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Sharing a Ride on the Commodities Roller Coaster: Common Factors in Business Cycles of Emerging Economies

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Abstract

Fluctuations in commodity prices are an important driver of business cycles in small emerging market economies (EMEs). We document how these fluctuations correlate strongly with the business cycle in EMEs. We then embed a commodity sector into a multi-country EMEs' business cycle model where exogenous fluctuations in commodity prices follow a common dynamic factor structure and coexist with other driving forces. The estimated model assigns to commodity shocks 42 percent of the variance in income, of which a considerable part is linked to the common factor. A further amplification mechanism is a "spillover" effect from commodity prices to risk premia.

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I. INTRODUCTION

In recent times, the world economy has witnessed large swings in the prices of commodity goods traded in international markets. These swings have been observed across distinct types of commodities, from agricultural products to fuels and metals. What have been the macroeconomic consequences of these swings for *small* emerging market economies (EMEs) that export these goods? EMEs, often portrayed as being vulnerable to external forces and price takers in commodity markets, may have been subject to more macroeconomic volatility through these fluctuations in the prices of the commodity goods that they export. Thus, an important open question in international macroeconomics is what are the main channels by which commodities price fluctuations affect business cycles in EMEs and how much have they mattered in practice.

Figure 1, panel (a), documents the large volatility experienced in the price indices across three distinct commodity goods traded: agricultural raw materials, metals, and fuels. Since the early 90s all three indices display increases and then contractions in the second half of that decade. But the largest swings are observed since the early 2000s when all three types of commodities experienced considerable price increases. This was interrupted by the global financial crisis, but the recovery came soon after it with considerable strength. Since early 2012, however, all three indices have exhibited downturns, some more pronounced than others, like the one recently experienced by fuels. Panel (b) of Figure 1 documents the strong cyclical comovement between commodity prices and economic activity in EMEs focusing on the years before and after the global financial crisis. It plots the cyclical components of a commodity price index that aggregates all commodity goods and real GDP for a pool of 13 small EMEs using simple averages.[1] In the years that preceded the crisis income increased up to 4 percentage points above its trend while commodities did so by more than 50 points. The financial crisis materialized in a sharp fall in economic activity, with real output bottoming close to 4 percentage points below trend in the second quarter of 2009. Commodity prices experienced a faster drop to minus 50 points and bottomed out one quarter before.

This first look at the recent historical account of commodity prices and EMEs business cycles suggests that fluctuations in the price of the commodity goods -easily comparable to a wild roller coaster ride, hence the title of our work- may be a key driver of their business cycles. This paper explores this hypothesis formally. It does so by using data and economic theory. First, we explore how much commodities matter in the average EME and then document the comovement between the international prices of these commodities and several macroeconomic variables across a pool of EMEs. We then build a structural, small open economy model that allows us to articulate a simple and tractable theory of how fluctuations in the commodity goods that these economies export can be drivers of business cycles. Lastly, we estimate the structural model to assess how important such drivers have historically been through the lens of this theory.

The paper makes three types of contributions. The first one is empirical as we document the share of commodities in total exports as well as the cyclical properties of commodity prices in EMEs. We do so by studying historical weights of these commodities in total exports in these economies, and then using them to construct country-specific commodity price indices from 44 distinct commodity goods. We then correlate these indices at the country level with macroeconomic variables and

[1] The countries included in this pool of 13 EMEs along with the criteria used when selecting them will be discussed at length in Section 2. The Appendix presents the same plot for the entire sample period of time, 1990-2014. The correlation between the two series in that longer period is 0.64.

assess their comovement along the business cycle. The second contribution is methodological as we build a fully dynamic and stochastic equilibrium model of EMEs' business cycles. The core of the model is built around the small open economy setup as in the seminal work by Mendoza (1991) and further analyzed by Schmitt-Grohé and Uribe (2003). But we extend it in several novel dimensions. First and foremost, we add a commodity endowment sector which takes the price of its goods as given from world markets. Exogenous fluctuations in the price of the commodity good that each EME in the model exports constitute one among other, more traditional, competing drivers of the business cycle. We allow prices of the various commodity goods produced across countries to have a common factor by adding a dynamic latent factor structure to the structural model. The final contribution is quantitative as we estimate this model with Bayesian methods using data of several EMEs. The multi-country setup allows us to quantify the role played by the common factor in commodity prices.

The economics behind the effect of a commodity price shock in our model are simple. When positive, it acts as an income shocks that pushes up consumer demand for domestic goods, increasing their relative price and appreciating the country's real exchange rate. The rental rate of capital also goes up following the subsequent increase in the demand for capital from domestic good producers. This creates incentives for more investment, expanding the investment good sector, and driving up the price of capital goods. Because non-commodity exports (imports) become relatively more (less) expensive, the trade balance deteriorates. Yet for realistically calibrated shares of exports as well as the relative size of domestically produced goods in total consumption and investment, the contraction in the trade balance does not offset the increase in absorption, and the non-commodity domestic sector expands. This pushes up equilibrium employment and expands wages. The model can thus deliver commodity price-driven business cycles akin to those observed in emerging economies where consumption and investment are procyclical, and real exchange rates and trade balance are countercyclical.

On the empirical front, we document three stylized facts. On one hand, the typical EME is a commodity exporter: the share of commodities in total exports in the average EME is more than double that of advanced economies. This is important insofar as business cycle models of EMEs have largely abstracted from this intrinsic characteristic of these economies. Second, the country-specific commodity prices in EMEs that we build exhibit strong comovement with other macro variables along the business cycle. They are procyclical and lead the cycle of production, consumption and investment. In addition, they are countercyclical to real exchange and measures of external risk premia, i.e., periods when commodity prices soar are accompanied by a real appreciation of the exchange rate and cheaper access to foreign capital markets. Lastly, we uncover a preponderant role of common factors when accounting for the dynamics of the country-specific commodity price indices that we build for several EMEs. This extends also to the dynamics of real gross domestic product. Such comovement does not come from EMEs exporting the same types of goods but is rather associated with different commodity good prices being driven by global forces.

The estimation of the structural model gives commodity price shocks a paramount role when accounting for aggregate dynamics in EMEs. The median share of the forecast error variance in output accounted by these shocks is 42 percent. Importantly, the bulk of the action from commodity prices is recovered by the model in the form of common shocks across economies. This common factor exhibits a marked increase in its volatility in the post-2005 period. Furthermore, it allows to bring the model much closer to the data in terms of business cycle comovement across EMEs, as well as other more traditional second moments studied in previous studies. For instance, the model predicts that real exchange rates are countercyclical, as observed unconditionally in the data. Fluctuations in commodity price have not always amplified the business cycle, though. A historical decomposition of the output gap reveals that, sometimes, they have acted as cushion devices against what the model identifies as domestic forces. This was particularly the case in the fast recovery after the world financial crisis when commodity prices rebounded and helped counterbalance negative domestic shocks in some EMEs.

Several robustness tests and extensions are conducted on the benchmark model. First, SVAR results yield close estimates to those from the structural model in terms of the share of variance associated with commodity price shocks. Second, extending the model to allow for the risk premia to be affected by country fundamentals increases the role of commodity prices because the estimation finds a sizeable effect on risk premia that acts as an amplification mechanism. Third, the important role of commodity price shocks remains even when we allow for the common factor in commodity prices to be correlated with other external forces, particularly external demand for non-commodity goods. Fourth, we show that, when turning off commodity price shocks in the structural estimation, total factor productivity shocks become the main driver but at a cost in terms of the model's performance. Notably, it counterfactually predicts that expansions (contractions) in economic activity are accompanied by real exchange rate depreciations (appreciations). Fifth, we estimate the dynamic factor model structure in isolation from the structural model, which allows us to extend the analysis to more countries/years and thus bring more information when estimating the common factor in commodity prices. The alternative common factors track down pretty well the one obtained in our benchmark estimation and continue to show that the volatility exhibited in the post-2005 subperiod stands out even when looking at a longer historical period. Sixth, we corroborate that our main results survive when using growth rates as alternative detrending filter. We conclude with a discussion about our modelling choices of an endowment in the commodity sector as well as the absence of fiscal policy in our model. While we view those two extensions as important for future work, we present evidence that supports the simplifying assumptions made.

This paper can be related to at least three strands of literature. The first and closest to our work is the literature that has used dynamic, equilibrium models to account for business cycles in small open and emerging economies. A set of papers in this literature has explored the role of terms of trade variations in driving aggregate fluctuations in EMEs (Mendoza, 1995; Kose, 2002). Our paper complements their analysis by focusing on commodity prices as one specific source of terms of trade variability. A more recent set of papers has explored the role of financial shocks and/or the amplifying effects of frictions, mostly of financial nature, when it comes to accounting for business cycles in EMEs (Neumeyer and Perri, 2005; Uribe and Yue, 2006; Aguiar and Gopinath, 2007; García-Cicco, et al., 2010; Fernandez-Villaverde, et al 2011; Chang and Fernández, 2013; and Fernández and Gulan, 2014). We extend this strand of the literature by postulating a link between commodity prices and financial conditions in EMEs and quantifying its relevance when accounting for aggregate fluctuations in these economies within a structural framework. In that sense, our work is related to Calvo and Mendoza (2000), who link the volatility in financial conditions for EMEs to the cyclicality of their terms of trade and other fundamentals in the context of informational frictions where uninformed investors cannot extract information from prices but rather do so from noisy information about specialists´ trades.[2]

[2]In a historical context, Eichengreen (1996) documented that during the crash of 1929 the sharp drop in the price of Brazilian coffee led foreign bankers to stop extending loans to Brazilian borrowers. And Min et.al. (2003) found that improved terms of trade are associated with lower yield spreads to the extent that such improvements imply an increase in

A second strand of literature that our work relates to has documented the presence of common factors in business cycles across world economies at both global and regional levels (Kose, et al., 2003; Mumtaz, et al., 2011). This has largely been investigated, separately, for developed economies (Kose, et al., 2008; Aruoba, et al., 2010; Crucini, et al., 2011; Kose, et al., 2012; Guerron-Quintana, 2013), and emerging economies (Broda, 2004; Bartosz, 2007; Akinci, 2013; Miyamoto and Nguyen, 2014). Within the literature of EMEs special attention has been given to two potential drivers of business cycles: fluctuations in external interest rates (Canova, 2005; Bartosz, 2007; Akinci, 2013) and terms of trade (Broda, 2004; Izquierdo, et al. 2008).[3] Our contribution to this literature is twofold. We provide further empirical evidence on the existence of common external forces coming from movements in commodity prices that drive business cycles in EMEs. With just a few exceptions (Guerron-Quintana, 2013; Miyamoto and Nguyen, 2014) most of this literature has not used structural models when evaluating the role of common factors. It is in that sense that we also contribute to this literature as we quantify this role by using a structural and estimated multi-country model in which common and idiosyncratic external forces interact.

A final strand of literature related to our work is one that documents the comovement of commodity prices. Since at least the work by Pindyck and Rotemberg (1990) it has been documented that prices of unrelated raw commodities have a persistent tendency to move together. More recently, this result has shown to be robust to the use of newer and more sophisticated statistical methods such as FAVAR models (Lombardi, et al., 2012; Byrne, et al., 2013) and networks analysis (Gomez, et al., 2011). Our work contributes to this literature by explicitly incorporating latent common factors in commodity prices and measuring their contribution to the business cycle of several EMEs through a full-fledged dynamic, stochastic equilibrium framework.

The rest of the paper is divided into seven sections including this introduction. Section 2 presents the set of stylized facts found. Section 3 builds the model. Section 4 discusses some of the details of the strategy used when taking the model to the data. Section 5 presents the main results of the estimated model and Section 6 reports robustness checks. Concluding remarks are given in Section 7. Additional material is gathered in a companion online Appendix.

II. Stylized Facts

A. Commodity Export Shares

The first empirical task that we set out to do is to assess the size of the share of commodity exports in a typical EME. To answer this we assemble a large (unbalanced) annual panel using the World

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export earnings and better repayment capacity. Cuadra and Sapriza (2006) link the volatility of terms of trade in EMEs to spreads in a dynamic model with strategic default model that delivers endogenous default risk, but do not explore the implications for the business cycle. Using FAVAR models, Bastourre et al. (2012) have also documented a strong negative correlation between commodity prices and emerging market spreads. In the context of the subprime crisis, Caballero et al. (2008) argued that persistent global imbalances and the volatility in both financial and commodity prices (such as oil) and asset prices that followed the crisis stemmed from a global environment where sound and liquid financial assets are in scarce supply. Morana (2013) has recently found that financial shocks have had a much larger role as drivers of the price of oil than previously noted in the literature.

[3]Within the group of emerging economies, some particular attention has been given to Latin America (Canova, 2005; Izquierdo, et al., 2008; Aiolfi, et al., 2011; Cesa-Bianchi, et al., 2012).

Bank's Word Development Indicators (WDI) covering 189 countries between the years 1960 and 2013. For each country, the panel contains the share of three broad groups of commodities in total exports: agricultural, fuel and metals. We average the sum of these three shares across time to obtain country-specific commodity export shares. Finally we separate countries between 61 EMEs and 128 non-EMEs. The latter is also further subcategorized into 74 advanced and 54 low-income economies[4].

The main descriptive statistics are reported in Table 1. The table documents the median export share across all groups of countries.[5] It also displays Mann-Whitney tests for the equivalence between non-EMEs groups and EMEs, and Kruskal-Wallis for equivalence across all groups. The results in this Table allow us to define a first stylized fact:

- **Stylized Fact 1**: **The share of commodities in total exports in the average EME is more than double that of advanced economies.** While the median export share in EMEs is 25.7 percent, that in advanced is 11.2 percent. Furthermore, a Mann - Whitney test easily rejects equivalence between the two distributions at 1 percent significance level. It fails to reject that hypothesis when emerging economies are compared to low-income economies, but continues to accept it when EMEs are compared to non-EMEs, mainly due to the smaller number of observations for low-income countries relative to those for advanced economies. Kruskal-Wallis tests easily reject equivalence across these three groups of countries. This stylized fact is robust to two alternative measures of the relative importance of commodities: the gross and net share of commodity exports to GDP (see Appendix).

B. Cyclicality

We now explore the cyclicality of commodity prices. For that purpose we use another, more refined, quarterly panel dataset of country-specific commodity price indexes for EMEs, between 1980.Q1 and 2014.Q4. The indices are constructed by averaging the time series of the international prices of 44 commodity goods using as constant weights the (country-specific) average shares of each of these commodities in total exports between 1999 and 2004. Formally, country n's commodity price index in quarter t, $P^{Co}_{n,t}$, is defined as

$$P^{Co}_{n,t} = \sum_{i=1}^{44} \theta_{i,n} P^{Co}_{t,i},$$

where $\theta_{i,n}$ is the export share of commodity good i in total commodity exports by n, and $P^{Co}_{t,i}$ is the real USD spot price of commodity i in world markets. The share $\theta_{i,n}$ is computed by averaging the shares of commodity i in total commodity exports by n between the years 1999 and 2004, using

[4]The 61 EMEs in our sample are classified as such following a simple criteria: we classify a country as an EME if there exists EMBI data on this country. The only two countries that we exclude from this list are China and India, as they clearly do not fall into the category of *small and open* emerging economy that is the focus of our analysis. Advanced and low-income economies are classified following World Bank's classification. See the Appendix for the list of countries, average country-specific commodity shares and further details.

[5]We report medians instead of means as the distributions in each of the groups are highly skewed and simple tests reject normality. The Appendix contains more descriptive statistics.

United Nations' Comtrade data. $P_{t,i}^{Co}$ is obtained by deflating the monthly commodity prices indices reported by the IMF's Primary Commodity Price Database with the US consumer price index.[6]

We study the comovement of these indices with several country-specific macro variables. Formally, we compute the serial correlation $\rho\left(X_{n,t}, P_{n,t+j}^{Co}\right)$ with $j = -4, ..., 4$; where X is sequentially replaced by real output, real private consumption, real investment, trade balance, real exchange rates, and external interest rate premia faced by EMEs in world capital markets. The latter is proxied with JP Morgan's EMBIG and CEMBI spreads. Unfortunately, imposing a minimum range of time series data to make the analysis valid reduces the sample to 13 EMEs.[7] Each of the serial correlations is depicted on a subplot in Figure 2. Statistics reported are simple averages across the 13 countries (EME13) in the sample and confidence bands denote +/-1.5 standard deviations. We also report results for a subset of these countries that we call LAC4 (Brazil, Chile, Colombia, and Peru) which will be explored more closely later in the paper. We use the Hodrick-Prescott filter to extract the cyclical component of all variables.[8] Inspecting the panels in this figure a second stylized fact can be defined:

- **Stylized Fact 2**: **The country-specific commodity prices in EMEs are procyclical and lead the cycle of output, consumption and investment. In addition, they are countercyclical to real exchange rates and measures of external risk premia.** The average contemporaneous correlation between the commodity price index and real GDP is about 0.5, and is slightly higher when the index is lagged one quarter (panel a, Figure 2). This correlation further increases to 0.6 if real income is computed by deflating GDP with the CPI (panel b). We present this alternative measure of real income because earlier works have demonstrated how real GDP tends to underestimate the changes in real domestic income following terms of trade movements (Kohli, 2004). The strong procyclicality and leading property of the indices remains when X is replaced with real consumption and investment (panels c and d). This explains why lagged values of the index are negatively correlated with the trade balance but commove contemporaneously, although the correlations are less tightly estimated (panel e).

Panel f in Figure 2 reveals that commodity price indices are negatively correlated with the real exchange rate, defined in terms of a basket of foreign goods. Hence increases in the indices are accompanied by real appreciations of the exchange rate. Lastly, panels g and h document a negative comovement between commodity price indices and interest rate spreads in EMEs. Such negative comovement is actually stronger when considering only corporate risk premia, as proxy by the CEMBI.

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[6] We choose to use constant weights as averages of the 1999-2004 period largely to be consistent with the structural model estimated later in the paper which assumes constant weights (i.e., a commodity endowment) during the period of estimation (2000-2014). We also deflate commodity prices with US CPI to be consistent with the model, where foreign prices are used as numeraire. More details of the construction of the indices, including the entire list of commodity goods and weights are presented in the Appendix.

[7] These countries are: Argentina, Brazil, Bulgaria, Chile, Colombia, Ecuador, Malaysia, Mexico, Peru, Russia, South Africa, Ukraine and Venezuela. The median commodity export share in this group is 28.7, only slightly above that of the 61 EMEs studied in the previous subsection. We only selected countries with (i) at least 32 consecutive quarterly observations of EMBI spreads and covering at least until 2014.Q1; (ii) whose median commodity export share is above the median for all 61 EMEs; and (iii) with quarterly time series for real GDP at least from 2000.Q1. Data on real output, real consumption, real investment and trade balance come from Haver Analytics; Real effective exchange rates, Nominal GDP and CPI are taken from IFS; and EMBI/CEMBI from Bloomberg. See the Appendix for further details.

[8] In the Appendix we present the same descriptive statistics using growth rates as alternative filter. The results are strongly robust.

Thus, when commodity prices are high (low), the cost of issuing debt in foreign capital markets decreases (increases) for EMEs.[9]

C. Common Factors

A last dimension that we explore empirically is the presence of common factors in our measures of commodity prices indices across EMEs. The evidence is presented in Figure 3, which plots time series of the (cyclical) indices for each of the EMEs in our sample. More formal principal component analysis is also conducted but is presented in the Appendix to economize space. This evidence leads us to define the third and final stylized fact of this section:

- **Stylized Fact 3**: **There is a preponderant role of common factors when accounting for the dynamics of commodity price indices across EMEs. This also extends to the dynamics of real gross domestic product.** A look at the time series dynamics in Figure 3 reveals the presence of strong comovement in the country-specific commodity price indices for the 13 EMEs that we study. Principal component analysis further corroborates this: the first principal component accounts for as much as 78 percent of the variance in the indices across these EMEs. This does not occur mechanically because the commodity exporting profiles of the countries in our sample are similar. In fact they differ substantially.[10] Instead, the main reason comes from the fact that the international prices of various commodity goods comove. To verify this, we extend the principal component analysis across the prices of the 44 commodity goods in our sample, which we group into five categories according to the Standard International Trade Classification (SITC, fourth revision)'s one level aggregation and also extract their cyclical component.[11,12] The first principal component accounts for 55 percent of the variance in commodity prices across the five SITC categories.

