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ABSTRACT

On the Exposure of the BRIC Countries to Global Economic Shocks

The financial crisis led to a deep recession in many industrial countries. While large emerging countries recovered relatively quickly from the financial crisis, their performance deteriorated in the recent years, despite the modest recovery in advanced economies. The higher divergence of business cycles is closely linked to the Chinese transformation. During the crisis, the Chinese fiscal stimulus prevented a decline in GDP growth not only in that country, but also in resource-rich economies. The Chinese shift to consumption-driven growth led to a decline in commodity demand, and the environment became more challenging for many emerging markets. This view is supported by Bayesian VARs specified for the BRIC (Brazil, Russia, India, and China) countries. The results reveal a strong impact of international variables on GDP growth. In contrast to the other countries, China plays a crucial role in determining global trade and oil prices. Hence, the change in the Chinese growth strategy puts additional reform pressure on countries with abundant natural resources.

JEL Classification: F44, E32, C32

Keywords: business cycle divergence, Chinese transformation, Bayesian VARs

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1. Introduction

Due to their fast catching-up in income, the BRIC countries (Brazil, Russia, India, and China) account for 30 percent of world GDP at the current edge, as expressed in PPPs (Figure 1). With a weight of about 56 percent of GDP, the BRIC evolution is dominated by the Chinese economy. The BRICs have been the primary source for global GDP growth before the financial crisis until the first years thereafter. The rebound from the crisis started earlier in many emerging markets, evolved much faster than in advanced economies and was often characterized by a V-shaped pattern of output growth (Didier et al., 2012). However, despite a modest recovery in the industrial countries seems to be on the way, GDP growth in the BRICs started to decline in the most recent years. Although differences across countries are striking, the slowdown is synchronized to some extent. While the acceleration of output is still high in India, the Chinese economy experienced lower growth, and countries like Brazil and Russia entered even a recession. In terms of the BRIC aggregate, growth fell not only below the post-crisis peak of 2010/11, but even below the rates in the pre-crisis decade (Figure 2). Due to the increasing role of emerging markets in the global economy, a stronger slowdown could constitute a major risk for global growth in the years ahead.

-Figures 1 and 2 about here-

External conditions are often blamed for this development. For instance, Almansour et al. (2014) argue that global factors account for one half of the variance of emerging markets growth. Tailwinds that supported the former catching up, like the vast acceleration of world trade, rocketing commodity prices and easy financial conditions did not continue and will probably not improve over the next years. The BRICs' slowdown might be traced back to the long-lasting effects of the crisis that have been temporarily whitewashed by expansionary policy measures. In particular, the Chinese authorities launched a huge fiscal program to compensate for the reduction in exports over the crisis period (Dreger and Zhang, 2014). That strategy prevented a sudden drop of output growth not only in China, but also in countries with strong exposure to natural resources. In the following years, the transformation towards consumption-driven growth lowered the Chinese demand for commodities, implying that external conditions became more challenging for other countries since then. In this sense, the change in

the Chinese growth strategy contributed to a higher divergence of international business cycles. The fiscal stimulus had a major impact on emerging markets, most notably on countries with abundant natural resources. In contrast, the effects on GDP growth in the main advanced economies have been relatively modest, probably with the exception of Japan (Dreger and Zhang, 2014).

This paper investigates the relative role of foreign factors for GDP growth in the BRIC economies. Foreign variables are captured by commodity prices, world trade and international financial conditions. Global shocks can be disseminated through various channels, like (a) the fiscal policy stance, as lower commodity prices put higher consolidation pressure on public budgets, (b) tighter monetary policy to combat capital outflows caused by a higher risk attitudes of investors, and (c) the real exchange rate, as the real depreciation of the BRIC currencies generates more inflation through higher import prices.

Our Bayesian VAR analysis suggests that the BRIC countries are heavily affected by the global economic conditions, albeit to a different degree. Commodity price movements are able to explain the downturn in Brazil and Russia to a huge extent. India is less affected by commodity markets, but the lower expansion of global trade depresses the country's GDP growth. Prices for raw materials and world trade are both relevant for output growth in China. However, in contrast to the other countries, the relationship appears to be bidirectional, as China heavily affects the global economy. Therefore, China is an important driver for economic growth in other emerging markets. In former years, China's investment-oriented strategy boosted emerging markets via higher commodity demand. The strong expansion provided a buffer to emerging markets during the period of the financial crisis. In the following years, the slowdown in China softened output growth at a global scale.

The remainder of the paper is organized as follows. Section 2 reviews previous papers on the catching-up of the BRICs and the current slowdown of growth. Section 3 discusses our Bayesian structural VAR model and its specification. Section 4 presents the empirical findings about China's role in determining global variables and the impacts of external shocks on GDP developments in BRIC countries. Finally, Section 5 concludes.

2 BRIC countries during the crisis period

Since the 1990s, the fast integration of the BRICs into the world economy has been triggered by the favorable global environment. Strong foreign demand, facilitated by advances in trade liberalization, lower global interest rates, and the acceleration in commodity prices accounted for half of the growth acceleration in the 2000s compared to the 1990s (Cubeddu et al., 2014). The large and sustained increase in commodity prices raised investment and GDP in commodity-exporting economies, many of which enjoyed unprecedented windfall profits. The effects have been most visible if countries are financially open. Higher growth in the years prior to the financial crisis reflected a combination of improved fundamentals and strong tailwinds that boosted demand and raised productivity in most countries.

By focusing on the acute financial crisis period, Blanchard et al. (2010) noted that emerging markets were severely hit by trade and financial shocks. For instance, capital outflows played a dominant role in Russia. Countries with high short-term foreign debt suffered larger declines in GDP compared to less leveraged economies. Interestingly, international reserves did not provide relevant buffers. Based on a decomposition of forecast errors, Fayad and Perrelli (2014) argued that lower demand from trading partners plays a key role to explain the slowdown, besides a general increase in the risk aversion of international investors. In addition, the crisis reduced the scope for expansionary fiscal and monetary policies in many emerging markets. According to Aslund (2013) the current decline in GDP growth is caused by the end of extraordinary commodity and credit booms, and overinvestment (China) or underinvestment (Brazil, Russia). Hence, the former acceleration was not sustainable, as structural factors are also important. Anand et al. (2014) concluded that the slowdown in China and India is related to lower potential output growth, mostly driven by a weaker evolution of TFP. In addition, the decline in the working-age population cuts long run growth in China and Russia. Hence, emerging markets should pursue structural reforms to ensure sustainable economic growth under more challenging global conditions (Didier et al., 2015).

The crucial role of China is often overlooked in the debate. The launch of the massive fiscal stimulus at the end of 2008 of six percent of GDP over a two years period might have delayed the adjustment in emerging markets and contributed to a higher divergence of business cycles. Due to the acceleration of investment China was able to keep the former high growth path for some time. It also provided a buffer for many emerging market countries, as the demand for commodities remained relatively strong. Because of high infrastructure investment to support

the process of faster industrialization and urbanization, China contributed to a large and growing demand for commodities. Over the crisis period, the strategy has been intensified to bolster the economy against negative global shocks. Roache (2012) concluded that shocks in aggregate activity in China could exert substantial impact on the prices of oil and some base metals even long before the crisis. Using a factor augmented VAR approach, Aastveit et al. (2015) argued that demand from emerging economies, especially from China, is more than twice as important to explain the fluctuations in the real oil price and in oil production than demand from developed countries. In 2011/2012, China started to rebalance the growth strategy towards a more sustainable development. Subsequently, many other countries experienced a growth decline. According to Gruss (2014) lower growth in China poses a key downside risk for the Latin American countries. As the shift is not a temporary phenomenon, policies trying to offset the economic slowdown with demand-side measures will be not successful. Hence, structural reforms to secure higher growth over the medium run are on the agenda.

