Journal of International Economics 108 (2017) S23–S41

Contents lists available at ScienceDirect

Journal of International Economics

journal homepage: www.elsevier.com/locate/jie

If the Fed sneezes, who catches a cold?

Luca Dedola[a,*], Giulia Rivolta[b], Livio Stracca[a]

[a]European Central Bank, Sonnemannstrasse 20, Frankfurt am Main 60314, Germany
[b]University of Brescia, Department of Economics and Management, Via San Faustino 74/b, Brescia 25122, Italy

ARTICLE INFO

Article history:
Received 26 August 2016
Received in revised form 22 December 2016
Accepted 1 January 2017
Available online 8 January 2017

JEL classification:
F3
F4

Keywords:
Identification of monetary shocks
International transmission
Capital mobility
Trilemma

ABSTRACT

This paper studies the international spillovers of US monetary policy shocks on a number of macroeconomic and financial variables in 36 advanced and emerging economies. In most countries, a surprise US monetary tightening leads to depreciation against the dollar; industrial production and real GDP fall, unemployment rises. Inflation declines especially in advanced economies. At the same time, there is significant heterogeneity across countries in the response of asset prices, and portfolio and banking cross-border flows. However, no clear-cut systematic relation emerges between country responses and likely relevant country characteristics, such as their income level, dollar exchange rate flexibility, financial openness, trade openness vs. the US, dollar exposure in foreign assets and liabilities, and incidence of commodity exports.

1. Introduction

This paper offers a re-examination of the international repercussions of U.S. monetary policy shocks. Does a monetary contraction in the U.S. lead to recessions or expansions in other countries? Does a monetary contraction improve or worsen financial conditions abroad? Does it lead to capital inflows or outflows? Are spillovers different across advanced and emerging economies, or across countries pegging their exchange rate to the dollar and those retaining monetary autonomy? These questions have long been studied and discussed, but empirical answers remain controversial, as recently argued by the former chairman of the Federal Reserve (Bernanke, 2015). A source of this lack of consensus is that most studies have tended to focus either on a limited set of countries (e.g., G7 countries, as in Kim, 2001) or on a limited set of variables (mainly output, inflation, short-term rates and bilateral dollar exchange rates as in e.g., Miniane and Rogers, 2007). In turn, the heterogeneity in the scope of these studies has made comparability of spillovers from their different estimates not very straightforward.

In this paper we contribute to this debate by documenting the effects of US monetary policy shocks on a broad set of macroeconomic and financial variables in 18 advanced and 18 emerging economies. We expand on previous work mainly in two dimensions. First, we identify US monetary policy shocks assuming that they have empirically plausible effects consistent with “textbook” monetary theory, also modelling their impact on a range of interest rates and asset prices. Second, and most importantly, in order to better understand the international transmission of monetary policy, we expand the set of the variables in countries other than the US included in our analysis. Going beyond measures of real activity and inflation, we also consider the responses of financial variables such as equity and housing prices, credit, and bank and portfolio flows. This allows us to better document any trade-off in terms of macroeconomic and financial stability for other countries brought about by a US monetary policy shock.

Our main findings are as follows. First, we find that a surprise US monetary tightening leads to a depreciation vis-à-vis the dollar in most countries in our sample, and drives them into recession. In a large majority of countries industrial production and real GDP fall, and unemployment rises; however, the trade balance improves. Inflation (both GDP deflator and CPI) also falls in a majority of countries, although these effects are less precisely estimated. Emerging economies tend to experience higher macroeconomic volatility. At the same time, and this is our second finding, the responses of financial variables are less clear cut and quite heterogeneous across countries. While many countries see their bond yields increase relative to the US, real equity and housing prices drop in just about half of the countries, mainly comprising emerging economies. Likewise, many countries experience opposite effects on domestic credit and capital flows, including borrowing from foreign banks. Finally,

* Corresponding author.

http://dx.doi.org/10.1016/j.jinteco.2017.01.002
0022-1996/

we do not find evidence of a systematic relation between likely relevant country characteristics (such as income level, exchange rate regime, financial openness, trade openness vs. the US, dollar exposure and incidence of commodity exports) and the distribution of cross-country responses to US monetary policy shocks.[1] For instance, across more and less financially open countries, asset prices and capital flows do not seem to react much differently.[2]

We carry out our analysis in two steps. First, we estimate US monetary policy shocks in a structural VAR identified with sign restrictions. These restrictions are consistent both with standard monetary theory, and with recent results in the empirical literature on the effects of monetary shocks. We then regress third country variables on estimated shocks. Hence, we are asking the following question: What are the consequences on the rest of the world of a US monetary policy shock, conditional on this shock having the assumed effects on the US economy?[3] Namely, we take for granted that monetary shocks have "textbook" effects on the US economy, such as that a tightening should reduce economic activity and inflation, while at the same time raising a range of interest rates.[4]

Specifically, in our first step we impose sign restrictions broadly consistent with the empirical findings in Gertler and Karadi (2015), which are representative of the literature. In addition to responses of output and inflation in line with previous evidence, these authors also estimate the effects of monetary policy shocks on several asset prices and interest rate spreads. This is an attractive feature for us, given our focus on the propagation of US monetary policy to international asset prices and interest rates. Moreover, their identification and results are robust to the presence of the lower bound on short-term interest rates in the aftermath of the Great Recession. By basing our restrictions on their estimates we can thus hope to make our results robust over the period that includes the global financial crisis. However, to further sharpen our identification, we also require that shocks also satisfy two further restrictions.[5] First, we impose that on impact the US effective nominal exchange rate appreciates following a US tightening. Second, that an aggregate of short-term rates in other major currencies react less than one-to-one to US rates. This ensures that we focus on those US monetary policy shocks which are not too positively correlated with any monetary policy shocks in other major countries. This concern is especially relevant in the aftermath of the recent financial crisis, when short-term rates in most advanced economies have been close to their lower bound, and more or less contemporaneously very expansionary conventional (and unconventional) monetary policies have been deployed. We find that under our identification assumptions, estimated impulse responses in the VAR are indeed robust to the inclusion of the 5 years from January 2009 to December 2013.

In our second step, armed with our estimated monetary policy shocks, we turn to the estimation of their effects on other countries. Similarly to other papers (e.g., Romer and Romer, 2004), we regress a host of variables for each country both at monthly and quarterly frequency on the estimated shocks. We then aggregate these estimates across countries on the basis of several structural characteristics. Namely, we compute median responses across countries in the same group. We group countries on the basis of the following characteristics: a) income levels — advanced and emerging economies; b) exchange rate regime — floaters and dollar pegs according to the de facto classification in Klein and Shambaugh (2010); c) financial openness according to the classification in Chinn and Ito (2006); d) US trade exposure and financial dollar exposure, the latter based on the currency composition of gross assets and liabilities in Benetrix et al. (2015); and e) incidence of commodity exports. Therefore, similar to Klein and Shambaugh (2010), we look at the role of receiving countries' structural characteristics and choice of policy regime in influencing the degree to which US monetary policy may impose (positive or negative) externalities abroad.[6]

Of course, our work is quite closely related to previous contributions in the literature on the global effects of U.S. monetary policy shocks (see Bernanke, 2015). A large body of this literature has shown that in the post-Bretton Woods period interest rates are more closely linked in countries that peg and in countries with open capital markets compared with countries that do not peg or impose capital restrictions.[7] Di Giovanni and Shambaugh (2008) look at the effect of foreign interest rates on domestic growth in a large group of countries, finding that the effect is stronger in countries with fixed exchange rate regimes, mainly on account of the stronger impact of foreign interest rates on domestic interest rates. Among VAR studies which try to control for systematic components in US interest rates, Canova (2005) and Mackowiak (2007) identify the effects of US monetary policy shocks on selected emerging economies. The former focuses on Latin American countries, finding that floaters and pegs display similar output but different inflation and interest rate responses. The latter finds that the impact on output and the price level in a few emerging economies are actually larger than in the US. Miniane and Rogers (2007), identifying US monetary shocks with contemporaneous exclusion restrictions, find no evidence that capital controls are effective in insulating other countries. Also in line with our results, they find that the exchange rate regime does not matter much for the macroeconomic transmission of US shocks, with countries having a fixed exchange rate regime being similarly affected as floaters in terms of output and inflation. Georgiadis (2016) shows, among other findings, that a floating exchange rate reduces the output spillover from US monetary policy shocks (the more so, the more trade and financially open the receiving countries). Most of these contributions do not consider, however, the potential financial dimension of spillovers, as we do in this paper. Similarly to us, Banerjee et al. (2016) document that a US contractionary monetary policy shock leads to a retrenchment in EME capital flows, a fall in EME GDP, and an exchange rate depreciation. In a theoretical model built to account for these findings, they show that macroeconomic spillovers may be exacerbated by financial frictions. Recently, Rey (2013) has shown that capital flows and stock prices in most countries, regardless of their exchange rate regime against the dollar, display strong comovements with the global cycle. The latter in turn is affected by US monetary policy. Miranda-Agrippino and Rey (2015) provide further evidence along the same lines, also using a large Bayesian VAR. Hence, monetary autonomy from the US is either not granted by a float or not sufficiently used. In this view,

[1] A caveat is that the spillovers from US monetary policy shocks are much less precisely estimated if we end our sample in the half of 2008.

[2] A reason why we do not find sharp differences across exchange rate regimes (beyond a more muted response of the bilateral dollar exchange rate in countries with lower exchange rate flexibility) could be that none of the countries in our sample has been all the time in a dollar peg.

[3] Thus a more precise title of the paper would be "If the Fed makes the US sneeze, who catches the cold?"

[4] See however Ramey (2016) for a critical appraisal of the literature on the domestic effects of US monetary policy shocks, challenging the robustness of the consensus view that we instead take as our starting point.

[5] This is a key reason why we do not use the shocks by Gertler and Karadi (2015) directly. See a thorough discussion in Section 2.2.

[6] We assign a country to a given group over the whole sample. However, to the extent that some country characteristics have not been very stable in our sample, this approach can bias our results toward finding less stark differences across country groups. Moreover, characteristics like the exchange rate regime or the degree of financial openness may be endogenously chosen to some extent as a function of the effects of US monetary policy shocks.

[7] See e.g., Klein and Shambaugh (2010). However, Rose (2011) finds that the macroeconomic and financial consequences of exchange rate regime choices are surprisingly inconsequential. Business cycles, capital flows, and other phenomena for peggers have been similar to those for inflation targeters during the global financial crisis and its aftermath.

the real choice confronting many countries is therefore a dilemma between monetary policy autonomy and unfettered capital mobility, rather than the classic Mundellian trilemma.[8]

The paper is organized as follows. We describe the empirical approach in Section 2, and present our data in Section 3. The US BVAR results are in Section 4; baseline results for all countries and for the subgroups are in Section 5. Section 6 concludes.

2. Empirical approach

We proceed in two steps. First, we estimate US monetary policy shocks using a large Bayesian VAR including several monthly US and global variables. We identify these shocks by imposing "textbook" sign restrictions on the effects of monetary policy shocks, drawing from the findings in the structural VAR literature. Second, following the literature (e.g., Romer and Romer, 2004), we obtain impulse responses by estimating simple autoregressive models for each variable in each country, including also contemporaneous and lagged values of the shocks. We then aggregate the resulting impulse responses across countries according to their own characteristics. A way to view our approach is the following. Conditional on assuming that there are US monetary policy shocks that have empirically plausible domestic effects consistent with standard theory, our aim is to investigate the consequences of these shocks for the rest of the world. Thus, we take for granted that these shocks have domestic effects on the US economy, such as that an interest rate hike (cut) should reduce (boost) economic activity and asset prices, and (at some point) also inflation. We rely on representative results in the literature to spell these effects out in an empirically plausible way with our priors, in particular Gertler and Karadi (2015) — henceforth GK, so that we can obtain estimates of the underlying monetary policy shocks.

2.1. The BVAR model

The empirical model used to estimate US monetary policy shocks is a Large BVAR. This tool has been introduced by Bańbura et al. (2010) to avoid the issue of over-fitting when dealing with systems of many variables, building on the seminal contributions by Litterman (1986) and Sims and Zha (1998). The rationale behind this approach is that by using informative priors it is possible to shrink the likely over-parametrized VAR toward a more parsimonious model. Therefore, the choice of the informativeness of the priors is crucial. In this work we follow the approach of Giannone et al. (2015); i.e., the appropriate degree of shrinkage is selected by treating hyper-parameters as any other unknown parameter, formulating a prior over them and using the data to determine their posterior values.

