

US Monetary Policy in a Globalized World*

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Abstract

This paper develops a global vector autoregressive model with time-varying parameters and stochastic volatility (TVP-SV-GVAR) to analyze the international effects of US monetary policy. We find significant changes in the international transmission of US monetary policy over time. For most variables, the global response to US monetary policy has been increasing since the 1980s, while effects on output were dampened following the period of the global financial crisis. Countries that are more strongly affected tend to be highly open economies with excessive fiscal and current account deficits and pronounced foreign currency exposure (either through reserves or the external balance). Improving these fundamentals, reducing trade openness or restricting capital inflows can mitigate spillovers. Last, and as a consequence of an increase of financial globalization, we show that also US rates react significantly to certain foreign shocks.

Keywords: Global vector autoregression, time-varying parameters, stochastic volatility, monetary policy, international spillovers

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

Ben Bernanke *Globalization and Monetary Policy*.
Speech at the Fourth Economic Summit, Stanford, 2007

1 Introduction

Economic theory has long recognized the interdependence of national economies. Models such as the Mundell-Fleming framework or microfounded New Keynesian approaches describe the effects that shocks to one economy may have on its trading partners (see [Obstfeld and Rogoff, 1995](#), for a textbook treatment). These models, however, have often been interpreted as only being valid for small open economies. Theory predicts that large and rather closed economies such as the US are more insulated from foreign shocks, especially if they pursue a flexible exchange rate regime that can serve as a buffer to external shocks. This line of thinking has also been reflected in the specifications used for monetary policy rules that describe the behavior of the US Federal Reserve in setting its monetary policy stance. One of the most prominent monetary policy reaction functions, the Taylor rule ([Taylor, 1993](#)), describes monetary policy directly in terms of the two major operational objectives of monetary policy, domestic inflation and economic growth. Among others, [Orphanides \(2003\)](#) finds that the simple Taylor rule serves as a particularly good description of Federal Reserve policies virtually since the founding of the institution. According to the standard Taylor rule, the US Fed sets monetary policy in response to developments of domestic macroeconomic variables and independently of other external factors. In recent years, however, the ability of monetary policy in the world's largest economies to independently control monetary policy objectives has been put into question ([Kamin, 2010](#)). Not surprisingly, monetary policymakers have taken an active interest in the extent to which increased globalization affects their ability to independently set monetary policy.

The implications of increased globalization on the policy behavior of the Fed itself have been significantly less researched. The trend in financial globalization may have increased the importance of external factors for domestic monetary conditions in the US. This, in turn, would imply less independence and control on setting domestic interest rates to successfully

shape domestic financial and economic conditions (Kamin, 2010). Monetary policy in a globalized world could be modeled directly by expanding the Taylor rule to feature international factors such as global output. Alternatively, one could think of the Fed reacting to external shocks via its response to domestic growth, which can be reasonably argued to be (at least partly) influenced by foreign shocks.

Accounting for changes in the economic environment and the reaction function of monetary policymakers appears essential when modeling monetary policy. Among researchers, a consensus has emerged concerning the fact that monetary policy in the US has changed over the last three decades (Sims and Zha, 2006; Boivin et al., 2010). Variation in the implementation of monetary policy and its effectiveness might be driven by several factors, including regulatory changes and changes in domestic and global macroeconomic and financial market conditions. In addition to changes in the reaction function of the Fed, changes in the economic environment can affect the outcome of monetary policy both in the US and globally. In particular uncertainty, understood as the time-varying component of the volatility of economic shocks, has been shown to be an important factor explaining the dynamics of real economic activity (Bloom, 2009; Fernández-Villaverde et al., 2011).

This paper uses of a new class of global macroeconomic models to assess the dynamic relationship between US monetary policy and the world economy over time. We augment the global vector autoregressive model put forth in Pesaran et al. (2004) to allow for changes in parameters and error variances. The newly developed time-varying parameter stochastic volatility global vector autoregressive (TVP-SV-GVAR) model is estimated using Bayesian methods for a global sample corresponding to approximately 80% of global output. To cope with such a data-rich environment efficiently from a computational point of view, we draw on recent contributions on Cholesky stochastic volatility models proposed by Lopes et al. (2013). Within this modeling framework, we examine both spillovers from US monetary policy to the global economy and vice-versa. We also address changes in spillovers over time to judge whether the transmission from and to the US has significantly changed in the last decades. In a second step, we shed light on the question to what extent differences in spillovers across world economies are driven by country-specific characteristics.