3 VAR and Bayesian SVAR analysis

3.1 Methodology

We estimate country-specific Bayesian structural VARs. Our benchmark specification includes a constant and a linear time trend, which we omit from the notation for convenience:

$$\sum_{l=0}^{L} \begin{bmatrix} A_{11}(l) & A_{12}(l) \\ A_{21}(l) & A_{22}(l) \end{bmatrix} y(t-l) = \varepsilon(t),$$

$$y(t) = \begin{bmatrix} y_1(t) \\ y_2(t) \end{bmatrix},$$

where $y_1(t)$ is a vector of domestic macroeconomic variables, $y_2(t)$ is a vector of global macroeconomic variables,

$$\varepsilon(t) = \begin{bmatrix} \varepsilon_1(t) \\ \varepsilon_2(t) \end{bmatrix},$$

where $\varepsilon(t)$ is uncorrelated with $y(t-l)$ for $l > 0$, and $E[\varepsilon(t)\varepsilon'(t)|y(t-l), l > 0] = I$, $E[\varepsilon(t)|y(t-l), l > 0] = 0$; $\varepsilon_1(t)$ is a vector of structural shocks of domestic origin, and $\varepsilon_2(t)$ is a vector of structural shocks emerging in the global economy.

The model is formulated separately for each BRIC country, where Brazil, Russia and India are treated as small open economies. For these countries the block exogeneity restriction is imposed, i.e. $A_{21}(l) = 0$ for all $l = 0,1,..,L$. Hence domestic variables do not have contemporaneous or lagged effects on global variables. This assumption is in line with the econometric evidence and similar to the approach presented by Cushman and Zha (1997) and Dungey and Pagan (2000). In the presence of a near-VAR system, OLS gives consistent estimates. However, some potential gains, such as reduced number of parameters and more precise estimates, come from estimation of the system using Seemingly Unrelated Regression Equations (SURE). Thus, for Brazil, Russia and India the near-SVAR model is specified and estimated by SURE techniques with Bayesian inference. The applied estimation approach for these countries also ensures a property of invariance for the common global macroeconomic shocks, i.e. although the models are formulated separately for each country, the external shocks hitting them are identical and we are thus able to compare the patterns of dynamics between countries under consideration. Based on Schwartz Information criteria and LM autocorrelation tests a lag length of two was chosen to estimate our models.

For China, however, the block exogeneity assumption might not hold, although there is no consensus on this issue in the literature. For example with respect to oil prices, on the one hand, Du et al. (2010) found that China's economic activity fails to affect the world oil price. This means that the latter is still exogenous with respect to China's macro-economy. On the other hand, investigating China's growing role in the global economy and world commodity markets, Cashin et al. (2016) concluded that indeed a permanent negative Chinese GDP shock of one percent will reduce global growth in the short run by 0.2 percent and oil prices by 2.8 percent. We contribute to that literature and estimate the effect of the Chinese economy on commodity prices, world trade and global financial market volatility, and, thus, implicitly on the other BRIC countries.

3.2 Choice of variables and preliminary data analysis

The vector of domestic variables $y_1(t)$ includes real government expenditures (*GSpend*), real GDP (*GDP*), the difference between the domestic short-term interest rate and corresponding US interest rate (*IR*), and real effective exchange rates (*REER*). The vector of external variables $y_2(t)$ includes the real oil price (*OIL*), the World Trade Index (*WT*), and the CBOE Volatility Index (*VIX*). All data are obtained at the quarterly frequency from Datastream, with exception of

merchandise world trade (2005=100), which comes from the CPB World Trade Monitor. The real oil price was calculated by dividing the price of oil by the US GDP deflator. All variables are defined in logarithms, except the interest rate differentials, and the time series for government expenditure, GDP and world trade are seasonally adjusted. The data are reported over the 2000:1-2015:2 sample. Therefore, periods before, during and after the global financial crisis are included[2].

The Augmented Dickey-Fuller (ADF) tests for the presence of unit roots indicate that all the variables are I(1), i.e. integrated of order one (Table 1). Although the volatility index should be stationary in principle, the unit root test indicates non-stationarity, probably due to the small sample size. Therefore, the VIX is treated as I(1), i.e. should not be excluded from the set of integrated variables.

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-Table 1 about here-

Next, possible cointegration relationships between the variables are explored by means of the trace statistic in line with the Johansen approach for the multivariate cointegration models. Restricted linear trend specifications have been chosen in order to allow the cointegrating relationships to be trend-stationary and have non-zero intercepts[3], the lag length of two was chosen according to the results of our autocorrelation tests. See Table 2 for the results.

-Table 2 about here-

There is the evidence of multiple cointegration relationships for each country: four for Brazil, three for China and India, and two for Russia. Thus, the use of differenced variables in estimations might lead to the loss of important information – such as long-run relationships. For instance, investigating some commodity exporting countries Chen and Rogoff (2002) and Cashin

[2] The starting date was also chosen so that the financial turmoil periods in Brazil, the 1997-1998 Asian Crisis, and the 1998 default in Russia do not enter the sample.
[3] The presence of long-run relationships for each country was also confirmed by the specification with an unrestricted constant, which allows for linear trends in the data, but it is assumed that the trends cancel in cointegrating relations.

et al. (2003) concluded that real exchange rates cointegrate with the real price of commodities.

Sims et al. (1990) argued that VAR models in levels with non-stationary variables might incur some loss in the estimator's efficiency but not its consistency if cointegration holds. Since the main objective of the VAR is to analyze the inter-relationships between the variables and not the coefficients, the system should be estimated in levels. Alternatively a Vector Error Correction Model (VECM) could be estimated. However, identifying the underlying structural parameters with any degree of accuracy within VECM is not an easy task, given the relatively small numbers of degrees of freedom. Therefore, we decided to not impose any cointegrating restrictions.

3.3 Identification of SVARs

The identification of the SVAR models is based on short-run restrictions in order to let the data reveal the pattern of the responses and the transmissions. Within the domestic block, the following ordering of the variables is assumed to hold: real government expenditure, real GDP, the real effective exchange rate and interest rate differential (*GSpend, GDP, REER, IR*). As a proxy for the fiscal stance the model includes government expenditures. Similar to Afonso et al. (2011) it is assumed that all reactions of fiscal policy within each quarter are purely automatic because of the presence of long decision and implementation lags. Blanchard and Perotti (2002) stated that they could not identify any automatic feedback from economic activity to government purchases of goods and services. Thus, government expenditure variable can be considered to be the most inertial variable in the model, and cannot react to current changes in the economy.

The production sector is reflected in real GDP. Cushman and Zha (1997) argued that signals in financial sector variables (interest rate differential, real exchange rate) are related to production only through lags due to inertia, cost adjustments and production planning. Since the commodities boom and the acceleration of world trade spurred export demand, commodity prices and world trade can affect real GDP within one quarter.

The interest rate differential reflects the monetary policy stance compared to the global financial conditions, the latter proxied by the US interest rate. As the only nominal variable in the model, it is the most fast-moving one. Given the lags in monetary policy transmission, the do-

mestic interest rate reacts faster to the shocks to output than the output reacts to changes in the interest rate. As pointed out by Bernanke et al. (1997) oil price shocks may affect monetary policy, which in turn may influence economic activity. Moreover, central banks might also tighten monetary policy in order to combat capital outflows caused by changing global financial conditions. Within a quarter, domestic interest rates might also react to unexpected shocks in the US interest rate and financial volatility. Due to price rigidities we assume that real effective exchange rate reacts to monetary policy only with the lag, but monetary policy, on the other hand, could react to REER developments simultaneously.

