malta	0	0.067	0	0	0.333	0.733	0	0	0.267	0.267	0	0	0.533	0.533	0	0	0	0.067	0	0
Philippines	0.044	0.077	0.143	0.22	0.176	0.253	0.077	0.099	0.231	0.275	0.011	0.055	0.253	0.286	0.044	0.099	0.187	0.231	0.077	0.143
Slovakia	0	0.029	0	0	0.371	0.4	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Slovenia	0	0	0	0	0	0	0	0	0	0	0	0	0.029	0.029	0	0	0.057	0.057	0	0
Both (Hybrid)																				
Bahrain	0	0.043	0	0.011	0.011	0.011	0	0	0.022	0.087	0	0	0	0	0	0	0	0.054	0	0.011
Belgium	0	0	0.171	0.286	0	0	0	0	0	0	0	0	0.029	0.029	0	0	0	0.029	0	0
Hong Kong	0.168	0.274	0.547	0.579	0.063	0.137	0	0.084	0.158	0.211	0.021	0.084	0	0	0.042	0.116	0.158	0.316	0.021	0.095
Estonia	0	0	0.114	0.343	0.086	0.143	0	0	0.029	0.057	0	0	0	0	0.057	0.057	0	0	0	0
Iceland	0.037	0.037	0.259	0.481	0	0	0	0	0	0	0	0	0	0	0	0	0	0.333	0	0
Israel	0.2	0.314	0.971	1	0	0	0.143	0.543	0	0	0	0	0	0	0	0	0	0.029	0	0
Latvia	0	0	0	0	0.029	0.029	0	0	0.029	0.086	0.029	0.029	0	0.029	0.029	0.057	0	0	0.029	0.086
Netherlands	0	0	0	0.032	0	0	0	0	0	0	0	0	0	0	0	0	0	0.032	0	0
Serbia	0	0.032	0.387	0.581	0	0	0.032	0.323	0	0	0.032	0.161	0	0.032	0	0	0	0	0	0
Seychelles	0	0.011	0.032	0.042	0.011	0.053	0.053	0.232	0.021	0.042	0.074	0.189	0.232	0.263	0.063	0.084	0.042	0.253	0.021	0.032
Singapore	0.211	0.263	0.695	0.821	0.084	0.2	0.284	0.411	0.042	0.095	0.126	0.137	0	0	0.084	0.168	0	0	0.326	0.453
Thailand	0.605	0.814	0.209	0.326	0.419	0.558	0	0	0.047	0.372	0	0	0.07	0.186	0	0	0.07	0.116	0	0.047
Viet Nam	0	0	0	0	0	0.182	0	0.091	0	0	0	0	0	0	0	0	0	0	0	0

	Panel B: Low-frequency model									
	horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG

	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																				
Australia	0.032	0.084	0.084	0.421	0.042	0.063	0.158	0.337	0.021	0.032	0.074	0.221	0	0.021	0	0.042	0	0.032	0.042	0.147
Bolivia	0	0	0	0.018	0	0	0	0	0.036	0.127	0	0	0.036	0.127	0	0	0.236	0.582	0	0
Canada	0.011	0.032	0.011	0.021	0	0	0.032	0.063	0	0	0.074	0.084	0	0	0.084	0.095	0.032	0.179	0.074	0.305
Chile	0	0.065	0	0.129	0	0.032	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Denmark	0	0	0	0	0	0.02	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Ecuador	0	0	0.8	0.867	0	0	0.933	0.933	0	0	0.867	1	0	0	1	1	0	0	0	0
Kazakhstan	0	0	0.865	0.946	0	0.027	0.216	0.324	0	0	0.135	0.162	0	0	0	0	0.054	0.162	0	0
New Zealand	0.394	0.485	0.258	0.606	0.197	0.318	0.303	0.591	0.167	0.318	0.636	0.864	0.136	0.167	0.242	0.455	0	0	0.091	0.091
Norway	0.011	0.042	0.105	0.147	0.053	0.105	0.053	0.074	0.105	0.137	0.095	0.179	0.137	0.147	0	0.021	0.137	0.179	0.074	0.137
Peru	0	0.011	0	0.011	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
South Africa	0	0	0	0.021	0	0	0	0	0	0	0	0.011	0.011	0.095	0.021	0.095	0	0	0.063	0.2
Venezuela	0	0	0.955	1	0	0	1	1	0	0	0.955	1	0	0	0.682	1	0	0	0	0
Importers																				
Czech Republic	0	0.032	0.226	0.323	0	0	0	0	0	0.065	0	0	0	0.097	0	0	0	0	0	0
Dominican Republic	0	0	0	0	0	0	0	0	0	0	0	0.027	0	0	0	0	0.027	0.189	0	0
Hungary	0	0	0	0.057	0	0	0	0.029	0	0	0	0	0	0	0	0	0	0.057	0	0
Luxembourg	0.067	0.267	0	0	0	0.067	0	0	0	0	0	0	0	0	0	0	0.133	0.2	0	0
Malta	0.067	0.333	0	0	0.6	0.733	0	0	0.333	0.533	0	0	0.133	0.267	0.067	0.2	0.133	0.2	0	0
Philippines	0.11	0.231	0	0.11	0.143	0.198	0	0	0.22	0.264	0	0.033	0.275	0.297	0	0	0.165	0.176	0.022	0.11
Slovakia	0	0	0	0.057	0	0	0	0.029	0	0	0	0.029	0	0	0.057	0.057	0	0	0.057	0.086
Slovenia	0.057	0.057	0	0	0	0.029	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Both (Hybrid)																				
Bahrain	0	0	0	0	0	0	0	0	0	0.011	0	0	0	0.076	0	0.043	0.065	0.228	0.011	0.12
Belgium	0.057	0.143	0	0.029	0	0	0	0.057	0	0	0	0	0	0.057	0	0.057	0	0	0	0
Hong Kong	0.168	0.221	0.053	0.116	0.179	0.211	0.011	0.105	0.042	0.126	0.074	0.168	0	0	0.105	0.211	0.137	0.2	0.147	0.274
Estonia	0	0	0.057	0.114	0	0	0.057	0.057	0	0	0.057	0.057	0	0	0.029	0.086	0	0	0	0
Iceland	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Israel	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.086	0.343	0	0
Latvia	0	0	0	0.029	0	0	0.029	0.057	0	0	0.029	0.057	0	0	0	0.057	0	0	0	0
Netherlands	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.032	0	0
Serbia	0	0.032	0.581	0.839	0	0	0.032	0.194	0	0	0	0	0	0	0	0	0	0	0	0
Seychelles	0	0	0.011	0.084	0	0	0	0.021	0	0.011	0	0	0.095	0.168	0	0	0.137	0.326	0	0
Singapore	0.295	0.474	0.084	0.095	0.074	0.305	0.137	0.158	0	0.042	0.137	0.179	0	0.063	0.137	0.147	0.011	0.105	0.474	0.474
Thailand	0.419	0.512	0.047	0.279	0.465	0.605	0	0.395	0.419	0.628	0.209	0.512	0.209	0.302	0	0.07	0	0	0	0.047
Viet Nam	0	0	0.727	0.909	0	0	1	1	0	0	0	0.636	0	0	0	0	0	0	0	0

Note: The table reports the rejection frequencies at different significance levels for rolling window mixed- and low-frequency Granger causality tests with a window size of 50 quarters for the horizons $h\epsilon\{1,2,3,4,6\}$. For each rolling window, the Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ replications. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0\colon CP \nRightarrow EG$ ($\nRightarrow$ means "does not Granger-cause"). All variables are mean-centred and annual log-differenced.

Table B.51 Rejection Frequencies at Different Significant Levels for Rolling Window Granger Causality Tests, Goldman Sachs

	Panel A: Mixed-frequency model																			
	horizon = 1				**horizon = 2**				**horizon = 3**				**horizon = 4**				**horizon = 6**			
	EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																				
Australia	0	0.032	0	0.074	0	0	0	0	0	0	0.021	0.021	0.011	0.021	0.063	0.116	0	0.032	0.063	0.189
Bolivia	0	0.018	0.109	0.218	0.018	0.073	0.018	0.055	0.073	0.091	0	0	0.055	0.109	0	0	0	0	0	0
Canada	0.011	0.011	0.537	0.632	0.011	0.053	0.432	0.516	0	0.063	0.368	0.453	0	0.053	0.179	0.295	0.011	0.021	0.105	0.2
Chile	0.097	0.097	0.194	0.29	0	0.129	0.323	0.355	0	0	0.387	0.387	0	0	0.129	0.387	0	0	0	0
Denmark	0	0.059	0.098	0.196	0	0	0.118	0.235	0	0	0.039	0.176	0	0.098	0.02	0.059	0	0.039	0	0
Ecuador	0.133	0.4	0.733	0.733	0.6	0.733	0.6	0.733	0.533	0.8	0.067	0.4	0	0.067	0	0	0.067	0.2	0	0
Kazakhstan	0	0	0.351	0.757	0	0	0.162	0.378	0.027	0.135	0.081	0.378	0	0.108	0	0	0	0	0	0
New Zealand	0	0.015	0.167	0.242	0	0.015	0.273	0.455	0.136	0.182	0.182	0.348	0.152	0.227	0	0.03	0	0	0	0.091
Norway	0	0	0.032	0.063	0.021	0.053	0	0.032	0.011	0.032	0.042	0.063	0.063	0.084	0	0	0.084	0.147	0	0
Peru	0.074	0.232	0.116	0.211	0	0.032	0.126	0.358	0	0.032	0.063	0.179	0.105	0.126	0	0	0.021	0.095	0	0
South Africa	0	0	0.358	0.568	0	0	0.337	0.484	0.011	0.074	0.105	0.232	0.074	0.126	0.032	0.053	0	0	0	0.011
Venezuela	0	0	0.591	0.818	0	0	0.591	0.864	0	0	0.273	0.636	0	0	0.045	0.091	0	0	0	0.045
Importers																				
Czech Republic	0.065	0.065	0.032	0.355	0	0.065	0	0.194	0	0.065	0	0.065	0	0.065	0.032	0.258	0	0	0	0.258
Dominican Republic	0	0	0	0	0	0	0.081	0.108	0.054	0.135	0	0	0	0	0	0	0	0	0	0
Hungary	0.029	0.057	0	0.057	0	0.029	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Luxembourg	0	0	0.2	0.267	0	0	0.333	0.6	0	0	0.333	0.4	0	0	0.2	0.4	0	0	0	0.067
Malta	0.6	0.733	0	0	0	0.2	0.267	0.467	0	0	0.467	0.6	0.133	0.4	0	0	0	0	0	0.067
Philippines	0.077	0.121	0.275	0.374	0.044	0.176	0.033	0.11	0.022	0.088	0	0.033	0.066	0.154	0	0.022	0.011	0.022	0	0
Slovakia	0	0	0.086	0.2	0	0	0.457	0.743	0	0	0	0.2	0	0	0.029	0.286	0	0	0.029	0.114
Slovenia	0	0	0.429	0.686	0	0.029	0	0.143	0	0	0	0	0	0.029	0	0.029	0	0	0	0
Both (Hybrid)																				
Bahrain	0	0.011	0	0.054	0	0.033	0.033	0.065	0.174	0.272	0	0.011	0.033	0.141	0	0.022	0.054	0.25	0	0
Belgium	0.114	0.114	0.714	0.771	0.086	0.114	0.8	0.886	0	0.029	0.114	0.257	0	0	0	0	0	0	0	0
Hong Kong	0.042	0.126	0.653	0.747	0.011	0.042	0.368	0.642	0.021	0.126	0.042	0.158	0.011	0.063	0.053	0.158	0.074	0.084	0.4	0.421
Estonia	0	0	0.286	0.429	0	0.029	0.343	0.543	0	0	0	0.029	0.286	0.714	0	0	0	0	0	0
Iceland	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Israel	0	0	1	1	0	0.057	0.914	0.914	0	0	0.657	0.8	0	0.029	0.314	0.4	0	0	0	0
Latvia	0	0	0.314	0.514	0	0	0.257	0.6	0	0	0.057	0.4	0	0.029	0	0	0	0	0	0
Netherlands	0.032	0.097	0.774	0.871	0	0	0.161	0.516	0	0.065	0	0.065	0	0.032	0	0	0	0	0	0
Serbia	0	0	0.194	0.194	0	0	0.129	0.226	0	0	0.161	0.258	0	0	0.129	0.194	0	0.032	0	0
Seychelles	0	0.011	0.179	0.221	0	0.032	0	0.084	0	0.032	0	0.021	0.011	0.105	0	0	0.074	0.116	0	0
Singapore	0.042	0.084	0.368	0.621	0.063	0.137	0.074	0.137	0.021	0.042	0	0	0	0.011	0	0	0.053	0.053	0	0
Thailand	0.488	0.628	0.209	0.558	0.372	0.512	0.023	0.233	0.349	0.349	0	0.047	0.372	0.372	0	0	0.163	0.326	0	0
Viet Nam	0.727	0.727	0	0	0.909	1	0	0	0	0	0	0	0	0	0	0	0	0	0.182	0.545

	Panel B: Low-frequency model									
	horizon = 1		**horizon = 2**		**horizon = 3**		**horizon = 4**		**horizon = 6**	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG

	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																				
Australia	0	0	0	0.032	0	0	0.032	0.042	0	0	0.011	0.042	0	0.011	0	0	0	0	0	0
Bolivia	0.018	0.073	0.418	0.509	0.127	0.2	0	0.073	0.091	0.145	0	0	0.127	0.2	0.018	0.055	0	0.018	0	0.2
Canada	0.053	0.179	0	0.021	0.011	0.137	0.095	0.116	0	0.063	0.221	0.358	0.021	0.084	0.274	0.4	0.053	0.105	0.2	0.305
Chile	0.258	0.419	0.581	0.935	0.129	0.258	0.387	0.806	0	0.097	0.226	0.258	0	0	0	0	0	0	0	0
Denmark	0	0	0	0	0	0	0.196	0.255	0	0.059	0.196	0.294	0.02	0.216	0.216	0.314	0	0	0	0.059
Ecuador	0.8	0.933	0.8	0.933	0.8	1	0.533	0.6	1	1	0.533	0.933	0.6	0.867	0.133	0.933	0	0	0	0
Kazakhstan	0	0	0.649	0.838	0	0	0	0.243	0	0	0	0	0	0	0	0	0	0	0.081	0.243
New Zealand	0.212	0.333	0.015	0.045	0.152	0.379	0.076	0.258	0.242	0.379	0.348	0.455	0.182	0.242	0.106	0.212	0	0	0.03	0.03
Norway	0	0.032	0.074	0.168	0.042	0.053	0.042	0.053	0.053	0.053	0.032	0.084	0.063	0.116	0.053	0.074	0.011	0.074	0	0
Peru	0.095	0.126	0.063	0.158	0.032	0.053	0.021	0.137	0	0	0.032	0.147	0	0	0	0.042	0	0	0	0.011
South Africa	0	0	0.126	0.263	0	0.032	0.116	0.284	0.053	0.126	0	0.095	0.126	0.147	0.042	0.105	0.011	0.032	0.011	0.032
Venezuela	0	0	0.273	0.545	0	0	1	1	0	0	0.955	1	0	0	1	1	0	0	0	0.136
Importers																				
Czech Republic	0	0.032	0.032	0.129	0.032	0.065	0	0	0.065	0.065	0	0	0.065	0.097	0	0.065	0	0.097	0.194	0.226
Dominican Republic	0	0.135	0	0.027	0	0	0	0.081	0	0	0	0	0	0	0	0	0	0	0	0
Hungary	0.029	0.114	0	0	0	0.057	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Luxembourg	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.533	0.133	0.467
Malta	0	0.067	0	0	0	0	0	0	0	0	0	0	0	0	0.067	0.2	0	0	0	0.6
Philippines	0.132	0.143	0.154	0.308	0.011	0.121	0.011	0.088	0.022	0.11	0	0.033	0.088	0.11	0	0	0.044	0.132	0	0
Slovakia	0	0	0.114	0.6	0	0	0.171	0.314	0	0	0.143	0.2	0	0	0.286	0.314	0	0	0.114	0.2
Slovenia	0	0	0	0.143	0	0	0	0.029	0	0.029	0	0.029	0	0	0.057	0.057	0	0	0	0
Both (Hybrid)																				
Bahrain	0.054	0.12	0.076	0.174	0.087	0.141	0.087	0.196	0	0.174	0.033	0.087	0.054	0.326	0.022	0.076	0.098	0.261	0.043	0.054
Belgium	0	0	0	0	0.029	0.057	0	0	0	0.057	0	0	0	0	0	0.086	0	0	0	0.143
Hong Kong	0.147	0.274	0.011	0.053	0.042	0.105	0.032	0.168	0	0.042	0.042	0.232	0.042	0.063	0.168	0.379	0	0.021	0.453	0.463
Estonia	0	0	0.6	0.829	0	0	0	0	0	0	0	0	0	0	0	0	0	0.171	0	0
Iceland	0	0.148	0	0	0	0.074	0	0	0	0.259	0	0	0.037	0.481	0	0	0	0	0	0
Israel	0	0	0.429	0.657	0	0	0.143	0.2	0	0.029	0.029	0.143	0	0.086	0	0	0	0	0	0
Latvia	0	0	0.371	0.629	0	0	0.029	0.314	0	0	0	0	0	0	0	0	0	0	0	0
Netherlands	0	0.129	0.065	0.29	0	0.032	0	0	0.097	0.194	0	0	0.161	0.194	0	0.065	0.065	0.194	0	0
Serbia	0.032	0.161	0.419	0.645	0.032	0.29	0.194	0.484	0	0	0.194	0.194	0	0	0.194	0.226	0	0	0	0
Seychelles	0	0.032	0	0.011	0	0.011	0	0.032	0	0.032	0.021	0.063	0.042	0.137	0.021	0.063	0.032	0.084	0	0.053
Singapore	0.179	0.347	0	0	0	0.053	0.011	0.021	0	0	0.011	0.074	0	0	0	0	0	0.116	0.053	0.158
Thailand	0.372	0.372	0.07	0.163	0.372	0.372	0.186	0.233	0.372	0.372	0.163	0.326	0.372	0.372	0	0.07	0	0.023	0	0
Viet Nam	0.091	0.636	0	0	0.182	0.727	0	0	0	0	0	0	0	0	0	0	0	0	0	0.455

Note: The table reports the rejection frequencies at different significance levels for rolling window mixed- and low-frequency Granger causality tests with a window size of 50 quarters for the horizons $h\epsilon\{1,2,3,4,6\}$. For each rolling window, the Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ replications. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0: CP \nRightarrow EG$ ($\nRightarrow$ means "does not Granger-cause"). All variables are mean-centred and annual log-differenced.

Table B.52 Rejection Frequencies at Different Significant Levels for Rolling Window Granger Causality Tests, Moody's

Panel A: Mixed-frequency model																				
	horizon = 1				horizon = 2				horizon = 3				horizon = 4				horizon = 6			
	EG $\nRightarrow$ CP		CP $\nRightarrow$ EG		EG $\nRightarrow$ CP		CP $\nRightarrow$ EG		EG $\nRightarrow$ CP		CP $\nRightarrow$ EG		EG $\nRightarrow$ CP		CP $\nRightarrow$ EG		EG $\nRightarrow$ CP		CP $\nRightarrow$ EG	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																				
Australia	0	0.053	0	0.021	0	0	0.105	0.284	0.074	0.147	0.126	0.179	0	0.011	0.253	0.263	0	0	0.011	0.074
Bolivia	0	0.055	0	0	0.036	0.091	0	0	0.055	0.055	0	0	0	0.055	0.018	0.109	0.091	0.218	0.036	0.145
Canada	0.074	0.168	0.6	0.642	0.053	0.053	0.295	0.463	0	0.042	0.095	0.158	0	0	0.074	0.095	0.095	0.126	0.074	0.095
Chile	0.097	0.258	0.032	0.161	0	0.129	0.032	0.226	0.161	0.194	0	0	0	0	0	0	0	0	0	0
Denmark	0.02	0.118	0.02	0.059	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Ecuador	0	0	0.733	0.867	0.067	0.4	0.467	0.467	0	0.067	0	0.333	0	0	0	0	0	0	0	0
Kazakhstan	0	0	0.892	0.946	0	0	0.514	0.919	0.027	0.027	0.135	0.351	0.027	0.216	0	0	0	0.108	0	0
New Zealand	0.152	0.227	0.212	0.348	0.061	0.167	0.106	0.288	0	0.045	0.136	0.379	0	0	0.136	0.167	0.045	0.091	0.182	0.379
Norway	0	0	0	0.021	0.011	0.021	0.042	0.074	0.053	0.074	0	0.011	0	0.011	0	0	0.105	0.137	0	0.032
Peru	0.053	0.063	0.011	0.011	0.053	0.074	0	0	0	0.021	0	0	0	0.021	0	0	0.011	0.074	0	0
South Africa	0	0	0.337	0.453	0.032	0.084	0.147	0.4	0.042	0.179	0.137	0.337	0.095	0.126	0	0.095	0	0	0	0.011
Venezuela	0.136	0.227	1	1	0	0.045	0.864	0.909	0	0	0.955	0.955	0	0	0.091	0.318	0	0	0	0
Importers																				
Czech Republic	0	0.032	0.097	0.258	0.065	0.065	0	0.097	0.097	0.161	0	0	0.129	0.161	0	0	0	0.032	0	0
Dominican Republic	0	0	0.027	0.108	0	0	0	0	0	0	0	0	0	0	0	0.054	0	0	0	0
Hungary	0	0	0	0	0	0.057	0	0	0	0	0	0	0	0	0	0	0.029	0.029	0	0.143
Luxembourg	0	0.067	0	0.067	0	0	0	0.133	0	0	0	0	0	0	0	0	0	0	0	0
Malta	0.067	0.333	0	0	0.133	0.333	0	0.133	0	0.333	0	0	0.333	0.667	0	0	0.2	0.267	0	0
Philippines	0.066	0.264	0.143	0.308	0.187	0.242	0	0.011	0.198	0.264	0	0	0.143	0.209	0	0.011	0.264	0.516	0	0.022
Slovakia	0	0	0	0	0.029	0.029	0.029	0.086	0	0	0	0	0	0	0	0	0	0	0	0.086
Slovenia	0.029	0.029	0.2	0.457	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Both (Hybrid)																				
Bahrain	0.12	0.163	0	0.033	0.065	0.109	0.152	0.228	0.054	0.087	0.054	0.185	0.011	0.109	0.228	0.304	0.098	0.141	0	0
Belgium	0	0	0.6	0.8	0	0	0.029	0.2	0	0	0	0	0	0	0	0	0	0.057	0	0.114
Hong Kong	0.021	0.053	0.632	0.768	0	0	0.274	0.537	0.021	0.063	0.084	0.147	0	0.021	0.126	0.221	0	0	0.242	0.305
Estonia	0	0	0.714	0.857	0.086	0.086	0.229	0.371	0	0	0	0	0	0	0	0	0	0.029	0	0
Iceland	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Israel	0	0	0.486	0.743	0	0.029	0.543	0.771	0	0	0.143	0.229	0	0	0	0	0	0.114	0	0.114
Latvia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Netherlands	0	0	0	0.29	0	0	0	0.032	0	0	0	0	0	0.032	0	0	0	0.032	0	0
Serbia	0.065	0.097	0.161	0.226	0	0	0.065	0.226	0	0	0.032	0.065	0.226	0.258	0	0	0	0	0	0
Seychelles	0	0.011	0	0.021	0	0.063	0	0.021	0	0.084	0	0.032	0	0.116	0	0	0.021	0.158	0	0
Singapore	0.232	0.337	0.684	0.779	0.053	0.084	0.242	0.368	0.011	0.042	0.032	0.095	0	0.042	0.011	0.126	0.042	0.053	0.116	0.326
Thailand	0.558	0.698	0.07	0.233	0.488	0.558	0	0	0.419	0.442	0	0.047	0.419	0.465	0	0	0	0	0	0.023
Viet Nam	0.909	1	0	0	0	0.364	0	0	0	0	0	0	0	0	0	0.182	0	0	0	0

Panel B: Low-frequency model										
	horizon = 1		horizon = 2		horizon = 3		horizon = 4		horizon = 6	
	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG	EG $\nRightarrow$ CP	CP $\nRightarrow$ EG

	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																				
Australia	0.042	0.084	0.063	0.232	0	0	0.021	0.095	0	0	0	0.021	0	0	0	0	0	0	0	0
Bolivia	0	0	0	0	0.018	0.055	0	0	0.073	0.164	0	0	0.091	0.182	0	0	0.418	0.491	0	0
Canada	0	0.042	0.021	0.042	0	0.011	0.084	0.126	0	0	0.116	0.147	0	0	0.126	0.147	0.221	0.221	0.063	0.095
Chile	0.065	0.065	0.129	0.194	0	0	0	0.065	0	0	0	0	0	0	0	0	0	0	0	0
Denmark	0	0.02	0	0	0	0.02	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Ecuador	0.467	0.933	1	1	0.4	0.733	0.667	0.867	0.067	0.533	0.4	0.933	0	0	0	0.2	0	0	0	0
Kazakhstan	0	0.027	1	1	0	0.108	0.568	0.919	0	0	0	0.135	0	0	0	0	0	0	0	0
New Zealand	0.273	0.303	0.136	0.212	0.227	0.258	0.242	0.333	0.136	0.197	0.439	0.606	0.045	0.076	0.258	0.394	0	0	0.045	0.091
Norway	0	0	0.032	0.084	0	0.021	0.011	0.032	0.011	0.095	0.021	0.074	0	0.105	0	0	0.021	0.116	0.053	0.147
Peru	0.095	0.284	0.095	0.179	0.053	0.074	0	0.042	0.011	0.032	0	0	0	0	0	0	0	0	0	0
South Africa	0	0.053	0.042	0.116	0.042	0.158	0	0.042	0.168	0.189	0.021	0.042	0.2	0.211	0.053	0.095	0	0	0	0.042
Venezuela	0	0	0.682	0.909	0	0	1	1	0	0	1	1	0	0	1	1	0	0	0	0
Importers																				
Czech Republic	0.032	0.065	0	0.129	0.065	0.065	0	0	0.097	0.097	0	0	0.129	0.129	0	0	0	0.032	0.032	0.161
Dominican Republic	0.216	0.459	0.054	0.135	0.162	0.351	0.027	0.27	0	0.108	0	0.189	0	0	0	0.081	0.027	0.189	0	0
Hungary	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Luxembourg	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.4	0	0
Malta	0	0	0	0	0	0.133	0	0	0	0.267	0.067	0.133	0	0.267	0.2	0.467	0.067	0.267	0	0.267
Philippines	0.143	0.154	0.066	0.242	0.055	0.22	0	0	0.176	0.319	0	0	0.275	0.319	0	0	0.121	0.132	0	0
Slovakia	0	0	0.057	0.2	0	0	0.114	0.143	0	0	0	0.029	0	0	0.086	0.114	0	0	0.114	0.143
Slovenia	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Both (Hybrid)																				
Bahrain	0.196	0.348	0.087	0.217	0.163	0.228	0.098	0.25	0	0.054	0.196	0.239	0	0.109	0.217	0.239	0	0.174	0.087	0.185
Belgium	0	0.029	0	0.029	0	0	0	0.029	0	0	0	0	0	0	0	0	0	0	0	0
Hong Kong	0.053	0.137	0	0.011	0.063	0.063	0	0.053	0.011	0.032	0.074	0.158	0	0	0.179	0.305	0	0	0.432	0.463
Estonia	0	0	0.229	0.543	0	0	0.057	0.057	0	0	0	0	0	0	0	0	0	0	0	0.029
Iceland	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Israel	0	0	0	0	0	0	0	0	0	0	0	0.029	0	0	0	0.057	0	0.2	0	0
Latvia	0	0	0.029	0.114	0	0	0.029	0.029	0	0	0	0.057	0	0	0	0	0	0	0	0
Netherlands	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.129	0	0
Serbia	0	0	0.548	0.677	0	0	0.161	0.355	0	0	0	0.032	0	0	0	0	0	0	0	0
Seychelles	0.011	0.032	0	0	0	0.011	0	0	0	0	0	0	0	0.084	0	0	0.147	0.337	0	0
Singapore	0.411	0.558	0.011	0.042	0.074	0.253	0.063	0.179	0	0.042	0.137	0.337	0	0	0	0.116	0.011	0.179	0.232	0.411
Thailand	0.488	0.512	0.163	0.233	0.512	0.651	0.14	0.349	0.442	0.535	0.116	0.395	0.419	0.488	0.023	0.093	0	0	0	0.116
Viet Nam	0	0.091	0	0	0	0.182	0	0	0	0	0	0	0	0.091	0	0	0	0	0	0

Note: The table reports the rejection frequencies at different significance levels for rolling window mixed- and low-frequency Granger causality tests with a window size of 50 quarters for the horizons $h\epsilon\{1,2,3,4,6\}$. For each rolling window, the Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ replications. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. $H_0\colon CP \nRightarrow EG$ ($\nRightarrow$ means "does not Granger-cause"). All variables are mean-centred and annual log-differenced.