Our results can be summarized as follows. First, a contractionary shock to US monetary policy tends to imply (a) a persistent global contraction in real activity, (b) a drop in international prices together with (c) a rise in global nominal interest rates, and (d) a real appreciation of the US dollar. The estimated effects are in line with the existing empirical literature on the effects of shocks to monetary policy originated in the US on other economies (see Feldkircher and Huber, 2016). Second, we find clear evidence for a changing transmission of US monetary policy shocks over time at the global level. For most variables, the global

response to US monetary policy has been larger during the period from the mid-1990s to mid-2000s, while effects on output dampened following the recent global financial crisis. Third, we find evidence for heterogeneity of the spillovers across economies. Highly open economies that face excessive fiscal and current account deficits are more strongly exposed to US shocks. Naturally, this holds also true for countries with a high share of (US-denominated) foreign exchange reserves or a pronounced foreign exchange component of the external balance. Improving these fundamentals, reducing trade openness, having a floating exchange rate or restricting capital inflows can mitigate spillovers. Last, and as a consequence of heightened financial globalization, US monetary policy itself is found to respond significantly to foreign shocks. An increase in foreign interest rates or a decrease in international output tends to trigger a decrease in US domestic rates. This effect seems to have increased over our sample. There is less evidence of other foreign shocks impacting US rates.

There is a large literature on the interaction between monetary policy in the US and other economies (see [Giancarlo Corsetti, 2010](#), for a recent survey). Recently, some papers also studied international linkages between US monetary policy and other countries. [Luca Dedola and Stracca \(2016\)](#) investigate the global effects of US monetary policy in a two-stage model using Bayesian VAR with sign restrictions. [Rey \(2015\)](#) proposed at the annual Jackson-Hole conference the analysis of a “Global Financial Cycle” that affects the global economy through financial variables, such as risk premia and term spreads. [Rey \(2015\)](#) and [Miranda-Agrippino and Rey \(2015\)](#) analyze the transmission channel of international monetary spillovers through these financial variables in a medium scale Bayesian VAR with real, financial and monetary variables.

Our paper differs from this literature by explicitly modeling the changing transmission mechanism of US monetary policy through time-varying coefficients and stochastic volatility in a global VAR (TVP-SV-GVAR). We identify US monetary policy shocks using a recursive identification proposed among others in [Christiano et al. \(2005\)](#). The rich cross-sectional structure in the GVAR and the flexible time-varying parameter and stochastic volatility structure allows for very complex dynamics in a sample of 36 countries representing a large part of the global economy.

The paper is structured as follows. Section 2 presents the econometric framework including the Bayesian estimation strategy and the prior specifications which makes estimation of the TVP-SV-GVAR model feasible. Section 3 presents the data, while section 4 discusses the results. Finally, section 5 concludes.

2 Econometric framework: The TVP-SV-GVAR specification

To assess the dynamic transmission mechanism between US monetary policy and the global economy, we develop a global VAR model featuring time-varying parameters and stochastic volatility (TVP-SV-GVAR model). The TVP-SV-GVAR model is estimated 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 using country weights to form a global model that is used to carry out impulse response analysis or forecasting.

2.1 The global vector autoregressive model with time-varying parameters

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

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

where w_{ic} is the weight corresponding to the j th variable of country c in country i 's specification. These weights are typically assumed to be related to bilateral trade exposure, sum up to 1 and only off-diagonal elements are non-zero ($\sum_{c=0}^N w_{ic} = 1$ and $w_{ii} = 0$). In line with the bulk of the literature on GVAR modeling, we assume that all variables and countries are linked by the same set of weights which is fixed over time (Dees et al., 2007a). 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 proposed TVP-SV-GVAR model allows for time variation in all coefficients as well as changes in residual variance and is thus capable to model a much richer set of dynamics at the international level.¹ We deviate from existing GVAR modeling efforts by specifying country-specific structural VAR models featuring exogenous regressors,

¹Moreover, note that in the empirical application we are not interested in interpreting 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.

