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Sharing a ride on the commodities roller coaster: Common factors in business cycles of emerging economies☆

Andrés Fernández[a,*], Andrés González[b], Diego Rodríguez[b]

[a] Inter-American Development Bank, Research Dpt., United States
[b] Banco de la República, Colombia

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ABSTRACT

We explore the hypothesis that fluctuations in commodity prices are an important driver of business cycles in small emerging market economies (EMEs). First, we document that commodity prices exhibit strong comovement with other macro variables along the business cycle of these economies; and that a common factor accounts for most of the time series dynamics of these commodity prices. Guided by these stylized facts, we embed a commodity sector into a dynamic, stochastic, multi-country business cycle model of EMEs where exogenous fluctuations in commodity prices coexist with other driving forces. Commodity prices follow a common dynamic factor structure in the model. When estimated with EMEs data, the model gives to commodity shocks, mostly in the form of perturbations to their common factor, a paramount role when accounting for aggregate dynamics: more than a third of the variance of real output across the EMEs considered is associated to commodity price shocks. The model also performs well when accounting for other business cycle facts. A further amplification mechanism is a "spillover" effect from commodity prices to interest rates. Yet, sometimes, positive commodity price shocks have also cushioned other negative domestic shocks, particularly during the fast recovery from the world financial crisis.

1. Introduction

In recent times, the world economy has witnessed large fluctuations in the prices of commodity goods traded in international markets, which have been observed across distinct types of commodities, from agricultural products to fuels and metals. What have been the macroeconomic consequences of these movements in prices 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 fluctuations in the prices of the commodity goods that they export. Thus, an important open question in international macroeconomics is the following: what are the main channels by which commodities price fluctuations affect business cycles in EMEs and how much have they mattered in practice?

This paper explores formally the hypothesis that fluctuations in the price of commodity goods - easily comparable to a wild roller coaster ride, hence the title of our work - may be a key driver of business cycles in EMEs. It does so by using data and economic theory. First, we 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.

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* Corresponding author.
E-mail address: andresf@iadb.org (A. Fernández).

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The paper makes three types of contributions. The first one is empirical as we document the cyclical properties of commodity prices in EMEs. We do so by constructing country-specific commodity price indices from 44 distinct commodity goods, using historical export weights. We then correlate these indices at the country level with macroeconomic variables and 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 model is a multi-EME version of Mendoza's (1991) small open economy setup. 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 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 and intuitive. When positive, it acts as an income shock 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 goods become relatively more expensive for the rest of the world, exports fall and the trade balance deteriorates. Yet for realistically calibrated shares of exports to income and shares of domestically produced goods in total consumption and investment, the contraction in exports 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 volatile, and real exchange rates and trade balance are countercyclical.

On the empirical front, we document two key stylized facts. On one hand, the country-specific commodity price indices that we build for several EMEs 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. On the other hand, 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 products. 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.

As an application of the structural model, we estimate it using data from Brazil, Chile, Colombia and Peru in the post-2000 period. The estimation gives commodity price shocks a paramount role when accounting for aggregate dynamics in these four EMEs. The median share of the forecast error variance in real output accounted by these shocks is 42%. 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 the four EMEs, as well as other more traditional second moments. For instance, the model predicts that real exchange rates are countercyclical, as observed unconditionally in the data. Though fluctuations in commodity prices have not always amplified the business cycle. 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 of the EMEs considered.

Several robustness tests and extensions are conducted. 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 country risk premia to be affected by fundamentals increases the role of commodity prices although the estimation does not find a sizeable quantitative role for this further amplification mechanism. Third, the important role of commodity price shocks remains 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 large cost in terms of the model's performance. Most notably, it counterfactually predicts that expansions (contractions) in economic activity are accompanied by real exchange rate depreciations (appreciations). Fifth, we corroborate that our main results from the estimated structural model survive when using alternative detrending filters. Sixth, we modify the structural model in order to allow for the possibility that commodity prices follow a unit root process.

Our work highlights the linkages between business cycle and commodity prices in commodity exporting EMEs. Even though there may be several channels through which commodity price fluctuations impact these economies, we concentrate on a particular channel: the demand channel coming from income shocks triggered by commodity price fluctuations. There are at least two additional channels that our conceptual framework is leaving out. First, because we model the commodity sector as an endowment enclave, we abstract from any supply channels that may operate, namely those involving the decision to produce and demand more labor and capital, which may be relevant for, respectively, labor-intensive and capital-intensive commodity exporters.[1] Another potentially relevant channel that our setup is leaving out is the role of monetary and fiscal policy. For the particular case of fiscal policy, shutting down this channel may be a nontrivial simplification because EME governments typically either own (at least part of) the commodity exporting firms sector or tax them heavily to finance spending elsewhere.[2] We conjecture, however, that adding such channels may end up reinforcing the central role of commodities for the business cycle of these economies either by reinforcing the reallocation of resources to the commodity sector or by reinforcing the role that procyclical policies may have during commodity price booms and busts.

[1] In a recent contribution Caputo and Irarrázabal (2017) investigate the role of commodity price shocks in an environment of commodity production calibrated to Chile. They also obtain that a positive shock to the commodity price leads to an increase in both non-commodity output and employment, mainly via the effect over total investment.

[2] A burgeoning literature has studied the interaction between commodity markets and fiscal/monetary policies in EMEs within a DSGE setup. See Kumhof and Laxton (2009), Bodenstein et al. (2011), Pieschacón (2012), Agénor (2016), Catão and Chang (2013), Hevia and Nicolini (2015), Fornero et al. (2016), Medina and Soto (2016), Ojeda-Joya et al. (2016), among others. These studies feature nominal rigidities and/or distinct policy rules with the goal of evaluating the relative performances of either monetary and/or fiscal policy rules. None of them looks at the relative contribution of commodity vs. various other possible shocks to output variance as the main object of analysis, like we do in this work. Likewise, none the above papers feature a latent factor model for the commodity sector which is central in our results.

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 (Kose, 2002; Mendoza, 1995). 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 financial frictions 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, 2015). 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 closely related to the recent contributions by Shousha (2016) and Drechsel and Tenreyro (2017) who also add commodity price shocks to a RBC-SOE model of an emerging economy and postulate endogenous movements in external debt spreads as a further propagating mechanism of fluctuations in commodity prices, finding also a key role of commodity price fluctuations for business cycles. Unlike us, none of these two works explicitly model nor quantify the role of a common factor driving business cycles across emerging economies.

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; Guerrón-Quintana, 2013), and emerging economies (Akinci, 2013; Maćkowiak, 2007; Broda, 2004; Miyamoto and Nguyen, 2017). Within the literature of EMEs, special attention has been given to two potential drivers of business cycles: fluctuations in external interest rates (Akinci, 2013; Maćkowiak, 2007; Canova, 2005) 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 (Guerrón-Quintana, 2013; Miyamoto and Nguyen, 2017) 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 FAVAR models (Lombardi et al., 2012; Byrne et al., 2013), networks analysis (Matesanz et al., 2014), and dynamic factor models (Alquist and Coibion, 2013; Delle Chiaie et al., 2017). 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.

2. Stylized facts

2.1. Cyclicality

We explore the cyclicality of commodity prices using a novel quarterly panel dataset with country-specific commodity price indexes for 61 EMEs, between 1980.Q1 and 2014.Q4.[4] 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 United Nations' Comtrade data. $P^{Co}_{t,i}$ is obtained by deflating the monthly commodity prices indices reported by the IMF's Primary Commodity Price Database with the US consumer price index.[5]

We study the comovement of these indices with several country-specific macro variables at the quarterly frequency. Formally, we compute the serial correlation $\rho\left(X_{n,t}, P^{Co}_{n,t+j}\right)$ with $j = -4, \ldots, 4$; 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. While the countries' commodity price indices $P^{Co}_{n,t}$ are computed for 61 EMEs, unfortunately, imposing a minimum range of time series data in the variables Xs reduces the sample to 13 EMEs.[6] Each of the serial correlations is depicted on a subplot in Fig. 1. Statistics reported are simple averages across the 13 countries (EME13) in the sample and confidence bands denote ±1.0 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. All correlations are computed using the cyclical component of each variable, which we extract using the Hodrick-Prescott filter, though later we consider alternative filtering methods.

[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).

[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. See the Appendix for the list of 61 countries and further details.

[5] 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 Online Appendix.

[6] 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 all 61 EMEs in our entire sample. 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.

