The transmission of US monetary policy shocks to EMEs: an empirical analysis

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Abstract

We consider the extent to which three channels that transmit US monetary policy shocks to emerging market economies (EMEs) may have changed in importance and relative strength – ie structurally changed – since the third quarter of 2008. We run linear regression models and estimate a factor-augmented vector autoregression (FAVAR) for the United States and each EME in our database, using macroeconomic variables from each economy. We find that the possibility of structural change in the policy rate, exchange rate, and long-term interest rate channels generally depends on the EME in question. Also, in the case of some tests, accounting for unconventional monetary policy (UMP) does seem to make a difference in this result. However, EMEs seem to have experienced some structural changes more uniformly in what can be interpreted as (i) second-round effects in the channels or, more likely, (ii) as changes in channels that we have not explicitly modelled but that are nonetheless being captured. Such changes highlight the potential for a renewed interdependence between the monetary policies in most EMEs and US monetary policy above and beyond the unprecedented stance of the latter. They also underscore the importance for most EMEs of taking appropriate policy measures, in particular as the United States begins tightening its monetary stance. Our results should be interpreted with some caution given the limited length of the time series.

Keywords: Monetary policy, central banking

JEL classification: E4, E5

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The opinions expressed in this paper are exclusively the responsibility of the authors and do not necessarily reflect the point of view of the Bank of Mexico.

Introduction

It is well understood that during recent decades the world's economies and financial systems have grown more integrated through a relentless globalization process (eg Friedman (2005)). However, what is perhaps less well understood is the extent to which the process has made monetary policy more globally integrated, thereby tightening the mechanisms by which monetary policy shocks are transmitted internationally.

Despite some significant efforts towards understanding this topic in general (eg Galí and Gertler (2009)) as well as with regard to emerging market economies (EMEs) (eg Kim and Yang (2009)), the issues pertaining to EMEs have received relatively little attention. In this context, several relevant questions arise: What is the nature of the mechanism by which US monetary policy effects are transmitted to EMEs? What is the strength and relative importance of the transmission channels? Have these channels changed after the recent financial crisis and, if so, to what extent? What are the implications, if any, for EME economic policies?

Against this backdrop, we undertake two studies. First, we quantitatively explore the policy rate, exchange rate, and long-term interest rate channels by which US monetary policy shocks are transmitted to EMEs.[2] Of course, these are but three among other channels present in the transmission mechanism. Second, we assess the extent to which the strength and relative importance of these channels might have changed since recent global financial crisis; we hereafter refer to such changes as structural changes. We apply these enquiries to a set of 15 EMEs (sometimes fewer, depending on data availability) that have varying exchange rate arrangements, monetary regimes, degrees of financial openness, and policy responses (the Appendix includes the list of EMEs along with descriptions of their exchange arrangements and monetary regimes).[3]

We estimate a set of linear regression models and a factor-augmented vector autoregression (FAVAR) for the United States and each EME in our database, using macroeconomic variables from each economy. This allows us to incorporate a wide range of time series data in our estimation. We assess the extent to which such channels might have changed beginning with the third quarter of 2008. To do so, we run a set of regressions with data for the 22 quarters preceding the crisis (Q1 2003 to Q2 2008) and, in a separate estimation, with data for the 22 subsequent quarters (Q3 2008 to Q4 2013). We use an auxiliary dummy variable to account for possible changes in the coefficients measuring the impact of the US policy rate (ie the federal funds rate) on relevant macroeconomic variables. Accordingly, we run a

[2] For a specific emerging market economy, we explore these channels by examining how changes in the US policy rate affect (i) the EME's policy rate; (ii) the EME's exchange rate; and (iii) the EME's long-term interest rates. As the transmission effect is not limited to these three variables, we aim to capture other effects through a vector autoregression.

[3] One might divide the EMEs in our database into the following groups: (i) South American EMEs, most of which have benefited from the commodities super cycle and the significant increase in demand, in particular from China; (ii) Asian EMEs, characterized by an efficient industrial block; (iii) eastern European EMEs, whose situation is greatly affected by the EU; (iv) Mexico, which is closely related to the US economy; and (v) others. Specific patterns within these groups do not appear in our exercises.

battery of statistical tests on the relevant coefficients. It is important to clarify that while in the regressions we use data from the quarters preceding the crisis and, in a separate estimation, we use data from the subsequent quarters; in the case of the FAVAR models we use the whole sample. The key difference is that in one FAVAR we allow for a structural change in some coefficients, while in the other FAVAR we do not allow for a structural change, as later explained in detail in the text.

In this context, a natural challenge is to measure the US monetary policy stance once the US federal funds rate reaches the zero lower bound. At that point, we sidestep the issue by using the rate proposed by Wu and Xia (2013) (hereafter, the Wu and Xia rate). That rate essentially coincides with the US federal funds rate as long as the policy rate is positive; but once the policy rate hits the zero lower bound, the Wu and Xia rate can become negative. When negative, its distance from zero is a quantitative measure of the effectiveness of unconventional monetary policy.

Having a better understanding of the international transmission mechanism of monetary policy is relevant for many reasons. Monetary policy in advanced economies has played an essential role in the recovery from the global financial crisis. In fact, it has been referred to as the "only game in town" (Rajan (2013)).[4] Also, it is crucial for investors and policymakers alike to understand the implications of global monetary policy for EMEs, where capital flows can be directly influenced by changes in the monetary policy stance in advanced economies. A better understanding of the international transmission of US monetary policy can therefore provide critical assistance to policy planning within EMEs.

Our main findings are as follows. We obtain initial evidence that the existence of a structural change in the policy rate, exchange rate, and long-term interest rate channels depends on the EME in question.[5] However, EMEs seem to have experienced some structural changes more uniformly. These changes can be thought of as second-round effects of the three channels modelled; but we believe they are more likely to be arising from other channels that our model is implicitly capturing.

Here, we will discuss our results in a broader perspective. Much of the attention given to the international transmission mechanisms of US monetary policy started with the global financial crisis. In responding to that crisis, the US central bank lowered the policy rate esentially to the zero lower bound, after which it began implementing unconventional monetary policies. Under those measures, significant capital flows swiftly entered and exited EMEs, depending to a great extent on diverse episodes and situations in the advanced economies. Moreover, there are possibly several factors influencing each of the three channels we study here.

For *the policy rate channel*, we could hardly have expected uniform results since: (i) the initial conditions of each EME before the global financial crisis differed substantially; (ii) we have observed diversity among EMEs in their management and policy responses to unconventional monetary policies; and (iii) except for the worst

4 Moreover, some researchers have cautioned about its limits and risks (eg Eichengreen et al (2011)).

5 As will be discussed shortly, structural change in the long-term interest rate channel appears to be more uniform across EMEs if the linear regression models account for unconventional monetary policies in the United States.

part of the crisis, very few EMEs have shared the same business cycle phase with the main advanced economies, in particular the United States. Moreover, business cycles have not been synchronized even among EMEs. Our mixed results are in line with the significant differences among EMEs in the factors considered above.

For *the exchange rate channel*, the results depend on the specific EME as well. Yet, arguably, there is less evidence of possible structural changes in this channel across the battery of tests. We could have expected this for two reasons. These EMEs have, for a relatively long time now, maintained a stationary inflationary process (Noriega, Capistrán and Ramos-Francia (2013)). Indeed, they have not monetized their fiscal deficits, ending fiscal dominance. This has produced a much lower pass-through.

Finally, *the long-run interest rate channel* also offers mixed evidence. Yet, a specific set of exercises (the linear regression models) give some evidence of a more generalized change when one accounts for unconventional monetary policies. We believe that there are several factors behind these results, some of which are above and beyond the global financial crisis. On the one hand, the secular decrease in long-term rates in the United States that began in the early 1980s accelerated, given the search-for-yield phenomenon, and was further supported by elements such as US demographic dynamics and the implementation of unconventional monetary policies. In fact, the fall in the long-term rates in EMEs has been faster than their corresponding individual deflationary processes.

In addition, these US factors have played a reinforcing role abroad. For example, the aggressive search for yield has certainly affected the interest rates in EMEs. In this context, a key question is what would be the scenario once normalization of the US federal funds rate starts. We believe that, although some bouts of volatility will quite possibly take place given the secular decrease in US long-term rates, the fierce competition among asset management companies for the highest yields in EMEs will continue. In the long-run rate channel, other factors have surely played a role as well and thus have led to differing results among EMEs. For instance, some of these economies have implemented capital controls, established macroprudential policies, allowed credit booms, or had more flexibility in their fiscal policies. All in all, it would have been a surprise to observe uniform results for the long-term interest rate channel.

However, joint tests (ie tests jointly assessing a structural change in various channels) tend to reject the null hypothesis of no structural change. We believe there are channels that may have changed but that we are not explicitly modelling. For instance, the global financial crisis and the associated financial regulatory response have significantly affected the balance sheets of banks and, to an extent, also of the so-called shadow banks. As is well known, their aim has been to improve their risk pricing. These factors are likely to have affected other channels, for instance, the credit channel. In sum, we find more uniform evidence of a structural change in channels we do not observe but that we think our models are capturing.

A caveat about our results is that only a few years have passed since the global financial crisis. Thus, the post-crisis period offers a relatively small number of observations for each time series, which makes statistical relationships difficult to detect. In addition, the relatively short time series limit the number of model specifications we can entertain. Thus, our results represent an exploration rather than robust statistical findings.

In the following sections we describe our data, model, and estimations, discuss our main results and offer some concluding remarks.

Data, model and estimations

As one would expect, there are more time series available for the United States than for any EME. In addition, data are not uniform across EMEs. Prominently, for some EMEs we do not have the time series of the long-term interest rate, which in those cases prevents us from studying that channel. Nonetheless, we have tried to keep databases as uniform as possible.

All time series have been transformed to a quarterly frequency and rendered stationary, the latter a needed condition for VAR model estimation. Let z_t be a generic variable. In general, except for interest rates, the percentage growth is taken (ie $\log(z_t)-\log(z_{t-1})$). Nonetheless, if a time series has negative values, then the difference is taken (ie z_t-z_{t-1}). To obtain evidence on stationarity, a stability test is performed on each of the VARs we estimate.[6] These tests are assessed but not reported.

Data

US data

We use a relatively large number of time series for the United States. The US federal funds rate is an indicator of the US monetary policy stance. To account for changes in the use of unconventional monetary policies or, equivalently, for those periods in which the US federal funds rate essentially hits the zero lower bound, we use the Wu and Xia rate in separate estimations. That rate coincides with the US federal funds rate if the latter is nonnegative. If the US federal funds rate hits the zero lower bound, the Wu and Xia rate can turn negative, and its distance from zero is a measure of the effectiveness of unconventional monetary policy.[7]

The following list shows each of the US variables we have used and their units of measure.[8] For convenience, we have divided them into three groups – financial, monetary and real – with no direct implications for our models.

Financial

- 3-month US interest rate (percentage);
- 10-year US interest rate (percentage);
- Morgan Stanley Capital International Index (MSCI) (percentage growth);
- Dow Jones 30 Industrial (percentage growth); and

[6] This test verifies that all of the norms of the matrix eigenvalues are strictly less than 1.

[7] Wu and Xia (2013) show that their model can be used to convey the macroeconomic effects of unconventional monetary policy at the zero lower bound.

[8] The source for each series is Haver Analytics except as otherwise noted.

- mortgage rate (percentage).

Monetary

- US federal funds rate (percentage);
- Wu and Xia rate (percentage);
- monetary base (percentage growth);
- M1 (percentage growth);
- M2 (percentage growth);
- currency in circulation (percentage growth);
- Federal Reserve balance sheet assets (percentage growth); and
- Federal Reserve balance sheet liabilities (percentage growth).

Real

- real GDP, seasonally adjusted (SA), billions of chained (2009) dollars (percentage growth);
- current account (per cent of GDP) (difference);
- exports (percentage growth);
- imports (percentage growth);
- consumer credit (percentage growth);
- micro-finance industry lending to private sector (percentage growth);
- private sector credit, over GDP (difference);
- central government budget, over GDP (difference);
- debt outstanding:
  - domestic economy, over SA GDP (difference);
  - households and non-profit institutions serving households, over quarterly GDP at a seasonally adjusted annual rate (SAAR) (difference);
  - non-financial corporations, over quarterly SAAR GDP (difference);
  - US financial corporations/institutions, over quarterly SAAR GDP (difference);
- gross capital formation (percentage growth);
- corporate gross operation surplus (percentage growth);
- gross disposable income (percentage growth);
- gross savings (percentage growth);
- earnings (percentage growth);
- manufacturing (percentage growth);
- manufacturing, excluding construction (percentage growth);
- housing prices (percentage growth);

- housing starts (percentage growth);
- housing permits (percentage growth);
- housing completions (percentage growth);
- shipments (percentage growth);
- retail value (percentage growth);
- wholesale (percentage growth);
- consumer confidence (percentage growth);
- consumer expectations (percentage growth);
- capacity (percentage growth);
- employment (percentage growth);
- labour force (percentage growth); and
- unit labour cost (percentage growth).

EME data

We divided the EME time series into the same three groups. We have taken as given the policy rate status provided by our main data source, Haver Analytics. The variables and units of measure are as follows.

Financial

- exchange rate (percentage growth);
- Morgan Stanley Capital International, MSCI (percentage growth);
- 3-month interest rate (percentage); and
- long-run (10-year) interest rate (percentage).

