Spillovers from US monetary policy: evidence from a time varying parameter global vector auto-regressive model

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Summary. The paper develops a global vector auto-regressive model with time varying parameters and stochastic volatility to analyse whether international spillovers of US monetary policy have changed over time. The model proposed enables us to assess whether coefficients evolve gradually over time or are better characterized by infrequent, but large, breaks. Our findings point towards pronounced changes in the international transmission of US monetary policy throughout the sample period, especially so for the reaction of international output, equity prices and exchange rates against the US dollar. In general, the strength of spillovers has weakened in the aftermath of the global financial crisis. Using simple panel regressions, we link the variation in international responses to measures of trade and financial globalization. We find that a broad trade base and a high degree of financial integration with the world economy tend to cushion risks stemming from a foreign shock such as US tightening of monetary policy, whereas a reduction in trade barriers and/or a liberalization of the capital account increase these risks.

Keywords: Globalization; Mixture innovation models; Spillovers; Zero lower bound

1. Introduction

Economists and policy makers have extensively argued about the implications of globalization for the design and conduct of monetary policy. Globalization has rendered monetary policy more complex. As former Federal Reserve Bank Chairman Ben Bernanke (Bernanke, 2007) noted,

'...effective monetary policy making now requires taking into account a diverse set of global influences, many of which are not fully understood'.

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Globalization has also laid the ground for the propagation of spillovers from one country to the rest of the world. Recently, international spillovers from monetary policy actions of one country to other economies have been dubbed 'a corollary of globalization' by Vice-President of the European Central Bank Vítor Constâncio (Constâncio, 2015).

With both trade and financial globalization on the rise, the main objective of this paper is to assess whether spillovers are currently different from those in the past and whether such differences can be linked to changes in globalization. These questions have received relatively little attention in the empirical literature on spillovers (for an exception, see Kamin (2013)). This is due to two reasons. First, the necessity to model several countries simultaneously gives rise to additional challenges involved in estimation and model specification. Second, a potentially large model of the world economy which accounts for changing spillovers needs to be able to accommodate movements in its coefficients. This, however, turns out to be computationally challenging by using standard econometric tools.

We propose a new econometric model that extends the global vector auto-regressive (GVAR) model put forth in Pesaran *et al*. (2004) to allow for movements in regression coefficients and error variances. To infer whether parameters change gradually or feature sudden breaks, we adapt recent techniques that were proposed in Huber *et al*. (2018) to the GVAR context. The resulting time varying parameter (TVP) GVAR model with mixture innovations is a flexible framework that enables estimation of global spillovers from a US monetary policy shock that potentially differ for each point in time in our observation sample.

The existence of significant spillovers from US monetary policy has been demonstrated in a range of empirical studies (see, among others, Kim (2001), Canova (2005), Dees *et al*. (2007) and Feldkircher and Huber (2016)). A consensus has also emerged concerning the fact that monetary policy and its transmission in the USA have changed over recent decades (Sims and Zha, 2006; Boivin *et al*., 2010; Boivin, 2006; Baumeister and Benati, 2013). As pointed out in Boivin *et al*. (2010), this could be driven by several factors, including regulatory changes as well as shifts in domestic macroeconomic and financial market conditions. Our paper thus contributes to the literature on asymmetric effects of US monetary policy depending on domestic economic conditions. There is also a related literature that examines asymmetry depending on whether monetary policy is tightened or loosened. Focusing on asset prices, this possibility has been examined among others in Kuttner (2001) and Rogers *et al*. (2014), who both found little evidence of asymmetry by using US data. In addition, there could be global drivers that determine effects of monetary policy, such as a global financial cycle that was proposed by Rey (2015) or more generally the degree of trade and financial globalization. Georgiadis and Mehl (2016) examined the relationship between monetary policy effectiveness—measured as the reaction of output to an unexpected change in the policy rate—and financial globalization. They found that a fall in a country's net foreign asset position in response to a monetary tightening strengthens *domestic* monetary policy effectiveness and that this 'valuation effect' offsets a dampening effect caused by the existence of a US-led global financial cycle—as argued in Bekaert *et al*. (2013) and Rey (2015). Considering this argument in the context of spillovers from US monetary policy, financial globalization would be expected to dampen spillovers from a US rate hike since the accompanying appreciation of the US dollar strengthens other countries' (dollar-held) asset positions. Rey (2015) and Miranda-Agrippino and Rey (2015), by contrast, stressed the importance of a global financial cycle and financial variables in general for the international propagation of macroeconomic shocks.

In this paper we ask two questions. First, do spillovers of US monetary policy shocks vary over time? And, second, what is the contribution of trade and financial globalization in determining the size of the international effects? As stated above, the model that we propose is capable of

answering the first question by allowing for movements in the coefficients that can be gradual or abrupt. This is of ample importance given the research question and sample period under study which features a rapid decrease of interest rates followed by a prolonged period of no interest rate changes (zero lower bound) and a gradual increase thereafter.

Our results can be summarized as follows. First, a contractionary shock to US monetary policy tends to imply a persistent global contraction in real activity and a drop in international consumer and equity prices. Also, currencies tend to depreciate against the US dollar. Second, for several variables, we find evidence for considerable time variation: spillovers to international output, exchange rates and equity prices have been stronger in the period before the global financial crisis. Last, we find that both trade and financial globalization can explain variation in the strength of spillovers. A broad trade base and a high degree of financial integration with the world economy cushion spillovers stemming from a tightening of US monetary policy, whereas a reduction of trade barriers and/or a liberalization of the capital account increase them.

The paper is structured as follows. Section 2 presents the econometric framework, including a detailed discussion on the novel mixture innovation specification that is adopted. Section 3 presents the data, whereas Section 4 discusses the results. Finally, Section 5 concludes the paper and a technical appendix provides information on the Bayesian estimation strategy and the prior specifications which makes estimation of the model feasible.

The data that are analysed in the paper and the programs that were used to analyse them can be obtained from

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https://rss.onlinelibrary.wiley.com/hub/journal/1467985x/series-a-
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2. Econometric framework

To assess the dynamic transmission mechanism between US monetary policy and the global economy, we develop a GVAR model featuring TVPs and stochastic volatility (SV). The TVP–SV–GVAR model is estimated by using a broad panel of countries and macroeconomic aggregates, thus providing a truly global and flexible representation of the world economy. In general, the structure of a GVAR model implies two distinct stages in the estimation process. In the first stage, $N+1$ country-specific multivariate time series models are specified, each of them including exogenous regressors that aim to capture cross-country linkages. In the second stage, these models are combined by using country weights to form a global model that is used to carry out impulse response analysis or forecasting.

2.1. A dynamic global macroeconomic model

Let the endogenous variables $y_{ij,t}$ $(j=1,\ldots,k_i)$ for country $i=0,\ldots,N$ be contained in a $k_i \times 1$ vector $\mathbf{y}_{it}=(y_{i1,t},\ldots,y_{ik_i,t})'$. In addition, all country-specific models feature a set of k_i^* weakly exogenous regressors $\mathbf{y}_{it}^*=(y_{i1,t}^*,\ldots,y_{ik_i,t}^*)'$, constructed as weighted averages of the endogenous variables in other economies:

$$y_{ij,t}^* = \sum_{c=0}^{N} w_{ic} y_{cj,t} \qquad \text{for } j=1,\ldots,k_i^*. \qquad (2.1)$$

Here, w_{ic} is the weight corresponding to the jth variable of country c in country i's specification. These weights are typically assumed to be related to bilateral trade exposure and sum to 1, and $w_{ii}=0$ for all i. In line with the bulk of the literature on GVAR modelling, we assume that all variables and countries are linked by the same set of weights which is fixed over time (Dees *et al.*,

2007). It could be argued that considering time varying weights would be an alternative way to model time variation within the GVAR framework. However, whereas this strategy would affect only the set of weakly exogenous variables, the TVP–SV–GVAR model proposed allows for time variation in all coefficients as well as changes in residual variances and is thus capable of modelling a much richer set of dynamics at the international level. Moreover, in the empirical application we are not interested in interpreting particular coefficients; rather we are interested in whether spillovers change over time leaving it open whether these changes are driven by changes in the economic relationship between countries or by changes how these countries react to foreign factors.

We deviate from existing GVAR modelling efforts by specifying country-specific VAR models that feature exogenous regressors, TVPs and SV, so that

$$\mathbf{y}_{it} = \sum_{p=1}^{P} \mathbf{B}_{ip,t}\mathbf{y}_{it-p} + \sum_{q=0}^{Q} \mathbf{\Lambda}_{iq,t}\mathbf{y}_{it-q}^{*} + \mathbf{u}_{it}. \tag{2.2}$$

Here the parameters are as follows.

(a) $\mathbf{B}_{ip,t}$ $(p=1,\ldots,P)$ is a $k_i \times k_i$ matrix of coefficients that are associated with the lagged endogenous variables.
(b) $\mathbf{\Lambda}_{iq,t}$ $(q=0,\ldots,Q)$ denotes a $(k_i \times k_i^{*})$-dimensional coefficient matrix corresponding to the k_i^{*} weakly exogenous variables in $\mathbf{y}_{it}^{*}$.
(c) $\mathbf{u}_{it} \sim \mathcal{N}(\mathbf{0}, \mathbf{\Sigma}_{it})$ is a heteroscedastic vector error term with

$$\mathbf{\Sigma}_{it} = \mathbf{A}_{i0,t}^{-1}\mathbf{D}_{it}(\mathbf{A}_{i0,t}^{-1})'. \tag{2.3}$$

We let $\mathbf{D}_{it} = \mathrm{diag}(\lambda_{i0,t},\ldots,\lambda_{ik_i,t})$ be a diagonal matrix and $\mathbf{A}_{i0,t}^{-1}$ denotes a $k_i \times k_i$ lower unitriangular matrix of covariance parameters that establishes contemporaneous relationships between the shocks in $\mathbf{u}_{it}$. Note that $\mathbf{u}_{it} = \mathbf{A}_{i0,t}^{-1}\varepsilon_{it}$, where ε_{it} is a Gaussian vector white noise process with zero mean and variance–covariance matrix $\mathbf{D}_{it}$.
(d) The variances $\lambda_{il,t}$ are assumed to follow a stationary auto-regressive process,

$$\log(\lambda_{il,t}) = \mu_{il} + \rho_{il}\{\log(\lambda_{il,t-1}) - \mu_{il}\} + \upsilon_{il,t}, \qquad \upsilon_{il,t} \sim \mathcal{N}(0, \varsigma_{il}^2), \tag{2.4}$$

where μ_{il} denotes the unconditional expectation of the log-volatility, ρ_{il} the corresponding persistence parameter and ς_{il}^2 is the innovation variance of the process.

The set of $N+1$ country-specific models can be linked to yield a global VAR model (Pesaran *et al.*, 2004). Collecting all contemporaneous terms of equation (2.2) and defining a $(k_i + k_i^{*})$-dimensional vector $\mathbf{z}_{it} = (\mathbf{y}_{it}', \mathbf{y}_{it}^{*\prime})'$, we obtain

$$\mathbf{C}_{it}\mathbf{z}_{it} = \sum_{\mathfrak{s}=1}^{\mathfrak{S}} \mathbf{L}_{i\mathfrak{s},t}\mathbf{z}_{it-\mathfrak{s}} + \mathbf{u}_{it}, \tag{2.5}$$

with $\mathbf{C}_{it} = (\mathbf{I}_{k_i}, -\mathbf{\Lambda}_{i0,t})$, $\mathbf{L}_{i\mathfrak{s},t} = (\mathbf{B}_{i\mathfrak{s},t}, \mathbf{\Lambda}_{i\mathfrak{s},t})$ and $\mathfrak{S} = \max(P,Q)$. A global vector $\mathbf{y}_t = (\mathbf{y}_{0t}', \ldots, \mathbf{y}_{Nt}')'$ of dimension $k = \Sigma_{i=0}^{N} k_i$ and a corresponding country-specific link matrix $\mathbf{W}_i$ $(i=1,\ldots,N)$ of dimension $(k_i + k_i^{*}) \times k$ can be defined such that equation (2.5) can be rewritten exclusively in terms of the global vector:

$$\mathbf{C}_{it}\mathbf{W}_i\mathbf{y}_t = \sum_{\mathfrak{s}=1}^{\mathfrak{S}} \mathbf{L}_{i\mathfrak{s},t}\mathbf{W}_i\mathbf{y}_{t-\mathfrak{s}} + \mathbf{u}_{it}. \tag{2.6}$$

Stacking the equations $N + 1$ times yields

$$\mathbf{G}_t \mathbf{y}_t = \sum_{\mathfrak{s}=1}^{\mathfrak{S}} \mathbf{F}_{\mathfrak{s}t} \mathbf{y}_{t-\mathfrak{s}} + \mathbf{u}_t, \tag{2.7}$$

where $\mathbf{G}_t = ((\mathbf{C}_{0t}\mathbf{W}_0)', \dots, (\mathbf{C}_{Nt}\mathbf{W}_N)')'$ and $\mathbf{F}_{\mathfrak{s}t} = ((\mathbf{L}_{0\mathfrak{s},t}\mathbf{W}_0)', \dots, (\mathbf{L}_{N\mathfrak{s},t}\mathbf{W}_N)')'$ denote stacked coefficient matrices. The error term $\mathbf{u}_t = (\mathbf{u}_{0t}', \dots, \mathbf{u}_{Nt}')'$ is normally distributed with mean 0 and block diagonal variance–covariance matrix $\mathbf{H}_t = \text{diag}(\boldsymbol{\Sigma}_{0t}, \dots, \boldsymbol{\Sigma}_{Nt})$. Equation (2.7) resembles a (very) large VAR model with drifting coefficients which, notwithstanding the problems that are associated with the high dimensionality of the parameter vector, can be estimated by using Bayesian techniques that have been developed to deal with multivariate linear models with TVPs.