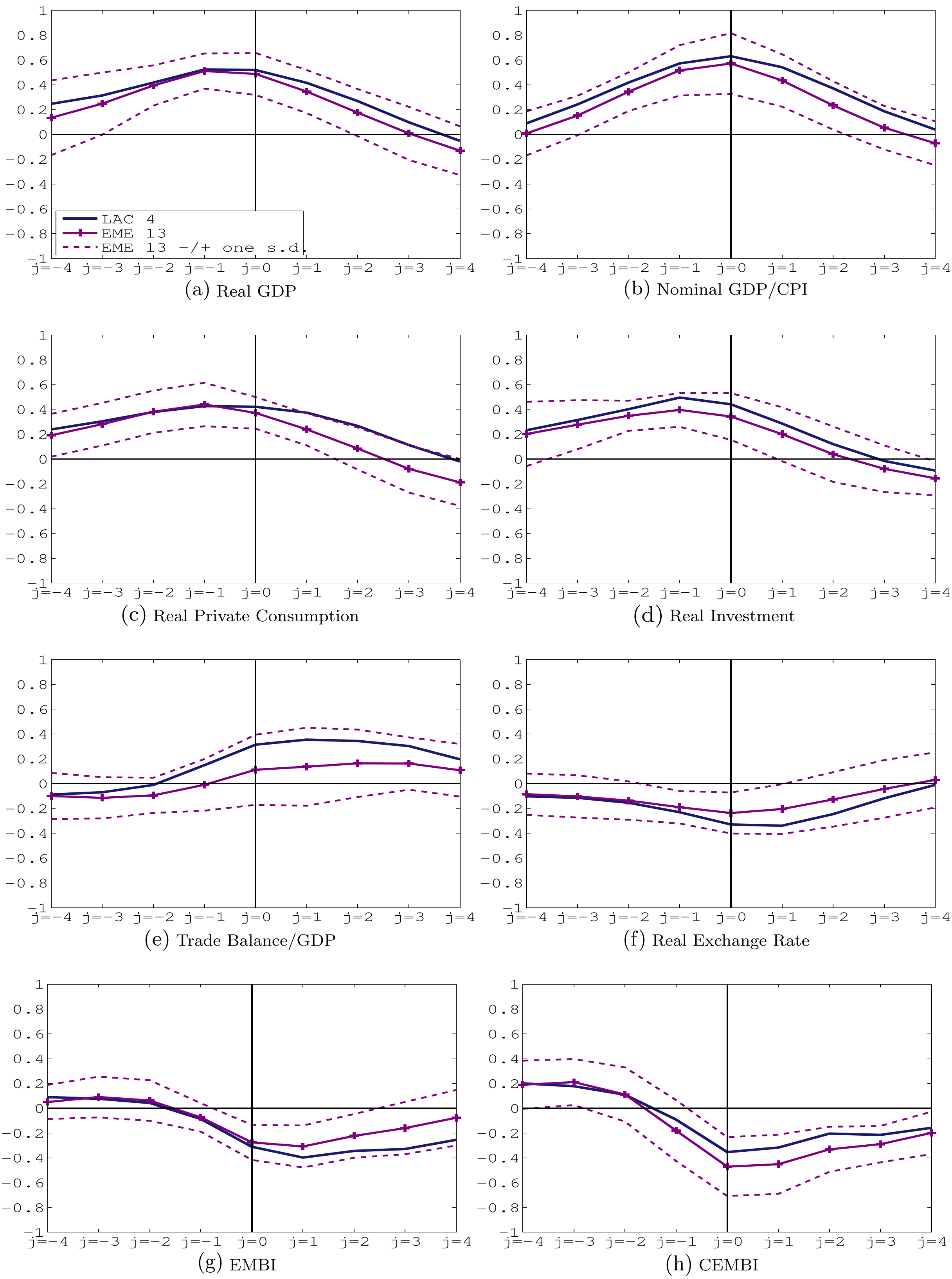

Fig. 1. Cyclicality of country-specific commodity price indices in emerging economies.
Note: Each panel reports in purple/solid with "+" lines the simple average correlation between the macroeconomic variable in the subtitle in period t and the country-specific commodity price indices in $t + j$ with $j = -4, \ldots, 4$, across 13 Emerging Market Economies (EMEs, 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 the Appendix for the exact range of quarterly time series used and for country coverage.

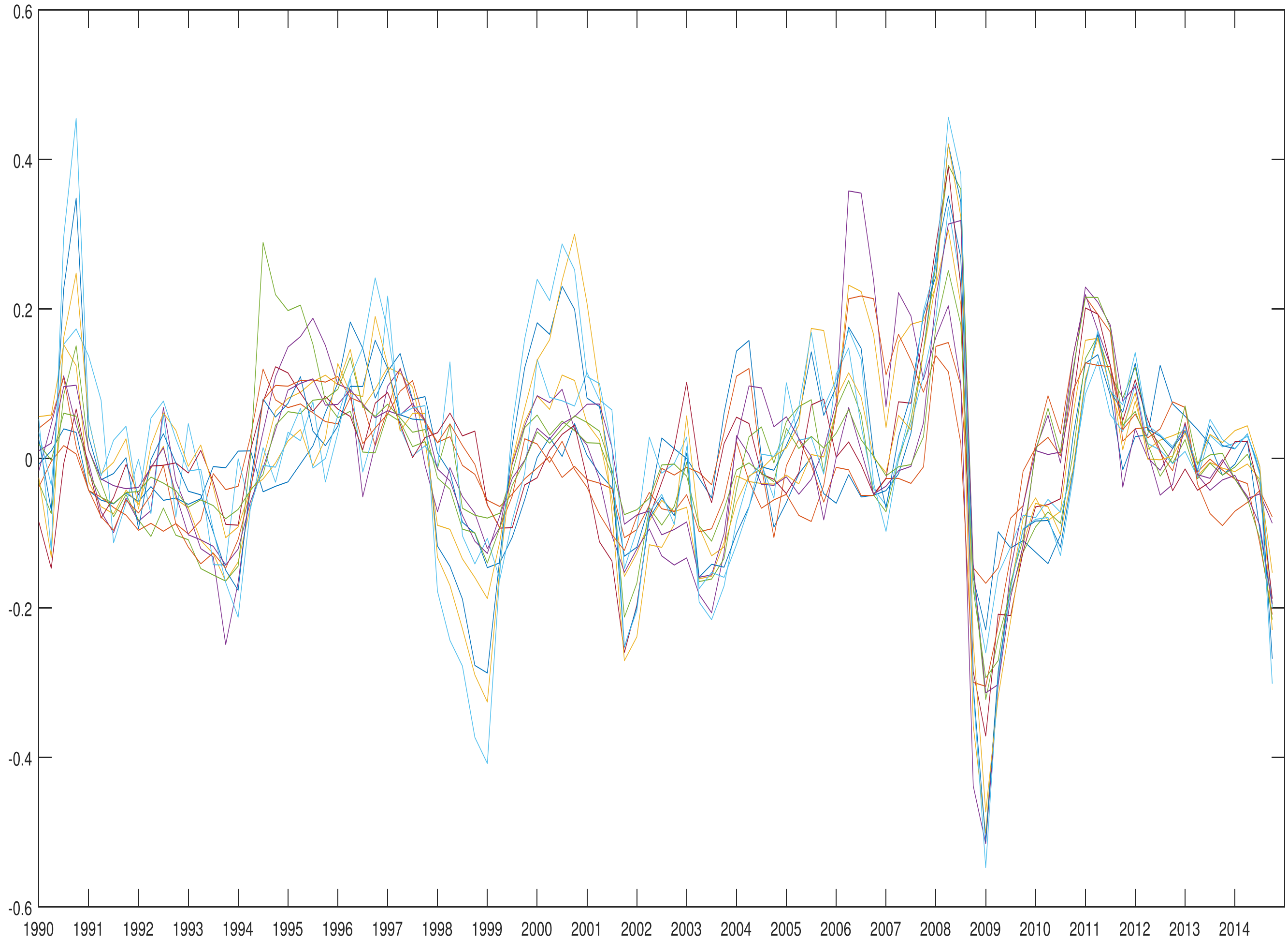

Fig. 2. Country-specific commodity price indices in emerging economies: 1990–2014.
Note: The figure reports the quarterly time series for the cyclical components of country-specific commodity price indices across the 13 EMEs studied (see Footnote 6). Cyclical components are computed as relative deviations 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. The Online Appendix contains a colored version of this plot that includes a legend.

A first stylized fact stands out when inspecting the panels in this figure: The cyclicality in the country-specific commodity prices in EMEs is strong. They 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, Fig. 1). 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 Fig. 1 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. Thus, when commodity prices are high (low), the cost of issuing debt in foreign capital markets decreases (increases) for EMEs.

It is well known that a caveat associated with the use of the HP filter is that it could generate spurious correlation across the filtered variables. Thus, as an alternative filtering device we consider the use of annual growth rates when computing the correlations described above. The results, reported in the Online Appendix for brevity, are strongly robust. The average contemporaneous correlation between the commodity price index and real GDP is about 0.6, even slightly above the one documented before. Likewise, the strong cyclicality and leading property of the indices remains and is therefore not a spurious by-product of the particular filter used.[7]

[7] The Online Appendix also reports descriptive statistics with two additional detrending methods: log-linear and log-quadratic filtering. Results are strongly robust: the above mentioned correlation is 0.5 in both cases.