Finally, we also conduct principal component analysis on the cyclical component of real GDP across the 13 EMEs in our sample. The results point also in the direction of a strong common factor, virtually as strong as that in the commodity price indices: the first principal component accounts for 76 percent of the variance of economic activity.[13]

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[9]For the interested reader, the Appendix presents the average time series plot of output gap and the country-specific commodity index across the 13 EMEs, and the 4 Latin American economies akin to Figure 1 (b) since the 1990s. The correlation between output gap and the country commodity price index is 0.48 and 0.52, respectively.

[10]For example, the serial correlation between the commodity price indexes for Colombia and Peru is 0.9 despite the fact that the commodity export patterns of the two countries differ in terms of the type of commodity goods that they export: while the two largest commodity export shares for Colombia are crude oil (53 percent) and coal (15 percent), those of Peru are gold (28 percent) and copper (22 percent). The Appendix presents the specific shares of each of the commodities in our sample for all 13 countries in our dataset.

[11]The groups are as follows (distribution of the 44 commodities in parentheses). Group 0: Food and live animals (19 commodities); Group 2: Crude material, inedible, except fuels (15 commodities); Group 3: Mineral fuels, lubricants and related materials (3 commodities); Group 4: Animal and vegetable oils, fats and waxes (5 commodities); Group 9: Commodities and transactions not classified elsewhere in the SITC (1 commodity).

[12]For the sake of space, the Appendix contains time series plots for the price of commodity goods and GDP, as well as information on the SITC aggregation that we use.

[13]A well-known potential caveat to conducting analysis of comovement between variables that have been filtered with the Hodrick-Prescott filter is that such a filter may generate spurious correlations. The Appendix presents the same set of results using growth rates as alternative filter. The results continue to point to strong comovement. If anything,

Taken together, the stylized facts presented in this section contribute to further improving the understanding of the main patterns exhibited by EMEs' business cycles. On one hand they shed light on the strong comovement between aggregate macro variables in these economies and the prices of the commodities that they export. In addition, because the relative share of these commodities is shown to be large in these economies and movements in their price are largely exogenous, they can be regarded as an important driver of EMEs' business cycles. The next section builds a dynamic general equilibrium model guided by these stylized facts where we formally articulate a mechanism by which exogenous changes in commodity prices turn into fluctuations in real economic activity.

III. Model

A. Setup

The setup of our model is a multi-country version of the small open economy framework first developed by Mendoza (1991), and further analyzed by Schmitt-Grohé and Uribe (2003). We take five departures from such framework. First, we add a country-specific commodity sector that faces fluctuations in the price of the good it sells in international markets. These fluctuations are exogenous, as we assume the countries are small players in these markets. The commodity good is an endowment that is entirely sold abroad and the income generated accrues directly to households who own the sector. Second, there are foreign (f) and (country-specific) home (h) goods, which are imperfect substitutes when consuming them or using them to produce investment goods. Home goods are produced domestically using capital and labor and a random domestic productivity shock. Foreign goods are imported from the rest of the world. Third, there is a sector that produces investment goods using home and foreign goods as inputs. As in the standard framework, households in each EME can issue non state-contigent, one-period bonds in international financial markets. Such bonds will pay a premium over the world interest rate. Both the premia and the world interest rate are exogenous and stochastic, acting as two additional driving forces.

The structure with which we model commodity prices constitutes a fourth novelty of our framework. We model them with a dynamic factor structure that incorporates a latent common factor in addition to idiosyncratic shocks. This structure is intended to capture the strong comovement across EMEs in the data documented in the previous section. Fifth, the multi-country structure of our framework comes from jointly modelling a collection of N EMEs that interact with the rest of the world as small open economies. The sole source of comovement across these EME comes from shocks to the common factor in the prices of the commodity goods that they sell in international markets.[14]

There are four agents in each EME considered in the model: households, firms, investment goods producers, and the rest of the world (which does not include the other EMEs in the model). Households consume final goods, defined as a bundle of home and foreign goods; decide how much labor

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they suggest an even higher degree of synchronization of the commodity price indices and real economic activity across EMEs.

[14]We are thus abstracting from trade linkages across the EMEs in the model, mostly for tractability. While in principle trade across EMEs can potentially be relevant for explaining their business cycle comovement, in the empirical application of the model we later provide evidence of the relatively low trade linkages among the EMEs chosen to estimate the model. We conjecture that, should trade linkages be added to our framework, the novel role of common factors that our work is highlighting would be further empasized.

and capital supply to domestic firms; and issue bonds in foreign markets. They also purchase investment goods to replace depreciated capital and increase the net stock of capital for which they face capital adjustment costs. Besides income from the commodity sector, they receive profits from firms, as well as capital and labor income. Firms maximize profits, defined as the revenue from selling home goods net of costs from renting labor and capital from households. Home goods are produced with a standard neoclassical technology and sold either to households, investment good producers or the rest of the world. Investment goods producers use a technology that combines home and foreign goods. They later sell these goods to households for capital accumulation purposes. The fourth agent is the rest of the world. It provides funding for households at a stochastic premium over the world interest rate. It also demands home goods for its own consumption as a function of both their relative price and a price elasticity, as well as exogenous external demand shocks. Finally, it provides (foreign) goods that are imported by households for their own consumption or used as inputs by investment good producers.

In the following subsections we formally describe the actions by each of the agents in a representative j^{th} EME in the model. However, we omit the country index to simplify the notation and only use it when common and idiosyncratic variables interact. The full set of equilibrium and optimality conditions is included in the Appendix.

B. Households

Households' lifetime utility is given by

$$E_0 \sum_{t=0}^{\infty} \beta^t U\left(C_t, L_t\right) \tag{1}$$

where E_0 is the expectation operator with information up to period $t = 0$, β is the intertemporal discount factor, $U(\cdot)$ is the concave period utility function, L_t is total hours worked, and C_t is consumption goods. We choose a GHH specification for $U(\cdot)$:

$$U\left(C_t, L_t\right) = \frac{1}{1-\sigma}\left[C_t - \psi^c \frac{L_t^{1+\gamma_c}}{1+\gamma_c}\right]^{1-\sigma}$$

where σ is the constant relative risk aversion coefficient and γ_c is the inverse of the Frisch elasticity of the labor supply. As it is well known, the key implication of these preferences is that the income effect does not affect the labor supply decision of the household. These preferences have been used extensively in previous works on business cycles of emerging economies (see, among others, Neumeyer and Perri, 2005; Uribe and Yue, 2006).

Households maximize (1) subject to the budget constraint and to the capital accumulation equation. The budget constraint is defined as:

$$p_t^c C_t + p_t^x X_t + R_{t-1} D_{t-1} = w_t L_t + r_t^k K_{t-1} + D_t + p_t^{Co} \bar{C}o + \xi_t \tag{2}$$

where p_t^c is the price of the consumption good, D_t is the stock of international debt at the beginning of each period, w_t is the real wage, R_t is the (gross) external real interest rate, p_t^x is the price of the investment good, p_t^{Co} is the unit price of a constant endowment flow of $\bar{C}o$ quantities of commodity goods, ξ_t are profits from the domestic production sector, r_t^k is the rent of capital and K_{t-1} is the stock of that capital. The full revenue from the commodity sector, $p_t^{Co}\bar{C}o$, is assumed to accrue to households. Thus, commodity price shocks will act as exogenous revenue fluctuations in the household's budget constraint.[15]

Consumption is assumed to be a bundle of domestic and imported goods with a constant elasticity of substitution (CES) function as follows

$$C_t = \left[(1-\alpha_c)^{\frac{1}{\eta_c}} \left(C_t^h\right)^{\frac{\eta_c - 1}{\eta_c}} + \alpha_c^{\frac{1}{\eta_c}} \left(C_t^f\right)^{\frac{\eta_c - 1}{\eta_c}} \right]^{\frac{\eta_c}{\eta_c - 1}} \tag{3}$$

where C^h and C^f denote domestic and imported consumption goods, η_c is the elasticity of substitution between the two and $\alpha_c \in (0,1)$ is a parameter that determines the share of imported goods in total consumption. Because we use the price of the imported good as the numeraire in the model, total expenditure in consumption goods will be

$$p_t^c C_t = p_t^h C_t^h + C_t^f \tag{4}$$

where p_t^h is the price of domestic goods in terms of the numeraire. All prices are then expressed in relative terms. As is well known, p_t^c will be a function $\phi\left(p_t^h, 1\right)$ with a CES functional form as well:

$$p_t^c = \phi\left(p_t^h, 1\right) = \left[(1-\alpha_c)\left(p_t^h\right)^{1-\eta_c} + \alpha_c (1)^{1-\eta_c} \right]^{\frac{1}{1-\eta_c}} \tag{5}$$

Hence, the real exchange rate is defined in the model as the inverse of the consumption good's price, $(p_t^c)^{-1}$.[16]

The capital accumulation equation is

$$K_t = (1-\delta) K_{t-1} + X_t \left(1 - s_t\left(\frac{X_t}{X_{t-1}}\right)\right) \tag{6}$$

[15] This modeling approach of the commodity sector is evidently simplistic as we do not incorporate a production sector that uses resources, nor do we incorporate a government sector that directly benefits from higher commodity prices (e.g., via higher tax revenues) which later spills over into the economy. The robustness section will include a discussion of the implications of these modeling assumptions as well as the empirical evidence that justified them.

[16] When computing the real exchange rate like this we are assuming that the law of one price (LOOP) holds between foreign goods in the EME considered and the rest of the world's domestic goods: $NER_t P_t^{h*} = P_t^f$, where NER_t is the nominal exchange rate, and P_t^{h*} and P_t^f are, respectively, the (nominal) price of the domestic good in the rest of the world (ROW) economy and the foreign good bought by the EME. The second assumption made is that the LOOP does not hold between P_t^{f*} and P_t^h, which are, respectively, the (nominal) foreign price faced by ROW and the domestic good price in EME: $NER_t P_t^{f*} \neq P_t^h$. Arguably, while ROW does indeed consume home goods of EME, these are just a marginal fraction from the perspective of that economy. Formally: $P_t^{c*} = \phi\left(P_t^{h*}, P_t^{f*}\right) \simeq \widetilde{\phi}\left(P_t^{h*}\right)$, where $\widetilde{\phi}\left(P_t^{h*}\right)$ is linear in P_t^{h*} and $\widetilde{\phi}(1) = 1$. In that case, the real exchange rate will be $NER_t P_t^{c*}/P_t^c$, which can be rewritten as $\phi\left(p_t^h, 1\right)^{-1}$.

where $s(\cdot)$ is a cost function with the following properties $s_t(1) = s_t'(1) = 0$, and $s_t''(\cdot) > 0$. In particular, we follow Christiano et al. (2010) and assume the following functional form

$$s_t\left(\frac{X_t}{X_{t-1}}\right) = \frac{1}{2}\left(e^{\left(\sqrt{a}\left(\frac{X_t}{X_{t-1}}-1\right)\right)} + e^{\left(-\sqrt{a}\left(\frac{X_t}{X_{t-1}}-1\right)\right)} - 2\right) \tag{7}$$

C. Production of h Goods

Firms in the domestic sector maximize profits, ξ_t, subject to a standard neoclassical production technology that uses capital and labor. Formally, the static optimization problem for domestic firms is

$$\max \xi_t = p_t^h Y_t - w_t L_t - r_t^k K_{t-1}$$

subject to a production technology:

$$Y_t = z_t K_{t-1}^{\alpha} L_t^{1-\alpha} \tag{8}$$

where Y_t denotes domestic output, z_t is the productivity shock.

D. Investment

The investment good, X_t, is produced with imported and home goods as intermediate inputs. The production technology for new investment goods is given by:

$$X_t = \left[(1-\alpha_x)^{\frac{1}{\eta x}}\left(X_t^h\right)^{\frac{\eta x-1}{\eta x}} + \alpha_x^{\frac{1}{\eta}c}\left(X_t^f\right)^{\frac{\eta x-1}{\eta x}}\right]^{\frac{\eta x}{\eta x-1}} \tag{9}$$

where X^h and X^f are domestic and imported goods used by the investment sector, η_x is the elasticity of substitution and $\alpha_x \in (0,1)$ is a parameter that determines the share of imported goods in total investment.

E. Market Clearing

The market clearing condition in the home goods market is:

$$Y_t = C_t^h + X_t^h + C_t^{h*} \tag{10}$$

where C_t^{h*} is external demand for home goods. We model it for country j as

$$C_{j,t}^{h*} = \left(p_{j,t}^h\right)^{-\epsilon_{j,e}} Y_t^* \tag{11}$$

where Y_t^* denotes the level of aggregate demand in the rest of the world, which we assume to be an exogenous process (and independent from country j), and $\epsilon_{j,e}$ is the parameter that governs the price elasticity of foreign demand.

Lastly, when taking the model to the data in the following sections, it will be convenient to define measures of real income and trade balance for the typical EME in the model. The former is measured as nominal GDP deflated by the consumer price index, GDP_t/p_t^c, where:

$$GDP_t = p_t^h Y_t + p_t^{Co}\overline{Co} \tag{12}$$

The trade balance, TB_t, is defined as the sum of commodity and non-commodity exports minus imports used as intermediate goods when producing final consumption and investment goods:

$$TB_t = p_t^{Co}\overline{Co} + p_t^h C_t^{h*} - C_t^f - X_t^f \tag{13}$$

F. Driving Forces

1. Common Factor Structure in Commodity Prices

The strong comovement of commodity prices across EMEs documented in the stylized facts is modeled with a dynamic factor structure. Following Geweke and Zhu (1996) and Jungbacker and Koopman (2008) we postulate a latent factor, f_t^{Co}, that evolves according to an AR(1) process:

$$f_t^{Co} = \phi_{Co} f_{t-1}^{Co} + \sigma^{f^{Co}} \varepsilon_t^{f^{Co}}, \quad \varepsilon_t^{f^{Co}} \sim N(0,1) \tag{14}$$

The (country-specific) commodity price, $p_{j,t}^{Co}$, is related to the common factor as follows:

$$p_{j,t}^{Co} = \omega_j^{Co} f_t^{Co} + \varepsilon_{j,t}^{Co} \tag{15}$$

where ω_j^{Co} is the loading factor associated to f_t^{Co} for the j^{th} economy, capturing the extent to which changes in the common factor percolate to changes in $p_{j,t}^{Co}$. The idiosyncratic component in (15), $\varepsilon_{j,t}^{Co}$, is assumed to behave as an AR(1) process

$$\varepsilon_{j,t}^{Co} = \rho^{Co} \varepsilon_{j,t-1}^{Co} + \sigma_j^{Co} \nu_{j,t}^{Co}, \quad \nu_{j,t}^{Co} \sim N(0,1) \tag{16}$$

where shocks $\nu_{j,t}^{Co}$ account for movements in $p_{j,t}^{Co}$ that are independent from the common factor. Hence, all else equal, the stronger and more persistence are the fluctuations in f_t^{Co} the more will the model assign movements in $p_{j,t}^{Co}$ to common factors across all EMEs considered.

An economic interpretation can be given to the two forces driving $p_{j,t}^{Co}$ in (15). The common factor, f_t^{Co}, can be thought of as a reduced way of capturing global demand shocks that simultaneously affect the price of several and distinct commodity price goods produced by EMEs. A boost in Chinese absorption, for example, could result in higher demand for various types of commodity goods, from soybeans to copper and crude oil, which in turn increase their international market prices due to the large market power of China and the sluggish response of global supply for these commodities.[17]

[17] There are, however, several more examples of global demand shocks that one can come up with, that are not associated with China. For that reason we prefer the flexibility of modeling the common factor as an unobserved/latent factor.

The idiosyncratic component, $\varepsilon_{j,t}^{Co}$, can be considered as stemming from supply side factors. For example, these could be coming from new technologies that reduce the price of a particular commodity produced by economy j. In addition, they could be interpreted as capturing new discoveries of the commodity endowment and hence observationally equivalent to endowment shocks. If, for instance, economy j discovers a new oil well, the idiosyncratic component would be capturing increases in that countries' revenue.[18]

2. Other Driving Forces

In addition to the two commodity price shocks described above, ε_t^{fCo} and $\nu_{j,t}^{Co}$, an EME in the model will face four additional sources of uncertainty. Two of them will be coming from interest rates. As in Neumeyer and Perri (2005), we assume that a large mass of international investors is willing to lend to the j^{th} EME any amount at a country-specific interest rate, R_{jt}, and that these loans are risky assets because there can be default on payments to foreigners. This assumption creates two sources of volatility in R_{jt}. On one hand, real interest rates may change as the perceived default risk changes in economy j. On the other, even if default risk stays constant, interest rate can change because the preference of international investors for risk changes over time. We capture these two sources of interest rate volatility by decomposing the interest rate faced by every EME as

$$R_{jt} = R_t^* S_{jt} \tag{17}$$

where R_t^* is assumed to be the world interest rate, which is not specific to any economy, and S_{jt} captures the country spread over R_t^*. We model the evolution of the external interest rate as an AR(1) process

$$\ln R_t^* = (1-\rho_{R^*}) \ln \overline{R^*} + \rho_{R^*} \ln R_{t-1}^* + \sigma^{R^*} \varepsilon_t^{R^*}, \qquad \varepsilon_t^{R^*} \sim N(0,1) \tag{18}$$

where shocks $\varepsilon_t^{R^*}$ capture changes in the risk appetite of foreign investors. As in Neumeyer and Perri (2005), default decisions are modeled assuming that private domestic lenders always pay their obligation in full but that in each period there is a probability that the local government will confiscate all the interest payments going from local borrowers to foreign lenders. Fluctuations in the perceived confiscation probability by investors are assumed to obey to idiosyncratic exogenous forces. Formally, we assume that S_{jt} is defined as

$$S_{jt} = z_{jt}^R \left\{ \exp\left[\Omega_{j,u}(D_{j,t} - \overline{D}_j)\right] - 1 \right\} \tag{19}$$

where the term $\exp\left[\Omega_{j,u}(D_{j,t} - \overline{D}_j)\right]$ is a debt-elastic interest rate mechanism whose role is to induce stationarity in the debt process, as in Schmitt-Grohé and Uribe (2003); and z_{jt}^R is the exogenous country risk spread.[19] The latter follows an AR(1) process:

$$z_{j,t}^R = \left(1 - \rho_j^{z^R}\right) \overline{z_j^R} + \rho_j^{z^R} z_{j,t-1}^R + \sigma_j^{z^R} \nu_{j,t}^R, \qquad \nu_{j,t}^R \sim N(0,1) \tag{20}$$

[18] In that sense, the structure of the DSGE would allow for an identification of these idiosyncratic sources of revenue fluctuations when taking the model to the data, as will be done later in the paper. Note that one could not pin down such exogenous revenue sources if the dynamic factor structure is estimated outside the DSGE model. In the robustness sections we will consider this case nonetheless.

[19] We will later consider an extension to this where the spread reacts to domestic economic conditions.

The remaining two driving forces are the world's foreign demand and the country-specific technology which we characterize as AR(1) processes[20,21]

$$\ln Y_t^* = (1-\rho_{Y^*})\ln \overline{Y^*} + \rho_{Y^*}\ln Y_{t-1}^* + \sigma^{Y^*}\varepsilon_t^{Y^*}, \quad \varepsilon_t^{Y^*} \sim N(0,1) \tag{21}$$

$$\ln z_{j,t} = \left(1-\rho_j^z\right)\ln \bar{z}_j + \rho_j^z \ln z_{j,t-1} + \sigma_j^z \varepsilon_{j,t}^z, \quad \varepsilon_{j,t}^z \sim N(0,1) \tag{22}$$

G. Equilibrium

Given initial conditions $K_{j,-1}$ and $D_{j,-1}$, and exogenous state-contingent sequences of idiosyncratic and common shocks $\{\nu_{j,t}^{Co}, \nu_{j,t}^{R}, \varepsilon_{j,t}^{z}, \varepsilon_t^{f^{Co}}, \varepsilon_t^{R^*}, \varepsilon_t^{Y^*}\}$ in economy j, an equilibrium is a set of state-contingent allocations[22]

$$\{C_{j,t}, L_{j,t}, D_{j,t}, X_{j,t}, K_{j,t}, C_{j,t}^h, C_{j,t}^f, X_{j,t}^h, X_{j,t}^f, Y_{j,t}, C_{j,t}^{h*}\}_{t=0}^{\infty}$$

and prices

$$\{p_{j,t}^c, p_{j,t}^x, R_{j,t}^*, w_{j,t}, r_{j,t}^k\}_{t=0}^{\infty}$$

such that, for all j's EMEs in the sample, given the laws of motion of shocks:

1. The allocations solve the consumer's problem given prices and the laws of motion for the capital stocks.

2. The allocations solve the firm's problem given prices.

3. Markets clear for capital, labor, investment goods, and home goods.

H. A Commodity Price Shock: Inspecting the Mechanism

What happens in economy j after it receives an income shock following an exogenous increase in the price of the commodity good that it exports? A full-blown quantitative answer to this question will be given in the following sections when we take the model to data from emerging economies. At this point, however, it is instructive to describe, in a qualitative manner, the main mechanism at work.