2.2. Modelling time variation in the regression coefficients

Up to this point, we have remained silent on the specific law of motion for the coefficients in the model. Since the number of parameters is typically large relative to the length of the sample T, a parsimonious way of modelling time variation is necessary to obtain precise estimates and to avoid overfitting.

Stacking the lagged endogenous and weakly exogenous variables in an m_i-dimensional vector, with $m_i = k_i P + k_i^*(Q + 1)$,

$$\mathbf{x}_{it} = (\mathbf{y}_{it-1}', \dots, \mathbf{y}_{it-P}', (\mathbf{y}_{it}^*)', \dots, (\mathbf{y}_{it-Q}^*)')' \tag{2.8}$$

and collecting all regression coefficients in a $k_i \times (m_i k_i)$ matrix,

$$\boldsymbol{\Psi}_{it} = (\mathbf{B}_{i1,t}, \dots, \mathbf{B}_{iP,t}, \boldsymbol{\Lambda}_{i0,t}, \dots, \boldsymbol{\Lambda}_{iQ,t})', \tag{2.9}$$

allows us to rewrite equation (2.2) as

$$\mathbf{y}_{it} = (\mathbf{I}_{k_i} \otimes \mathbf{x}_{it}')\boldsymbol{\psi}_{it} + \mathbf{u}_{it}. \tag{2.10}$$

For convenience, define $\boldsymbol{\psi}_{it} = \text{vec}(\boldsymbol{\Psi}_{it})$ and collect the free covariance parameters in $\mathbf{A}_{i0,t}$ in an $l_i = k_i(k_i - 1)/2$-dimensional vector $\mathbf{a}_{i0,t}$. For each individual coefficient in $\boldsymbol{\xi}_{it} = (\mathbf{a}_{i0,t}', \boldsymbol{\psi}_{it}')'$, we assume a random walk law of motion,

$$\xi_{ij,t} = \xi_{ij,t-1} + \eta_{ij,t}, \qquad \text{for } j = 1, \dots, s_i, \tag{2.11}$$

where $s_i = l_i + k_i(m_i k_i)$ and $\eta_{ij,t}$ denotes a white noise shock with time varying variance $\vartheta_{ij,t}$.

In principle, allowing all coefficients of the model to move freely yields a highly parameterized model that is prone to overfitting. This issue is intensified in the context of a multicountry GVAR model, calling for some form of regularization of the variation in the parameters over time. To achieve this, we follow Huber *et al.* (2018) and assume that $\vartheta_{ij,t}$ evolves according to

$$\vartheta_{ij,t} = (1 - d_{ij,t})\vartheta_{ij,0} + d_{ij,t}\vartheta_{ij,1}, \tag{2.12}$$

whereby $\vartheta_{ij,1} \gg \vartheta_{ij,0}$ and $\vartheta_{ij,0}$ is set close to 0. In this paper we follow Huber *et al.* (2018) and set $\vartheta_{ij,0} = 10^{-5}\hat{\sigma}_{ij}$, with $\hat{\sigma}_{ij}$ denoting the ordinary least squares standard deviation of a time invariant VAR model. Moreover, let $d_{ij,t}$ denote a binary random variable that follows an independent Bernoulli distribution with

$$d_{ij,t} = \begin{cases} 1 & \text{with probability } p_{ij}, \\ 0 & \text{with probability } 1 - p_{ij}. \end{cases} \tag{2.13}$$

This specification is commonly referred to as a mixture innovation model (Giordani and Kohn, 2008; Koop *et al.*, 2009) and nests a wide variety of competing models. For instance, if $d_{ij,t} = 1$

for all t, we obtain a standard TVP specification whereas, in the case of $d_{ij,t} = 0$ for all t, we end up having a nearly constant parameter specification (as the variance of $\eta_{ij,t}$ will be relatively small). Cases in between are also possible, implying that our framework flexibly accommodates situations where parameters might be time varying during certain intervals of time, while being effectively constant during other periods. Especially in the context of GVAR models, selecting appropriate model features *a priori* is a daunting task given the high dimensionality of the parameter space. Our approach avoids this by effectively selecting data-based restrictions on the law of motion of each coefficient separately.

Estimation of such a mixture innovation model would be unfeasible given the high dimensionality of the parameter space of the GVAR model. Hence, following Huber *et al.* (2018), we approximate the latent indicators $d_{ij,t}$ by proposing a simple thresholding rule during Markov chain Monte Carlo (MCMC) sampling. More specifically, the lth draw of $d_{ij,t}$ is approximated through

$$d_{ij,t}^{(l)} = \begin{cases} 1 & \text{if } |\Delta\xi_{ij,t}^{(l)}| > c_{ij}^{(l-1)}, \\ 0 & \text{if } |\Delta\xi_{ij,t}^{(l)}| \leqslant c_{ij}^{(l-1)}, \end{cases} \tag{2.14}$$

with $|\Delta\xi_{ij,t}^{(l)}|$ and $c_{ij}^{(l-1)}$ denoting draws of the (time varying) coefficients and of a latent threshold c_{ij} respectively. This approximation captures the notion that, if the period-on-period change in the respective parameter is large, the unconditional probability (after integrating $\xi_{ij,t}$ out) that $d_{ij,t} = 1$ is also large. The key advantage of this approach is its computational simplicity. Compared with standard TVP–VAR models in the spirit of Cogley *et al.* (2005) and Primiceri (2005), the computational burden is increased only slightly, whereas the model is much more flexible. Relative to standard mixture innovation models, our approach avoids Kalman-filter-based algorithms to infer the full history of the indicators altogether.

This coefficient-specific law of motion for the regression parameters enables us to investigate changes in the domestic and international transmission mechanisms rigorously. Moreover, we account for heteroscedasticity by making the country-specific variance–covariance matrix of $\mathbf{u}_{it}$ time varying. This is to ensure that changes in the parameters reflect changes in the underlying macroeconomic relationships and are not confounded by a wrongly assumed constant error variance. Our model captures a range of properties that are essential to assess changes in domestic and international transmission mechanisms of monetary policy shocks and can accommodate important features which are commonly observed in macroeconomic and financial time series.

We use Bayesian methods to carry out inference in the model that was proposed above. Given the risk of overparameterization that is inherent in the specification that is used, we rely on Bayesian shrinkage methods to achieve a simpler representation of the data. The time varying nature of the parameters in the model and the presence of the weakly exogenous variables in equation (2.2) present further complications that are tackled in the estimation procedure. More details on the exact prior specification and the proposed MCMC algorithm as well as convergence criteria are detailed in Appendix A. Here, it suffices to note that we repeat the algorithm that is outlined in Appendix A 40000 times, where the first 30000 draws are discarded. From the draws retained, we single out unstable draws, which gives us a final sample of 500 posterior draws on which inference is based.

3. Data and model specification

This section introduces the data and provides details on the specification of the model. We use quarterly data for 35 countries spanning the period from 1990, first quarter, to 2016, fourth quarter. The countries that are covered in our sample are shown in Table 1.

Table 1.   Country coverage of the GVAR model†

European countries	Other developed economies	Emerging Asian countries	Latin American countries	Middle East and African countries
Austria (AT)	Australia (AU)	*China (CN)*	*Argentina (AR)*	Turkey (TR)
Belgium (BE)	*Canada (CA)*	*India (IN)*	*Brazil (BR)*	Saudi Arabia (SA)
Germany (DE)	*Japan (JP)*	Indonesia (ID)	*Chile (CL)*	South Africa (ZA)
Spain (ES)	*USA (US)*	Malaysia (MY)	*Mexico (MX)*	
Finland (FI)		*Korea (KR)*	*Peru (PE)*	
France (FR)		Philippines (PH)		
Greece (GR)		Singapore (SG)		
Italy (IT)		Thailand (TH)		
Netherlands (NL)				
Portugal (PT)				
Denmark (DK)				
Great Britain (GB)				
Switzerland (CH)				
Norway (NO)				
Sweden (SE)				

†Country codes are in parentheses. Empirical results are shown for the countries in italics.

The country-specific models include real gross domestic product growth Δgdp, inflation Δcp measured by the log-difference of the consumer price level and the log-difference of the nominal exchange rate, Δer, *vis-à-vis* the US dollar, with an increase denoting an appreciation of the dollar. We include (3-months) short-term nominal interest rates ir in all economies, except for euro area countries, Great Britain, Japan and the USA. For these countries, we use shadow interest rates instead, since, in these economies and over the time period that is covered, interest rates stayed at the zero lower bound for considerable time. We use the shadow rates of Krippner (2013), which are publicly available from `https://www.rbnz.govt.nz/research-and-publications/research-programme/additional-research/measures-of-the-stance-of-united-states-monetary-policy/comparison-of-international-monetary-policy-measures`. These standard macroeconomic data are augmented by financial variables to take into account their potential role as shock propagators (Rey, 2015; Miranda-Agrippino and Rey, 2015). Specifically, we include the term spread sp, constructed as the difference between 10-year government bond yields and short-term interest rates, and changes in stock market prices, Δeq. Not all variables are available for each of the countries that we consider in this study. This concerns mostly long-term interest rates (that are used to calculate the term spread) and equity prices.

The vector of *domestic* variables for a typical country i is given by

$$\mathbf{y}_{it} = (\Delta \text{gdp}_{it}, \Delta \text{cp}_{it}, \text{ir}_{it}, \text{sp}_{it}, \Delta \text{er}_{it}, \Delta \text{eq}_{it})'. \tag{3.1}$$

We follow the bulk of the literature on GVAR modelling by including changes in oil prices, Δpoil, as a global control variable. With the exception of exchange rates, we construct foreign counterparts for all domestic variables. The weights to calculate foreign variables are based on average bilateral annual trade flows in the period from 2000 to 2014. Recent contributions (Eickmeier and Ng, 2015; Dovern and van Roye, 2014) suggest the use of financial data to compute foreign variables that are related to the financial side of the economy (e.g. interest rates or credit volumes). However, reliable data on financial flows—such as portfolio flows or foreign direct investment—are not available for the country coverage that we consider in this study.

See the appendix of Feldkircher and Huber (2016) for the results of a sensitivity analysis with respect to the choice of weights in Bayesian GVAR specifications in the framework of models with fixed parameters. For a typical country i the set of *weakly exogenous* and global control variables comprises

$$\mathbf{y}_{it}^* = (\Delta\text{gdp}_{it}^*, \Delta\text{cp}_{it}^*, \text{ir}_{it}^*, \text{sp}_{it}^*, \Delta\text{poil})'. \tag{3.2}$$

The US model, which we normalize to correspond to $i=0$, deviates from the other country specifications in that oil price inflation is determined within that country model, and the change in the trade-weighted exchange rate, Δer^*, is included as an additional control variable, so that the vector of endogenous and weakly exogenous variables for the USA is given by

$$\mathbf{y}_{0t} = (\Delta\text{gdp}_{0t}, \Delta\text{cp}_{0t}, \text{ir}_{0t}, \text{sp}_{0t}, \Delta\text{eq}_{0t}, \Delta\text{poil}_t)', \tag{3.3}$$

$$\mathbf{y}_{0t}^* = (\Delta\text{gdp}_{0t}^*, \Delta\text{cp}_{0t}^*, \Delta\text{er}_{0t}^*, i_{0t}^*, \text{sp}_{0t}^*, \Delta\text{eq}_{0t}^*)'. \tag{3.4}$$

Finally, for all countries considered, we set the lag length of endogenous and weakly exogenous variables equal to a quarter. Despite the parsimonious lag structure, the model adequately captures the serial correlation of the underlying data. Fig. 10 in Appendix A.3 provides evidence on the lack of serial dependence of the residuals. In Fig. 10, we further show evidence of convergence of the MCMC algorithm, the distribution of trade weights and evidence of weak cross-country correlation of the residuals.

3.1. Structural identification

In this paper, we consider structural generalized impulse response functions (see Koop *et al.* (1996) and Pesaran and Shin (1998)) to trace the global effects of a US monetary policy shock. In the GVAR framework, using structural generalized impulse responses proves to be a standard choice since identifiyng all k shocks is usually unfeasible. Moreover, since we are interested in only the causal effects of a US-based monetary policy shock, identifying the remaining shocks in the system is not necessary.

To identify the monetary policy shock, we follow Dees *et al.* (2007), Eickmeier and Ng (2015) and Feldkircher and Huber (2016) and adopt sign restrictions imposed on the contemporaneous responses of the US macroeconomic quantities. This implies that the reactions of $\mathbf{y}_t$ to the US monetary policy shock coincide with the structural impulse responses, whereas responses of $\mathbf{y}_t$ to shocks outside the US country model are generalized impulse responses (for a detailed discussion, see Dees *et al.* (2007)).