Within the global block we keep the variables in the lower triangularized fashion of the order real oil prices, world trade, and financial volatility (*OIL, WT, VIX*). The volume of world trade is affected by oil prices through the demand changes in oil-importing and oil-exporting countries (Husain et al., 2015). Lower oil prices might also reduce the distance elasticity of trade, and thus could promote globalization. Financial volatility can react immediately to unexpected shocks hitting both oil prices and global trade.

The identification of the structural form requires $\mathrm{N}(\mathrm{N}-1)/2$ restrictions to hold. For all models we assume that global variables are not affected contemporaneously by domestic shocks (and with the lags for Brazil, India and Russia due to the block exogeneity assumption). Other restrictions come from the Cholesky orderings within the two blocks. As twenty one zero restrictions are imposed, the model is exactly identified. The following matrix summarizes the set of the contemporaneous restrictions:

$$\begin{vmatrix} \varepsilon_{GSpend} \\ \varepsilon_{GDP} \\ \varepsilon_{REER} \\ \varepsilon_{IR} \\ \varepsilon_{OIL} \\ \varepsilon_{WT} \\ \varepsilon_{VIX} \end{vmatrix} = \begin{vmatrix} 1 & 0 & 0 & 0 & a_{15} & a_{16} & a_{17} \\ a_{21} & 1 & 0 & 0 & a_{25} & a_{26} & a_{27} \\ a_{31} & a_{32} & 1 & 0 & a_{35} & a_{36} & a_{37} \\ a_{41} & a_{42} & a_{43} & 1 & a_{45} & a_{46} & a_{47} \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & a_{65} & 1 & 0 \\ 0 & 0 & 0 & 0 & a_{75} & a_{76} & 1 \end{vmatrix} . \begin{vmatrix} u_{GSpend} \\ u_{GDP} \\ u_{REER} \\ u_{IR} \\ u_{OIL} \\ u_{WT} \\ u_{VIX} \end{vmatrix}$$

Alternative Cholesky orderings within the two blocks, as well as imposing over-identifying restrictions on contemporaneous effects of external shocks to national variables do not change the main results significantly.[4]

The identification approach involves simultaneity among the contemporaneous variables. Therefore, the shape of the posterior density of the model parameters tends to be non-Gaussian. In order to obtain accurate statistical inferences from the parameter estimates we estimate the model employing the Metropolis-within-Gibbs sampling method[5]. Bayesian methods provide an explicit, straightforward approach to incorporate uncertainty into modelling and forecasting. Monte Carlo integration and Gibbs Sampling can be efficiently used for the SVAR models that have a restricted covariance matrix of the reduced-from residuals as well as restrictions on the lagged coefficients (such as SVAR models with block exogeneity). In order to get initial estimates for the Gibbs sampler the model is estimated by seemingly unrelated regression (SURE) techniques. The maximum of the log of the marginal posterior density for the matrix with contemporaneous restrictions is computed using the Broyden, Fletcher, Goldfarb and Shanno (BFGS) approach. See Press et al. (1989) for details. The prior degrees of freedom are equal to (N+1)/2, where N is the number of variables in the model. The covariance matrix of residuals Σ_β is also used as increment for the Random Walk Metropolis.

[4] The results are available upon request.
[5] For more details please refer to Doan (2010).

The following algorithm is applied. Firstly, we compute the log likelihood for the structural model given the covariance matrix of residuals at the current draw for the coefficients (Σ_β). After drawing a candidate set of structural parameters and computing the log likelihood, a Metropolis acceptance test is performed to determine whether to reject or accept the candidate draw. In case of acceptance, the diagonal elements are drawn for the structural covariance matrix using the set of structural parameters and the covariance matrix Σ_β. Finally, the coefficients are drawn from the Seemingly Unrelated Regressions, and the covariance matrix of residuals at the current draw for the coefficients is computed for the next round.

4. Empirical findings

4.1 Weak exogeneity tests

The block exogeneity assumption for Brazil, India and Russia implies that these countries individually do not have dominant influence on the global markets. They are seen as price-takers, and their contribution to world trade can be neglected. On the other hand, China might play a significant role in affecting global conditions. Statistical evidence on the small open economy assumption for Brazil, India and Russia can be obtained by weak exogeneity tests. Gujarati (2006) pointed that when the variables are integrated, one may not be able to use F-statistic to jointly test the Granger causality, since the test statistics do not have a standard distribution. Thus, instead of standard Granger causality tests, the less strong concept of weak exogeneity is used. A variable is said to be weakly exogenous if it does not adjust to temporary deviations from the cointegration relationships.

For each country three model variants are considered, one for each global variable. The specifications include all domestic variables - $y_1(t)$, and a specific global variable - $\mathrm{z(t)}$, where $\mathrm{z} \in \{\mathrm{OIL}, \mathrm{WT}, \mathrm{VIX}\}$. The cointegration rank is determined by the trace test statistic. The marginal model for $\Delta \mathrm{z}(t)$ can be written as[6]:

$$\Delta \mathrm{z}(t) = \alpha_2 \beta' X(t-1) + \Gamma_{21} \Delta \mathrm{X}(t-1) + \varepsilon_2(t),$$

where $X(t) = \begin{pmatrix} y_1(t) \\ \mathrm{z(t)} \end{pmatrix}$.

[6] Deterministic components are omitted for notational convenience.

Then, the global variable in question $z(t)$ is tested for being weakly exogenous, i.e. that $\alpha_2 = 0$. The associated test statistic is asymptotically χ^2distributed with *r* degrees of freedom, see Table 3 for the results. According to the evidence, the global variables can be considered as weakly exogenous for all countries, except China, where the oil price and world trade variables demonstrate adjusting behavior to deviations from long-run relationships. Thus, while the near SVAR specification is appropriate for the other countries, it would not be the optimal choice for the Chinese case, and we do not impose block exogeneity assumption for China's model.

-Table 3 about here-

In the following we will present the findings of the role of China in determining global variables and proceed by the analysis of the relative role of external factors in the GDP developments for the BRIC economies.

4.2 China's role in determining global variables

In order to analyze the role of China in determining global variables, we proceed with the parsimonious VAR model, where only the global variables (world trade, oil price, financial volatility) and China's GDP are included. Table 4 contains the results of the VAR diagnostic tests, the cointegration rank and the weak exogeneity tests, as well as standard Granger causality tests.

-Table 4 about here-

To avoid potentially unreasonable restrictions the model is simply kept in its reduced form, where global variables are ordered first (WT, OIL, VIX, and China's GDP). Alternative orderings only slightly change the pattern of impulse responses. Figure 3 shows that China indeed plays a significant role in determining oil prices and global trade. According to Figure 4 the one-time unexpected shock in China's output could explain up to 22.3 percent of oil price forecast error variations after two years. The effect on world trade is less pronounced (5.3 percent after eight quarters), but nevertheless significant. Hence, the change in the Chinese growth strategy puts

reform pressure on countries with abundant natural resources. China's effect on financial risk aversion is found to be insignificant.

-Figures 3 and 4 about here-

4.3 GDP responses in the BRIC countries

The impulse responses of the Bayesian SVAR models are listed in Figures 5 to 8 and show a sensible adjustment pattern after global shocks.

-Figures 5, 6, 7 and 8 about here-

Oil-exporting countries – Brazil and Russia – react positively to higher real oil prices. We observe the opposite effect for oil-importing China, and for India, the oil price shock tends to be insignificant. An acceleration of global trade is associated with an output increase within the first two years in all countries under consideration. While a real exchange rate appreciation exerts a positive effect on Brazil's and Russia's GDP, it has a negative impact for China and India. For the latter countries this might be explained by losses in export competiveness in a highly competitive global trade environment. Government expenditures have a high expansionary effect on GDP for Brazil. For the other countries the impact is positive as well, however, with less magnitude. The unexpected tightening of monetary policy (compared to the US) and an increase in financial uncertainty lead to fall in output in all BRIC countries. Overall, one can conclude that external variables played a significant role in the development of GDP in the BRIC countries.