Table B.53 Rejection Frequencies at Different Significant Levels for Rolling Window Granger Causality Tests, Thompson Reuters

Panel A: Mixed-frequency model																				
	horizon = 1				**horizon = 2**				**horizon = 3**				**horizon = 4**				**horizon = 6**			
	EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG		EG ⇏ CP		CP ⇏ EG	
	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																				
Australia	0.032	0.063	0.011	0.042	0.021	0.084	0.147	0.295	0	0	0.221	0.337	0	0	0.358	0.442	0	0.011	0.021	0.053
Bolivia	0	0.018	0	0.164	0	0.018	0	0.018	0	0.109	0	0	0	0.018	0	0	0	0.164	0	0
Canada	0	0	0.221	0.442	0	0	0.147	0.4	0	0	0.105	0.168	0.042	0.074	0.084	0.326	0	0	0.095	0.211
Chile	0.194	0.581	0.065	0.323	0	0	0.387	0.387	0	0	0	0.097	0	0	0	0.032	0	0.097	0	0
Denmark	0	0	0.118	0.294	0.039	0.059	0.137	0.353	0	0	0	0.118	0	0	0	0.118	0.118	0.314	0.098	0.196
Ecuador	0.2	0.2	0.6	0.933	0	0	0.333	0.467	0	0	0.333	0.4	0	0.067	0.133	0.333	0	0	0	0
Kazakhstan	0	0	0.973	0.973	0	0	0.703	0.865	0	0	0	0	0	0	0	0.054	0	0	0	0
New Zealand	0	0.015	0.076	0.091	0.091	0.121	0.182	0.379	0	0	0.197	0.727	0	0	0	0.076	0	0	0.015	0.045
Norway	0	0	0.084	0.242	0.021	0.032	0.105	0.221	0	0.011	0	0.021	0	0	0	0.011	0.084	0.137	0	0
Peru	0.211	0.253	0.211	0.368	0.063	0.168	0.137	0.221	0.011	0.032	0	0.021	0	0	0	0.011	0.042	0.126	0	0
South Africa	0	0	0.347	0.421	0	0	0.305	0.421	0	0	0.063	0.305	0	0	0	0.021	0.042	0.053	0.021	0.063
Venezuela	0	0.045	0.182	0.409	0	0	0.227	0.318	0	0	0.5	1	0	0	0.545	0.909	0	0	0	0.045
Importers																				
Czech Republic	0.065	0.129	0.226	0.355	0.097	0.129	0.097	0.548	0.065	0.097	0.097	0.194	0.097	0.097	0	0.065	0	0	0.29	0.29
Dominican Republic	0	0	0	0.027	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Hungary	0	0.029	0.514	0.829	0	0	0.457	0.657	0	0	0	0	0	0	0	0	0	0	0	0
Luxembourg	0.267	0.4	1	1	0	0	1	1	0	0	0	0.6	0	0	0	0.067	0	0.067	0	0
Malta	0	0.4	0	0	0	0	0	0	0.133	0.267	0	0	0.133	0.533	0	0	0	0	0.133	0.2
Philippines	0.033	0.11	0.044	0.11	0.099	0.209	0.099	0.264	0	0.033	0.044	0.088	0.055	0.231	0	0.022	0	0.011	0.033	0.099
Slovakia	0	0	0.143	0.2	0	0	0.143	0.686	0	0	0.143	0.543	0	0	0.086	0.086	0	0	0.143	0.229
Slovenia	0	0.029	0.657	0.743	0.029	0.057	0.571	0.829	0	0	0	0.057	0	0	0	0	0	0	0.029	0.057
Both (Hybrid)																				
Bahrain	0	0.054	0.011	0.076	0.033	0.098	0.065	0.163	0.033	0.163	0.054	0.185	0.098	0.207	0.207	0.228	0.033	0.054	0.033	0.13
Belgium	0	0	1	1	0	0.029	0.8	0.829	0	0	0.029	0.286	0	0	0	0.029	0	0	0	0
Hong Kong	0.011	0.032	0.358	0.474	0.011	0.011	0.158	0.326	0	0.021	0.074	0.137	0.021	0.063	0.179	0.211	0	0.032	0.326	0.495
Estonia	0	0	1	1	0	0.029	0.743	0.886	0	0	0.057	0.2	0	0	0	0.343	0	0	0	0
Iceland	0	0	0.333	0.519	0	0	0	0.037	0	0	0	0	0	0	0	0	0	0	0	0.148
Israel	0	0	0.857	0.943	0	0	0.6	0.686	0	0	0.229	0.229	0	0	0.229	0.229	0.143	0.229	0	0.143
Latvia	0	0	0.029	0.171	0	0	0.171	0.571	0	0	0.514	0.657	0	0	0.2	0.657	0	0	0	0
Netherlands	0	0	0.839	0.968	0	0	0.71	0.774	0	0	0	0.065	0	0	0	0	0	0	0	0
Serbia	0.194	0.258	0.29	0.516	0	0	0.161	0.419	0	0	0	0.097	0	0	0	0	0	0	0	0
Seychelles	0	0.011	0.042	0.137	0	0.011	0	0.116	0	0	0.095	0.105	0.011	0.032	0.011	0.063	0.021	0.021	0	0
Singapore	0.116	0.221	0.263	0.347	0.011	0.074	0	0.042	0.042	0.168	0	0	0	0	0	0.021	0.095	0.232	0	0.011
Thailand	0.349	0.465	0.209	0.535	0.349	0.395	0	0.07	0.372	0.395	0	0.07	0.372	0.372	0	0.093	0.023	0.116	0	0.023
Viet Nam	1	1	0	0	0.818	0.818	0	0	0	0	0	0	0	0	0	0	0	0.273	0.091	0.636

Panel B: Low-frequency model										
	horizon = 1		**horizon = 2**		**horizon = 3**		**horizon = 4**		**horizon = 6**	
	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG	EG ⇏ CP	CP ⇏ EG

	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%	5%	10%
Exporters																				
Australia	0.011	0.042	0	0.042	0	0	0.063	0.168	0	0	0.053	0.158	0	0	0	0.021	0	0	0	0
Bolivia	0.036	0.255	0.218	0.473	0	0.091	0	0.018	0.073	0.182	0	0	0.055	0.182	0	0.036	0.018	0.018	0	0.036
Canada	0	0.011	0	0	0	0	0.011	0.042	0	0	0.168	0.211	0	0	0.2	0.358	0	0.063	0.189	0.358
Chile	0.097	0.161	0.355	0.71	0	0.129	0.258	0.323	0	0	0.032	0.194	0	0	0	0	0	0	0	0
Denmark	0	0.039	0	0	0	0.02	0	0.059	0	0	0.157	0.255	0	0	0.216	0.353	0	0	0.216	0.373
Ecuador	0.867	0.933	1	1	0.733	0.933	0.6	0.6	1	1	0.133	0.333	0	0	0	0.133	0	0	0	0
Kazakhstan	0	0	0.649	0.865	0	0	0.081	0.189	0	0	0.081	0.081	0	0	0	0.054	0	0	0	0
New Zealand	0.076	0.197	0	0.061	0	0.045	0.03	0.061	0	0	0.121	0.227	0	0	0.106	0.106	0	0	0.015	0.03
Norway	0	0	0.095	0.211	0	0	0	0	0	0	0	0	0	0	0	0	0	0.084	0.032	0.095
Peru	0.242	0.253	0.095	0.316	0.158	0.211	0.011	0.042	0.063	0.126	0	0.042	0	0.011	0	0	0	0	0	0
South Africa	0	0.011	0.4	0.474	0	0	0.158	0.305	0	0	0.105	0.116	0	0.011	0.011	0.053	0.011	0.053	0	0.032
Venezuela	0	0	0.273	0.409	0	0	0.909	1	0	0	1	1	0	0	1	1	0	0	0	0.545
Importers																				
Czech Republic	0.065	0.065	0	0.129	0.065	0.065	0	0	0.097	0.097	0	0	0.097	0.097	0.032	0.065	0	0	0.065	0.226
Dominican Republic	0.108	0.27	0	0	0	0.162	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Hungary	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.057
Luxembourg	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.267	0.867	0	0
Malta	0	0	0	0	0	0	0	0	0	0	0	0	0	0.267	0.067	0.133	0.133	0.333	0	0
Philippines	0.132	0.143	0.11	0.176	0.066	0.11	0.011	0.022	0.099	0.11	0	0.033	0.121	0.165	0	0	0.011	0.176	0	0
Slovakia	0	0	0.314	0.686	0	0	0.286	0.4	0	0	0.086	0.114	0	0	0.229	0.286	0	0	0.143	0.257
Slovenia	0	0	0.086	0.257	0	0	0	0	0	0	0	0	0	0	0	0	0	0.057	0	0
Both (Hybrid)																				
Bahrain	0.065	0.12	0	0.022	0.033	0.152	0.087	0.098	0.033	0.076	0.152	0.239	0.087	0.185	0.217	0.239	0	0.011	0.25	0.272
Belgium	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0
Hong Kong	0.295	0.421	0.011	0.032	0.095	0.179	0.032	0.084	0.021	0.095	0.105	0.284	0.095	0.116	0.253	0.442	0.021	0.063	0.463	0.463
Estonia	0	0	0.486	0.971	0	0	0	0.029	0	0	0	0	0	0	0	0	0	0	0	0
Iceland	0	0	0	0.148	0	0	0	0.037	0	0	0	0.037	0	0	0.037	0.037	0	0	0	0
Israel	0	0	0.171	0.371	0	0	0.114	0.171	0	0	0	0.057	0	0	0	0	0	0	0	0
Latvia	0	0	0.629	0.714	0	0	0.371	0.6	0	0	0	0	0	0	0	0	0	0	0	0
Netherlands	0	0	0	0.226	0	0	0	0	0	0	0	0	0	0	0	0	0.032	0.097	0	0
Serbia	0.032	0.161	0.742	0.871	0.258	0.323	0.29	0.677	0	0.032	0.194	0.194	0	0	0.065	0.226	0	0	0	0
Seychelles	0.011	0.074	0.032	0.095	0	0.042	0.032	0.063	0	0	0.042	0.063	0	0.021	0	0.032	0	0	0	0
Singapore	0.347	0.516	0	0	0	0.116	0	0	0	0.011	0	0	0	0	0	0	0.063	0.095	0.063	0.126
Thailand	0.512	0.651	0	0	0.419	0.465	0	0.116	0.372	0.395	0.047	0.302	0.372	0.372	0.023	0.047	0	0	0	0.023
Viet Nam	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0	0.091

Note: The table reports the rejection frequencies at different significance levels for rolling window mixed- and low-frequency Granger causality tests with a window size of 50 quarters for the horizons $h\epsilon\{1,2,3,4,6\}$. For each rolling window, the Wald statistic p-values are computed based on the non-robust covariance matrix and Gonçalves and Kilian's (2004) bootstrap with $N = 999$ replications. "CP" denotes the commodity prices, while "EG" denotes the economic growth variables. H_0: $CP \nRightarrow EG$ ($\nRightarrow$ means "does not Granger-cause"). All variables are mean-centred and annual log-differenced.

Table B.54 Rejection Frequencies at Different Significant Levels for Rolling Window Granger Causality Tests, IMF

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	-3.308	8.121***	5.279**	0.237	7.396***	4.670***
Bolivia	-1.599	-1.759	0.109	1.127	-1.004	1.248*
Canada	17.568***	11.321***	14.419***	2.514**	13.409***	16.462***
Chile	-0.584	4.871**	9.247***	0.250	4.442**	8.709***
Denmark	2.799*	12.149***	9.089***	-0.273	9.839***	6.854***
Ecuador	6.059***	8.489***	11.320***	5.291***	7.489***	10.188***
Kazakhstan	5.307**	13.661***	16.661***	4.689***	14.38***	17.909***
New Zealand	0.640	0.056	1.140	1.523*	-3.283	-2.360
Norway	3.225*	29.607***	7.436***	2.570**	31.719***	8.929***
Peru	-10.710	-4.851	-11.727	-9.474	-6.067	-12.625
South Africa	19.225***	40.478***	52.065***	7.753***	41.425***	52.142***
Venezuela	2.558*	0.245	2.090	6.125***	-0.047	2.544**
Importers						
Czech Republic	4.710**	-0.634	1.212	1.555*	0.330	2.449**
Dominican Republic	-1.803	-1.848	-1.440	-0.150	-1.732	-1.267
Hungary	2.654*	0.622	2.326*	0.990	1.571	3.567**
Luxembourg	-0.534	-0.774	-0.692	-0.374	0.071	0.161
Malta	-0.774	-1.507	-0.881	-0.670	-1.428	-0.769
Philippines	9.351***	32.930***	16.614***	-1.983	35.115***	18.372***
Slovakia	2.778*	1.234	4.832**	1.457*	0.773	4.481***
Slovenia	2.820*	0.863	2.736*	-0.531	1.497	3.528**
Both (Hybrid)						
Bahrain	-1.118	-3.276	-4.286	-1.423	-3.789	-4.183
Belgium	7.940***	7.357***	8.426***	-0.039	6.215***	7.088***
Hong Kong	35.339***	25.603***	32.124***	0.042	7.878***	12.761***
Estonia	5.877***	8.835***	11.120***	1.991**	11.283***	13.896***
Iceland	-0.398	-2.335	-2.883	-0.100	-2.370	-2.772
Israel	23.883***	21.615***	22.506***	-0.465	16.485***	16.854***
Latvia	2.428*	0.400	3.031*	5.886***	1.500	4.621***
Netherlands	5.771***	7.334***	2.343*	-0.187	8.340***	3.207**
Serbia	2.511*	0.168	4.853**	2.628**	-0.930	3.591**
Seychelles	0.173	-4.023	-1.947	-0.312	-3.097	-0.969
Singapore	23.99***	31.085***	28.805***	0.894	15.780***	13.352***
Thailand	5.204**	14.177***	15.216***	0.574	1.963*	2.282**
Viet Nam	0.865	0.941	1.648	1.172	0.643	1.296*

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Clark and McCracken's (2001) critical values. All variables are mean-centred and annual log-differenced.

Table B.55 Tests for Out-of-Sample Forecasting Ability – Regression Based Forecast Models, Reuters/Jeffries

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	-2.132	7.884***	6.336***	3.093**	1.106	0.633
Bolivia	-0.128	-1.906	0.481	1.724*	-2.068	0.550
Canada	8.789***	6.992***	8.207***	0.775	6.172***	7.588***
Chile	0.035	10.623***	16.775***	1.019	8.701***	14.507***
Denmark	3.247*	12.215***	8.443***	-0.812	7.785***	4.460***
Ecuador	8.882***	9.252***	12.590***	7.266***	6.810***	9.644***
Kazakhstan	4.812**	16.258***	19.135***	5.726***	18.387***	21.570***
New Zealand	5.587**	1.861	2.748*	6.752***	-2.264	-0.994
Norway	1.350	20.727***	0.343	1.747*	23.124***	1.939*
Peru	-0.148	1.378	-0.608	-2.672	0.336	-1.756
South Africa	16.957***	52.166***	51.700***	1.225*	44.284***	43.732***
Venezuela	5.261**	2.506	5.766***	6.068***	2.158*	5.653***
Importers						
Czech Republic	3.346**	-1.839	-0.520	1.607*	-1.716	-0.238
Dominican Republic	-1.082	0.366	0.798	-0.559	-0.903	-0.398
Hungary	0.962	-0.450	1.118	0.320	-0.313	1.418*
Luxembourg	-0.720	-1.142	-1.098	-0.529	-0.751	-0.707
Malta	-1.298	-1.810	-1.305	-1.312	-1.774	-1.236
Philippines	3.998**	28.492***	11.190***	-2.242	29.011***	11.579***
Slovakia	2.221	3.787**	7.700***	1.206	1.656	5.348***
Slovenia	2.670*	1.502	3.113*	-0.459	1.474	3.148**
Both (Hybrid)						
Bahrain	-1.289	-3.428	-1.697	-1.390	-3.222	-1.022
Belgium	7.348***	7.942***	8.568***	0.146	6.354***	6.801***
Hong Kong	15.848***	18.518***	25.324***	-4.054	9.352***	15.541***
Estonia	4.759**	5.067**	6.713***	2.401**	6.588***	8.447***
Iceland	-0.315	-2.268	-2.907	-0.013	-2.546	-3.051
Israel	19.480***	24.109***	24.531***	-1.122	19.644***	19.828***
Latvia	2.770*	-0.347	1.993	3.309**	-0.094	2.512**
Netherlands	4.195**	4.317**	-0.360	-0.601	4.933***	0.113
Serbia	1.613	0.543	4.548**	1.834*	-0.662	3.377**
Seychelles	0.305	-3.833	-2.185	0.455	-2.552	-1.009
Singapore	22.547***	26.768***	26.117***	-0.275	18.786***	17.936***
Thailand	4.431**	14.180***	14.891***	0.065	6.068***	6.481***
Viet Nam	1.669	2.473	3.365**	1.346*	1.186	1.880*

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Clark and McCracken's (2001) critical values. All variables are mean-centred and annual log-differenced.