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[20] Strictly speaking, the processes for R^* and Y^* can also be considered two additional common factors. We prefer not to label them as such because they will be considered observable processes in the estimation of the model, unlike f^{Co} which will be treated as a latent variable.

[21] Note that we are not allowing for any comovement between the common factor in commodity prices and the other external shocks. This will be relaxed later in an extension to this setup.

[22] The set of equilibrium allocations and prices that characterize the solution of the model is found by applying a first-order Taylor approximation to the set of equilibrium conditions. The log-linearized system is then solved using perturbation methods in DYNARE. The Appendix presents the list of stationary equations as well as further technical details of the solution.

Because the commodity sector is an endowment, real economic activity will only be affected in this economy insofar as the market of home goods is affected. Looking at a (log-linearized version) of the market clearing condition for this market:

$$\widehat{y}_t = \underbrace{\left(C^h/Y\right)\widehat{c}_t^h}_{Effect\,1} + \underbrace{\left(X^h/Y\right)\widehat{x}_t^h}_{Effect\,2} + \underbrace{\left(C^{h*}/Y\right)\widehat{c}_t^{h*}}_{Effect\,3},$$

where lower-case variables are log deviations from their steady state levels ($x_t = \ln\left(X_t/X\right)$), reveals that economic activity will be affected by three distinct effects. A first effect is a demand channel by which households increase demand for home goods as an optimal response to the income shock, which will mainly depend on the intratemporal elasticity of substitution between h and f goods, η_c, the relative sizes of consumption and the commodity sector in steady state which defines the magnitude of the revenue shock. Such a demand channel will have real effects to the extent that it will boost the relative price of these goods, p^h, making it optimal for home good producers to increase production. This effect is illustrated in panel a. of Figure 4 by an outward shift in the demand curve for consumption of h-goods that takes the new equilibrium from point A to B.

A second effect operates via the market of new investment goods. The increased production of domestic goods raises the rental price of capital (panel b.) and pushes the demand for new investment goods by households increasing their price (panel c.). This generates a further outward shift in the demand curve for home goods from point B to C (panel a.). Here, again, the strength of this channel will mainly depend on the intratemporal elasticity of substitution between h and f goods in the production of investment goods, η_x, the relative sizes of investment in steady state, and the magnitude of the revenue shock.

A third and final effect comes from the fall in the demand for home goods by the rest of the world. This negative effect comes from the fact that these goods become relatively more expensive for foreigners. Its strength will mainly be a function of the price elasticity of foreign demand, $\epsilon_{j,e}$, and the relative size of (non-commodity) exports in steady state. Assuming that such a shock generates an inward shift in the demand curve for home goods that is not enough to counterbalance the previous two effects, as depicted in panel a. from points C to D, the net effect on real economic activity of the commodity price shock will be positive. In that case, the labor market will accommodate an increase in labor demand from home good producers (panel d).[23] This assumption will be later verified when the model is taken to the data. Importantly, note that this new equilibrium in the model will be one where the real exchange rate will appreciate, driven by the increase in the relative price of home goods, p^h. Thus, periods of booming commodity prices are characterized as episodes of real exchange rate appreciation. Note, however, that the opposite happens with productivity shocks. Indeed, a positive productivity shock in the home goods sector would bring about an outward shift in the supply curve, decreasing p^h, and causing a real depreciation.[24]

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[23] The reader may wonder about the shift in the labor supply curve depicted in Figure 4, panel d, given that GHH preferences are used which, in principle, eliminate the income effect. That indeed the case in one-good models. In models with multiple goods, however, there is another income effect that appears from the relative price shifts. In this particular case foreign goods become relatively cheaper, inducing an inward shift in the labor supply curve. The log-linearized version of the model presented in the Appendix illustrates this.

[24] Another possible offsetting mechanism which may dampen the size of the income shock triggered by the increase in commodity prices is an increase in the price of consumption goods. If the latter increases more than nominal GDP the income effect will be offset and real income, GDP/p^c, could actually fall. As will be shown when we evaluate the quantitative implication of the model, this will not be the case and a positive commodity shock will be associated to an

The next couple of sections will take this model to the data of emerging economies with the main goal of quantitatively assessing the strength of this mechanism by which commodity price shocks lead to movements in real economic activity and compare it with alternative driving forces of the business cycle in these economies.

IV. Taking the Model to the Data: Preliminaries

A. A Representative Group of EMEs

From the pool of 13 countries studied in Section 2, we pick a sample of four EMEs to estimate the model: Brazil, Chile, Colombia and Peru.[25] We do so mainly for two reasons. First, we want to have a reduced number of countries for tractability, while still keeping a pool that is both large enough and with sufficient heterogeneity in the production of commodities to pin down the common factors in the model. Second, as will be shown below, the calibration of the model's steady state presents some further data requirements that are binding for several of the 13 countries studied in Section 2.

This pool of four economies is, nonetheless, representative of the type of economies modeled in our theoretical framework. The four countries are all well-known commodity exporters, with a median commodity export share of 35.4.[26] In addition, as depicted in Figure 2, this group of countries reproduces well the average cyclicality documented in the larger pool of 13 EMEs.[27]

There is also strong evidence in favor of common factors affecting the macro dynamics in these four countries. The first principal component explains 81 and 67 percent of the variance in the commodity prices indices and real output across the four economies, respectively (see Appendix). Such numbers are even more supportive of the kind of common factors embedded in our model given that the commodity exporting profiles are fairly heterogenous and trade linkages are small across these four economies. The Appendix presents evidence on the average commodity export shares for these four countries. Brazil's largest shares are soybeans and iron; Chile's top two are copper and fish; Colombia's two are oil and coal; and in Peru they are gold and copper. This little overlap between the commodities that these four countries export implies that the common factor in commodities does not come mechanically from these countries exporting the same type of commodity goods. Data on the low trade linkages between the four economies and their other trading partners is also presented. For example, Brazil's exports to Chile, Colombia and Peru are, respectively, 0.2, 0.1, and 0.1 percent of output. In contrast, trade with its main trading partners is nearly twenty times that share (10.1 percent). Similar orders of magnitudes are observed for the other three countries.

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increase in real income.

[25]The four countries account for roughly half of Latin American output.

[26]Specific commodity export shares are: Brazil (17.9); Chile (69.7); Colombia (28.6); and Peru (60.5) (see appendix).

[27]The Appendix presents evidence of the cyclical movement in commodity prices and spreads for these four countries before and after the Great Recession, along the lines of Figure 1. It is shown how, as in the bigger pool of 13 EMEs, these four countries experienced expansions in output above trend in the period before the Lehman collapse, which reversed completely afterwards.

B. Calibrating the Steady State of the Model

When assigning values to the parameters in the model, we follow a strategy that uses both calibration and formal estimation methods. A subset of the parameters in the model, in particular those that determine the steady state of the model, are either taken from previous studies or calibrated so as to match certain long-run ratios in the data. The latter is an important prerequisite to disciplining the quantitative exercise.

The calibrated parameters are summarized in Table 3. Table 4 presents the long-run ratios from the data used in the calibration and the model's performance in matching them. We assume that, across all households in the economies considered, the risk aversion coefficient, σ, equals 2 and the Frisch elasticity, $1/\gamma_c$, equals 1.72, in line with what is usually found in the business cycle literature. We assume an annual depreciation rate of capital, δ, of 10 percent across all countries, and calibrate the share of capital in the production function, α, to match the consumption and investment ratios to GDP in each country. Following Adolfson et al. (2007) we set the price elasticity of exports, $\epsilon_{j,e}$, to 1.18. The elasticities of substitution in the consumption and investment bundles, η_c and η_x, are set to 0.43 consistent with the cross country evidence of Akinci (2011) for a sample of developing economies. The parameters α_c and α_x are calibrated so that the steady state of the model reproduces the average import shares of consumption and investment goods in the data presented in Table 4. This disciplines the model as it assigns realistic relative weight to f and h goods.

The parameters $\overline{D}$ and $\bar{Co}$ are calibrated so as to match the long run shares of external debt and commodity exports. This is crucial as it disciplines the calibration of the model so that it delivers realistic income effects from shocks to interest rates and commodity price shocks. The share of non-commodity exports to GDP adjusts to satisfy the balance of payments identity. We assume that the steady state (gross annual) real world interest rate, $\overline{R^*}$, is 1.01 and we calibrate $\overline{z^R}$ to match the long-run value of the country risk premium for each economy. The discount factor, β, for each economy is pinned down as the inverse of the country's real interest rate. We calibrate the scale parameter in the labor supply, ψ^c, so that labor in the steady state is set to 0.3. We normalize to 1 the steady state levels of productivity and foreign demand, $\bar{z}_j$ and $\overline{Y^*}$. Finally, we set the debt elastic parameter Ω_u to a minimum level, 0.001, that guarantees that debt is stationary.

The performance of the model in matching the key long-run shares is quite acceptable. By and large, the most important of all long-run ratios is the relative size of the commodity sector, which the model matches by construction in all four countries. As mentioned above, this guarantees that the size of the commodity price shocks will have realistic revenue implications for the countries considered. Likewise, for the propagation mechanism of a commodity price shock to be properly quantified, it is also of paramount importance that the model matches the relative size of home/foreign components in the total flows of consumption and investment of the countries considered,[28] as well as the relative size of total investment. As presented in Table 3, these dimensions are also among those that the model can perfectly match. Likewise, for the possible financial channel to be properly disciplined we make sure that the model reproduces the historical interest rate premia faced by these four countries as well as their historic debt-to-output ratio observed. The model cannot perfectly match the

[28]Lack of precise numbers for these relative shares of home/foreign goods in total consumption and investment was one stringent data limitation that prevented us from extending the formal quantitative analysis to other EMEs in our sample.

relative size of consumption, imports or exports. It nonetheless accomplishes a pretty satisfactory fit along those lines, as documented in Table 3,[29]

C. Bayesian Estimation

The remaining subset of parameters are estimated using Bayesian techniques. They govern the short-run dynamics of the model, but not its steady state. Namely, the persistence of the six driving forces $\{\phi_{Co}, \rho_j^{Co}, \rho_j^{z^R}, \rho_{R^*}, \rho_{Y^*}, \rho_j^z\}$, the standard deviation of their shocks $\{\sigma^{f^{Co}}, \sigma_j^{Co}, \sigma_j^R, \sigma^{z^R}, \sigma^{Y^*}, \sigma_j^z\}$; the loading factors in the dynamic common factor $\{\omega_j^{Co}\}$,[30] and those that determine the cost of adjusting the capital stock $\{a_j\}$.

The estimation uses as observables quarterly time series data on seven variables from each of the four countries considered which have a direct mapping onto the variables in the structural model: real private consumption, real income, real investment, the trade balance to GDP ratio, the EMBIG spread, the real effective exchange rate, and the commodity price index described in Section 2. Two additional observables are the 3-month real US TBills rate, and the United States's real GDP, as proxies for the world interest rate and foreign aggregate demand, respectively. Thus, in total, the estimation uses 30 quarterly time series that are mapped onto the model's counterpart through measurement equations.[31] We add measurement errors in consumption, investment, the trade balance share, and the real effective exchange rate as they are the only variables that do not have a shock directly linked to them. The model is estimated on a balanced panel that covers the period 2000.Q1 to 2014.Q3.[32] In the measurement equations the data are expressed in log-deviations from the Hodrick-Prescott trend and are measured in percent. Interest rates and EMBI are measured in logs of gross rates. As mentioned earlier, Kohli (2004) demonstrated how real GDP may underestimate the increase in real domestic income following terms of trade movements. We thus follow Kehoe and Ruhl (2008) by using real income instead of real GDP in the set of observables.

Fairly agnostic priors for the Bayesian estimation are used. We assume a Beta distribution with mean 0.5 and standard deviation 0.15 for the autoregressive coefficients in all six driving forces, implying a prior highest density region at 90 percent probability between 0.25 and 0.74. The prior distribution for a_j is Gamma with mean 0.5 and standard deviation 0.25, implying a large range for possible values of the parameter that governs the capital adjustment costs. Finally, we choose a rather conservative approach and center the prior distribution for ω_j^{Co} around zero (i.e., no common factor) using a normal distribution with standard deviation 0.015. In this case, the 90 percent bounds of the prior higher density region are $(-0.025, 0.025)$.

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[29] Some readers may object the fact that we do not incorporate a tradable/non-tradable setup in our modeling strategy. Such framework would place even more stringent data requirements in terms of precise time series data on production in both the tradable (commodity and non-commodity) and non-tradable goods when carrying the estimation, which may not be fulfilled in practice. Still we believe that the calibration of the model, as depicted in Table 3, retains the flavor of a non-tradable sector given that, in the steady state, the overwhelming majority of consumption and investment is done in home goods (on average 85.1 and 62.4, respectively, across the four countries).

[30] As is standard in the literature, and without loss of generality, we normalize to 1 the loading factor associated with f_t^{Co} in one of the N economies in the model (Colombia, in the empirical application) to identify the sign and scale of the common factor.

[31] The Appendix presents the set of measurement equations.

[32] Lack of data availability on some of the observables prevents us from covering a longer historical period.

V. ESTIMATION RESULTS

A. Posterior Distributions and Common Factors

Posterior distributions of the estimated parameters are reported in Table 4.[33] Overall the posterior densities are considerably different from the loose priors that we choose, implying that the data contain information on the estimated parameters. In terms of the parameters associated with the driving forces, the most salient result comes from the somewhat persistent and highly volatile commodity shock processes estimated. The AR(1) coefficient in the common factor of commodity prices, ϕ_{Co}, has a posterior mean of 0.73. Moreover, shocks to this common factor display an estimated standard deviation, $\sigma^{f^{Co}}$, of 6.31 percent, at least an order of magnitude larger than those of the other global shocks, implying an important role for common factors in commodity prices. The AR(1) parameters associated with the idiosyncratic components in commodity prices, ρ_j^{Co}, are also high, in most cases higher than those for the other idiosyncratic forces. It is also worth noticing that the presence of commodity prices reduces both the persistence of the (country-specific) productivity processes and the size of their shocks relative to previous business cycle studies. The posterior mean estimates of ρ_z range between 0.39/$0.40/0.51$ in Brazil/Chile/Colombia to 0.68 for Colombia; while the estimated standard deviation σ^z ranges between 0.84/0.89 percent in Peru/Colombia to 1.32/3.39 for Brazil/Chile.

Another salient result comes from the estimated coefficient of the loading factor, ω_j^{Co}, which captures the degree of transmission from movements in the common factor to the country-specific commodity price index, relative to the normalized value chosen for Colombia (equal to one). Chile displays the strongest response in relative terms, with the highest value for the posterior mode of ω_j^{Co}, equal to 1.53. Peru and Brazil, in contrast, display relatively less resposnes with posterior modes for ω_j^{Co} equal to 0.89, and 0.53, respectively.

Another salient result comes from the estimated coefficient of the loading factor, ω_j^{Co}, which captures the degree of transmission from movements in the common factor to the country specific commodity prices. Relative to the normalized (to one) value chosen for Colombia, Chile exhibits the highest value for this parameter with a mode of 1.53, followed by Peru (0.89) and Brazil (0.53).

Lastly, Figure 5 presents the time series evolution of the common factor, f^{Co}, which we back out from the Kalman filter smoothing evaluated at the posterior mean, together with the 90 percent distribution bands. It is tightly estimated and displays large and short-lived deviations from trend, particularly in the second half of the sample, post 2005, reaching sometimes near 30 percentages points –hence the allusion of a roller coaster ride that we make in the title of this paper. The increase in the years preceding the financial crisis was followed by a sharp fall during the crisis, and then a vigorous recovery within the next two years, only to fall once more in the last couple of years of the estimated period.

[33]The Appendix reports prior and posterior plots, including those of the measurement errors used in the estimation, which we do not report here for the sake of space.

B. Business Cycle Drivers

We now use the estimated model to document the main business cycle drivers. Our first tool to accomplish this is the forecast error variance decomposition (FEVD) of real income across the four countries and at various forecast horizons, summarized in Table 5. Statistics are reported using posterior means. The upper left panel of the table presents the unconditional contribution of each of the six structural shocks in each country. The other three panels report FEVD conditioning the exercise at alternative forecast horizons of one, four, and twelve quarters. Commodity price shocks play a large role, only comparable to that of productivity shocks. Their share in the unconditional FEVD of income displays a median of 42 percent, ranging from 27.5 percent in Brazil, up to 77.1 percent in Chile, with Colombia (43.5) and Peru (40.4) in between the two. Moreover, in all four countries a considerable amount of this share is related to the common factor in commodity prices. The remaining three external shocks to foreign demand, the world interest rate and spreads do not account for a large share of output's unconditional FEVD, with the exception of Brazil where interest rate shocks do play a role. The latter is related to the considerably large stock of external debt that Brazil has, relative to that in the other countries (see Table 3 again).

The role of commodity price shocks is positively related to the forecasting horizon in the FEVD decomposition exercise, as documented in the remaining panels of Table 5. When the FEVD is computed at only one quarter forecast horizon (upper right panel in Table 5) the share of output's FEVD accounted by commodity price shocks decreases considerably, between $1/10$ and $1/2$ to the one computed unconditionally. The domestic productivity shock accounts for the highest share, while the contribution of the remaining four shocks continues to be very small. This is related to the strong persistence in commodity prices, as documented earlier.

In order to gauge when, historically, commodity price shocks have contributed the most to business cycles in the four countries considered, we decompose the observed time series of output into the structural shocks of the model. For expositional purposes, the results, reported in Figure 6, group the six shocks into four groups: (i) *Productivity Shocks* captures domestic TFP shocks (ε^z); (ii) *Commodity Shocks* includes both common factor and country-specific shocks to commodity prices ($\varepsilon^{f^{Co}}$, ν_j^{Co}); (iii) *Spread Shocks* which alludes to the idiosyncratic risk premia shock ($\nu_{j,t}^R$); and *Foreign Shocks* that puts together the world's demand and riskless interest rate shocks ($\varepsilon^{Y^*}, \varepsilon^{R^*}$). A common feature across the four decompositions is the preponderant role of only two groups: *Commodity* and *Productivity Shocks*, in line with the FEVD results above. The influence of the former group is more marked in the second half of the sample, with Chile being the exception where its influence is large throughout all the sample. This coincides with the period where the common factor exhibits the largest fluctuations, as discussed earlier in the context of Figure 5. The contribution is more marked in the two-year boom that preceded the world financial crisis, which abruptly turned negative during the 2009 recession when, again, these shocks' contribution was large. Remarkably, these shocks turned positive in the recovery of the 2009 recession and helped all the countries recover from the recession. This helped counterbalance the negative contributions of domestic shocks, mostly in Brazil and Chile.

Table 6 presents further evidence on the role played by the structural shocks in the model from a historical perspective by running various counterfactual experiments on the historical decompositions. Starting from the second row, the table reports the standard deviation (S.D) of real income predicted by the model if one or several shocks are counterfactually turned off. The numbers are reported as percentage of the observed S.D reported in the first row of the table. One of the most salient findings is that the counterfactual S.D. falls considerably relative to the observed one when the two commodity price shocks are turned off (second to last row in Table 6). A fall that is only compared to the one observed when productivity shocks are shut down. In Peru, for instance, the counterfactual S.D when commodity price shocks are turned off is 69.6 percent of the observed one, followed by Chile (72.1), Colombia (75.4), and, to a lesser extent, Brazil (91.1). Similar qualitative results pointing to the preponderant role of commodity price shocks are obtained when all shocks are turned off except the two commodity price shocks (last row in Table 6). Chile stands out from the other countries where the counterfactual volatility actually increases above the observed one, implying that in this country other shocks have helped smooth out the massive volatility coming from commodity prices.[34]

C. Quantifying the Mechanism

Earlier, when presenting the model, we gave a qualitative description of the mechanism in place following a commodity price shock. We now provide a quantification of this mechanism using the estimated model's impulse response function (IRF) following an unexpected 1 S.D. common shock in commodity prices, which equates to assuming that the common factor increases by 6.3 percentage points above its long run mean. The IRFs of several of the variables in the model for all four countries are reported in Figure 7. Numbers reported are percentage deviations from steady state values.