time-varying parameters and stochastic volatility, so that

$$A_{i0,t}y_{it} = \sum_{p=1}^P B_{ip,t}y_{it-p} + \sum_{q=0}^Q \Lambda_{iq,t}y_{it-q}^* + \varepsilon_{it}, \quad (2.2)$$

where

- $A_{i0,t}$ is a $k_i \times k_i$ matrix of structural coefficients used to establish contemporaneous relationships between the variables in y_{it} . We assume that $A_{i0,t}$ is a lower triangular matrix with a diagonal of ones. This choice ensures that the errors of the model are orthogonal to each other by imposing a Cholesky structure on the specification;
- $B_{ip,t}$ ($p = 1, \dots, P$) is a $k_i \times k_i$ matrix of coefficients associated with the lagged endogenous variables;
- $\Lambda_{iq,t}$ ($q = 0, \dots, Q$) denotes a $k_i \times k_i^*$ dimensional coefficient matrix corresponding to the k_i^* weakly exogenous variables in y_{it}^* ;
- $\varepsilon_{it} \sim \mathcal{N}(0, D_{it})$ is a heteroskedastic vector error term with $D_{it} = \text{diag}(\lambda_{i0,t}, \dots, \lambda_{ik_i,t})$. The assumption of a diagonal D_{it} simplifies the computational burden of model estimation enormously, since the k_i equations can be viewed as separate estimation problems and hence easily parallelized to achieve computational gains.²

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

$$x_{it} = (y_{it-1}, \dots, y_{it-P}, y_{it}^*, \dots, y_{it-Q}^*)' \quad (2.3)$$

and storing all coefficients in a $k_i \times (m_i k_i)$ matrix Ψ_{it} ,

$$\Psi_{it} = (B_{i1,t}, \dots, B_{iP,t}, \Lambda_{i0,t}, \dots, \Lambda_{iQ,t})' \quad (2.4)$$

allows us to rewrite [equation \(2.2\)](#) as

$$A_{i0,t}y_{it} = (I_{k_i} \otimes x_{it}') \text{vec}(\Psi_{it}) + \varepsilon_{it}. \quad (2.5)$$

²The ordering of the variables will be discussed in [section 3](#) and is the same used to identify the structural shocks later on. See the Appendix for further details on the computational challenges involved in obtaining posterior distributions for model quantities of interest.

Collecting the elements of $A_{i0,t}$ which are not zero or unity in a $l_i = k_i(k_i - 1)/2$ -dimensional vector $a_{i0,t}$, the law of motion of $a_{i0,t}$ is assumed to be given by

$$a_{i0,t} = a_{i0,t-1} + \epsilon_{it}, \quad \epsilon_{it} \sim \mathcal{N}(0, V_i) \quad (2.6)$$

where V_i is a diagonal variance-covariance matrix with $V_i = \text{diag}(v_{i1}^2, \dots, v_{il_i}^2)$. The diagonal nature stems from the fact that we estimate the model on an equation-by-equation basis, thus effectively disregarding the contemporaneous relationships between parameters in the model. Likewise, we assume that the $K_i = k_i^2 m_i$ autoregressive coefficients in Ψ_{it} evolve according to

$$\text{vec}(\Psi_{it}) = \text{vec}(\Psi_{it-1}) + \eta_{it}, \quad \eta_{it} \sim \mathcal{N}(0, S_i), \quad (2.7)$$

with $S_i = \text{diag}(s_{i1}^2, \dots, s_{iK_i}^2)$ being a $K_i \times K_i$ variance-covariance matrix. Finally, the the variances $\lambda_{il,t}$ are assumed to follow a stationary autoregressive process,

$$\log(\lambda_{il,t}) = \mu_{il} + \rho_{il}(\log(\lambda_{il,t-1}) - \mu_{il}) + v_{il,t}, \quad v_{il,t} \sim \mathcal{N}(0, \varsigma_{il}^2), \quad (2.8)$$

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.

Some features of the model in [equation \(2.2\)](#) deserve a more detailed explanation. All parameters are allowed to vary over time, which implies that we can explicitly account for changes in domestic and international transmission mechanisms with our specification. Moreover, we also account for heteroskedasticity by making the country-specific variance-covariance matrix of ϵ_{it} time-varying. Our model can thus simultaneously accommodate many features which are commonly observed in macroeconomic and financial time series data. Moreover, the inclusion of weakly exogenous foreign variables accounts for cross-country linkages and enables us to investigate the propagation of economic shocks across both space and time. Given the marked increase in globalization and the stronger degree of business cycle synchronization experienced globally over the last decades, this is an essential ingredient when modeling the transmission of shocks at the global level.