4.4 Forecast error variance decomposition (FEVD) of BRIC's GDP

In order to determine the ability of external shocks to explain domestic GDP fluctuations at different horizons we perform a standard forecast error variance decomposition exercise.

Variance decomposition separates the variation of an endogenous variable into the components of the VAR. Thus, it provides information about the relative importance of each innovation. The forecast error variance decomposition analysis suggests that the BRICs are heavily affected by the global economy, albeit in different manner and to a different degree.

-Figures 9, 10, 11 and 12 about here-

Commodity prices can explain the downturn in Brazil and Russia to a large extent – the average share of the total variance of the forecast error for GDP attributable to the variance of oil shocks during the first two years is 14 and 30 percent. The time path of the responses is different in these two countries - in Russia it gains immediately about 27 percent of the FEVD with the peak of 43 percent already achieved in the second quarter, comparing with an initially small but persistently increasing role in Brazil up to 23.6 percent in the 8th quarter. The effects of global trade are not emerging instantaneously and gain importance for both countries after some time. To sum up, Brazil and Russia are found to be more prone to external shocks than India and China, where the relative proportion of domestic shocks to external ones in GDP's FEVD is higher on average during the first two years. Output in India is insignificantly affected by the oil price evolution. However, a slower expansion of world trade depresses Indian GDP growth. Prices for raw materials and the expansion of world trade are both relevant to explain output growth in China. However, in contrast to other countries, the relationship for China is bidirectional.

Conclusions

In this contribution, we started from the observation that the financial crisis led to a deep recession in many industrial countries. However, the downturn in large emerging markets turned out to be less persistent. Despite the modest recovery in advanced economies, GDP growth declined in emerging markets in the last years. We argued that the higher divergence of business cycles is closely linked to the Chinese transformation. During the crisis, the Chinese fiscal stimulus prevented a decline in GDP growth not only in that country, but also in resource-rich economies. The Chinese shift to consumption-driven growth led to a decline in commodity

demand, and the environment became more challenging for many emerging markets. We have been able to support this view by means of Bayesian VARs which we specified for the BRIC (Brazil, Russia, India, China) countries. Our results revealed a strong impact of international variables on GDP growth. As a stylized fact and in contrast to the other countries, China plays a crucial role in determining global trade and oil prices. Hence, we concluded that the change in the Chinese growth strategy puts additional reform pressure on countries with abundant natural resources.

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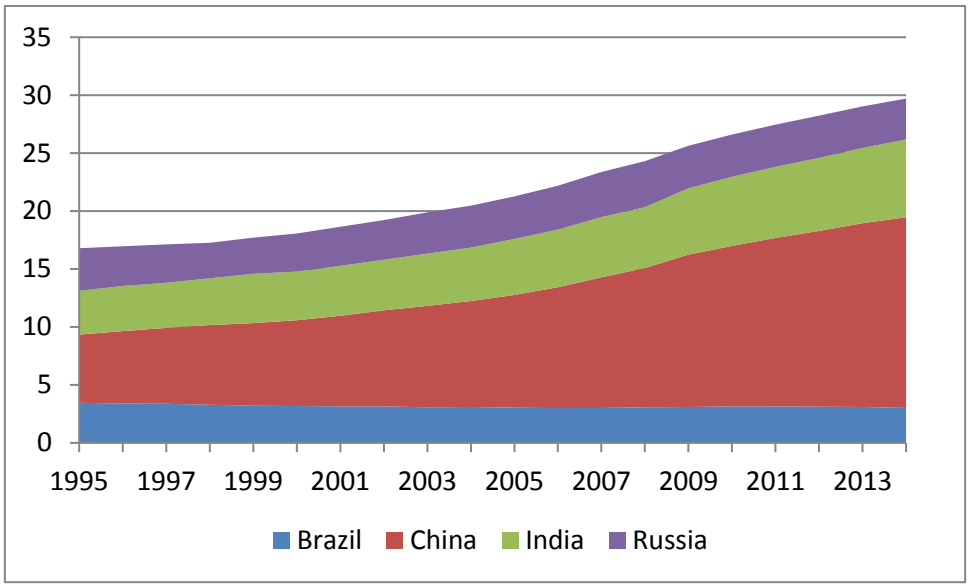
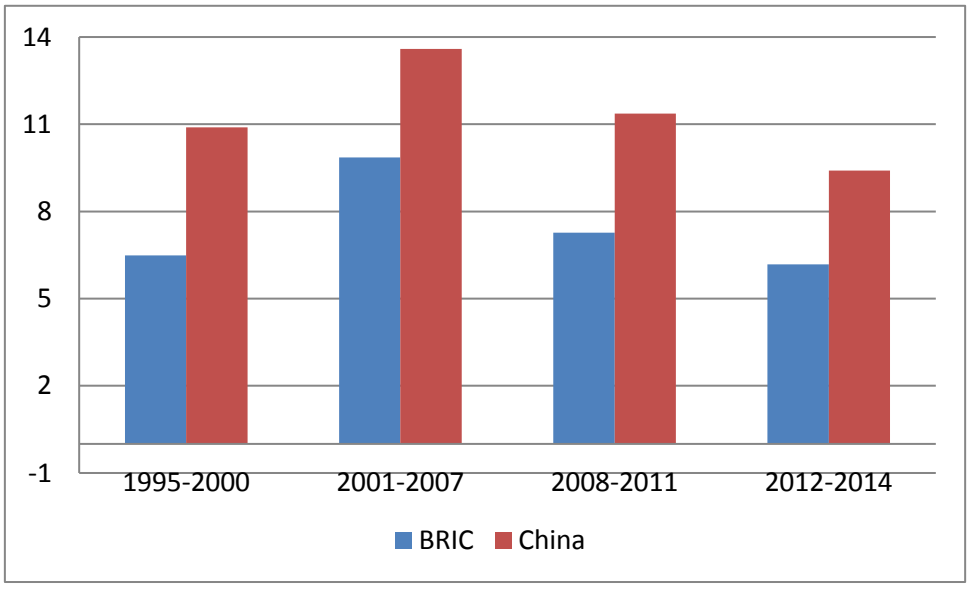
Figure 1: Share of the BRICs in the world economy



Note: Calculation based on PPPs. IMF World Economic Outlook.

Figure 2: Economic growth in the BRICs and China



Note: GDP growth rates based on PPPs. IMF World Economic Outlook.