Monetary

- policy interest rate (percentage);
- M1 (percentage growth);
- M2 (percentage growth); and
- M3 (percentage growth).

Real

- real GDP (quarter percentage growth);
- current account, over GDP (difference);
- manufacturing (percentage growth);
- trade balance, over GDP (difference);
- private consumption expenditure (percentage growth);
- public consumption expenditure (percentage growth); and
- gross capital formation (percentage growth).

Exchange rate arrangements, monetary policy frameworks and financial openness

Each economy varies in the way it sets and conducts its exchange rate and monetary policies. In a broader context, it is relevant to understand the policy framework being implemented in the EME of interest, including its monetary and fiscal policies, capital flow management, and macro-prudential policies. Presumably, the state of these policies is partially captured by the macroeconomic variables we have incorporated into our VAR models. Yet, it is plausible that some aspects not captured explicitly may nonetheless be relevant in interpreting our results.

In this context, we first provide an IMF (2013) classification of the de facto exchange rate arrangements and monetary policy framework for each of the economies in our database (Table 1). We have omitted those classifications in which none of our economies appear. Hong Kong SAR is at one end of the exchange rate arrangements, having a currency board. At the other end, we find countries, such as Chile and Mexico, with an independently floating regime. For the monetary policy framework, Hong Kong SAR is again distinctive, having the US dollar as its anchor; most EMEs in our database maintain an inflation-targeting regime.

IMF de facto exchange rate arrangements and monetary policy regimes[1] Table 1

	Monetary policy framework		
Exchange rate arrangement	Exchange rate anchor USD	Inflation targeting framework	Other
Currency board arrangement	Hong Kong SAR		
Other conventional fixed peg arrangement			
Crawling peg	China		
Managed floating with no pre-determined path for the exchange rate		Colombia Indonesia Peru Thailand	India
Independently floating		Brazil Chile Czech Republic Israel South Korea Mexico Poland South Africa	

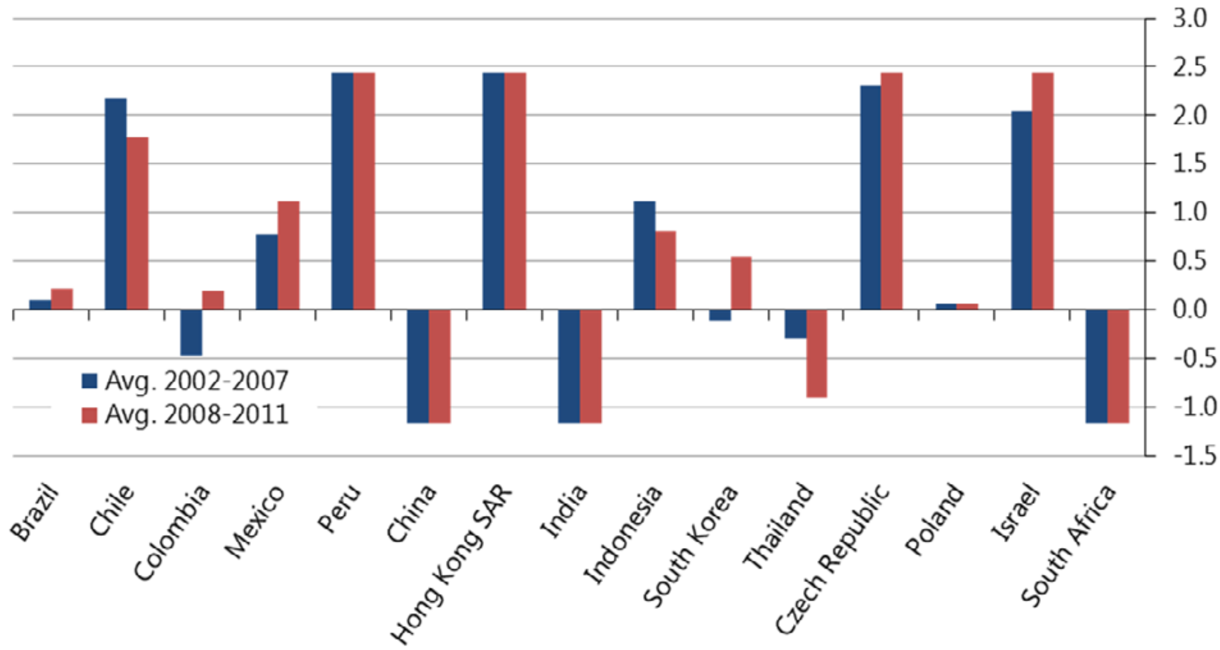
[1] See the Appendix for a more detailed explanation of the monetary policy regimes.

Source: IMF (2013).

Second, we present the Chinn and Ito index (2006) for each country (Figure 1). This index is a *de jure* measurement of the financial openness in an economy. We are interested in knowing the level of financial openness in each EME and whether it has changed in the sample period. Notable cases of an increase in financial openness are Colombia and Korea, whereas Thailand is a case in which financial openness diminished. In general, however, economies that began the sample period either financially open, such as Hong Kong SAR and Israel, or relatively closed, such as China and India, have remained that way.

Chinn-Ito index[1]

Figure 1



[1] Averages for the 2002–07 and 2008–11 periods. The index is a de jure measure of financial openness. A larger and positive (negative) value means greater (lesser) financial openness. Values based on the restrictions on cross-border financial transactions in the IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions*.

Source: Chinn and Ito (2006).

Third, in the Appendix, we provide the IMF's definitions of de facto exchange arrangements shown in Table 1 as well as brief descriptions of exchange arrangements and monetary regimes, as contained in HSBC (2011), for the EMEs in our database. The HSBC information complements, although in some respects does not exactly coincide with, the IMF's classifications.

In sum, a thorough assessment of our results requires an understanding of these arrangements and regimes. As noted, we do not include this information in the model; they are for interpretation purposes only.

Model

We implement an approach similar to that in Bernanke et al (2005) and Boivin et al (2009) by using a factor-augmented VAR (FAVAR). There are at least three important aspects to their approach. First, monetary policy decisions involve the analysis of large amounts of information. The use of principal component analysis (PCA) allows for a systematic use of a wide range of data. Second, some of the variables on which monetary policy depends are unobservable to agents, eg potential output growth. Presumably, the use of PCA partially captures latent variables. Third, we could have attempted to calibrate, say, a dynamic stochastic general equilibrium (DSGE) model. Indeed, there are close relationships between

DSGE and VAR models (eg Fernandez-Villaverde et al (2007)). However, DSGE models require strong identifying restrictions. Moreover, although the VAR approach has not been free of criticism, it provides enough flexibility and has few identification assumptions.

There are several steps to our model's estimation. The first one entails the extraction of the main principal components from our time series set. On one hand, consider x_t, an $nx1$ vector that contains all of the time series (at time t) in the EME database of interest, except for the variables in question, such as the policy rate and exchange rate of the EME at hand. On the other hand, consider x_t^{US}, an $nx1$ vector containing all the time series (at time t) in the US database, except for the federal funds rate.

Next, consider the following approximations:

$$x_t \cong E(x_t) + \upsilon' c_t$$
$$x_t^{us} \cong E(x_t^{us}) + \upsilon'^{us} c_t^{us}$$

where c_t is an $mx1$ vector with the m^{th} first principal components at time t associated with the PCA decomposition of x_t, with $m << n$. Similarly, c_t^{us} is an $mx1$ vector with the $(m^{us})^{th}$ first principal components associated with the PCA decomposition of the time series x_t^{us}, with $m^{us} << n^{us}$.[9] The vector $\upsilon(\upsilon^{us})$ contains the factor loadings associated with the PCA decomposition of the time series in $x_t(x_t^{us})$.

Second, once the c_t^{us} and c_t time series have been obtained, the following vector is constructed by stacking the vectors obtained above along with the economic variables of interest, specifically:

$$y_t = \left[c_t^{US} \; ff_t \; c_t \; lri_i \; i_t \; dfx_t \right]',$$

where we have the $(m^{us})^{th}$ and m^{th} first principal components (c_t^{US} and c_t), the percentage change in the foreign exchange rate (dfx_t), the EME's long-run interest rate (lri_t) and its policy rate (i_t), and the US federal funds rate (ff_t). The long-run rate is not available for all the economies in our database. In a different exercise, we substitute the US federal funds rate for the Wu and Xia rate. Thus, the VAR model for y_t is posited as:

$$y_t = \psi(L) y_{t-1} + \varepsilon_t \tag{1}$$

where $\psi(L)$ is a lag polynomial, and ε_t has mean 0 and variance-covariance matrix Σ.

Third, the VAR model is estimated under some identifying assumptions, for which we provide more details in the next subsection.

9 If m equals n, then $x_t - E(x_t) = \upsilon' c_t$. Similarly, if m^{us} equals n^{us}, then $x_t^{us} - E(x_t^{us}) = \upsilon'^{us} c_t^{us}$.

Our main tests are based on the following extension of model (1):

$$\tilde{y}_t = \tilde{\psi}(L)\,\tilde{y}_{t-1} + u_t \qquad (2)$$

where $\tilde{y}_t = [y_t \;\; ff_t \times d_t]'$ and d_t is a dummy variable defined as follows:

$$d_t = \begin{cases} 0 & \text{if } t < 3Q.2008 \\ 1 & \text{if } t \geq 3Q.2008 \end{cases}$$

Similarly, $\tilde{\psi}(L)$ is a lag polynomial that accommodates the coefficients associated with the $ff_t \times d_t$ variable. The error term u_t has mean 0 and variance-covariance matrix Ω.

To see how the dummy captures a possible structural change in our original VAR model (1), consider the following simplified version of the model. In it, we have included only the variation in the exchange rate (dfx_t), the EME policy rate (i_t), and the US federal funds rate (ff_t). The fourth variable is an auxiliary $(ff_t \times d_t)$, as defined above. Hence, we have:

$$\begin{aligned}
dfx_t &= a_{11}dfx_{t-1} + a_{12}i_{t-1} + a_{13}ff_{t-1} + a_{14}ff_{t-1} \times d_{t-1} + u_{1,t} \\
i_t &= a_{21}dfx_{t-1} + a_{22}i_{t-1} + a_{23}ff_{t-1} + a_{24}ff_{t-1} \times d_{t-1} + u_{2,t} \\
ff_t &= a_{31}dfx_{t-1} + a_{32}i_{t-1} + a_{33}ff_{t-1} + a_{34}ff_{t-1} \times d_{t-1} + u_{3,t} \\
ff_t \times d_{t-1} &= a_{41}dfx_{t-1} + a_{42}i_{t-1} + a_{43}ff_{t-1} + a_{44}ff_{t-1} \times d_{t-1} + \varepsilon_{4,t},
\end{aligned}$$

which can be rewritten as:

$$\begin{aligned}
dfx_t &= a_{11}dfx_{t-1} + a_{12}i_{t-1} + (a_{13} + a_{14} \times d_{t-1})\,ff_{t-1} + u_{1,t} \\
i_t &= a_{21}dfx_{t-1} + a_{22}i_{t-1} + (a_{23} + a_{24} \times d_{t-1})\,ff_{t-1} + u_{2,t} \\
ff_t &= a_{31}dfx_{t-1} + a_{32}i_{t-1} + (a_{33} + a_{34} \times d_{t-1})\,ff_{t-1} + u_{3,t}n \\
ff_t \times d_{t-1} &= a_{41}dfx_{t-1} + a_{42}i_{t-1} + (a_{43} + a_{44} \times d_{t-1})\,ff_{t-1} + u_{4,t}.
\end{aligned}$$

Our first test considers whether the coefficient a_{24} is statistically significantly different from zero. If so, that would provide us with some evidence that the policy rate channel had a structural change after the third quarter of 2008.[10] That is, the coefficient, which measures the contemporaneous effect of ff_t on i_t, would have changed from a_{23} to $a_{23} + a_{24}$. This happens provided that a_{24} is statistically different from zero. Thus, we posit as a null hypothesis that $a_{24} = 0$.

Similarly we can test for a structural change in the exchange rate channel by statistically assessing whether $a_{14} = 0$. In addition, we can jointly test whether there has been a change in both channels by considering the null hypothesis: $a_{14} = a_{24} = 0$.

[10] An underlying assumption is that we know a priori the period in which the structural change could have taken place.

In sum, we perform each of these tests as linear restrictions on the following system:[11]

$$\tilde{y}_t = \tilde{\psi}(L)\,\tilde{y}_{t-1} + u_t .$$

As an important exercise, we plot the impulse-response functions (IRFs) given a 25 basis point orthogonal shock to the US federal funds rate under model (1) and again under model (2). This allows for a visual comparison of, on one hand, the IRFs from the model in which it is assumed that there has been no structural change in the channels, with, on the other hand, the IRFs from the model that allows for a possible structural change between the Q1 2003–Q2 2008 period and the Q3 2008–Q4 2013 period.[12] Compared with the linear tests, the IRFs depend on the specific Cholesky decomposition considered.

An underlying assumption is that the rest of the coefficients do not go through a structural change. Although it would be desirable to consider a more general hypothesis in order to capture a possible structural change in all of $\psi(L)$, this exercise is limited by the relatively short length of the time series.[13]

Estimation

For estimation purposes, we first divide our database into two sets, as suggested by our notation. In one set, we have the US time series, except for the US federal funds rate (or the Wu and Xia rate). The second set contains an EME's time series except for its policy rate, exchange rate, and long-term interest rate.