The set of $N + 1$ country specific models can be linked together 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 $z_{it} = (y'_{it}, y'^*_{it})'$, we obtain

$$C_{it}z_{it} = \sum_{s=1}^{\ominus} L_{is,t}z_{it-s} + \epsilon_{it} \quad (2.9)$$

with $C_{it} = (A_{i0,t}, -\Lambda_{i0,t})$, $L_{is,t} = (B_{is,t}, \Lambda_{is,t})$ and $\mathfrak{S} = \max(P, Q)$. A global vector $y_t = (y'_{0t}, \dots, y'_{Nt})'$ of dimension $k = \sum_{i=0}^N k_i$ and a corresponding country-specific linkage matrix W_i ($i = 1, \dots, N$) of dimension $(k_i + k_i^*) \times k$ can be defined so as to rewrite [equation \(2.9\)](#) exclusively in terms of the global vector,

$$C_{it}W_i y_t = \sum_{s=1}^{\mathfrak{S}} L_{is,t} W_i y_{t-s} + \varepsilon_{it}. \quad (2.10)$$

Stacking the equations $N + 1$ times yields

$$G_t y_t = \sum_{s=1}^{\mathfrak{S}} F_{st} y_{t-s} + e_t \quad (2.11)$$

where $G_t = ((C_{0s,t}W_0)', \dots, (C_{Ns,t}W_N)')$ and $F_{st} = ((L_{0s,t}W_0)', \dots, (L_{Ns,t}W_N)')$. The error term $e_t = (\varepsilon'_{0t}, \dots, \varepsilon'_{Nt})'$ is normally distributed with variance-covariance matrix $H_t = \text{diag}(D_{0t}, \dots, D_{Nt})$. Equation (2.11) resembles thus a (very) large VAR model with drifting coefficients which, notwithstanding the problems associated with the high dimensionality of the parameter vector, can be estimated using Bayesian techniques developed to deal with multivariate linear models with time-varying parameters.

2.2 Bayesian estimation of the TVP-SV-GVAR model

We use Bayesian methods to carry out inference in the TVP-SV-GVAR model proposed above. Given the risk of overparameterization that is inherent to the specifications used, we rely on Bayesian shrinkage methods to achieve 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.

In a Bayesian framework we need to elicit priors on the coefficients in [equation \(2.5\)](#). [Crespo Cuaresma et al. \(2016\)](#) show that prior elicitation at the individual country levels translates into a specific prior structure at the global level, providing additional shrinkage through the trade weights used. We impose a normally distributed prior on Ψ_{i0} , the initial state of Ψ_{it} ,

$$\text{vec}(\Psi_{i0}) \sim \mathcal{N}(\text{vec}(\underline{\Psi}_i), \underline{V}_{\Psi_i}), \quad (2.12)$$

with $\underline{\Psi}_i$ a $k_i \times m_i$ prior mean matrix and \underline{V}_{Ψ_i} a $k_i m_i \times k_i m_i$ prior variance-covariance matrix. In addition, we specify a prior for the free parameters of the state equation. We impose a Gamma distributed prior on the elements of the variance-covariance matrix S_i in [equation \(2.7\)](#). As noted in [Frühwirth-Schnatter and Wagner \(2010\)](#), this choice proves to be convenient since

it does not bound the posterior distribution of s_{ij}^2 artificially away from zero and provides significantly more shrinkage. Specifically, we assume a prior distribution for s_{ij}^2 which is given by

$$s_{ij}^2 \sim \mathcal{G}\left(\frac{1}{2}, \frac{1}{2B_s}\right), \quad j = 1, \dots, K_i, \quad (2.13)$$

where B_s is a scalar hyperparameter controlling the tightness of the prior. The normal prior on Ψ_{i0} and the set of Gamma priors on S_i allow us to achieve shrinkage along two important dimensions. First, the prior on the initial state provides the possibility of shrinking the parameters towards the prior mean which we assume to be zero. Second, the Gamma prior can be specified such that the model is effectively pushed towards a constant coefficient specification a priori, therefore allowing to control the degree of variation of the autoregressive parameters. To see how this prior setup exerts shrinkage on the state variances, note that it is straightforward to show that the Gamma prior on s_{ij}^2 induces a normal prior on the signed standard deviations $\pm\sqrt{s_{ij}^2}$ centered on zero with variance B_s . Smaller values of B_s push the specification *a priori* towards a constant parameter model.