2.2. Common factors

A second dimension that we explore empirically is the presence of common factors in our measures of commodity price indices across EMEs. The evidence is presented in Fig. 2, which plots time series of the (cyclical) indices for each of the 13 EMEs in our sample (for a list, see Footnote 6). More formal principal component analysis is also conducted but is presented in the Appendix to economize space.[8]

This evidence leads us to a second stylized fact: 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 Fig. 2 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% 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.[9] Instead, the main reason comes from the fact that the international prices of various commodity goods commove strongly. 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's (SITC, fourth revision) one level aggregation and also extract their cyclical component.[10] The first principal component accounts for 55% of the variance in commodity prices across the five SITC categories.

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% of the variance of economic activity.

The Online Appendix presents the same set of results using growth rates as alternative filter. The results continue to point to strong comovement. If anything, they suggest an even higher degree of synchronization of the commodity price indices and real economic activity across EMEs. The first principal component accounts for 80% of the variance in the commodity price indices across the 13 EMEs and 76% of their respective variance of economic activity.[11]

Taken together, the stylized facts presented in this section contribute to further improving the understanding of the main patterns exhibited by EMEs' business cycles by shedding light on the strong comovement between aggregate macro variables in these economies and the prices of the commodities that they export. 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.[12]

3. Model

3.1. 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.[13] 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 stochastic productivity level. 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-contingent, 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 departure from the canonical framework. We model them with a dynamic factor structure that incorporates a latent common factor in addition to idiosyncratic shocks in order to capture the strong comovement across EMEs documented in the previous section. Fifth, the multi-country structure of our framework comes from jointly modeling 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 as well as perturbations to the world interest rate and the rest of the world's demand for non-commodity goods.[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). In the following subsections we formally describe the actions by each of the agents and their interactions in a representative jth EME in the model. 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 Online Appendix.

[8] We refer the reader of the black-and-white only printed version of our work to the Online Appendix where a colored version of Fig. 2 can be found, which also includes a legend to differentiate the 13 EMEs.

[9] 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%) and coal (15%), those of Peru are gold (28%) and copper (22%). The Appendix presents the specific shares of each of the commodities in our sample for all 13 countries in our dataset.

[10] 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.

[11] The same descriptive statistics are, respectively, 92% and 67%, using a log-linear detrending method; and 78% and 65% using a log-quadratic method (see the Online Appendix).

[12] In the working paper version of our work, Fernández et al. (2015), we document a third stylized fact that we omit here for space considerations. Using a large (unbalanced) annual panel covering 189 countries between the years 1960 and 2013, including the 61 EMEs defined earlier, we document how the share of commodities in total exports in the average EME is more than double that of advanced economies (see also the Online Appendix for details).

[13] The commodity sector is modeled as a special case of the enclave commodity sector in Catão and Chang (2013).

[14] We are thus abstracting from trade linkages across the EMEs in the model, mostly for tractability. While, in theory, trade across EMEs can potentially be relevant for explaining their business cycle comovement, in the empirical application we 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 emphasized.

3.2. Households

Households' lifetime utility is given by

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

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 $U(\cdot)$, $U(C_t, L_t) = \left[C_t - \psi^c L_t^{1+\gamma_c}/(1+\gamma_c)\right]^{1-\sigma}/(1-\sigma)$, where σ is the constant relative risk aversion coefficient and γ_c is the inverse of the Frisch elasticity of the labor supply.[15] Households maximize Eq. (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.[16]

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, $(p_t^c)^{-1}$.[17] The capital accumulation equation is

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

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:

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

3.3. Production of h goods

Firms in the economy produce h goods. They maximize profits, $\xi_t = p_t^h Y_t - w_t L_t - r_t^k K_{t-1}$, subject to a standard neoclassical production technology that uses capital and labor:

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

where Y_t denotes domestic output, z_t is the stochastic productivity level.

3.4. 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

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

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.

3.5. 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, modeled for country j as

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

[15] 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).

[16] 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 and extensions section will include a discussion of the implications of these modeling assumptions as well as the empirical evidence that justified them.

[17] 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: $P_t^{c*} = \phi\left(P_t^{h*}, P_t^{f*}\right) \simeq \tilde{\phi}\left(P_t^{h*}\right)$, where $\tilde{\phi}\left(P_t^{h*}\right)$ is linear in P_t^{h*} and $\tilde{\phi}(1) = 1$. 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 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, real GDP and the trade balance are defined as

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

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

3.6. Driving forces

3.6.1. 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 jth economy, capturing the extent to which changes in the common factor percolate to changes in $p_{j,t}^{Co}$. The idiosyncratic component in Eq. (15), $\varepsilon_{j,t}^{Co}$, is assumed to behave as an AR(1) process

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

where shocks $\nu_{j,t}^{Co}$ account for movements in $p_{j,t}^{Co}$ that are independent from the common factor.

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.[18]

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.[19]

3.6.2. Other driving forces

In addition to the two commodity price shocks described above, $\varepsilon_t^{f^{Co}}$ 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 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

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

where shocks $\varepsilon_t^{R^*}$ capture changes in the risk appetite of foreign investors. In turn, S_{jt} is defined as

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

where the term $\exp\left[\Omega_{j,u} \left(D_{j,t} - \bar{D}_j \right) \right]$ 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.[20] The latter follows an AR(1) process:

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

The remaining two driving forces are the world's foreign demand and the country-specific technology which we characterize as AR(1) processes[21,22]

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

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

3.7. 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.[23] At this point, however, it is instructive to describe, in a qualitative manner, the main mechanism at work.

Because the commodity sector is an endowment, real economic activity will only be affected in this economy insofar as the market

[18] There are, however, several more examples of global demand shocks and/or explanations for the comovement of prices of different commodity goods. The so-called financialization of commodity markets is one of them. For that reason we prefer the flexibility of modeling the common factor as an unobserved/latent factor.

[19] 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.

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

[21] 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.

[22] 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.

[23] See also the working paper version for a formal definition of the equilibrium in the model.

of home goods is affected. Looking at a (log-linearized version) of the market clearing condition for this market:

$$y_t = \underbrace{\left(C^h/Y\right) c_t^h}_{Effect\ 1} + \underbrace{\left(X^h/Y\right) x_t^h}_{Effect\ 2} + \underbrace{\left(C^{h*}/Y\right) c_t^{h*}}_{Effect\ 3},$$

where lower-case variables are log deviations from their steady state levels ($x_t = \ln(X_t/X)$), 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, and the relative sizes of consumption and the commodity sector in steady state. The latter 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 Fig. 3 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 outward, 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).[24] 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. Lastly, note that the opposite effect 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.[25]

4. Taking the model to the data: preliminaries

4.1. 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.[26] We do so mainly for two reasons. First, we want to have a reduced number of countries for tractability. Second, as will be shown below, the estimation of the model and calibration of its steady state present some further data requirements that are binding for several of the 13 countries studied in Section 2 which prevent us from carrying the estimation for the entire sample of EMEs.

This pool of four countries is representative of the type of economies modeled above. They are all well-known commodity exporters, with a median commodity export share of 35.4.[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% of the variance in the commodity price index and real output across the four economies, respectively. 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 (see Online Appendix). Lastly, as depicted in Fig. 1, this group of countries reproduces well the average cyclicality documented in the larger pool of 13 EMEs.

4.2. 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 1. Table 2 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% 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 2. This disciplines the model as it assigns realistic relative weight to f and h goods.[28]

The parameters $\bar{D}$ and $\bar{Co}$ are calibrated so as to match the long run shares of external debt and commodity exports. This is crucial as

24 The reader may wonder about the shift in the labor supply curve depicted in Fig. 3, panel d, given that GHH preferences are used which, in principle, eliminate the income effect. That indeed is the case in one-good models. In models with multiple goods, as the one in our work, 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 Online Appendix illustrates this.

25 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 increase in real income.

26 The four countries account for roughly half of Latin American output.

27 Specific commodity export shares are Brazil (17.9); Chile (69.7); Colombia (28.6); and Peru (60.5) (see the Online Appendix for further details).

28 In the Online Appendix, we test the robustness of our results to alternative higher parameterization values for η_c and η_x, while simultaneously modifying α_c and α_x so as to keep the same steady state shares targeted. While, in principle, raising this elasticity diminishes the income effect associated to commodity price shocks, in practice the quantitative effects are modest.