Table B.56 Tests for Out-of-Sample Forecasting Ability – Regression Based Forecast Models, Goldman Sachs

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	-1.121	0.573	0.500	0.132	0.828	0.659
Bolivia	5.831***	-0.087	5.098**	5.063***	0.892	6.279***
Canada	22.820***	-1.961	-0.186	2.131**	1.158	3.451**
Chile	2.843*	7.459***	12.235***	7.028***	6.893***	11.550***
Denmark	4.874**	0.799	-1.796	-0.298	4.218**	1.223*
Ecuador	2.977*	0.666	1.864	3.089**	0.214	1.341*
Kazakhstan	14.071***	12.453***	14.729***	4.03***	10.756***	12.446***
New Zealand	3.115*	1.331	1.930	0.785	-0.098	0.041
Norway	0.043	13.224***	-3.541	-0.381	16.950***	-1.247
Peru	-5.310	-1.031	-8.798	-7.390	-0.495	-8.703
South Africa	16.053***	32.423***	28.923***	6.378***	39.693***	35.730***
Venezuela	5.304**	1.471	2.955*	3.202**	-0.310	1.319*
Importers						
Czech Republic	5.517**	-0.318	0.459	1.271*	-0.099	0.965
Dominican Republic	-1.896	-3.308	-3.094	-1.215	-3.154	-2.948
Hungary	3.801**	-3.990	-2.581	0.027	-2.110	-0.194
Luxembourg	-0.997	-1.285	-1.321	-0.084	-0.623	-0.674
Malta	-0.525	-0.916	-0.230	-0.467	-0.805	-0.093
Philippines	19.827***	52.242***	32.046***	6.039***	59.994***	38.284***
Slovakia	1.222	3.070*	6.064***	5.572***	6.477***	10.908***
Slovenia	6.192***	-1.465	-0.309	0.196	0.809	2.469**
Both (Hybrid)						
Bahrain	-4.415	-6.019	-6.623	-3.243	-3.818	-4.369
Belgium	20.404***	4.601**	4.438**	-0.199	6.612***	6.302***
Hong Kong	30.761***	22.516***	33.411***	-1.353	13.050***	22.076***
Estonia	9.583***	-0.796	0.660	3.638**	2.582*	4.602***
Iceland	-0.182	-0.260	0.103	-0.251	-0.208	0.116
Israel	23.354***	10.465***	10.983***	1.991**	13.388***	13.702***
Latvia	6.357***	-3.734	-1.831	7.776***	-2.123	0.504
Netherlands	13.696***	8.937***	2.656*	0.551	10.331***	3.728***
Serbia	-0.292	-3.073	-2.093	0.084	-4.159	-3.053
Seychelles	6.587***	-2.103	-1.420	-1.706	-1.655	-0.729
Singapore	17.691***	44.548***	45.477***	-0.541	33.923***	34.498***
Thailand	4.634**	31.360***	34.559***	-1.579	12.570***	14.214***
Viet Nam	-0.142	-0.828	-0.588	-0.311	-0.938	-0.671

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Clark and McCracken's (2001) critical values. All variables are mean-centred and annual log-differenced.

Table B.57 Tests for Out-of-Sample Forecasting Ability – Regression Based Forecast Models, Moody's

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	-0.797	9.461***	9.005***	2.380**	4.603**	4.876***
Bolivia	1.911	0.201	4.255**	3.719***	0.505	4.433***
Canada	19.587***	3.989**	6.072***	1.073	5.060***	7.048***
Chile	1.032	3.993**	8.324***	2.002**	5.344***	9.976***
Denmark	2.245	4.412**	1.479	-0.502	7.448***	4.083***
Ecuador	5.779***	1.723	3.301**	5.136***	1.476	3.020**
Kazakhstan	11.302***	21.637***	25.013***	7.027***	17.737***	20.582***
New Zealand	0.174	-0.711	-0.175	2.058**	-2.503	-2.028
Norway	-0.597	15.704***	-3.042	-0.519	18.425***	-0.966
Peru	-7.045	1.114	-5.054	-4.838	-0.920	-6.780
South Africa	13.380***	43.846***	44.235***	4.848***	50.596***	50.571***
Venezuela	3.589**	0.788	2.873*	4.559***	0.076	2.387**
Importers						
Czech Republic	3.832**	0.139	1.279	0.800	1.351	2.706**
Dominican Republic	-2.152	-1.133	-0.962	-1.173	-1.874	-1.663
Hungary	2.559*	-1.690	-0.215	0.022	-0.324	1.458*
Luxembourg	-1.014	-1.348	-1.301	-0.384	-0.271	-0.220
Malta	-0.820	-1.389	-0.785	-0.844	-1.315	-0.716
Philippines	13.140***	39.776***	19.984***	0.754	39.408***	18.874***
Slovakia	2.237	4.011**	7.745***	3.597**	7.549***	12.491***
Slovenia	3.404**	1.299	2.765*	-0.314	2.954**	4.696***
Both (Hybrid)						
Bahrain	-4.052	-6.383	-6.218	-1.992	-4.603	-3.383
Belgium	10.902***	7.235***	7.642***	-0.523	6.568***	6.670***
Hong Kong	26.902***	35.117***	45.254***	-0.478	16.023***	23.856***
Estonia	7.903***	3.396*	5.233**	3.204**	6.679***	9.003***
Iceland	-0.089	-1.331	-1.756	-0.083	-1.088	-1.388
Israel	21.483***	21.629***	21.998***	-0.690	20.200***	20.137***
Latvia	2.266	-2.030	0.202	6.314***	-0.277	2.610**
Netherlands	5.116**	9.246***	3.285**	-0.116	11.802***	5.279***
Serbia	0.332	-2.553	0.046	1.454*	-2.753	-0.076
Seychelles	1.401	-4.567	-3.450	-0.358	-2.587	-1.282
Singapore	17.733***	37.937***	37.835***	0.643	24.622***	24.292***
Thailand	2.134	18.109***	19.423***	-0.218	6.839***	7.422***
Viet Nam	0.896	0.689	1.324	0.575	-0.017	0.500

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Clark and McCracken's (2001) critical values. All variables are mean-centred and annual log-differenced.

Table B.58 Tests for Out-of-Sample Forecasting Ability – Regression Based Forecast Models, Thompson Reuters

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	-0.122	0.671	1.141	-0.075	1.182	1.827*
Bolivia	7.987***	0.469	4.512**	3.971***	0.490	5.579***
Canada	25.743***	4.084**	5.943***	1.977**	3.195**	5.146***
Chile	3.724**	7.081***	11.719***	5.120***	10.710***	16.322***
Denmark	6.775***	7.473***	4.734**	0.056	9.203***	5.529***
Ecuador	3.250*	1.556	3.001*	2.571**	1.122	2.511**
Kazakhstan	18.708***	21.913***	24.781***	2.397**	9.424***	11.371***
New Zealand	0.707	3.600*	3.584**	0.363	-1.152	-1.132
Norway	2.861*	15.364***	-0.667	-1.456	14.186***	-3.536
Peru	-5.871	0.562	-5.650	-4.628	0.726	-5.559
South Africa	24.212***	51.095***	46.869***	10.051***	56.063***	52.360***
Venezuela	1.317	0.068	2.228	2.265**	-0.131	2.147**
Importers						
Czech Republic	9.547***	0.273	1.158	1.206	1.355	2.572**
Dominican Republic	-1.026	0.701	0.902	-0.568	0.316	0.514
Hungary	12.533***	-1.678	-0.022	-0.226	-0.951	0.942
Luxembourg	0.064	-0.989	-0.979	-0.075	-0.255	-0.292
Malta	-0.569	-1.014	-0.403	-0.704	-1.252	-0.606
Philippines	17.057***	61.462***	37.287***	5.792***	53.634***	30.601***
Slovakia	2.135	6.467***	10.124***	4.605***	9.059***	13.951***
Slovenia	12.355***	2.148	3.874**	0.462	2.989**	4.851***
Both (Hybrid)						
Bahrain	-2.934	-4.242	-4.106	-2.527	-5.111	-4.598
Belgium	34.741***	14.032***	14.911***	-0.049	8.070***	8.233***
Hong Kong	13.190***	42.603***	52.883***	-1.521	27.562***	36.617***
Estonia	13.632***	4.662**	6.926***	3.434**	5.951***	8.238***
Iceland	0.722	-1.666	-1.881	-0.033	-1.112	-1.466
Israel	8.668***	22.803***	23.447***	1.373*	22.325***	23.167***
Latvia	6.981***	-1.752	0.602	9.894***	-0.091	2.804**
Netherlands	21.104***	8.677***	3.219*	0.276	11.638***	5.337***
Serbia	1.633	-2.423	-0.121	2.489**	-2.538	-0.164
Seychelles	1.095	-1.105	-0.036	-1.494	-1.148	0.156
Singapore	11.295***	50.715***	52.077***	-1.152	33.588***	35.000***
Thailand	3.305**	29.267***	30.351***	-0.251	10.508***	10.902***
Viet Nam	0.344	-0.126	0.236	-0.029	-0.695	-0.344

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Clark and McCracken's (2001) critical values. All variables are mean-centred and annual log-differenced.

Table B.59 Tests for Out-of-Sample Forecasting Ability – Regression Based Forecast Models, IMF

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	1.137	-4.364***	-4.643***	1.176	-4.329***	-4.617***
Bolivia	0.408	-2.201**	-3.357***	0.388	-2.198**	-3.354***
Canada	0.042	-3.633***	-3.687***	0.463	-3.606***	-3.665***
Chile	-0.365	-3.320***	-3.582***	-0.413	-3.381***	-3.644***
Denmark	0.762	-2.264**	-2.250**	1.035	-2.254**	-2.241**
Ecuador	-2.237**	-2.756***	-2.745***	-2.235**	-2.772***	-2.757***
Kazakhstan	-1.615	-0.714	-1.012	-1.423	-0.685	-0.973
New Zealand	-0.368	-2.799***	-2.997***	-0.323	-2.789***	-2.984***
Norway	0.303	-3.505***	-2.574**	0.348	-3.483***	-2.547**
Peru	3.052***	-4.271***	-3.877***	3.059***	-4.141***	-3.835***
South Africa	-0.206	-6.435***	-7.846***	0.078	-6.397***	-7.823***
Venezuela	-1.553	-3.003***	-2.690***	-1.424	-2.972***	-2.663***
Importers						
Czech Republic	-0.172	-2.729***	-3.195***	0.085	-2.69***	-3.167***
Dominican Republic	0.571	-1.008	-1.047	0.545	-1.010	-1.048
Hungary	-0.353	-2.279**	-2.339**	-0.015	-2.262**	-2.325**
Luxembourg	1.182	0.739	0.406	1.174	0.755	0.416
Malta	2.492**	-4.872***	-4.821***	2.518**	-4.797***	-4.777***
Philippines	1.512	-6.495***	-5.720***	1.857*	-6.431***	-5.641***
Slovakia	-1.332	-1.568	-1.798*	-1.247	-1.561	-1.792*
Slovenia	1.535	-2.643***	-2.590***	2.246**	-2.629***	-2.579***
Both (Hybrid)						
Bahrain	2.566**	-0.165	-2.286**	2.565**	-0.119	-2.228**
Belgium	0.322	-2.084**	-1.866*	0.924	-2.070**	-1.855*
Hong Kong	-0.029	-3.775***	-4.135***	0.753	-3.710***	-4.074***
Estonia	-2.369**	-1.947*	-1.903*	-2.213**	-1.942*	-1.898*
Iceland	0.942	-0.094	-0.569	0.860	-0.094	-0.576
Israel	-0.200	-2.680***	-3.080***	1.261	-2.597***	-2.991***
Latvia	-2.546**	-2.545**	-2.455**	-2.507**	-2.546**	-2.455**
Netherlands	0.521	-1.896*	-2.117**	1.229	-1.856*	-2.055**
Serbia	1.676*	-1.657*	-3.054***	1.718*	-1.543	-2.971***
Seychelles	0.865	-6.757***	-8.032***	0.930	-6.742***	-8.016***
Singapore	0.042	-3.459***	-3.546***	0.589	-3.384***	-3.477***
Thailand	0.204	-1.322	-1.560	0.458	-1.236	-1.468
Viet Nam	-0.208	0.189	-0.032	-0.177	0.185	-0.021

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Diebold and Mariano's (1995) critical values. All variables are mean-centred and annual log-differenced.

Table B.60 Tests for Out-of-Sample Forecasting Ability – Combination Forecast Models, Reuters/Jeffries

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	0.602	-5.046***	-5.293***	0.633	-5.020***	-5.307***
Bolivia	-0.269	-2.315**	-3.328***	-0.298	-2.319**	-3.325***
Canada	0.722	-3.621***	-3.684***	1.059	-3.591***	-3.661***
Chile	-0.636	-3.496***	-3.736***	-0.670	-3.583***	-3.817***
Denmark	1.726*	-2.252**	-2.242**	2.024**	-2.242**	-2.232**
Ecuador	-2.064**	-2.771***	-2.745***	-2.047**	-2.788***	-2.756***
Kazakhstan	-1.242	-0.632	-0.950	-1.194	-0.636	-0.945
New Zealand	-0.819	-2.853***	-3.021***	-0.704	-2.835***	-3.003***
Norway	0.615	-3.578***	-2.617***	0.723	-3.534***	-2.566**
Peru	2.189**	-5.853***	-4.186***	2.313**	-5.798***	-4.158***
South Africa	-0.809	-6.513***	-7.960***	-0.330	-6.482***	-7.943***
Venezuela	-1.987**	-3.039***	-2.740***	-1.783*	-2.994***	-2.714***
Importers						
Czech Republic	-0.133	-2.873***	-3.280***	0.037	-2.854***	-3.265***
Dominican Republic	1.039	-0.970	-1.008	0.939	-0.974	-1.010
Hungary	0.109	-2.270**	-2.336**	0.336	-2.262**	-2.328**
Luxembourg	1.085	0.808	0.479	1.069	0.835	0.499
Malta	2.906***	-4.445***	-4.861***	2.959***	-4.283***	-4.798***
Philippines	1.677*	-6.861***	-6.174***	2.011**	-6.828***	-6.147***
Slovakia	-0.419	-1.552	-1.790*	-0.166	-1.534	-1.773*
Slovenia	1.721*	-2.639***	-2.591***	2.488**	-2.621***	-2.577***
Both (Hybrid)						
Bahrain	1.393	-0.496	-2.779***	1.423	-0.459	-2.744***
Belgium	0.827	-2.067**	-1.854	1.307	-2.040**	-1.834*
Hong Kong	1.004	-3.736***	-4.115***	1.844	-3.655***	-4.041***
Estonia	-2.385**	-1.930	-1.890	-2.153**	-1.926	-1.886
Iceland	0.715	-0.145	-0.570	0.664	-0.157	-0.585
Israel	-0.116	-2.637***	-3.007***	1.438	-2.548**	-2.906***
Latvia	-2.180**	-2.535**	-2.447**	-2.169**	-2.535**	-2.447**
Netherlands	1.020	-1.852	-2.089**	1.778*	-1.786*	-1.994**
Serbia	1.737*	-1.644	-2.982***	1.810*	-1.512	-2.876***
Seychelles	0.688	-6.785***	-8.022***	0.786	-6.764***	-7.997***
Singapore	-0.486	-3.551***	-3.641***	0.801	-3.441***	-3.542***
Thailand	0.045	-1.368	-1.613	0.470	-1.275	-1.518
Viet Nam	0.194	0.407	0.272	0.271	0.459	0.335

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Diebold and Mariano's (1995) critical values. All variables are mean-centred and annual log-differenced.