A set of normal priors are imposed on the initial state of $a_{i0,t}, a_{i0,0}$

$$\text{vec}(a_{i0,0}) \sim \mathcal{N}(\text{vec}(\underline{a}_i), \underline{V}_{a_i}), \quad (2.14)$$

where \underline{a}_i and \underline{V}_{a_i} denote the prior mean and prior variance covariance matrices of the initial state. Similarly to the prior on S_i , we impose a set of Gamma priors on the elements of V_i

$$v_{ir}^2 \sim \mathcal{G}\left(\frac{1}{2}, \frac{1}{2B_v}\right), \quad r = 1, \dots, l_i. \quad (2.15)$$

Here, we let B_v denote the shrinkage hyperparameter used to penalize variation in the covariances of the model.

Finally, we use the prior setup proposed in [Kastner and Frühwirth-Schnatter \(2013\)](#) and subsequently used in [Huber \(2016\)](#) on the coefficients of the log-volatility process in [equation \(2.8\)](#). A normal prior is imposed on μ_{il} ($l = 1, \dots, k_i$) with mean $\underline{\mu}_i$ and variance \underline{V}_{μ_i}

$$\mu_{il} \sim \mathcal{N}(\underline{\mu}_i, \underline{V}_{\mu_i}). \quad (2.16)$$

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

$$\frac{\rho_{il} + 1}{2} \sim \text{Beta}(a_0, b_0), \quad (2.17)$$

which implies

$$E(\rho_{il}) = \frac{2a_0}{a_0 + b_0} - 1,$$

$$\text{Var}(\rho_{il}) = \frac{4a_0b_0}{(a_0 + b_0)^2(a_0 + b_0 + 1)}.$$

For typical data sets arising in macroeconomics, the exact choice of the hyperparameters a_0 and b_0 in [equation \(2.17\)](#) is quite influential, since data do not tend to be very informative about the degree of persistence of log-volatilities.

We impose a non-conjugate gamma prior for ζ_{ij}^2 , ($j = 1, \dots, k_i$),

$$\zeta_{ij}^2 \sim \mathcal{G}\left(\frac{1}{2}, \frac{1}{2B_\zeta}\right). \quad (2.18)$$

Mirroring the properties of the prior used for the other state equations, this choice does not bound ζ_{il}^2 away from zero, thus providing more shrinkage than standard typical conjugate inverted gamma priors do. Moreover, such a prior setting can improve sampling efficiency considerably ([Kastner and Frühwirth-Schnatter, 2013](#)).

Using the prior setting described above, a Markov chain Monte Carlo (MCMC) algorithm to draw samples from the (country-specific) parameter posterior distribution can be designed. Let us denote the full history of the time-varying elements in [equation \(2.9\)](#) up to time T as

$$\text{vec}(\Psi_i^T) = (\text{vec}(\Psi_{i1})', \dots, \text{vec}(\Psi_{iT})')',$$

$$a_i^T = (a'_{i1}, \dots, a'_{iT})',$$

$$\lambda_i^T = (\lambda_{i1}, \dots, \lambda_{iT})'.$$

The MCMC algorithm consists of the following blocks

- $\text{vec}(\Psi_i^T)$ and a_i^T are sampled through the well known algorithm provided in [Carter and Kohn \(1994\)](#) and [Frühwirth-Schnatter \(1994\)](#).

Table 1: Country coverage of GVAR model

Europe	Other developed economies	Emerging Asia	Latin America	Mid-East and Africa
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)	New Zealand (NZ)	Malaysia (MY)	Mexico (MX)	
Finland (FI)	United States (US)	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)				

Notes: ISO-2 country codes in parentheses. Empirical results shown for countries in bold.

- Conditional on $\text{vec}(\Psi_i^T)$ and a_i^T , the variances in [equation \(2.6\)](#) and [equation \(2.7\)](#) can be sampled from a generalized inverse Gaussian distribution,³ i.e.,

$$s_{ij}^2 | \text{vec}(\Psi_i^T) \sim \text{GIG} \left(\frac{1}{2} - \frac{T}{2}, \sum_{t=1}^T (\Psi_{ij,t} - \Psi_{ij,t-1})^2, \frac{1}{2B_s} \right), \quad (2.19)$$

$$v_{ij}^2 | a_i^T \sim \text{GIG} \left(\frac{1}{2} - \frac{T}{2}, \sum_{t=1}^T (a_{ij,t} - a_{ij,t-1})^2, \frac{1}{2B_v} \right), \quad (2.20)$$

with $\Psi_{ij,t}$ denoting the j th element of the $\text{vec}(\Psi_{it})$.