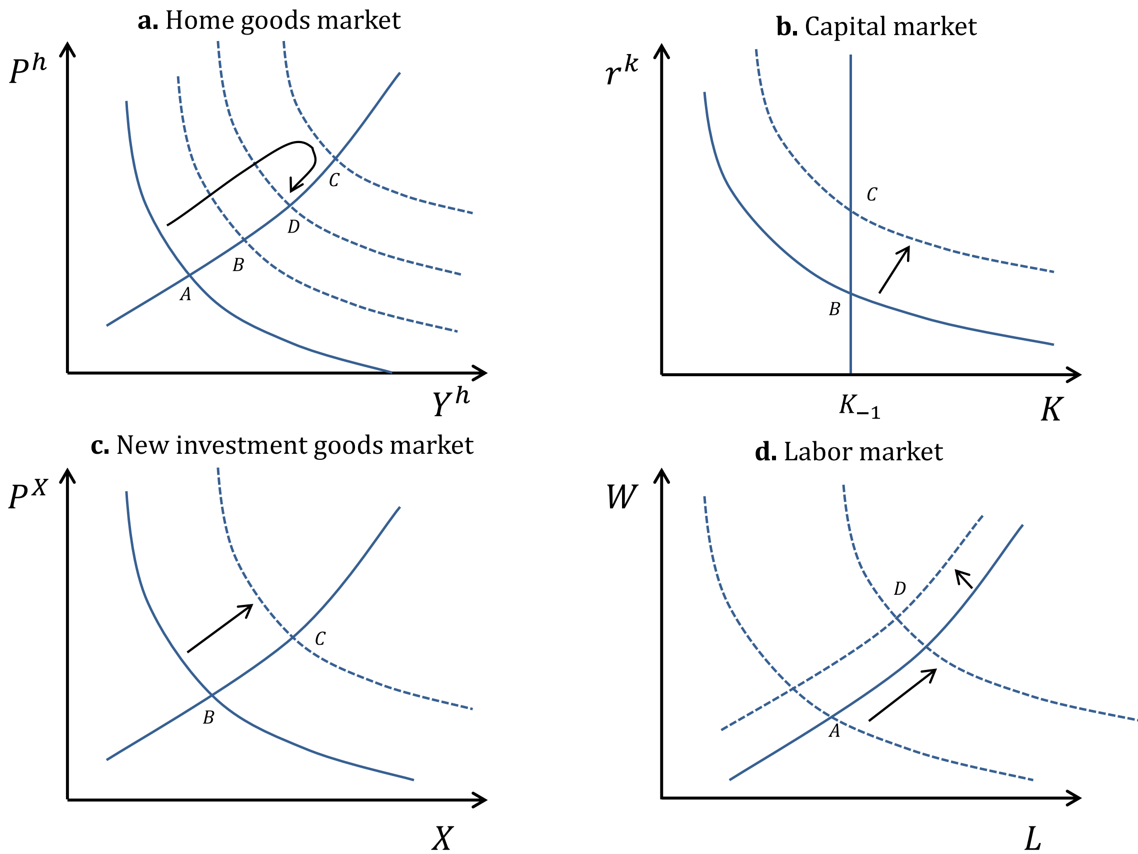

Fig. 3. The propagation mechanism of a positive commodity price shock in the model.
Note: Each panel depicts, qualitatively, movements in supply and demand curves in four markets in the model following a positive shock to commodity prices.

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

Table 1
Calibrated parameters.

Parameters		Brazil	Chile	Colombia	Peru
Calibrated parameters common to all countries					
$\overline{R^*}$	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 annual rate (%)	10.00	10.00	10.00	10.00
ϵ_e	Price elasticity of exports	1.18	1.18	1.18	1.18
$\bar{Y}^*$	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
$\overline{Co}$	Steady state level of commodity endowment	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.

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. 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. This guarantees that the size of 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, as well as the relative size of total investment. As presented in Table 2, these dimensions are also among those that the model can perfectly match for the countries considered.[29] Also, 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. Lastly, the model accomplishes a satisfactory fit of the relative shares of consumption, imports and exports to GDP.[30]

4.3. Bayesian estimation

The remaining 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 $\left\{\phi_{Co}, \rho_j^{Co}, \rho_j^{z^R}, \rho_{R^*}, \rho_{Y^*}, \rho_j^z\right\}$, the standard deviation of their shocks $\left\{\sigma^{f^{Co}}, \sigma_j^{Co}, \sigma^{R^*}, \sigma_j^{z^R}, \sigma^{Y^*}, \sigma_j^z\right\}$; the loading factors in the dynamic common factor $\left\{\omega_j^{Co}\right\}$,[31] and those that determine the cost of adjusting the capital stock $\{a_j\}$. Fairly agnostic priors are used, as depicted in Table 3.

The short-run dynamics of the model are obtained using well-known methods. In particular, 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. Thus both the model's solution and data are expressed in terms of cyclical deviations when the Bayesian estimation is performed.[32]

The measurement equations in the state space representation of the model use 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' 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 through measurement equations.[33] 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.[34] 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, GDP_t/p_t^c.

5. Estimation results

5.1. Posterior distributions and common factors

Posterior distributions of the estimated parameters are reported in Table 3.[35] Overall the posterior densities are considerably different from the loose priors that we choose, implying that the data contain information on the estimated parameters. The AR(1) coefficient in the common factor of commodity prices, ϕ_{Co}, has a posterior mean of 0.73. Importantly, shocks to this common factor display a large estimated standard deviation, $\sigma^{f^{Co}}$, of 6.31%, 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 Peru; while the estimated standard deviation σ^z ranges, in percentage terms, between 0.84/0.89 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 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).

29 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.

30 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 2, 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).

31 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. Further tests based on rank conditions inform us that the vector of parameters to be estimated is locally identifiable given the set of observables considered (see further details in the Online Appendix).

32 The solution is found in DYNARE. The Online Appendix presents the list of stationary equations as well as further technical details of the solution.

33 The Online Appendix presents the set of measurement equations.

34 Lack of data availability on some of the observables prevents us from covering a longer historical period. Nonetheless, in the working paper version of our work, we explore the estimation of the dynamic factor model outside the rigid structure imposed by the DSGE on the data. We estimate the dynamic factor model (DFM) structure, Eqs. (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. The end of the Section on Robustness and extensions discusses the results of this case.

35 The Online 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.

Table 2
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	82.007	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 and Soto (2007); Colombia see González et al. (2011); for Peru see Castillo et al. (2009).

Table 3
Priors and posteriors of estimated parameters.

Parameters	Prior			Posterior				
	Type	Mean	S.D.	Mode	S.D.	Mean	90% HPD interval	
Global								
ρ_{Y^*}	Beta	0.500	0.150	0.8099	0.0471	0.8048	0.7322	0.8833
ρ_{R^*}	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
σ^{Y^*}	Invg	0.013	Inf	0.0061	0.0006	0.0062	0.0053	0.0071
σ^{R^*}	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.

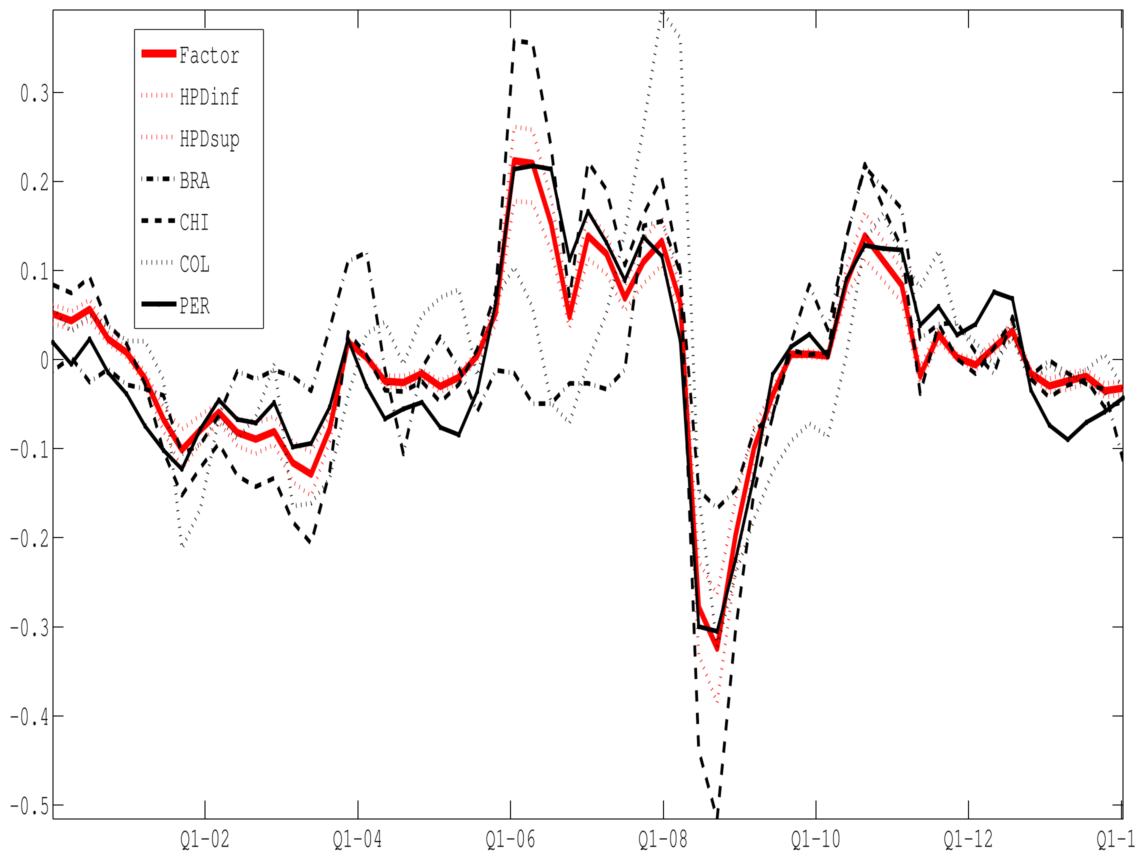

Fig. 4. Estimated common factor in country-specific commodity price indices.
Note: The common factor (in solid/red) is the latent variable obtained from the Kalman filter smoothing evaluated at the mean of the posterior distribution. HPDinf and HPDsup stand for the 10th and 90th percentiles of the posterior distribution, respectively. The remaining lines depict the commodity price indices of the four countries studied in the estimated model.