Table B.61 Tests for Out-of-Sample Forecasting Ability – Combination Forecast Models, Goldman Sachs

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	1.106	-4.763***	-4.915***	1.118	-4.751***	-4.907***
Bolivia	-0.605	-2.533**	-3.394***	-0.490	-2.502**	-3.358***
Canada	0.743	-3.610***	-3.677***	1.081	-3.580***	-3.653***
Chile	0.009	-2.304**	-2.658***	0.100	-2.198**	-2.563**
Denmark	1.171	-2.251**	-2.234**	1.375	-2.239**	-2.223**
Ecuador	-1.529	-2.604***	-2.618***	-1.493	-2.606***	-2.618***
Kazakhstan	-0.898	-0.780	-1.225	-0.644	-0.661	-1.104
New Zealand	0.210	-2.701***	-2.943***	0.293	-2.673***	-2.918***
Norway	1.409	-3.442***	-2.471**	1.425	-3.421***	-2.449**
Peru	1.892*	-4.525***	-4.351***	1.899*	-4.293***	-4.297***
South Africa	0.048	-6.453***	-7.845***	0.324	-6.425***	-7.821***
Venezuela	-1.236	-3.079***	-2.797***	-1.122	-3.051***	-2.772***
Importers						
Czech Republic	0.649	-2.526**	-2.963***	0.825	-2.483**	-2.923***
Dominican Republic	1.813*	-0.875	-0.891	1.828*	-0.865	-0.882
Hungary	0.454	-2.255**	-2.325**	0.806	-2.243**	-2.315**
Luxembourg	1.120	0.788	0.462	0.821	0.750	0.427
Malta	1.050	-4.533***	-4.674***	1.037	-4.451***	-4.633***
Philippines	-0.044	-6.644***	-5.719***	0.082	-6.527***	-5.589***
Slovakia	-0.371	-1.739*	-1.933*	-0.318	-1.735*	-1.927*
Slovenia	0.537	-2.629***	-2.576***	0.886	-2.609***	-2.561**
Both (Hybrid)						
Bahrain	3.274***	0.009	-1.874*	3.319***	0.048	-1.829*
Belgium	1.066	-2.059**	-1.848*	2.037**	-2.022**	-1.821*
Hong Kong	0.368	-3.759***	-4.14***	1.028	-3.657***	-4.047***
Estonia	-1.728*	-1.970**	-1.924*	-1.614	-1.963**	-1.918*
Iceland	1.326	-0.041	-0.534	1.357	-0.023	-0.525
Israel	-0.220	-2.630***	-3.039***	0.256	-2.541**	-2.947***
Latvia	-2.356**	-2.567**	-2.473**	-2.192**	-2.566**	-2.473**
Netherlands	0.893	-1.736*	-1.906*	1.274	-1.634	-1.764*
Serbia	2.010**	0.489	-1.132	1.997**	0.609	-0.955
Seychelles	1.043	-6.731***	-7.929***	1.390	-6.708***	-7.900***
Singapore	0.199	-3.448***	-3.564***	0.807	-3.347***	-3.482***
Thailand	1.127	-1.149	-1.406	1.401	-1.070	-1.333
Viet Nam	1.105	1.602	1.253	1.187	1.646	1.294

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Diebold and Mariano's (1995) critical values. All variables are mean-centred and annual log-differenced.

Table B.62 Tests for Out-of-Sample Forecasting Ability – Combination Forecast Models, Moody's

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	0.649	-4.809***	-5.050***	0.723	-4.767***	-5.034***
Bolivia	-1.208	-2.509**	-3.473***	-1.125	-2.501**	-3.459***
Canada	0.354	-3.638***	-3.703***	0.723	-3.607***	-3.679***
Chile	-0.005	-2.644***	-3.036***	0.039	-2.581***	-2.992***
Denmark	1.189	-2.263**	-2.248**	1.459	-2.251**	-2.237**
Ecuador	-2.382**	-2.739***	-2.728***	-2.368**	-2.752***	-2.738***
Kazakhstan	-1.707*	-0.810	-1.157	-1.344	-0.734	-1.074
New Zealand	-0.072	-2.734***	-2.939***	-0.013	-2.710***	-2.911***
Norway	1.572	-3.449***	-2.455**	1.571	-3.432***	-2.437**
Peru	2.436**	-4.445***	-3.885***	2.469**	-4.324***	-3.838***
South Africa	0.183	-6.418***	-7.857***	0.661	-6.378***	-7.832***
Venezuela	-1.251	-2.988***	-2.709***	-1.081	-2.955***	-2.675***
Importers						
Czech Republic	0.645	-2.558**	-2.997***	0.823	-2.520**	-2.962***
Dominican Republic	1.424	-0.862	-0.903	1.383	-0.856	-0.898
Hungary	0.297	-2.263**	-2.332**	0.809	-2.250**	-2.321**
Luxembourg	0.984	0.707	0.381	0.872	0.696	0.365
Malta	1.530	-4.331***	-4.596***	1.512	-4.182***	-4.513***
Philippines	0.319	-6.913***	-6.031***	0.508	-6.871***	-5.977***
Slovakia	-0.681	-1.678*	-1.888*	-0.580	-1.675*	-1.884*
Slovenia	1.058	-2.638***	-2.587***	1.614	-2.620***	-2.573**
Both (Hybrid)						
Bahrain	2.864***	-0.223	-2.233**	2.776***	-0.185	-2.181**
Belgium	1.183	-2.057**	-1.848*	1.581	-2.029**	-1.828*
Hong Kong	-0.191	-3.765***	-4.142***	0.673	-3.648***	-4.038***
Estonia	-2.225**	-1.963**	-1.918*	-2.055**	-1.955*	-1.912*
Iceland	1.140	-0.104	-0.576	1.126	-0.097	-0.575
Israel	-0.591	-2.680***	-3.092***	1.566	-2.567**	-2.973***
Latvia	-2.719***	-2.559**	-2.466**	-2.625***	-2.561**	-2.467**
Netherlands	1.029	-1.766*	-1.949*	1.504	-1.666*	-1.812*
Serbia	2.026**	-0.833	-2.406**	2.051**	-0.605	-2.218**
Seychelles	0.615	-6.773***	-7.998***	0.833	-6.736***	-7.953***
Singapore	-0.186	-3.452***	-3.569***	0.361	-3.354***	-3.486***
Thailand	0.303	-1.273	-1.546	0.560	-1.186	-1.463
Viet Nam	-0.154	0.355	0.099	-0.107	0.353	0.111

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Diebold and Mariano's (1995) critical values. All variables are mean-centred and annual log-differenced.

Table B.63 Tests for Out-of-Sample Forecasting Ability – Combination Forecast Models, Thompson Reuters

	Panel A: Mixed-frequency model			Panel B: Low-frequency model		
	AR Benchmark	RW Benchmark	RWWD Benchmark	AR Benchmark	RW Benchmark	RWWD Benchmark
Exporters						
Australia	1.056	-4.654***	-4.799***	1.098	-4.631***	-4.782***
Bolivia	-0.504	-2.500**	-3.367***	-0.408	-2.471**	-3.333***
Canada	-0.320	-3.662***	-3.716***	0.235	-3.637***	-3.696***
Chile	-0.250	-2.574**	-2.870***	-0.150	-2.473**	-2.779***
Denmark	0.768	-2.258**	-2.249**	1.141	-2.249**	-2.241**
Ecuador	-2.210**	-2.631***	-2.647***	-2.180**	-2.634***	-2.649***
Kazakhstan	-1.031	-0.602	-1.019	-0.305	-0.390	-0.810
New Zealand	0.374	-2.734***	-3.009***	0.454	-2.716***	-2.994***
Norway	1.745*	-3.545***	-2.522**	1.839*	-3.513***	-2.481**
Peru	2.243**	-4.552***	-3.729***	2.289**	-4.403***	-3.649***
South Africa	-0.097	-6.434***	-7.810***	0.284	-6.381***	-7.771***
Venezuela	-0.753	-3.003***	-2.756***	-0.686	-2.983***	-2.739***
Importers						
Czech Republic	0.214	-2.575***	-3.043***	0.712	-2.469**	-2.947***
Dominican Republic	1.259	-0.986	-1.011	1.295	-0.981	-1.005
Hungary	0.050	-2.272**	-2.338**	0.910	-2.253**	-2.323**
Luxembourg	0.908	0.686	0.371	0.671	0.682	0.364
Malta	2.027**	-4.64***	-4.774***	2.035**	-4.557***	-4.743***
Philippines	-0.273	-6.762***	-5.884***	-0.054	-6.656***	-5.762***
Slovakia	-0.568	-1.708*	-1.910*	-0.486	-1.698*	-1.899*
Slovenia	-0.448	-2.67***	-2.607***	0.158	-2.649***	-2.591***
Both (Hybrid)						
Bahrain	3.427***	-0.358	-2.400**	3.478***	-0.329	-2.365**
Belgium	-0.944	-2.095**	-1.878*	2.017**	-2.060**	-1.851*
Hong Kong	0.466	-3.743***	-4.127***	0.989	-3.688***	-4.078***
Estonia	-2.069**	-1.967**	-1.922*	-1.892*	-1.956*	-1.912*
Iceland	0.979	-0.012	-0.517	1.020	0.012	-0.501
Israel	-0.540	-2.679***	-3.081***	-0.153	-2.637***	-3.039***
Latvia	-2.761***	-2.584***	-2.487**	-2.643***	-2.586***	-2.488**
Netherlands	0.703	-1.825*	-2.045**	1.641	-1.679*	-1.841*
Serbia	1.729*	-0.152	-1.581	1.771*	0.009	-1.406
Seychelles	1.094	-6.712***	-7.913***	1.207	-6.693***	-7.887***
Singapore	0.659	-3.489***	-3.599***	1.523	-3.419***	-3.545***
Thailand	0.142	-1.250	-1.487	0.344	-1.181	-1.420
Viet Nam	-0.020	1.101	0.646	0.663	1.215	0.725

Note: The table reports the re-scaled MSFE differences between the model and the benchmark forecasts. Negative values imply that the commodity-based model forecasts better than the benchmark model. Asterisks denote rejections of the null hypothesis that the benchmark model is better in favour of the alternative hypothesis that the commodity-based model is better at 1% (***), 5% (**) and 10% (*) significance levels, respectively, using Diebold and Mariano's (1995) critical values. All variables are mean-centred and annual log-differenced.

Table B.64 Tests for Out-of-Sample Forecasting Ability – Combination Forecast Models, IMF

Appendix C

C.1 Stock market indexes for each country/region

Country/Region	Code (Datastream)	Stock Market Index	Data Period (From/To)	
Panel A: Northern Africa				
Egypt	EYSHRPRCF	EGX 30 BENCHMARK INDEX (EP)	Jan 1998	Mar 2018
Morocco	MCSHRPRCF	CFG 25 STOCK PRICE INDEX (EP)	Dec 1987	Mar 2018
Tunisia	TUSHRPRCF	TSE TUNINDEX	Dec 1997	Mar 2018
Panel B: Sub-Saharan Africa				
Kenya	KNSHRPRCF	NAIROBI S.E. INDEX (EP)	Jan 1990	Mar 2018
Malawi	MISHRPRCF	MALAWI STOCK EXCHANGE: ALL SHARE INDEX	Aug 2008	Feb 2018
Mauritius	MUSHRPRCF	MAURITIUS SE SEMDEX INDEX	Jul 1989	Mar 2018
Uganda	UGSHRPRCF	UGANDA SE ALL SHARE INDEX	Aug 2004	Mar 2018
Tanzania	TNSHRPRCF	DSE ALL SHARE INDEX	Dec 2006	Mar 2018
Zambia	ZMSHRPRCF	LUSAKA SE ALL SHARE INDEX	Jan 1997	Mar 2018
Namibia	WASHRPRCF	NSX LOCAL INDEX	Jul 2002	Jan 2017
South Africa	SASHRPRCF	DATASTREAM TOTAL MARKET STOCK PRICE INDEX (MONTHLY AVERAGE)	Jan 1973	Mar 2018
Ghana	GHSHRPRCF	GSE COMPOSITE INDEX	Jan 2011	Mar 2018
Nigeria	NGSHRPRCF	NIGERIAN S.E. - 30 STOCK PRICE INDEX (EP)	Dec 2009	Mar 2018
West African Economic and Monetary Union	BESHRPRCF	BRVM 10 INDEX	Sep 1998	Mar 2018
Panel C: Latin America and the Caribbean				
Mexico	MXSHRPRCF	SHARE PRICE INDEX OR IPC	Jan 1981	Mar 2018
Argentina	AGSHRPRCF	MERVAL STOCK MARKET INDEX	Jul 1993	Mar 2018
Brazil	BRSHRPRCF	BOVESPA SHARE PRICE INDEX (EP)	Jan 1982	Mar 2018
Chile	CLSHRPRCF	STOCK MARKET INDEX	Sep 1993	Mar 2018
Colombia	CBSHRPRCF	STOCK PRICE INDEX	Dec 1985	Mar 2018
Panel D: Northern America				
Canada	CNSHRPRCF	TORONTO STOCK EXCHANGE COMPOSITE SHARE PRICE INDEX (EP)	Dec 1964	Mar 2018
United States of America (the US)	USSHRPRCF	DOW JONES INDUSTRIALS SHARE PRICE INDEX (EP)	Jan 1950	Mar 2018
Panel E: Central Asia				
Kazakhstan	KZSHRPRCF	KASE SHARES INDEX (EP)	Jul 2000	Mar 2018
Panel F: Eastern Asia				
China (Mainland)	CHSHRPRCF	SHANGHAI SE COMPOSITE INDEX - CLOSE	Jan 1997	Mar 2018
Hong Kong	HKSHRPRCF	HANG SENG SHARE PRICE INDEX (EP)	Dec 1964	Mar 2018
Japan	JPSHRPRCF	TOKYO STOCK EXCHANGE - TOPIX (EP)	Jan 1957	Mar 2018
South Korea	KOSHRPRCF	KOSPI STOCK PRICE INDEX (EP)	Dec 1974	Mar 2018
Taiwan	TWSHRPRCF	TAIWAN STOCK EXCHANGE WEIGHTED TAIEX PRICE INDEX (EP)	Dec 1984	Mar 2018
Panel G: South-eastern Asia				
Indonesia	IDSHRPRCF	JAKARTA STOCK EXCHANGE COMPOSITE (EP)	Dec 1989	Mar 2018
Malaysia	MYSHRPRCF	FTSE BURSA MALAYSIA KLCI - PRICE CLOSE (EP)	Jan 1980	Mar 2018
Philippines	PHSHRPRCF	STOCK MARKET COMPOSITE INDEX - TOTAL (SUSP)	Jan 2005	Mar 2018
Thailand	THSHRPRCF	BANGKOK STOCK EXCHANGE PRICE INDEX (EP)	Apr 1975	Mar 2018