Chile stands out as the country where more real effects following the shock are observed. This is consistent with the results from Tables 5 and 6 that showed the common factor to be particularly relevant for this country. Nonetheless, the other three countries exhibit non-trivial real effects as well.

The two upper panels in Figure 7 display the IRFs of two proxies of real economic activity: on the left we have plotted the quantities of the home good produced (Y_t), and on the right our proxy for real income as defined above, GDP_t/p_t^c . The production of home goods follows a hump-shape behavior, peaking between one tenth and a little over a half of a percentage point. The income shock triggered by the movement in commodity prices manifests as a rise of real income, which considerably increases on impact, ranging from 3 to 0.5 percentage points across all four countries and then steadily decreases.

The following panels in Figure 7 quantify the three effects that were described before. First, total private consumption increases, mainly driven by the boost in consumption of home goods. The latter increases between half and one full percentage point relative to its steady state. It is also noteworthy the strong persistence in the response of consumption, which is inherited from the long memory that commodity shocks display, as already mentioned.

The second effect, the response of investment, is also pronounced and hump-shaped. At its peak, approximately 4 to 5 quarters after the shock, it deviates between 0.5 and 3.5 percentage points from its long run value. As argued earlier, higher investment is driven to a large extent by the increase in the rental price of capital as home good producers demand more capital goods. The real rental return increases on impact between 10 and 100 basis points.

[34]One could conjecture that this result is capturing, albeit in a reduced form, countercylical policies being implemented. Further discussion on the role of fiscal policy is presented below.

The third effect materializes as a fall in the non-commodity exports that ranges between one half and three percentage points. This is linked to the fact that the equilibrium relative price of home goods increases substantially, leading to an appreciation of the real exchange rate across countries between half and two percentage points. The contractionary effect of the fall in the non-commodities exports is, however, not large enough to offset the expansion driven by the first two effects. Thus home good production expands. In the process, the labor market expands as labor demand increases, driving up real wages.

A last feature that stands out in the set of IRFs is that the common factor shock has the ability to generate sufficient comovement across countries. We now turn to a more formal evaluation of the model's performance.

D. Model's Performance

Table 7 presents the cross correlation of real income across the four countries predicted by the model (numbers below the main diagonal) and compares it to the one observed in the data (numbers above). Two model-based correlations are presented: the one coming from the benchmark model with a common factor in commodity prices ("CF"), and another one coming from an alternative model where we turn off the common factor in commodity prices ("NCF"). To be precise, in the latter case we set $\omega_j^{Co} = 0 \; \forall j$ and reestimate the model.

The benchmark model does a good job when bringing the model closer to the observed cross-correlations of output in the data. The average cross correlation in the data is 0.45 while the one predicted from the benchmark model is 0.32. This is, to a very large extent, the work of the common factor in commodity prices. In the alternative model, where the only common drivers are world interest rates (R^*) and external demand (Y^*), the average cross correlation falls to virtually zero (0.02). This is corroborated by the considerable fall of the alternative model's marginal likelihood relative to that of the benchmark model, reported in the bottom of Table 7.

Figure 8 presents further evidence on the model's performance by focusing on the serial correlation of macro variables with commodity price indices within countries. Again, data and benchmark model-based statistics are compared. The model is capable of reproducing the procyclicality of commodity prices relative to income, consumption and investment processes. It also captures the trade balance to GDP process appropriately, and the countercyclicality of real exchange rates. Overall the fit is pretty acceptable considering that these moments were not used as specific targets in the estimation.

Finally, Table 8 documents the model's performance along other second moments that are more standard in the literature and compares them to the data. The first two columns document the volatility of the main variables. The model does a proper job in accounting for the absolute and relative (to income) volatility of all the variables: a volatility of commodity prices that is close to four times that of income; that of investment and the real exchange rate that is nearly twice that of income; and that of consumption and the trade balance share that is less than that of income. The model also replicates in a satisfactory way the moderately high serial autocorrelation of the series as well as their comovement with real income. In particular, it reproduces the strong positive comovement of commodity prices, investment and consumption with income, and the negative one observed for the real exchange rate.[35]

VI. Robustness and Extensions

A. Comparison with a SVAR

As a first robustness, we compare the results of our DSGE model in terms of the relevance of commodity price shocks as drivers of business cycles in emerging economies against those from a SVAR. As suggested in a recent contribution by Schmitt-Grohé and Uribe (2015), comparing SVAR and DSGE may be a good way to assess the performance of the DSGE when assigning a role to terms of trade fluctuations when explaining business cycles given that usually SVARs are believed to impose less structure in the data than DSGE do. Hence they ”let the data speak” more freely. Their approach, however, does not look particularly at commodity price shocks as ours do.

The strategy we follow is simple. We first estimate the SVAR in Schmitt-Grohé and Uribe (2015) with data for the same four economies on which we estimate the DSGE model, except that we replace terms of trade in their SVAR with our measure of country-specific commodity price indices. Then we compare the results of each of the two models, SVAR and DSGE, when it comes to gauging the relevance of commodity price shocks. Formally, for each country, we estimate the VAR model:

$$x_t = \mathbf{A}x_{t-1} + u_t$$

and $x_t = [p_t^{Co}, tb_t, y_t, c_t, i_t, rer_t]$, where p_t^{Co} is the country-specific commodity price that we build, tb_t is the trade balance share, y_t is real income, c_t is real private consumption, i_t is real investment, and rer_t is the real exchange rate; u_t is a 6x1 vector with mean zero and variance-covariance matrix Σ, and $\mathbf{A}$ is a 6x6 matrix of coefficients. Additionally, we impose the same two identifying restrictions in Schmitt-Grohé and Uribe (2015). First, we assume that p_t^{Co} is exogenous and that it follows an AR(1) process, thus the off-diagonal elements of the first row of $\mathbf{A}$ are zero, $a_{1j} = 0$ for $j = 2, ..., 6$; where a_{1j} denotes the element (i, j) of the matrix $\mathbf{A}$. Second, we assume that $u_t = \Pi\epsilon_t$, where Π is a 6x6 matrix and ϵ_t is a 6x1 vector of white noise mean zero and identity variance-covariance matrix. Thus, $\Sigma = \Pi\Pi'$. Also following this work, we pick Π to be the lower-triangular Cholesky decomposition of Σ. The SVAR is estimated with filtered quarterly data for each of the four countries on which we estimate the DSGE model, in the same sample period.

The results from the comparison between the two models are gathered in Figure 9 and Table 9. Figure 9 plots the time series of y_t in the data against the predicted series for this variable from the DSGE and the SVAR models. The predicted time series for each model are obtained by recovering the history of structural shocks in commodity prices and simulating both models using only these shocks. Thus, they are the predicted time series by the two models if only commodity shocks had

[35] When constructing Table 8, we use as real income the nominal GDP to CPI to be consistent with the measure of income used in the estimation (see the earlier discussion on the Bayesian estimation). The Appendix contains, nonetheless, similar statistics obtained using real GDP, deflated with the GDP deflator. The results are qualitatively robust to this change.

been realized. Table 9 presents some further comparison between the two models in terms of the unconditional variance decomposition of output explained by shocks to p_t^{Co}, among others.

The main result coming out of this evidence is that, while not identical, the implications of the SVAR are aligned with those from the DSGE in that they also point to a considerable role of commodity price shocks. In the SVAR, the mean/median share of shocks to p_t^{Co} in the unconditional variance of output across the four countries is $49/50$ percent, while that from the DSGE is $47/42$ percent. Some differences remain nonetheless. The predicted share in Brazil (15) and Chile (53) by the SVAR are lower than their DSGE counterparts, (28 and 77, respectively), while in Colombia they are pretty much aligned (47 and 44, respectively), and in Peru the SVAR predicted share is 80 percent, twice that of the DSGE.

The resemblance between the two models increases when the predicted time series from the two models using only p_t^{Co} shocks are compared, as can be seen from Figure 9, particularly during the second half of the sample. The median correlation between the two predicted series is equal to or above 0.9 for all countries but Chile (panel E in Table 9). In the latter case, however, the DSGE model tracks much more closely the observed data than the SVAR (panel D). Furthermore, the mean standard deviation of the simulated time series from the SVAR, relative to that in the data, is 67 percent while that of the DSGE model is 62 percent (panel B). Lastly, we also conduct the opposite experiment, where all shocks but those from commodities are used to simulated both models (panel C). In this case, again, both models continue to yield very close results.

Overall, the results point to the SVAR being consistent with our benchmark results derived from the DSGE in that commodity prices account for a large part of output fluctuations in EMEs. This contrasts with the results from Schmitt-Grohé and Uribe (2015) who find that only a median share of 10 percent of output's variance can be associated with terms of trade shocks. While there could be several reasons why our results differ from theirs, we think a good place to start is the link between commodity prices and terms of trade. This clearly goes beyond the scope of this paper so we postpone it for future research.[36]

B. Countercyclical Spreads

To keep the analysis simple and the transmission mechanism of a commodity shock as tractable as possible, the benchmark model assumed the simplest structure for risk spreads as AR(1) processes (see Eq. 20). However, spreads may react to country fundamentals as well, and vice versa. As argued by the seminal works of Neumeyer and Perri (2005) and Uribe and Yue (2006), this may generate an amplification mechanism for real shocks. In this subsection we follow these works and extend the benchmark framework to allow explicitly for such interaction between country fundamentals and spreads. We do so by modifying the model in two dimensions. First, the risk premia process is now allowed to be affected by country fundamentals, in the form of TFP and commodity prices. Formally, Eq. (20), in log deviations, becomes now:

$$\hat{z}_{j,t}^{R} = \rho_j^{z^R} \hat{z}_{j,t-1}^{R} - \eta_j^{z} E_t \hat{z}_{j,t+1} - \eta_j^{Co} E_t \hat{p}_{j,t+1}^{Co} + \sigma_j^{z^R} \nu_{j,t}^{R}, , \quad \nu_{j,t}^{R} \sim N(0,1) \tag{23}$$

[36]We think that it is beyond the scope of this paper to replicate our SVAR results to the nearly 40 economies considered by Schmitt-Grohé and Uribe (2015). However, the Appendix contains further evidence that the share of output's variance associated with terms of trade shocks found in this work is positively related to the size of commodity exports in total exports.

where a hat, "$\hat{}$", denotes log deviations from steady state values; and $\{\eta_j^z, \eta_j^{Co}\}$ govern the degree of responsiveness of spreads to expected deviations of productivity and commodity price indices from their long-run values. The inclusion of TFP in (23) is not new and follows Neumeyer and Perri (2005). What is novel is the inclusion of (expected) changes in commodity prices. This captures the idea that commodity prices contain information on the creditworthiness of the borrower EME to the extent that they are a determinant of its repayment capacity. Thus expectations of *future* commodity prices may determine *current*'s risk premium.

Evidently, Eq. (23) models risk premia in a reduced form and is not derived from first principles. A more complete model of the determination of fluctuations in country risk and their interaction with commodity prices is beyond the scope of this paper because our main goal is to analyze the extent to which this interaction matters for business cycles. There is, nonetheless, both theoretical and historical evidence of this link. Calvo and Mendoza (2000) link volatility in financial conditions for EMEs in world markets to the cyclicality of their terms of trade and other fundamentals in the context of informational frictions where uninformed investors cannot extract information from prices but rather do so from noisy information about specialists' trades. In a historical context, Eichengreen (1996) documented that during the crash of 1929 the sharp drop in the price of Brazilian coffee led foreign bankers to stop extending loans to Brazilian borrowers. And Min et al. (2003) found that improved terms of trade are associated with lower yield spreads to the extent that such improvements imply an increase in export earnings and better repayment capacity.[37]

The second modification to the model aims at capturing the other direction of the linkage: from spreads to economic activity. We do so following the literature by introducing a working capital constraint that creates a direct supply-side effect of changes in spreads, and hence in interest rates, on the demand for labor by home-good producers. Formally, this is done by assuming that the profit function of the firm is now:

$$\xi_t = p_t^h Y_t - w_t \left[1 + \psi \left(R_t - 1\right)\right] L_t - r_t^k K_{t-1}$$

where ψ is the fraction of the wage bill that must be set aside in advance in order to produce.

We take this modified model to the data including the three new parameters per country, η_j^z, η_j^{Co}, and ψ_j, in the estimation. The posterior results are presented in the Appendix. We report the unconditional FEVD for output in the top-left corner of Table 10. The most interesting result is that in three of the four countries (Brazil, Colombia, and Peru) the share of variance associated with commodity prices raises relative to the benchmark case presented in the upper-left panel of Table 5. This is related to the fact that the estimated coefficients η_j^{Co} and ψ_j are significantly different from 0. The

[37] Cuadra and Sapriza (2006) link the volatility of terms of trade in EME to spreads in a dynamic model with strategic default that delivers endogenous default risk, but do not explore the implications for the business cycle. Using FAVAR models, Bastourre et al. (2012) have also documented a strong negative correlation between commodity prices and emerging market spreads. In the context of the subprime crisis, Caballero et al. (2008) argued that persistent global imbalances and the volatility in both financial and commodity prices (such as oil) and asset prices that followed the crisis stemmed from a global environment where sound and liquid financial assets are in scarce supply. Morana (2013) has recently found that financial shocks have had a much larger role as drivers of the price of oil than previously noted in the literature. Recently, González et al. (2015) and Beltran (2015) have related spreads to commodity prices using the financial accelerator model.

share in Chile is slightly reduced, but continues to be the largest of all countries. Overall, this provides evidence in favor of the theories that link commodity prices to spreads and see the latter as a propagation mechanism for commodity price fluctuations.[38]

C. Correlated External Driving Forces

In this subsection we relax the assumption made in the benchmark case that external driving forces are uncorrelated. In particular, we now allow for the common factor in commodity prices, f_t^{Co}, to be correlated with external demand, Y_t^*. The assumption of uncorrelated forces in the benchmark case is tenable insofar as the primitive shocks to the common factor in commodity prices are driven by, e.g., demand shocks that stem from one region, say China, while external demand for non-commodities produced in EMEs is coming from a separate region, say the United States. Still, one could also argue that demand shocks in the United States may also impact market equilibrium prices in commodities, thus calling for an approach that takes into account a correlation between these two forces.

Formally, we model this by modifying the stochastic processes for these two driving forces, (21) and (14), as

$$\ln Y_t^* = (1 - \rho_{Y^*}) \ln \overline{Y^*} + \rho_{Y^*} \ln Y_{t-1}^* + \nu_t^{Y^*}$$

$$f_t^{Co} = \phi_{Co} f_{t-1}^{Co} + \nu_t^{f^{Co}}$$

where $\nu_t^{Y^*}$ and $\nu_t^{f^{Co}}$ are assumed to be correlated as follows

$$\begin{bmatrix} \nu_t^{Y^*} \\ \nu_t^{f^{Co}} \end{bmatrix} = \begin{bmatrix} 1 & 0 \\ \sigma^{Co,Y^*} & 1 \end{bmatrix} \begin{bmatrix} \varepsilon_t^{Y^*} \\ \varepsilon_t^{f^{Co}} \end{bmatrix}$$

and $\{\varepsilon_t^{Y^*}, \varepsilon_t^{f^{Co}}\} \sim N(0, 1)$. This simple identification strategy assumes that primitive shocks to external demand may have a contemporaneous effect on commodity prices, governed by the parameter σ^{Co,Y^*}, but not vice versa. This parameter is included in the set of estimated parameters in the modified model.

We report the unconditional FEVD for output in the top-right corner of Table 10. Posterior results are presented in the Appendix. The main result is that the large share of commodity price shocks remains, relative to the benchmark case. There is indeed evidence that the two forces are correlated, i.e. σ^{Co,Y^*} is estimated to be positive and statistically significant (with a posterior mean of 0.021), and the share associated with external demand shocks increases to an average of 5.5 percent. While this is still a moderate share, it is nonetheless a considerable increase relative to the benchmark case where the average share of these shocks was only one tenth of a percentage point. Lastly, there is also an improvement in the modified model's overall fit judging by the increase in the marginal likelihood relative to the benchmark case.

[38] An earlier version of this work extended this model in order to account for a second common factor in spreads across EMEs. In that version, both factors interacted to affect country-specific spreads and commodity prices. Results were similar to those in this extension and are available upon request.

D. Absence of Commodity Shocks

We now explore what happens if commodity shocks are turned off in the benchmark estimation. This case is interesting as the reduced model becomes much similar to the benchmark small open economy RBC model. Here, again for the sake of space, we document only the FEVD of output in bottom left panel of Table 10 and the marginal likelihood of the reduced model and report in the Appendix the posterior results as well as other second moments generated by the model.[39]

Two results stand out. First, in three of the four countries the overwhelming share of output's variance is now explained by TFP shocks. The exception is Brazil, where the role of commodity shocks is not replaced by TFP shocks as in the other three countries, but by interest rate shocks. We believe these results reconcile our results with those coming from previous works on business cycles of emerging economies where TFP and interest rate shocks were at center stage.

The second result is the large fall in the model's performance, measured by its marginal likelihood, relative to the benchmark. It drops from 4530.8 to 4410.6. The Appendix reports the model's fit in terms of second moments. A noticeable failure of this reduced model is the fact that it fails to reproduce the negative comovement between the real exchange rate and real income observed in the data. Instead it counterfactually predicts that expansions (contractions) in economic activity are accompanied by real exchange rate depreciations (appreciations). As explained in the context of Figure 4, positive productivity shocks in the home goods sector would bring about an outward shift in the supply curve of home goods, decreasing p^h, and causing a real depreciation.

E. Alternative Dynamic Factor Models

A choice made in the estimation strategy of the benchmark model is the inclusion of the parameters of the dynamic factor structure $\{\phi_{Co}, \rho_j^{Co}, \sigma^{f^{Co}}, \sigma_j^{Co}, \omega_j^{Co}\}$ in the full set of parameters from the DSGE model that are estimated using full-information methods. As argued before, this choice allows the structural model to identify country-specific, commodity-related, revenue shocks via the idiosyncratic component of the commodity process. This choice, however, entails a tradeoff as it constrains the information on commodity prices used to only the sample period where the DSGE model is estimated, after the year 2000. One might also wonder how the estimation of the dynamic factor model would be if it were done outside the rigid structure imposed by the DSGE on the data.

In this subsection we undertake several extensions that relax this initial choice. First, we estimate the dynamic factor model (DFM) structure, (14) - (16), in isolation from the structural model, using only data from country-specific commodity price indices. This allows us to use data on commodity prices that goes as far back as 1980 when our time series on single commodity goods prices begin. For this case we also consider using the full range of 13 EMEs in our empirical analysis. We also

[39]To be precise, we estimate a separate and independent AR(1) process for commodity prices. But this process no longer perturbs the budget constraint of the household as in the benchmark case. Thus, the households commodity revenue is now deterministic and equal to $p^{Co}\bar{C}o$ in every period. These assumptions imply that the modified and benchmark models will have the same non-stochastic steady state. Importantly, they also allow the estimation of this reduced model to have the same set of observables as the benchmark model, thereby rendering the comparison of the marginal likelihood across the two models valid.

study the case of only the 4 countries studied in the DSGE model, starting the estimation in 1980 and 2000.[40]

The estimated common factor for each of these alternative cases, together with the benchmark case, that we reproduce for comparison, are reported in Figure 10. The most noticeable result that can be drawn from this figure is that all robustness cases track down pretty well the common factor estimated in our benchmark estimation. Another noticeable result is that the volatility exhibited in the post-2005 subperiod continues to stand out even when one looks at a longer historical period.

F. Alternative Filtering

In the empirical analysis of Section 2, as well as in the estimation of the structural model, we detrend the data using a Hodrick-Prescott filter. We do so in order to keep the analysis within a business cycle frequency. However, as is well known, results may differ depending on the filter used when detrending the data. In this subsection we assess the sensitivity of our results when using simple growth rates as alternative detrending method. For the sake of space, the Appendix reproduces the same set of stylized facts as those reported in Figure 2, and the bottom-right panel in Table 10 reports the results of the FEVD of output from the structural model when the observable data used in the estimation are on growth rates.[41]

The stylized facts remain strongly robust to the change in filtering methodology. Commodity price indices continue to correlate vigorously with the other macro variables considered. Likewise, the results of the FEVD derived from the structural model continue to point to a preponderant role of commodity price shocks as a key driving force of the business cycle in the economies considered. The only economy for which the share of variance drops relative to the benchmark case is Chile, although it continues to remain above half. In the remaining three countries the share associated with commodity price shocks increases relative to the benchmark results. The median share of output explained by commodity shocks is now 68.1 percent. Thus, if anything, the benchmark result in terms of the preponderant role of commodity price shocks for business cycles strengthens when considering an alternative data detrending technique.