Lastly, Fig. 4 presents the time series evolution of the common factor, f^{Co}, which we back out from the Kalman smoother evaluated at the posterior mean, together with the 90% 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 deviations from trend sometimes above 30 percentages points – hence the allusion of a roller coaster ride 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.

5.2. Business cycle drivers

We now use the estimated model to document the main business cycle drivers. Our first tool to accomplish this is the (unconditional) forecast error variance decomposition (FEVD) of real income across the four countries, summarized in Table 4, panel (a). The panel presents the contribution of each of the six structural shocks in each country's real income variance, using posterior means. 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%, though with considerable variability across countries, ranging from 27.5% in Brazil, up to 77.1% 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.[36] 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 non-trivial 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 2).[37]

[36] The admittedly high number for Chile may signal elements that our simplified structural setup is missing. Among others, we do not incorporate a countercyclical fiscal rule as the one that Chile has had in place for most of the period which, in practice, may have reduced the extent to which commodity price fluctuations affect the macroeconomy.

[37] The working paper version of this work presents FEVD results when conditioning the exercise at alternative forecast horizons of one, four, and twelve quarters. The role of commodity price shocks is positively related to the forecasting horizon in the FEVD decomposition exercise. This is related to the strong persistence in commodity prices, as documented earlier. The working paper also contains various counterfactual experiments, namely standard deviation (S.D.) of real income predicted by the model if commodity price shocks are turned off, finding a considerable reduction in macro volatility across countries.

Table 4
Forecast error variance decomposition of output: baseline and alternative models.

Shocks	Brazil	Chile	Colombia	Peru	Mean	Median	Shocks	Brazil	Chile	Colombia	Peru	Mean	Median
	(a) Baseline							(b) Countercyclical spreads					
Commodity price shocks							Commodity price shock						
ν^{Co}	16.3	0.5	17.6	4.9	9.8	10.6	ν^{Co}	24.4	0.6	18.0	5.6	12.2	11.8
ε^{fCo}	11.1	76.6	25.8	35.5	37.3	30.7	ε^{fCo}	17.1	74.0	27.7	37.6	39.1	32.7
Sum	27.5	77.1	43.5	40.4	47.1	42.0	Sum	41.5	74.6	45.7	43.2	51.3	44.5
All other shocks							All other shocks						
ν^R	30.0	0.03	4.43	0.3	8.7	2.4	ν^R	24.8	0.1	4.9	0.5	7.6	2.7
ε^z	22.0	22.9	45.4	58.3	37.2	34.2	ε^z	17.6	25.2	41.1	54.7	34.7	33.2
ε^{R^*}	20.3	0.03	5.8	0.92	6.8	3.4	ε^{R^*}	16.0	0.1	7.5	1.5	6.3	4.5
ε^{Y^*}	0.19	0.01	0.9	0.1	0.3	0.1	ε^{Y^*}	0.1	0.	0.8	0.1	0.3	0.1
Sum	72.5	22.9	56.5	59.6	52.9	58.1	Sum	58.5	25.4	54.3	56.8	48.8	55.6
	Marginal likelihood			4530.9				Marginal likelihood			4502.40		
	(c) Correlation between external forces							(d) Basic RBC model (no commodities)					
Commodity price shocks							Commodity price shock						
ν^{Co}	16.5	25.6	17.8	4.8	16.2	17.2	ν^{Co}	–	–	–	–		
ε^{fCo}	7.9	64.3	18.5	27.0	29.4	22.8	ε^{fCo}	–	–	–	–		
Sum	24.5	89.9	36.3	31.8	45.6	34.1	Sum	–	–	–	–		
All other shocks							All other shocks						
ν^R	30.3	0.6	4.4	0.3	8.9	2.5	ν^R	42.4	0.	3.8	0.3	11.6	2.1
ε^z	22.6	0.	47.1	62.5	33.1	34.9	ε^z	33.3	100.0	88.5	99.0	80.2	93.8
ε^{R^*}	20.5	0.	6.0	1.0	6.9	3.5	ε^{R^*}	24.1	0.	6.6	0.7	7.9	3.7
ε^{Y^*}	2.1	9.4	6.1	4.4	5.5	5.3	ε^{Y^*}	0.3	0.	1.1	0.1	0.4	0.2
Sum	75.6	10.1	63.7	68.2	54.4	66.0	Sum	100.0	100.0	100.0	100.0	100.0	100.0
	Marginal likelihood			4535.70				Marginal likelihood			4410.60		

Note: The panels report real output's unconditional forecast error variance decomposition (FEVD), calculated at the posterior mean for four alternative reduced models: panel (a): Benchmark Model; panel (b): Countercyclical spreads; panel (c): Correlation between common factor in commodity prices and foreign output; panel (d): No commodity shocks. Shocks are as follows: Idiosyncratic commodity price shock (ν^{Co}); Common factor in commodity price shock ($\varepsilon^{f^{Co}}$); Spread shock (ν^R); Domestic productivity shock (ε^z); World riskless interest rate shock (ε^{R^*}); World demand shock (ε^{Y^*}). Marginal likelihoods are reporte at the bottom of each panel, they are computed using Geweke's modified harmonic mean.

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 Fig. 5, 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 ($\varepsilon^{f^{Co}}, \nu^{Co}$); (iii) *Spread Shocks* which alludes to the idiosyncratic risk premia shock (ν^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, though in Chile 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 Fig. 4. 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, commodity price shocks turned positive in the recovery from the 2009 recession, supporting a swift recovery. They also helped counterbalance the negative contributions of domestic shocks, mostly in Brazil and Chile.

5.3. *Quantifying the mechanism*

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 Fig. 6. Numbers reported are percentage deviations from steady state values.

Chile stands out as the country where more real effects following the shock are observed -note the change in scale used for this country. This is consistent with earlier results 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 Fig. 6 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 quantify the three effects that were described before in the context of Fig. 3. 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. The second effect, the response of investment, is also pronounced and hump-shaped.