Panel H: Southern Asia				
Bangladesh	BSSHRPRCF	BANGLADESH ALL SHARE PRICE INDEX (EP)	Jan 1990	Mar 2018
India	INSHRPRCF	BOMBAY STOCK EXCHANGE NATIONAL 100 SHARE PRICE INDEX (EP)	Jan 1987	Mar 2018
Iran	IASHRPRCF	TEHERAN STOCK EXCHANGE PRICE INDEX (TEPIX)(1369SH=100)	Mar 2007	Feb 2018
Sri Lanka	LKSHRPRCF	COLOMBO ALL SHARE PRICE INDEX (EP)	Jan 1985	Mar 2018
Panel I: Western Asia				
Israel	ISSHRPRCF	TEL AVIV STOCK EXCHANGE GENERAL PRICE INDEX	Apr 1992	Mar 2018
Saudi Arabia	SISHRPRCF	STOCK PRICE INDEX	Jan 1998	Mar 2018
Turkey	TKSHRPRCF	ISE NATIONAL 100 SHARE PRICE INDEX	Jan 1988	Mar 2018
Panel J: Eastern Europe				
Czech Republic	CZSHRPRCF	PX-50 SHARE PRICE INDEX (EP)	Sep 1993	Mar 2018
Hungary	HNSHRPRCF	BUX SHARE PRICE INDEX (EP)	Jan 1991	Mar 2018
Poland	POSHRPRCF	WARSAW GENERAL SHARE PRICE INDEX (EP)	Apr 1991	Mar 2018
Russia	RSSHRPRCF	MICEX SHARE PRICE INDEX	Sep 1997	Mar 2018
Slovakia	SXSHRPRCF	SAX 12 SHARE PRICE INDEX (EP)	Aug 1993	Mar 2018
Ukraine	URSHRPRCF	PFTS INDEX (EP)	Oct 1997	Mar 2018
Panel K: Northern Europe				
Finland	FNSHRPRCF	HELSINKI STOCK EXCHANGE ALL SHARES PRICE INDEX (EP)	Dec 1957	Mar 2018
Iceland	ICSHRPRCF	SE ICEX ALL SHARE PRICE INDEX (EP)	Dec 1992	Mar 2018
Ireland	IRSHRPRCF	PRICE INDEX: ORDINARY STOCKS & SHARES - FIRST WORKING DAY	Jan 1958	Mar 2018
Lithuania	LNSHRPRCF	LITHUANIAN STOCK PRICE INDEX (EP)	Dec 1999	Mar 2018
Norway	NWSHRPRCF	OSLO STOCK EXCHANGE BENCHMARK INDEX	Dec 1995	Mar 2018
United Kingdom	UKSHRPRCF	FT ALL SHARE INDEX (EP)	Apr 1962	Mar 2018
Panel L: Southern Europe				
Croatia	CTSHRPRCF	STOCK EXCHANGE SHARE INDEX - CROBEX	Jan 1997	Mar 2018
Greece	GRSHRPRCF	ATHENS STOCK EXCHANGE GENERAL SHARE PRICE INDEX (EP)	Jan 1985	Mar 2018
Italy	ITSHRPRCF	MILAN COMIT GENERAL SHARE PRICE INDEX (EP)	Jan 1969	Mar 2018
Portugal	PTSHRPRCF	PSI GENERAL STOCK PRICE INDEX (EP)	Jan 1988	Mar 2018
Spain	ESSHRPRCF	MADRID S.E - GENERAL INDEX	Jan 1958	Mar 2018
Panel M: Western Europe				
Belgium	BGSHRPRCF	BRUSSELS STOCK EXCHANGE CASH MARKET RETURN INDEX (EP)	Dec 1979	Mar 2018
France	FRSHRPRCF	SHARE PRICE INDEX - SBF 250	Jan 1958	Mar 2018
Germany	BDSHRPRCF	DAX SHARE PRICE INDEX, EP	Sep 1959	Mar 2018
Netherlands	NLSHRPRCF	AMSTERDAM SE ALL SHARE STOCK PRICE INDEX (EP)	Dec 1964	Mar 2018
Switzerland	SWSHRPRCF	SPI SHARE PRICE INDEX (EP)	Sep 1987	Sep 2017
Panel N: Europe				
Euro Zone	EMSHRPRCF	DATASTREAM EURO SHARE PRICE INDEX (MONTHLY AVERAGE)	Jan 1973	Mar 2018
Panel O: Australia and New Zealand				
Australia	AUSHRPRCF	S&P/ASX 200 (METHODOLOGY BREAK MARCH 2000)	Feb 1971	Mar 2018
New Zealand	NZSHRPRCF	NEW ZEALAND STOCK EXCHANGE ALL SHARE PRICE INDEX (EP)	Jun 1986	Mar 2018

Source: Datastream (2018)

Table C.1 Stock Market Indexes for each Country/Region

Global Shock Variable	Data Period (From/To)	
World oil prices	Jan 1986	Mar 2018
Oil supply shocks	Jan 1973	Mar 2018
Oil demand shocks	May 1985	Mar 2018
World commodity prices (all items)	Jan 1951	Mar 2018
World metal prices	Jan 1951	Mar 2018

Table C.2 Global Shock Variables

C.2 Unit root tests and other preliminary statistics

Panel A: Stock Indexes	Obs.	Mean	Median	Minimum	Maximum	Std. Dev.	Skewness	Kurtosis	p-KS	p-AD	p-JB
Argentina	296	0.004	0.016	-0.537	0.379	0.113	-0.966	6.571	0.000	0.001	0.001
Australia	565	0.004	0.007	-0.604	0.195	0.069	-1.735	14.851	0.000	0.001	0.001
Bangladesh	335	0.003	0.001	-0.364	0.564	0.092	0.582	9.929	0.000	0.001	0.001
Belgium	459	0.008	0.009	-0.321	0.199	0.057	-0.901	8.048	0.000	0.001	0.001
Brazil	434	0.009	0.013	-1.120	0.657	0.171	-0.838	9.157	0.000	0.001	0.001
Canada	639	0.004	0.008	-0.320	0.187	0.054	-0.994	7.146	0.000	0.001	0.001
Chile	294	0.006	0.002	-0.269	0.174	0.059	-0.453	5.220	0.000	0.015	0.001
China (Mainland)	254	0.006	0.007	-0.281	0.278	0.079	-0.300	4.780	0.000	0.001	0.001
Colombia	387	0.009	0.002	-0.366	0.413	0.086	0.086	5.469	0.000	0.001	0.001
Croatia	254	0.002	0.004	-0.531	0.367	0.088	-1.291	10.872	0.000	0.001	0.001
Czech Republic	294	0.005	0.011	-0.402	0.451	0.088	-0.071	7.223	0.000	0.001	0.001
Egypt	240	0.006	0.010	-0.424	0.351	0.097	-0.429	5.315	0.000	0.014	0.001
Euro Zone	542	0.005	0.005	-0.337	0.185	0.062	-0.409	4.951	0.000	0.010	0.001
Finland	723	0.007	0.004	-0.324	0.253	0.061	-0.237	6.119	0.000	0.001	0.001
France	722	0.004	0.005	-0.286	0.217	0.057	-0.556	5.874	0.000	0.001	0.001
Germany	702	0.006	0.009	-0.286	0.216	0.061	-0.534	4.916	0.000	0.001	0.001
Ghana	86	0.001	-0.001	-0.188	0.185	0.062	0.341	4.245	0.000	0.036	0.030
Greece	396	0.009	0.009	-0.429	0.419	0.105	0.125	5.172	0.000	0.001	0.001
Hong Kong	639	0.008	0.011	-0.576	0.608	0.094	-0.599	10.524	0.000	0.001	0.001
Hungary	326	0.007	0.012	-0.494	0.448	0.097	-0.711	7.940	0.000	0.001	0.001
Iceland	303	0.003	0.014	-1.388	0.210	0.105	-8.000	102.718	0.000	0.001	0.001
India	364	0.008	0.008	-0.352	0.350	0.088	-0.115	4.440	0.000	0.032	0.001
Indonesia	339	0.002	0.006	-0.523	0.431	0.108	-0.882	8.025	0.000	0.001	0.001
Iran	131	0.009	0.000	-0.796	0.201	0.097	-4.196	36.979	0.000	0.001	0.001
Ireland	722	0.007	0.011	-0.311	0.264	0.059	-0.457	6.755	0.000	0.001	0.001
Israel	311	0.006	0.011	-0.235	0.278	0.066	-0.489	5.050	0.000	0.001	0.001
Italy	590	0.004	0.003	-0.289	0.239	0.070	-0.215	3.881	0.000	0.014	0.001
Japan	734	0.006	0.004	-0.310	0.229	0.056	-0.218	5.098	0.000	0.001	0.001
Kazakhstan	212	0.011	0.014	-0.464	0.452	0.104	-0.311	7.855	0.000	0.001	0.001
Kenya	338	0.000	0.003	-0.291	0.426	0.074	0.498	8.130	0.000	0.001	0.001
Lithuania	219	0.010	0.008	-0.453	0.373	0.077	-0.985	12.009	0.000	0.001	0.001
Malawi	114	-0.002	0.002	-0.458	0.290	0.072	-1.532	18.607	0.000	0.001	0.001
Malaysia	458	0.003	0.010	-0.416	0.398	0.081	-0.459	7.489	0.000	0.001	0.001
Mauritius	344	0.006	0.003	-0.297	0.179	0.052	-0.475	8.070	0.000	0.001	0.001
Mexico	446	0.008	0.017	-0.935	0.767	0.133	-2.063	19.245	0.000	0.001	0.001
Morocco	363	0.009	0.011	-0.160	0.206	0.047	-0.012	4.625	0.000	0.022	0.001

Namibia	174	0.012	0.015	-0.129	0.162	0.052	-0.103	3.095	0.000	0.563	0.500
Netherlands	639	0.004	0.008	-0.307	0.194	0.054	-0.927	6.462	0.000	0.001	0.001
New Zealand	381	0.002	0.008	-0.434	0.237	0.065	-0.895	8.430	0.000	0.001	0.001
Nigeria	99	-0.001	0.013	-0.277	0.134	0.075	-0.829	4.153	0.000	0.026	0.005
Norway	267	0.007	0.016	-0.450	0.204	0.067	-1.554	10.852	0.000	0.001	0.001
Philippines	158	0.009	0.014	-0.287	0.163	0.059	-0.907	6.378	0.000	0.017	0.001
Poland	323	0.009	0.012	-0.433	0.700	0.115	0.379	8.414	0.000	0.001	0.001
Portugal	362	0.003	0.008	-0.337	0.215	0.064	-0.611	5.317	0.000	0.001	0.001
Russia	246	0.003	0.013	-1.043	0.341	0.135	-2.239	17.540	0.000	0.001	0.001
Saudi Arabia	242	0.006	0.011	-0.302	0.179	0.071	-0.795	5.098	0.000	0.001	0.001
Slovakia	295	0.004	0.005	-0.360	0.775	0.089	2.479	25.908	0.000	0.001	0.001
South Africa	542	0.006	0.011	-0.359	0.201	0.073	-0.592	4.801	0.000	0.002	0.001
South Korea	519	0.005	0.000	-0.464	0.520	0.089	-0.177	8.028	0.000	0.001	0.001
Spain	722	0.005	0.006	-0.325	0.247	0.060	-0.274	5.529	0.000	0.001	0.001
Sri Lanka	398	0.006	0.004	-0.192	0.301	0.072	0.359	4.217	0.000	0.001	0.001
Switzerland	360	0.007	0.012	-0.272	0.176	0.051	-0.700	6.057	0.000	0.001	0.001
Taiwan	399	0.007	0.008	-0.483	0.437	0.103	-0.143	6.469	0.000	0.001	0.001
Tanzania	135	0.002	0.000	-0.137	0.144	0.041	0.008	4.852	0.000	0.001	0.003
Thailand	515	0.004	0.006	-0.395	0.340	0.088	-0.441	5.995	0.000	0.001	0.001
Tunisia	243	0.005	0.001	-0.181	0.155	0.042	-0.007	5.035	0.000	0.003	0.001
Turkey	362	0.004	0.011	-0.549	0.541	0.151	-0.032	4.710	0.000	0.001	0.001
Uganda	163	0.006	0.015	-0.452	0.208	0.077	-1.494	10.405	0.000	0.001	0.001
Ukraine	245	-0.006	0.002	-0.656	0.415	0.132	-1.101	7.948	0.000	0.001	0.001
United Kingdom	671	0.004	0.007	-0.252	0.435	0.059	0.101	8.613	0.000	0.001	0.001
United States of America (the US)	807	0.006	0.008	-0.264	0.135	0.041	-0.690	5.966	0.000	0.001	0.001
West African Economic and Monetary Union	234	0.003	0.000	-0.238	0.231	0.067	-0.116	4.128	0.000	0.001	0.008
Zambia	254	0.008	0.007	-0.304	0.306	0.088	0.275	4.938	0.000	0.001	0.001
Panel B: Global Shock Variables											
World oil prices	1695	0.000	0.000	-0.021	0.019	0.005	-0.354	5.110	0.000	0.001	0.001
Oil supply shock	2387	0.000	0.000	-0.021	0.010	0.003	-0.792	7.890	0.000	0.001	0.001
Oil demand shock	1737	0.000	0.000	-0.061	0.031	0.008	-1.240	12.567	0.000	0.001	0.001
World commodity prices (all items)	3548	0.000	0.000	-0.010	0.005	0.001	-0.691	11.758	0.000	0.001	0.001
World metal prices	3548	0.000	0.000	-0.021	0.009	0.002	-1.148	17.114	0.000	0.001	0.001

Note: The table reports the primary statistics obtained for monthly stock price returns and weekly global shock variables. All series are log-differenced, as specified in Section 4.5. "p-KS" signifies a p-value of the Kolmogorov–Smirnov test for normality. "p-AD" signifies a p-value of the Anderson–Darling test for normality. "p-JB" signifies a p-value of the Jarque–Bera test for normality.