- The history of log volatilities is sampled using the algorithm outlined in [Kastner and Frühwirth-Schnatter \(2013\)](#).⁴

3 Data, model specification and prior implementation

This section introduces the data and the priors placed on the parameters of the model framework. We extend the dataset used in [Dees et al. \(2007a,b\)](#) with respect to both variable coverage and time span. In our analysis we use quarterly data for 36 countries spanning the period from 1979:Q2 to 2013:Q4. The countries covered in our sample are shown in [Table 1](#).

³The corresponding proof can be found in the Appendix.

⁴Further details of the sampling algorithm by [Kastner and Frühwirth-Schnatter \(2013\)](#) can be found in the Appendix.

The country-specific TVPVAR-SV models include real GDP growth (Δy), inflation measured by the log-difference of the consumer price level (Δp), the log-difference of the real exchange rate (Δe) vis-à-vis the US dollar, short-term interest rates (i) and the term spread, constructed as the difference between long-term and short-term interest rates (s). Note that not all variables are available for each of the countries we consider in this study. However, with the exception of long-term interest rates (that are used to calculate the term-spread), the coverage of all variables is above 80%.⁵

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

$$\mathbf{x}_{it} = (\Delta y_{it}, \Delta p_{it}, \Delta e_{it}, i_{it}, s_{it})'. \quad (3.1)$$

We follow the bulk of the literature on GVAR modeling by including changes in oil prices (Δpoil) as a global control variable. With the exception of the bilateral real exchange rate, 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 1980 to 2003.⁶ For a typical country i the set of *weakly exogenous* and global control variables comprises

$$\mathbf{x}_{it}^* = (\Delta y_{it}^*, \Delta p_{it}^*, i_{it}^*, s_{it}^*, \Delta \text{poil})'. \quad (3.2)$$

The US model, which we normalize to correspond to $i = 0$, deviates from the other country specifications in that the oil price is determined within that country model and the change in the trade weighted real exchange rate (Δe^*) is included as an additional control variable, so that its vectors of endogenous and weakly exogenous variables are given by

$$\mathbf{x}_{0t} = (\Delta \text{poil}_t, \Delta y_{0t}, \Delta p_{0t}, i_{0t}, s_{0t})', \quad (3.3)$$

$$\mathbf{x}_{0t}^* = (\Delta y_{0t}^*, \Delta p_{0t}^*, \Delta e_{0t}^*, i_{0t}^*, s_{0t}^*)'. \quad (3.4)$$

⁵We also corrected for outliers in countries that witnessed extraordinarily strong crisis-induced movements in some of the variables contained in our data. We accounted for these potentially influential observations by smoothing the relevant time series after defining outliers as those observations that exceed 1.5 times the interquartile range in absolute value.

⁶Note that recent contributions (Eickmeier and Ng, 2015; Dovern and van Roye, 2014) suggest using financial data to compute foreign variables related to the financial side of the economy (e.g., interest rates or credit volumes). Since our data sample starts in the early 1980s, reliable data on financial flows – such as portfolio flows or foreign direct investment – are not available. 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.

We rely on the recursive identification design proposed – among others – in [Christiano et al. \(2005\)](#) to pin down the US monetary policy shock. For that purpose, we order the block of variables that do not react instantaneously to a monetary policy shock first ($\Delta poil_t, \Delta y_{0t}, \Delta p_{0t}$), followed by the policy instrument (i_{0t}) and a block that reacts immediately if the monetary policy shock hits the economy (s_{0t}) ([Christiano et al., 1999](#)). This ensures identification of the monetary policy shock.⁷ All results considered below are based on generalized structural impulse response functions that aim to identify the structural responses to the shock of interest while integrating out other shocks (see [Dees et al., 2007a](#)). This yields generalized impulse responses to shocks that are not explicitly identified and structural responses to the US monetary policy shock. For all countries considered, we set the lag length of endogenous and weakly exogenous variables equal to one. In principle and in a standard vector-autoregressive one-country model, such a parsimonious lag structure might not be able to adequately capture serial correlation of all variables in the model. However the GVAR model, due to its high cross-sectional dimension, may already allow for very complex dynamics of each modeled variable. In fact, a residual analysis based on the posterior median shows that the residuals tend to be serially uncorrelated for the vast majority of countries and variables considered.