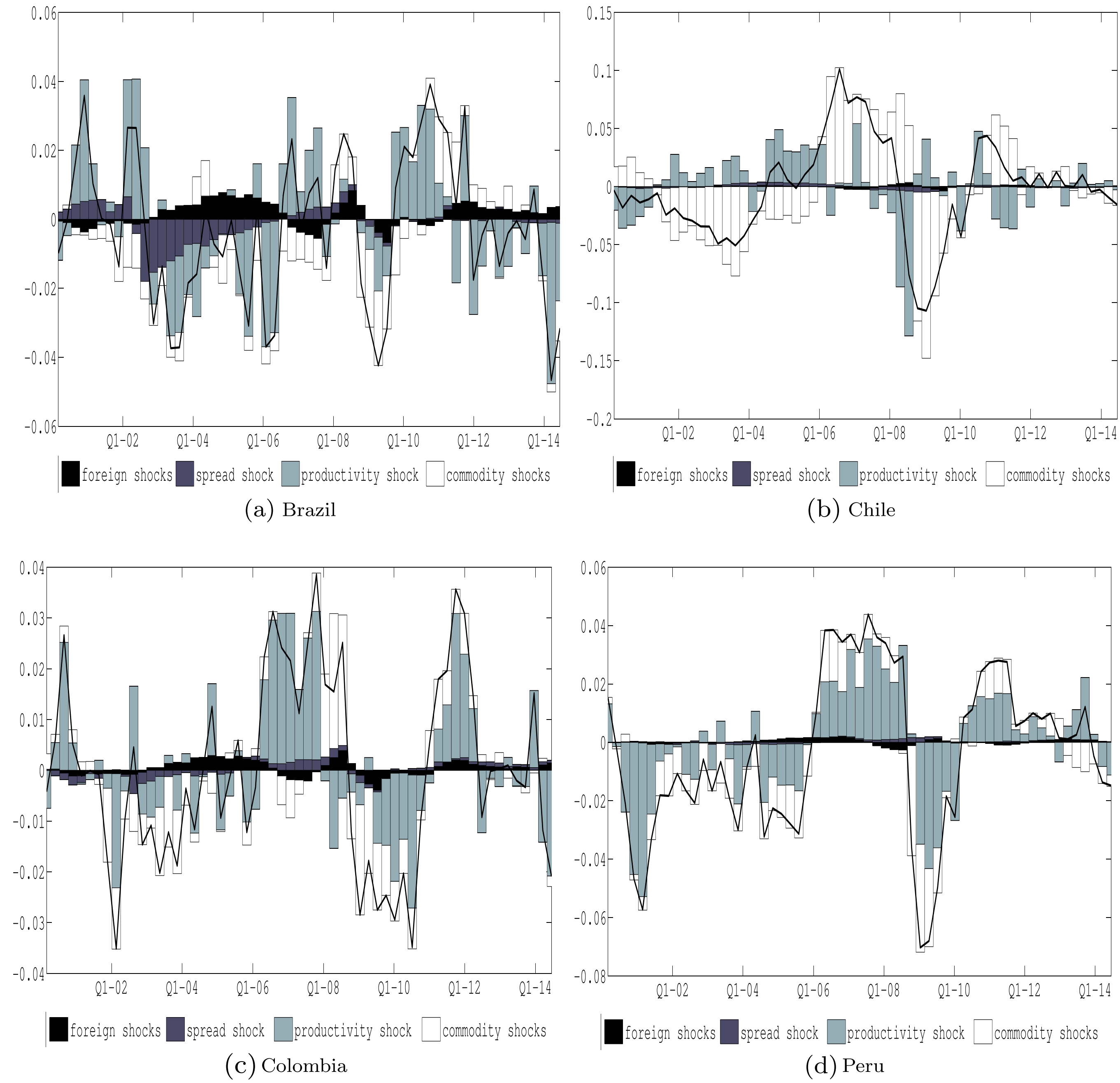

Fig. 5. Historical decomposition of real income.
Note: Groups of shocks are as follows: Foreign shocks are world riskless interest rate shock (ϵ^{R^*}) and world demand shock (ϵ^{Y^*}); spread shock is (ν^R); productivity shock is (ϵ^z) ; and commodity shocks are idiosyncratic commodity price shock (ν^{Co}) and common factor in commodity price shock $\left(\epsilon^{f^{Co}}\right)$.

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. 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.

5.4. Model's performance

We now study the performance of the model when accounting for various dimensions of the data. Table 5 focuses on 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

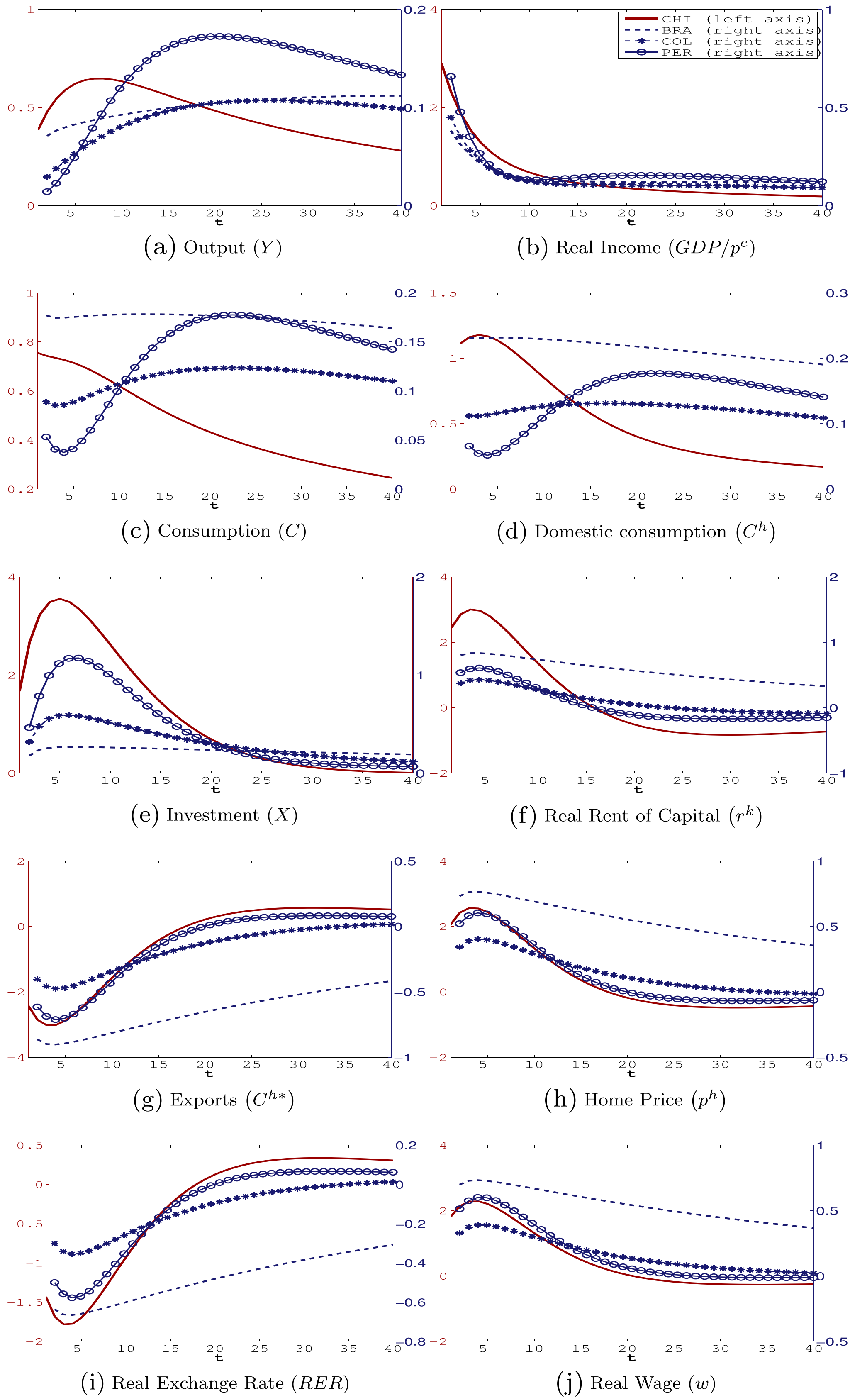

Fig. 6. Impulse responses to a common shock in commodity prices.
Note: The subplots present the impulse response functions following a common factor shock in commodity prices. Units are percentage deviations from steady state levels. The unit of time in the horizontal axis is a quarter.

Table 5
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*)		4530.9		

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% 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 likelihoods for both models, in the lower panel, are computed using Geweke's modified harmonic mean.

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*′′), i.e. we set $\omega_j^{Co} = 0$, $\forall j$, and reestimate the model. The baseline model does a good job when bringing the average simulated cross-correlations of output across countries, 0.32, close to that in the data, 0.45. 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 5.

Fig. 7 presents further evidence on the model's performance by focusing on the serial correlation of macro variables with commodity price indices studied earlier (Fig. 1). 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.[38]

6. Robustness and extensions

6.1. 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 (2017), comparing DSGE to SVAR models may be a good way to assess the performance of DSGEs when quantifying the role of terms of trade fluctuations on business cycles. Usually SVARs are believed to "let the data speak" more freely relative to DSGEs, which impose considerable more structure on the data. Their approach, however, does not look particularly at commodity price shocks as ours does.

We first estimate the same SVAR in Schmitt-Grohé and Uribe (2017) with data for the same four economies on which we estimate the DSGE model, except that we replace the terms of trade variable 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 $x_t = \left[p_t^{Co}, tb_t, y_t, c_t, i_t, rer_t\right]$, 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 distributed with mean zero and variance-covariance matrix Σ. We impose the same two identifying restrictions in Schmitt-Grohé and Uribe (2017). First, we assume that p_t^{Co} is exogenous and that it follows an AR(1) process. Second, we assume that $u_t = \Pi\epsilon_t$, where ϵ_t is a white noise, mean zero vector with identity variance-covariance matrix, and Π is picked 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.

Fig. 8 plots the time series of y_t in the data against the predicted series for this variable from the DSGE and the SVAR models 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 been realized. The main result coming out of this evidence is that the implications of the SVAR are very well aligned with those from the DSGE, and both point to a considerable role of commodity price shocks. The correlation between the two predicted series is equal to or above 0.9 for all countries but Chile. In the latter case, however, the DSGE model tracks much more closely the observed data than the SVAR. The strong resemblance between both models is further confirmed by comparing the two models' implied unconditional output variance share associated to commodity prices. 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%, while that from the DSGE is 47/42% (see Online Appendix for the full set of results).

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 (2017) who find that only a median share of 10% of output's variance can be associated with terms of trade shocks. This difference has been further studied in subsequent work by Fernández et al. (2017). Indeed, they have documented on a much wider panel of 138 countries spanning the last half century of data that world shocks mediated by commodity prices account for about one third of movements in aggregate activity. This number is three times as large as those obtained in single world price specifications (e.g. terms of trade), suggesting

[38] The working paper version of our work further documents the model's relatively good performance along other second moments that are more standard in the literature and compares them to the data (i.e., volatility, serial autocorrelation, etc.).