Table C.3 Sample Statistics of Differenced Series

	Stock Market Indexes and Global Shock Variables: Low-Frequency			
	ADF with intercept, no trend	ADF with no intercept, no trend	PP with intercept, no trend	PP with no intercept, no trend
Panel A: Stock Indexes				
Argentina	-15.582***	-15.609***	-15.584***	-15.611***
Australia	-22.954***	-22.975***	-22.959***	-22.980***
Bangladesh	-15.383***	-15.406***	-15.383***	-15.406***
Belgium	-18.891***	-18.911***	-19.048***	-19.068***
Brazil	-21.249***	-21.274***	-21.349***	-21.376***
Canada	-22.991***	-23.009***	-22.991***	-23.009***
Chile	-13.658***	-13.682***	-13.622***	-13.645***
China (Mainland)	-14.178***	-14.206***	-14.381***	-14.407***
Colombia	-14.568***	-14.587***	-14.551***	-14.570***
Croatia	-14.750***	-14.779***	-14.813***	-14.841***
Czech Republic	-13.697***	-13.720***	-13.599***	-13.624***
Egypt	-13.272***	-13.300***	-13.656***	-13.680***
Euro Zone	-17.729***	-17.746***	-17.904***	-17.920***
Finland	-21.858***	-21.873***	-22.671***	-22.684***
France	-23.793***	-23.809***	-23.895***	-23.911***
Germany	-25.641***	-25.659***	-25.658***	-25.676***
Ghana	-6.717***	-6.757***	-6.718***	-6.758***
Greece	-16.615***	-16.636***	-16.656***	-16.677***
Hong Kong	-23.523***	-23.542***	-23.498***	-23.516***
Hungary	-16.384***	-16.409***	-16.362***	-16.388***
Iceland	-11.858***	-11.878***	-12.874***	-12.890***
India	-17.150***	-17.173***	-17.151***	-17.174***
Indonesia	-14.082***	-14.103***	-14.087***	-14.108***
Iran	-9.054***	-9.089***	-9.053***	-9.089***
Ireland	-22.647***	-22.663***	-23.289***	-23.303***
Israel	-16.042***	-16.068***	-16.034***	-16.060***
Italy	-22.005***	-22.024***	-22.259***	-22.276***
Japan	-24.157***	-24.174***	-24.353***	-24.369***

Kazakhstan	-5.837***	-5.851***	-8.822***	-8.843***
Kenya	-14.686***	-14.707***	-15.102***	-15.122***
Lithuania	-10.690***	-10.714***	-11.058***	-11.081***
Malawi	-8.891***	-8.931***	-9.146***	-9.182***
Malaysia	-12.167***	-12.181***	-19.289***	-19.308***
Mauritius	-10.040***	-10.055***	-14.833***	-14.851***
Mexico	-19.311***	-19.333***	-19.281***	-19.303***
Morocco	-17.884***	-17.909***	-18.395***	-18.416***
Namibia	-14.097***	-14.137***	-14.136***	-14.177***
Netherlands	-23.193***	-23.211***	-23.216***	-23.234***
New Zealand	-18.232***	-18.256***	-18.237***	-18.261***
Nigeria	-8.432***	-8.476***	-8.432***	-8.476***
Norway	-11.534***	-11.555***	-11.534***	-11.555***
Philippines	-10.654***	-10.688***	-10.830***	-10.862***
Poland	-16.318***	-16.343***	-16.450***	-16.474***
Portugal	-16.630***	-16.653***	-16.697***	-16.719***
Russia	-12.925***	-12.951***	-12.975***	-13.002***
Saudi Arabia	-12.786***	-12.813***	-12.981***	-13.007***
Slovakia	-13.378***	-13.401***	-13.143***	-13.168***
South Africa	-18.579***	-18.596***	-18.326***	-18.345***
South Korea	-21.321***	-21.342***	-21.406***	-21.425***
Spain	-24.443***	-24.460***	-25.044***	-25.059***
Sri Lanka	-16.482***	-16.503***	-16.931***	-16.950***
Switzerland	-18.914***	-18.938***	-18.944***	-18.969***
Taiwan	-17.984***	-18.007***	-18.040***	-18.062***
Tanzania	-13.760***	-13.813***	-13.555***	-13.599***
Thailand	-20.049***	-20.068***	-20.033***	-20.053***
Tunisia	-14.316***	-14.346***	-14.293***	-14.324***
Turkey	-17.538***	-17.562***	-17.526***	-17.551***
Uganda	-12.356***	-12.395***	-12.433***	-12.469***
Ukraine	-10.800***	-10.823***	-11.091***	-11.112***

United Kingdom	-23.625***	-23.642***	-23.609***	-23.627***
United States of America (the US)	-27.566***	-27.583***	-27.558***	-27.575***
West African Economic and Monetary Union	-14.223***	-14.254***	-14.257***	-14.287***
Zambia	-6.714***	-6.727***	-14.864***	-14.889***
Panel B: Global Shock Variables				
World oil prices	-18.413***	-18.437***	-18.41***	-18.436***
Oil supply shock	-21.711***	-21.731***	-21.711***	-21.731***
Oil demand shock	-17.548***	-17.57***	-17.781***	-17.811***
World commodity prices (all items)	-21.364***	-21.377***	-21.812***	-21.825***
World metal prices	-21.630***	-21.643***	-21.953***	-21.966***

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table C.4 Unit Root Tests, Low-frequency

	Stock Market Indexes: Mixed-Frequency							
	ADF with intercept, no trend				PP with intercept, no trend			
	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,4)	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,4)
Argentina	-18.532***	-16.407***	-14.286***	-17.216***	-18.681***	-16.650***	-14.392***	-17.314***
Australia	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Bangladesh	-18.729***	-17.145***	-15.460***	-18.737***	-19.022***	-17.411***	-15.490***	-18.744***
Belgium	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Brazil	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Canada	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Chile	-18.471***	-16.243***	-14.231***	-17.159***	-18.616***	-16.482***	-14.336***	-17.255***
China (Mainland)	-17.146***	-15.152***	-13.396***	-16.195***	-17.160***	-15.180***	-13.419***	-16.294***
Colombia	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Croatia	-17.146***	-15.152***	-13.396***	-16.195***	-17.160***	-15.180***	-13.419***	-16.294***
Czech Republic	-18.471***	-16.243***	-14.231***	-17.159***	-18.616***	-16.482***	-14.336***	-17.255***
Egypt	-16.646***	-14.735***	-12.931***	-15.964***	-16.647***	-14.738***	-12.952***	-16.073***
Euro Zone	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Finland	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
France	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Germany	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Ghana	-10.048***	-7.924***	-9.699***	-9.228***	-10.052***	-7.857***	-9.816***	-9.237***
Greece	-20.184***	-18.885***	-16.875***	-20.437***	-20.508***	-19.155***	-16.894***	-20.441***
Hong Kong	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Hungary	-19.634***	-17.179***	-15.182***	-18.304***	-19.820***	-17.420***	-15.301***	-18.316***
Iceland	-18.687***	-16.502***	-14.442***	-17.381***	-18.843***	-16.753***	-14.552***	-17.470***
India	-19.516***	-18.050***	-16.108***	-19.654***	-19.774***	-18.316***	-16.126***	-19.648***
Indonesia	-18.955***	-17.210***	-15.552***	-18.854***	-19.239***	-17.513***	-15.576***	-18.862***
Iran	-11.923***	-10.062***	-8.596***	-9.111***	-11.911***	-9.960***	-8.607***	-9.091***
Ireland	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Israel	-18.851***	-16.716***	-14.636***	-17.540***	-18.998***	-16.965***	-14.746***	-17.610***
Italy	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Japan	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Kazakhstan	-15.090***	-14.086***	-12.178***	-13.907***	-15.095***	-14.079***	-12.186***	-14.231***
Kenya	-18.918***	-17.205***	-15.541***	-18.916***	-19.194***	-17.502***	-15.566***	-18.917***
Lithuania	-15.630***	-14.339***	-12.682***	-14.215***	-15.636***	-14.337***	-12.699***	-14.497***
Malawi	-11.332***	-9.453***	-7.674***	-8.262***	-11.342***	-9.387***	-7.729***	-8.242***
Malaysia	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Mauritius	-19.033***	-17.295***	-15.667***	-18.984***	-19.271***	-17.579***	-15.690***	-18.999***
Mexico	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Morocco	-19.411***	-17.955***	-15.852***	-19.297***	-19.584***	-18.237***	-15.883***	-19.356***

Namibia	-13.790***	-11.845***	-10.887***	-12.287***	-13.775***	-11.821***	-10.887***	-12.507***
Netherlands	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
New Zealand	-20.375***	-18.893***	-16.463***	-20.061***	-20.572***	-19.226***	-16.476***	-20.066***
Nigeria	-10.408***	-9.122***	-11.010***	-9.206***	-10.408***	-9.099***	-11.048***	-9.230***
Norway	-17.633***	-15.704***	-13.726***	-16.745***	-17.638***	-15.727***	-13.725***	-16.843***
Philippines	-13.043***	-11.456***	-9.886***	-10.946***	-13.042***	-11.433***	-9.889***	-11.005***
Poland	-19.219***	-17.073***	-15.046***	-17.920***	-19.380***	-17.314***	-15.162***	-17.987***
Portugal	-19.391***	-17.945***	-15.829***	-19.313***	-19.553***	-18.235***	-15.861***	-19.366***
Russia	-16.795***	-14.955***	-13.134***	-15.954***	-16.779***	-14.980***	-13.150***	-16.078***
Saudi Arabia	-16.653***	-14.877***	-13.015***	-15.914***	-16.633***	-14.879***	-13.034***	-16.040***
Slovakia	-18.569***	-16.324***	-14.264***	-17.192***	-18.695***	-16.550***	-14.374***	-17.295***
South Africa	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
South Korea	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Spain	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Sri Lanka	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Switzerland	-19.334***	-17.729***	-15.830***	-19.380***	-19.540***	-18.020***	-15.863***	-19.395***
Taiwan	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Tanzania	-12.298***	-10.536***	-8.783***	-9.246***	-12.273***	-10.494***	-8.795***	-9.224***
Thailand	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
Tunisia	-16.679***	-14.874***	-13.044***	-15.928***	-16.660***	-14.894***	-13.062***	-16.048***
Turkey	-19.391***	-17.945***	-15.829***	-19.313***	-19.553***	-18.235***	-15.861***	-19.366***
Uganda	-13.292***	-11.801***	-10.019***	-11.193***	-13.280***	-11.828***	-10.026***	-11.294***
Ukraine	-16.667***	-14.901***	-13.161***	-7.110***	-16.655***	-14.928***	-13.183***	-16.091***
United Kingdom	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
United States of America (the US)	-20.169***	-18.878***	-16.897***	-20.520***	-20.392***	-19.249***	-16.917***	-20.548***
West African Economic and Monetary Union	-16.203***	-15.293***	-12.782***	-14.822***	-16.192***	-15.293***	-12.812***	-15.099***
Zambia	-17.146***	-15.152***	-13.396***	-16.195***	-17.160***	-15.180***	-13.419***	-16.294***

Note: The table reports the test statistics obtained from the unit root tests for monthly commodity price series at a given month of each quarter period τ.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table C.5 Unit Root Tests with an Intercept and no Trend, Mixed-frequency

	Stock Market Indexes: Mixed-Frequency							
	ADF with no intercept, no trend				PP with no intercept, no trend			
	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,4)	CP(τ,1)	CP(τ,2)	CP(τ,3)	CP(τ,4)
Argentina	-18.564***	-16.435***	-14.311***	-17.245***	-18.715***	-16.674***	-14.416***	-17.340***
Australia	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Bangladesh	-18.757***	-17.170***	-15.483***	-18.765***	-19.055***	-17.434***	-15.513***	-18.770***
Belgium	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Brazil	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Canada	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Chile	-18.503***	-16.271***	-14.255***	-17.189***	-18.650***	-16.507***	-14.36***	-17.281***
China (Mainland)	-17.179***	-15.182***	-13.422***	-16.227***	-17.194***	-15.209***	-13.445***	-16.323***
Colombia	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Croatia	-17.179***	-15.182***	-13.422***	-16.227***	-17.194***	-15.209***	-13.445***	-16.323***
Czech Republic	-18.503***	-16.271***	-14.255***	-17.189***	-18.650***	-16.507***	-14.360***	-17.281***
Egypt	-16.681***	-14.766***	-12.958***	-15.998***	-16.682***	-14.769***	-12.979***	-16.102***
Euro Zone	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Finland	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
France	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Germany	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Ghana	-10.109***	-7.971***	-9.757***	-9.273***	-10.113***	-7.909***	-9.881***	-9.280***
Greece	-20.210***	-18.910***	-16.897***	-20.463***	-20.540***	-19.177***	-16.916***	-20.465***
Hong Kong	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Hungary	-19.663***	-17.205***	-15.206***	-18.332***	-19.851***	-17.443***	-15.324***	-18.341***
Iceland	-18.718***	-16.530***	-14.466***	-17.409***	-18.876***	-16.777***	-14.575***	-17.496***
India	-19.543***	-18.075***	-16.130***	-19.681***	-19.805***	-18.338***	-16.148***	-19.674***
Indonesia	-18.984***	-17.235***	-15.575***	-18.882***	-19.272***	-17.535***	-15.599***	-18.889***
Iran	-11.970***	-10.103***	-8.629***	-9.146***	-11.956***	-10.007***	-8.641***	-9.127***
Ireland	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Israel	-18.881***	-16.743***	-14.660***	-17.569***	-19.031***	-16.989***	-14.769***	-17.636***
Italy	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Japan	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Kazakhstan	-15.127***	-14.119***	-12.207***	-13.940***	-15.131***	-14.114***	-12.215***	-14.258***
Kenya	-18.947***	-17.231***	-15.564***	-18.944***	-19.227***	-17.525***	-15.589***	-18.944***
Lithuania	-15.666***	-14.371***	-12.711***	-14.248***	-15.672***	-14.370***	-12.728***	-14.525***
Malawi	-11.382***	-9.496***	-7.708***	-8.299***	-11.392***	-9.433***	-7.763***	-8.279***
Malaysia	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Mauritius	-19.061***	-17.320***	-15.690***	-19.012***	-19.302***	-17.601***	-15.713***	-19.025***
Mexico	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Morocco	-19.438***	-17.980***	-15.874***	-19.323***	-19.614***	-18.259***	-15.904***	-19.380***