For simplicity, we assume that the US model is indexed by $i=0$. Introducing a $k_0 \times k_0$ matrix $\mathbf{R}_{0t}$ (with $\mathbf{R}_{0t}\mathbf{R}_{0t}' = \mathbf{I}_{k_0}$) and multiplying equation (2.2) from the left with $\tilde{\mathbf{A}}_{00,t} = \mathbf{R}_{0t}\mathbf{A}_{00,t}$ yields

$$\tilde{\mathbf{A}}_{00,t}\mathbf{y}_{it} = \sum_{p=1}^{P} \tilde{\mathbf{B}}_{0p,t}\mathbf{y}_{it-p} + \sum_{q=0}^{Q} \tilde{\mathbf{\Lambda}}_{iq,t}\mathbf{y}_{it-q}^* + \tilde{\mathbf{R}}_{i0,t}\varepsilon_{it}, \tag{3.5}$$

with $\tilde{\mathbf{B}}_{0p,t} = \mathbf{R}_{0t}\mathbf{A}_{00,t}\mathbf{B}_{0p,t}$ and $\tilde{\mathbf{\Lambda}}_{iq,t} = \mathbf{R}_{0t}\mathbf{A}_{00,t}\mathbf{\Lambda}_{iq,t}$. Note that the introduction of the rotation matrix $\mathbf{R}_{0t}$ leaves the likelihood function untouched.

Traditional sign restrictions are implemented by simulating rotation matrices $\mathbf{R}_{i0,t}$, computing the corresponding structural impulse responses, and if a set of restrictions is fulfilled the associated rotation matrix is kept. We implement this approach by using the algorithm outlined in Rubio-Ramírez *et al.* (2010). For each rotation, we construct a $(k \times k)$-dimensional matrix $\mathbf{R}_t$ that features $\mathbf{R}_{0t}$ in the first $k_0 \times k_0$ block and equals an identity matrix elsewhere. More specifically, $\mathbf{R}_t$ is given by

$$\mathbf{R}_t = \begin{pmatrix} \mathbf{R}_{0t} & 0 & \cdots & 0 \\ 0 & \mathbf{I}_{k_1} & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & \mathbf{I}_{k_N}. \end{pmatrix}. \tag{3.6}$$

This matrix is then used to recover the structural form of the global VAR model.

Before proceeding to the actual sign restrictions that are included, a few words on the specific choice of the rotation matrices are in order. First, consistent with the literature that deals with sign restrictions in GVAR models, the shock is only *locally* identified in the US model. This implies that the structure of the rotation matrix in equation (3.6) is sufficient to identify the impact vector with respect to the US monetary policy shock. In principle, we could also simulate a full $k \times k$ rotation matrix or introduce a separate rotation matrix $\mathbf{R}_{it}$ for each country. However, doing so would increase the computational burden as well as potentially lead to higher estimation uncertainty. Second, since we are exclusively interested in identifying a US-based monetary policy shock, we do not identify additional shocks outside the US country model. Such a modelling strategy would increase the number of restrictions significantly, leading to a situation where finding suitable rotation matrices becomes almost impossible. Third, note that $\mathbf{R}_t$ is time specific. This is a consequence of the fact that the full variance–covariance matrix is time varying, implying that the contemporaneous relationships across shocks are subject to change. Thus, a rotation matrix that fulfils the sign restrictions at time t might not satisfy the restrictions at time $t+1$. To circumvent this issue, we follow the literature and simulate a rotation matrix for each point in our sample (for a recent example, see Gambetti and Musso (2017)). We then assess whether the sign restrictions are fulfilled, in which case we keep the rotation matrix.

We elicit the restrictions on the basis of Feldkircher and Huber (2016) and Peersman (2005). These are imposed on the US country model and are provided in Table 2.

The constraints above are based on a typical aggregate demand-and-supply diagram and are consistent with most dynamic stochastic general equilibrium models. The unexpected rate increase in the USA is assumed to decrease output, consumer price and equity price growth. This assumption is based on empirical evidence for the reaction of stock markets to monetary-policy-induced interest rate changes (Thorbecke, 1997; Rigobon and Sack, 2004; Bernanke and Kuttner, 2005; Li *et al.*, 2010; Rogers *et al.*, 2014).

The identification of monetary policy shocks in a zero lower bound environment deserves some further discussion. As noted above, we use shadow rates instead of actual short-term interest rates as the policy instrument. These are estimated from a term structure model and reflect what short-term rates would have been in the absence of the zero lower bound (see, for example, Krippner (2013)). Hence, shadow rates constitute an overall measure of the monetary policy stance that is equally valid during both normal periods and times where the zero lower

Table 2. Sign restrictions†

Shock	Δy	Δp	i_s	Δeq
Monetary policy	↓	↓	↑	↓
Aggregate demand	↑	↑	↑	—
Aggregate supply	↑	↓	↓	—

†The restrictions are imposed as greater than or equal to or less than or equal to and on impact only.

bound is binding. It could be argued that our results thus blend effects of conventional monetary policy (i.e. interest rate changes) and unconventional monetary policy tools such as quantitative easing, which have been launched in the wake of the global financial crisis. However, since we use a TVP framework (with SV), our analysis enables us to attribute macroeconomic effects of the monetary policy shock to conventional monetary policy during normal times and to unconventional monetary policy during the zero lower bound period (in which the shadow rate becomes negative).

Hence, our econometric framework coupled with a generally valid policy instrument yields a consistent analysis of monetary policy with no need to change the policy instrument or the identification of the shock over different subsamples. To facilitate pinning down the shock of interest, we further identify an aggregate demand-and-supply shock based on standard macroeconomic reasoning (see Feldkircher and Huber (2016) and Peersman (2005)). Note that our assumptions are minimalistic in a sense that they apply to growth rates, are imposed on impact only and are introduced exclusively to the US economy. This is to ensure that our results are not driven by the identifying assumptions. (The early literature on US monetary policy shocks relied heavily on recursive identification, such as in Christiano *et al.* (2005). More recently, some researchers have proposed the use of external instruments, based on either the narrative approach (Romer and Romer, 2004) or high frequency information (Gertler and Karadi, 2015; Rogers *et al.*, 2014, 2018). Miranda-Agrippino and Ricco (2015), however, have shown that using these measures often leads to output and/or price puzzles.)

4. The international dimension of US monetary policy

We start showcasing our model framework by presenting the time variation of two exemplary coefficients with the aim of providing some intuition of the proposed mixture innovation mechanism. In the next step, we briefly investigate how US monetary policy affects international macroeconomic variables. We then move on to assess whether the effects have strengthened or weakened over time. Finally, we relate country characteristics to the extent that the monetary policy shock affects international output.

4.1. *Illustrating our modelling approach*
In this section, we provide additional intuition by considering two examples of TVPs in the framework of our application.
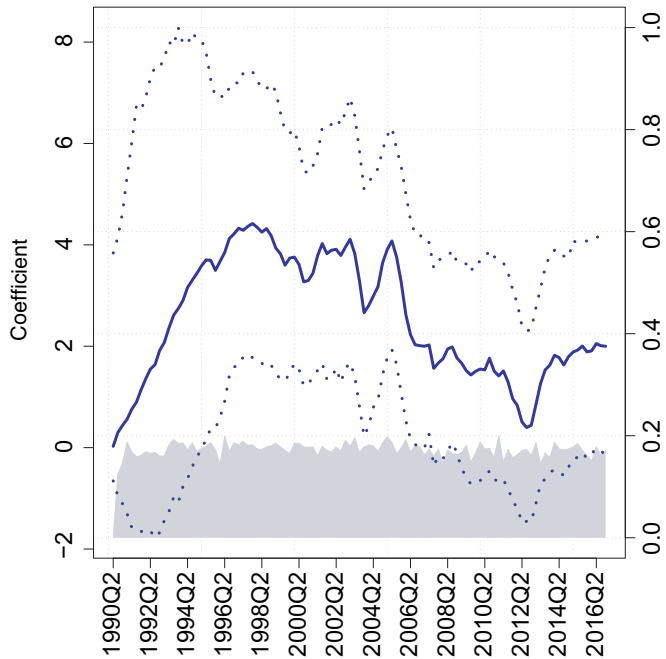
Fig. 1 shows the marginal posterior distribution of the dynamic regression coefficients as well as the probability that a certain coefficient is time varying for a given point in time (the grey shaded area; left-hand scale). The dotted curves refer to the 16th (and 84th) percentiles of the respective posterior distribution whereas the full curve is the posterior median.

Fig. 1(a) shows the evolution of the coefficient that is associated with weakly exogenous term spreads in the output equation for Greece. This plot serves to demonstrate that our flexible specification of the error variance in the state equation enables us to detect situations where coefficients remained approximately constant over a certain time frame (i.e. the period up to the global financial crisis) and then exhibit sudden shifts (during the crisis period). After the shift, Fig. 1(a) suggests that the corresponding coefficient remained approximately constant. The posterior moving probability (the grey shaded area) suggests that, during the crisis, strong evidence in favour of time variation is present whereas in the remaining periods the moving probability is approximately 0.

As a second illustrative example, Fig. 1(b) displays the path of the coefficient on the intercept term of the output equation for Brazil. In contrast with Fig. 1(a), we find that the moving
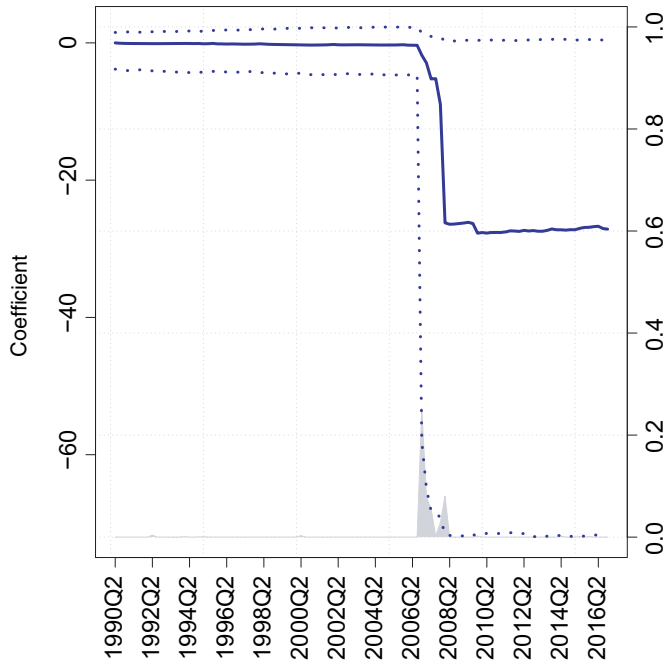
Fig. 1. Illustrative marginal posterior distribution of reduced form coefficients (the plots show the marginal posterior distribution of the dynamic regression coefficients alongside the probability that a given coefficient is time varying at a certain point in time (in grey, left-hand scale): ——, posterior median; ·········, 68% credible intervals): (a) Greece; (b) Brazil

probability is around 0.2 during the estimation period. The corresponding posterior tends to display a rather strong degree of time variation. At first glance, this may seem counterintuitive, since the state innovation variances are pushed to 0 in 80% of the posterior draws. However, it is worth noting that Fig. 1(b) refers to the marginal distribution of the coefficients. These are obtained after integrating out the indicators that control the amount of time variation, effectively leading to a situation where the unconditional variance of the shocks to the regression parameters is non-zero (and potentially moderate).

4.2. Does the global economy respond to US monetary policy shocks?

First, we investigate the international responses to an unexpected tightening of US monetary policy normalized to a 25-basis-point increase in US short-term interest rates (measured by the Krippner shadow rate) throughout the sample period. Whereas the shock on impact is fixed to 25 basis points for the USA, spillovers that are generated by the shock are allowed to vary if macroeconomic relationships change over time. The results are summarized in Fig. 2, which shows posterior medians of time-averaged responses for the largest three countries from each region as defined in Table 1. To provide some information on the behaviour of the whole region, we moreover show credible sets that correspond to regional (time-averaged) responses. These reflect the variation in responses within each country group. All results except those for the short-term rates and the term spread are shown in cumulative terms.

Figs 2(a), 2(f), 2(k) and 2(p) show that US output declines by approximately 0.5% in response to the rate increase. In quantitative terms, this result lies between estimates of linear models that were surveyed in Coibion (2012) and those of Baumeister and Benati (2013) and Feldkircher and Huber (2018), who used a TVP VAR model with SV. Looking at the other countries, output contracts and responses tend to be quite persistent, corroborating the findings by Feldkircher and Huber (2016). Also note that credible sets mostly lie below zero, indicating that responses are (on average viewed over the sample span) statistically significantly different from 0. Considering the different world regions, most responses are very homogeneous and fall inside the credible sets that are spanned by the respective cross-country means. An exception is Argentina which shows a much more pronounced reaction to the monetary tightening in the USA than do its regional peers.

Figs 2(b), 2(g), 2(l) and 2(g) show responses of consumer prices. With the exception of Latin America, all responses and regional credible sets lie below zero. Contractions in consumer prices range between 0.05% and 0.4% in other developed and western European economies. They tend to be more pronounced in emerging Asia and especially so in China. In Latin America, Mexico shows a (modest) positive price response.