Figure 3. Impulse responses for VAR model with global variables and China's GDP

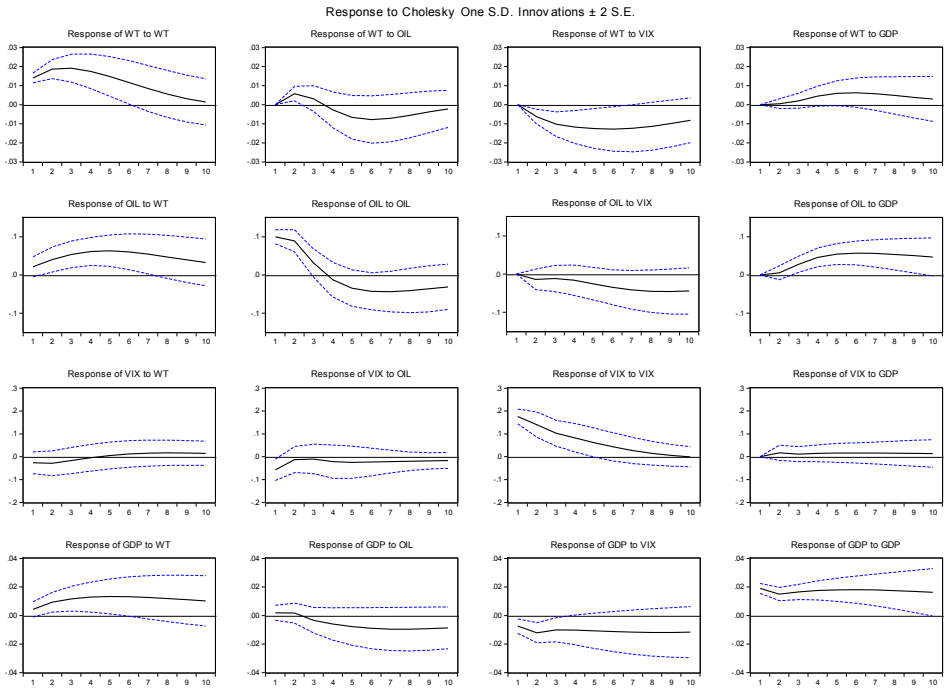


Figure 4. FEVD for VAR model with global variables and China's GDP

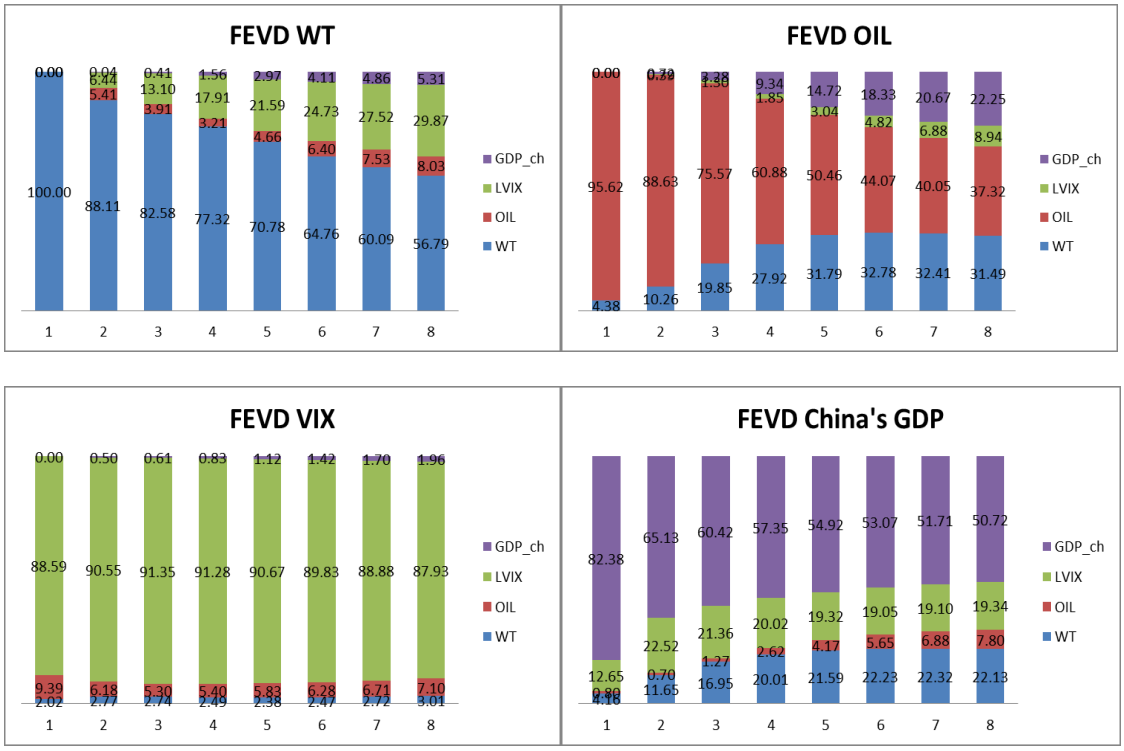


Figure 5. Impulse Responses of Brazil's GDP

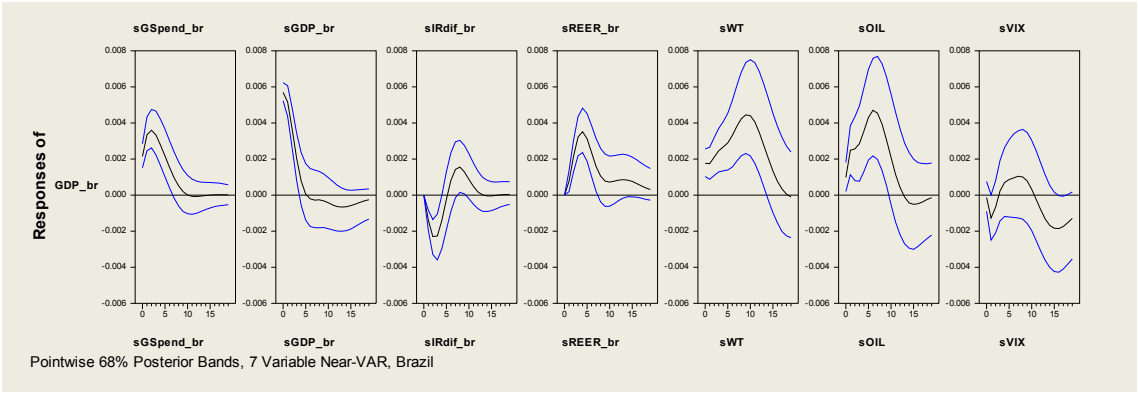


Figure 6. Impulse Responses of Russia's GDP

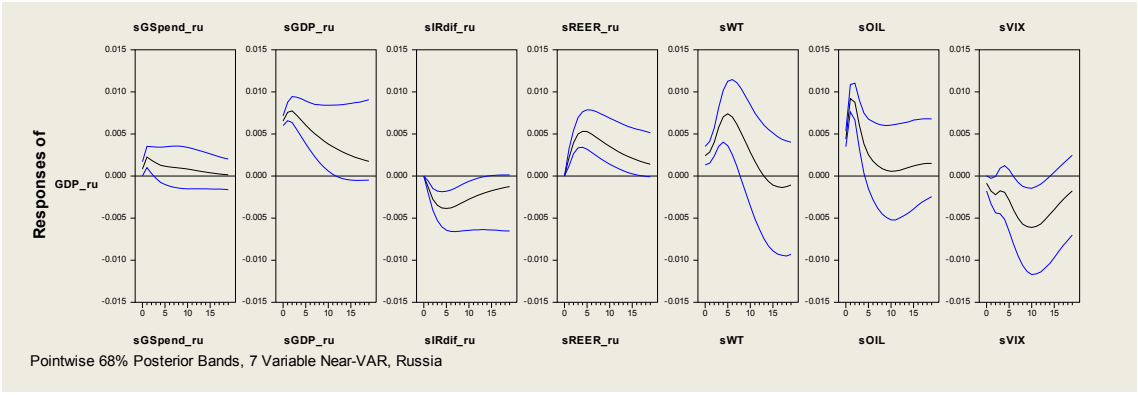


Figure 7. Impulse Responses of India's GDP

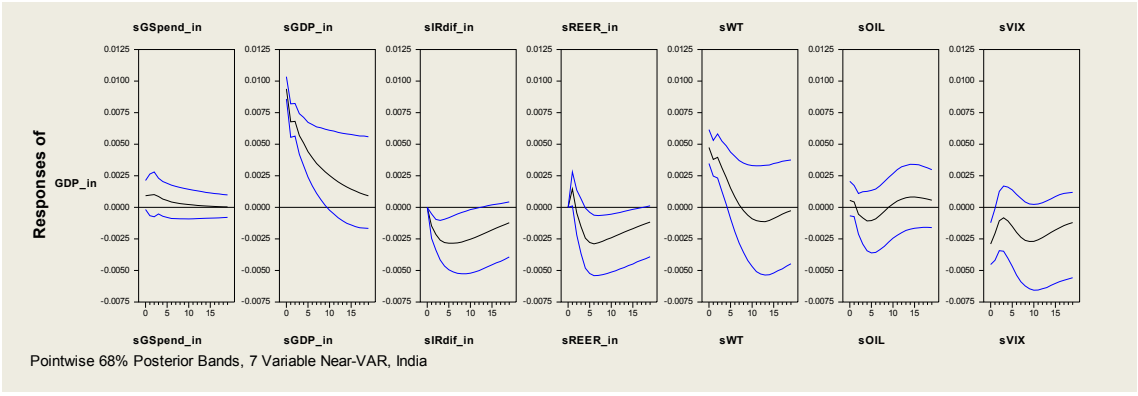


Figure 8. Impulse Responses of China's GDP

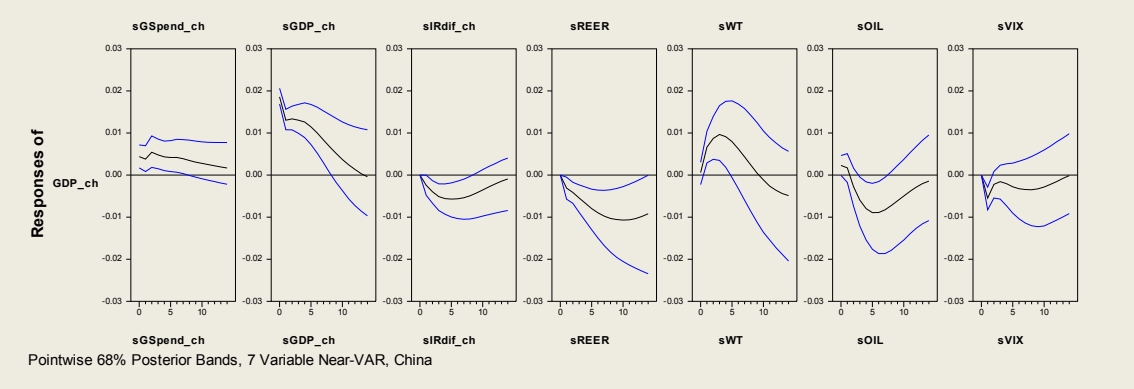


Figure 9. FEVD of Brazil's GDP

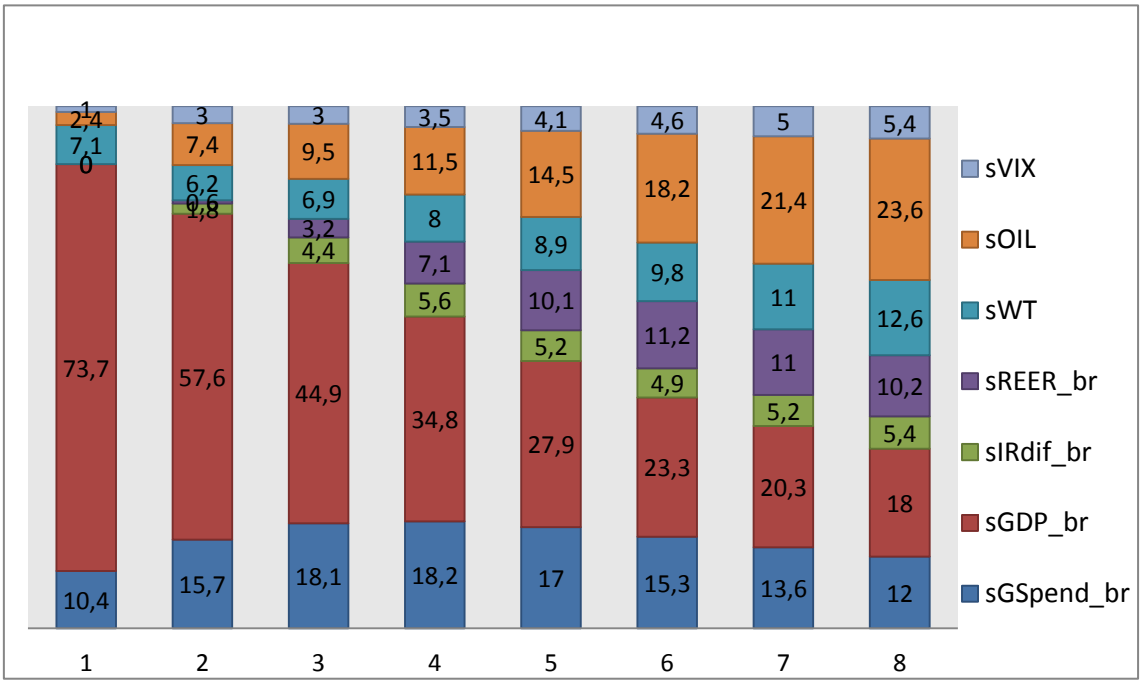


Figure 10. FEVD of Russia's GDP

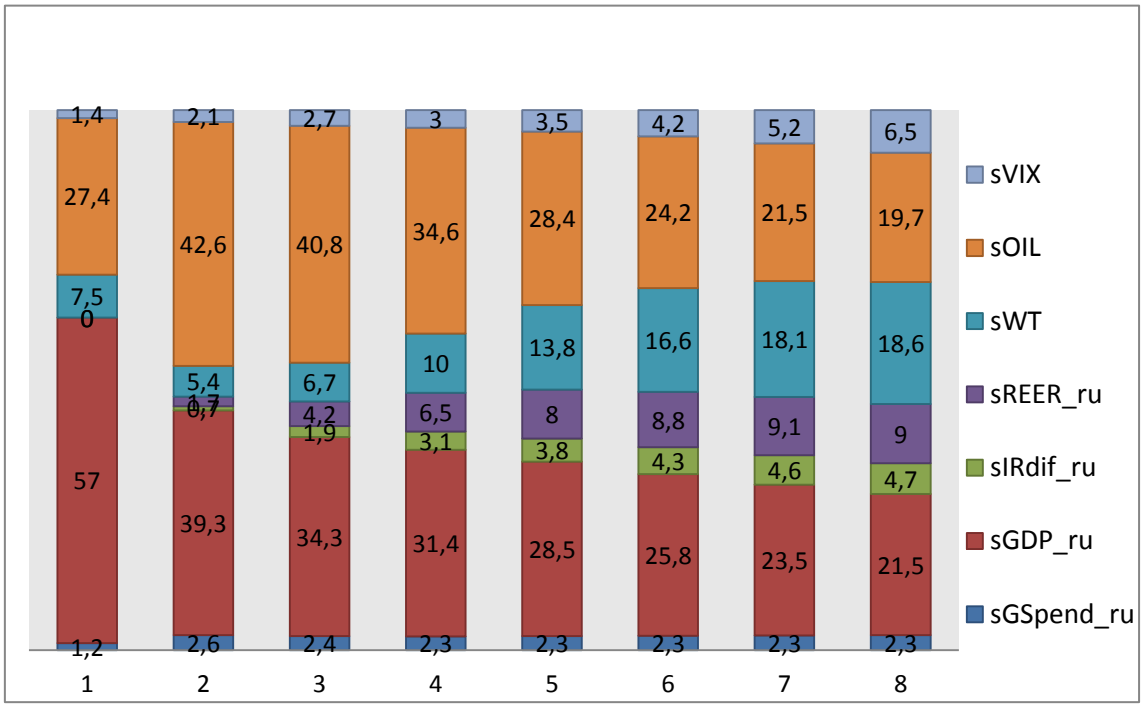


Figure 11. FEVD of India's GDP

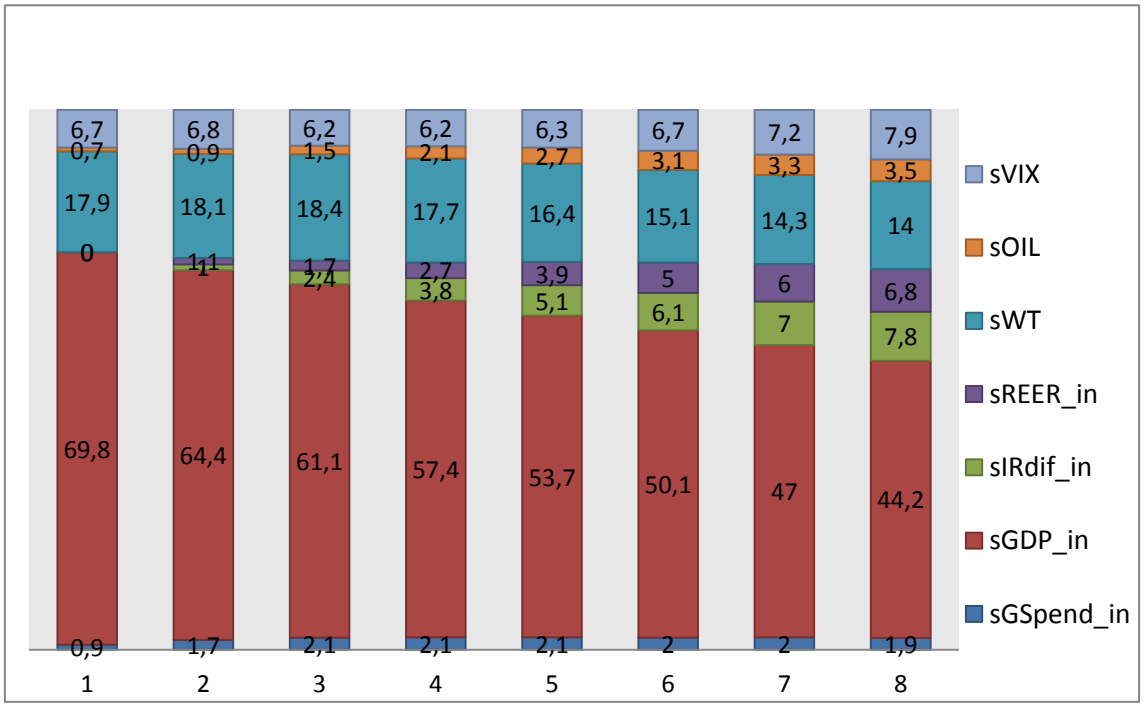


Figure 12. FEVD of China's GDP

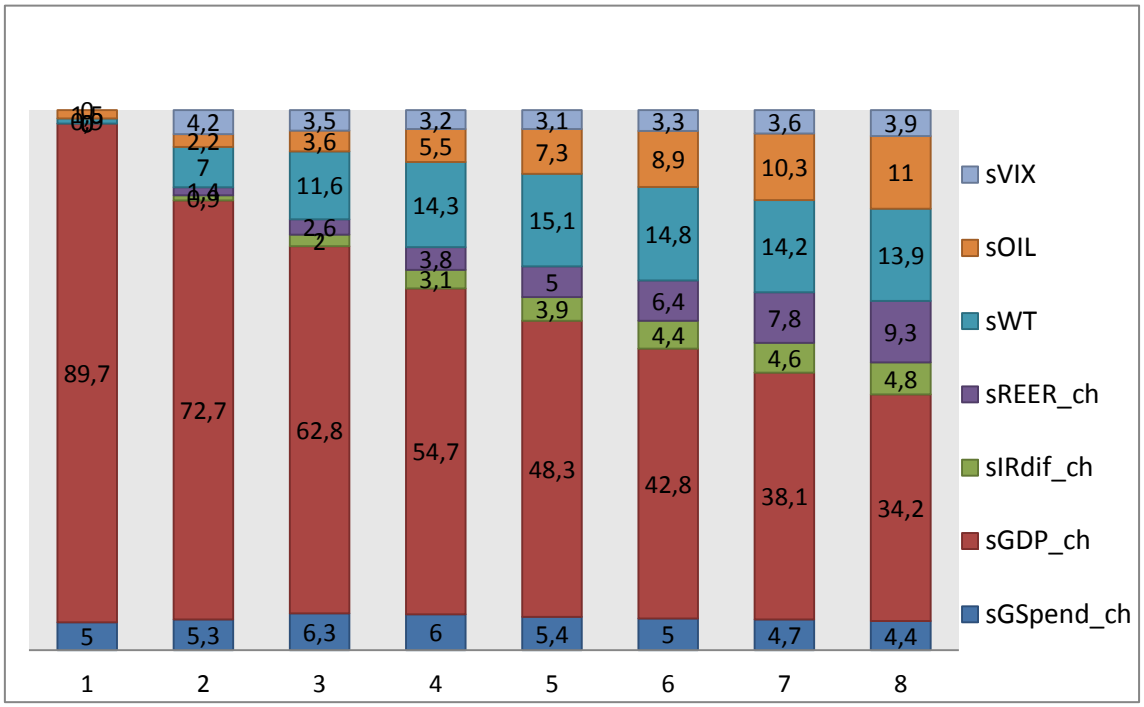


Table 1. ADF test for included variables

	Brazil				**Russia**			