Specifically, the reduced form VAR model for n variables,

$$Y_t = BY_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \Sigma)$$

is conceived as a hierarchical model, where hyper-parameters are assigned diffuse hyper-priors so that maximizing their posterior simply amounts to maximizing the marginal likelihood with respect to them. As regards priors, a Normal — Inverse-Wishart distribution is used for the coefficients and the variance–covariance matrix, namely

$$\Sigma \sim IW(\psi I_n; n+2)$$
$$vec(B) \mid \Sigma \sim N(b, \Sigma \otimes \Omega)$$

where $b(\gamma)$ and $\Omega(\gamma)$ are functions of a small vector of hyper-parameters γ. The scale parameter ψ is also a hyper-parameter, which follows a diffuse prior, $\psi \sim IG(0.02^2, 0.02^2)$. Bayesian shrinkage is achieved through the combination of Minnesota, sum-of-coefficients and dummy-initial-observation priors for the VAR coefficients. The Minnesota prior assumes that the limiting form of each VAR equation is a random walk with drift. The sum-of-coefficients prior and the dummy-initial-observation prior are necessary to account for unit root and cointegration. Because the posterior does not admit analytical characterization, even under Gaussianity of the likelihood function, an MCMC algorithm is used for inference, based on a Metropolis step to draw the vector of hyper-parameters and on a standard Gibbs sampler to draw the model's parameters conditional on the former. From the conditional posterior distribution we extract 20,000 draws, of which the first 10,000 are discarded and the last 10,000 are used for inference on monetary policy shocks. Further details on the prior specification and estimation procedure can be found in Giannone et al. (2015).

This framework allows to estimate the VAR in levels, with variables expressed in annualized terms. Specifically, our model consists of $n = 13$ monthly variables, both US-specific and international variables. We wish to include many US and global variables for two reasons. First, we want to identify the monetary policy shocks by imposing sign restrictions in the spirit of the findings in the structural VAR literature for as many variables as possible. This implies that in addition to output and inflation, we need to include several interest rates and spreads in our VAR for which authors like GK find an effect of monetary policy. Second, given the open-economy focus of our study, as well as including the US nominal effective exchange rate, we also need to control for global drivers of economic and financial fluctuations. The US economy variables comprise industrial production (IP), the CPI, the Federal Funds rate, the 1-year government bond yield, the S&P500 index, the dollar nominal effective exchange rate against 20 trading partners, the corporate bond spread, the mortgage spread and the commercial paper spread. The last three variables are the same as in GK. The global variables consist of the CRB commodity price index, world industrial production (excluding construction) as calculated by the OECD, (ex-US) world stock prices, and the difference between a "G7" ex-US short-term interest rate and the US 3-month T-bill rate. The former rate is computed as an average of the short-term rates of the four major currency areas (Canada, Euro Area, Japan, UK).[9] As variables are monthly and enter the VAR in levels, the model is estimated with $p = 13$ lags.

2.2. Identification

We find it convenient to impose priors to identify US monetary policy shocks through sign restrictions on the impulse response functions, following the well-known methods pioneered by Faust (1998), Canova and De Nicoló (2002), and Uhlig (2005). A notable difference from these early contributions is that our goal is not to obtain new independent evidence on the effects of monetary policy shocks through minimal identifying assumptions. Rather, we use sign restrictions to impose plausible, "textbook" assumptions on the overall domestic effects of monetary policy shocks. Conditional

[8] Ostry and Ghosh (2016) point out that there may be a need for policy coordination if US monetary policy creates trade-offs for the receiving countries that they cannot (costlessly) undo with their own macroeconomic policy. Nevertheless, Woodford (2007) shows that globalisation does not, in general, imply a loss of monetary control in a model with frictionless international asset markets. Fahri and Werning (2014) provide an elegant model of such a dilemma.

[9] The 3-month T-bill rate is used for UK, the call money rate for Japan, the 3-month Euribor for the euro area and the general T-bill rate for Canada.

on the assumed domestic effects, we then investigate the cross-border repercussions of US monetary policy shocks. Operationally, we impose restrictions consistent with the effects of US monetary policy estimated by GK, which are illustrative of the findings in the VAR literature on monetary non-neutralities. These authors use external instruments, based on high-frequency financial data (see also e.g., Gürkaynak et al., 2005), to identify monetary policy shocks, including the period over which US short-term interest rates have been at their lower bound. While their findings in terms of responses of macroeconomic variables such as industrial production and inflation are well in line with those from other VAR studies and standard monetary theory (see e.g., Galí, J, 2015 textbook), a distinct feature of their results is that they also estimate the responses of a broad range of government and private bond interest rates.

Given our interest in the financial transmission through asset prices, we rely on their findings to model the contemporaneous responses of these variables. Another advantage of drawing on the GK estimates is that they identify monetary policy shocks whose effects are reasonably robust to the presence of the lower bound on short-term interest rates. Thus, we can also hope that our identification is equally robust to the inclusion of the period that encompasses the recent financial crisis. While we will look at results both including or excluding this most recent period after 2008, the latter could be important to identify the transmission of US monetary policy shocks. On the one hand, to the extent that the systematic reaction of monetary policy has been constrained by the lower bound on short-term rates, this has effectively resulted in a series of contractionary monetary shocks. This intuition is borne out by standard New Keynesian models in which systematic monetary policy follows a rule for the short-term interest rate and is constrained by the lower bound.[10] On the other hand, when the lower bound binds, the current level of the short-term rate may not be a good gauge of the stance of monetary policy by itself, if the central bank is able to credibly rely on forward guidance and thus still affect longer-dated interest rates. Neglecting this aspect may then result in an overestimation of the size of contractionary shocks over this period. Our identification in this respect possesses a key safeguard as we require that a contractionary shock not only increases the short-term rate (relative to its normal level in line with macroeconomic conditions), but that also the 1-year rate and a series of interest rate spreads go up.[11] Therefore, any lack of accommodation in short-term rates over the more recent period will be interpreted as a contractionary shock only if associated with increases in all these other longer-dated interest rates (and as we discuss below also with both dollar appreciation and an increase in the US interest differential with other major currencies).

In principle, we could have used the same external instruments as in GK to identify US monetary policy shocks with our reduced form VAR residuals. We prefer to pursue a different approach for several reasons.[12] First, our approach does not require literally imposing the same identifying assumptions as in GK. Indeed, it is well-known that sign restrictions yield a plurality of orthogonalizations of reduced form VAR residuals. In this respect, our assumptions are less restrictive, but still compatible with the effects of monetary policy in a wide range of both empirical and theoretical models, including estimated and calibrated DSGE models. Second, we obtain a longer series of monetary policy shocks as we impose our restrictions on the whole sample starting in 1980, rather than the shorter one for which their external instruments are available. Finally, we also want to focus on US monetary policy shocks which should be at most weakly correlated with monetary policy shocks in other major countries. This is especially a concern in the aftermath of the recent financial crisis, when short-term interest rates in most advanced economies have been at their lower bound, as more or less contemporaneously very expansionary conventional (and unconventional) monetary policies have been deployed. The inclusion of the interest rate differential with ex-US G7 is also likely to make our results more robust to the risk of giving too much weight to contractionary shocks during the more recent period. This is similar to the argument above regarding the inclusion of other longer-dated interest rates and spreads. Any deviation of the US short-term rate over this period from its estimated systematic relation with the underlying state of the economy is going to be mapped into a discretionary lack of accommodation and thus a contractionary monetary policy shock only if associated with a higher interest rate than in the other major economies. We thus recover shocks that also satisfy, at least on impact, the following requirements. First, a measure of short-term rates in other major currencies should react less than one-to-one to US rates; second, the US nominal effective exchange rate appreciates.

In more detail, we impose the following restrictions:

$$\begin{aligned}
&FFR > 0 \quad \text{for} \quad t = 1, \ldots, 6\\
&IP_{US} < 0 \quad \text{for} \quad t = 2, \ldots, 6\\
&CPI_{US} \leq 0 \quad \text{for} \quad t = 4\\
&1Y - GBY_{US} > 0 \quad \text{for} \quad t = 1, \ldots, 4\\
&MS_{US} > 0 \quad \text{for} \quad t = 2\\
&CPS_{US} > 0 \quad \text{for} \quad t = 1, 2, 3\\
&SP_{US} < 0 \quad \text{for} \quad t = 1\\
&NEER_{US} > 0 \quad \text{for} \quad t = 1\\
&DiffIR < 0 \quad \text{for} \quad t = 1
\end{aligned}$$

Here FFR is the Fed Funds rate, IP_{US} is the US industrial production, CPI_{US} is the US consumer price index, $1Y - GBY_{US}$ are 1-year government bond yields, MS_{US} is the mortgage spread, CPS_{US} is the commercial paper spread, SP_{US} is the S&P500 index, $NEER_{US}$ is the nominal effective exchange rate, and $DiffIR$ is the difference between the G7 interest rate and the US short-term rate. The first six restrictions are consistent with the results in GK as reported in their Figs. 2–8. Moreover, a persistent contraction in industrial production is a fairly widespread finding in the literature on the effects of US monetary policy shocks. Similarly, we impose that inflation be negative after four months, striking a compromise between studies imposing a fall on impact (e.g., Uhlig, 2005) and the evidence of a delayed response.[13]

We also impose that US stock prices fall on impact, and that the nominal effective exchange rate appreciates, while $DiffIR < 0$. As discussed above, the last two restrictions help in ensuring the identification of a US-specific monetary policy shock. The fall in the

[10] Of course, the risk here is that the effects of these contractionary shocks are also commingled with those of other underlying shocks. However, as we show below, our results are reasonably similar for the two samples including or excluding the period 2009–2013.

[11] In this respect, we are focusing on what Gürkaynak et al. (2005) dub a "path" shock to interest rates.

[12] Indeed, we could have used GK instruments directly in IV estimates of regressions of third-countries' variables on US interest rates. However, the results in Ramey (2016) are a source of concern in this respect. This author shows that the GK instruments and shocks may be rather weak and lead to inconclusive results in a single equation setting like the one we use below, even when applied to US data, especially concerning the estimated effects on output and inflation. In practice, we do find that the GK shocks result in an increase in US industrial production in a regression like Eq. (2), in sharp contrast to the VAR impulse responses in their paper. This is a further reason to seek alternative and potentially sharper instruments with our approach.

[13] We did not conduct a search specification on the number and timing of periods for which we impose our restrictions. As it is clear from the estimated impulse responses in Fig. 1, we obtain effects that are in general more persistent than assumed. For instance, the CPI falls already on impact even though it is restricted to do so only after 4 months. Therefore, marginal changes to these restrictions are likely to have little effects on our baseline results.

interest differential does not require interest rates in other major currencies to fall, but only that they increase by less than their US counterparts on impact. Observe that these assumptions are conservative for our purposes, as we are constraining interest rates in major currencies to increase by less than US rates and thus to be more accommodative, other things equal. This can then result in an attenuation of the effects of US monetary policy on the rest of the world.

Finally, the impulse response functions of the remaining four variables we include are left unrestricted. Namely, the US corporate bond spread, commodity prices, world industrial production, and world stock prices are free to react to the shock according to the data. These last three variables then will provide initial unrestricted evidence of the aggregate effects of US monetary policy shocks on the rest of the world.

The algorithm to estimate the posterior distribution of impulse response functions and of monetary policy shocks is standard (see Uhlig, 2005). As discussed above, we obtain 10,000 draws from the conditional posterior distributions of the reduced-form coefficients and variance–covariance matrix. Recall that any candidate contemporaneous response to the vector of structural shocks can be calculated as

$$H = PQ,$$

where P is the Choleski factor of the variance matrix of the reduced form innovations, $\Sigma = PP'$, and Q an orthogonal matrix obtained from the following decomposition

$$X = QR,$$

where X is the realization of a matrix of independent $N(0,1)$. Thus, for each of our 10,000 reduced-form draws, 5000 random orthogonalizations Q of the variance–covariance matrix are evaluated, discarding those that do not satisfy the sign restrictions. This amounts to assuming a uniform prior over the set of all possible Q matrices, multiplied by an indicator function taking a value one for those Qs that satisfy the restrictions. These Qs are then used to compute candidate monetary shocks from the associated reduced form residuals; namely we have that

$$\varepsilon_{US,t}^{MP} = (PQ)^{-1}\varepsilon_t = (PQ)^{-1}\left(Y_t - BY_{t-1}\right).$$

The algorithm always finds at least one suitable orthogonalization for more than 99% of the draws from the conditional posterior distributions. Therefore, our restrictions do not implausibly constrain the reduced form BVAR posterior.

A subtle issue with this approach is that since it admits many equally plausible structural models (Q matrices), it recovers many equally plausible time series of monetary policy shocks ($\varepsilon_{US,t}^{MP}$). These series can be quite different one from the other. On the one hand, only one of these models is the "true" one and all the others are "wrong". This would imply that the shocks obtained from all the other "wrong" structural models are linear combinations of the "true" monetary policy shocks and other shocks. From this perspective, this is a limitation of our approach.[14] On the other hand, our methodology implies that we do not need to commit to one single structural model, that we would regard as (un)likely as all the others which satisfy our sign restrictions. In other words, all these models are observationally equivalent, so that we would not learn which one is "true" even with an infinite amount of data. In this vein, the monetary shocks we obtain are all plausible candidate shocks, as are all the associated Q matrices that satisfy sign restrictions, and we take fully into account the uncertainty over all of these shocks, in the same fashion as we do when we estimate impulse responses (as implied by the uncertainty over the Q matrices). The hope is that these potentially different shocks nevertheless have similar enough effects to provide informative evidence. Moreover, any serious contamination by other shocks should be quite apparent already in the BVAR results. As we show below (see Section 4), these results are quite conventional and similar to those obtained in the literature.