Second, by decomposing the US variables using PCA and also decomposing the EME's variables using PCA, we obtain a pair of sets of principal components. The components associated with the US data are denoted by c_t^{US}, and those associated with the EME's data by c_t.

Third, we construct the vector $y_t = [c_t^{US} \;\; ff_t \;\; c_t \;\; lri_t \;\; i_t \;\; dfx_t]'$, having used the first three principal components from each set.[14] Accordingly, the c_t^{US} and c_t entries in y_t stand for vectors, and the rest stand for scalars for each t.

11 These linear restrictions, under the null hypothesis, follow an F distribution, with the number of restrictions (v_1), and the number of observations minus the number of parameters being estimated (v_2), as degrees of freedom; ie its distribution is $F(v_1, v_2)$.

12 We have estimated models (1) and (2) under the identifying assumptions of orthogonal shocks, as will be explained later.

13 In the VAR model, if one has n variables, $n + n^2 + n(n-1)/2$ parameters have to be estimated. Adding one more variable implies that such number increases by $4n-1$. Assessing a possible structural change in all of $\psi(L)$ calls for n^2 additional coefficients. Regime switching calls for n^2 additional coefficients plus the transition probabilities.

14 We normally choose three components, for the following reasons: (i) On average, three components explain more than 50% of the accumulated variance of the whole data set; in general, the increment of accumulated variance tended to markedly drop by the fourth component. And (ii) using the same number of lags for each EME has the advantage of maintaining a level of comparability across EMEs.

Fourth, we assess the number of lags by means of Schwartz's Bayesian information criterion (SBIC). In most cases, SBIC points to a lag of 1. This is reasonable given that we are using principal components, which can proxy the lagged variables' dynamics.[15] For those cases in which the SBIC indicates a greater lag, we instead increase the number of EME and US components until the test points to a lag of 1.[16]

Fifth, the systems $y_t = \psi(L) y_{t-1} + \varepsilon_t$ and $\tilde{y}_t = \tilde{\psi}(L) \tilde{y}_{t-1} + u_t$ are estimated. As a final step, tests and IRFs are calculated and assessed.

Also, we redo the steps above with the Wu and Xia rate instead of the US federal funds rate to allow the model to capture the effects of unconventional monetary policies.

A VAR estimation has some identifying restrictions. In particular, one has to make a choice regarding the identification of shocks. In our case, we assume orthogonal shocks by applying the Cholesky decomposition to the variance-covariance matrices. To this end, we order our VAR variables by their speed of adjustment, as suggested by our notation $y_t = [c_t^{US} \;\; ff_t \;\; c_t \;\; lri_t \;\; i_t \;\; dfx_t]'$. Thus, on impact, we have assumed that the adjustment of the exchange rate is followed by that of interest rates and then of the EME's components. Then the US federal funds rate adjusts, followed by the US components.[17] Thus, in general, the EME's variables respond on impact to changes in US variables but not the other way around.

For identification purposes, we could have estimated a structural VAR (SVAR). We have nonetheless preferred a Cholesky decomposition, for two reasons: (i) it is equivalent to a just-identified SVAR and thus generally implies a less restrictive scheme; and (ii) we find it more suitable, at least initially, as a SVAR would generally require stronger identification assumptions and consequently more a priori knowledge about the variables' relationships.

Results

To set the stage, we first consider the cross-correlations between the US federal funds rate and each of our three variables of interest for each EME – the policy rate (Figure 2), the exchange rate (Figure 3) and long-term interest rates (Figure 4) – for the Q1 2003–Q2 2008 and the Q3 2008–Q4 2013 periods. The autocorrelations have

15 In fact, Stock and Watson (1999) have referred to (1) as a dynamic model.

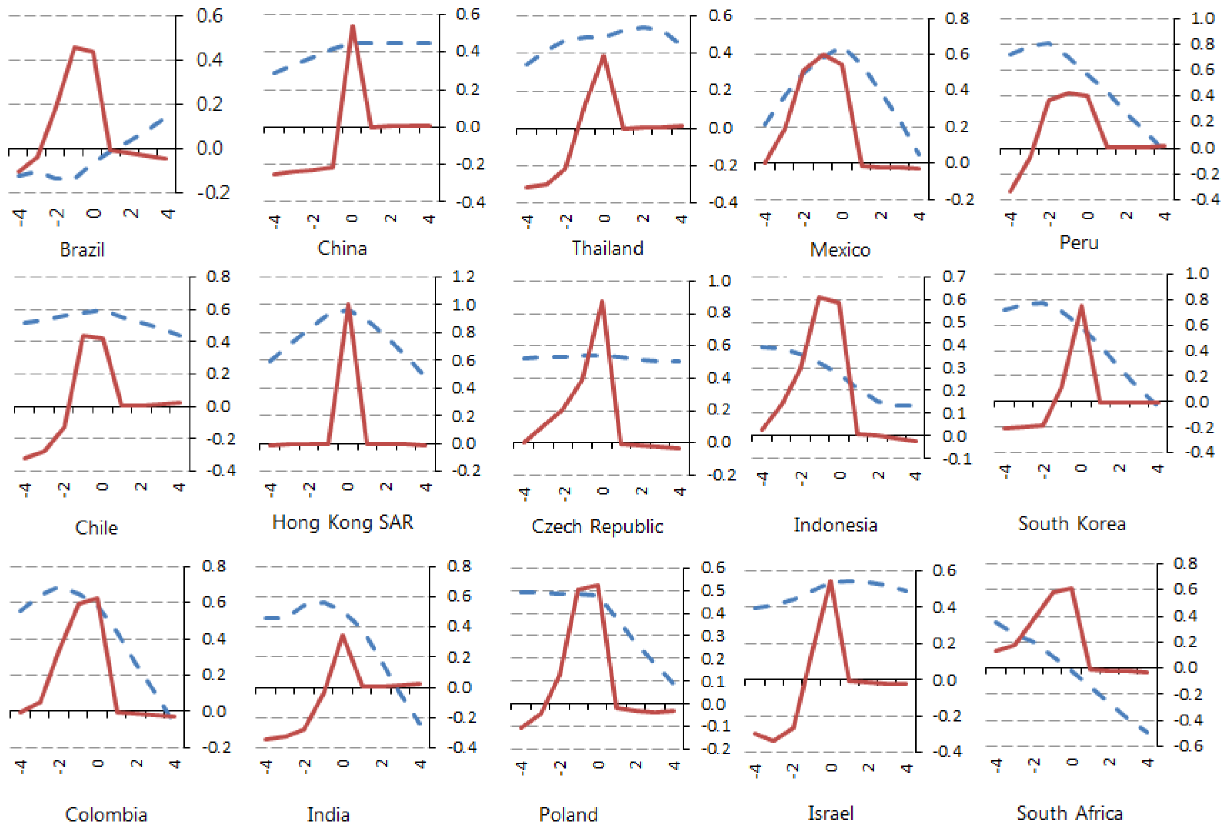
16 Specifically, for Chile, Peru and China, the SBIC calls for a lag greater that 1. The issue is that, given our time series' short length, increasing the lag might not be feasible because of the large number of coefficients that would need to be estimated. Adding lags makes estimated coefficients grow exponentially, and adding components makes them grow linearly; efficiency calls for the latter approach. Thus, to mitigate the problem, we increase the number of components, which tends to reduce the lag indicated by SBIC.

17 Such order might be better suited to some EMEs than others. We nonetheless keep the same order to maintain comparability across EMEs.

lags and leads of up to four quarters. These statistics provide us with initial clues on possible changes in the relationships between the variables.[18]

Autocorrelations of the US federal funds rate with each EME's policy rate[1]

Figure 2



[1] The maximum lead and lag time is four quarters. The blue (dashed) lines correspond to the Q1 2003–Q2 2008 period, the red (solid) lines to the Q3 2008–Q4 2013 period.

First, the autocorrelations with the policy rates (which tend to peak when the lag is set to zero) suggest that some economies, including Brazil, could have experienced a change in such relationships between the two periods. Second, they also indicate that some economies, including Colombia and Mexico, might have maintained such relationships over the two periods.

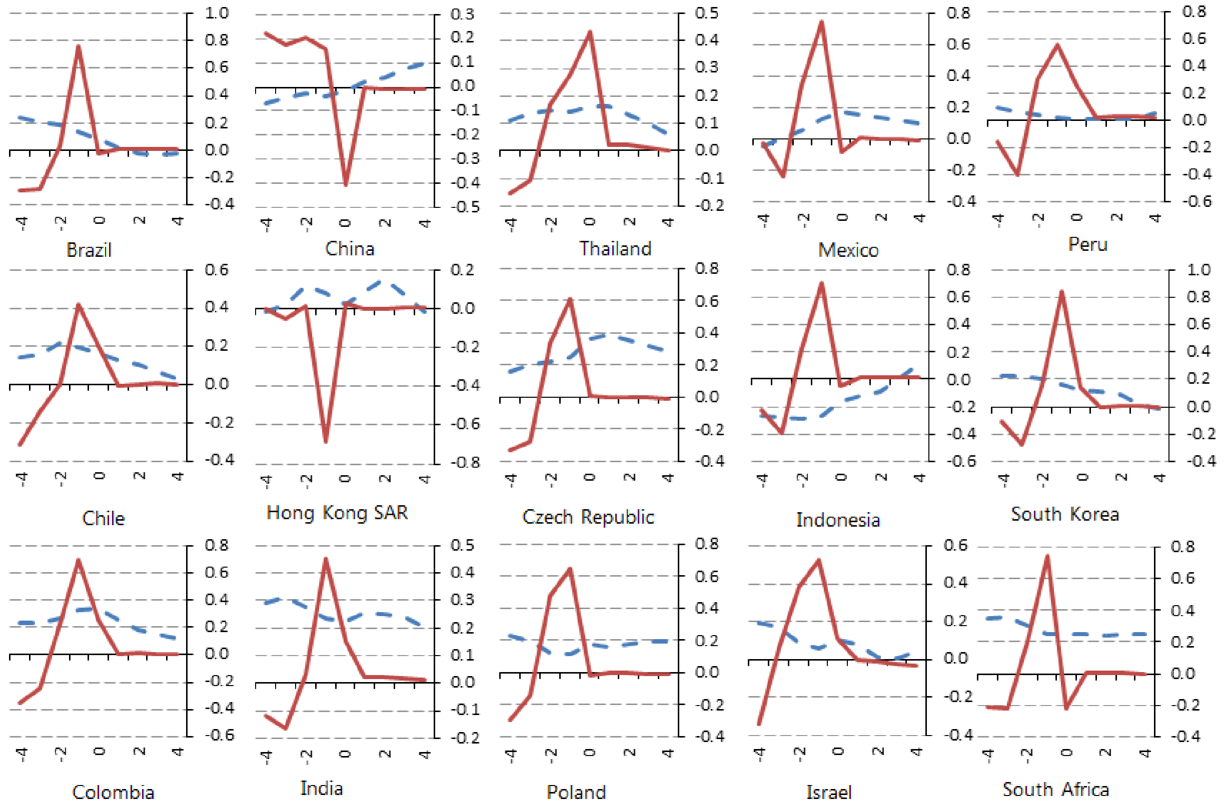
Second, the autocorrelations with the exchange rate are generally smaller in magnitude than those with policy rates. In the Q3 2008–Q4 2013 period, their values tended to be less stable through the leads and lags. Interestingly enough, some of the autocorrelations in this period share a similar pattern in several economies, peaking at around two lags and then dropping towards the zero lag mark.[19]

[18] They of course capture unconditional moments.

[19] This is notable for Chile, Colombia, Mexico and Peru; Indonesia and Thailand; the Czech Republic and Poland; and Israel and South Africa.

Autocorrelations of the US federal funds rate with variations in each EME's exchange rate[1]

Figure 3



[1] The maximum lead and lag time is four quarters. The blue (dashed) lines correspond to the Q1 2003–Q2 2008 period, the red (solid) lines to the Q3 2008–Q4 2013 period.

Third, the autocorrelations with the long-term rates on average increased in the second period, eg for the Czech Republic, Indonesia, Korea, Mexico, Poland and South Africa. In contrast, two economies have seen these autocorrelations decrease: India and, except for the zero lag, Thailand.

As a next step, we analyse the autocorrelations between the Wu and Xia rate and the three variables: the policy rate (Figure 5), the exchange rate (Figure 6) and long-run interest rates (Figure 7). Note that for the first period, Q1 2003–Q2 2008, the autocorrelations coincide with those in our last exercise, as the US federal funds rate and the Wu and Xia rate are essentially the same during the period.