G. Cyclical Production of Commodities: A Discussion

Arguably the assumption of an endowment flow of commodity goods in the model is not without loss of generality. Ideally a more complete model of an emerging economy would include production in this sector along other features associated to the economics of commodities like storage capacity, optimal management of inventories, foreign ownership of capital, monopolistic rents, specu-

[40]For consistency, the country-specific commodity price indices that start in 1980 continue to use the same commodity weights as those used in the benchmark case. Given that the annual weights do not systematically change much across years for most of the countries considered we think this is an innocuous assumption.

[41]The Appendix also presents the modified measurement equations used in the estimation together with the posterior results. When carrying the estimation, all observables are expressed in quarter-to-quarter growth rates and demeaned. This makes the marginal likelihood of this model not directly comparable to that in the benchmark model so we do not report it.

lation, etc.[42] Yet we believe that such realistic features would likely make the analytical framework less tractable, and would obscure the main propagation channel that we think is the most important from a business cycle perspective: the revenue shock that comes with movements in the world markets of commodity goods.

Nonetheless, we think that further discussion about the validity of this assumption is warranted. Thus, in this subsection we present the available evidence regarding the cyclicality of commodity production over the business cycle and discuss its implications. Figure 11, panels (a) to (h), report the serial correlation between the volumetric production of the two most important commodities in each of the four countries considered in the estimation of the structural model with the international market price of each commodity. The sample period over which the correlations are computed is the same as the one used when estimating the structural model. The results indicate no systematic comovement whatsoever between quantities and prices. Indeed, in most cases, commodity production does not react to changes in prices at business cycle frequencies.[43] A more formal analysis with the use of Granger-causality tests between the two variables reveals no causality in either direction (see Appendix). This not only provides evidence of the absence of a response of commodity production to changes in commodity prices, but also backs the assumption that these countries are price takers in international markets.

The remaining two subplots in Figure 11, (i) and (j), present further evidence related to the propagation mechanism of the revenue shock implied by the model. Subplot (i) reports the serial correlation between the simple average of the non-commodities GDP for the four countries studied in the structural estimation and their country-specific commodity price indices.[44] Subplot (j) looks instead at the nontraded GDP and its comovement with commodity prices across countries. Because constraints on data availability are less stringent for non-traded output, the latter is done for a wider pool of EMEs. Regardless of which of the two measures of real economic activity one looks at, there is a considerable comovement between them and country commodity price indices. This goes in line with the main prediction of the model's propagation mechanism following a commodity shock: that economic activity, mainly the one that is unrelated to the commodity sector, comoves with commodity prices, driven by changes in real income which, in turn, come from exogenous fluctuations in those prices.

H. The Role of Fiscal Policy: A Discussion

A potentially important agent that is absent from the theoretical framework is the government, which may be the recipient of a large part of the revenue from commodity exports, either because it owns

[42]For a model of commodity price dynamics with some of these features, mainly storage and speculation behavior, see Deaton and Laroque (1996).

[43]Data from the volumetric quantities of each of the two commodities is not reported in UN Comtrade. We use national sources for each commodity. The Appendix reports the same statistics presented in this subsection for the entire period for which data on volumetric quantities is available. Results are qualitatively robust.

[44]Non-commodity production for Brazil is computed as the difference between total GDP and GDP in agriculture and mining. In Chile, it is defined as total GDP excluding copper mining and fishing sectors. Non-commodity production in Colombia is computed as the total GDP minus GDP in coffee and mining (which inludes crude oil production). Finally, in Peru it is defined as total GDP excluding mining and hydrocarbon sectors.

a stake in the commodity sector, or because it collects taxes levied on this sector.[45] This, in turn, may bear importance for the source of the propagation mechanism following a commodity price shock. Indeed, it would imply that government expenditure may be an important part of the channel through which the real effect propagates.[46] More generally, this possibility opens the door for the conduct of fiscal policy to become a crucial element. If policy is conducted with a fiscal rule, this may help dampen commodity shocks. On the other hand, absent a fiscal rule, governments may amplify the business cycle effects of those shocks.[47]

We view the analysis, both positive and normative, of the role of fiscal policy in the face of commodity price fluctuations as an important topic that deserves a careful analysis, which goes well beyond the scope of this paper.[48] However, a simple first look at the data suggests that this mechanism does not have first order importance. This assertion is supported by the serial correlations reported in Figure 12 where, once more, we correlate country commodity price indices with economic activity across EMEs, except that this time we compare the results when such activity is measured with and without the flow of government expenditure. The left panel reports the results for LAC4, and the center panel reports results for a wider pool of 12 EMEs. The right panel directly computes the correlations between real government expenditure and commodity prices in the two sets of countries. Real government is not positively correlated with commodity prices, and real economic activity, if anything, marginally *increases* its comovement with commodity prices when public expenditure is removed.[49]

It may be argued that the measures of government expenditure used above do not incorporate direct transfers from the government to households and that these can be an alternative mechanism for channeling resources from the commodity sector.[50] In this case, however, our model would be

[45] For the case of the 4 economies considered in the structural estimation the role of the government may mostly come from taxes, as it does not participate actively in the ownership of the companies in the main commodity exports. In Brazil, the soybean industry is almost entirely privately owned, and the iron sector underwent a large-scale privatization in 1997. A similar story occured in Chile with the privatization of several copper mines to the point that private ownership is beyond 60 percent of production. Fishing, Chile's second commodity export sector, is privately owned. In Colombia, Ecopetrol, a mostly state-owned company, is responsible for only a third of oil production. And coal, the country's second commodity export, is private. Lastly, Peru also underwent a large-scale privatization of its mining industry.

[46] Evidently, changes in taxes may very well represent another channel.

[47] For the specific case of the four countries considered in the estimation the presence of fiscal rules, particulary balanced budget rules, is a relatively recent or non existent characteristic. Brazil does not have one; Chile started one in 2001 with statutory basis in 2006; Colombia implemented one only in 2011; and Peru has one since 2000 but only in 2005 converged to a stable debt ceiling (see Budina et al., 2012).

[48] See, for instance, Guerra-Salas (2014) or Caputo and Irrarazabal (2015) who incorporate a government sector that directly benefits from higher commodity prices. Agenor (2015) has recently investigated the possibility that governments conduct fiscal policy to optimally manage the commodity price shock, although with an emphasis on low-income economies.

[49] Further empirical evidence for a larger number of countries against the hypothesis that the reaction of government spending in a commodity-exporting economy amplifies the effect of commodity prices on the business cycles is presented Cespedes and Velasco (2013).

[50] There is evidence that during the previous decade, which coincided with the commodity boom period, large-scale cash transfer programs to poor households were put in place by goverments in these economies. Familias en Accion, and Bolsa de Familia are examples in Colombia and Brazil, respectively.

isomorphic to this case as households are the ultimate recipients of revenues from the commodity sector.

VII. Concluding Remarks

This paper has shed light on the nature and relative importance of external forces as drivers of aggregate fluctuations in emerging market economies with a special focus on commodity prices. It has involved both a careful study of the stylized facts in the data and an attempt to structurally identify these external forces by estimating a dynamic, stochastic, equilibrium model. We have found support for the view that these external forces are relevant and that their sources can mostly be traced back to exogenous changes in the prices of the commodity goods that these economies export which are viewed through the lens of the theory as large income shocks. A salient characteristic of these movements -often comparable to a wild roller coaster ride- is that they share a common factor. The latter cannot be solely attributed to these economies exporting similar commodity goods. Indeed, the common factor arises also because there is a marked tendency for the price of different commodity goods to move in tandem. Furthermore, the real effects generated by the fluctuations in the prices of these commodity goods can be amplified by the fact that they are often accompanied by movements in interest rates in opposite directions. Lastly, while most often movements in these relative prices have amplified the business cycle of EMEs, there are instances where they have served as cushion devices against other forces. This was the case during the recovery after the world financial crisis when a rapid reversal of commodity prices helped to counterbalance negative shocks of domestic and external sources.

The simplicity of the theoretical framework with which we have looked into the data has served us well for the kind of question that we set out to answer. However, its simplicity has also left aside important issues that are worth exploring in subsequent work. One important topic left aside is to try to uncover the role of government in the mechanism through which changes in commodity prices affect the real economy. Also worth exploring is the type of optimal fiscal and monetary policies that may be implemented to counteract the effect of those shocks.

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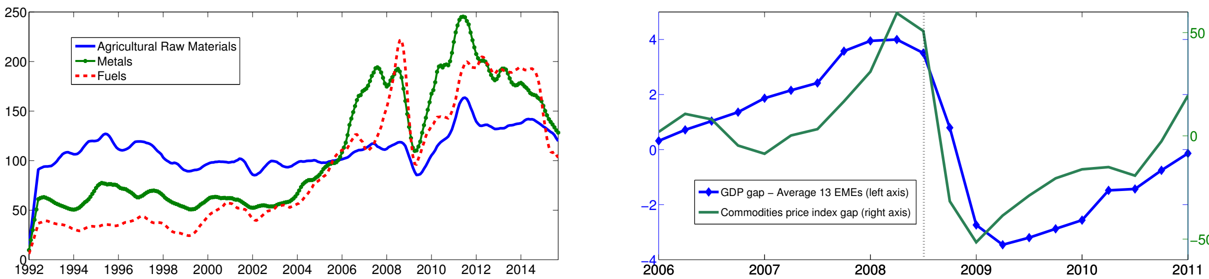
VIII. Figures and Tables

Figure 1: The Great Recession in Emerging Economies



(a) Real Commodity Price Indices

(b) Output Gap and Commodity Price Gap

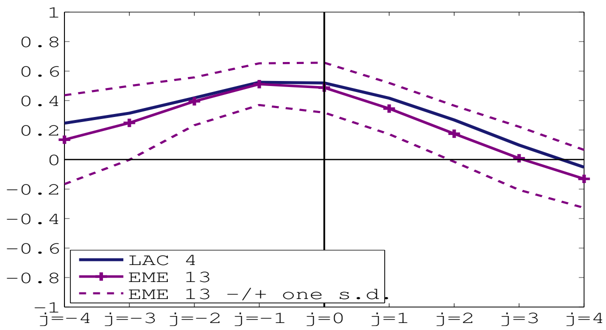
Note: Left panel plots the six months centered moving average of commodity prices indices for three separate commodities: agricultural materials (solid/blue), metals (diamonds/green) and fuels (dashed/red) with 2005 as the base year (100). Right panel plots the quarterly cyclical components of a consolidated price index for all commodities (solid/green) and the simple average of the quarterly real GDP gap across 13 emerging economies (diamond/blue). The dotted vertical line represents the quarter when Lehman declared bankruptcy. Cyclical components are computed with a Hodrick-Prescott filter with smoothing parameter of 1600. See text for the list of countries. Source: commodity prices come from IMF, real RDP are from Haver.

Table 1: Commodity Export Shares

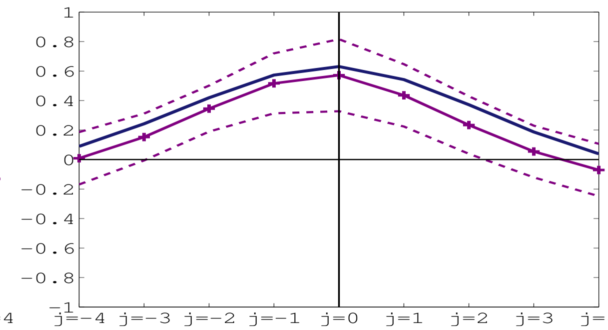
	All Countries	**Advanced**	**Emerging**	**Low Income**	**Non-Emerging**
Median (%)	19.3	11.2	25.7	26.5	14
Mann-Whitney Test (against Emerging)		3.046 *** (0.0023)		-0.689 (0.4906)	-2.309 ** (0.0209)
Kruskall-Wallis Test		8.048 ** (0.0179)			
No. of Countries	189	74	61	54	128
No. of observations (annual)	5514	2296	2205	1013	3309

Note: The first row reports the median commodity export share in total exports across various groups of countries (in percentages). Second and third rows report Mann-Whitney and Kruskal – Wallis tests statistics (p-values in parenthesis). Mann-Whitney tests are computed always relative to emerging economies. The Kruskal – Wallis test reports the results using three groups: advanced, emerging and low income. Data come from WDI in the form of an unbalanced annual panel of three kinds of commodity exports (agricultural, fuels, and metals). EMEs are classified as such if data on EMBI exist. Advanced and low-income are categorized according to World Bank's WDI indicators. The Appendix contains further information on the countries in each group and a description of the dataset. Statistical significance at 1, 5, and 10 percent is reported with (***), (**), and (*), respectively.

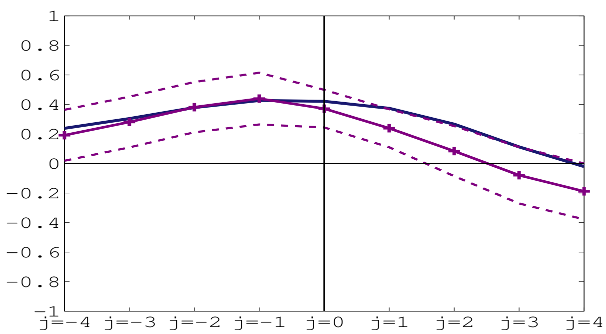
Figure 2: Cyclicality of Country-Specific Commodity Price Indices in Emerging Economies