We conclude this subsection by briefly discussing some aspects of the priors on impulse responses elicited by this procedure. Recently, Baumeister and Hamilton (2015) have argued that sign restrictions can unduly constrain impulse response posteriors. These authors show that the effects of this kind of priors do not vanish asymptotically, a property shared by all priors on impulse responses in under-identified models (i.e., which admit more than one matrix Q). Moreover, sign restrictions imply priors on impulse responses that tend to put sizable probability mass on values close to zero for restricted variables, and literally at zero for unrestricted variables. At the same time, possibly large but implausible responses also receive non-zero probability. Therefore, it is important to have some sense of how much posteriors differ from implicit priors. In the web appendix we report the posterior distributions of the impact responses of restricted and unrestricted variables. We find that these posteriors tend to have densities with little mass close to zero or on extreme values, and thus are in most cases reasonably different from priors.

2.3. Estimation of the impact on countries other than the US

The above procedure, in addition to yield the posterior distribution for the impulse response functions from the BVAR, allows us to obtain the posterior distribution of US monetary policy shocks. Both equally result from the uncertainty over reduced form parameters (Σ, B), as summarized by their respective posterior distributions, and over matrices Q, as summarized by our prior. Armed with these shocks, for each variable y in country i, y_i, we compute a vector of impulse responses at horizon h

$$IRF_{y,i,h} = \frac{\partial y_{i,t+h}}{\partial \varepsilon_{US,t}^{MP}}, \tag{1}$$

for all the countries in our sample other than the US. Under the assumption that the monetary policy shocks are exogenous, we can arbitrarily approximate the true impulse responses by regressing each variable y_{it} on an infinite series of $\varepsilon_{US,t-j}^{MP}, j = 0, \ldots, \infty$. Following the literature (e.g., Romer and Romer, 2004), given the finite sample constraint, we obtain the impulse response coefficients by estimating, for a given series of shocks $\varepsilon_{US,t}^{MP}$, the following distributed lag model for each variable:

$$y_{it} = \alpha_{i,j} + \phi_i(L)y_{i,t-1} + \beta_i(L)\,\varepsilon_{US,t}^{MP} + \varepsilon_{it}, \tag{2}$$

where we also include monthly or quarterly dummies and a time trend. Variables are transformed as in the BVAR. A unitary shock, $\varepsilon_{US,t}^{MP} = 1$, thus amounts to a one-standard deviation structural shock, as in the BVAR impulses responses.

We characterize uncertainty of estimates of Eq. (2) by computing their distributions over the realizations of our estimated shocks $\varepsilon_{US,t}^{MP}$, to take into account that the latter are generated regressors. In particular, we assume that conditional on both lags of $y_{i,t}$ and a given realization of the series of monetary policy shocks $\varepsilon_{US,t}^{MP}$, the error term in Eq. (2) is Gaussian $N(0, \sigma^2)$. Together with a conjugate

[14] However, we can always find ways to select just one model that we would consider most plausible. For instance we could pick the matrix Q that in addition to satisfying our restrictions also maximizes the correlation of our shocks with other monetary policy shocks, such as the GK shocks or the Romer and Romer (2004) shocks, or both.

(Normal-IG) prior on the vector of coefficients $\Gamma = (\alpha_{i,j}, \phi_i, \beta_i)$ and on σ^2, this implies that the posterior for these coefficients is also a standard Normal-IG. Therefore, we can easily draw from it to simulate the posterior of the impulse responses, conditional on the given series of shocks. Repeating this procedure for a number of draws of time series of the monetary shocks allows to simulate the posterior distribution of the impulse responses taking into account also overall uncertainty about the estimation of shocks.

In practice, we proceed as follows. We extract 10,000 time series of the US monetary policy shocks following the procedure described in the previous sub-section, and for each of them we also extract 10 draws from the conditional (Normal-IG) posterior of the regression parameters Γ, σ^2.[15] Given the combined parameter and shock uncertainty, relying on an uninformative prior over Γ, σ^2 would result in very imprecise estimates. Thus, we pick the prior hyper-parameters in the following way. First, similarly to the BVAR, we set $\sigma^2 \sim IG(\nu = 3, 0.02^2)$. This implies that the variance of the Normal prior on the coefficients of Eq. (2) is then equal to $\sigma^2 I$. Second, we set the mean of the Normal prior over Γ equal to the OLS estimates which are obtained in the regression using the time series of the cross-sectional median values of our estimated monetary policy shocks.[16] We document in the online appendix the consequences of using this prior for monthly US variables (such as industrial production, CPI, equity prices, the nominal and real effective exchange rate and 3-month and 12-month interest rates), for which we have also evidence from the BVAR. First, the posterior distribution of impulse responses we obtain is sufficiently different from the prior; therefore the latter, though informative, does not unduly affect the former. Second, it is also interesting that the posterior distributions of US variables computed with the single-equation procedure based on Eq. (2) are similar to those of the impulse responses of the same variables obtained from the BVAR (reported in Fig. 1). While this property may not be so relevant for countries other than the US, at least it ensures some degree of consistency in our approach. We conclude from this exercise that this prior choice strikes a reasonable balance between making the US estimates based on Eq. (2) close to their BVAR counterparts, a minimal consistency requirement, and imposing too tight a constraint on the posterior inference.[17]

The flexibility of this approach represents a key advantage given our quite heterogeneous panel of data. It allows us to consider variables at both monthly and quarterly frequency for each country i, as discussed in the next section, also using samples shorter than those for which we estimate our shocks. This heterogeneity in frequencies and sample length prevents us to estimate directly impulse responses by including other countries' variables in our BVAR. Nevertheless, relative to a VAR, the single equation approach does not allow us to fully take into account the possible influence of third variables, including the presence of common stochastic trends. But in this respect, we need to strike a balance between data constraints and an optimal model specification. Moreover, in addition to yielding results country by country for each variable, it makes it convenient also to aggregate them across countries on the basis of several characteristics. These aggregations are obtained by taking the point-by-point median response for each variables across countries belonging to a given group. The choice of the median is robust to the presence of outliers which could unduly influence the aggregation otherwise. Note that we do not pool the data, due to significant heterogeneity in country-specific results, which could give rise to an aggregation bias (see Pesaran and Smith, 1995).[18]

We group countries on the basis of the following characteristics: a) income levels — advanced and emerging economies; b) exchange rate regime; c) financial openness; d) US trade and dollar financial exposure; and e) the incidence of commodity exports; the details of these characteristics are described in the next section.

3. Data description

The tables in the web appendix describe in detail all variables used in the empirical analysis and their sources. The Bayesian VAR model to identify US monetary policy shocks consists of 13 monthly variables which were discussed above (see also Table B1 in the web appendix).

In order to study the international effects of US monetary policy, a large number of country-specific variables are regressed on the estimated monetary policy shocks and the impulse response functions are computed. Our sample consist of 36 countries, namely: Australia, Austria, Belgium, Brazil, Canada, Chile, China, Colombia, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, India, Italy, Japan, Korea, Latvia, Lithuania, Malaysia, Mexico, Netherlands, Norway, Philippines, Poland, Portugal, Russia, South Africa, Spain, Sweden, Thailand, Turkey and UK. We consider euro area countries individually for all variables but short-term rates and bilateral US dollar exchange rates. These series refer only to euro area aggregates after 1999 (or the date of euro adoption).

For each country we consider both monthly and quarterly variables. *Monthly variables* include the following: (i) the bilateral dollar exchange rate;[19] (ii) the real effective exchange rate; (iii) the short-term interest rate differential with the US; (iv) CPI; (v) industrial production; (vi) real stock prices (deflated with the CPI); the nominal trade balance (scaled by the average of the sum of import and export over the whole sample); and (viii) the differential of long-term government bond yields vis-á-vis the US.

Quarterly variables include the following: (i) real GDP; (ii) the GDP deflator; (iii) the unemployment rate; (iv) real housing prices (deflated by CPI); (v) real domestic credit (deflated by CPI); (vi)–(vii) total portfolio inflows and outflows, and (viii) total bank inflows, all scaled by GDP. Finally, as a gauge of macroeconomic volatility we also report results for the sum of the absolute changes in unemployment and inflation (as measured by the GDP deflator), the so-called "misery index".[20] Details about the source of each series are provided in the tables in the online appendix.

The series of monetary policy shocks extracted from the BVAR starts in February 1981 (as we use 13 lags in the model) so that the regressions can be estimated from that date on. When coming to quarterly regressions the monetary policy shocks are aggregated taking their quarterly average. These regressions are estimated starting from Q2 1981. As some variables are not available over the whole sample, we are forced to run some regressions over shorter samples, as documented in the online appendix.

[15] Using 100 draws instead of 10 does not materially alter our results but greatly increases the computational time.

[16] Of course this time series of shocks does not correspond to any of the actual series we estimate with our procedure; it just represents a convenient way to initialize our priors.

[17] In a small sample like ours, there is no guarantee that impulse responses for the same variables obtained from the BVAR and from Eq. (2) would be similar — see e.g., Kilian and Kim (2011). Indeed, results here depend on the prior over the coefficients of Eq. (2). When we experimented with less informative priors, either by increasing the variance of the Normal prior for the coefficients or setting the mean of the latter to zero, we found that the implied posteriors were now fairly different from those estimated in the BVAR.

[18] A further reason preventing us to use panel techniques relates to computational difficulties inherent in our Bayesian approach to deal with the uncertainty in shock estimates. Bayesian panel data analysis requires at least the use of Gibbs sampling (if not full MCMC methods) to simulate posterior distributions conditional on a given monetary shock time series. But this is hardly feasible given the large number of draws we need to simulate from the empirical distribution of our shocks.

[19] It is defined as the amount of local currency needed for 1$ so that an increase in the exchange rate represents an appreciation of the US dollar.

[20] The misery index as the sum of the inflation and unemployment rates is a simple measure of economic welfare introduced by Arthur Okun in the 1970s. Of course, it is only a crude measure, and has been criticised in later contributions. Therefore, we interpret it only as a synthetic indicator of macroeconomic volatility.

3.1. Country characteristics

The second step of our analysis consists of aggregating the impulse response functions of single-country variables according to several country-specific characteristics. One important qualification to keep in mind is that some of the country characteristics, notably the exchange rate regime, may be selected endogenously. In particular, countries with stronger links with the US and/or small open economies may be more likely to peg to the dollar. Therefore, a possible nexus between the exchange rate regime and the strength of monetary policy spillovers from the US may reflect not only the regime as such, but possibly also the country characteristics that predict the exchange rate regime.

We group countries on the basis of the following characteristics: a) income levels — advanced and emerging economies; b) dollar exchange rate regime; c) financial openness; d) US trade and dollar financial exposure; and e) the incidence of commodity exports. Here we describe the indicators used to derive country groups and their sources.

Advanced vs. emerging economy
The classification into advanced or emerging economy is consistent with the one in the IMF World Economic Outlook. In this case we refer to the latest classification rather than the average over the sample.

Exchange rate regime
The classification of the exchange rate regime is not a straightforward one since there is more than one meaningful classification (see Rose, 2011). We mainly draw from Klein and Shambaugh (2010), who also have some information on the base country. Hence we rank countries according to their exchange rate flexibility with the USD following the index in Klein and Shambaugh (2010).

Financial openness
We measure financial openness with the Chinn–Ito index, which is an indicator of de iure financial openness.

US trade exposure
We consider countries' trade linkages with the United States (exports to and imports from the US as a share of domestic GDP).

Dollar exposure
This is computed on the basis of data in Benetrix et al. (2015) on the currency composition of gross foreign assets and liabilities. Here we focus on gross rather than net exposure, although the choice is not uncontroversial.

Commodity exporters
We define commodity exporters based on the incidence of net exports of primary goods over total exports plus imports. Primary goods include fuels (oil, gas, coal), metals, food and other raw materials.

Table 1 displays the average sample values of the indicators used to derive country groups (with the exclusion of the income level and the share of commodity exports). Unless differently specified (namely in the case of advanced vs emerging countries and commodity exporters), countries are split in two different groups depending on whether the value of their indicators fall above or below the median value over the whole sample for which these characteristics are computed. Table 2 reports the list of countries comprising the respective groups. These groupings are then combined to derive subgroups of countries with several interesting common characteristics, so that we also consider emerging floaters or peggers, emerging financially open and emerging less-financially open countries, and so on.