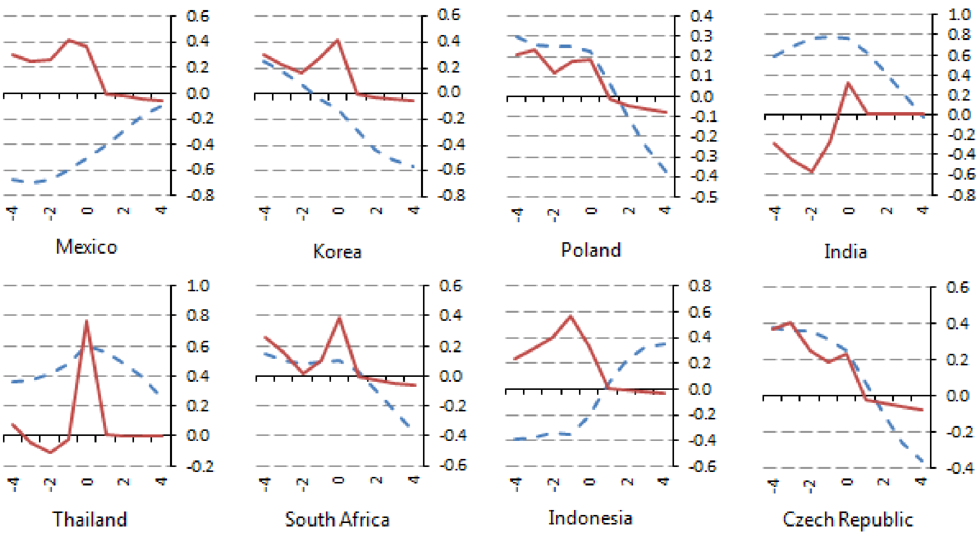
First, regarding the autocorrelations with the policy rates, those with the Wu and Xia rate have decreased in magnitude. This suggests that the policy rates channel could have lost some importance for the second period, Q3 2008–Q4 2013.

Second, the autocorrelations with the exchange rates do not have a stable pattern in the second period.

Third, the autocorrelations with the long-run rates show an overall increase in value, both for the leads and lags; the two exceptions are India and Thailand, which seem to have swapped signs. This result suggests that the long-run rate channel might have gained a greater role in the period after the crisis.

Autocorrelations of the US federal funds rate with each EME's long-run interest rate[1]

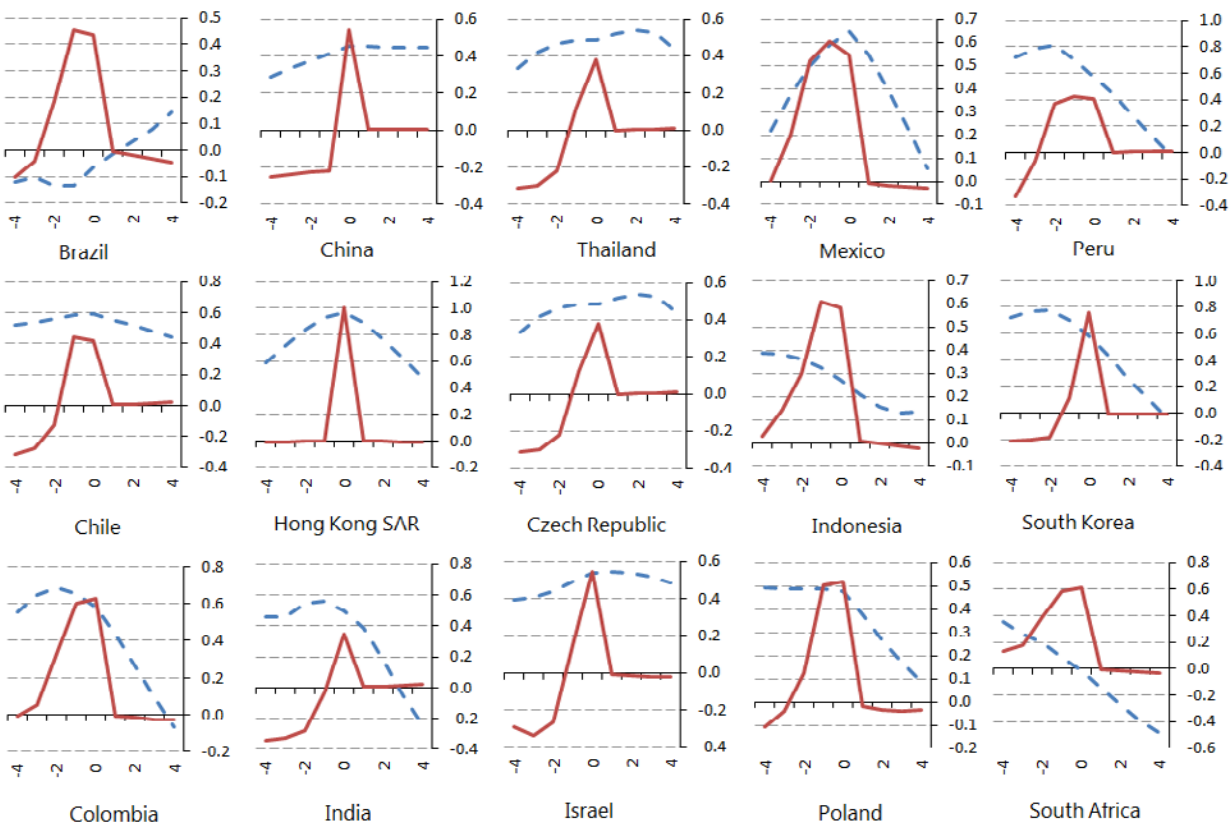
Figure 4



[1] The maximum lead and lag time is four quarters. The blue (dashed) lines correspond to the Q1 2003–Q2 2008 period, the red (solid) lines to the Q3 2008–Q4 2013 period.

Autocorrelations of the Wu and Xia rate with each EME's policy rate[1]

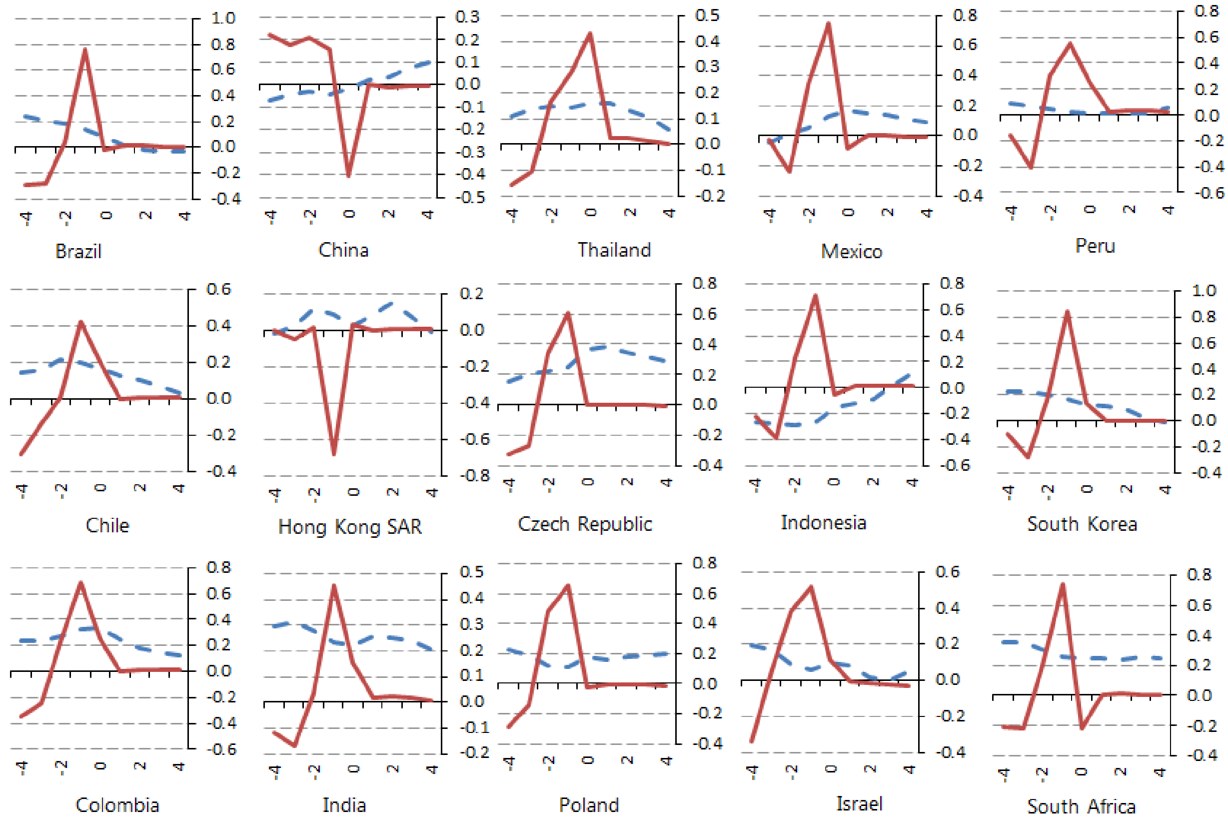
Figure 5



[1] The maximum lead and lag time is four quarters. The blue (dashed) lines correspond to the Q1 2003–Q2 2008 period, the red (solid) lines to the Q3 2008–Q4 2013 period. See Wu and Xia (2013) for the Wu and Xia rate.

Autocorrelations of the Wu and Xia rate with variations in each EME's exchange rate[1]

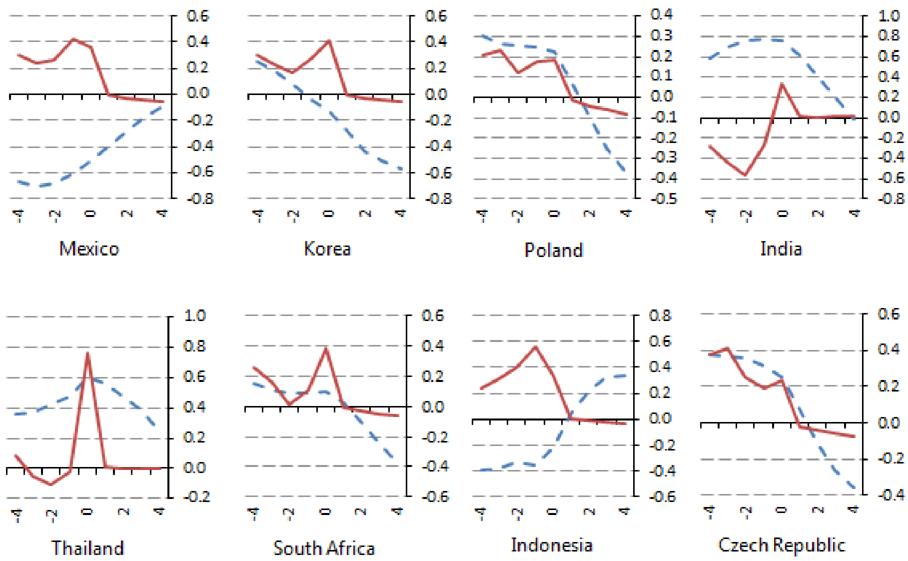
Figure 6



[1] The maximum lead and lag time is four quarters. The blue (dashed) lines correspond to the Q1 2003–Q2 2008 period, the red (solid) lines to the Q3 2008–Q4 2013 period. See Wu and Xia (2013) for the Wu and Xia rate.

Autocorrelations of the Wu and Xia rate with each EME's long-run interest rate[1]

Figure 7



[1] The maximum lead and lag time is four quarters. The blue (dashed) lines correspond to the Q1 2003–Q2 2008 period, the red (solid) lines to the Q3 2008–Q4 2013 period. See Wu and Xia (2013) for the Wu and Xia rate.

In sum, although autocorrelations are only a first step, they provide some initial evidence that the eruption of the global financial crisis and the policy responses could have affected to some extent the channels by which US monetary policy shocks are transmitted to EMEs.

Linear regression models

As a next step, we assess the following linear regression models. Specifically, we posit the following three benchmark regressions for each of the corresponding economies:[20]

$$
\begin{aligned}
i_t &= \beta_{10} + \beta_{11} c_t + \beta_{12} ff_t + e_{1,t} \\
dfx_t &= \beta_{20} + \beta_{21} c_t + \beta_{22} ff_t + e_{2,t} \\
lri_t &= \beta_{30} + \beta_{31} c_t + \beta_{32} ff_t + e_{3,t}
\end{aligned}
\tag{3}
$$

where, as above, c_t is a vector with the first three principal components of the EME's macroeconomic variables, ff_t is the US federal funds rate, $\beta_{.,0}$ are constants and $e_{.,t}$ are the error terms. We separately estimate them for Q1 2003–Q2 2008 and for Q3 2008–Q4 2013. Our focus is on the statistics of $\beta_{.,2}$ between those two periods, in particular analysing the associated *t*-statistics and R^2 (Table 2).

The first regression can be seen as a rule that is more general than the Taylor rule. We depart from the Taylor rule for two reasons. First, a given country may not strictly follow a Taylor rule to set its policy rate. Second, in the same vein, using principal components allows us to have a model that is similar across economies and the variables of interest, and thus allowing us to make closer comparisons.

For the policy rate, the results are as follows. For three EMEs, the coefficients lost their statistical significance; for four others, they attained statistical significance; for the rest, the coefficients maintained their status. Moreover, while the explained variance generally increased between periods, the *t*-statistics decreased on average. In sum, the relevance of the policy rate channel seems to depend on the economy in question.

Second, the coefficients associated with the exchange rate channel are in general not statistically significant in either period. The exceptions to this seem to be the coefficient for China, which maintained statistical significance; for Indonesia and Poland, which attained statistical significance; and for South Africa, which has swapped its sign. Hence, this test indicates that, in the majority of cases, the exchange rate channel does not seem to have had a structural change.

Third, for long-term rates, three economies (out of the eight EMEs for which we have the requisite rate data) have a significant coefficient in the first period – India, Mexico and Thailand. Mexico and India lose statistical significance in the second period, while Thailand maintains it. South Africa's coefficient attains statistical significance in the second period.