Before proceeding to the empirical results, we discuss the specific choices of the hyperparameters needed to construct our prior distributions. Since the GVAR comprises $N + 1$ countries, each country could be endowed with a country-specific set of hyperparameters to elicit the prior. We simplify the elicitation of the prior by imposing equal hyperparameters across countries. For the prior over the initial state Ψ_{i0} , we set $\text{vec}(\underline{\Psi}_i) = 0$ and $\underline{V}_{\Psi_i} = 10I_{k_i m_i}$. Similarly we set $\text{vec}(\underline{a}_i) = 0$ and \underline{V}_{a_i} equal to a diagonal matrix with 10 on its main diagonal. This setup renders the prior on the initial conditions fairly uninformative and proves not to be influential in the empirical application.

The prior on the innovation variances of the state equations in [equation \(2.6\)](#) and [equation \(2.7\)](#) is set such that $B_v = B_s = 0.1$. Since this choice turns out to be highly relevant in practice, we perform an extensive prior sensitivity analysis. In contrast to [Primiceri \(2005\)](#), who elicits the prior on the variance of the state innovations using a pre-sample of data, we evaluate different hyperparameters on a grid of values, ranging from values which translate into a much tighter prior than [Primiceri \(2005\)](#)'s setup to a specification with a prior which is quite loose. Given that we are interested in allowing the data to be as informative as possible

⁷A number of alternative identification strategies of structural VARs have been proposed in the recent empirical literature, including the narrative approach ([Romer and Romer, 2004](#)), the use of high frequency data ([Gertler and Karadi, 2015](#)), or external information ([Stock and Watson, 2012](#)). [Miranda-Agrippino and Rey \(2015\)](#) contrast results using a recursive identification and external information available to the Fed around FOMC meetings, and find qualitatively similar impulse-responses in both cases. The identification used in this paper imposes the same recursive structure on all countries in the GVAR model.

with respect to the drifting behavior of the coefficients, we strongly favor hyperparameters that are loose. We still impose enough discipline on the parameter dynamics such that the resulting posterior quantities do not show explosive behavior. The grid of parameter values we evaluate is given by $(0.001, 0.01, 0.1, 0.5, 1, 4)$, where we pick 0.1 as our reference value for both B_v and B_s . In our experience, higher values should be avoided since they often lead to excessively unstable posterior draws.

Finally, the prior on the mean of the log-volatility equation is set such that $\underline{\mu}_i = 0$ and $\underline{V}_{\mu_i} = 10$, which is uninformative given the scale of our data. For the autoregressive parameter ρ_{il} we set a_0 and b_0 equal to 25 and 1.5, respectively. This prior places a lot of mass on high persistence regions of the parameter space. Since the data are usually not informative about the autoregressive parameter corresponding to latent factors, the posterior distribution can be significantly shaped by this choice. A sensitivity analysis using hyperparameters that place more prior mass on stationary regions of ρ_{il} leads to qualitatively similar results to those presented in this section. The last piece missing is the prior on ς_{il} , where we only have to elicit B_σ , which is set equal to unity.

We compute all relevant quantities by performing Monte Carlo integration by drawing 1,500 samples from a total chain of 30,000 draws, where the first 15,000 draws are discarded. Standard diagnostic checks indicate convergence to the stationary distribution, with inefficiency factors for the autoregressive coefficients and volatilities well below 20 for most country models.⁸

4 The international dimension of US monetary policy

Using the estimated TVP-SV-GVAR, we investigate in this section how US monetary policy affects international output, prices, short-term interest rates and exchange rates. In a second step, we then investigate the drivers of these spillovers. Finally, we ask the reverse question and analyze whether and how US monetary policy responds to foreign shocks.

4.1 Does the global economy respond to US monetary policy shocks?

First, we analyze international effects of US monetary policy using the newly developed TVP-SV-GVAR model. In contrast to existing literature, we are not only able to assess cross-country differences but whether spillovers have changed over time. Hitherto the empirical literature using linear models has found significant effects of US monetary policy on global output.