The use of average characteristics is in line with the approach in e.g., Miniane and Rogers (2007), in which point impulse response estimates are directly regressed on average characteristics over the sample such as the intensity of capital controls. However, to the extent that countries' characteristics have not been very stable in our sample, this approach can bias our results toward finding less stark differences across country groupings. Unfortunately, for many countries we simply don't have the degrees of freedom to consider time-varying characteristics in the individual regressions (Eq. (2)), as this would imply a proliferation of interactions of the regressors with a time-varying index for the different country characteristics. This approach would make more sense using panel techniques; however, as already argued above, panel techniques raise computational difficulties if we want to take into account the model uncertainty in our estimates of the US monetary policy shocks.[21]

4. The domestic effects of US monetary policy shocks

We begin by presenting our results for a contractionary US monetary policy shock in Fig. 1 for the BVAR estimated until the end of 2013, the full sample period. As it is customary, the figure reports the 16th, 50th (median) and 84th percentiles of the point by point posterior distribution of the impulse responses (the dotted red lines) in response to a one-standard deviation structural shock, as well as the mean response.[22] It is clear from the figure that the typical response to the shocks is estimated to be stronger and longer-lasting than assumed. The federal fund rate and the 1-year rate rise persistently, with the median responses peaking around 10 basis points. These responses are significant (i.e., the 16th percentile is above zero) for each of the first 10 months. The interest rate hike is associated with a shorter-lived widening in the mortgage spread, the commercial paper spread and the corporate bond spread, where only the latter's response (which we leave unrestricted) is not significant even on impact. As a result, the US price level, industrial production and stock prices fall significantly on impact and in later periods, with the effects dissipating (the 16th percentile becoming positive) one year to 4 years out. The trough median responses are smaller for the CPI (around −0.05%), and larger for stock prices (over −1%); the peak median decline in industrial production is around −0.3%.

Finally, most international variables respond as would be expected according to standard textbook predictions. The fall in the interest differential closely mirrors the hike in US rates, and is thus consistent with interest rates in other major currencies barely responding; the dollar effective exchange rate strongly appreciates, with a median response around 0.5%. The appreciation however is not significant at the 6-month horizon and thereafter, as the 16th percentile returns below zero. Turning to the unconstrained variables, despite the dollar appreciation, industrial production and stock prices fall in the rest of the world, while the large median decrease in commodity prices is always bracketed between a positive 16th percentile and negative 68th percentile. The contraction in world industrial production and stock prices is similar in magnitude to that in their US counterparts, albeit somehow less persistent. These responses are consistent with a transmission involving substantial complementarity between US and foreign manufacturing goods or a limited degree of expenditure switching — see e.g., Corsetti et al. (2010).

The impulse responses estimated excluding the most recent period after 2008 are broadly similar to those in Fig. 1, qualitatively and in most cases quantitatively — see Fig. 2. Notable exceptions

[21] An alternative could be to use informative priors on the time variation, obviously at the risk of unduly constraining posterior inference.

[22] Although we follow the conventional way to report results from the impulse response posterior, it should be clear that neither the median nor any other percentiles should be interpreted as arising from a specific orthogonalization matrix Q. Indeed, each point can result not only from a different orthogonalization matrix Q, but also from a different realization of the reduced form (B, Σ) — see Uhlig (2005).

Table 1
Average sample values of country characteristics.

Country	Lane and Shambaugh (2010) dollar peg	Chinn and Ito (2006) capital openness	Benetrix et al. (2015) dollar exposure	Trade exposure
Australia	0	1.422	51,771	4,09%
Austria	0	1.903	42,402	2,30%
Belgium	0	1.713	97,406	9,41%
Brazil	0	−1.147	34,089	3,71%
Canada	0.147	2.439	97,251	38,30%
Chile	0.059	−0.325	76,571	8,64%
China	0.618	−1.318	35,067	4,93%
Colombia	0	−1.149	44,366	8,29%
Czech Republic	0	1.559	33,450	2,58%
Denmark	0	1.719	69,907	2,88%
Estonia	0	2.390	18,112	3,11%
Finland	0	1.903	46,878	2,95%
France	0	1.410	46,067	2,57%
Germany	0	2.439	39,907	3,72%
Greece	0	0.487	14,873	1,13%
Hungary	0	0.272	24,311	2,53%
India	0.176	−1.169	24,344	2,24%
Italy	0	1.334	26,168	2,35%
Japan	0	2.348	46,915	4,54%
Korea	0.147	−0.361	41,999	11,56%
Latvia	0.045	2.307	24,785	1,31%
Lithuania	0.364	2.258	23,753	2,42%
Malaysia	0.353	1.025	68,766	22,54%
Mexico	0.147	0.435	44,349	27,30%
Netherlands	0	2.439	95,219	5,82%
Norway	0	1.237	72,906	3,20%
Philippines	0.147	−0.409	50,386	12,53%
Poland	0	−0.854	26,576	1,15%
Portugal	0	1.056	18,342	2,35%
Russia	0.045	−0.320	60,264	2,31%
South Africa	0	−1.309	30,956	4,28%
Spain	0	1.279	22,996	1,93%
Sweden	0	1.821	68,828	3,95%
Thailand	0.382	−0.245	41,979	10,94%
Turkey	0	−0.827	31,094	2,02%
UK	0	2.390	207,213	4,47%

concern the response of the mortgage spread and especially the commercial paper spread, which is now smaller than when the period after 2008 is included. Conversely, the response of the corporate bond spread is now significant at the 6-month horizon. However, with the exception of the commercial paper spread, these responses would be included in the percentiles estimated over the whole sample.

We conclude this section by reporting on a few exercises we carry out to provide further corroboration of our results. First, we re-estimate the BVAR impulse responses by dropping the interest rate differential (not shown here to save on space). We find that most of these impulse responses are similar to those in Fig. 1, but there are some notable differences. In particular, the responses of interest rates are now significant for many more periods, with the 16th percentile staying positive for more than 40 months. Moreover, the responses of several variables are somehow larger than in Fig. 1, especially those of the international variables. When we reestimate the VAR over the sample ending in 2008 again omitting the interest rate differential, instead results are very similar to those in Fig. 2. As discussed above, this difference underscores the importance of including the short-term interest rate differential in our analysis to make results more robust to the inclusion of the most recent period with interest rates at their lower bound. Indeed, this interest rate differential has been as stable over this period as US short-term rates.

Second, we report in Fig. 3 the distribution of correlations of our shocks with the (point estimates of the) GK shocks and also the extended series of the Romer and Romer (2004) shocks as computed by Barakchian and Crowe (2013). The correlations are mostly positive in both cases. As shown in Table 3, median values range between 0.12 and 0.21, depending on the shocks and the samples. These values are similar to that of the correlation between the GK and RR shocks, equal to 0.19.[23]

Third, we computed impulse responses of the monthly US VIX index to our identified shocks, again using a specification like Eq. (2).[24] We could not include the VIX directly in the BVAR because it is available only after the early 1990s. This could be an important omission in light of the results in Rey (2013), where the VIX, taken as a proxy for the "global financial cycle", is shown to be correlated with capital flows and asset prices across countries, as well as to increase in response to a US monetary policy tightening. Fig. 4 reports the impulse responses of the VIX to our monetary policy shocks, estimated again over both samples. Similarly to the impulses responses in the previous figures, the (blue) dotted lines represent the point-by-point 16th, 50th and 84th percentiles. It is clear that an unexpected monetary tightening in the US, as measured by our shocks, results in a substantial (around 7% on impact, in response to a one-standard deviation structural monetary policy shock) and fairly persistent increase in the VIX, in line with the results in Rey (2013). The responses are also broadly similar across the 1990–2008 and 1990–2013 samples. This finding, together with our result that US and especially global stock prices fall in response to a US interest rate hike, shows that our estimated monetary policy shocks are consistent with salient features of the effect of US monetary policy on key global financial variables as documented by Miranda-Agrippino and Rey (2015).

[23] We also checked the first order autocorrelation coefficient of our estimated monetary policy shocks, whose median value is around 0.05 over the whole sample.
[24] These results are broadly insensitive to the prior we use.

Table 2
Country classifications[a].

Income level		Exchange rate regime		Capital openness		Dollar exposure		Trade openness		Commodity exporters	
Adavanced	Emerging	Floaters	Dollar pegs	More	Less	More	Less	More	Less	Exporers	Non-exporters
Australia	Brazil	Australia	China	Australia	Brazil	Belgium	Australia	Australia	Austria	Australia	Austria
Austria	Chile	Austria	India	Austria	Chile	Canada	Austria	Belgium	Czech Republic	Brazil	Belgium
Belgium	China	Belgium	Malaysia	Belgium	China	Chile	Brazil	Brazil	Denmark	Canada	China
Canada	Colombia	Brazil	Mexico	Canada	Colombia	China	Colombia	Canada	Estonia	Chile	Czech Republic
Denmark	Czech Republic	Canada	Philippines	Czech Republic	Greece	Czech Republic	Estonia	Chile	Finland	Colombia	Denmark
Finland	Estonia	Chile	Thailand	Denmark	Hungary	Denmark	Finland	China	France	Norway	Estonia
France	Hungary	Colombia		Estonia	India	France	Greece	Colombia	Greece	Russia	Finland
Germany	India	Czech Republic		Finland	Korea	Germany	Hungary	Germany	Hungary	South Africa	France
Greece	Latvia	Denmark		France	Malaysia	Japan	India	Japan	India		Germany
Italy	Lithuania	Estonia		Germany	Mexico	Korea	Italy	Korea	Italy		Greece
Japan	Malaysia	Finland		Italy	Norway	Malaysia	Latvia	Malaysia	Latvia		Hungary
Korea	Mexico	France		Japan	Philippines	Netherlands	Lithuania	Mexico	Lithuania		India
Netherlands	Philippines	Germany		Latvia	Poland	Norway	Mexico	Netherlands	Norway		Italy
Norway	Poland	Greece		Lithuania	Portugal	Russia	Philippines	Philippines	Poland		Japan
Portugal	Russia	Hungary		Netherlands	Russia	South Africa	Poland	South Africa	Portugal		Korea
Spain	South Africa	Italy		Spain	South Africa	Spain	Portugal	Sweden	Russia		Latvia
Sweden	Thailand	Japan		Sweden	Thailand	Sweden	Thailand	Thailand	Spain		Lithuania
UK	Turkey	Korea		UK	Turkey	UK	Turkey	UK	Turkey		Malaysia
		Latvia									Mexico
		Lithuania									Netherlands
		Netherlands									Philippines
		Norway									Poland
		Poland									Portugal
		Portugal									Spain
		Russia									Sweden
		South Africa									Thailand
		Spain									Turkey
		Sweden									UK
		Turkey									
		UK									

[a] The sources and references for our classifications are the following:

- Income level from IMF World Economic Outlook.
- Commodity exporters based on the incidence of net exports of primary goods (fuels, metals, food and other raw materials) over total exports plus imports, UNCTAD data.
- For other classifications see Table 1.

Table 3
Correlations between Romer and Romer's (2004), Gertler and Karadi's (2015) and our shocks.[a]

Shocks	Mean	Median	Max	Min
R&R with MPS 2013	0.13	0.13	0.39	−0.13
R&R with MPS 2008	0.12	0.12	0.34	−0.12
G&K with MPS 2013	0.15	0.15	0.38	−0.07
G&K with MPS 2008	0.21	0.21	0.43	−0.06
G&K with R&R		0.19		

[a] The Romer and Romer's shocks are the updated series from Barakchian and Crowe (2013).

To summarize, these exercises together lend reasonable support to our benchmark identification and the effects of the resulting monetary policy shocks.

5. Evidence on the global transmission of US monetary policy shocks

In the next subsection we provide a broad overview of the country specific responses to the US monetary policy shocks. In Section 5.2 we explore whether these responses have any commonality that can be attributed to shared country characteristics. Before going into the details of the results, it is useful to provide an overview of the key findings. First, a surprise US monetary tightening leadsto a dollar