Figs 2(c), 2(h), 2(m) and 2(r) depict the (non-cumulative) response of interest rates to the monetary policy shock. Here, the responses tend to differ markedly across regional aggregates. For example, western European countries lower interest rates to provide stimulus and to offset output losses. These responses are tightly estimated and homogeneous within the region. By contrast, short rates in Asian economies (including Japan) appear to display little reactions to the US monetary policy tightening. This could be driven by the comparably low degree of capital account openness. In general, Shambaugh (2004) found that domestic interest rates in countries with a low degree of capital account openness respond less strongly to foreign interest rate changes. By contrast and looking at US monetary policy shocks, Miniane and Rogers (2007) did not find evidence that domestic rates in countries with high capital account openness respond more swiftly compared with those financially more open economies.
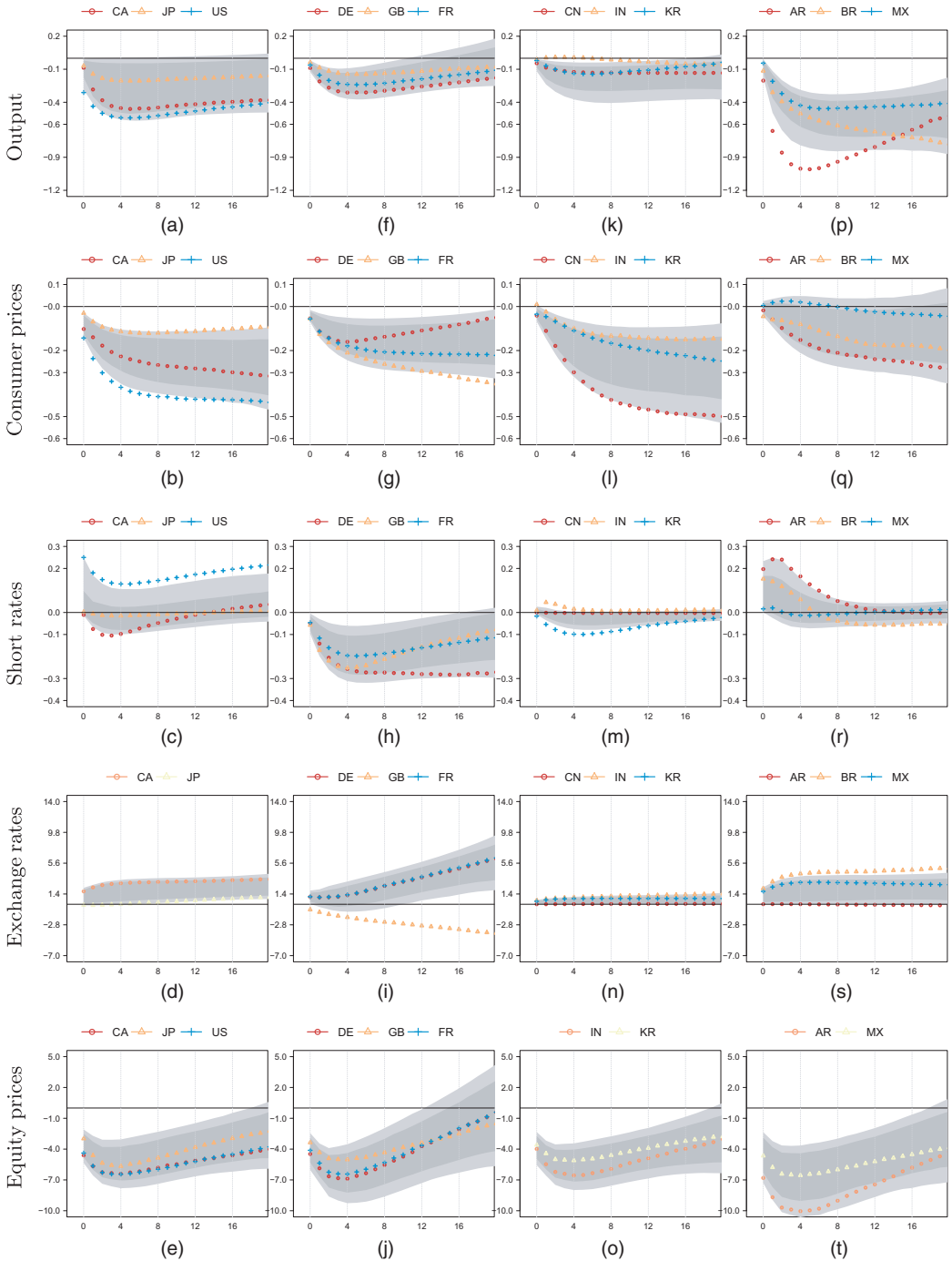
Fig. 2. Responses to a plus 25-basis-point US monetary policy shock (the figure shows the posterior median of time-averaged responses for selected countries; the shaded areas correspond to the 68% (light grey) and 50% (dark grey) credible sets of the regional time-averaged responses; regions are defined as in Table 1 and all responses in cumulative terms except those of short rates): (a)–(e) other developing economics; (f)–(j) western Europe, countries; (k)–(o) emerging Asian countries; (p)–(t) Latin American countries

Figs 2(d), 2(i), 2(n) and 2(s) show the cumulative responses of the exchange rate *vis-à-vis* the US dollar. As expected, responses for countries with a flexible exchange rate regime tend to be positive, indicating a weakening of the respective local currency against the dollar. Advanced economies and Latin American countries respond most strongly to the rate increase, whereas Asian currencies tend to be more insulated—a result which is paralleled in their interest rates responses. Again, this could be driven by the comparably low degree of financial openness since this renders exchange rates less sensitive to foreign rate changes (Kamin, 2013). More specifically, in western Europe, exchange rates depreciate as the interest rate differential widens.

Figs 2(e), 2(j), 2(o) and 2(t) show time-averaged responses of global equity prices. The 25-base-point increase in US rates triggers a 4% decline in US equity prices (on average over the sample period), which is roughly in line with the findings reported in Li *et al.* (2010). As monetary policy is tightened in the USA, equity prices contract world wide. This finding is consistent with Hausman and Wongswan (2011) and Ehrmann and Fratzscher (2009). Responses in other developed economies, western Europe and Asia are very homogeneous, whereas those in Latin America show more variation. For completeness, we show results for term spreads and responses in the figures in the on-line appendix. International term spreads show a homogeneous negative response. They also adjust quickly after the initial decrease. That term spreads behave in a similar fashion could be explained by the high cross-country correlation of short-term rates and bond yields for advanced economies (Kamin, 2013).

Summing up, we find that a US monetary tightening decreases international output, consumer prices and equity prices. International interest rates also respond to the US monetary policy shock, but to a varying degree. The same holds true for exchange rates *vis-à-vis* the dollar. These observations hold on average, viewed over the whole sample period. The estimated effects for the domestic economy are in line with the rich literature on US monetary policy shock. Moreover, it is worth emphasizing that average reactions across real and financial quantities exhibit considerable differences in their shapes, pointing towards heterogeneous timing patterns in the international transmission of US monetary policy shocks. Whereas real quantities generally display a weak immediate reaction, financial quantities such as equity prices tend to display a strong impact response. These results provide confidence in our econometric framework and identification strategy.

4.3. Have spillovers changed over time?

In this section we examine whether spillovers have changed over time. For that, we first construct a simple measure of time variation, namely the robust version of the coefficient of variation, given by

$$\text{CV}_{ij}(\mathfrak{h}) = \frac{\varpi_{ij}(\mathfrak{h})^{75} - \varpi_{ij}(\mathfrak{h})^{25}}{\varpi_{ij}(\mathfrak{h})^{50}}. \tag{4.1}$$

Here, $\varpi_{ij}(\mathfrak{h}) = (\varpi_{ij,1}(\mathfrak{h}), \ldots, \varpi_{ij,T}(\mathfrak{h}))'$ denotes the impulse responses of the ith variable at impulse forecast horizon $\mathfrak{h}$ with respect to the jth structural shock in the system (i.e. the US monetary policy shock) over time. The superscript indicates the 25th, 50th and 75th quantiles of $\varpi_{ij}(\mathfrak{h})$. We compute the marginal posterior distribution of the CV-statistic by using Monte Carlo integration. In what follows we present the posterior median CV and 68% credible set for impact responses ($\mathfrak{h} = 0$) with red bars in Fig. 3 denoting coefficients of variation that are statistically different from 0.

Fig. 3 demonstrates that there is pronounced time variation of impact responses for most variables and countries that are considered in this study. Examining cross-country differences
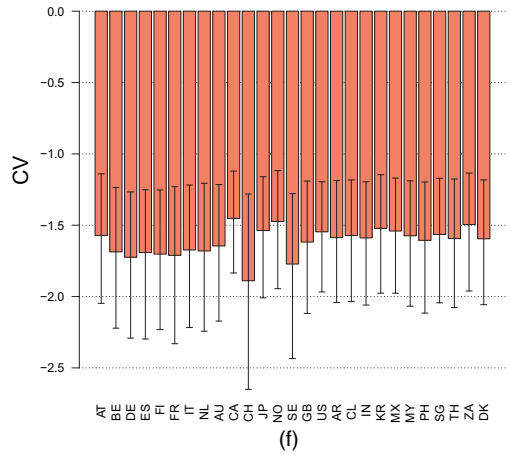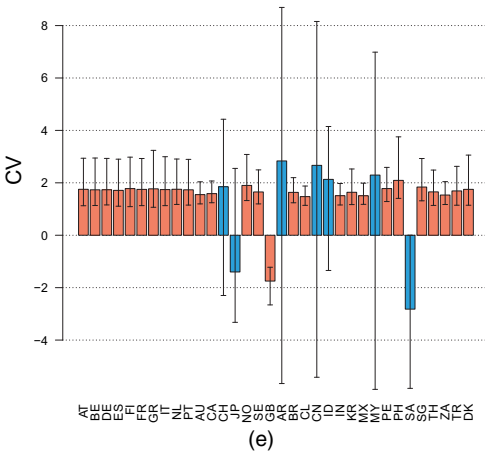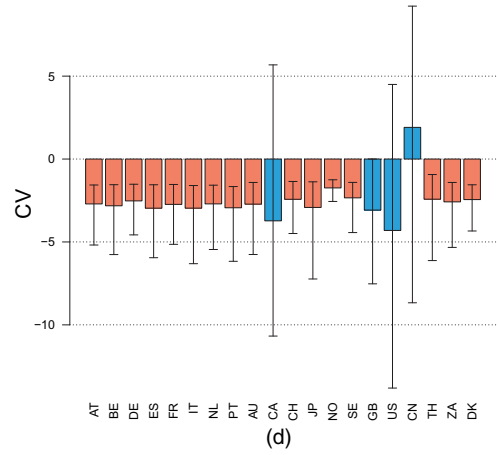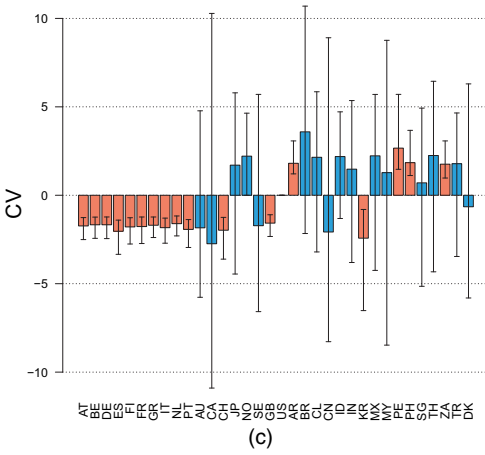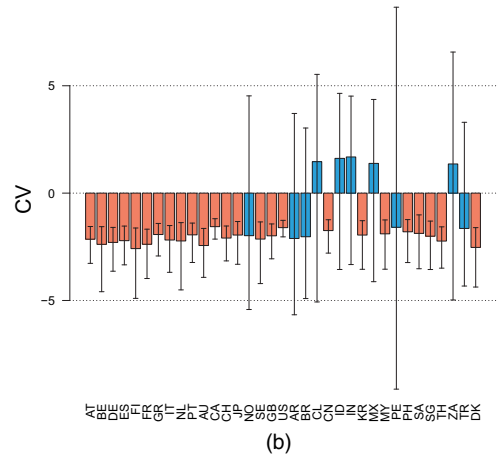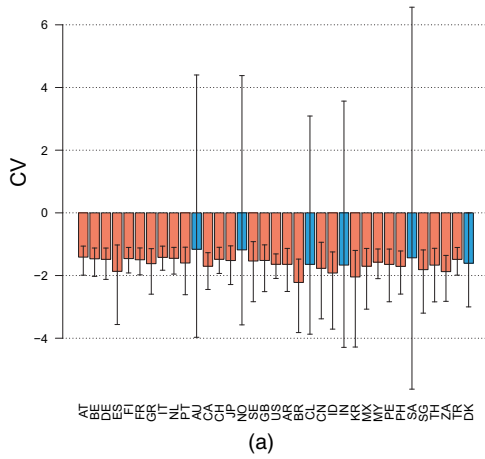
Fig. 3. Coefficient of variation CV of impact responses (the plots show the posterior posterior median of the coefficient of variation of impact responses over time; ⊥, 68% credible intervals; ■, coefficients of variation that are statistically different from 0): (a) output; (b) consumer prices; (c) short rates; (d) term spread; (e) exchange rate; (f) equity prices

reveals that no systematic pattern emerges with respect to which countries exhibit time variation in their impulse response functions. Both advanced as well as emerging countries display strong time variation in their impact responses.