For the period in which the US federal funds rate hits the zero lower bound, we consider the following model:

[20] The long-term rate model is estimated only for the eight economies for which we have data on long-term rates.

$$
\begin{aligned}
i_t &= \gamma_{10} + \gamma_{11} c_t + \gamma_{12} xx_t + \varepsilon_{1,t} \\
dfx_t &= \gamma_{20} + \gamma_{21} c_t + \gamma_{22} xx_t + \varepsilon_{2,t} \\
lri_t &= \gamma_{30} + \gamma_{31} c_t + \gamma_{32} xx_t + \varepsilon_{3,t}
\end{aligned} \tag{4}
$$

where, instead of the US federal funds rate, we use the Wu and Xia rate, denoted by xx_t. Hence, our focus is now on estimates $\gamma_{.,2}$. We report the associated *t*-statistics and the R^2 in Table 3.

Coefficient statistics for the US federal funds rate[1] — Table 2

		EME macroeconomic variable and regression period					
		Policy Rate		Exchange Rate		Long-run interest rate	
		Period 1	Period 2	Period 1	Period 2	Period 1	Period 2
Brazil	*t*	–1.815	**2.667**	1.409	0.141		
	R^2	0.767	0.570	0.434	0.803		
Chile	*t*	**4.130**	0.664	0.755	–0.010		
	R^2	0.743	0.564	0.052	0.218		
Colombia	*t*	**4.145**	**3.567**	0.257	1.328		
	R^2	0.443	0.576	0.148	0.381		
Mexico	*t*	**4.588**	**4.304**	–1.686	–1.381	**–2.470**	1.533
	R^2	0.528	0.855	0.284	0.696	0.334	0.246
Peru	*t*	**2.286**	1.784	1.799	0.296		
	R^2	0.734	0.790	0.223	0.404		
China	*t*	**2.872**	**4.096**	**–2.057**	**–2.391**		
	R^2	0.580	0.802	0.501	0.508		
Hong Kong SAR	*t*	**155.590**	–	1.500	0.073		
	R^2	0.999	1.000	0.098	0.627		
India	*t*	0.714	1.566	–0.934	0.513	**4.792**	1.468
	R^2	0.680	0.538	0.289	0.476	0.823	0.331
Indonesia	*t*	0.714	**2.822**	–0.578	**–2.451**	–1.110	1.415
	R^2	0.285	0.599	0.360	0.805	0.569	0.346
South Korea	*t*	**2.156**	**3.838**	0.366	0.253	–1.380	0.820
	R^2	0.453	0.722	0.156	0.835	0.387	0.317
Thailand	*t*	**2.746**	0.955	0.559	1.688	**2.684**	**4.297**
	R^2	0.412	0.325	0.222	0.443	0.429	0.606
Czech Republic	*t*	**2.444**	**9.227**	1.420	–1.041	1.464	1.045
	R^2	0.295	0.893	0.130	0.686	0.214	0.099
Poland	*t*	1.805	**2.711**	0.863	**–2.487**	1.000	0.093
	R^2	0.645	0.783	0.127	0.754	0.606	0.339
Israel	*t*	**2.895**	**3.646**	1.222	0.333		
	R^2	0.215	0.483	0.400	0.250		
South Africa	*t*	0.071	**3.445**	**2.375**	**–3.288**	1.129	**2.253**
	R^2	0.543	0.675	0.199	0.537	0.412	0.256

[1] For equation (3). Period 1 = Q1 2003–Q2 2008. Period 2 = Q3 2008–Q4 2013. *T*-statistics are for the coefficients $\beta_{.,2}$; those in bold are for statistically significant coefficients at a 90% confidence level.

For the policy rate channel, the coefficient for only two economies – Indonesia and South Africa – attained statistical significance. For the rest of the EMEs, the coefficient either lost or maintained significance.

Coefficient statistics for the Wu and Xia rate[1]

Table 3

		EME macroeconomic variable and regression period					
		Policy Rate		Exchange Rate		Long-run interest rate	
		Period 1	Period 2	Period 1	Period 2	Period 1	Period 2
Brazil	t	**–1.996**	**2.381**	1.648	–0.371		
	R^2	0.771	0.541	0.445	0.805		
Chile	t	**5.411**	–0.633	0.412	0.808		
	R^2	0.811	0.563	0.028	0.249		
Colombia	t	**4.987**	**3.988**	0.266	**2.113**		
	R^2	0.523	0.618	0.148	0.463		
Mexico	t	**4.445**	**9.515**	–1.905	–0.984	**–3.108**	**5.768**
	R^2	0.515	0.953	0.295	0.679	0.409	0.720
Peru	t	**2.714**	–0.252	1.645	0.421		
	R^2	0.760	0.749	0.215	0.407		
China	t	**3.702**	**2.205**	**–2.693**	**–2.254**		
	R^2	0.645	0.689	0.555	0.493		
Hong Kong SAR	t	**38.444**	**3.118**	1.463	–0.202		
	R^2	0.981	0.510	0.095	0.627		
India	t	1.116	–1.044	–1.032	0.588	**5.793**	–1.478
	R^2	0.688	0.501	0.292	0.479	0.853	0.332
Indonesia	t	1.009	**5.719**	–0.640	–0.943	–1.395	**5.986**
	R^2	0.299	0.804	0.362	0.746	0.589	0.773
South Korea	t	**3.008**	0.383	0.845	0.907	–0.928	**4.254**
	R^2	0.521	0.472	0.174	0.842	0.361	0.666
Thailand	t	**2.414**	–1.383	0.646	0.594	**2.538**	**2.440**
	R^2	0.382	0.362	0.225	0.358	0.417	0.379
Czech Republic	t	**3.024**	**6.348**	1.085	–0.780	1.802	**4.147**
	R^2	0.357	0.808	0.105	0.677	0.243	0.549
Poland	t	**2.457**	1.551	0.751	–0.547	1.475	**2.314**
	R^2	0.675	0.724	0.121	0.665	0.622	0.505
Israel	t	**2.875**	1.203	0.826	1.126		
	R^2	0.212	0.132	0.387	0.301		
South Africa	t	0.093	**9.870**	**2.281**	–1.222	0.962	**4.350**
	R^2	0.543	0.920	0.194	0.290	0.405	0.551

[1] For equation (4). Period 1 = Q1 2003–Q2 2008. Period 2 = Q3 2008–Q4 2013. *T*-statistics are for the coefficients $\gamma_{.,2}$; those in bold are for statistically significant coefficients at a 90% confidence level. See Wu and Xia (2013) for the Wu and Xia rate.

For the exchange rate channel, only two EMEs have a statistically significant coefficient in the first period – China, which attained it, and Colombia, which maintained it. South Africa's coefficient lost its significance between the two periods.

In these exercises, the majority of EMEs do not present strong evidence of a structural change in the exchange rate channel.

For the long-term rate channel, five of the eight EMEs for which we have data gained a significant coefficient in the second period. This performance contrasts somewhat with the results obtained with the US federal funds rate. The result, however, does not strictly hold in the tests we perform, as discussed next.

Tests

The hypotheses we test are posited in terms of linear restrictions in (2), as follows:

1. $H_0 : a_{i_t, ff_t \times d_t} = 0$. A direct test of no structural change in the *policy rate channel*.
2. $H_0 : a_{dfx_t, ff_t \times d_t} = 0$. A direct test of no structural change in the *exchange rate channel*.
3. $H_0 : a_{lri_t, ff_t \times d_t} = 0$. A direct test of no structural change in the *long-term rate channel*.
4. $H_0 : a_{i_t, ff_t \times d_t} = 0$ and $a_{lri_t, ff_t \times d_t} = 0$. A joint direct test of no structural change in the policy rate, exchange rate, *and* long-run rate channels (the "joint all rates test").
5. $H_0 : a_{r, ff_t \times d_t} = 0$ for all rows r. An indirect test of no structural change in any channel (the "joint all variables test"). This test involves other channels whose mechanisms are, in our view, embedded in the principal components. By construction, tests 4 and 5 are more stringent than tests 1, 2 and 3.[21]

For the US policy rate, we run the tests using the US federal funds rate and, on a separate set of FAVARs, the Wu and Xia rate. The latter accounts for the possible effects of unconventional monetary policy and allows us to compare the effects of solely traditional monetary policy with those of unconventional policy. We present our estimates in that order.

The *p*-value results of our tests (Table 4) show that, first, eight EMEs reject null hypothesis 1 ($a_{i_t, ff_t \times d_t} = 0$), two at a 10% confidence level and six at a 5% confidence level, providing some evidence of a change in the policy rate channel.

Second, seven EMEs reject null hypothesis 2 ($a_{dfx_t, ff_t \times d_t} = 0$), one at a 10% confidence level and six at a 5% confidence level, which indicates a possible change in the exchange rate channel.

Third, for the long-run rate channel, four of eight EMEs reject null hypothesis 3 at a 5% confidence level.

[21] Under the null hypothesis, all of the them have an *F* distribution with (i) the number of restrictions and (ii) the number of observations minus the number of parameters being estimated as degrees of freedom. These tests are invariant to the specific Cholesky decomposition one opts to use.

Fourth, 11 EMEs reject null hypothesis 4 (the joint all rates test), one at a 10% confidence level and ten at a 5% confidence level.

Interestingly enough, some economies, including the Czech Republic and Indonesia, reject some individual tests but fail to reject the joint all rates test. Conversely, some others, including Poland, fail to reject rate tests 1, 2 and 3 but reject the joint all rates tests. These results are a consequence of the behaviour of the confidence regions in the joint tests compared with that of the confidence intervals in the individual tests.

P-values of test results for the US federal funds rate[1] — Table 4

	EME macroeconomic variable				
	Policy Rate	Exchange Rate	Long-run interest rate	Joint all rates	Joints all variables
Brazil	0.5602	0.0060		0.0184	0.0001
Chile	0.0005	0.1601		0.0007	0.0004
Colombia	0.0090	0.0129		0.0019	0.0000
Mexico	0.0016	0.1306	0.0353	0.0020	0.0002
Peru	0.0561	0.0164		0.0143	0.0004
China	0.0000	0.0086		0.0000	0.0000
Hong Kong SAR	0.0330	0.4491		0.0905	0.0008
India	0.6543	0.3749	0.0010	0.0090	0.0000
Indonesia	0.5389	0.0428	0.2984	0.2085	0.0000
South Korea	0.6157	0.0356	0.0010	0.0020	0.0000
Thailand	0.6602	0.7729	0.0020	0.0181	0.0000
Czech Republic	0.0273	0.5930	0.1953	0.1338	0.0000
Poland	0.0684	0.3806	0.2523	0.0302	0.0018
Israel	0.8682	0.0633		0.1351	0.0000
South Africa	0.4836	0.9771	0.1296	0.3308	0.0010

[1] For the *F*-tests of the linear restrictions set on each FAVAR, as in equation (2). Darkest shading indicates a *p*-value of less than 0.05; lighter shading, a *p*-value between 0.05 and 0.1.

Fifth, all EMEs reject null hypothesis 5 (the joint all variables test) at a 5% confidence level. Our presumption is that these tests capture other channels we are not measuring directly. Hence, we take these results as evidence of a possible structural change in channels we are not explicitly modelling but that we are nonetheless capturing.

In sum, whether an EME experienced a structural change in any of the three channels we are assessing depends on the EME in question.

We next consider the same battery of tests using the Wu and Xia rates in place of the US federal funds rate (Table 5). First, for the policy rate, five EMEs fail to reject null hypothesis 1, one at a 10% confidence level and four at a 5% confidence level. Thus, whether a policy rate channel changed or not depends on the EME in question.

Second, for the exchange rate, four EMEs reject null hypothesis 2, one at a 10% confidence level and three at a 5% confidence level. Here too, although less than one third of EMEs reject this test, it seems that a change in the channel generally depends on the specific economy being considered.

Third, of the eight EMEs tested for the long-run rate channel, three reject null hypothesis 3, one (Thailand) at a 10% confidence level and two (India and Korea) at a 5% confidence level. Again, these results indicate a country-dependent change.

P-values of test results for the Wu and Xia rate[1] Table 5

	EME macroeconomic variable				
	Policy Rate	Exchange Rate	Long-run interest rate	Joint all rates	Joints all variables
Brazil	0.0500	0.1344		0.0995	0.0019
Chile	0.0178	0.4308		0.0432	0.0081
Colombia	0.1117	0.3804		0.2132	0.0220
Mexico	0.0208	0.0263	0.4784	0.0600	0.0007
Peru	0.1269	0.1383		0.1351	0.0046
China	0.0001	0.0510		0.0003	0.0000
Hong Kong SAR	0.2982	0.1561		0.2691	0.0048
India	0.4301	0.3702	0.0064	0.0362	0.0000
Indonesia	0.6560	0.0067	0.8165	0.0128	0.0001
South Korea	0.7414	0.0240	0.0350	0.0294	0.0038
Thailand	0.4728	0.7515	0.0802	0.3104	0.0160
Czech Republic	0.0130	0.7368	0.7209	0.0334	0.0040
Poland	0.3420	0.5417	0.6867	0.6594	0.2394
Israel	0.9503	0.2702		0.5243	0.0304
South Africa	0.4235	0.6066	0.7384	0.8167	0.0660

[1] For the *F*-tests of the linear restrictions set on each FAVAR, as in equation (2). Darkest shading indicates a *p*-value of less than 0.05; lighter shading, a *p*-value between 0.05 and 0.1. See Wu and Xia (2013) for the Wu and Xia rate.