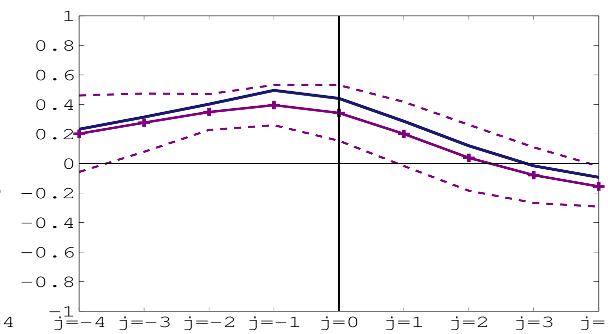
(a) Real GDP



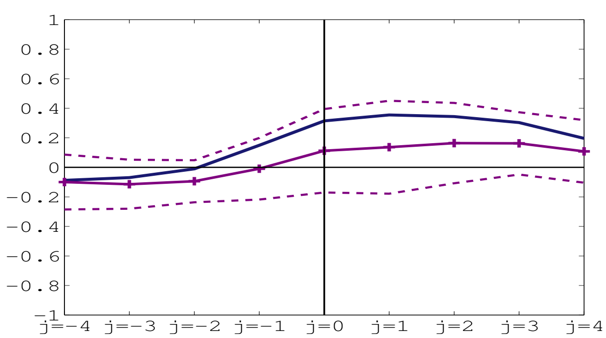
(b) Nominal GDP/CPI



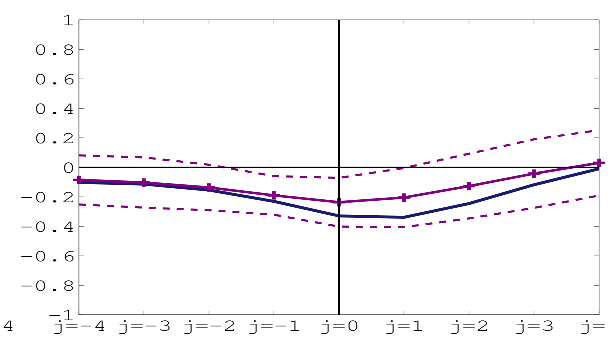
(c) Real Private Consumption



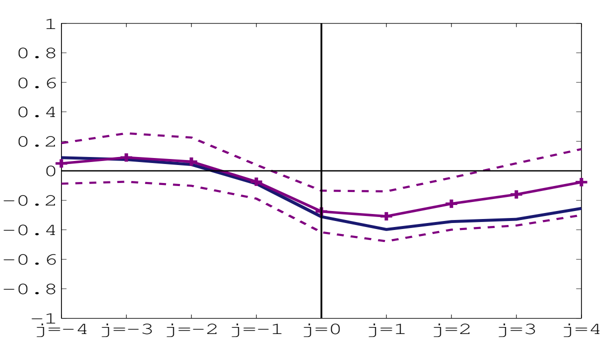
(d) Real Investment



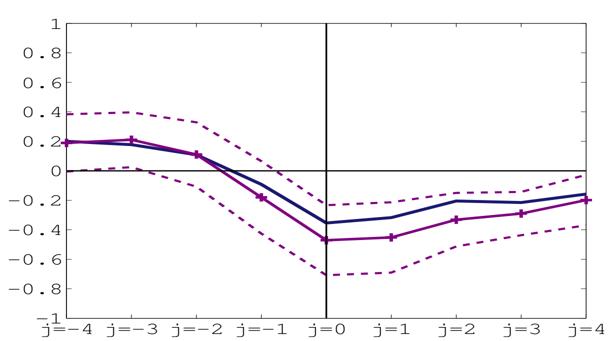
(e) Trade Balance/GDP



(f) Real Exchange Rate



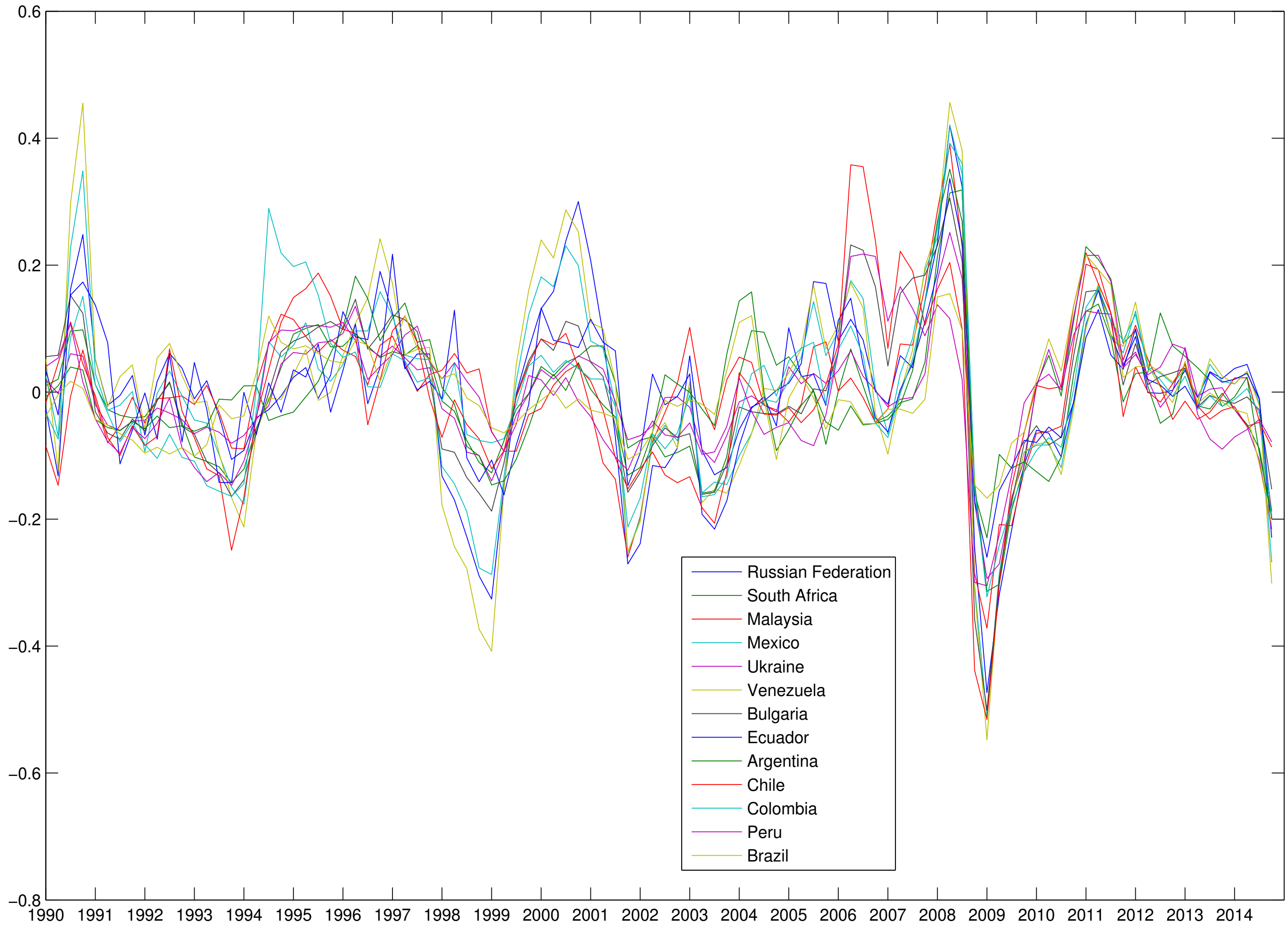
(g) EMBI



(h) CEMBI

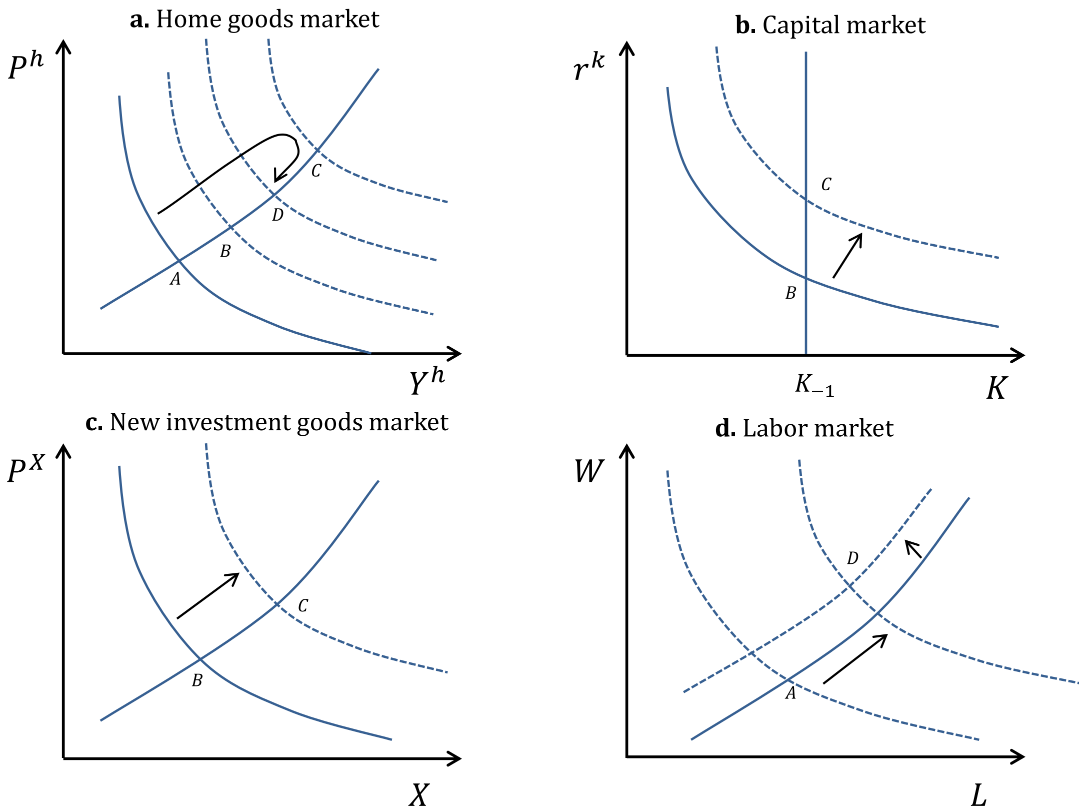
Note: Each panel reports in purple/solid with "+" lines the simple average correlation between the macroeconomic variable in the subtitle in period t and the country-specific commodity price indices in $t+j$ with $j = -4, \ldots, 4$, across 13 Emerging Market Economies (EME, see text for the complete list). Dashed/purple lines report plus/minus one S.D. bands. Blue/solid line report the results for a smaller sub-group of EMEs in Latin America (Brazil, Chile, Colombia, and Peru). All statistics are computed with the cyclical components obtained using a Hodrick-Prescott filter ($\lambda = 1600$) on a quarterly unbalanced panel between 1990.Q1 and 2014.Q4. See Appendix for the exact range of quarterly time series used and for country coverage.

Figure 3: Country-Specific Commodity Price Indices in Emerging Economies: 1990-2014



Note: The panel reports the quarterly time series for the cyclical components of country-specific commodity price indices across 13 EMEs. Cyclical components are computed as relative deviations from a Hodrick-Prescott trend (lambda of 1600). See Section 2 and the Appendix for the sources, and description of the construction of the indices.

Figure 4: The Propagation Mechanism of a Positive Commodity Price Shock in the Model



Note: Each panel depicts, qualitatively, movements in supply and demand curves in four markets in the model following a positive shock to commodity prices.

Table 2: Calibrated Parameters

Parameters		Brazil	Chile	Colombia	Peru
	Calibrated parameters common to all countries				
$\overline{R^{\star}}$	Gross steady state external annual interest rate	1.01	1.01	1.01	1.01
Ω_u	Interest rate elasticity of debt to GDP ratio	0.001	0.001	0.001	0.001
γ_c	Inverse of labor supply elasticity	0.58	0.58	0.58	0.58
σ	Relative risk aversion coefficient	2.00	2.00	2.00	2.00
η_c	Elasticity of substitution - consumption bundle	0.43	0.43	0.43	0.43
η_x	Elasticity of substitution - investment bundle	0.43	0.43	0.43	0.43
δ	Capital depreciation anual rate (%)	10.00	10.00	10.00	10.00
ϵ_e	Price elasticity of exports	1.18	1.18	1.18	1.18
$\overline{Y}^{*}$	Steady state aggregate demand of ROW	1.00	1.00	1.00	1.00
$\bar{z}$	Steady state productivity level	1.00	1.00	1.00	1.00
	Calibrated parameters to match long run relations				
$\bar{D}$	Steady state level of external debt	26.80	3.64	9.41	11.25
ψ^c	Scale parameter in labor supply	6.94	16.44	7.65	10.69
α_c	Governing import share in consumption	0.36	0.48	0.23	0.09
α_x	Governing import share in investment	0.13	0.58	0.60	0.82
α	Capital share in production	0.32	0.39	0.34	0.37
$\overline{p^{Co}}$	Steady state level of commodity price	0.85	3.16	0.37	1.36
$100\left(1-\overline{z^R}\right)$	Annual steady state spread (%)	5.39	1.45	3.9	3.56
β	Discount factor	$\frac{1}{\overline{R^{\star}}\,\overline{z^R}}$	$\frac{1}{\overline{R^{\star}}\,\overline{z^R}}$	$\frac{1}{\overline{R^{\star}}\,\overline{z^R}}$	$\frac{1}{\overline{R^{\star}}\,\overline{z^R}}$

Note: World interest rate $\overline{R^{\star}}$, country spread $\overline{z^R}$ and the depreciation rate δ are presented in annual terms. ROW stands for rest of the world.

Table 3: Long-Run Ratios: Model and Data

Long-run ratios (%)	Brazil		Chile		Colombia		Peru	
	Model	Data	Model	Data	Model	Data	Model	Data
Consumption / GDP	78.08	82.11	76.66	72.67	77.06	82.00	76.64	77.67
Investment / GDP	17.74	17.74	23.16	23.16	21.00	21.00	22.22	22.22
Imports / GDP	10.79	11.31	32.54	31.33	17.65	19.00	18.06	18.11
Exports / GDP	14.96	11.45	32.72	35.49	19.59	16.00	19.21	18.19
Imported Invest. / Invest.	3.58	3.58	40.00	40.00	40.00	40.00	66.84	66.84
Home Invest. / Invest.	96.42	96.42	60.00	60.00	60.00	60.00	33.16	33.16
Imported Cons. / Cons.	13.00	13.00	30.36	30.36	12.00	12.00	4.19	4.19
Home Cons. / Cons.	87.00	87.00	69.64	69.64	88.00	88.00	95.81	95.81
Commodities Exports / GDP	8.49	8.49	26.19	26.19	6.40	6.40	12.27	12.27
External Real Interest Rate (%, annual)	6.44	6.44	2.46	2.46	4.94	4.94	4.60	4.60
External Debt / GDP	66.39	66.39	7.54	7.54	40.00	40.00	25.31	25.31

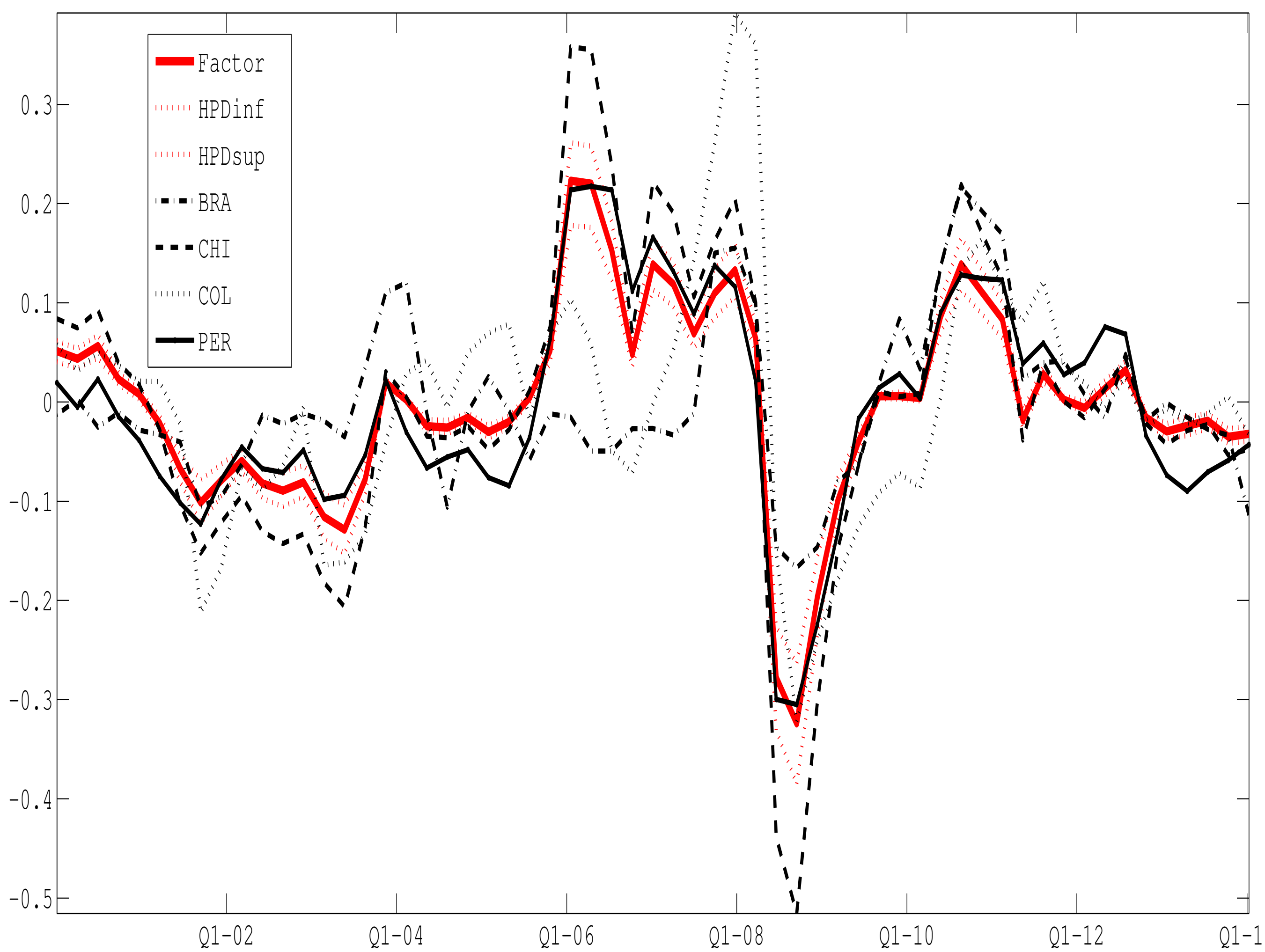
Note: The long-run values are equal to the numbers reported in the model descriptions of the DSGE models currently used for policy analysis at the central bank in Brazil, Chile, Colombia and Peru. For Brazil see de Castro et al. (2011), Chile see Medina et al. (2007); Colombia see Gonzalez et al. (2011); for Peru see Castillo (2006).

Table 4: Priors and Posteriors of Estimated Parameters

Parameters	Type	Prior Mean	Prior S.D.	Posterior Mode	Posterior S.D.	Posterior Mean	90% HPD interval	
				Global				
$\rho_{Y^\star}$	beta	0.500	0.150	0.8099	0.0471	0.8048	0.7322	0.8833
$\rho_{R^\star}$	beta	0.500	0.150	0.6362	0.0459	0.6229	0.5471	0.7042
ϕ_{Co}	beta	0.500	0.150	0.7337	0.0455	0.7310	0.6558	0.8046
$\sigma^{Y^\star}$	invg	0.013	Inf	0.0061	0.0006	0.0062	0.0053	0.0071
$\sigma^{R^\star}$	invg	0.013	Inf	0.0024	0.0002	0.0025	0.0021	0.0028
$\sigma^{f^{Co}}$	invg	0.013	Inf	0.0631	0.0089	0.0623	0.0478	0.0769
				Brazil				
ρ^{z^R}	beta	0.50	0.150	0.3482	0.0627	0.3430	0.2376	0.4448
ρ^{Co}	beta	0.50	0.150	0.7136	0.0717	0.7022	0.5923	0.8139
ρ^{z}	beta	0.50	0.15	0.3853	0.0921	0.3895	0.2421	0.5387
a	gamma	0.50	0.250	0.4313	0.1945	0.4971	0.1804	0.8165
ω^{Co}	norm	0.00	1.0	0.5315	0.1057	0.5629	0.3760	0.7466
σ^{z^R}	invg	0.01	Inf	0.0047	0.0004	0.0048	0.0041	0.0056
σ^{Co}	invg	0.01	Inf	0.0450	0.0040	0.0461	0.0390	0.0531
σ^{z}	invg	0.01	Inf	0.0129	0.0012	0.0132	0.0112	0.0154
				Chile				
ρ^{z^R}	beta	0.50	0.150	0.6996	0.0965	0.6740	0.5213	0.8422
ρ^{Co}	beta	0.50	0.150	0.4952	0.1743	0.5058	0.2631	0.7495
ρ^{z}	beta	0.50	0.150	0.3974	0.0930	0.4009	0.2430	0.5535
a	gamma	0.50	0.250	1.1702	0.3180	1.2267	0.6837	1.7537
ω^{Co}	norm	0.00	1.0	1.5358	0.1688	1.5987	1.2902	1.8999
σ^{z^R}	invg	0.01	Inf	0.0021	0.0002	0.0021	0.0018	0.0024
σ^{Co}	invg	0.01	Inf	0.0061	0.0029	0.0120	0.0036	0.0212
σ^{z}	invg	0.01	Inf	0.0332	0.0030	0.0339	0.0287	0.0389
				Colombia				
ρ^{z^R}	beta	0.50	0.150	0.5392	0.0928	0.5363	0.3942	0.6833
ρ^{Co}	beta	0.50	0.150	0.6805	0.0681	0.6741	0.5624	0.7846
ρ^{z}	beta	0.50	0.150	0.5033	0.0925	0.5097	0.3578	0.6676
a	gamma	0.50	0.250	0.4761	0.1330	0.5267	0.2885	0.7580
σ^{z^R}	invg	0.01	Inf	0.0025	0.0002	0.0026	0.0022	0.0029
σ^{Co}	invg	0.01	Inf	0.0579	0.0052	0.0597	0.0504	0.0688
σ^{z}	invg	0.01	Inf	0.0087	0.0008	0.0089	0.0076	0.0103
				Peru				
ρ^{z^R}	beta	0.50	0.150	0.3272	0.0846	0.3295	0.1946	0.4611
ρ^{Co}	beta	0.50	0.150	0.7039	0.0661	0.6945	0.5774	0.8051
ρ^{z}	beta	0.50	0.150	0.6667	0.0922	0.6815	0.5120	0.8521
a	gamma	0.50	0.250	1.1874	0.2581	1.2149	0.7438	1.6496
ω^{Co}	norm	0.00	1.0	0.8895	0.1083	0.9399	0.7294	1.1261
σ^{z^R}	invg	0.01	Inf	0.0023	0.0002	0.0023	0.0020	0.0027
σ^{Co}	invg	0.01	Inf	0.0236	0.0022	0.0238	0.0195	0.0283
σ^{z}	invg	0.01	Inf	0.0082	0.0007	0.0084	0.0071	0.0096

Note: This table shows the priors and posteriors based on 200,000 draws from the Metropolis-Hastings (MH) algorithm, discarding the first 100,000 draws. The mean and covariance matrix of the proposal density for the MH algorithm were the maximum of the posterior distribution and the negative inverse Hessian around that maximum obtained with Nelder-Mead simplex based optimization routine. The computations were conducted using Dynare 4.4.2. HPD stands for higher posterior density.

Figure 5: Estimated Common Factor in Country-Specific Commodity Price Indices



Note: The common factor (in solid/red) is the latent variable obtained from the Kalman filter smoothing evaluated at the mean of the posterior distribution. HPDinf and HPDsup stand for the 10^{th} and 90^{th} percentiles of the posterior distribution, respectively.

Table 5: Forecast Error Variance Decomposition of Real Income

shocks	Brazil	Chile	Colombia	Peru
		Commodity price shocks		
ν^{Co}	16.3	0.5	17.6	4.9
$\varepsilon^{f^{Co}}$	11.1	76.6	25.8	35.5
	27.5	77.1	43.5	40.4
		All other shocks		
ν^{R}	30.0	0.03	4.43	0.3
ε^{z}	22.0	22.9	45.4	58.3
$\varepsilon^{R^\star}$	20.3	0.03	5.8	0.92
$\varepsilon^{Y^\star}$	0.19	0.01	0.9	0.1
	72.5	22.9	56.5	59.6

(a) Unconditional

shocks	Brazil	Chile	Colombia	Peru
		Commodity price shocks		
ν^{Co}	7.2	0.7	10.0	3.3
$\varepsilon^{f^{Co}}$	4.3	55.2	11.1	19.8
	11.5	55.9	21.1	23.1
		All other shocks		
ν^{R}	1.94	0.03	0.4	0.01
ε^{z}	85.2	44.0	78.0	76.8
$\varepsilon^{R^\star}$	1.3	0.04	0.5	0.05
$\varepsilon^{Y^\star}$	0.02	0.00	0.07	0.0
	88.5	44.1	78.9	76.9

(b) Conditional (one quarter ahead forecast)

shocks	Brazil	Chile	Colombia	Peru
		Commodity price shocks		
ν^{Co}	11.3	0.6	12.3	3.3
$\varepsilon^{f^{Co}}$	7.3	69.0	15.6	21.1
	18.6	69.6	27.9	24.4
		All other shocks		
ν^{R}	4.2	0.04	0.7	0.01
ε^{z}	74.0	30.4	70.2	75.6
$\varepsilon^{R^\star}$	3.2	0.04	1.09	0.04
$\varepsilon^{Y^\star}$	0.03	0.01	0.11	0.01
	81.4	30.5	72.1	75.6

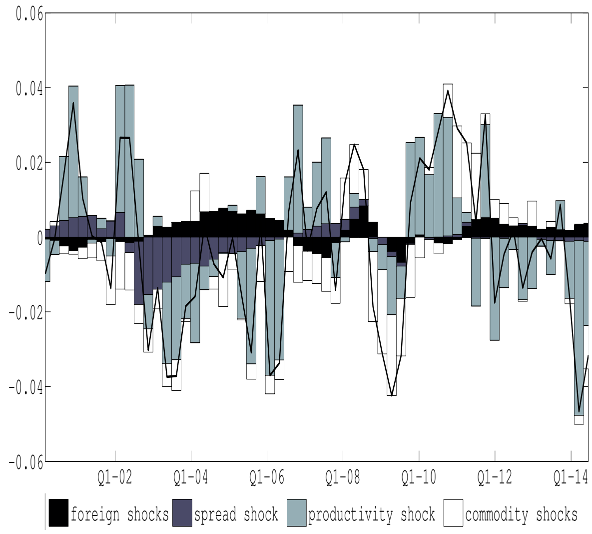
(c) Conditional (four quarters ahead forecast)

shocks	Brazil	Chile	Colombia	Peru
		Commodity price shocks		
ν^{Co}	13.4	0.5	13.5	3.4
$\varepsilon^{f^{Co}}$	9.0	74.3	18.7	22.9
	22.3	74.8	32.2	26.3
		All other shocks		
ν^{R}	9.4	0.04	1.5	0.06
ε^{z}	61.3	25.1	63.8	73.3
$\varepsilon^{R^\star}$	6.9	0.04	2.3	0.4
$\varepsilon^{Y^\star}$	0.06	0.01	0.19	0.01
	77.7	25.2	67.8	73.7

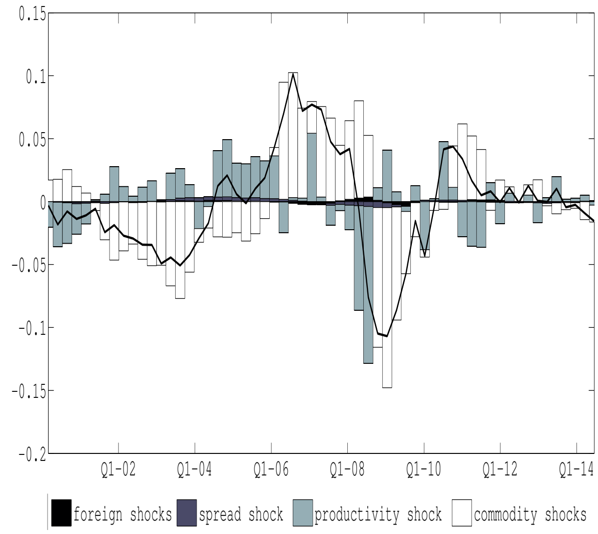
(d) Conditional (twelve quarters ahead forecast)

Note: The panels report the forecast error variance decomposition (FEVD) calculated at the posterior mean of output for four alternative forecast horizons. Shocks are as follows: Idiosyncratic commodity price shock $\left(\nu^{Co}\right)$; Common factor in commodity price shock $\left(\varepsilon^{f^{Co}}\right)$; Spread shock $\left(\nu^{R}\right)$; Domestic productivity shock (ε^{z}); World riskless interest rate shock $\left(\varepsilon^{R^\star}\right)$; World demand shock $\left(\varepsilon^{Y^\star}\right)$.