Turning to variable-specific results reveals most time variation for equity prices, output and exchange rates. This is mirrored in the largest number of countries with coefficients of variation that are significantly different from 0. Responses of term spreads also show considerable variation but are available for only a small set of countries. Especially equity impact responses vary considerably. That there is comparably less time variation concerning short-term interest rates implies that central banks adjusted their policy rates in response to the US monetary policy induced output loss in a consistent fashion over the sample period.

Looking at time variation of impact responses provides only an indication of the amount of time variation in international spillovers. Spillovers could also change at longer horizons or, more generally, the shapes of spillover responses could change over time. Hence, we show the full set of output responses in Fig. 4 for selected countries and the regional average over time. Light yellow responses correspond to the beginning of the sample (i.e. 1990, second quarter) and dark red responses to the end of the sample (i.e. 2016, fourth quarter). For brevity we focus on the three variables for which most country responses show time variation, namely equity prices, output and exchange rates. Results for the remaining variables are provided in Appendix A. We also show the posterior median of the time-averaged response along with 68% credible bounds. Responses that fall outside these bounds can be regarded as significantly different from their time average.

Looking at Fig. 4 yields further insights on the amount of time variation in international output responses. In general, we find considerable time variation not only on impact dynamics but also up to a time horizon of 10–15 quarters. In the longer term (up to 20 quarters), however, responses tend to be covered by the credible sets. This might indicate that long-run responses are shaped by fundamental relationships that do not vary that much over time. Taking a regional angle, we see that western European countries that adopted the euro over the sample period show considerable variation in their responses. This carries over to other developed economies, including the US and Latin American economies. By contrast, responses in Asia are generally more time invariant.

Next, we investigate time variation of exchange rate responses that are provided in Fig. 5. Here we observe that responses vary considerably for western European economies but mainly so in the medium to long term, which is in contrast with output responses. Consistent with our previous assessment, currencies in emerging Asia show smaller reactions to the US monetary policy shock. Fig. 5 reveals that this finding holds true throughout the sample period. Latin American currencies, by contrast, show again strong reactions and impulse responses vary throughout the impulse response horizon.

Next, Fig. 6 shows the responses of equity prices. Inspecting the shape of equity price responses yields a very homogeneous picture. Throughout all regions, we observe substantial variation of responses up to 10 quarters. The strongest (i.e. most negative) responses are obtained for the beginning of the sample period. Also the variation of responses over the sample period for a given country is much larger compared with output responses, which indicates a considerable degree of time variation.

Last, we aim to answer the question whether international spillovers have strengthened or weakened over time. For that matter we assess time profiles of international trough output and equity price effects and the peak effect (i.e. maximum depreciation against the US dollar) of international exchange rates. Simple cross-country averages are depicted in Fig. 7.

Note that the exchange rates are on an inverted scale to ease visual comparison with the other
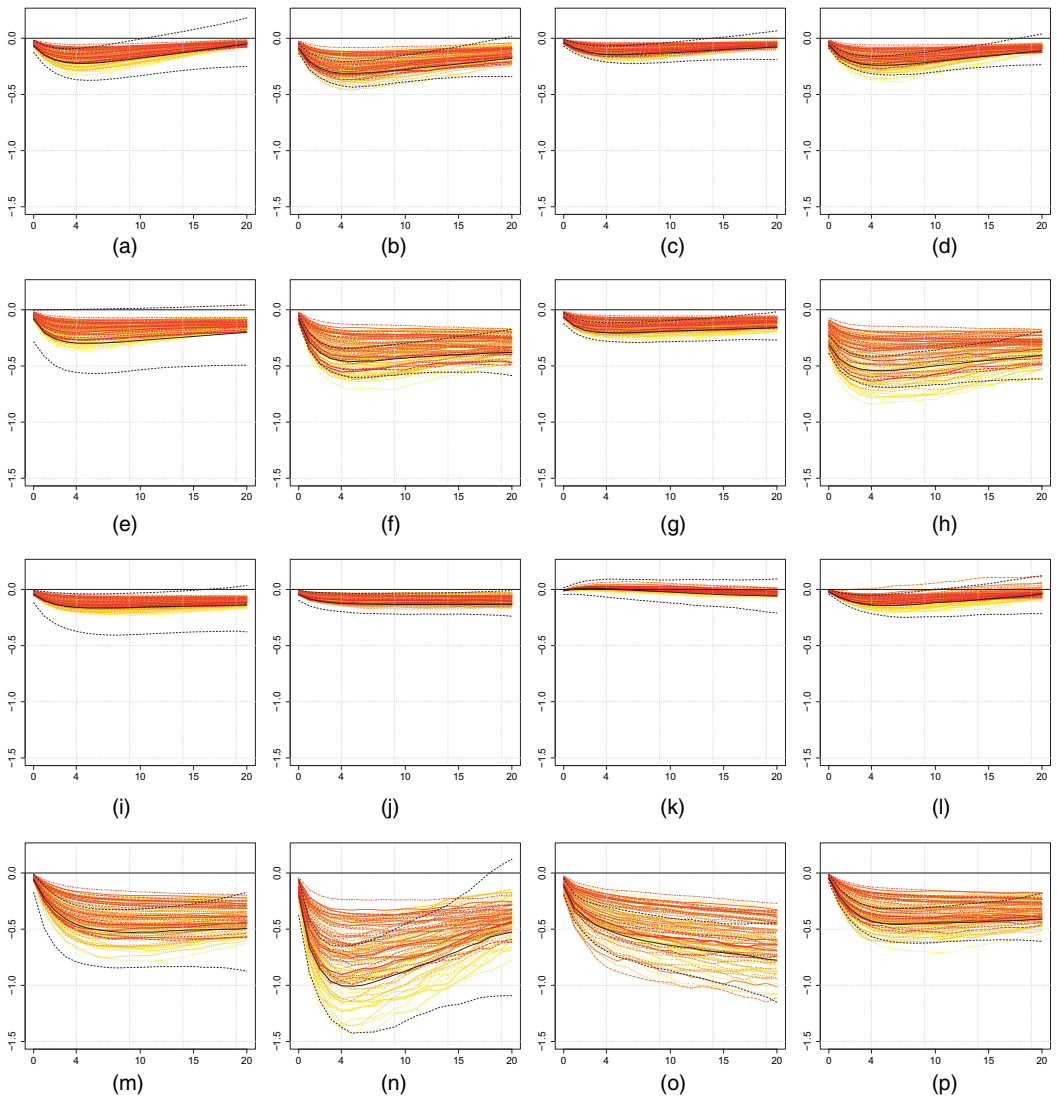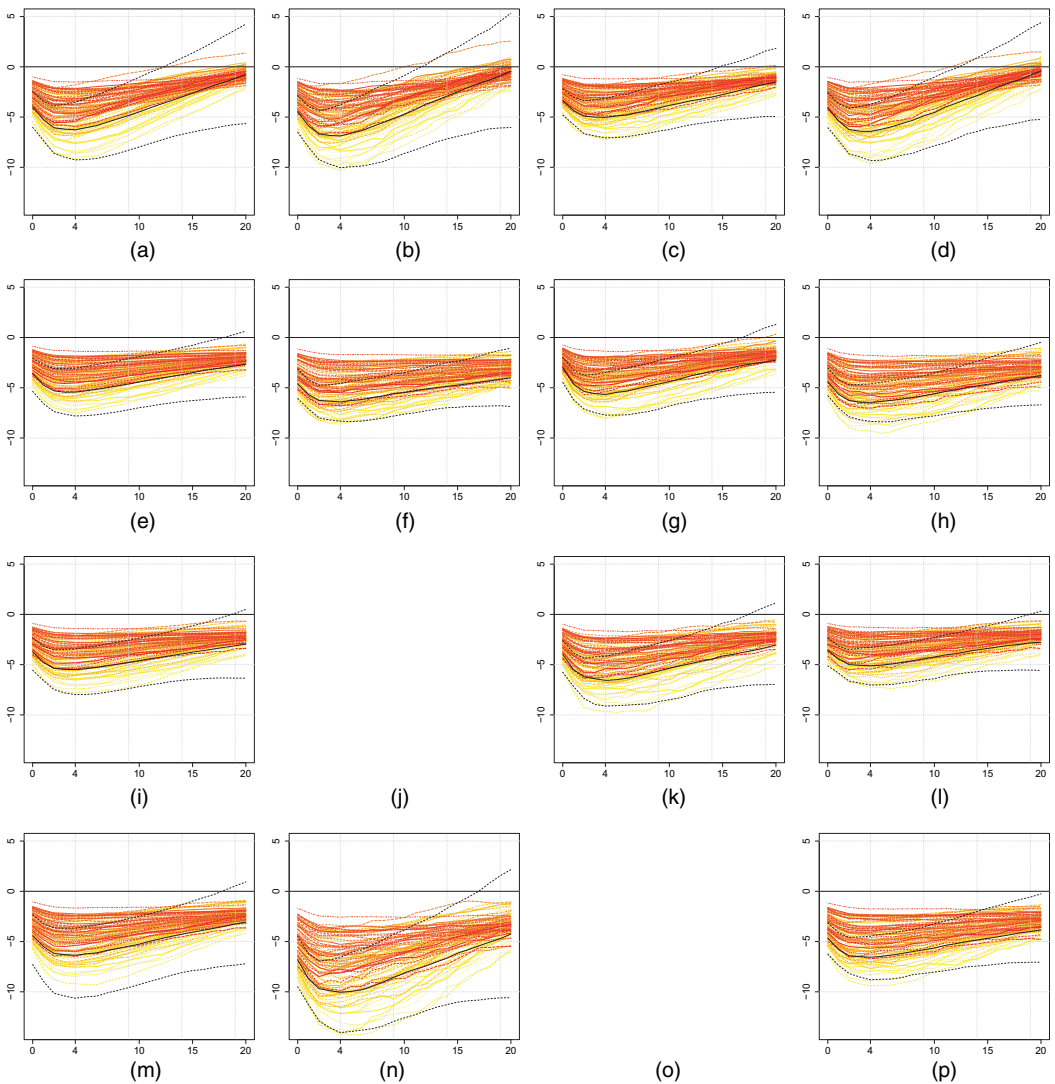
Fig. 4. Output responses over time for selected countries (the plots show posterior median responses over the sample period; ———, beginning of the sample (i.e. 1990, second quarter): ———, end of the sample (i.e. 2016, fourth quarter); ———, posterior median of the time-averaged response along with 68% credible bounds): (a) western Europe; (b) DE; (c) GB; (d) FR; (e) other developing economies; (f) CA; (g) JP; (h) US; (i) emerging Asian countries; (j) CN; (k) IN; (l) KR; (m) Latin American countries; (n) AR; (o) BR; (p) MX