	Level	*Prob*	*Diff*	*Prob*	*Level*	*Prob*	*Diff*	*Prob*
Gspend	-0.48	0.89	-4.42	0.00	-1.69	0.43	-4.09	0.00
GDP	-1.30	0.63	-4.11	0.00	-2.04	0.27	-2.42	0.02
IR	-1.05	0.26	-4.97	0.00	-1.71	0.08	-3.81	0.00
REER	0.44	0.81	-7.40	0.00	2.03	0.99	-6.61	0.00
	China				**India**			
	Level	*Prob*	*Diff*	*Prob*	*Level*	*Prob*	*Diff*	*Prob*
Gspend	-1.09	0.71	-1.82	0.07	-0.45	0.89	-9.72	0.00
GDP	-1.27	0.64	-2.70	0.01	0.48	0.98	-7.75	0.00
IR	-1.15	0.23	-4.83	0.00	-0.27	0.58	-5.94	0.00
REER	1.65	0.97	-6.10	0.00	0.91	0.90	-7.08	0.00

	Global variables			
	Level	*Prob*	*Diff*	*Prob*
OIL	-1.74	0.40	-6.85	0.00
WT	-1.38	0.59	-3.98	0.00
VIX	-0.56	0.47	-8.34	0.00

Following ADF specifications were applied:

-for levels of Gspend, GDP, WT, OIL - ADF with intercept;

-for levels of IR, REER,VIX and all differences - ADF with no intercept and no trend;

Lag length was chosen according to Schwarz criterion

Table 2. Cointegration rank test[7]

	Brazil		**China**		**India**		**Russia**	
*r***	*Trace*	*P-Value****	*Trace*	*P-Value*	*Trace*	*P-Value*	*Trace*	*P-Value*
0	162.66	0.01	169.86	0.00	197.06	0.00	172.15	0.00
1	119.74	0.04	119.80	0.04	128.89	0.01	122.99	0.02
2	87.29	0.06	87.15	0.06	84.87	0.09	83.47	0.11
3	60.90	0.09	56.71	0.17	55.51	0.21	52.75	0.30
4	38.11	0.14	35.26	0.24	29.16	0.56	29.25	0.55
5	15.90	0.51	19.64	0.25	9.52	0.94	14.78	0.60
6	3.06	0.86	9.28	0.17	3.24	0.84	6.71	0.39

* The model for each county includes domestic (GSpend, GDP, IR, REER) and global variables (WT, OIL, LVIX)

** *r* is the rank

*** *P*-values for rank test with the null hypothesis that the number of cointegrating vectors is less or equal to r against a general alternative

Restricted linear trend specifications have been chosen in order to allow the cointegrating relationships to be trend-stationary and have non-zero intercepts, the lag length of two was chosen according to the autocorrelation tests.