Fourth, for the joint all rates test, eight EMEs reject null hypothesis 4, two at a 10% confidence level and six at a 5% confidence level.

Fifth, for the joint all variables test, all EMEs except Poland reject null hypothesis 5. Again, this result underscores the possible structural changes in what we think are some of the channels we have not modelled explicitly.

In sum, our tests provide some evidence that a share of economies seem to have experienced a change in their policy rate, exchange rate, or long-run interest rate channels. Moreover, evidence of possible structural changes appears for channels that we are not explicitly modelling. Another plausible interpretation is that such changes are second-round effects of the respective channels. That is, variations in the US federal funds rate affect the US components, which in turn affect EMEs through one or more of the three channels we test.

Importantly, the results for the Wu and Xia rate do not markedly differ from those for the US federal funds rate. This contrasts with the test of the linear regression models.

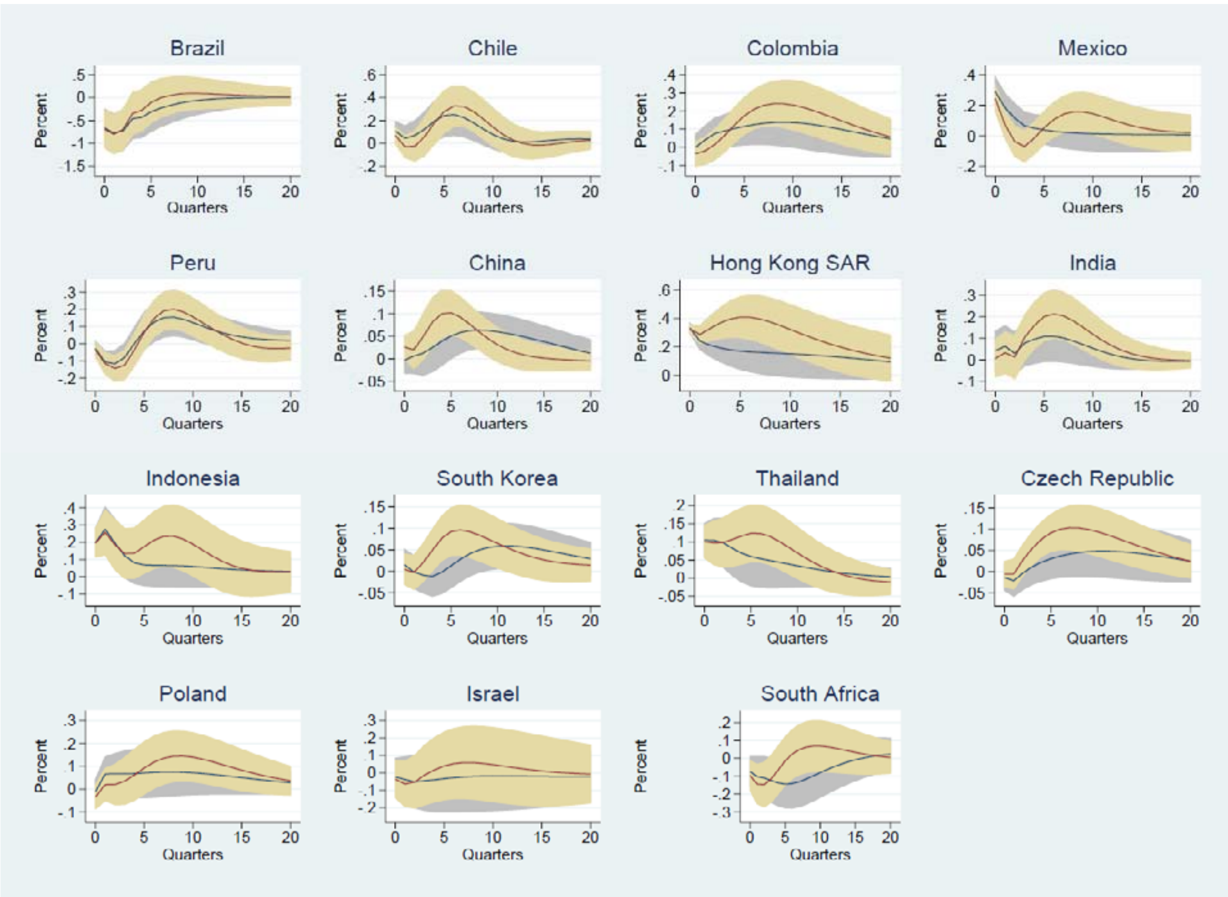
Impulse-Response Functions

The tests just presented are a statistical way to assess evidence of a possible structural change in the selected transmission channels. The VAR's IRFs provide an additional perspective: They show possible changes in the paths and magnitudes of the complete variables between the periods we have analysed. And, of course, IRFs capture variations above and beyond changes in each of the specific coefficients we have associated with each channel.

There are at least two key IRF features to watch. The first is the relationship between the responses of the policy rate and those of the exchange rate. In general, they should maintain some degree of consistency. One can think of this consistency in terms of uncovered interest rate parity, as explained in further detail below. Second, and perhaps most important, is the comparison of the response dynamics under (i) the assumption in model (1) of no structural change after the crisis and (ii) the assumption in model (2) of a possible structural change after the crisis, that is, in the Q3 2008–Q4 2013 period.

Impulse response functions of each EME's policy rate to a US federal funds rate shock[1]

Figure 8



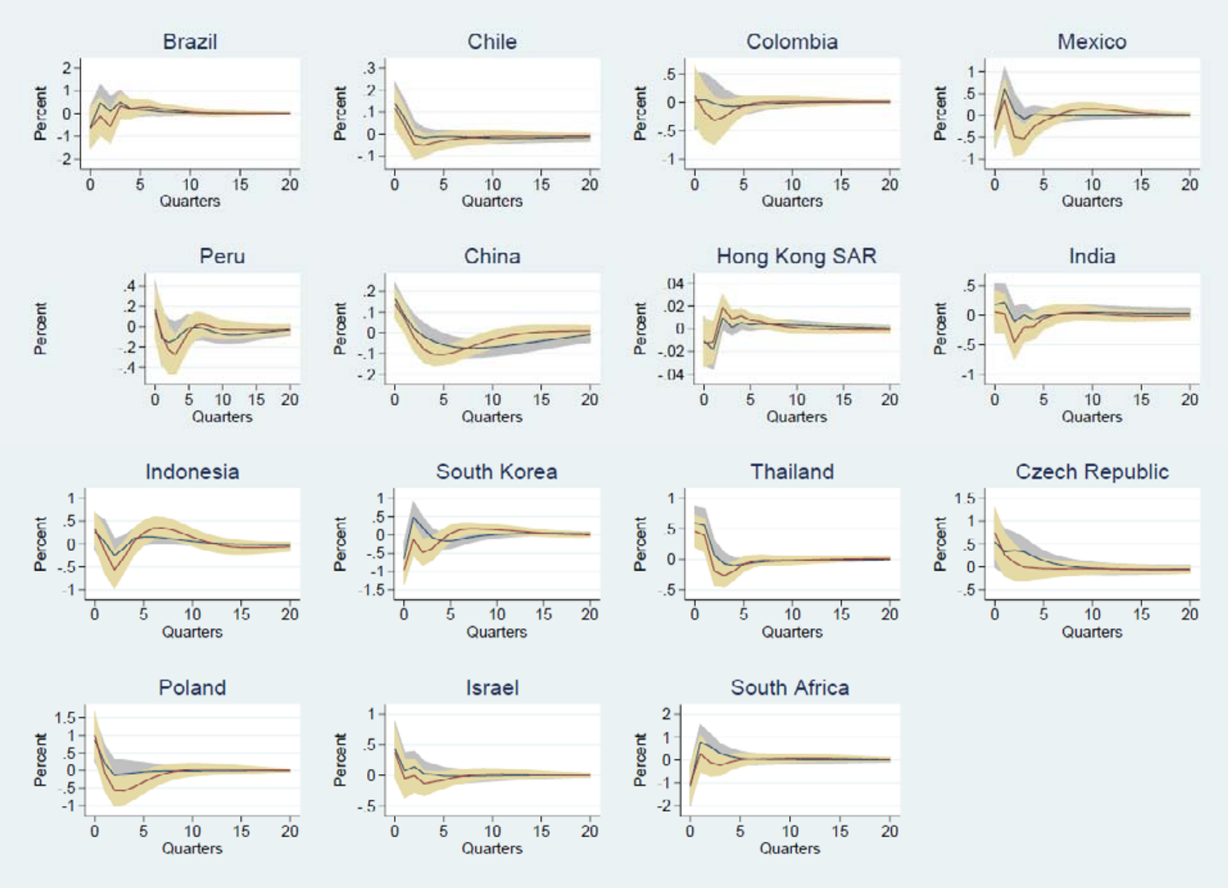
[1] The impulse is a 25 basis point orthogonal shock to the US federal funds rate. The blue lines are obtained from model (1), which assumes no change in the policy rate channel. The red lines are obtained from model (2), which allows for a change in the policy rate channel and potentially in other channels; any such structural changes are assumed to have occurred in the Q3 2008–Q4 2013 period. Confidence intervals are set at the 70% level.

In the case of the policy rate channel (Figure 8), the IRFs arise from an orthogonal shock of 25 basis points to the US federal funds rate. The responses from model (1) and model (2) are both shown.

A bird's-eye view of Figure 8 suggests that, given a positive shock to the US federal funds rate, the policy rate response strongly depends on the country in question. For instance, the magnitude of the responses across EMEs varies somewhat. In most cases, the shock eventually leads to an increase in the policy rate.

Impulse response functions of variations in each EME's exchange rate to a US federal funds rate shock[1]

Figure 9



[1] The impulse is a 25 basis point orthogonal shock to the US federal funds rate. The blue lines are obtained from model (1), which assumes no change in the exchange rate channel. The red lines are obtained from model (2), which allows for a change in the exchange rate channel and potentially in other channels; any such structural changes are assumed to have occurred in the Q3 2008–Q4 2013 period. Confidence intervals are set at the 70% level.

First, the difference in the effect of the US shock between the pre-crisis and post-crisis periods varies across some economies. Moreover, if we interpret our confidence intervals literally (at the 70% level), the channels for Indonesia and Poland attain statistical significance only in the second period.

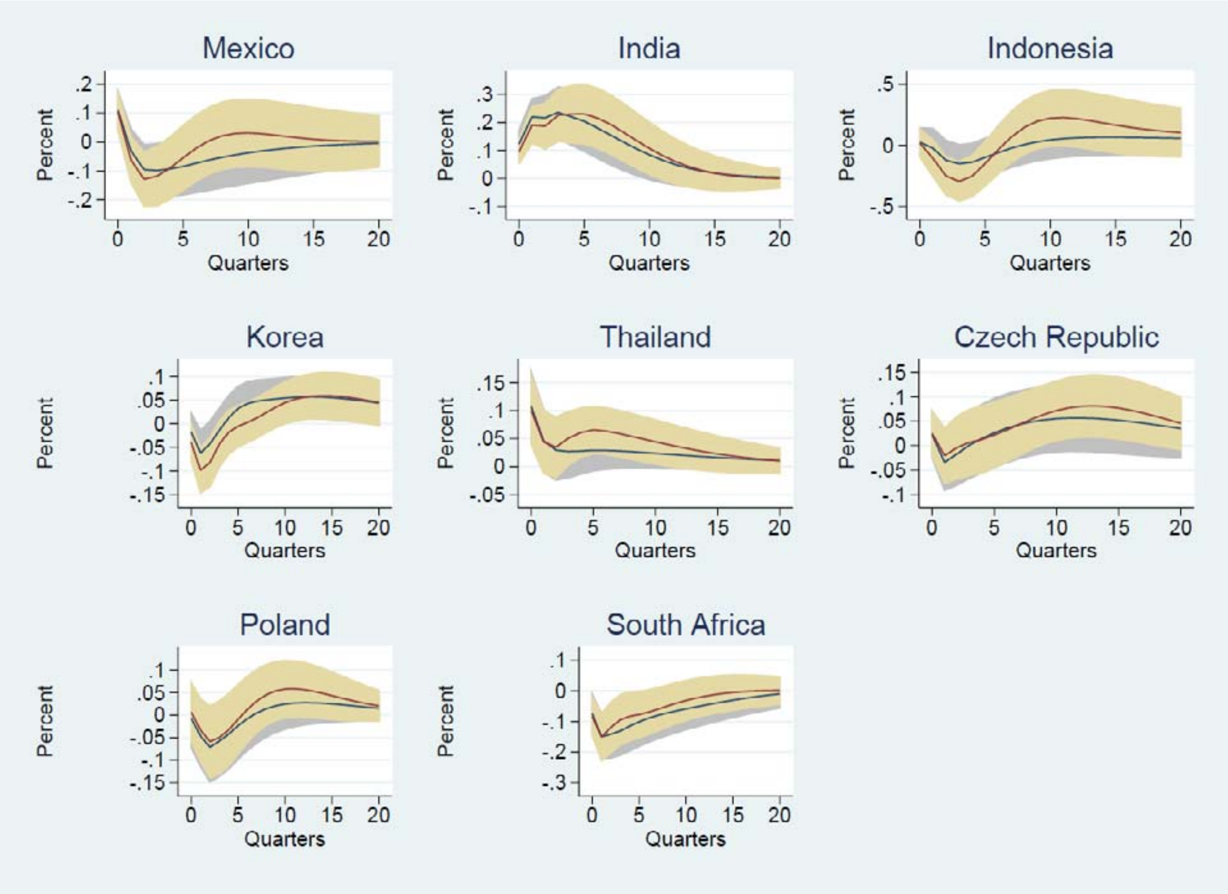
Second, in a some cases – including Colombia, Hong Kong SAR, India, Indonesia and Thailand – the IRFs suggest that the policy rate responses became more persistent in the second period. For these EMEs, the change in the policy rate channel seems to make a difference along all of the responses.