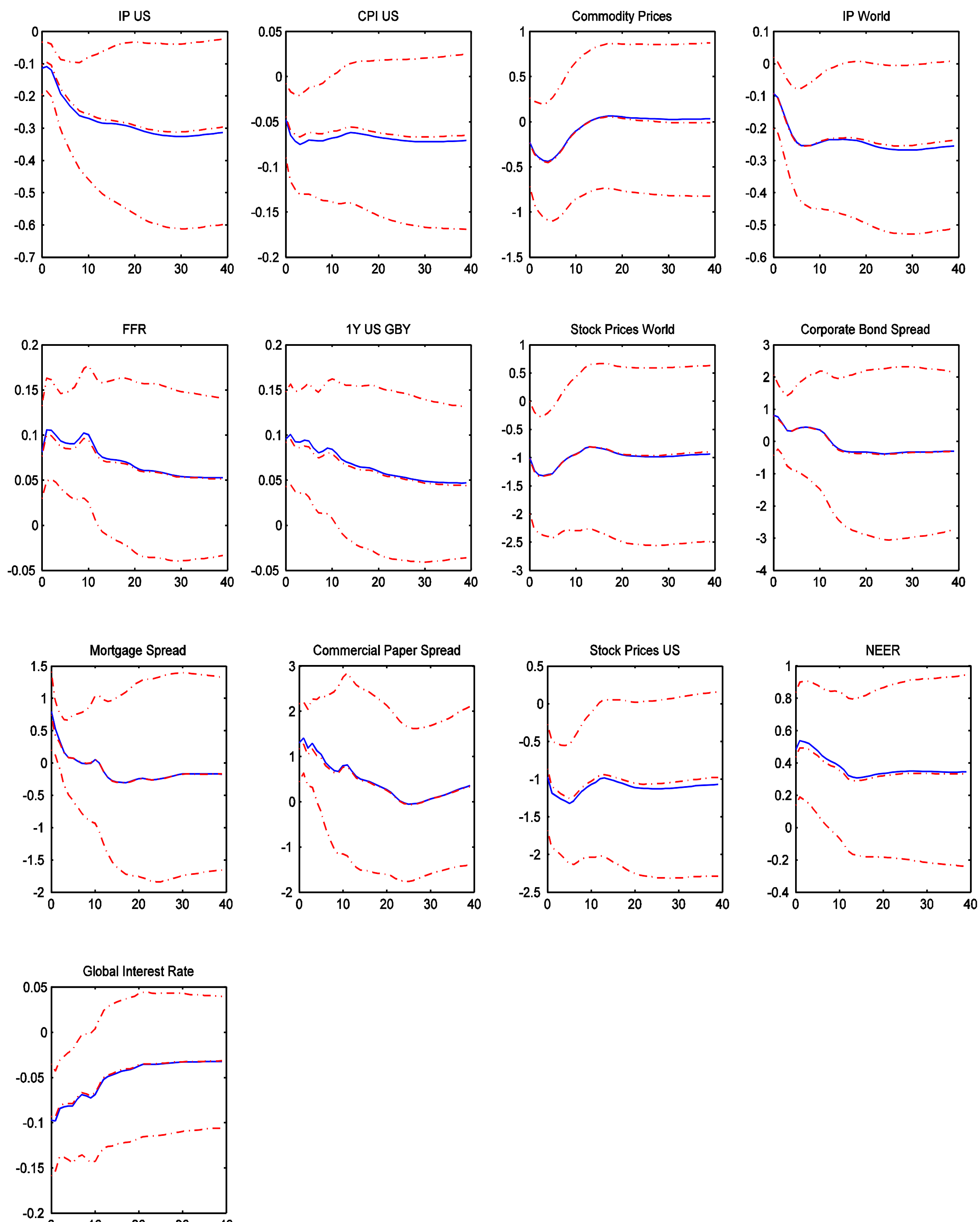


Fig. 1. IRFs from baseline BVAR estimated over the sample 1980–2013.

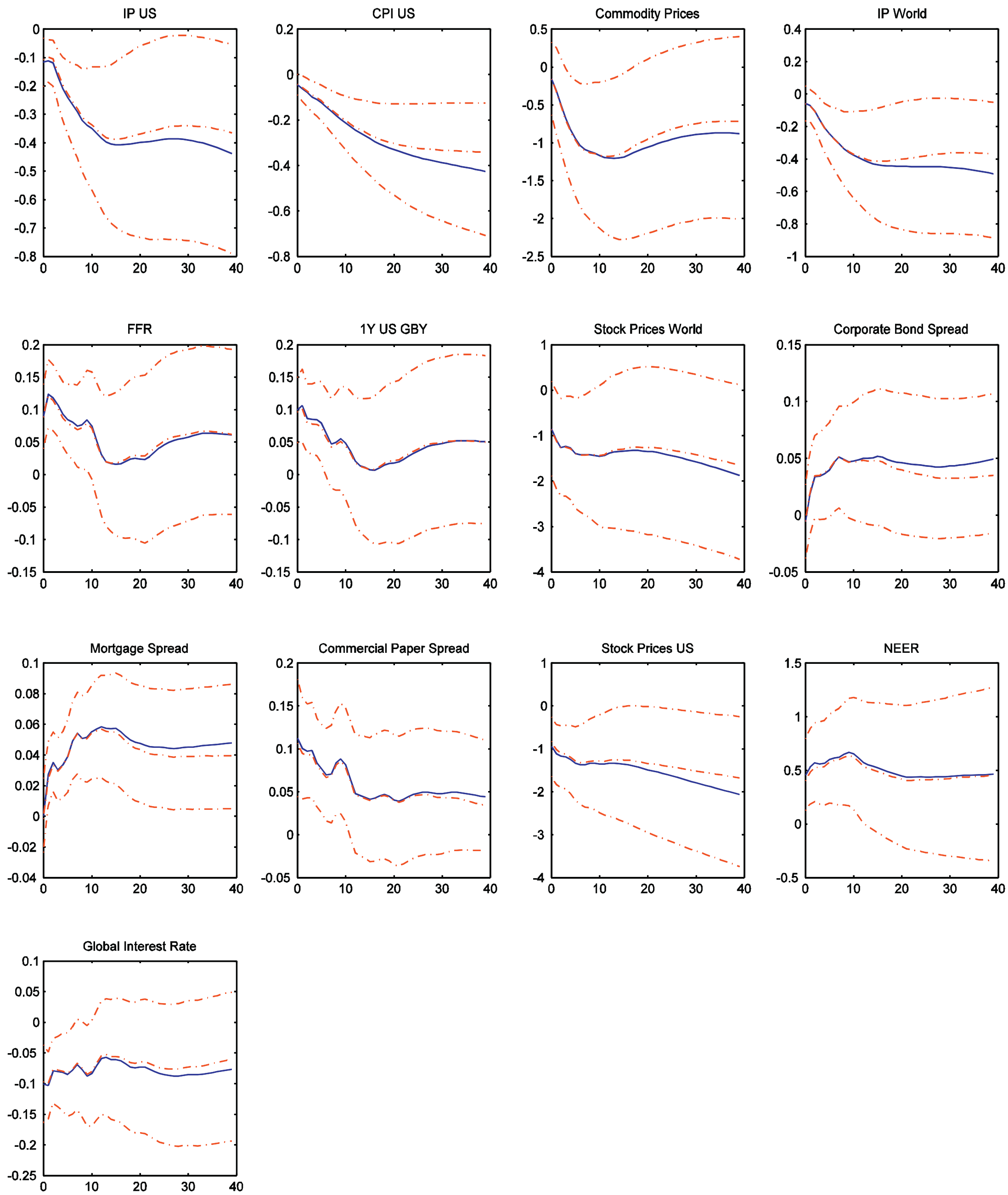


Fig. 2. IRFs from baseline BVAR estimated over the sample 1980–2008.

appreciation vis-á-vis most countries in our sample. In a large majority of countries industrial production and real GDP fall, and unemployment rises; however the trade balance improves. Inflation (both GDP deflator and CPI) also falls in a majority of countries, although the effects are less statistically significant. Emerging economies tend to experience more volatile macroeconomic effects. At the same time, and this is our second finding, the responses of financial variables are more heterogeneous: while most countries see their bond yields increase relative to the US, real equity and housing prices drop in about half the countries. Likewise, many countries experience opposite effects on real credit and capital flows, including borrowing from foreign banks. Finally, we do not find any of systematic relations between the most likely country characteristics (income level, exchange rate regime, financial openness, trade openness vs. the US, dollar exposure and commodity exporting) and the distribution of cross-country responses to US monetary policy shocks (in Section 5.2). While a dollar peg at least mutes the effects on the nominal and real exchange rate, asset prices and capital flows do not seem to react differently between more and less financially open countries.

5.1. The cross-country distribution of the effects of US monetary policy shocks

We summarize the effects across countries of US monetary policy shocks in Fig. 5. For each variable the figure reports a chart with the maximum absolute value over an horizon of 5 years of the median responses to a one-standard deviation monetary shock, country by country.[25] The responses in advanced economies are depicted in blue

[25] Of course, as noted before, the median response may well reflect the effects of different monetary policy shocks across countries and variables.

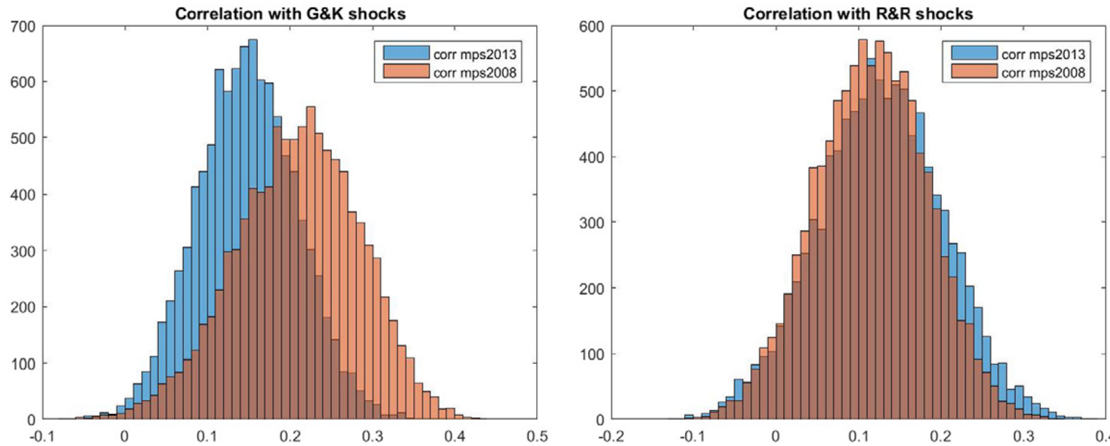


Fig. 3. Correlations between our estimated shocks and (i) Gertler and Karadi (2015) shocks, and (ii) updated Romer and Romer shocks from Barakchian and Crowe (2013).

bars, those of emerging economies in red bars. The peak impulse response for the euro area is reported in yellow, and the overall country average in black to the far right-hand side of each chart.

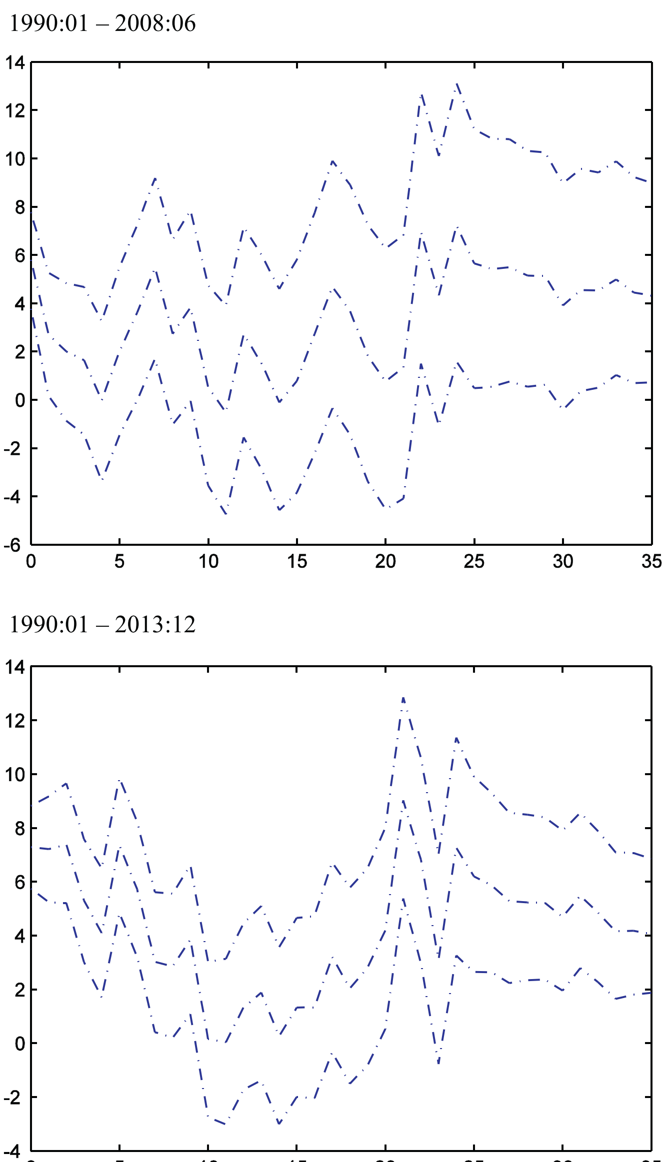


Fig. 4. VIX responses to US monetary policy shocks, monthly regressions.

Recall that euro area countries are not included individually in the case of the bilateral dollar exchange rate and of short-term rates. The first panel of Fig. 5 shows the maximum absolute responses of monthly variables, while the second panel shows the maximum absolute responses for quarterly variables — see Section 3 for the list of variables included.

Starting with the first panel in Fig. 5, it is apparent that virtually all countries experience a nominal bilateral depreciation (a positive value) with the US dollar.[26] The largest significant depreciation, almost 1.4%, occurs in Hungary. The response is not significant for a few currencies, in particular for countries managing their exchange rate vis-á-vis the US dollar, such as China. However, the euro depreciation, while showing a large median peak of 1.3%, is not significant too. The widespread bilateral dollar depreciation transpires into a broad based real depreciation (a negative value) in more than half of the countries, mostly advanced ones. However, only in a few countries the responses are now statistically significant.[27] Sweden experiences the largest significant peak depreciation, −0.5%; the largest significant real appreciation, 0.5%, takes place in China.