[7] Cointegration rank and weak exogeneity tests have been performed with the CATS in RATS software

Table 3. Weak exogeneity tests

		Brazil		**China**		**India**		**Russia**	
	r	Trace	P-Value	Trace	P-Value	Trace	P-Value	Trace	P-Value
I*	0	76.38	0.01	98.12	0.01	95.15	0.02	96.61	0.01
	1	50.07	0.03	64.13	0.05	44.06	0.69	64.17	0.05
	2	26.47	0.12	39.73	0.01	20.51	0.95	34.29	0.28
	3	10.32	0.26	20.23	0.22	5.02	1.00	15.93	0.51
	4	1.27	0.26	8.03	0.26	0.92	1.00	5.45	0.54
Test WE**		$\chi^2(2)$ =0.114, [0.944]		$\chi^2(3)$ =11.044, [0.011]		$\chi^2(1)$ =1.004, [0.316]		$\chi^2(2)$ =2.251, [0.324]	
	r	Trace	P-Value	Trace	P-Value	Trace	P-Value	Trace	P-Value
II***	0	67.81	0.07	86.72	0.07	96.82	0.01	90.93	0.03
	1	44.91	0.09	47.80	0.52	53.19	0.29	59.18	0.12
	2	24.15	0.20	28.51	0.60	22.46	0.89	36.28	0.20
	3	6.50	0.64	11.73	0.83	6.58	0.99	15.80	0.52
	4	1.51	0.22	3.94	0.75	1.72	0.97	7.87	0.27
Test WE		$\chi^2(2)$ =3.027, [0.220]		$\chi^2(1)$ =5.251, [0.022]		$\chi^2(1)$ =1.173, [0.279]		$\chi^2(1)$ =0.527, [0.468]	
	r	Trace	P-Value	Trace	P-Value	Trace	P-Value	Trace	P-Value
III****	0	69.94	0.05	94.19	0.02	97.73	0.01	86.73	0.07
	1	46.49	0.07	54.12	0.25	56.90	0.17	51.17	0.37
	2	26.43	0.12	30.14	0.50	29.36	0.55	27.07	0.68
	3	11.75	0.17	13.65	0.69	8.27	0.97	13.57	0.70
	4	0.61	0.44	4.20	0.71	4.18	0.72	4.36	0.69
Test WE		$\chi^2(2)$ =2.252, [0.324]		$\chi^2(1)$ =5.293, [0.021]		$\chi^2(1)$ =0.063, [0.802]		$\chi^2(1)$ =1.118, [0.290]	

* The model I for each county includes domestic variables (GSpend, GDP, IR, REER) and real oil prices

** LR test for weak exogeneity performed based on obtained cointegrating rank, P-values in brackets

*** The model II for each county includes domestic variables (GSpend, GDP, IR, REER) and world trade

**** The model III for each county includes domestic variables (GSpend, GDP, IR, REER) and VIX

Restricted linear trend specifications have been chosen in order to allow the cointegrating relationships to be trend-stationary and have non-zero intercepts, the lag length of two was chosen according to the autocorrelation tests.

Table 4. Diagnostic, cointegration rank, weak exogeneity and Granger causality tests for the VAR model with global variables and China's GDP

Diagnostic tests [8]*:*

VAR Residual Serial Correlation LM Tests

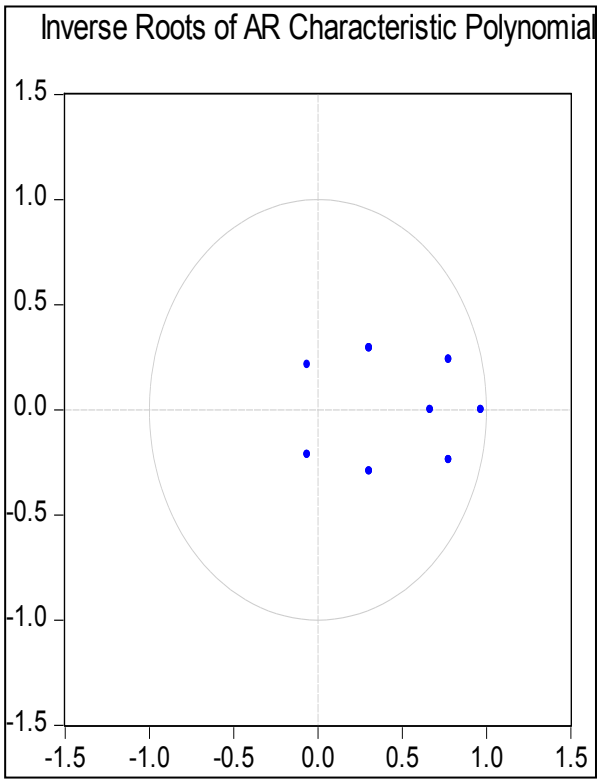
Lags	LM-Stat	Prob
1	20.38	0.20
2	14.25	0.58
3	21.67	0.15

Probs from chi-square with 16 df.

VAR Lag Order Selection Criteria

Lag	AIC	SC	HQ
0	-6.05	-5.47	-5.82
1	-11.73	-10.58	-11.28
2	-12.36*	-10.64*	-11.69*
3	-12.05	-9.76	-11.16
4	-12.01	-9.14	-10.89
5	-11.92	-8.48	-10.58

* indicates lag order selected by the criterion
AIC: Akaike information criterion
SC: Schwarz information criterion
HQ: Hannan-Quinn information criterion



Cointegration rank and weak exogeneity tests:

Cointegration rank test

r	Trace	P-Value
0	63.74	0.05
1	27.68	0.65
2	11.52	0.84
3	0.43	1.00

Test of weak exogeneity

r	5% C.V.	GDP_ch	WT	OIL	VIX
1	3.84	7.20	2.96	16.74	0.01
		[0.01]	[0.09]	[0.00]	[0.93]

Granger Causality tests:

Dependent variable: OIL			
Excluded	Chi-sq	df	Prob.
WT	7.07	2	0.03
VIX	4.94	2	0.08
GDP_ch	14.20	2	0.00

Dependent variable: WT			
Excluded	Chi-sq	df	Prob.
OIL	11.46	2	0.00
VIX	11.81	2	0.00
GDP_ch	4.70	2	0.10

[8] Based on Schwarz, Akaike and Hannan-Quinn information criterions the lag length of 2 was chosen. According to the autocorrelation LM test the residuals don't show the signs of autocorrelation up to the third lag. Inverse roots of AR characteristic polynomial lie inside unit circle, and thus, the model is stable

All	*30.71*	*6*	*0.00*

All	*45.90*	*6*	*0.00*

Dependent variable: VIX			
Excluded	Chi-sq	df	Prob.
OIL	2.05	2	0.36
WT	0.87	2	0.65
GDP_ch	1.10	2	0.58
All	*4.03*	*6*	*0.67*

Dependent variable: GDP_ch			
Excluded	Chi-sq	df	Prob.
OIL	3.50	2	0.17
WT	7.73	2	0.02
VIX	6.96	2	0.03
All	*15.53*	*6*	*0.02*