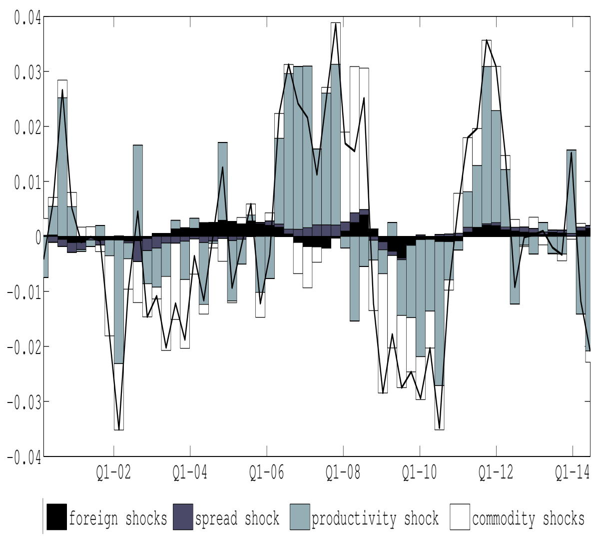
Figure 6: Historical Decomposition of Real Income



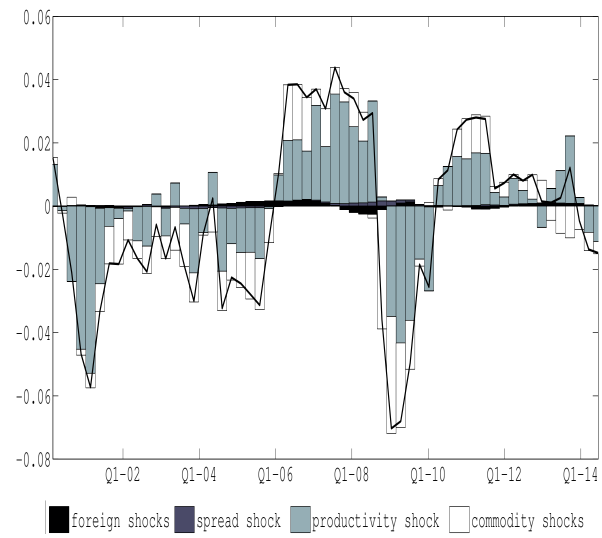
(a) Brazil



(b) Chile



(c) Colombia



(d) Peru

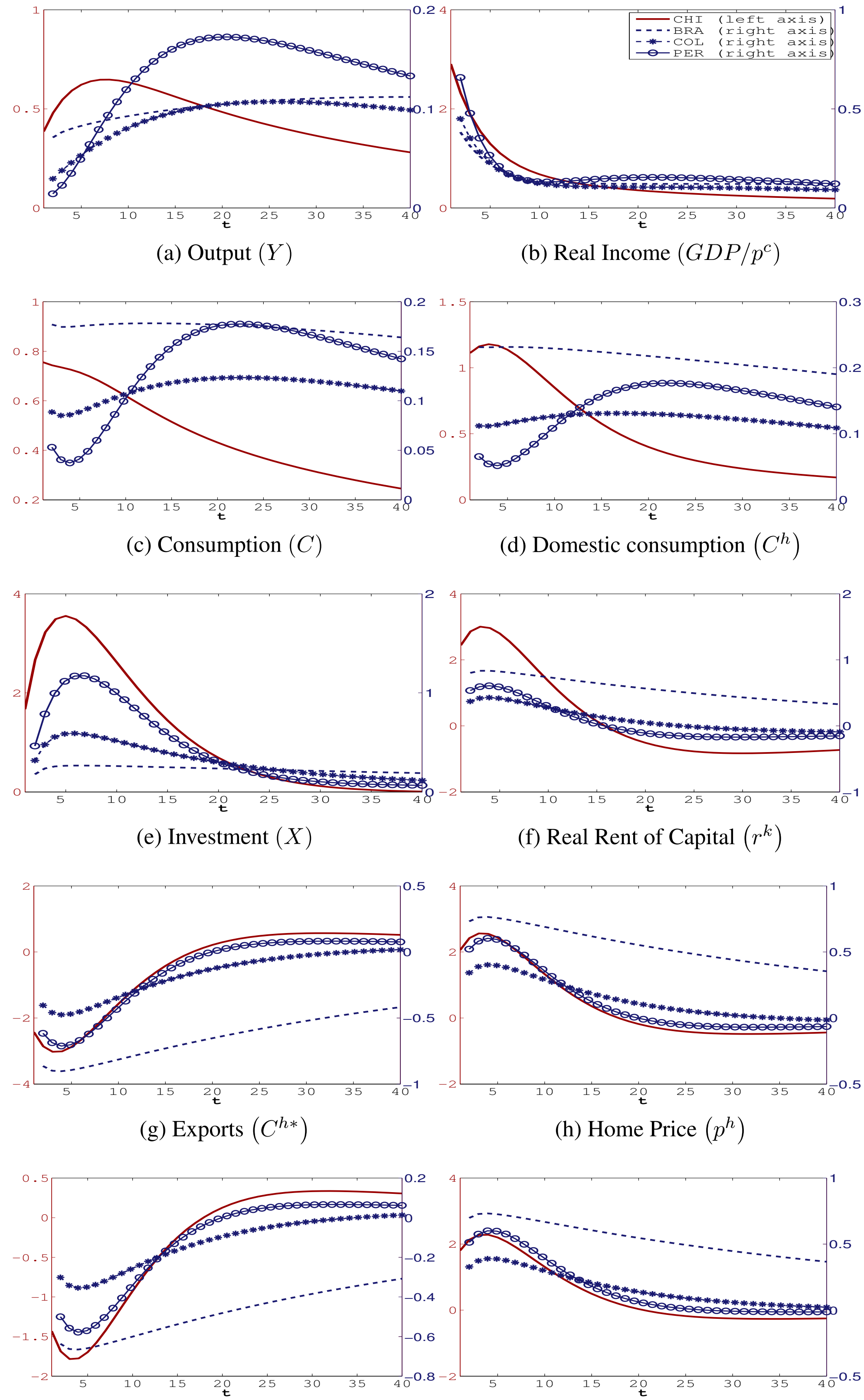
Note: Groups of shocks are as follows: Foreign shocks are World riskless interest rate shock $\left(\varepsilon^{R^\star}\right)$ and World demand shock $\left(\varepsilon^{Y^\star}\right)$; Spread shock is $\left(\nu^{R}\right)$; Productivity shock is $\left(\varepsilon^{z}\right)$; and Commodity Shocks are Idiosyncratic commodity price shock $\left(\nu^{Co}\right)$ and Common factor in commodity price shock $\left(\varepsilon^{f^{Co}}\right)$.

Table 6: Counterfactual Historical Decomposition Experiments

shocks	Brazil	Chile	Colombia	Peru
S.D. of observed real income cycle	0.0218	0.0418	0.0184	0.0273
Counterfactual S.D. excluding each shock separately (in percent of the observed S.D) (%)				
ν^{R} (Idiosyncratic spread)	92.7	98.8	97.7	99.9
ν^{Co} (Idiosyncratic commodity price)	103.1	99.7	98.2	94.0
ε^{z} (Domestic productivity)	53.0	114.9	55.1	43.9
$\varepsilon^{R^\star}$ (World interest rate)	105.6	101.2	100.4	99.8
$\varepsilon^{Y^\star}$ (World demand)	99.4	99.3	98.1	100.6
$\varepsilon^{f^{Co}}$ (Common factor in commodity prices)	89.1	72.2	80.3	75.4
Counterfactual S.D. excluding two commodities price shocks (%)				
$\left\{\nu^{Co},\, \varepsilon^{f^{Co}}\right\}$	91.1	72.1	75.4	69.6
Counterfactual S.D. excluding all shocks except commodities price shocks (%)				
$\left\{\nu^{R},\, \varepsilon^{z},\, \varepsilon^{R^\star},\, \varepsilon^{Y^\star}\right\}$	41.5	114.7	47.9	44.1

Note: The table presents in rows two and below the counterfactual standard deviation (S.D) of real income implied by the model when one or several of the structural shocks are turned off. The numbers reported are in percentage of the observed S.D. presented in the first row computed over the entire sample period of estimation: 2000.Q1 to 2014.Q3.

Figure 7: Impulse Responses to a Common Shock in Commodity Prices



(a) Output (Y)

(b) Real Income (GDP/p^c)

(c) Consumption (C)

(d) Domestic consumption $\left(C^h\right)$

(e) Investment (X)

(f) Real Rent of Capital $\left(r^k\right)$

(g) Exports $\left(C^{h*}\right)$

(h) Home Price $\left(p^h\right)$

(i) Real Exchange Rate (RER)

(j) Real Wage (w)

Note: The subplots present the impulse response functions following a common factor shock in commodity prices. Units are percentage deviations from steady state levels. The unit of time in the horizontal axis is a quarter.

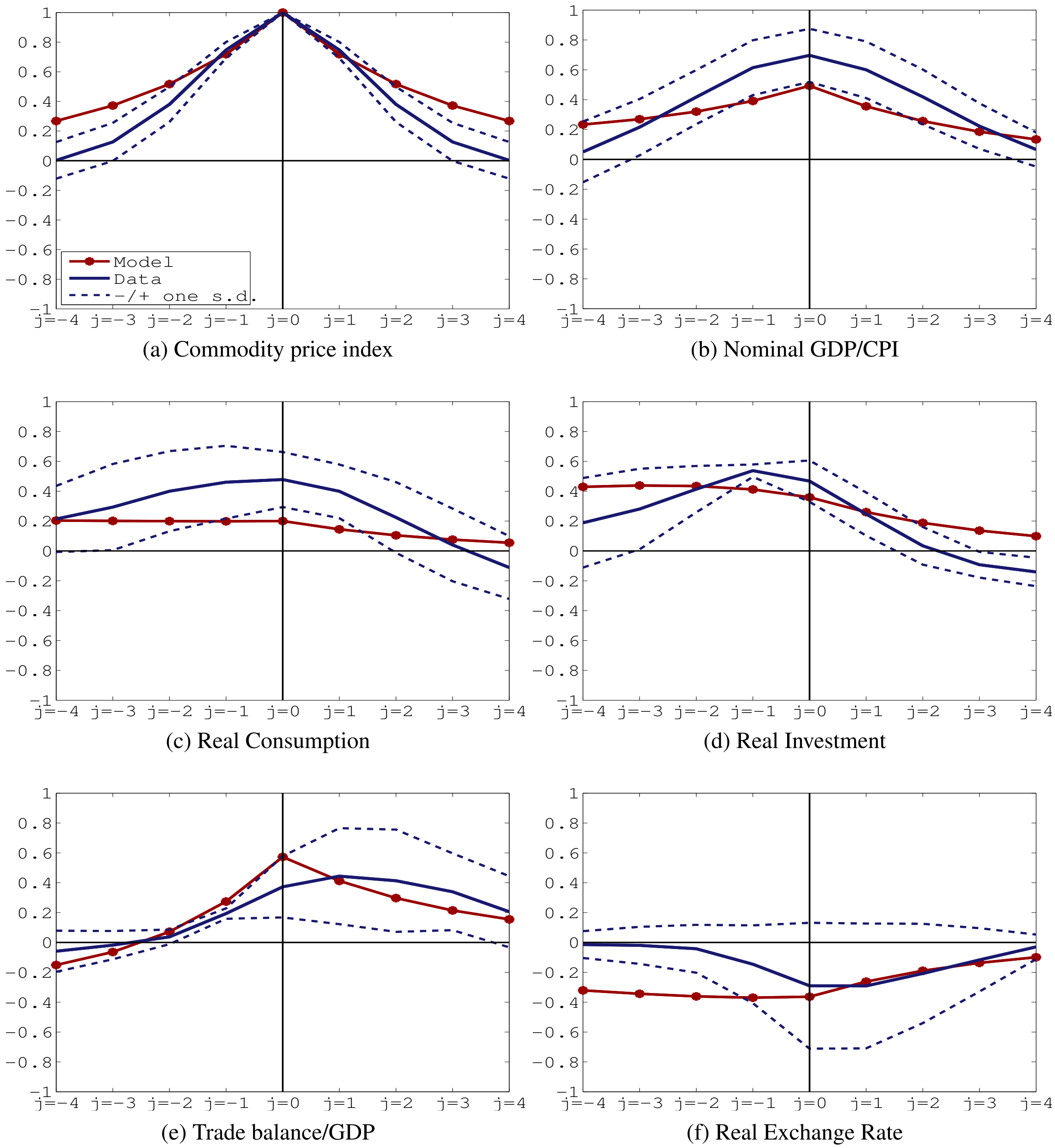
Table 7: Cross-Correlations of Real Income across Countries: Data and Model

Shocks	Brazil	Chile	Colombia	Peru
Brazil	1	0.31**	0.21	0.36***
Chile	0.00^{NCF}, 0.22^{CF}	1	0.53***	0.70***
Colombia	0.10^{NCF}, 0.27^{CF}	0.00^{NCF}, 0.40^{CF}	1	0.58***
Peru	0.04^{NCF}, 0.22^{CF}	0.00^{NCF}, 0.47^{CF}	0.02^{NCF}, 0.32^{CF}	1

Marginal likelihood	
Model with no common factor (NCF):	4445.3
Model with common factor (CF):	4530.8

Note: The upper panel reports the covariance matrix of real income across the four countries used in the estimation of the structural model. Numbers above the main diagonal are computed from the data during the same range of time on which the model is estimated (2000.Q1 to 2014.Q3). Statistical significance at 1, 5, and 10 percent is reported with (***), (**), and (*), respectively. Numbers below the main diagonal are those predicted by the model. Those with superscript "CF" come from the benchmark model with a common factor in commodity prices. Those with "NCF" are generated with a model with no common factor. Marginal likelihood is computed using Geweke's modified harmonic mean.

Figure 8: Unconditional Serial Correlations with Commodity Prices: Data and Model



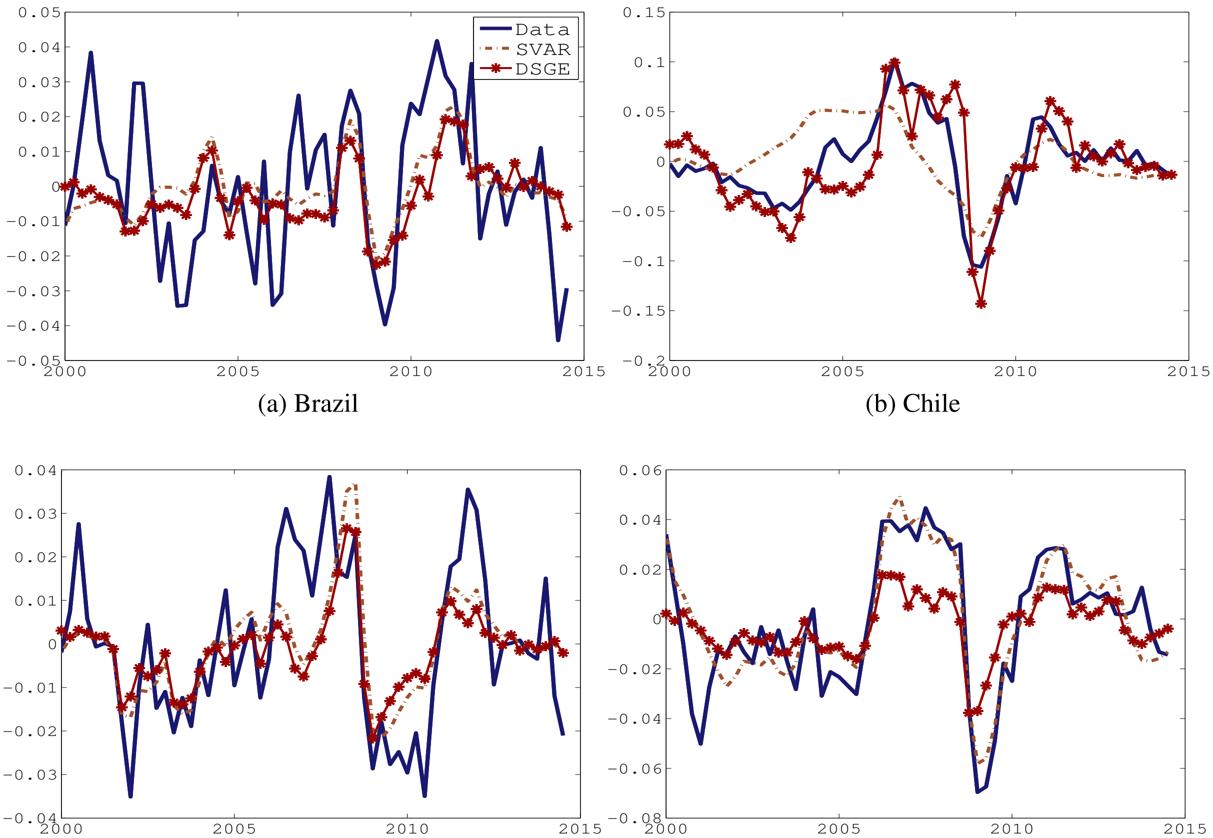
Note: Data corresponds to the simple average across the four countries used in the estimation of the model. Model-based correlations are computed with the estimated benchmark model.