Cross-country heterogeneity in the responses of other asset prices is larger. Short-term interest rates tend to moderately fall relative to the US in advanced countries; e.g., the peak differential is −11 basis points in the euro area. They increase, sometimes sizably, in emerging ones, such as Chile, where the peak differential is 62 basis points. The responses of differentials in longer-term yields are more similar across countries, displaying a generalized small increase (Greek bonds experience the largest significant positive differential, 56 basis points). However the differential turns negative in a few emerging economies (the largest relative fall, −80 basis points, occurs in Turkey). Finally, against the background of a 1% drop in the US, stock prices fall in most emerging markets and several advanced economies; some countries however experience significant increases.[28]

Conversely, the sign of the responses of macroeconomic variables is quite similar across economies. Industrial production and the CPI fall in most countries, while the trade balance improves. The decline in industrial production is fairly significant in a majority of countries (the largest in Lithuania at −1.2%), and among advanced economies (−0.8% in Japan). Euro area IP also significantly contracts, by −0.5%.

[26] The exception is Estonia, whose appreciation however is not significant as it is bracketed between the 16th and 84th percentile. Brazil's responses for the dollar exchange rate, the CPI and the short-term differential are not shown as they are very large (standing at −9.86, −10.66 and 493.18, respectively) and imprecisely estimated.

[27] The real effective exchange rate is now reported for all members for the euro area separately. On aggregate, the euro depreciates in real terms but not significantly.

[28] Lithuanian stock prices fall significantly the most, −2.8% (Norwegian stocks by −1.3% the most among advanced countries); the largest significant increase occurs in China, almost 4%; the euro area increase is not significant.

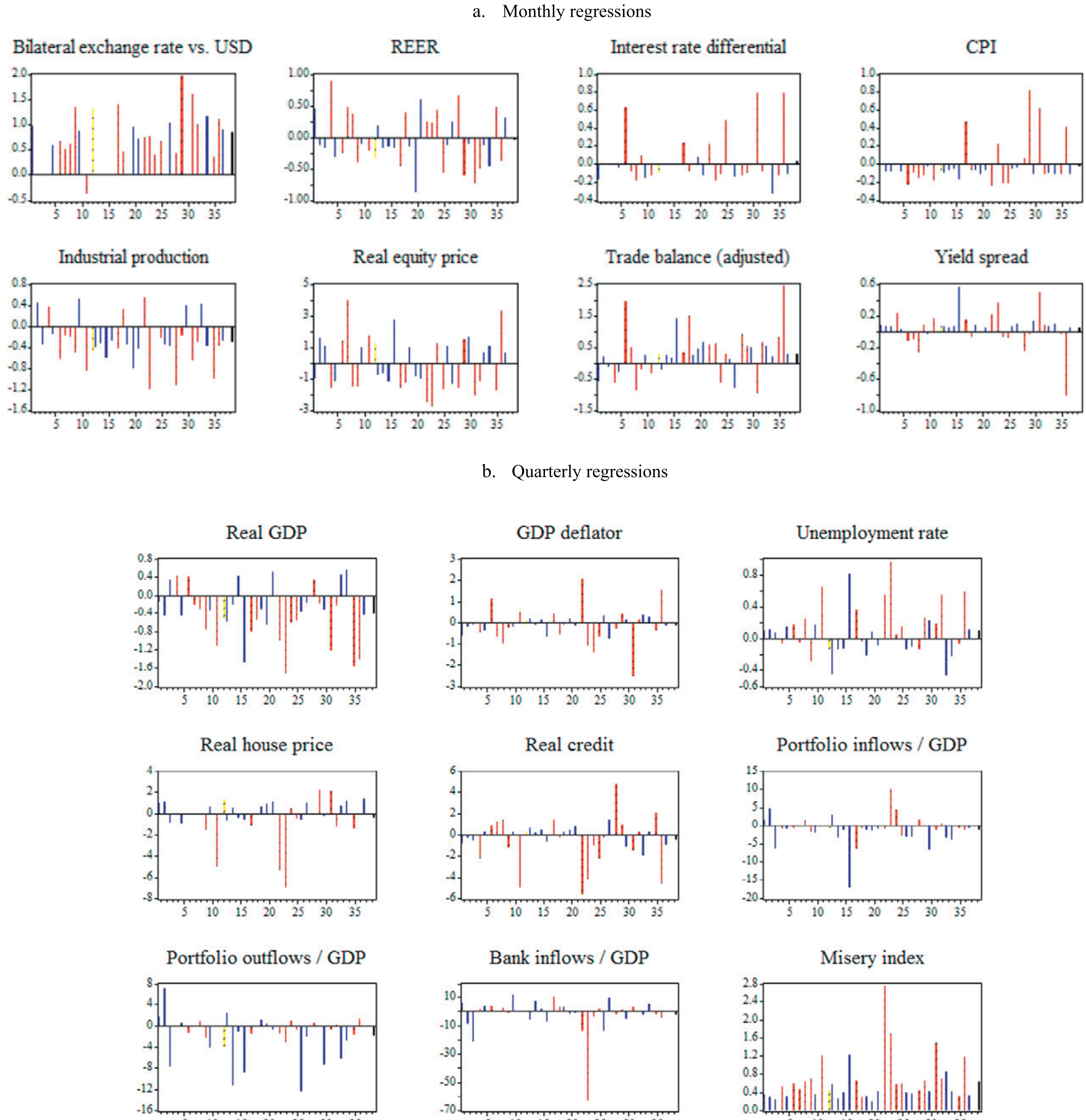


Fig. 5. Country-specific median peak impulse responses to a one standard deviation contractionary US monetary policy shock. a. Monthly regressions. b. Quarterly regressions. Sample period: 1980 - 2013. Note that we exclude inflation, the nominal exchange rate and the interest rate differential in Brazil as well as bank inflows into China due to very high values. Blue bars refer to advanced countries, red bars to emerging countries. The peak impulse response for the euro area is reported in yellow, and the overall country average in black to the very right.

Interestingly, these falls are bigger than the US own response. In contrast, the few positive responses are never significant. The CPI displays a similar, generalized reduction. For instance, nominal prices significantly fall in the euro area, by −0.1% (Malaysian CPI decreases the most, by −0.2%). The trade balance improves in most countries, both advanced and emerging; however, countries like Norway and Russia experience especially significant and large deteriorations (by −0.8% and −1% respectively).[29]

Turning to the second panel of Fig. 5, we also find that the effects on quarterly macroeconomic variables such as the real GDP and its deflator, and unemployment, do not greatly differ in sign across countries. The real GDP contraction is statistically significant in a majority of countries, including the euro area as a whole, where the peak effect is −0.5%. Unemployment rises in around half of the countries (the largest responses of real GDP and unemployment, at −1.75% and 1%, occur in Lithuania). The fall in the GDP deflator is also less widespread than the real GDP contraction, and more muted (Malaysia, whose CPI also falls significantly, experiences the largest drop, −1.4%). Both variables are barely affected in the euro area, the GDP deflator despite the significant decline in the CPI reported above.

We find a great deal more heterogeneity across countries in the responses of financial variables and capital flows. Real housing prices decline in many emerging economies, but are large and significant especially in the Baltic countries. Advanced economies

[29] The largest significant improvement, 2.4%, takes place in Turkey; among advanced economies in Greece, by 1.4%.

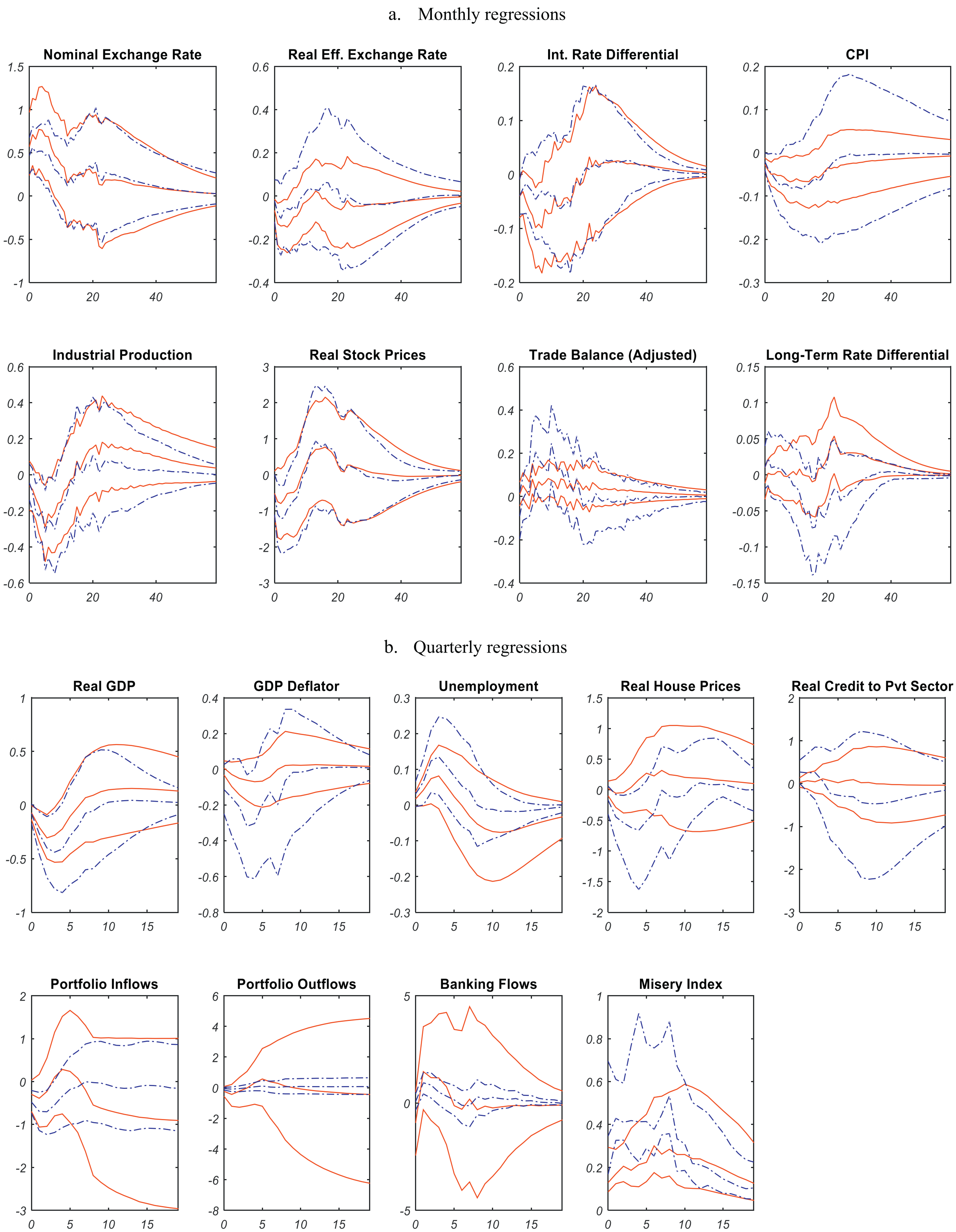


Fig. 6. Responses of advanced (solid red line) and emerging economies (dotted blue line) to US monetary policy shocks. a. Monthly regressions. b. Quarterly regressions.

tend to experience small but generally little significant increases, including the euro area. The response of real private credit varies a great deal across countries, falling in several emerging economies, although with little statistical significance, but also in advanced economies like Belgium, where it declines by a significant −0.6%. However, the generally positive responses in advanced economies are also rather muted. Finally, capital flows, including borrowing from foreign banks (all scaled by nominal GDP), display quite