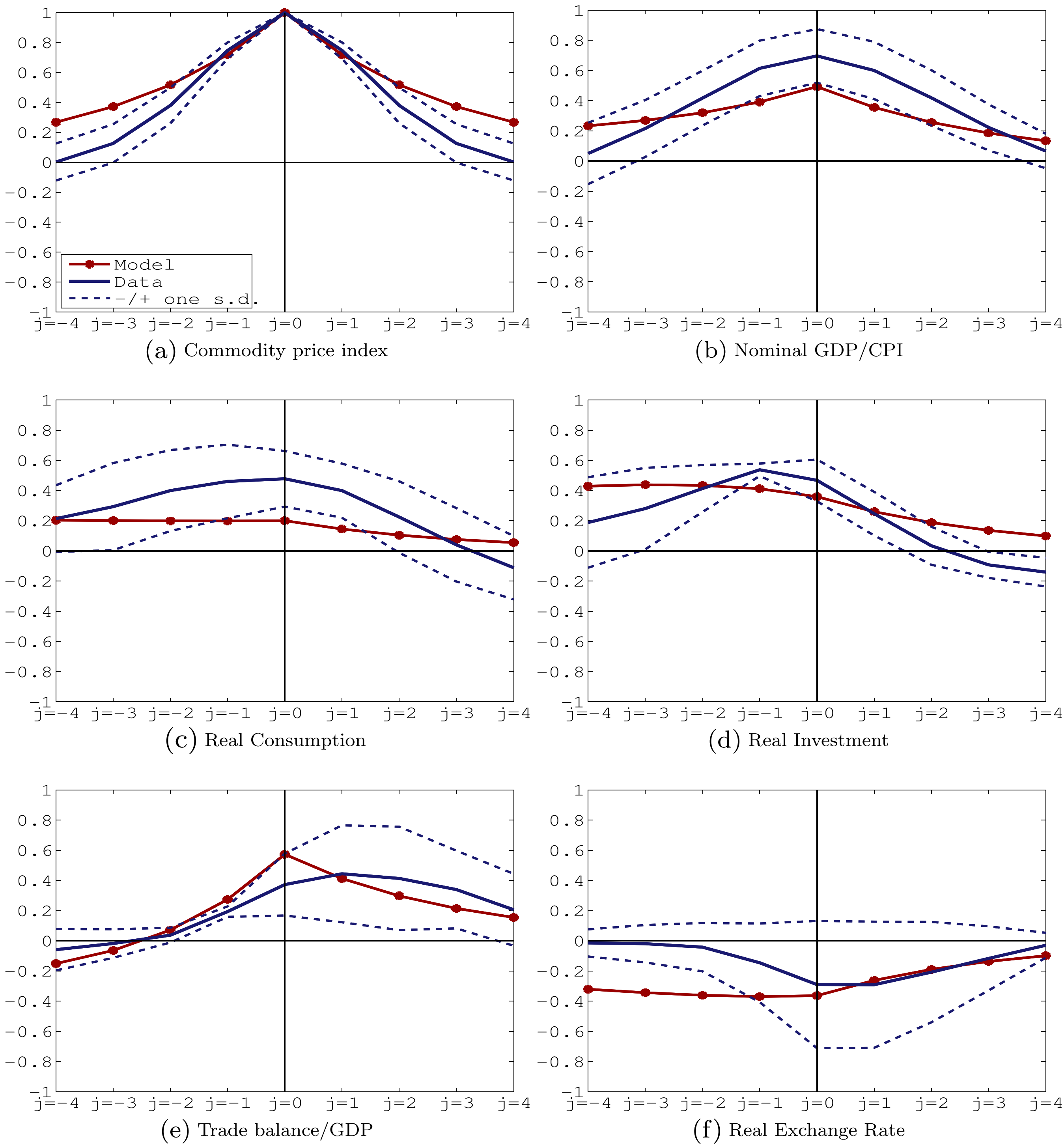

Fig. 7. Unconditional serial correlations with commodity prices: data and model.
Note: Data corresponds to the simple average across the four countries used in the estimation of the model. Model-based correlations are computed with the estimated benchmark model.

that one-world-price specifications significantly underestimate the importance of world shocks for domestic business cycles.

6.2. Countercyclical spreads

As argued by Neumeyer and Perri (2005) and Uribe and Yue (2006), spreads may react to country fundamentals and vice versa, thereby generating an amplification mechanism for real shocks. More recently, Shousha (2016) and Drechsel and Tenreyro (2017) have postulated, within a DSGE environment, that commodity prices ought to be included among those fundamentals. 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

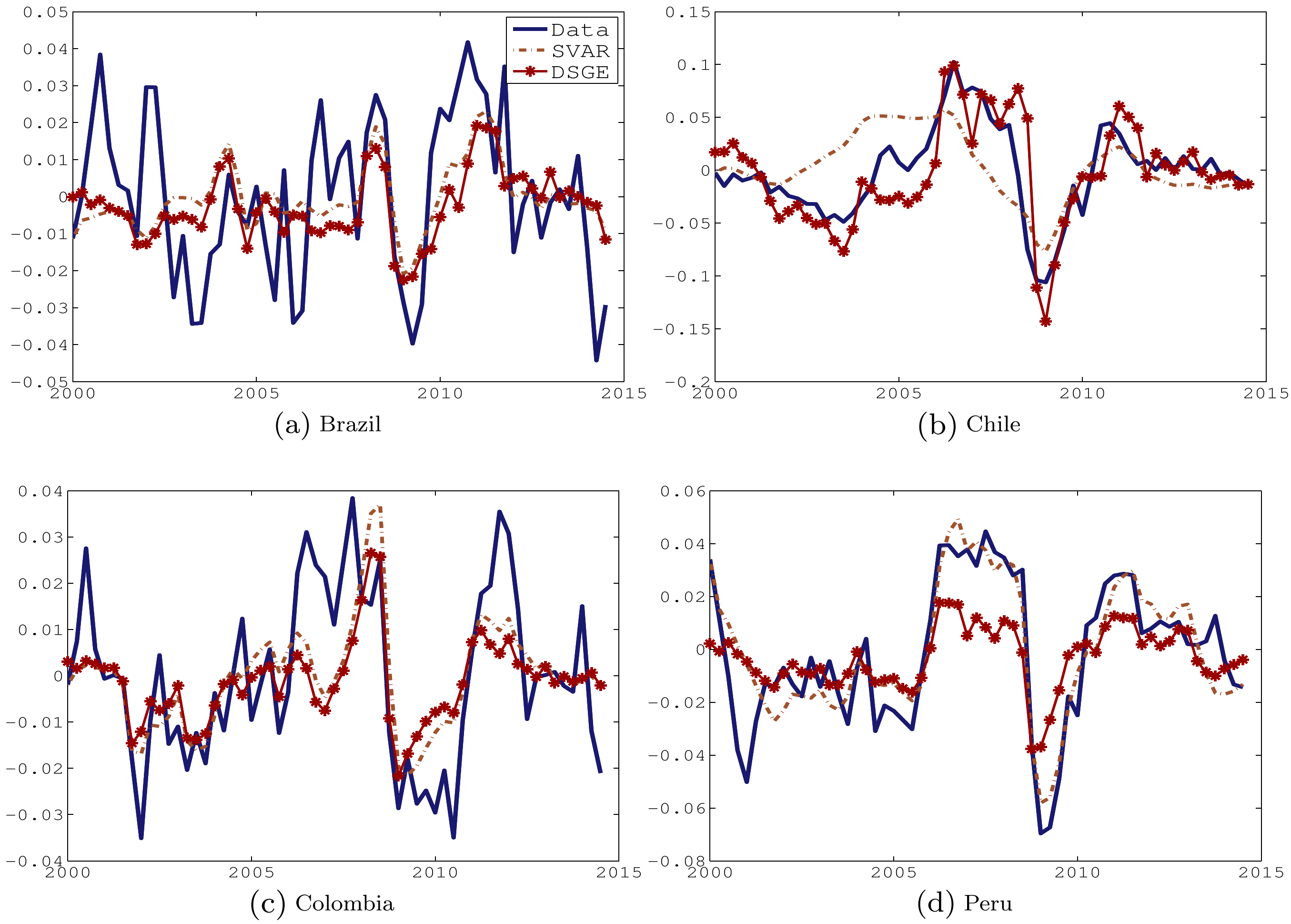

Fig. 8. Comparison with an SVAR model.
Note: Solid/blue lines represent the filtered real income series. Starred/red and dotted/orange lines report the simulated real GDP series from the DSGE and the SVAR models respectively when only commodity price shocks are on. The SVAR is estimated with the cyclical 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 it follows an autoregressive process of order one and by applying a lower triangular Cholesky decomposition on the variance-covariance matrix of the residuals (see text for details).

country fundamentals, in the form of TFP and commodity prices. Formally, Eq. (20), in log deviations, now becomes

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

where a hat, "^", denotes log deviations from steady state values; and $\left\{\eta_j^z, \eta_j^{Co}\right\}$ 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 Eq. (23) is not new and follows Neumeyer and Perri (2005). In addition to this, we include (expected) changes in commodity prices, capturing the idea that they 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.[39]

[39] 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 Beltrán (2015) have related spreads to commodity prices using the financial accelerator model.