variables. All three variables show a very similar pattern of trough responses over time. Roughly, these results indicate stronger international effects in the period from 1990, second quarter, to 2008, fourth quarter, compared with the post-crisis period. From a domestic perspective, this finding could be related to diminishing effectiveness of asset purchases in the USA, which we capture with the shadow rate during the zero lower bound period which lasted until 2015, fourth quarter. In fact, there is ample empirical evidence for abating effects of US asset purchase programmes on the US macroeconomy either due to diminishing effects on US investment growth (Stein, 2012; Feldkircher and Huber, 2018) or via signalling effects that have the
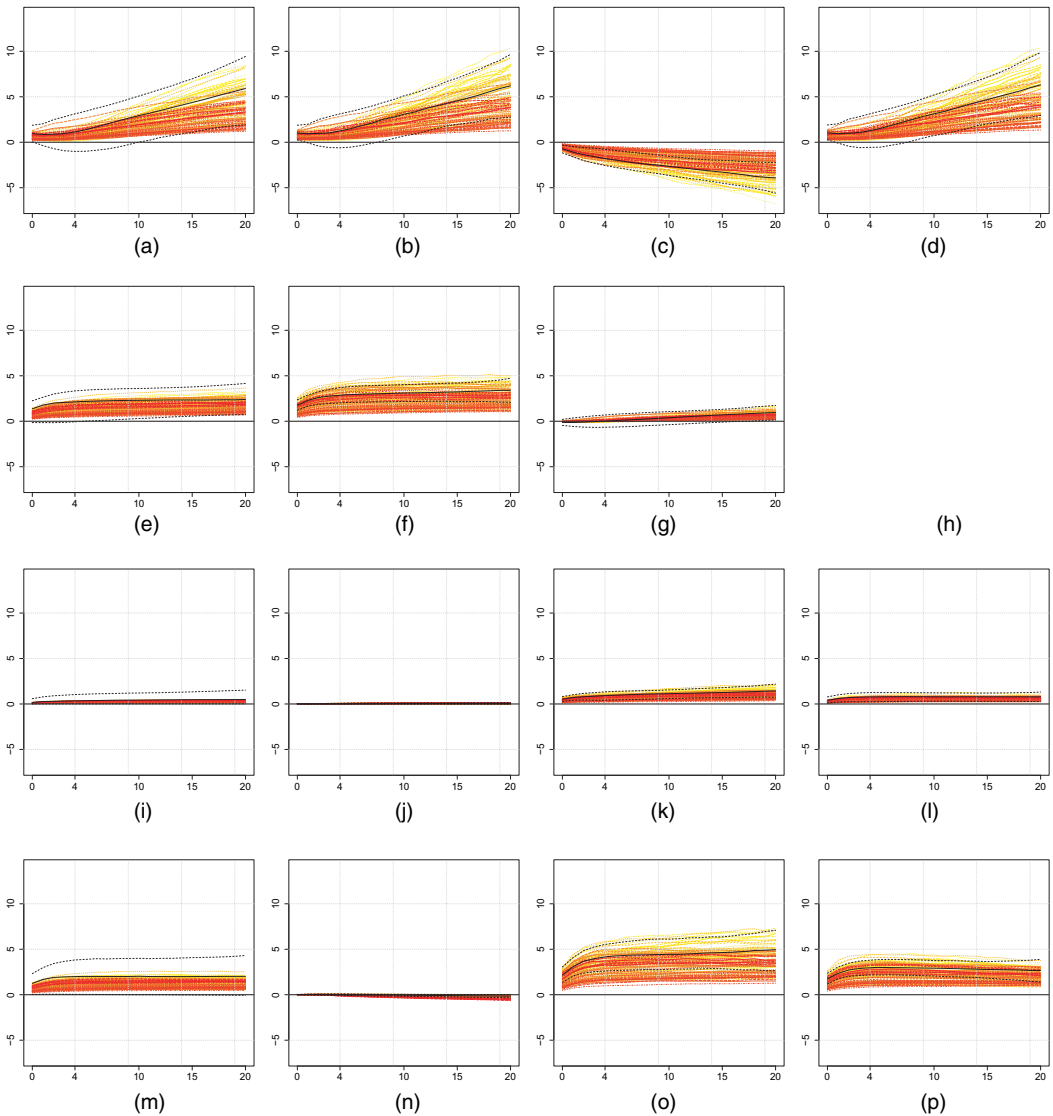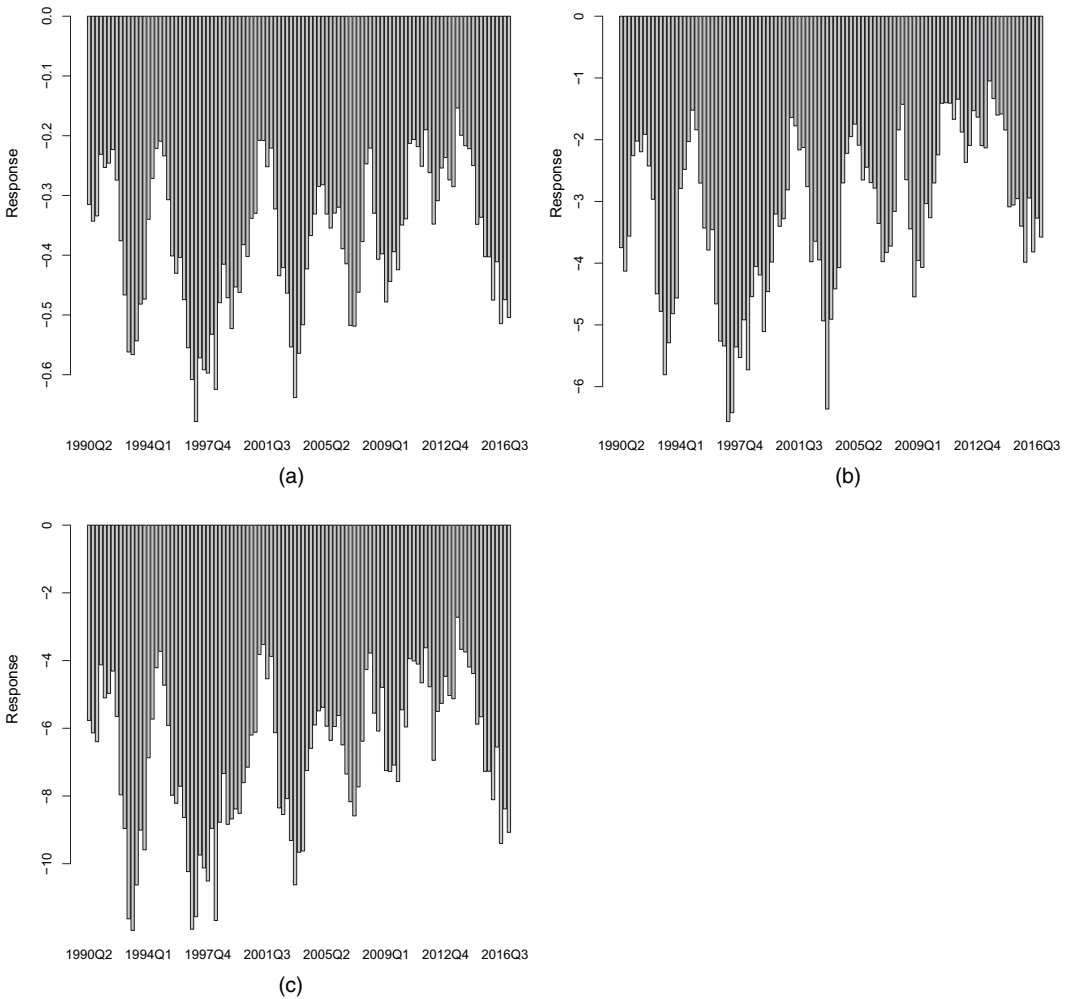
Fig. 5. Exchange rate responses over time for selected countries (the plots show posterior median responses over the sample period; ——, beginning of the sample (i.e. 1990, second quarter); ——, end of the sample (i.e. 2016, fourth quarter); ——, posterior median of the time-averaged response along with 68% credible bounds (an increase implies an appreciation of the US dollar against the respective local currency): (a) western Europe; (b) DE; (c) GB; (d) FR; (e) other developing economies; (f) CA; (g) JP; (h) US; (i) emerging Asian countries; (j) CN; (k) IN; (l) KR; (m) Latin American countries; (n) AR; (o) BR; (p) MX

strongest effect when financial markets are strained (Engen *et al.*, 2015). Weaker effects on the US economy could trigger weaker international effects. (That this is not necessarily so is demonstrated in Fratzscher *et al.* (2018) who showed that the last two US asset purchase programmes had a particularly strong effect on foreign markets.) Besides that, changes in the global macroeconomic environment such as declines in trade and financial globalization could also account for this finding—a more systematic analysis will be carried out in the next section. Last, note that there is no steady decline in the strength of the effects; rather they appear to evolve in cycles.

Fig. 6. Equity price responses over time for selected countries (the plots show posterior median responses over the sample period; ——, beginning of the sample (i.e. 1990, second quarter); ——, end of the sample, i.e. 2016, fourth quarter); ——, posterior median of the time-averaged response along with 68% credible bounds): (a) western Europe; (b) DE; (c) GB; (d) FR; (e) other developing economies; (f) CA; (g) JP; (h) US; (i) emerging Asian countries; (j) CN; (k) IN; (l) KR; (m) Latin American countries; (n) AR; (o) BR; (p) MX

Summing up, we find evidence for time variation in US domestic responses as well as for international reactions of selected quantities. Specifically, output, exchange rates and, most notably, equity price responses vary considerably throughout all regions. To be precise, variation in international output and equity price effects mostly regard short- to medium-term dynamics (up to 10 quarters), whereas exchange rate responses vary also in the longer term (up to 20 quarters). Focusing on trough and peak responses, we find no evidence for a steady increase or decline in the strength of spillover effects in the data—rather these evolve in cycles. Roughly, we find weaker international effects in the aftermath of the global financial crisis compared with

(a)

(b)

(c)

Fig. 7. Trough and peak responses over time (the plots show simple cross-country averages of trough responses of output and equity prices over time, as well as peak responses of exchange rates (i.e. maximum depreciations against the US dollar) on an inverted scale): (a) output; (b) exchange rates; (c) equity prices

the rest of the sample period. This could be explained either domestically by noting diminishing effects of large-scale asset purchase programmes on the US economy or by global trends such as changes in trade and financial globalization, which we shall analyse in more detail in the next section.

4.4. The influence of financial and trade globalization on spillovers
So far we have established that spillovers from US monetary policy are significant, different across countries and time varying. In this section, we assess whether the size of these international effects can be related to two phenomena that have shaped the global economy over recent decades: the increase in trade and financial globalization.

In a survey, Kamin (2013) reviewed the channels through which globalization can shape international responses. In a nutshell, and as pointed out in Mishkin (2009), the effect of more

globalization can either enhance or decrease spillovers. As countries become more integrated with the global economy, their macroeconomic variables generally become more exposed to external shocks. By the same token, countries with a broad trade base and a high degree of financial market integration reduce the risk of suffering from an adverse shock that originates from a single country. A second observation that was made in Kamin (2013) is that globalization may alter the transmission channels of monetary policy. As trade becomes more important, monetary policy could work more through exchange rates and net exports and less through its effects on domestic demand by steering long-term interest rates. Also, as government bond yields tend to be increasingly determined on global financial markets, their sensitivity to domestic changes in short-term rates might have abated (Kamin, 2013).

We measure trade and financial globalization by using the new version of the *Konjunktur-forschungsstelle* globalization indices KOF that were described in Gyglia *et al.* (2018). This updated database enables us to distinguish between *de facto* (activities) and *de jure* measures (policies). That these can differ substantially has been pointed out in Kose *et al.* (2009) and corroborated empirically in the context of economic growth by Quinn *et al.* (2011). *De facto* trade globalization is captured by variables that measure the exchange of goods and services over long distances and *de jure* globalization by measures that reflect the degree of trade regulation and tariff barriers. Financial globalization is measured by capital flows and stocks of foreign assets and liabilities, as well as *de jure* measures such as the foreign exchange rate regime and capital account and current account restrictions.

The indices, aggregated for the global economy, are depicted in Fig. 8. Looking first at *de facto* globalization measures, the picture shows that both trade and financial globalization have steadily increased up until 2008. With the onset of the global financial crisis, a sharp drop in world trade is reflected in a pronounced decline in trade globalization and a slowdown in financial globalization. Considering *de jure* measures yields a more diverse picture. Whereas trade *de jure* globalization follows closely its *de facto* counterpart, financial *de jure* globalization evolves differently. Capital account openness increased only until 1995 and then moved in cycles. Consistent with the other measures, *de jure* financial openness has declined from 2008 onwards.

To assess the effect of globalization measures on spillovers from the US monetary policy shock, we collect cumulative trough effects on output and consumer and equity prices, as well as peak effects (i.e. maximum appreciations of the US dollar) for the exchange rate. We focus on these variables since time variation of spillovers in interest rates and term spreads are comparably low. Similar exercises to explain cross-country differences in spillovers have been carried out recently in Georgiadis and Mehl (2016) and Dedola *et al.* (2017). (Using a large cross-country data set with both monthly and quarterly macroeconomic time series, Dedola *et al.* (2017) did not find a systematic relationship between a country's response and country characteristics. By contrast, Georgiadis (2016) found a range of potential determinants that can account for differences in the extent of spillovers. These include, among other characteristics, the receiving country's degree of trade and financial openness and the exchange rate regime. The importance of these determinants differs, however, across advanced and emerging countries and there are non-linear relationships between the determinants also.) Note, however, that the time varying approach yields trough or peak responses for each point over the sample period, which enables us to estimate a panel, as opposed to cross-country regressions that were previously employed in the literature, and to perform more reliable inference, as pointed out in Dedola *et al.* (2017). We convert the annual globalization indices to quarterly frequency by simply repeating the annual observations over the quarterly frequency domain. The results of a simple panel regression by using country fixed effects are depicted in Table 3.
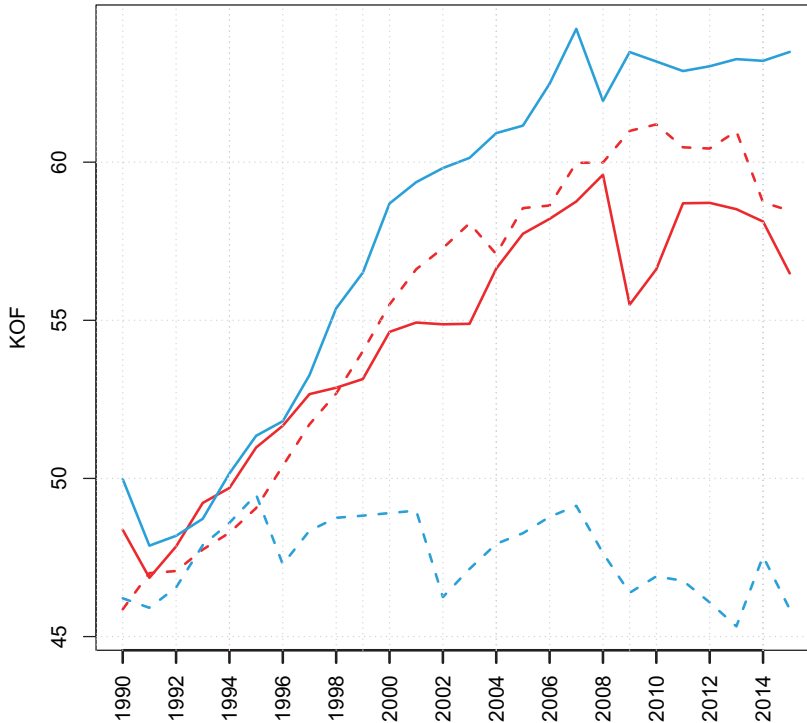
Fig. 8. KOF globalization indices (world): the plot shows the evolution of world trade *de facto* (———) and *de jure* (- - - -) and financial *de facto* (———) and *de jure* (- - -) globalization indices (data retrieved from https://www.kof.ethz.ch/en/forecasts-and-indicators/indicators/kof-globalisation-index.html)