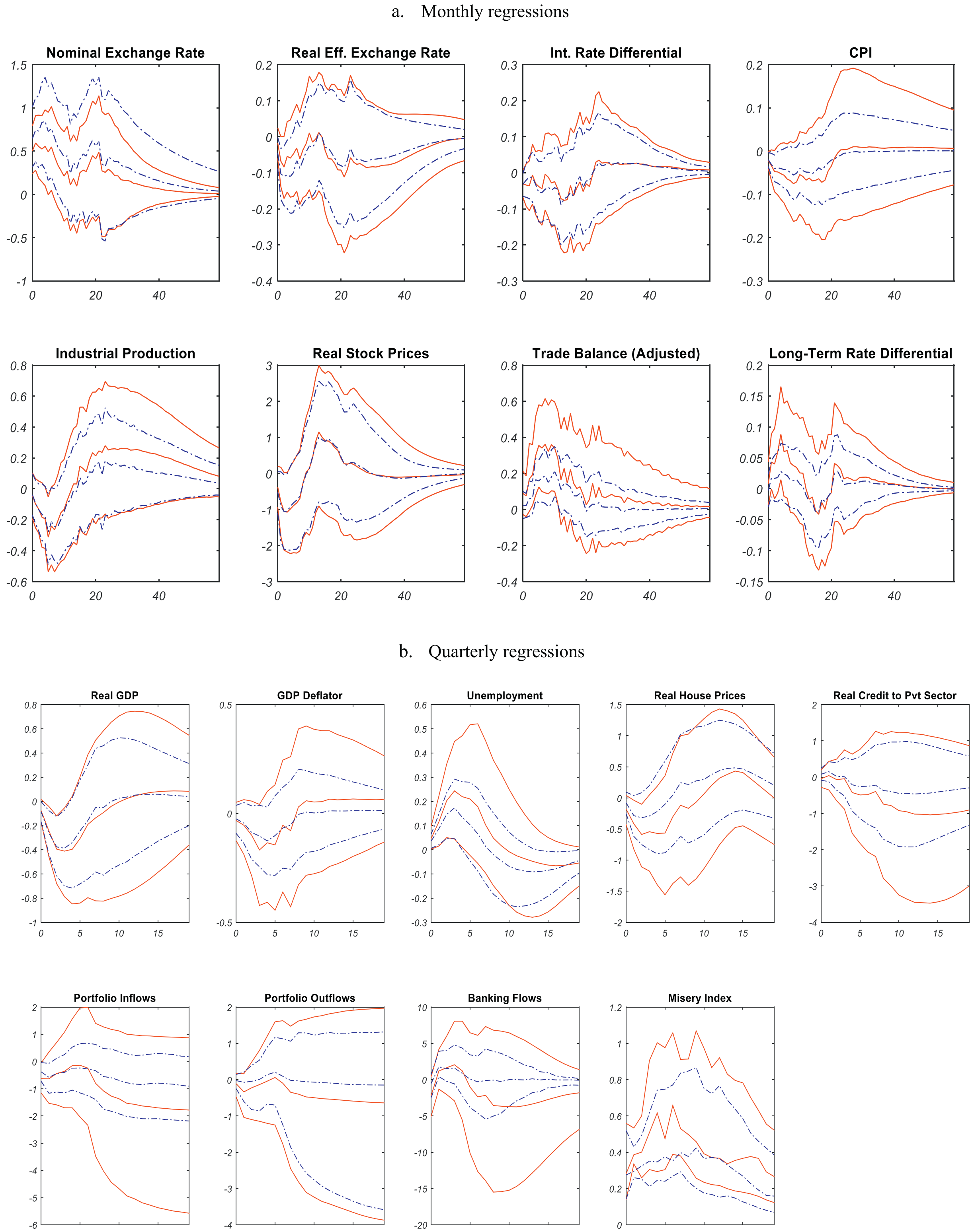
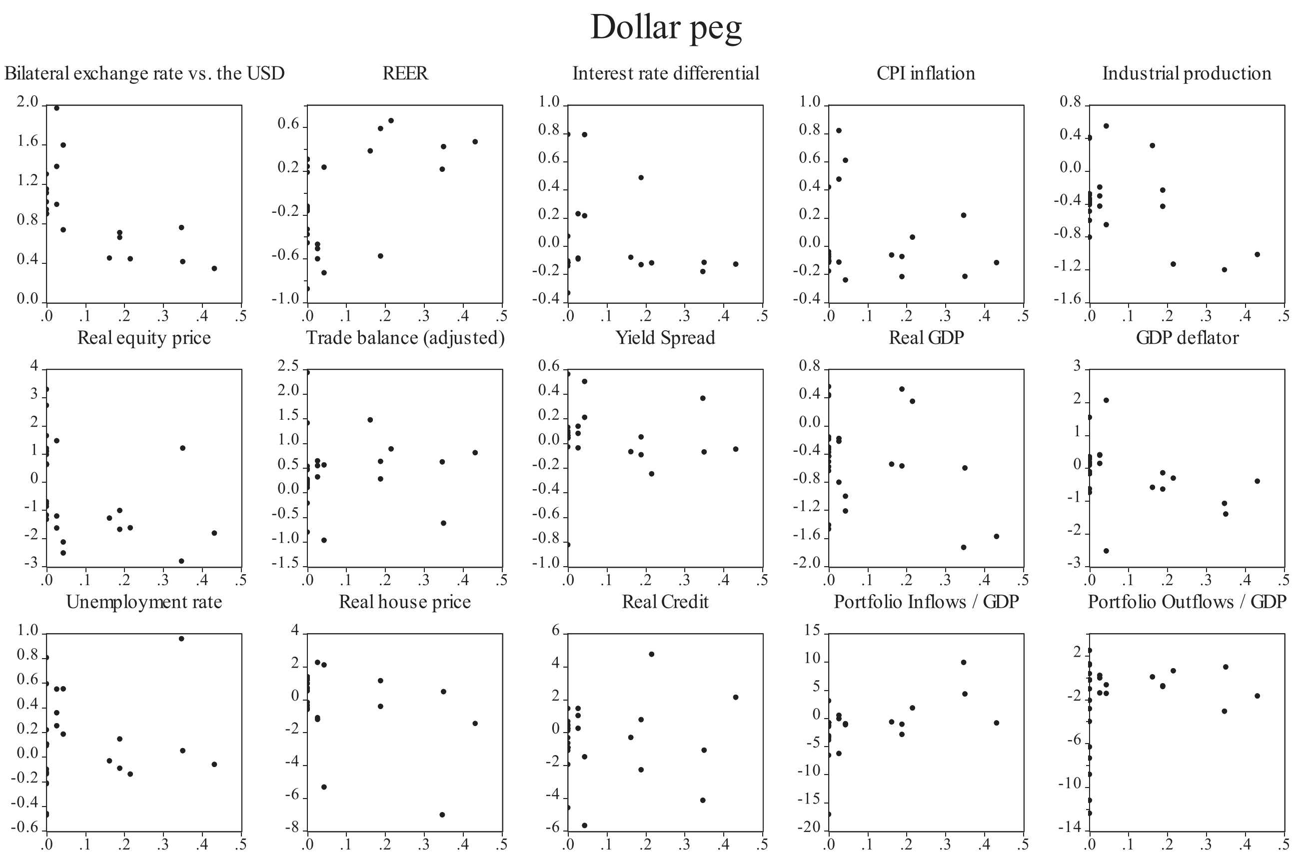


Fig. 7. Responses of EMEs with lower (solid red line) and higher capital mobility (dotted blue line) to US monetary policy shocks. a. Monthly regressions. b. Quarterly regressions.

different effects in sign and size. Cumulated portfolio inflows by foreign residents but also outflows by domestics decline in a majority of countries, including many advanced economies. For instance, even the euro area experiences a significant decline in portfolio inflows of around −1%. The decline in outflows, though also large at over −4%, is not significant. Total borrowing from foreign



Note: the scatter plots report peak impulse responses from the second stage regression (y axis) against the average country characteristic (dollar peg) over the whole sample (x axis).

Fig. 8. Country-specific median peak impulse response against average country characteristics.

banks displays many positive and negative responses across both advanced and emerging countries. However positive responses tend to be significant in a majority of the former, negative responses in a majority of the latter.[30] For instance, borrowing from foreign banks significantly soars in Denmark by 11% and drops by −4.7% in Turkey.

To summarize, a US surprise tightening brings about a widespread dollar appreciation and a fall in broad macroeconomic activity, with an improvement in trade balances in nominal terms. Inflation also tends to fall in most countries, although less sharply. Emerging economies experience more volatile macroeconomic effects, as summarized by the "misery index". Among financial variables, the increase in long-term government bond differentials and, in a more limited fashion, the drop in equity prices are also fairly generalized. Conversely, the response of short-term rates differentials and housing prices is more heterogeneous. By the same token, several countries experience opposite effects on real private credit and capital flows, especially borrowing from foreign banks.

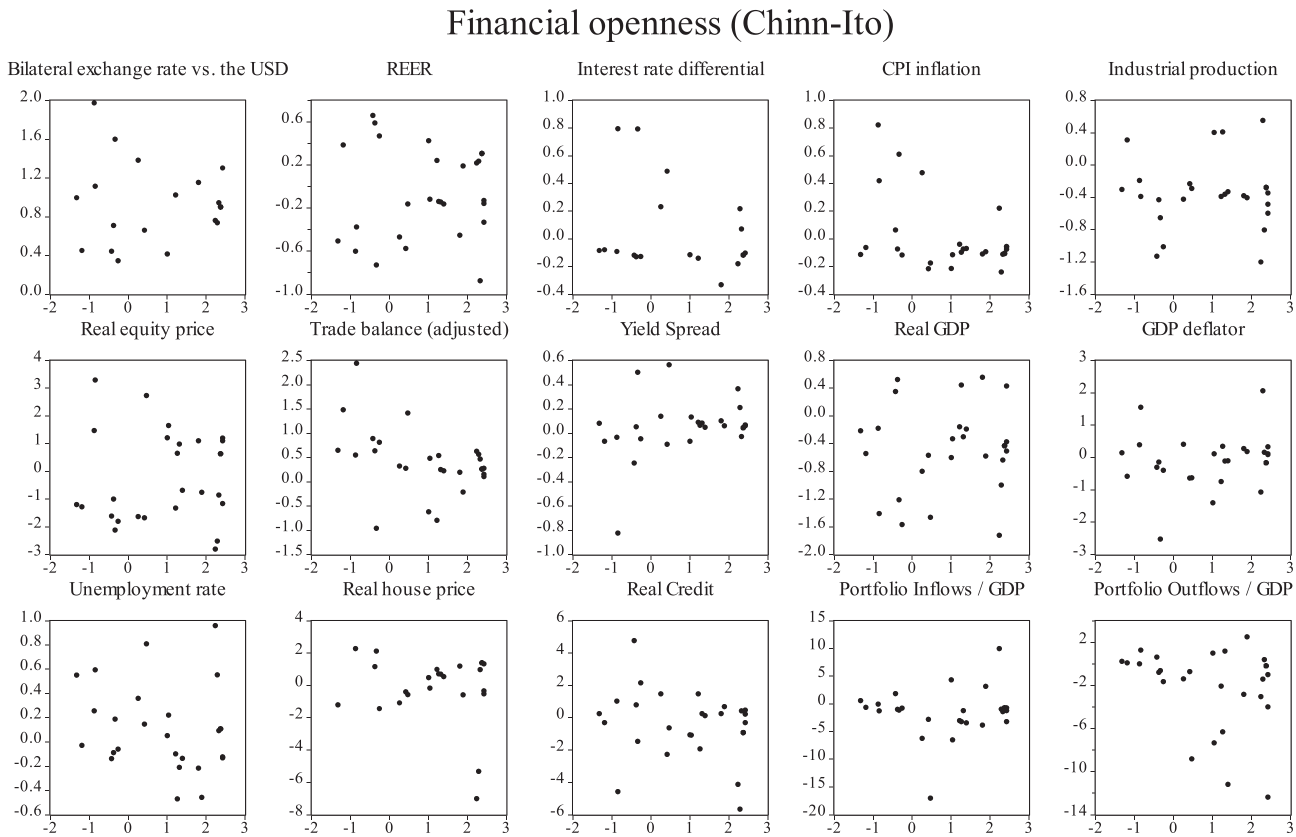
5.2. Country characteristics and the effects of US monetary policy shocks

In the following, we find it convenient to organize the results for both monthly and quarterly regressions by country groupings. Therefore, for each figure, the first panel (a) will show impulse responses obtained from monthly regressions, while the second panel (b) will depict impulse responses obtained from quarterly regressions. As before, the figures report the point-by-point 16th, 50th and 84th percentiles of the impulses responses taking the median across country groups. The latter are described in Table 2.

5.2.1. Advanced vs. emerging countries

We start by presenting results by splitting countries on the basis of their income levels (see first and second column in Table 2), as displayed in Fig. 6. The percentiles of distribution of the median responses of the 18 AEs are shown in the solid (red) lines, while those of the 18 EMEs are shown in dotted (blue) lines. These responses confirm and extend our previous results that a US monetary policy shock has substantial cross-border effects. Panel (a) shows that in the "typical" country in the rest of the world, an unexpected interest rate tightening is associated with depreciation both nominally against the US dollar and on a real trade-weighted basis — whereas a fall again indicates depreciation. While for both AEs and EMEs all percentiles of the bilateral dollar rate are positive on impact and for some periods after (the bilateral depreciation is significant for both groups), for the real exchange rate the percentiles are all negative only in AEs (the real depreciation is significant only for this group). Industrial production significantly declines across the board, while stock prices do so significantly only in EMEs. These variables seem to react similarly to their BVAR analogs, and thus to the US counterparts, though in a less persistent fashion. The responses of other variables are also very similar across the AEs and EMEs groups. The trade balance significantly improves in both groups of countries, while the decline

[30] Lithuania displays the very large negative response in the chart, which is however not significant. We do not report the much larger Chinese negative response, standing at −117.21.



Note: the scatter plots report peak impulse responses from the second stage regression (y axis) against the average country characteristic (financial openness) over the whole sample (x axis).

Fig. 9. Country-specific median peak impulse response against average country characteristics.

in the CPI and short-term interest differentials tends to be estimated more precisely in AEs. The response of the long-term interest differential is not precisely estimated for either group.

The responses of quarterly variables displayed in panel (b) confirm and further sharpen these results. In both AEs and EMEs groups, the contraction in industrial production is also associated with a fall in broad-based output as measured by real GDP, and an increase in unemployment. The median fall in the GDP deflator is never significant, however, in either group. The median increase in real credit is marginally significant only in EMEs. But some interesting quantitative differences emerge from the responses of other financial variables. The responses on housing prices and portfolio inflows are quite dispersed in the AEs group, but they tend to significantly decline among emerging economies. Conversely, the effects on portfolios outflows by domestic residents are not precisely estimated in either country groups.