Namibia	-13.831***	-11.880***	-10.918***	-12.323***	-13.814***	-11.856***	-10.918***	-12.538***
Netherlands	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
New Zealand	-20.401***	-18.918***	-16.485***	-20.087***	-20.600***	-19.247***	-16.498***	-20.089***
Nigeria	-10.463***	-9.169***	-11.066***	-9.254***	-10.463***	-9.149***	-11.107***	-9.276***
Norway	-17.666***	-15.734***	-13.752***	-16.777***	-17.671***	-15.755***	-13.751***	-16.870***
Philippines	-13.085***	-11.493***	-9.918***	-10.980***	-13.084***	-11.476***	-9.921***	-11.038***
Poland	-19.249***	-17.100***	-15.069***	-17.948***	-19.413***	-17.338***	-15.184***	-18.013***
Portugal	-19.418***	-17.970***	-15.851***	-19.340***	-19.584***	-18.257***	-15.882***	-19.390***
Russia	-16.829***	-14.986***	-13.161***	-15.986***	-16.813***	-15.010***	-13.177***	-16.107***
Saudi Arabia	-16.687***	-14.908***	-13.042***	-15.948***	-16.667***	-14.910***	-13.061***	-16.068***
Slovakia	-18.601***	-16.352***	-14.288***	-17.222***	-18.728***	-16.575***	-14.397***	-17.321***
South Africa	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
South Korea	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Spain	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Sri Lanka	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Switzerland	-19.361***	-17.754***	-15.853***	-19.407***	-19.571***	-18.042***	-15.885***	-19.420***
Taiwan	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Tanzania	-12.343***	-10.576***	-8.816***	-9.281***	-12.316***	-10.538***	-8.828***	-9.259***
Thailand	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
Tunisia	-16.713***	-14.905***	-13.071***	-15.961***	-16.694***	-14.924***	-13.089***	-16.077***
Turkey	-19.418***	-17.970***	-15.851***	-19.340***	-19.584***	-18.257***	-15.882***	-19.390***
Uganda	-13.334***	-11.838***	-10.050***	-11.228***	-13.319***	-11.872***	-10.057***	-11.327***
Ukraine	-16.701***	-14.932***	-13.188***	-7.125***	-16.689***	-14.958***	-13.210***	-16.119***
United Kingdom	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
United States of America (the US)	-20.195***	-18.903***	-16.919***	-20.547***	-20.422***	-19.269***	-16.939***	-20.572***
West African Economic and Monetary Union	-16.238***	-15.326***	-12.809***	-14.854***	-16.227***	-15.326***	-12.839***	-15.126***
Zambia	-17.179***	-15.182***	-13.422***	-16.227***	-17.194***	-15.209***	-13.445***	-16.323***

Note: The table reports the test statistics obtained from the unit root tests for monthly commodity price series at a given month of each quarter period τ.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table C.6 Unit Root Tests with no Intercept and no Trend, Mixed-frequency

C.3 National commodity export prices and stock market returns

	National Commodity Export Prices			
	ADF with intercept, no trend	**ADF with no intercept, no trend**	**PP with intercept, no trend**	**PP with no intercept, no trend**
Argentina	-14.513***	-14.524***	-14.557***	-14.569***
Australia	-13.373***	-13.357***	-13.853***	-13.854***
Bangladesh	-16.656***	-16.673***	-16.637***	-16.654***
Belgium	-16.861***	-16.863***	-17.046***	-17.052***
Benin	-12.604***	-12.618***	-12.642***	-12.656***
Brazil	-14.745***	-14.736***	-14.898***	-14.895***
Burkina Faso	-13.155***	-13.169***	-13.155***	-13.169***
Canada	-16.035***	-16.042***	-16.076***	-16.083***
Chile	-14.365***	-14.375***	-14.297***	-14.309***
China (Mainland)	-9.899***	-9.896***	-13.686***	-13.695***
Colombia	-14.714***	-14.726***	-14.728***	-14.740***
Croatia	-16.546***	-16.554***	-16.561***	-16.608***
Czech Republic	-10.196***	-10.175***	-16.939***	-16.943***
Egypt	-13.499***	-13.510***	-13.722***	-13.733***
Finland	-12.838***	-12.799***	-22.376***	-22.375***
France	-14.861***	-14.863***	-15.426***	-15.434***
Germany	-14.217***	-14.218***	-14.743***	-14.754***
Ghana	-15.260***	-15.267***	-17.010***	-17.041***
Greece	-14.048***	-14.046***	-14.083***	-14.082***
Guinea-Bissau	-15.347***	-15.362***	-14.758***	-14.777***
Hong Kong	-16.878***	-16.876***	-16.905***	-16.915***
Hungary	-15.901***	-15.904***	-15.887***	-15.890***
Iceland	-16.893***	-16.912***	-17.354***	-17.371***
India	-13.705***	-13.697***	-13.710***	-13.672***
Indonesia	-12.847***	-12.855***	-13.215***	-13.225***
Iran	-15.174***	-15.190***	-14.534***	-14.553***
Ireland	-15.024***	-15.008***	-15.000***	-15.122***
Israel	-11.955***	-11.962***	-12.035***	-12.041***
Italy	-10.509***	-10.481***	-14.974***	-14.978***
Ivory Coast / Cote d'Ivoire	-15.447***	-15.464***	-15.293***	-15.311***
Japan	-15.631***	-15.637***	-15.795***	-15.809***
Kazakhstan	-14.907***	-14.921***	-14.466***	-14.483***
Kenya	-19.859***	-19.870***	-19.890***	-19.902***
Lithuania	-9.167***	-9.157***	-14.268***	-14.268***
Malawi	-15.244***	-15.191***	-15.480***	-15.477***
Malaysia	-13.790***	-13.801***	-13.911***	-13.925***
Mali	-16.904***	-16.911***	-17.085***	-17.099***
Mauritius	-16.560***	-16.578***	-16.542***	-16.560***
Mexico	-15.040***	-15.055***	-14.464***	-14.482***
Morocco	-9.949***	-9.950***	-13.895***	-13.904***
Namibia	-15.284***	-15.284***	-15.548***	-15.555***
Netherlands	-13.525***	-13.528***	-13.570***	-13.573***
New Zealand	-9.756***	-9.736***	-14.266***	-14.262***
Niger	-15.075***	-15.089***	-14.694***	-14.710***
Nigeria	-15.164***	-15.180***	-14.526***	-14.546***
Norway	-15.262***	-15.276***	-14.923***	-14.940***
Philippines	-16.125***	-16.126***	-16.239***	-16.244***
Poland	-13.350***	-13.348***	-13.720***	-13.727***
Portugal	-9.370***	-9.362***	-13.745***	-13.749***
Russia	-15.364***	-15.377***	-15.284***	-15.299***
Saudi Arabia	-15.167***	-15.183***	-14.526***	-14.546***
Senegal	-12.117***	-12.120***	-12.499***	-12.504***
Slovakia	-10.432***	-10.426***	-16.903***	-16.918***
South Africa	-14.828***	-14.834***	-15.242***	-15.255***
South Korea	-17.022***	-17.026***	-17.075***	-17.088***
Spain	-15.944***	-15.949***	-15.465***	-15.444***
Sri Lanka	-20.564***	-20.568***	-20.576***	-20.568***

Switzerland	-15.777***	-15.793***	-15.903***	-15.920***
Tanzania	-17.199***	-17.193***	-17.425***	-17.434***
Thailand	-14.288***	-14.303***	-14.628***	-14.643***
Togo	-7.687***	-7.693***	-15.078***	-15.091***
Tunisia	-13.641***	-13.652***	-13.661***	-13.673***
Turkey	-17.113***	-17.085***	-17.071***	-17.082***
Uganda	-15.852***	-15.870***	-15.851***	-15.869***
Ukraine	-13.653***	-13.651***	-13.979***	-13.982***
United Kingdom	-14.713***	-14.727***	-14.499***	-14.514***
United States of America (the US)	-14.122***	-14.124***	-14.513***	-14.520***
Zambia	-14.717***	-14.729***	-14.797***	-14.809***

Note: The table reports the test statistics obtained from the unit root tests.
, ** and * denote the statistical significance at the 10%, 5% and 1% level respectively.*

Table C.7 Unit Root Tests, National Commodity Export Prices

	National Commodity Export Prices	
	SP ⇏ CP	CP ⇏ SP
	Africa	
Panel A: Northern Africa		
Egypt	0.078	0.520
Morocco	0.030	0.044
Tunisia	0.010	0.400
Panel B: Sub-Saharan Africa		
Kenya	0.294	0.876
Malawi	0.034	0.912
Mauritius	0.630	0.614
Uganda	0.576	0.868
Tanzania	0.140	0.888
Zambia	0.174	0.002
Namibia	0.196	0.788
South Africa	0.018	0.478
Ghana	0.452	0.652
Nigeria	0.442	0.580
West African Economic and Monetary Union		
Benin	0.012	0.010
Burkina Faso	0.074	0.058
Guinea-Bissau	0.008	0.146
Ivory Coast / Cote d'Ivoire	0.094	0.134
Mali	0.076	0.338
Niger	0.084	0.100
Senegal	0.060	0.006
Togo	0.052	0.048
	Americas	
Panel C: Latin America and the Caribbean		
Mexico	0.690	0.286
Argentina	0.074	0.116
Brazil	0.002	0.832
Chile	0.006	0.066
Colombia	0.226	0.942
Panel D: Northern America		
Canada	0.024	0.460
United States of America (the US)	0.152	0.918
	Asia	
Panel E: Central Asia		
Kazakhstan	0.002	0.398
Panel F: Eastern Asia		
China (Mainland)	0.650	0.808
Hong Kong	0.102	0.658
Japan	0.004	0.606
South Korea	0.034	0.664
Panel G: South-eastern Asia		
Indonesia	0.056	0.558
Malaysia	0.186	0.730
Philippines	0.010	0.940
Thailand	0.006	0.694
Panel H: Southern Asia		
Bangladesh	0.044	0.734
India	0.002	0.336
Iran	0.040	0.226
Sri Lanka	0.814	0.378
Panel I: Western Asia		
Israel	0.186	0.784
Saudi Arabia	0.026	0.110

Turkey	0.034	0.260
	Europe	
Panel J: Eastern Europe		
Czech Republic	0.096	0.346
Hungary	0.020	0.472
Poland	0.066	0.870
Russia	0.174	0.418
Slovakia	0.162	0.274
Ukraine	0.028	0.634
Panel K: Northern Europe		
Finland	0.002	0.958
Iceland	0.052	0.184
Ireland	0.078	0.676
Lithuania	0.056	0.374
Norway	0.032	0.682
United Kingdom	0.002	0.920
Panel L: Southern Europe		
Croatia	0.064	0.532
Greece	0.006	0.468
Italy	0.008	0.480
Portugal	0.014	0.412
Spain	0.258	0.774
Panel M: Western Europe		
Belgium	0.034	0.426
France	0.074	0.180
Germany	0.002	0.320
Netherlands	0.002	0.944
Switzerland	0.006	0.044
	Oceania	
Panel N: Australia and New Zealand		
Australia	0.012	0.948
New Zealand	0.004	0.478

Note: The table contains bootstrapped p-values for the full sample LF Granger causality tests. The LF approach uses monthly measures of national commodity export prices and monthly stock market returns. "SP" denotes stock market returns, while "CP" denotes national commodity export prices. $H_0: SP \nRightarrow CP$ ($\nRightarrow$ means "does not Granger-cause"). We follow Ghysels et al. (2016) and use bootstrapped p-values with N = 499 replications (Gonçalves and Kilian, 2004). All variables are mean-centred and log-differenced.

Table C.8 P-values for Full Sample Tests of Non-Causality, National Commodity Export Prices

	SP ⇏ CP		CP ⇏ SP	
	Significance Level		Significance Level	
	5%	10%	5%	10%
		Africa		
Panel A: Northern Africa				
Egypt	0.681	0.862	0.052	0.103
Morocco	0.695	0.944	0.260	0.520
Tunisia	0.282	0.658	0	0.009
Panel B: Sub-Saharan Africa				
Kenya	0	0	0.200	0.382
Malawi	0.849	0.962	0	0
Mauritius	0	0	0	0.012
Uganda	0	0	0	0
Tanzania	0.317	0.333	0.175	0.286
Zambia	0.447	0.699	0.789	0.951
Namibia	0.023	0.125	0.011	0.114
South Africa	0.547	0.587	0.058	0.156
Ghana	0	0	0.205	0.333
Nigeria	0	0	0	0.022
West African Economic and Monetary Union				
Benin	0.867	0.912	0.637	0.655
Burkina Faso	0.416	0.690	0.496	0.646
Guinea-Bissau	0.265	0.487	0.142	0.442
Ivory Coast / Cote d'Ivoire	0	0.124	0.416	0.566
Mali	0	0.186	0	0.035
Niger	0.239	0.496	0.088	0.292
Senegal	0.327	0.575	0.531	0.726
Togo	0.150	0.442	0.150	0.540
		Americas		
Panel C: Latin America and the Caribbean				
Mexico	0.119	0.251	0	0.041
Argentina	0.424	0.507	0.028	0.104
Brazil	1	1	0.225	0.423
Chile	0.580	0.776	0	0
Colombia	0.212	0.497	0	0
Panel D: Northern America				
Canada	0.422	0.431	0	0
United States of America (the US)	0.316	0.440	0	0

		Asia		
Panel E: Central Asia				
Kazakhstan	0.980	1	0	0.078
Panel F: Eastern Asia				
China (Mainland)	0	0	0.033	0.073
Hong Kong	0.360	0.444	0.093	0.382
Japan	0.880	1	0.009	0.067
South Korea	0.591	0.818	0	0.022
Panel G: South-eastern Asia				
Indonesia	0.406	0.461	0	0.006
Malaysia	0	0.009	0.004	0.040
Philippines	0.600	0.707	0	0
Thailand	0.573	0.938	0	0
Panel H: Southern Asia				
Bangladesh	0.117	0.276	0	0.018
India	1	1	0.011	0.197
Iran	0.032	0.194	0.290	0.468
Sri Lanka	0.005	0.062	0	0
Panel I: Western Asia				
Israel	0.007	0.139	0	0
Saudi Arabia	0.316	0.590	0.291	0.735
Turkey	0.017	0.186	0	0
		Europe		
Panel J: Eastern Europe				
Czech Republic	0.608	0.713	0	0.021
Hungary	0.428	0.535	0	0
Poland	0.694	0.745	0.013	0.159
Russia	0.773	0.824	0	0.076
Slovakia	0.944	0.944	0.049	0.098
Ukraine	0.822	0.864	0	0.017
Panel K: Northern Europe				
Finland	0.742	0.844	0	0
Iceland	0.150	0.415	0.007	0.068
Ireland	0.347	0.440	0	0
Lithuania	0.229	0.695	0	0.038
Norway	0.744	0.744	0	0.039
United Kingdom	0.436	0.511	0.240	0.493
Panel L: Southern Europe				
Croatia	0.821	0.829	0	0.024

Greece	0.392	0.485	0	0
Italy	0.440	0.524	0	0
Portugal	0.576	0.638	0.023	0.175
Spain	0.427	0.538	0	0
Panel M: Western Europe				
Belgium	0.404	0.444	0	0
France	0	0.027	0	0.031
Germany	0.427	0.676	0.004	0.027
Netherlands	0.591	0.778	0.351	0.467
Switzerland	0.397	0.693	0.413	0.553
			Oceania	
Panel N: Australia and New Zealand				
Australia	0.538	0.569	0.164	0.244
New Zealand	0.656	0.855	0.124	0.242

Note: The table shows the rejection frequencies at different significant levels for rolling window LF Granger causality tests of non-causality. "SP" denotes stock market returns, while "CP" denotes national commodity export prices. $H_0: CP \nRightarrow SP$ *($\nRightarrow$ means "does not Granger-cause"). We follow Ghysels et al. (2016) and use bootstrapped p-values with N = 499 replications.*

Table C.9 Rejection Frequencies at Different Significant Levels for Rolling Window Low-Frequency Granger Causality Tests, National Commodity Export Prices

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