Table 8: Other Second Moments: Model and Data

	S.D. (%)	Relative S.D to GDP/CPI	Autocorrelation		Cross correlation with GDP/CPI $(t+j)$								
			1	2	$j=-4$	$j=-3$	$j=-2$	$j=-1$	$j=0$	$j=1$	$j=2$	$j=3$	$j=4$
Model													
P^{Co}	11.08	3.43	0.72	0.52	0.13	0.19	0.26	0.36	0.49	0.39	0.32	0.27	0.23
GDP/CPI	3.80	1.00	0.79	0.66	0.53	0.59	0.66	0.79	1.00	0.79	0.66	0.59	0.53
$Cons$	3.32	0.89	0.83	0.75	0.61	0.63	0.67	0.75	0.92	0.75	0.65	0.59	0.55
Inv	7.40	2.00	0.91	0.80	0.55	0.61	0.69	0.77	0.80	0.64	0.54	0.49	0.46
TB/GDP	1.57	0.40	0.76	0.57	-0.38	-0.36	-0.33	-0.26	-0.09	-0.13	-0.17	-0.19	-0.21
RER	6.27	1.64	0.75	0.63	-0.23	-0.24	-0.24	-0.19	-0.01	-0.13	-0.18	-0.21	-0.22
Data													
P^{Co}	12.14	4.24	0.75***	0.39***	0.11*	0.29***	0.49***	0.67***	0.73***	0.64***	0.45***	0.26***	0.10
GDP/CPI	2.86	1.00	0.80***	0.56***	0.15**	0.36***	0.56***	0.80***	1.00***	0.80***	0.56***	0.36***	0.15**
$Cons$	1.34	0.47	0.82***	0.60***	0.42***	0.56***	0.68***	0.72***	0.69***	0.53***	0.28***	0.06	-0.11**
Inv	5.14	1.80	0.72***	0.42***	0.35***	0.45***	0.52***	0.54***	0.49***	0.30***	0.07	-0.05	-0.14**
TB/GDP	1.99	0.70	0.68***	0.41***	-0.13*	-0.07	0.05	0.25***	0.46***	0.52***	0.52***	0.46***	0.31***
RER	5.44	1.90	0.63***	0.19***	-0.09	-0.08	-0.11*	-0.22***	-0.27***	-0.30***	-0.25***	-0.17**	-0.10

Note: This table reports other moments of the data and the model for the commodity price index, nominal GDP over CPI, consumption, investment, trade balance over GDP and real exchange rate. The panels report the absolute standard deviation, the standard deviation relative to GDP/CPI, the autocorrelation coefficient of orders one and two and the cross correlation for every variable in period t with GDP/CPI in period $t+j$ with $j=-4,\ldots,4$. Statistical significance at 1, 5, and 10 percent is reported with (***), (**), and (*), respectively.

Figure 9: Comparison with an SVAR model



(a) Brazil

(b) Chile

(c) Colombia

(d) Peru

Note: Solid/blue lines represent the filtered real income series using the Hodrick-Prescott filter. Starred/red and dotted/orange lines report the simulated real GDP series from the DSGE and the SVAR models respectively when only the structural commodity price shocks are used and the remaining shocks are shut down. The SVAR is estimated with the cyclical components obtained using a Hodrick-Prescott filter (lambda of 1600) on quarterly series of the commodity price index, trade balance over real income, real income, real private consumption, real investment and real exchange rate between 2000.Q1 and 2014.Q4. The identification scheme imposes that the commodity price index is exogenous by assuming that the commodity price follows an autorregresive process of order one and by applying a lower triangular Cholesky decomposition on the variance-covariance matrix of the residuals (see text for details).

Table 9: Comparison with an SVAR

	Brazil	Chile	Colombia	Peru	Median	Mean
	A. Unconditional Variance Decomposition of Output: Share of Pcom shocks					
SVAR	15	53	47	80	50	49
DSGE	28	77	44	40	42	47
	B. SD(Output-counterfactual)/SD(Output-data): Only Pcom shocks					
SVAR	41	74	63	91	69	67
DSGE	42	115	48	44	46	62
	C. SD(Output-counterfactual)/SD(Output-data): All shocks except Pcom					
SVAR	88	90	71	48	80	74
DSGE	91	72	75	70	74	77
	D. corr(Output-counterfactual,Output-data): Only Pcom shocks					
SVAR	0.45	0.55	0.71	0.88	0.63	0.65
DSGE	0.41	0.78	0.67	0.82	0.73	0.67
	E. corr(Output-counterfactual-SVAR,Output-counterfactual-DSGE): Only Pcom shocks					
	0.88	0.27	0.96	0.91	0.89	0.75

Note: PCom stands for country-specific commodity price. SD(X) stands for the standard deviationof variable "X". This table reports some statistics that compare the estimated SVAR and the DSGE model. See the footnote in Figure 9 for further information on the SVAR estimation procedure. In the table, "counterfactual" stands for the simulated series from the model either when all shocks but the commodity price one are shut down or when only the commodity price shock is shut down.

Table 10: Forecast Error Variance Decomposition of Output: Alternative Models

(a) Countercyclical spreads

shocks	Brazil	Chile	Colombia	Peru
	Commodity price shock			
ν^{Co}	24.4	0.6	18.0	5.6
$\varepsilon^{f^{Co}}$	17.1	74.0	27.7	37.6
	41.5	74.6	45.7	43.2
	All other shocks			
ν^{R}	24.8	0.1	4.9	0.5
ε^{z}	17.6	25.2	41.1	54.7
$\varepsilon^{R^\star}$	16.0	0.1	7.5	1.5
$\varepsilon^{Y^\star}$	0.1	0.0	0.8	0.1
	58.5	25.4	54.3	56.8
Marginal likelihood				4,502.40

(b) Correlation between external forces

shocks	Brazil	Chile	Colombia	Peru
	Commodity price shock			
ν^{Co}	16.5	25.6	17.8	4.8
$\varepsilon^{f^{Co}}$	7.9	64.3	18.5	27.0
	24.5	89.9	36.3	31.8
	All other shocks			
ν^{R}	30.3	0.6	4.4	0.3
ε^{z}	22.6	0.0	47.1	62.5
$\varepsilon^{R^\star}$	20.5	0.0	6.0	1.0
$\varepsilon^{Y^\star}$	2.1	9.4	6.1	4.4
	75.6	10.1	63.7	68.2
Marginal likelihood				4,535.70

(c) Basic RBC model (no commodities)

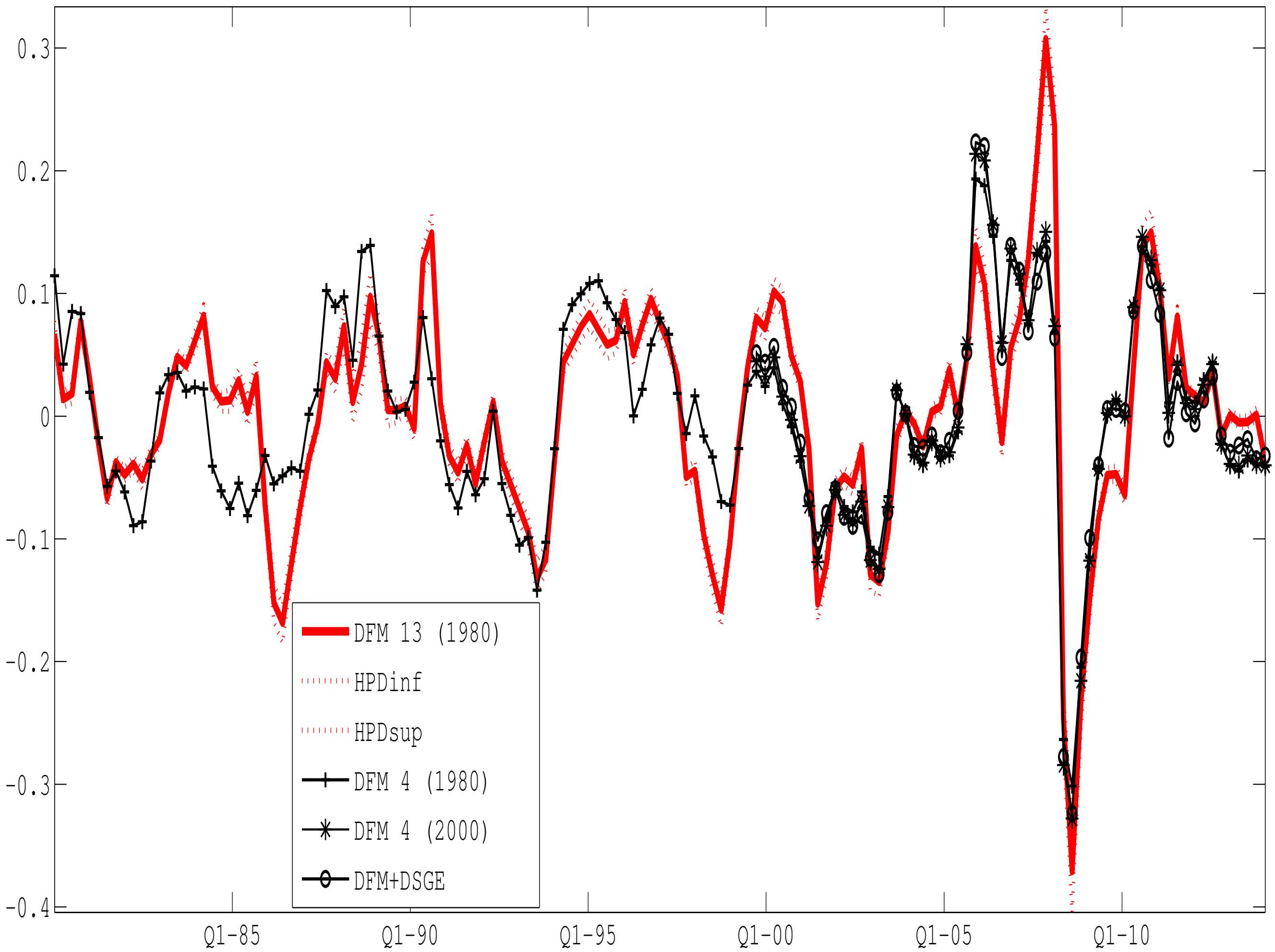
shocks	Brazil	Chile	Colombia	Peru
	Commodity price shock			
ν^{Co}	-	-	-	-
$\varepsilon^{f^{Co}}$	-	-	-	-
	-	-	-	-
	All other shocks			
ν^{R}	42.4	0.0	3.8	0.3
ε^{z}	33.3	100.0	88.5	99.0
$\varepsilon^{R^\star}$	24.1	0.0	6.6	0.7
$\varepsilon^{Y^\star}$	0.3	0.0	1.1	0.1
	100.0	100.0	100.0	100.0
Marginal likelihood				4,410.60

(d) Benchmarck model first differences

shocks	Brazil	Chile	Colombia	Peru
	Commodity price shock			
ν^{Co}	15.5	1.8	5.0	0.8
$\varepsilon^{f^{Co}}$	44.7	56.6	71.2	76.4
	60.1	58.4	76.1	77.3
	All other shocks			
ν^{R}	22.1	0.0	2.7	0.1
ε^{z}	10.9	41.6	18.9	22.4
$\varepsilon^{R^\star}$	6.6	0.0	1.4	0.1
$\varepsilon^{Y^\star}$	0.2	0.0	0.9	0.1
	39.9	41.6	23.9	22.8

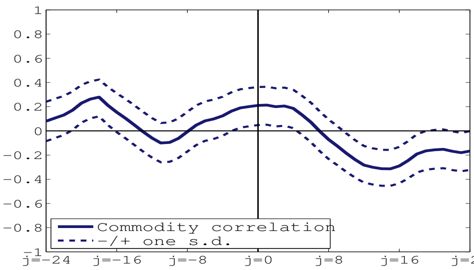
Note: The panels report the forecast error variance decomposition (FEVD) of real income calculated at the posterior mean for four alternative reduced models. Panel (a): Countercyclical spreads; Panel (b): Correlation between common factor f^{co} and foreign output y^{*}; Panel (c): No commodity shocks; Panel (d): Benchmark model in first differences. Marginal Likelihood are computed with Geweke's modified harmonic mean. See footnote in Table 5 for description of the shocks.

Figure 10: Extended Dynamic Factor Models

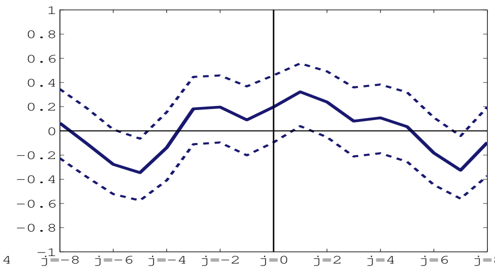


Note: The solid/red line is the common factor obtained with a dynamic factor model (DFM) that uses only quarterly data on country-specific commodity prices between 1980.Q1 and 2014.Q3 for the 13 EMEs listed in the text. Dashed/red lines are highest posterior density estimates of confidence bands. DFM 4 (1980) and DFM 4 (2000) reproduce the common factors of two separate DFMs ran with LAC4 countries (Brazil, Chile, Colombia, and Peru) starting in 1980.Q1 and 2000.Q1, respectively. The latter is the starting point of the DSGE estimation. Lastly, DFM+DSGE reproduces the common factor from our benchmark estimation from the DSGE model presented in Figure 5. The common factor is the latent variable obtained from the Kalman filter smoothing evaluated at the mean of the posterior distribution.

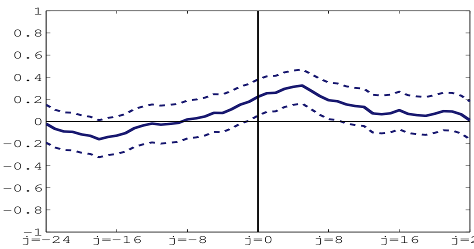
Figure 11: Correlation of Commodity and Non-Commodity Production and International Commodity Prices



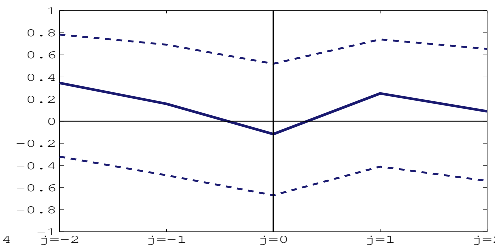
(a) Brazil - Tons of Iron (monthly)



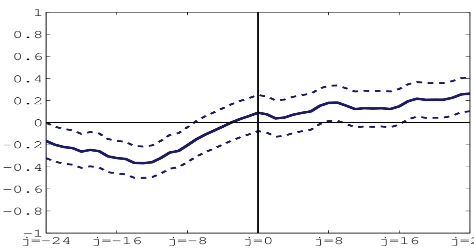
(b) Brazil - Tons of Soybeans (quarterly)



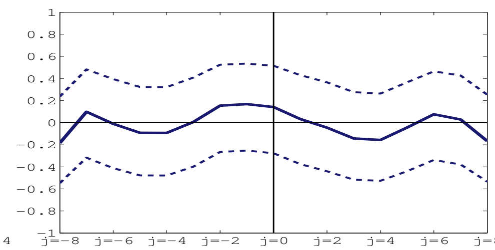
(c) Chile - Tons of Copper (monthly)



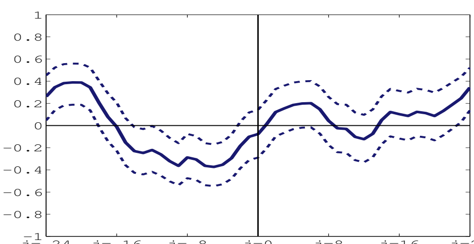
(d) Chile - Tons of Fish (annual)



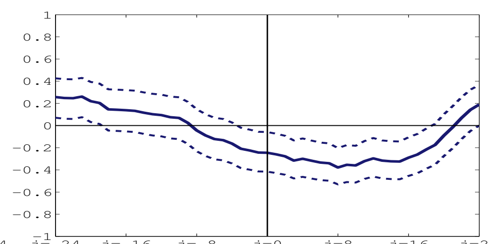
(e) Colombia - Barrels per Day of Oil (monthly)



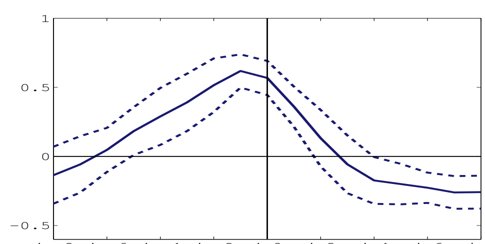
(f) Colombia - Tons of Coal (quarterly)



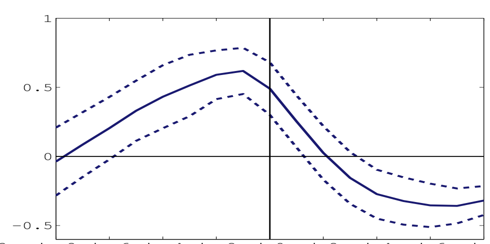
(g) Peru - Fine Grams of Gold (monthly)



(h) Peru - Tons of Copper (monthly)



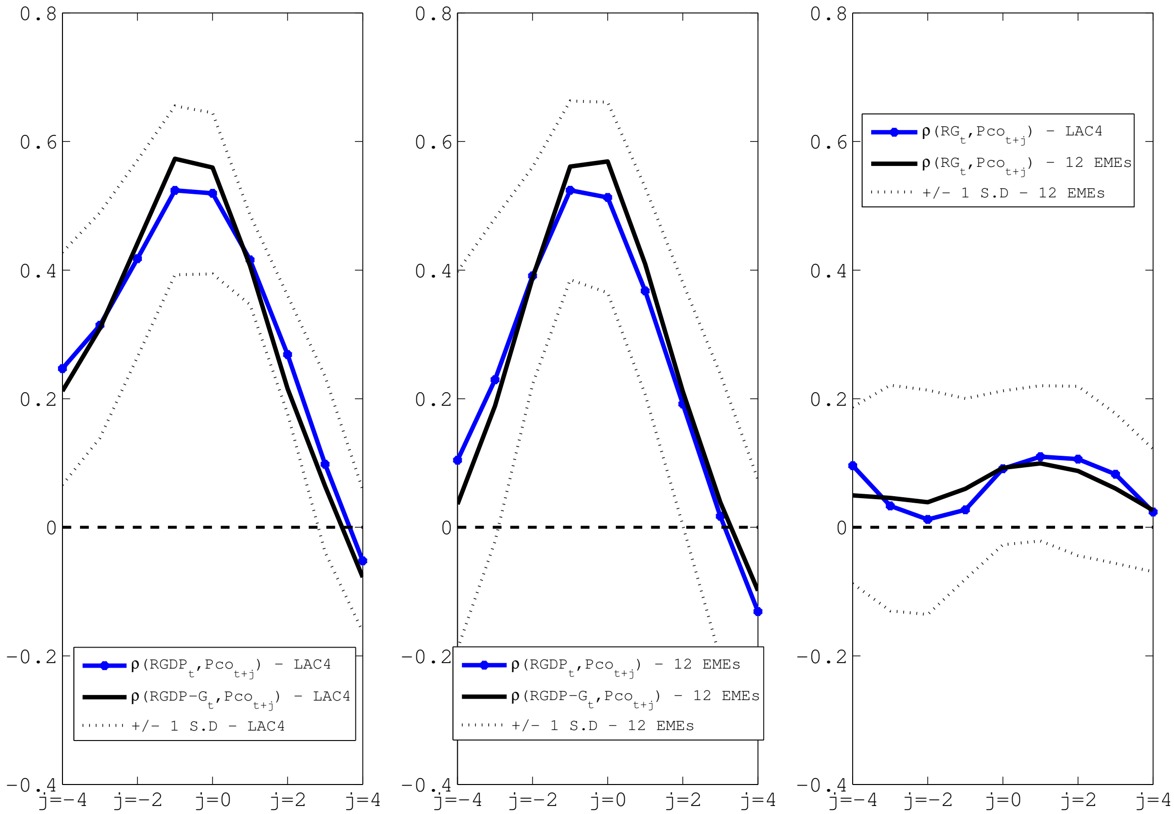
(i) LAC-4 - Non-commodities GDP (quarterly)



(j) EME-11 - Non-tradable GDP (quarterly)

Note: Panels (a) to (h) report the serial correlations between the volumetric production of commodities in each of the two main commodities of LAC4 countries in t and the international market price of each commodity in $t + j$ with $j = -4, \ldots, 4$. The correlations were computed with the highest frequency available which is reported in the subtitle. Panels (i) and (j) present averages of the correlations between the non-commodity GDP and non-tradable GDP in t, respectively, with country-specific commodity price indices across LAC4 and EME11 $t + j$ with $j = -4, \ldots, 4$. Sources: National Statistical Agencies and author's calculations. See text for further details.

Figure 12: The Role of Government Expenditure



Note: Left panel reports in starred/blue the average serial correlation between real GDP (RGDP) in t and country-specific commodity prices (Pco) in $t+j$ with $j=-4,\ldots,4$, for LAC4 countries along with the 1 S.D error bands (dotted line). Middle panel expands this to 12 EMEs. Right Panel plots the average correlation of real government expenditure (RG) in t and Pco in $t+j$ with $j=-4,\ldots,4$ across EME12. Solid dark line presents the same statistic but substracting RG from RGDP.