A first important result then is that the repercussions of a US monetary policy shock on economic activity are qualitatively and quantitatively similar across advanced and emerging economies, since a US tightening brings about a recession and an increase in unemployment in both groups. Inflation and interest rate dynamics are also broadly similar, but with higher dispersion and volatility among EMEs than among AEs. As a result, macroeconomic volatility as captured by the sum of absolute changes in inflation and unemployment (the "misery index"), tends to be marginally higher for EMEs than for AEs. On the other hand, some negative financial spillovers are estimated more sharply for EMEs, especially concerning asset prices, and foreign capital outflows.

5.2.2. Other country characteristics: financial openness, currency regime, commodity exports and US trade and dollar exposure

We turn next to the analysis of the effects of other country-specific dimensions on the transmission of US monetary policy shocks with a view of exploring some of the possible reasons behind the asymmetric response across countries. Specifically, among emerging markets, we consider differences in the exchange rate regime, the degree of capital mobility, trade openness toward the US and US dollar exposure, and the incidence of commodity exports.[31] Surprisingly, we find that none of the chosen characteristics appears to explain country heterogeneity.[32] To save on space, in Fig. 7 we thus present results only for the degree of capital mobility (results for other characteristics are available in the online appendix). Although impulse responses are sometimes different between groups, they overlap in most cases, so that their difference is never statistically significant. This includes also the responses of capital and banking flows between the two groups of EMEs with higher and lower capital mobility in panel (b) of Fig. 7. By the same token, we do not detect large discrepancies in the exchange rate regime vs. the US dollar (see the online appendix). Even in the case of the bilateral exchange rate,

[31] We focus on emerging markets because financial openness tends to be uniformly higher in advanced countries. AEs are also all classified as floating relative to the dollar according to the Klein–Shambaugh metric. Likewise, most commodity exporters are EMEs.

[32] Results do not change for these last three characteristics when we look at different degrees of exposure and commodity exports among all countries, or among advanced economies only.

which as expected reacts less in dollar pegs than in other countries, the difference is not large and indeed not statistically significant. Also the interest rate reaction does not seem to significantly differ between pegs and floats, differently from previous results in the literature (e.g., Shambaugh, 2004).

Our results so far, however, are predicated on the assumption that country characteristics are constant across the sample and can be summarised as 0–1 dummies. Several country characteristics, such as differences in income levels, are relatively persistent, but not all of them are necessarily so. For example, there is some time variation in the foreign exchange regime. Some of the countries we classify as dollar pegs over the whole sample, in reality have had also spells of floating rates. This is the case of India and Mexico, for instance, which seem thus closer to an intermediate exchange rate regime over the whole sample. By the same token, several countries have an intermediate degree of financial openness, often the results of incremental measures of financial liberalization. Time variation in some characteristics can thus account for the lack of sharp differences between some groups, especially concerning the effects of the exchange rate regime.

We try to address some of these concerns in a robustness exercise in which we relate the effects of US monetary policy shocks to the actual sample averages of the two key country characteristics for the question of the trilemma, as measured by the actual values of the indices of dollar exchange rate flexibility (Klein–Shambaugh) and financial openness (Chinn–Ito), as reported in Table 1. The indexes can now take a whole range of values, capturing in a less coarse way the degree of variation across countries. Figs. 8 and 9 report scatter plots of the median peak impulse responses of each variables against the sample average of these two characteristics.

Fig. 8 now shows that countries with a higher value of the index (i.e., countries which have a less flexible dollar exchange rate on average) display a smaller effect of US monetary shocks on the bilateral exchange rate. Likewise, these countries tend to experience a smaller depreciation of their trade-weighted real exchange rate, and in several cases even an appreciation (a positive value in the plot). Moreover, some of the largest drops in industrial production and also equity prices occur in countries with a less flexible dollar exchange rate. More surprisingly, these countries, however, also display a more negative short-term interest rate differential and inflation. At the same time, and in line with our previous results, we find no evidence of a systematic effect of the exchange rate regime on real GDP, unemployment, housing prices, credit, and capital flows. Finally, Fig. 9 confirms that there is also apparently no link between the (now finer) degree of financial openness and any variable responses, with perhaps the exception of residents outflows, which tend to be more negative for countries with higher capital mobility (a higher value of the index).

Overall, these results confirm our previous findings that the degree of capital mobility and the exchange rate regime do not seem to be key drivers of most of the spillover effects of US monetary policy shocks.

6. Concluding remarks

In this paper we investigate the global effects of US monetary policy shocks using a two stage approach. First, we obtain estimates of US monetary policy shocks by using an identification scheme based on sign restrictions, consistent with the domestic effects of these shocks in standard monetary models. This allows us to model the responses of a range of interest rates and spreads to a US monetary policy shock. Further, a number of macroeconomic and financial variables at monthly and quarterly frequency are regressed on the estimated shocks to compute impulse responses in 18 advanced and 18emerging economies. Countries are grouped on the basis of characteristics like their income levels, the openness of their capital accounts, their dollar exchange rate regime, dollar exposure in their foreign assets and liabilities, trade openness vs. the US, and incidence of commodity exports. All these variables are plausible candidates to explain the cross country incidence of the effect of US monetary policy shocks.

We find that a surprise US monetary tightening leads to depreciation vis-á-vis the US dollar in most countries in our sample; industrial production and real GDP fall, and unemployment rises, despite an improving trade balance. Inflation (GDP deflator and CPI) also tends to fall in a majority of countries. Emerging economies tend to experience more macroeconomic volatility. Responses of financial variables are more heterogeneous and muted: bond yields increase relative to the US yields in most countries, while real equity and housing prices drop in about half the countries. Finally, we do not find evidence of a systematic relation between country characteristics and the distribution of cross-country responses to US monetary policy shocks. While less flexibility in the dollar exchange rate regime at least seems to limit the response of the nominal and real exchange rate, which can be expected, asset prices and capital flows do not react differently between more and less financially open emerging economies.

A main policy implication of this finding is that, conditional on monetary policy shocks, neither the exchange rate regime nor financial openness, at least the way we measure them, appear to matter much for the international transmission of US monetary policy. In particular, in line with Miniane and Rogers (2007), we do not find compelling evidence that capital controls may provide an effective protection against US monetary spillovers. Nevertheless, it should be kept in mind that we do find evidence of significant country heterogeneity, which suggests that spillovers are indeed asymmetric — though the asymmetry is not well explained by the country characteristics we have explored in the paper. Further research may focus on refining the measurement of the country characteristics we have used, and exploring whether other country characteristics matter to explain the apparent heterogeneity in the effects.

Acknowledgments

The views expressed in the paper belong to the authors and are not necessarily shared by the European Central Bank (ECB). We thank Peter Karadi for sharing his data and Michele Lenza his BVAR codes, as well as B. Born, F. Canova, R. Clarida, G. Corsetti, T. Dahlhaus, M. Del Negro, J. Frankel, G. Georgiades, L. Goldberg, O. Jeanne, S. Kalemli-Ozcan, P. Lane, G. Primiceri, H. Rey, G. Ricco, J. Rogers, V. Vanasco, I. van Robays, and participants in seminars at CEMLA, the IMF, NSU, NY Fed, CEPR Annual IMF Programme Meeting 2016, the 9th International Research Forum on Monetary Policy, NBER IFM Summer Institute 2015 and ISoM 2016, and the 3rd University of Ghent Workshop on Empirical Macroeconomics for useful comments and suggestions. All remaining errors are our own.

Appendix A. Supplementary data

Supplementary data to this article can be found online at http://dx.doi.org/10.1016/j.jinteco.2017.01.002.

References

Bańbura, M., Giannone, D., Reichlin, L., 2010. Large Bayesian vector autoregressions. J. Appl. Econ. 25 (1), 71–92.
Banerjee, R., Devereaux, M.B., Lombardo, G., 2016. Self-oriented monetary policy, global financial markets and excess volatility of international capital flows. J. Int. Money Financ. 68, 275–297.
Barakchian, C., 2013. Monetary policy matters: evidence from new shocks data. J. Monet. Econ. 60 (8), 950–966.
Baumeister, C., Hamilton, J., 2015. Sign restrictions, structural vector autoregressions, and useful prior information. Econometrica 83 (5), 1963–1999.
Benetrix, A.S., Shambaugh, J.C., Lane, P.R., 2015. International currency exposures, valuation effects and the global financial crisis. J. Int. Econ. 96 (1), 98–109.
Bernanke, B., 2015. Federal Reserve policy in an international context. IMF 16th Jacques Polak Annual Research Conference.

Canova, F., 2005. The transmission of US shocks to Latin America. J. Appl. Econ. 20 (2), 229–251.
Canova, F., De Nicoló, G., 2002. Monetary disturbances matter for business cycle fluctuations in the g-7. J. Monet. Econ. 49 (6), 1131–1159.
Chinn, M., Ito, H., 2006. What matters for financial development? Capital controls, institutions, and interactions. J. Dev. Econ. 81 (1), 163–192.
Corsetti, G., Dedola, L., Leduc, S., 2010. Optimal monetary policy in open economies. Handbook of Monetary Economics,
Di Giovanni, J., Shambaugh, J.C., 2008. The impact of foreign interest rates on the economy: the role of the exchange rate regime. J. Int. Econ. 74 (2), 341–361.
Fahri, E., Werning, I., 2014. Dilemma not trilemma? Capital controls and exchange rates with volatile capital flows. IMF Economic Review (Special Volume in Honor of Stanley Fischer). 62. pp. 569–605.
Faust, J., 1998. The robustness of identified VAR conclusions about money. Carnegie Rochester Conference Series on Public Policy. 49. pp. 207–244.
Galí, J., 2015. Monetary Policy, Inflation and the Business Cycle: An Introduction to the New Keynesian Framework and Its Applications. Second Edition, Princeton University Press.
Gertler, M., Karadi, P., 2015. Monetary policy surprises, credit costs and economic activity. Am. Econ. J. Macroecon. 7 (1), 44–76.
Georgiadis, G., 2016. Determinants of global spillovers from US monetary policy. J. Int. Money Financ. 67 (C), 41–61.
Giannone, D., Lenza, M., Primiceri, G., 2015. Prior selection for vector autoregressions. Rev. Econ. Stat. 97 (2), 436–451.
Gürkaynak, R., Sack, B., Swanson, E., 2005. Do actions speak louder than words? The response of asset prices to monetary policy actions and statements. Int. J. Cent. Bank. 1 (1), 55–93.
Kilian, L., Kim, Y.J., 2011. How reliable are local projection estimators of impulse responses? Rev. Econ. Stat. 93 (4), 1460–1466.
Kim, S., 2001. International transmission of U.S. monetary policy shocks: evidence from VAR's. J. Monet. Econ. 48 (2), 339–372.
Klein, M.W., Shambaugh, J.C., 2010. Exchange Rate Regimes in the Modern Era. The MITPress, Boston.
Lane, P., Shambaugh, J.C., 2010. Financial exchange rates and international currency exposures. Am. Econ. Rev. 100 (1), 518–540.
Litterman, R.B., 1986. Forecasting with Bayesian vector autoregressions — five years of experience. J. Bus. Econ. Stat. 4 (1), 25–38.
Mackowiak, B., 2007. External shocks, U.S. monetary policy and macroeconomic fluctuations in emerging markets. J. Monet. Econ. 54 (8), 2512–2520.
Miniane, J., Rogers, J., 2007. Capital controls and the international transmission of U.S. money shocks. J. Money Credit Bank. 39 (5), 1003–1035.
Miranda-Agrippino, S., Rey, H., 2015. World asset markets and the global financial cycle. NBER Working Paper No. 21722.
Ostry, J.D., Ghosh, A.R., 2016. On the obstacles to international policy coordination and how to overcome them. J. Int. Money Financ. 67 (C), 25–40.
Pesaran, M.H., Smith, R., 1995. Estimating long-run relationships from dynamic heterogeneous panels. J. Econ. 68 (5), 79–113.
Ramey, V., 2016. Macroeconomic shocks and their propagation. NBER Working Paper 21978.
Rey, H., 2013. Dilemma not trilemma: the global financial cycle and monetary policy independence. Federal Reserve Bank of Kansas City Economic Policy Symposium,
Romer, C.D., Romer, D.H., 2004. A new measure of monetary shocks: derivation and implications. Am. Econ. Rev. 94 (4), 1055–1084.
Rose, A.K., 2011. Exchange rate regimes in the modern era: fixed, floating, and flaky. J. Econ. Lit. 49 (3), 652–672.
Shambaugh, J.C., 2004. The effect of fixed exchange rates on monetary policy. Q. J. Econ. 119 (1), 301–352.
Sims, C.A., Zha, T., 1998. Bayesian methods for dynamic multivariate models. Int. Econ. Rev. 39 (4), 949–968.
Uhlig, H., 2005. What are the effects of monetary policy on output? Results from an agnostic identification procedure. J. Monet. Econ. 52 (2), 381–419.
Woodford, M., 2007. Globalisation and monetary control. NBER Working Paper No. 13329.