First, we see that *de facto* measures of both trade and financial globalization reduce trough output and equity price responses as well as maximum depreciations against the dollar. This implies that economies with a broad trade base that are well integrated with the global financial economy are less exposed to spillover effects of a US monetary policy shock. Regarding financial globalization and output effects, this finding implicitly corroborates the results of Georgiadis and Mehl (2016), who postulated that the appreciation of the US dollar strengthens other countries' (dollar-held) asset positions, which mitigate negative effects on international gross domestic product. Moreover, as argued in Mishkin (2006), financial globalization can help to promote institutional reforms and, in turn, financial stability, thereby contributing to more output stability. By contrast, policies that reduce trade barriers and open the capital account can amplify spillovers to output, equity markets and the exchange rate. Since the uncovered interest rate parity condition that links interest rate differentials to expected exchange rate movements assumes substitutability of assets and capital account openness, a high degree of openness implies that interest rate differentials have a stronger effect on exchange rates. As a robustness check, we regressed the yearly average of trough or peak responses on the measures of globalization to investigate whether our results are driven by our frequency conversion of the annual indices to the quarterly domain. The results do not change qualitatively and are available on request.

Up to this point, we have shown that the international effects of a US monetary policy shock vary over time and that the degree of trade and financial globalization explains cross-country variation of these spillovers. A natural further question is to ask whether the relationship between spillovers and globalization has also changed over time: in other words, whether for example *de*

Table 3. Effects of globalization on strength of spillovers†

	Results for the following dependent variables:		
	y	*eq*	*er*
Trade (*de facto*)	0.002‡	0.059‡	−0.040‡
	(0.001)	(0.008)	(0.005)
Financial (*de facto*)	0.001§	0.031‡	−0.037‡
	(0.0004)	(0.006)	(0.004)
Trade (*de jure*)	−0.001‡	−0.024‡	0.035‡
	(0.0003)	(0.006)	(0.003)
Financial (*de jure*)	−0.003‡	−0.045‡	0.025‡
	(0.0003)	(0.005)	(0.003)
Observations	3605	2781	3502
R^2	0.039	0.130	0.144
Adjusted R^2	0.029	0.121	0.135
F-statistic	36.464‡	102.771‡	145.970‡

†Output *y* and equity prices eq refer to trough values of the cumulative impulse response; exchange rates er to peak values (i.e. maximum appreciations of the US dollar against the local currency). Country fixed effects are employed.
‡$p < 0.01$.
§$p < 0.05$.

facto trade globalization has always acted as a cushion to the monetary policy shock. We do this by running the same panel regression as before but using an extending window for the observations. We start using only the first four observations and then go forward until the end of the sample period. The estimated coefficients along with 1-standard-error bounds are depicted in Fig. 9, with Figs 9(a), 9(d), 9(g) and 9(j) showing results on international output effects, Figs 9(b), 9(e), 9(h) and 9(k) those on exchange rates and Figs 9(c), 9(f), 9(i) and 9(l) those on global equity prices.

A striking observation that emerges from these regressions is the changing correlation of the globalization indices and the trough or peak effects in the very early part of the sample until 1992. This holds equally true for output, exchange rate and equity price effects and for all measures of globalization. The reason for this is the changing dynamics of the globalization indices, which are also evident in Fig. 8. They fall at the beginning of the sample and then start to rise during 1992. The effect of *de jure* trade measures changes more strongly and frequently compared with the other globalization indicators. All estimated coefficients converge at the end of the sample period to the results that are shown in Table 3.

Summing up, we find that both trade and financial globalization can account for differences in international responses of output, equity prices and exchange rates. The distinction between *de facto* measures (i.e. outcomes) and *de jure* measures (i.e. policies) is crucial: loosening *de jure* measures amplifies spillovers to output and equity prices. It also triggers a greater sensitivity of the exchange rate to tightening of monetary policy in the USA. By contrast, a diverse trade base and a high degree of financial integration cushion the international effects of the US rate hike. The distinction between trade and financial globalization is less important since both measures have the same mitigating or amplifying effect on spillovers. These results hold true over most of the sample period, which indicates a stable relationship between international spillovers and globalization. An exception to this is the early part of our sample: a period that showed dynamics in globalization that were different from those of the rest of the observation period.
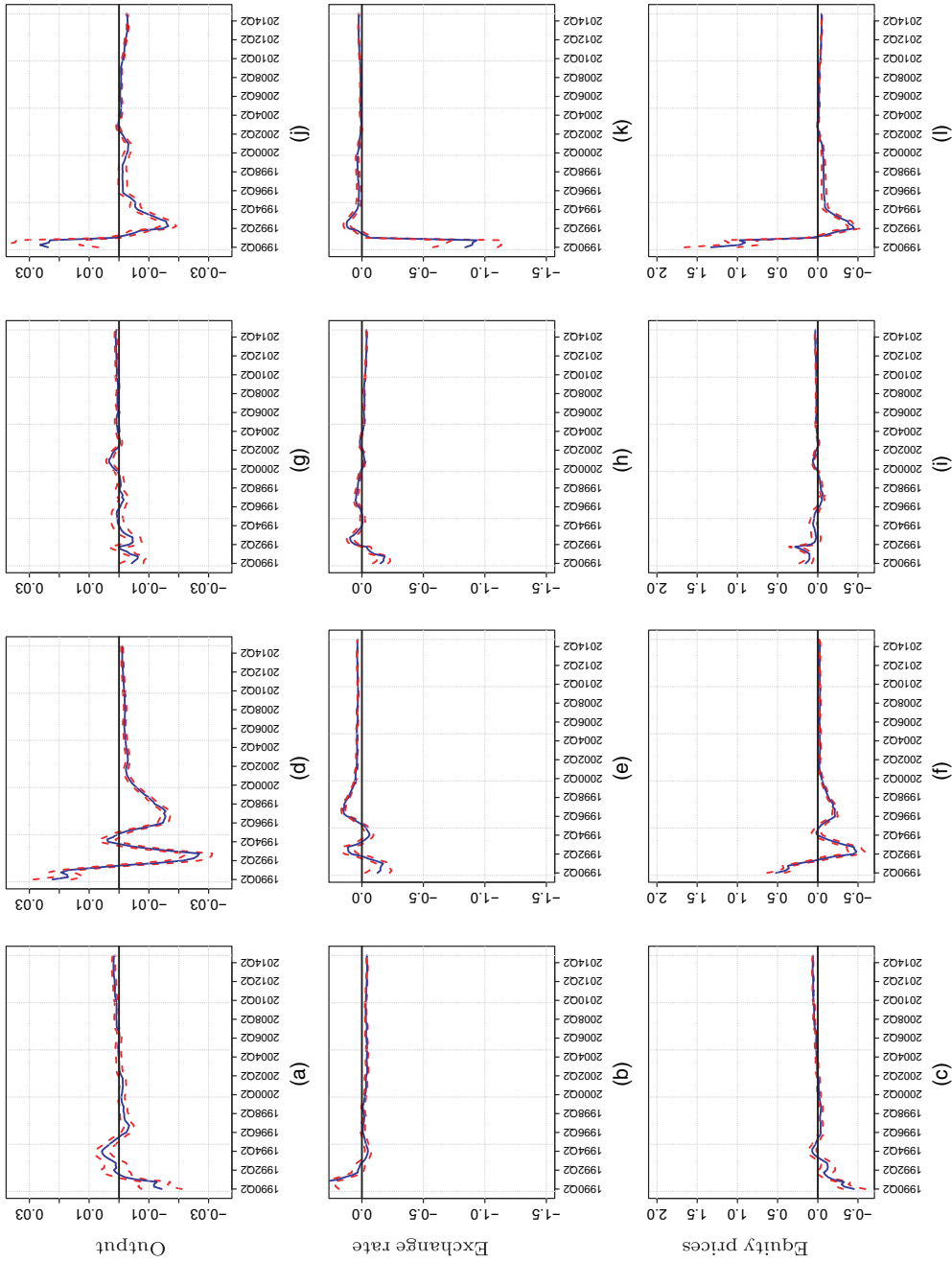
Fig. 9. Estimated coefficients of a rolling panel regression (the figure shows coefficient estimates of rolling panel regressions by using the same specification as in Table 3): (a)–(c) trade (*de facto*); (d)–(f) trade (*de jure*); (g)–(i) financial (*de facto*); (j)–(l) financial (*de jure*)

5. Closing remarks

This paper analyses the international effects of US monetary tightening taking explicitly into account that the extent of spillovers might have changed over time. For that, we develop a TVP GVAR model with SV and use shadow interest rates with sign restrictions to identify the monetary policy shock. Our econometric model yields a consistent framework for the analysis of monetary policy during normal times, as well as during the zero lower bound period. We further assess the global drivers of these spillovers, focusing on new measures of trade and financial globalization.

Our results indicate significant international effects caused by an unexpected tightening of US policy rates. In general, a US monetary policy contraction tends to decrease global output and this response appears to be quite persistent: a result which is in line with Feldkircher and Huber (2016). Global consumer prices tend to fall, and most currencies strengthen against the dollar. International short-term interest rates show a more diverse picture: they decrease in western Europe to compensate shortfalls in output, and they follow the rate hike in Latin America. Taking a regional angle, these effects are quite modest in emerging Asia compared with the rest of the world. We also find evidence for a significant and negative effect on global equity prices. These results relate to average effects over the sample period, from 1990, second quarter, to 2016, fourth quarter, and are consistent with the bulk of the literature on international effects of US monetary policy shocks.

More importantly, our results yield significant evidence for a changing international transmission of monetary policy shocks over time. This holds true for effects on output, exchange rates and especially so on equity prices. More precisely, for output and equity prices short-run dynamics tend to vary considerably, whereas long-run responses are less time sensitive. This implies that a (hypothetical) monetary policy shock has long-run consequences in 1990 that are similar to those in 2016, but their immediate effects might differ. By contrast, the effect on exchange rates varies also over the longer term. These results hint at complex dynamics of the international transmission mechanism and heterogeneity across countries. Considering the strength of international effects, our findings suggest that these evolve in cycles rather than follow a continuous trend. Roughly speaking, our results point at weaker international effects in the period after the global financial crisis compared with the earlier part of the sample period. Partially, this could be explained by domestic factors, such as diminishing effects of quantitative easing on the US economy (see, for example, Engen *et al*. (2015) and Stein (2012)).

Other potential driving forces of the strength of international spillovers are related to global developments in trade and financial globalization. Over our sample period, both forms of globalization have steadily increased until the outbreak of the global financial crisis which triggered a relaunch of prohibitive policies. This pattern is mostly consistent with trough or peak values that we observe for the variables with considerable time variation. A more systematic analysis that draws on the newly available KOF globalization indices shows distinct effects of globalization outcomes, *de facto* measures and policies, and *de jure* measures. An increase in *de facto* measures of both trade and financial globalization mitigates international spillovers. This implies that a broad trade base and a high degree of financial integration with the world economy cushion risks stemming from a foreign shock such as a rate increase in the USA. By contrast, an increase in *de jure* measures (i.e. a reduction of trade barriers and/or a liberalization of the capital account) is associated with stronger international responses to the US shock. In this sense, our analysis yields a more precise assessment of the theoretically ambiguous effects of globalization on the macroeconomy as pointed out in Mishkin (2009). Except for the early part of our sample, these relationships between globalization and spillovers are stable over time.

The distinction between trade and financial globalization as a determinant of the strength of spillovers is less important.