Third, in contrast, various EMEs appear to have maintained similar responses in each period.

In sum, these EME-dependent results are broadly in line with our test results.

Impulse response functions of each EME's long-term interest rate to a US federal funds rate shock[1]

Figure 10



[1] The impulse is a 25 basis point orthogonal shock to the US federal funds rate. The blue lines are obtained from model (1), which assumes no change in the long-term interest rate channel. The red lines are obtained from model (2), which allows for a change in the long-term interest rate channel and potentially in other channels; any such structural changes are assumed to have occurred in the Q3 2008–Q4 2013 period. Confidence intervals are set at the 70% level.

Moreover, some responses do not strictly conform with what one could expect under the assumption of uncovered interest rate parity. Specifically, the responses of Brazil and South Africa are negative and statistically significant.[22] In the periods considered, these economies have some of the largest average inflation rates within our sample. Moreover, both economies score low in the Chinn and Itto index of financial openness, a factor that could play a role in the data and thus in our estimations.

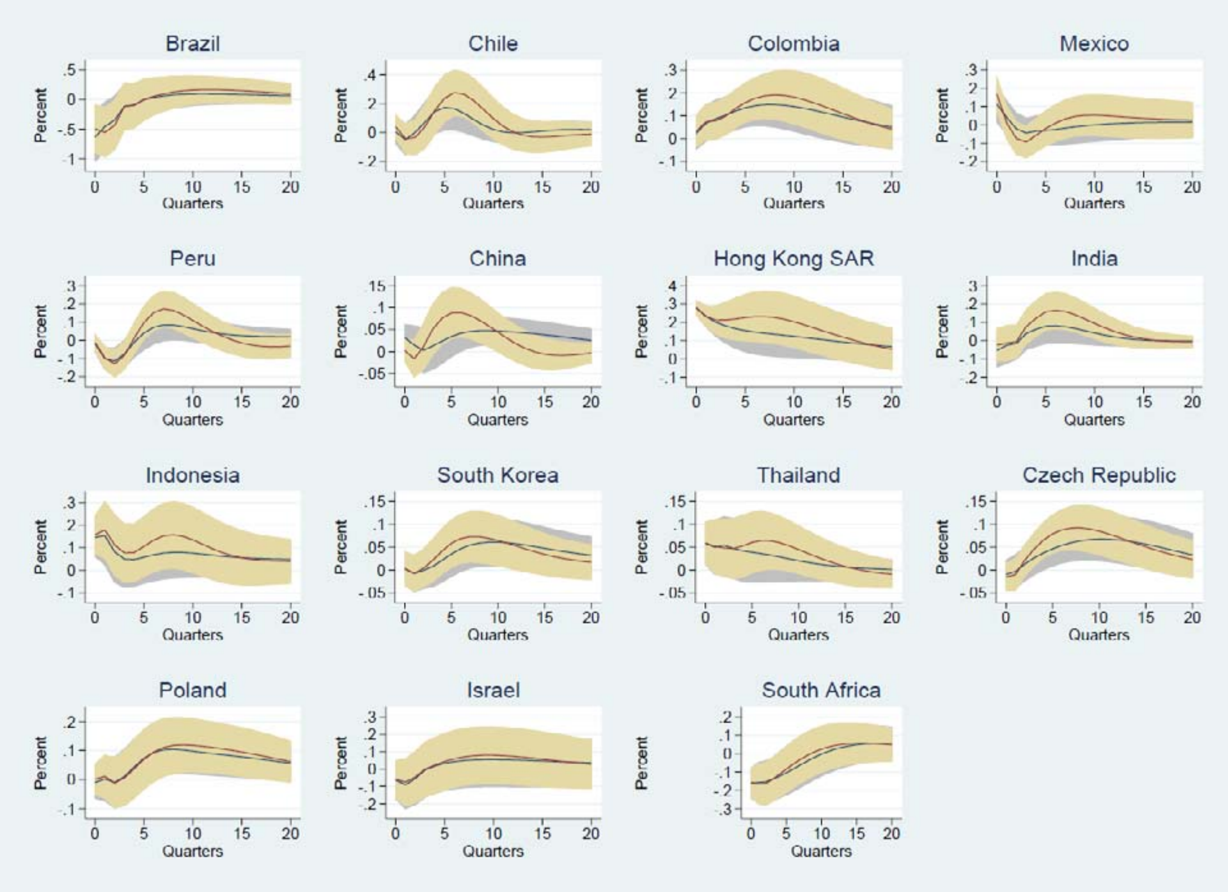
In the case of the exchange rate channel (Figure 9), the IRFs again arise from an orthogonal shock of 25 basis points to the US federal funds rate. The responses are

[22] In addition, their exchange rate responses are not statistically significant.

from model (1), which assumes no change in the channel, and from model (2), which allows for a change in the post-crisis period.

The results indicate, first, that only a very few economies may have faced a change in their exchange rate channel. Overall, the potential structural changes in the exchange rate channel do not seem very important. Moreover, many of the responses are not statistically significant in any of the periods.

Impulse response functions of each EME's policy rate to a Wu and Xia rate shock[1] Figure 11



[1] The impulse is a 25 basis point orthogonal shock to the Wu and Xia rate. The blue lines are obtained from model (1), which assumes no change in the policy rate channel. The red lines are obtained from model (2), which allows for a change in the policy rate channel and potentially in other channels; any such structural changes are assumed to have occurred in the Q3 2008–Q4 2013 period. Confidence intervals are set at the 70% level. See Wu and Xia (2013) for the Wu and Xia rate.

Second, the magnitude of the responses varies widely across economies. For example, the exchange rate of Hong Kong SAR essentially stays put, while South Africa responds with 100 basis points of currency depreciation on impact.

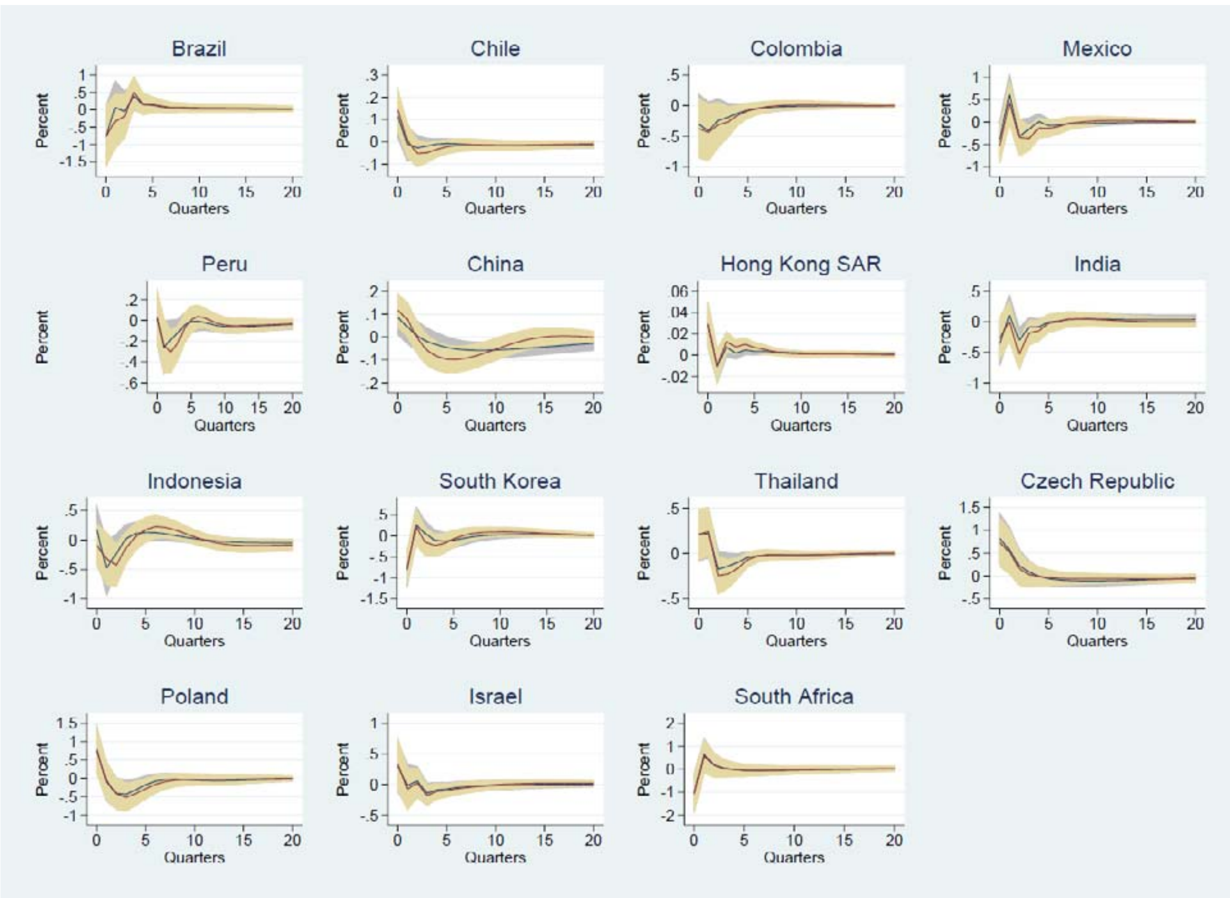
Third, only three EMEs undergo a statistically significant depreciation of the currency at some point. Moreover, in some economies, the possible structural change in the exchange rate channel has implications for their responses to the shock, but only after several quarters have elapsed.

In general, the prevalent exchange rate arrangement and monetary policy regime are relevant in explaining the effects we have measured with the IRFs. For example, Hong Kong SAR, which anchors its currency to the US dollar, experiences only a negligible variation in its exchange rate. In contrast, inflation targeters tend to have more ample and lagged variations in their exchange rate responses, as the exchange rates serve as buffers to external shocks.

For the IRFs involving the long-run rate, the evidence of a change in the related channel is not conclusive (Figure 10). In general, the paths are maintained.

Moving on to the IRFs that use the Wu and Xia rate in place of the US federal funds rate (Figure 11), we find that, first, the evidence of a possible structural change in the policy rate channel depends on the EME in question. For most cases, unconventional monetary policies seem to attenuate the differences in the responses before and after the crisis.

Impulse response functions of variations in each EME’s exchange rate to a Wu and Xia rate shock[1] Figure 12



[1] The impulse is a 25 basis point orthogonal shock to the Wu and Xia rate. The blue lines are obtained from model (1), which assumes no change in the exchange rate channel. The red lines are obtained from model (2), which allows for a change in the exchange rate channel and potentially in other channels; any such structural changes are assumed to have occurred in the Q3 2008–Q4 2013 period. Confidence intervals are set at the 70% level. See Wu and Xia (2013) for the Wu and Xia rate.

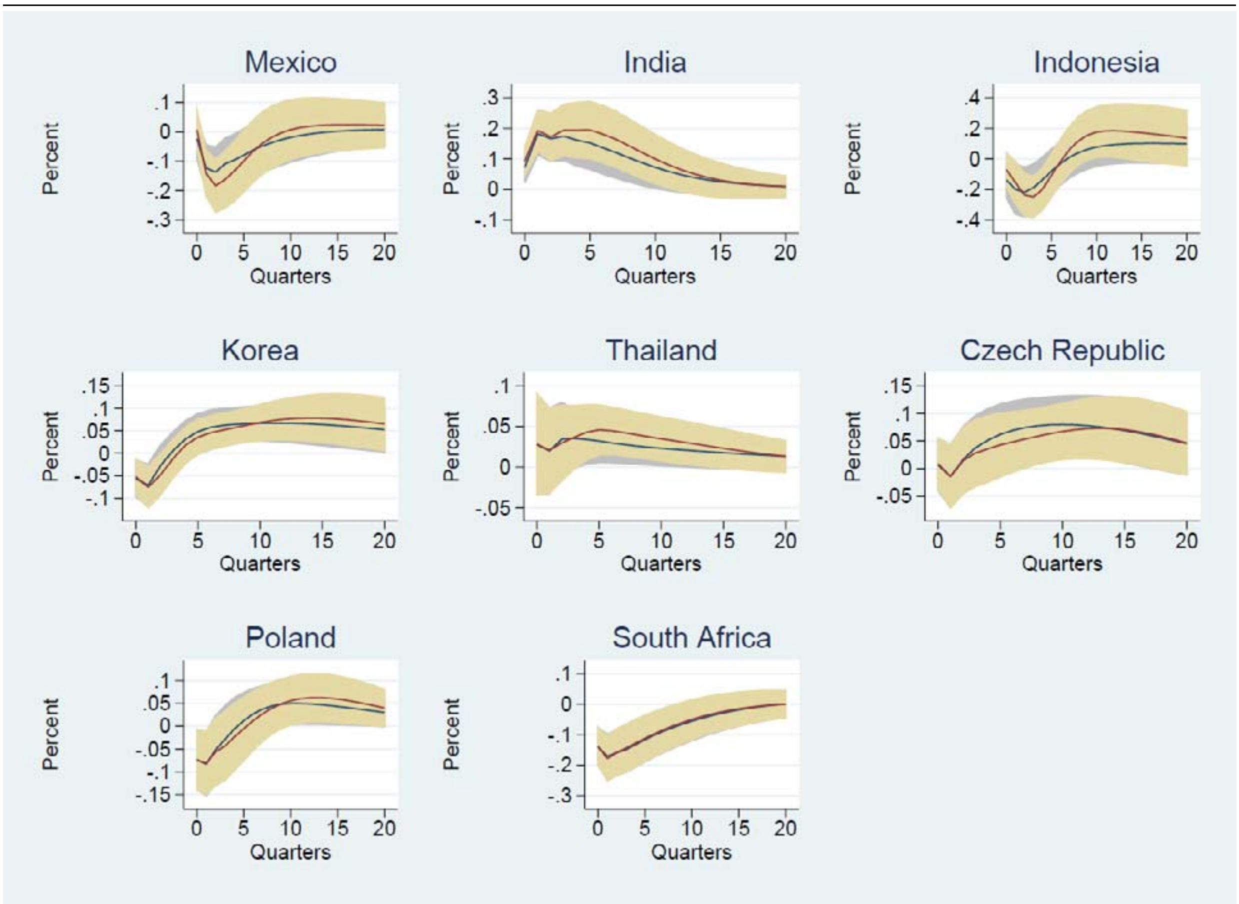
Second, except for a few cases, the exchange rate channel does not structurally change for most EMEs. The responses to shocks to the Wu and Xia rate are in general very similar, and in some cases the response’s paths are identical (Figure 12).