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 on the demand for labor by home-good producers. Formally, we now assume that the profit function of the firm is

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

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 (posterior results are presented in the Appendix). We report the unconditional FEVD for output in panel (b) of Table 4. While this extension raises the importance of commodity price shocks, thereby validating the link between commodity prices and spreads as another propagation mechanism for commodity price fluctuations, the quantitative effects do not appear to be strong. The median output variance share associated to these shocks raises from 42 to 45% across the countries considered.

6.3. 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 allow for the common factor in commodity prices, f_t^{Co}, to be correlated with external demand, Y_t^*.[40] Formally, we model this by modifying the stochastic processes for these two driving forces, Eqs. (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 $\left\{\varepsilon_t^{Y^*}, \varepsilon_t^{f^{Co}}\right\}$ are structural and independent disturbances whose variance we estimate. 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 panel (c) of Table 4. Posterior results are presented in the Appendix. The main result is that the large share of commodity price shocks remains, as in 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%. 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 three tenths 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.

6.4. 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 more similar to the benchmark small open economy RBC model. Here, again for the sake of space, we document only the FEVD of output in panel (d) of Table 4 and the marginal likelihood of the reduced model.[41]

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 findings reconcile our baseline 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 4530.9 to 4410.6. Indeed, as documented in the working paper version, 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 discussed earlier in the context of Fig. 3.

6.5. Alternative filtering

It was argued that the stylized facts presented in Section 2 were robust to the type of filtering technique used to extract the cyclical component in the data. We now extend this robustness analysis to the structural estimation of the model by modifying the measurement equations in such a way that deviations in both model and data are expressed in terms of growth rates, one of the alternative filtering methods used in Section 2. In doing so we also address two potential concerns associated with our benchmark structural specification. On one hand, we make sure that deviations in both model and data are expressed using the same filter (growth rates). On the other hand, by using growth rates, no future information is used when detrending the data which is later used in the estimation, thereby not violating the information assumption in the Kalman filter used for the Bayesian estimation of the model, which uses current and past information only.

Panel (b) in Table 6 reports the results of the FEVD of output from the structural model when the observable data used in the estimation are filtered using growth rates.[42] For comparison, panel (a) reports again the baseline results. The new FEVD results 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 50%. In

[40] 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, as we do in this extension.

[41] 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 household's commodity revenue is now deterministic and equal to $p^{Co}\bar{Co}$ 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.

[42] The Online Appendix 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. Note that this makes the marginal likelihood of this model not directly comparable to that in the benchmark model.

Table 6
Forecast error variance decomposition of output: further alternative models.

Shocks	Brazil	Chile	Colombia	Peru	Mean	Media	Shocks	Brazil	Chile	Colombia	Peru	Mean	Media
			(a) Baseline							(b) Benchmarck model in first differences			
Commodity price shocks							Commodity price shock						
ν^{Co}	16.3	0.5	17.6	4.9	9.8	10.6	ν^{Co}	15.5	1.8	5.0	0.8	5.8	3.4
$\varepsilon^{f^{Co}}$	11.1	76.6	25.8	35.5	37.3	30.7	$\varepsilon^{f^{Co}}$	44.7	56.6	71.2	76.4	62.2	63.9
Sum	27.5	77.1	43.5	40.4	47.1	42.0	Sum	60.1	58.4	76.1	77.3	68.0	68.1
All other shocks							All other shocks						
ν^{R}	30.0	0.03	4.43	0.3	8.7	2.4	ν^{R}	22.1	0.0	2.7	0.1	6.2	1.4
ε^{z}	22.0	22.9	45.4	58.3	37.2	34.2	ε^{z}	10.9	41.6	18.9	22.4	23.5	20.7
ε^{R^*}	20.3	0.03	5.8	0.92	6.8	3.4	ε^{R^*}	6.6	0.0	1.4	0.1	2.0	0.8
ε^{Y^*}	0.19	0.01	0.9	0.1	0.3	0.1	ε^{Y^*}	0.2	0.0	0.9	0.1	0.3	0.2
Sum	72.5	22.9	56.5	59.6	52.9	58.1	Sum	39.9	41.6	23.9	22.8	32.1	31.9
	Marginal likelihood			4530.9				Marginal likelihood			4421.8		

Note: The panels report real output's unconditional forecast error variance decomposition (FEVD) calculated at the posterior mean for two alternative models: panel (a): Benchmark case with Hodrick-Prescott filtered data; panel (b): Benchmark model with data in first differences. Marginal Likelihood are computed with Geweke's modified harmonic mean. See footnote in Table 4 for description of the shocks.

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%, relative to 42% in the benchmark case. 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.[43]

6.6. General discussion

Despite the various extensions and robustness considered, there are several others that we are leaving out. An obvious one is the presence of commodity production. Because we model the commodity sector as an endowment enclave, we abstract from any supply channels that may operate, namely those involving the decision to produce and demand more labor and capital, which may be relevant for, respectively, labor-intensive and capital-intensive commodity exporters. We decided to do so for two reasons. On one hand, the scant high frequency data on production of commodities across the EMEs prevent us from carrying out a full-information estimation of a model with commodity production, similar to the one we do in the benchmark case. On the other hand, such a scant data reveal no evidence that, *in the short run*, volumetric measures of commodity production in the EMEs considered commove with changes in the international market prices of these commodities (see Online Appendix).[44] In the medium to long run, evidently, such relationship must exist and we abstract from modeling it. Doing so would not be a trivial endeavour as one would have to take into account the complex institutional elements of the "natural resource curse" in Frankel (2010). In a recent contribution, however, Caputo and Irarrázabal (2017) investigate the role of commodity price shocks in an environment of commodity production calibrated to Chile. They also obtain that a positive shock to the commodity price leads to an increase in both non-commodity output and employment, mainly via the effect over total investment, in line also with the income channel that our model emphasizes.

Another potentially relevant channel that our setup is leaving out is the role of monetary and fiscal policy. For the particular case of fiscal policy, shutting down this channel may be a nontrivial simplification because EME governments typically either own (at least part of) the commodity exporting firms or tax them heavily to finance spending elsewhere. There is, however, a well established fact that both types of policies tend to be procyclical in EMEs (Kaminsky et al., 2005). We conjecture, therefore, that adding such channels may end up reinforcing the central role of commodities for the business cycle of these economies via the multiplier effect that both policies would have on the cycle. But we acknowledge that more work remains in terms of disentangling how important, quantitatively, such a propagating force is.

Other potential shortcoming of our baseline setup is that, due to lack of availability, we have only used data post-2000 when estimating the parameters of the dynamic factor in commodity prices. This follows from our estimation strategy, which included the parameters of the dynamic factor structure in the full set of parameters from the DSGE model that are estimated using full-information methods. One might also wonder how different the results of the estimation of the dynamic factor model would be if it is done outside the rigid structure imposed by the DSGE on the data. In the Online Appendix we present results from an additional extension where we estimate the dynamic factor model (DFM) structure in isolation from the structural model, using only data from country-specific commodity price indices. This allows us to use data as far back as 1980 when our time series on commodity price indices begin. The most noticeable result that can be drawn from this extension is that the new common factor tracks 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.

Last, but not least, our baseline setup has abstracted from the possibility that commodity prices display a unit root, implying that small shocks can have long-lasting effects. The profession does not seem to have reached an empirical consensus, though recent evidence appears to favor the rejection of such hypothesis, at least when one accounts for the presence of structural breaks (Mariscal and Powell, 2014). Still, at least in theory, the possibility of a unit root process opens up various other channels by which commodity prices can impact the macroeconomy of EMEs. One of these will likely be the way that long run price expectations of commodity prices can have an effect over the supply side decisions of economic agents. Yet, proper modeling of this additional channel would require institutional features as well as a more in depth modeling of the way agents form expectations about the long run behavior of commodity prices, particularly how they extract information

[43] We also re-estimated the baseline model using two additional detrending methods: (i) a log-linear filter; and (ii) a log-quadratic filter. For the sake of space we report the entire set of variance decomposition results in the Online Appendix. We note here, however, that results are quite robust in terms of the preponderant role of commodity price shocks when accounting for business cycles. With either of the two filters, the median share of these shocks when accounting for the variation in real GDP is above a third.

[44] This is also in line with recent evidence in the crude oil market that points to low elasticity of supply, at least in the conventional (non-shale) industry (Bjørnland et al., 2017).

about the transitory and permanent shocks that perturb commodity prices (see, Fornero and Kirchner, 2014; Leduc et al., 2016), to name but a few, which we consider is far beyond the scope of this paper. Nevertheless, the Online Appendix postulates a simple extension with non-stationary country-specific commodity price indexes. Overall, the results are consistent with those in the baseline case in terms of the preponderance of commodity price shocks. If anything, the results change in the direction of assigning an even higher share of real income's variance to commodity shocks.

7. 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 quantitatively relevant and that their sources can mostly be traced back to exogenous changes in the prices of the commodity goods that these economies export and can be viewed, through the lens of the simple theory that we provide, 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 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.

Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.jinteco.2017.11.008.

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