Acknowledgements

Appendix A: Technical appendix

A.1.   Prior specification

In a Bayesian framework, we need to elicit priors on the coefficients in equation (2.10). Crespo Cuaresma *et al.* (2016) showed that prior elicitation at the individual country level translates into a specific prior structure at the global level, providing additional shrinkage through the trade weights that are used. We impose a normal–gamma prior (Griffin and Brown, 2010) on $\boldsymbol{\xi}_{i0}$, the initial state of $\boldsymbol{\xi}_{it}$:

$$\boldsymbol{\xi}_{i0} \sim \mathcal{N}(\mathbf{0}_{s_i}, \underline{\mathbf{V}}_{\xi_i}), \tag{A.1}$$

with $\mathbf{0}_{s_i}$ being an s_i-dimensional vector of 0s and $\underline{\mathbf{V}}_{\xi_i}$ an $s_i \times s_i$ diagonal prior variance–covariance matrix. (For a recent application of this prior to the VAR case, see Huber and Feldkircher (2017).) We assume that each diagonal element of $\underline{\mathbf{V}}_{\xi_i}$, labelled $\underline{v}_{\xi_i j}$, features a gamma prior with

$$\underline{v}_{\xi_i j} \sim \mathcal{G}(\theta_i, \theta_i \kappa_i / 2). \tag{A.2}$$

We let θ_i denote a scalar hyperparameter that is specific to each country that serves to control the tail be-haviour of the prior. It allows for non-zero regression coefficients in the presence of a large global shrinkage parameter $\kappa_i \sim \mathcal{G}(q_0, q_1)$ that pushes all elements in $\boldsymbol{\xi}_i$ to 0. Again, q_0 and q_1 are prior hyperparameters that are typically set to small values. In the empirical application we set $\theta_i = 0.1$ and $q_0 = q_1 = 0.01$. These choices are based on VAR evidence provided in Huber and Feldkircher (2017).

We can think of this prior specification as an approximation to spike-and-slab priors (George *et al.*, 2008), that fail to perform well if s_i is large. The main goal of this specification is to select whether a given regressor should be included or excluded at time $t = 0$. Note that the question whether some covariate should enter the model over time depends not only on whether $\underline{v}_{\xi_i j}$ is close to 0 but also whether that parameter is time varying. Since we introduce a flexible law of motion for $\boldsymbol{\xi}_i$, our model selects important regressors and enables a stochastic model specification search over time.

For the parameter governing time variation of coefficients $\vartheta_{ij,1}^{-1}$, we use a gamma prior with

$$\vartheta_{ij,1}^{-1} \sim \mathcal{G}(n_0, n_1), \tag{A.3}$$

where $n_0 = 3$ and $n_1 = 0.03$ are hyperparameters. Moreover, on the thresholds we introduce a prior that is uniformly distributed and depends on $\vartheta_{ij,1}$:

$$c_{ij} | \vartheta_{ij,1} \sim \mathcal{U}(\pi_{ij,0}\sqrt{\vartheta_{ij,1}}, \pi_{ij,1}\sqrt{\vartheta_{ij,1}}). \tag{A.4}$$

We set $\pi_{ij,0} = 0.1$ and $\pi_{ij,1} = 3$, effectively bounding the thresholds away from zero. This implies that high frequency noise in the latent states is always set equal to zero, effectively reducing uncertainty stemming from this source without seriously distorting inference.

Moreover, we use the prior set-up that was proposed in Kastner and Frühwirth-Schnatter (2013) and subsequently used in Huber (2016) on the coefficients of the log-volatility process in equation (2.4). A normal prior is imposed on μ_{il} ($l = 1, \ldots, k_i$) with mean $\underline{\mu}_i$ and variance $\underline{V}_{\mu_i}$:

$$\mu_{il} \sim \mathcal{N}(\underline{\mu}_i, \underline{V}_{\mu_i}). \tag{A.5}$$

For the persistence parameter ρ_{il}, we elicit a beta-distributed prior:

$$\frac{\rho_{il} + 1}{2} \sim \text{beta}(a_0, b_0). \tag{A.6}$$

For typical data sets arising in macroeconomics, the exact choice of the hyperparameters $a_0 = 25$ and $b_0 = 5$ in equation (A.6) is quite influential, since data do not tend to be very informative about the degree of persistence of log-volatilities. Our proposed choice translates into quite a persistent log-volatility process and appears to be in concordance with earlier literature (see Primiceri (2005)).

Finally, we impose a non-conjugate gamma prior for ς_{ij}^2 ($j = 1, \ldots, k_i$):

$$\varsigma_{ij}^2 \sim \mathcal{G}\left(\frac{1}{2}, \frac{1}{2B_\varsigma}\right). \tag{A.7}$$

This choice does not bound ς_{il}^2 away from zero, thus providing more shrinkage than standard typical conjugate inverted gamma priors do. Here, we follow the recommendations that were provided in Kastner and Frühwirth-Schnatter (2013) and set $B_\varsigma = 1$.

A.2. Posterior inference

Using the prior setting that was described in Section 2, an MCMC algorithm to draw samples from the (country-specific) parameter posterior distribution can be designed. Denote the full history of the time varying elements in equation (2.5) up to time T as

$$\boldsymbol{\xi}_i^T = \text{vec}(\boldsymbol{\xi}_{i1}', \ldots, \boldsymbol{\xi}_{iT}')',$$
$$\boldsymbol{\lambda}_i^T = (\lambda_{i1}, \ldots, \lambda_{iT})'.$$

The MCMC algorithm consists of the following blocks.

(a) $\boldsymbol{\xi}_i^T$ is sampled through the well-known algorithm that was provided in Carter and Kohn (1994) and Frühwirth-Schnatter (1994).

(b) Conditionally on $\boldsymbol{\xi}_i^T$, the hierarchy of the model implies that the posterior of $\vartheta_{ij,1}^{-1}$ ($j = 1, \ldots, s_i$) is independent of the data and follows a gamma distribution:

$$\vartheta_{ij,1}|\cdot \sim \mathcal{G}^{-1}\left\{\frac{T_1}{2} + \frac{1}{2} + n_0, n_1 + \frac{\sum_{t=1}^{T} d_{ij,t}(\xi_{ij,t} - \xi_{ij,t-1})^2}{2}\right\}, \tag{A.8}$$

with $T_1 = \Sigma_{t=1}^{T} d_{ij,t}$ denoting the number of observations that display time variation in the jth parameter of country i, and the dot notation is a generic notation that indicates conditioning on all remaining model parameters and the data.

(c) The posterior of $\underline{v}_{ij}^2$ follows a generalized inverted Gaussian distribution:

$$\underline{v}_{\xi_{ij}}|\cdot \sim \mathcal{GIG}\left(\theta_i - \frac{1}{2}, \xi_{ij,0}^2, \theta_i \kappa_i^2\right). \tag{A.9}$$

We sample from the generalized inverse Gaussian distribution by using the R package `GIGrvg` (Leydold and Hörmann, 2015).
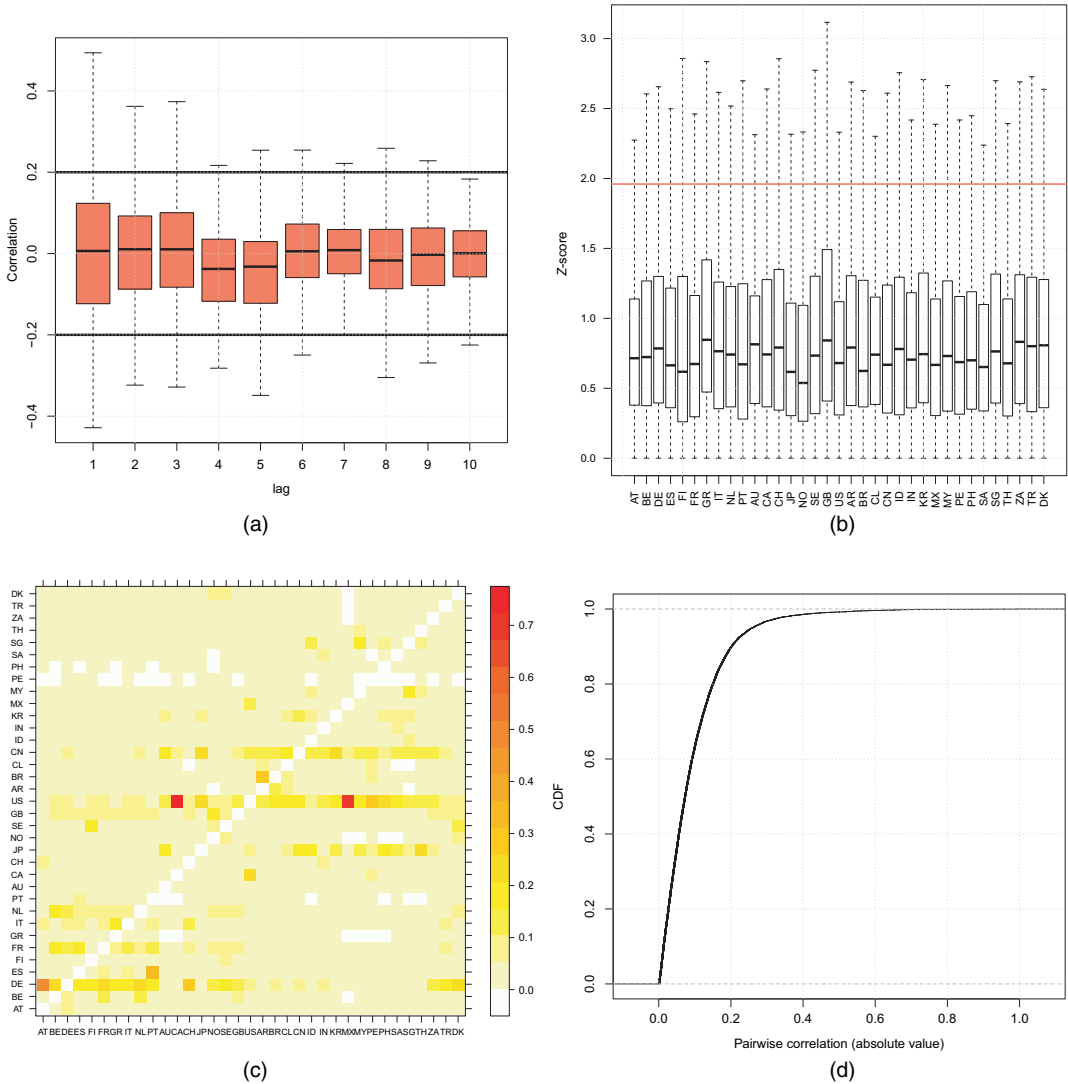
Fig. 10. Diagnostics of the estimated GVAR model: (a) auto-correlation function of the cross-country residuals; (b) Z-scores of Geweke's convergence diagnostic (Geweke, 1992); (c) distribution of trade weights; (d) empirical cumulative density function of average pairwise cross-country residual correlations (in absolute values)

(d) The global shrinkage parameter κ_i is sampled from a gamma distribution given by

$$\kappa_i | \cdot \sim \mathcal{G}\left(q_0 + \theta_i s_i, q_1 + \frac{\theta_i}{2} \sum_{j=1}^{s_i} \underline{v}_{\xi_{ij}}\right). \tag{A.10}$$

(e) We use a griddy Gibbs algorithm (Ritter and Tanner, 1992) to obtain draws from the posterior distribution of the thresholds c_j. Because of the hierarchical structure of the model, the likelihood is given by

$$p(\boldsymbol{\xi}_i^T | \cdot) \propto \prod_{t=1}^{T} \frac{1}{\sqrt{\vartheta_{ij,t}}} \exp\left\{ -\frac{(\xi_{ij,t} - \xi_{ij,t-1})^2}{2\vartheta_{ij,t}} \right\}. \tag{A.11}$$

We combine equation (A.11) with the uniform prior in equation (A.4) and evaluate the conditional posterior of c_{ij} over a fine grid of potential values to approximate the inverse cumulative distribution function of the posterior. This approximation is then consequently utilized to perform inverse transform sampling.

(f) The history of log-volatilities is sampled by using the algorithm that was outlined in Kastner and Frühwirth-Schnatter (2013).

To speed up computation and to permit equation-by-equation estimation, we exploit a Cholesky ordering for estimating the model (Koop *et al.*, 2018). This implies that, for each equation j within a given country model, we include the endogenous variables of the preceding $j-1$ equations to draw the covariance parameters and the VAR coefficients jointly. This technique, in principle, is not order invariant with respect to random permutations in $\mathbf{y}_{it}$. However, we found in limited experiments that randomly assigning the order of the variables in $\mathbf{y}_{it}$ yields only minor differences in the estimated impulse responses. Note that this is a short cut that enables us to exploit parallel computing fully.

A.3. Convergence properties of the Markov chain Monte Carlo algorithm and residual diagnostics

In Fig. 10 we show several diagnostic checks based on the residuals of the country models. These are based on 10000 posterior draws after a burn-in phase of 30000 draws. From Fig. 10(a) we see that the residuals are generally not serially auto-correlated. In Fig. 10(b) we show boxplots of Z-scores of Geweke's convergence diagnostic (Geweke, 1992) per country. These indicate that the MCMC algorithm has converged to its target distribution since most (absolute) values of the statistic are below the 1.96-threshold. Fig. 10(c) illustrates the distribution of the trade weights. One assumption underlying the GVAR framework is that the weights are relatively small (see Pesaran *et al.* (2004)). We see that most countries are well integrated with the rest of the world and weights tend to be small (i.e. equally distributed). Germany for Austria, and the USA (for Canada and Malaysia) are notable exceptions. Last, we show in Fig. 10(d) that the cross-sectional dependence of the country residuals is generally weak. The cumulative density function of the pairwise correlations across the country residuals show that 90% of the mass lies below 30%, indicating weak cross-sectional dependence (Burriel and Galesi, 2018).

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Supporting information

Additional 'supporting information' may be found in the on-line version of this article:

'Supplementary material: Spillovers from US monetary policy: evidence from a time-varying parameter GVAR model'.