Third, in the case of the long-run rate channels, structural changes in the periods considered (Figure 13) are also country dependent. This is broadly in line with the results we have previously shown, in particular in Table 3. Two economies – Indonesia and, especially, South Africa – have, on impact, a stronger response; and surprisingly, the effects have the opposite sign from the one expected.

We set all IRFs' confidence intervals at a 70% level. Had we considered a higher level, confidence intervals would have been much wider, and evidence of no structural change in the transmission channels would have been stronger. This comment is applicable to all the IRFs we consider in this paper. Thus, the tests are more likely to reject the null hypothesis of no change, whereas the confidence intervals in the IRFs are more likely to be supportive of the null hypothesis. The IRF confidence intervals are not directly comparable to those of the tests because the former capture the uncertainty associated with the shocks to the economies, and the latter the uncertainty related to the estimated coefficients. Thus, in this vein, the tests on the linear regression and the FAVAR coefficients are more comparable to each other than either is to the IRFs.

Impulse response functions of each EME's long-term interest rate to a Wu and Xia rate shock[1]

Figure 13



[1] The impulse is a 25 basis point orthogonal shock to the Wu and Xia rate. The blue lines are obtained from model (1), which assumes no change in the long-term interest rate channel. The red lines are obtained from model (2), which allows for a change in the long-term interest rate channel and potentially in other channels; any such structural changes are assumed to have occurred in the Q3 2008–Q4 2013 period. Confidence intervals are set at the 70% level. See Wu and Xia (2013) for the Wu and Xia rate.

Moreover, it is relevant to assess the degree of consistency between the policy rate responses and the exchange rate responses and the extent to which it might have changed after the third quarter of 2008. As mentioned, it is useful to consider such consistency under uncovered interest rate parity (ie $i_t - ff_t = E_t(log(fx_{t+1} / fx_t)) = E_t(dfx_{t+1})$. We discuss this consistency in terms of the IRFs for the federal funds rate. Hence, if a positive shock to ff_t takes place, the most common response is for the policy rate, i_t, to increase and for the change in the exchange rate, dfx_{t+1}, to increase (ie depreciate) as well.[23] One can think of three other responses that, while plausible, are less likely.[24]

In this context, we have the following comments. First, consider EMEs with a floating exchange rate and an inflation targeting regime, eg Colombia and Mexico. In these economies, the policy rates respond with an increment, while the exchange rates depreciate, in accordance with uncovered interest rate parity.

Second, Indonesia is an example of how the policy rate's response might offset the US federal funds shock. Thus, the appreciation of the exchange rate might be initially seen as contrary to what is expected. Yet, the marked increment in the policy rate might be offsetting the US federal funds shock, thus making the exchange rate appreciate, as explained above. These two examples are all fairly consistent with uncovered interest rate parity.

Third, some of the results are quantitatively uneven. For instance, Korea's exchange rate appreciates on impact by about 80 basis points. This could have been reasonable provided that Korea's policy rate response had been on the order of 2 basis points – qualitatively plausible but hardly quantitatively so.

Nonetheless, it is counterintuitive if on impact the policy rate decreases *and* the exchange rate appreciates (assuming the expected exchange rate stays constant). We have three plausible explanations for this combination: (i) the model might be over-parameterised, thus not capturing the dynamics of the variables; (ii) the EME in question might have a low degree of financial openness, and its macroeconomic variables could be reflecting such restrictions; and (iii) if prices adjust only sluggishly, the EME's exchange rate might for some time appreciate even though US inflation might be increasing while the EME's inflation rate is decreasing.[25]

Final Remarks

All in all, we have found mixed results. In effect, whether an EME has undergone a structural change in the policy rate, exchange rate, or long-term rate channels depends on the EME in question. However, the evidence is not uniform across the

[23] An extension of the uncovered interest rate parity in terms of a risk premium is possible; ie $i_t - ff_t = E_t(dfx_{t+1}) + rp_t$. We leave such a possibility to future research.

[24] First, the policy rate, i_t, increases more than proportionally, and the exchange rate, dfx_{t+1}, appreciates. Second, the policy rate increases less than proportionally and the exchange rate depreciates. Third, the interest rate decreases and the exchange rate depreciates more than proportionally. The exact response depends on the prevalent exchange rate arrangement, the monetary policy regime, and other related policies.

[25] This last argument is akin to Dornbusch's overshooting model. A fourth option is to consider an extension of uncovered interest rate parity in terms of a risk premium, the one described in footnote 23.

various exercises and tests we have performed. Notably, some of the structural change we have documented might have taken place through second-round effects or, quite possibly, through other channels, which we believe our model is only implicitly capturing.

Although mixed, the results have some policy implications. First, regardless of the response channel, an increase in the sensitivity of EMEs to US monetary policy shocks could lead to higher dependence on US economic developments and accordingly to a higher impact of US policy on EMEs' policy cycles.

Second, in the same vein, EMEs might nowadays be facing more stringent policy trade-offs. This could partially explain what we have recently seen in some cases, which some perceived as authorities having decided to implement policy responses due to events in the US monetary policy stance. Understanding the degree to which such trade-offs might have changed is relevant for EME policymakers, who must remain aware of the policy trade-off magnitudes they are facing.

Third, since several EMEs may have gone through similar shifts in transmission channels, it is crucial to examine the extent to which some of the policies implemented by one EME can have an impact on other EMEs. For instance, the relative monetary stance and related macroeconomic policies in one EME can affect other EMEs, potentially deflecting capital flows; the possibility that such interaction could restrict an EME's policy even further is certainly latent.[26] This underscores the importance of having a sound policy framework in which government officials have readily available policy options with which to implement a cohesive and flexible policy response.

Furthermore, we believe there are at least three issues that require further scrutiny. First, as indicated above, some of our results suggest possible structural changes in channels not explicitly modelled. Disentangling such channels and learning their relative strength in the international transmission mechanism remains an important research endeavour.[27]

Second, what are the reasons behind the apparent structural change in some channels? Several are possible. For instance, are we seeing an upsurge in correlations between key international macroeconomic variables, which would increase the correlation among central bank responses? In contrast, have EMEs' central bank policy functions become more sensitive to changes in US monetary policy? Both effects probably played a part in such general changes, but it is important to understand their relative contribution in each EME.

Third, what are the economic implications of the structural change in a given transmission channel? If the importance of some channels has indeed grown or declined, how concerned should we be about the eventual tightening of US monetary policy beyond the withdrawal of unconventional policy measures? Once more, this brings home for EMEs the importance of having strong macroeconomic fundamentals, including a sensible policy framework.

[26] In addition, based on this point and as a more concrete exercise, it seems useful to explore an extension of our model in which data from two or perhaps more EMEs are used under the same FAVAR. This could, in principle, allow for the assessment of possible cross effects between EMEs.

[27] Relatedly, it would be relevant to further explore our models' specifications.

Appendix

IMF classification of de facto exchange rate regimes and monetary policy frameworks

The IMF (2013) classification system is based on IMF members' de facto exchange arrangements as identified by IMF staff, which may differ from the members' officially announced arrangements. The scheme ranks exchange arrangements on the basis of their degree of flexibility and the existence of formal or informal commitments to exchange rate paths. It distinguishes forms of exchange arrangements, in addition to arrangements with no separate legal tender, to help assess the implications of the choice of exchange arrangement for the degree of independence of monetary policy. The classification system presents members' exchange rate regimes against alternative monetary policy frameworks in order to highlight the role of the exchange rate in broad economic policy and to illustrate that different exchange arrangements can be consistent with similar monetary frameworks. The following sections explain the exchange arrangement categories.

Exchange rate anchor

The monetary authority stands ready to buy or sell foreign exchange at given quoted rates to maintain the exchange rate at its predetermined level or within a range (the exchange rate serves as the nominal anchor or intermediate target of monetary policy). These regimes cover those with no separate legal tender as well as currency board arrangements, fixed pegs with or without bands, and crawling pegs with or without bands.

Monetary aggregate target

The monetary authority uses its instruments to achieve a target growth rate for a monetary aggregate, such as reserve money, M1 or M2, and the targeted aggregate becomes the nominal anchor or intermediate target of monetary policy.

Inflation targeting framework

This involves the public announcement of medium-term numerical targets for inflation, with an institutional commitment by the monetary authority to achieve these targets. Additional key features include increased communication with the public and the markets about the plans and objectives of monetary policymakers and increased accountability of the central bank for its inflation objectives. Monetary policy decisions are guided by the deviation of forecasts of future inflation from the announced inflation target, with the inflation forecast acting (implicitly or explicitly) as the intermediate target of monetary policy.

Other

The country has no explicitly stated nominal anchor but rather monitors various indicators in conducting monetary policy. This category is also used when no relevant information on the country is available.

Descriptions of exchange rate markets based on HSBC (2011)

1. Brazil. Brazil has a free-floating regime implementing occasional interventions. The *National Monetary Council* sets the exchange rate regulations. The *Central Bank of Brazil* intervenes through the spot, swaps, and futures markets. In time of significant appreciation, it has also implemented capital controls.
2. Chile. The Chilean peso is a floating non-deliverable currency. The *Central Bank of Chile* will intervene occasionally in the exchange rate market. The exchange rate is determined in the interbank foreign exchange rate market.
3. China. The *People's Bank of China* keeps a managed float with reference to a basket of currency. The renminbi is non-deliverable and partially convertible (with respect to the capital account). On 19 June 2010, China's central bank announced a change in its exchange rate regime to increase its flexibility. The offshore renminbi market is developing quite swiftly.
4. Colombia. The *Bank of the Republic* maintains a flexible exchange rate regime with intervention rules to procure a certain level of international reserves, to limit excessive volatility, and to moderate excessive appreciation or depreciation of the nominal exchange rate. The mechanisms for intervening include discretionary purchases/sales of US dollars in the spot market.
5. Czech Republic. The *Czech National Bank* oversees a freely floating exchange rate. The Czech koruna is fully convertible. The Czech Republic joined the European Union in 2004, but there is no definite date for it to adopt the euro.
6. Hong Kong SAR. The *Hong Kong Monetary Authority* upholds a currency board system under which its monetary base is fully backed by foreign reserves. The Hong Kong dollar is a convertible and freely tradable currency.
7. India. The *Reserve Bank of India* oversees a managed floating regime. The Indian rupee is convertible on the current account but has some restrictions with respect to the capital account.
8. Indonesia. The *Bank of Indonesia* maintains a managed floating currency regime. The Indonesian rupiah is tradable but non-deliverable.
9. Israel. The *Bank of Israel* maintains a freely floating currency. It intervenes at times of disorderly market conditions.
10. Korea. The *Bank of Korea* preserves a floating exchange rate regime and might intervene under excess volatility. The won is fully convertible and tradable on a non-deliverable basis in the offshore market.
11. Mexico. The *Foreign Exchange Commission* formed by the central bank and the Ministry of Finance are responsible for exchange rate policy. The *Bank of Mexico* preserves a freely floating foreign exchange regime. It mostly intervenes using rules-based mechanisms to procure orderly conditions and liquidity in the exchange rate market.
12. Peru. The *Central Reserve Bank of Peru* intervenes in the foreign exchange market to prevent excess volatility in the exchange rate by buying and selling US dollars.
13. Poland. The *National Bank of Poland* maintains a freely floating currency. The zloty is freely convertible and is one of the most commonly traded currencies

against the euro. Poland joined the European Union in 2004, but there is no definite date for it to adopt the euro.

14. South Africa. The *South African Reserve Bank* oversees a managed floating exchange rate system. The rand is not yet fully convertible. The exchange rate controls have been gradually relaxed.

15. Thailand. The *Bank of Thailand* operates a managed floating currency regime and intervenes regularly to avoid volatility in the exchange rate market. The baht is to a large extent convertible and deliverable.

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