⁸Further information on the convergence properties of the sampler for our empirical application can be found in the Appendix.

Most studies assessing the effects of macroeconomic shocks in the US economy on the world use either stylized linear two-country vector autoregressions (see for example [Kim, 2001](#); [Canova, 2005](#)) or systems of country-specific models. Both approaches have been mostly confined to linear models with fixed parameters and are thus not able to track changes in the transmission channel or the external environment. [Canova \(2005\)](#), for example, finds large and significant output responses to US monetary policy shocks in Latin America. In line with the results in [Kim \(2001\)](#), the transmission tends to be driven by the strong response of domestic interest rates to US monetary expansions rather than by the trade balance. [Ehrmann and Fratzscher \(2009\)](#) show that US monetary policy shocks impact strongly on short-term interest rates and ultimately on equity markets in a large number of economies. Several recent contributions draw on the framework put forth in [Pesaran et al. \(2004\)](#) and use a global system of vector autoregressions to investigate the propagation of different monetary and fiscal policy shocks across the globe (see for instance [Dees et al., 2007a; 2010](#); [Feldkircher and Huber, 2016](#)). Employing this framework and using a Bayesian set-up, [Feldkircher and Huber \(2016\)](#) find significant and rather persistent spillovers from US monetary policy shocks on international output. Examining conditional forecasts of different future policy paths for the federal funds rate, [Feldkircher et al. \(2015\)](#) find strong direct spillover effects for output in emerging economies, while second-round effects play a more prominent role in advanced economies.

We investigate the international responses to an unexpected US monetary policy tightening normalized to 100 basis points (bp) throughout the sample period. While the shock on impact is fixed to 100 bp for the US, spillovers generated by the shock are allowed to vary if macroeconomic relationships or residual variances change over time. A hypothetical monetary policy shock during the period of the global financial crisis, when economic and financial conditions are weak and macroeconomic uncertainty is high, might impact differently on international output than during tranquil times, warranting a time-varying parameter framework. The results are summarized in Figures 1 to 4, which show the posterior mean of the corresponding (cumulative) impulse response for selected countries, along 25% and 75% percentiles of the posterior distribution of the cross-country means (gray shaded regions). These can be interpreted as the uncertainty surrounding the impulse responses of a typical country from a given region. Responses are shown over the whole sample period and for the one and eight quarter forecast horizon.

Figure 1 shows the cumulative response of output to the monetary policy shock originated in the US. The estimated effects for the US economy itself are in line with the empirical literature on US monetary policy (see, e.g., [Christiano et al., 1999](#); [Coibion, 2012](#)). In most economies, including the US itself, output contracts and responses tend to be rather persistent,

corroborating the findings by [Feldkircher and Huber \(2016\)](#), who use a linear, time-invariant version of the Bayesian GVAR model. Looking at different world regions, most responses are very homogeneous and fall inside the credible sets spanned by the respective cross-country means. Canada shows a very pronounced negative response that is even stronger than the domestic reaction of output in the US itself. Also for Germany and Japan, spillovers of the US monetary policy shock are strong. Countries with positive impact responses and that deviate from their regional peers include Australia, Indonesia, Mexico and Spain. While in the medium term, Australia and Indonesia seem rather isolated from the shock with responses hovering around zero, the response in Mexico and Spain becomes negative and thus in line with that of their regional peers. Considerable time variation is evident from the graphs. Starting in the mid-1990s, most economies show a downward-trending medium-term response, which reaches a trough around the episode of the global financial crisis, after which responses become less pronounced again. Taken at face value, this finding reveals stronger effects on international output in the most recent part of our sample as compared to earlier periods.

Figure 2 shows the cumulative effects on international inflation. Responses in other developed and Western European economies are very homogeneous. They are mostly negative in the short-run and peter out very quickly. By contrast, there is considerably more variation across countries in emerging Asia and Latin America. In emerging Asia, the monetary policy tightening tends to trigger negative reactions of prices on impact, while responses in the medium-term are close to zero. Showing an increase of prices, Indonesia is an exception in the region. Impact responses in Latin America tend to be positive and pronounced for all countries but Argentina. Responses are rather persistent, especially for Chile and Peru (positive) and Argentina (negative). Modeling changes in parameters and variances seems particularly important when assessing spillovers for emerging markets, which show more pronounced reactions during the 2000s and smaller responses in the aftermath of the global financial crisis.

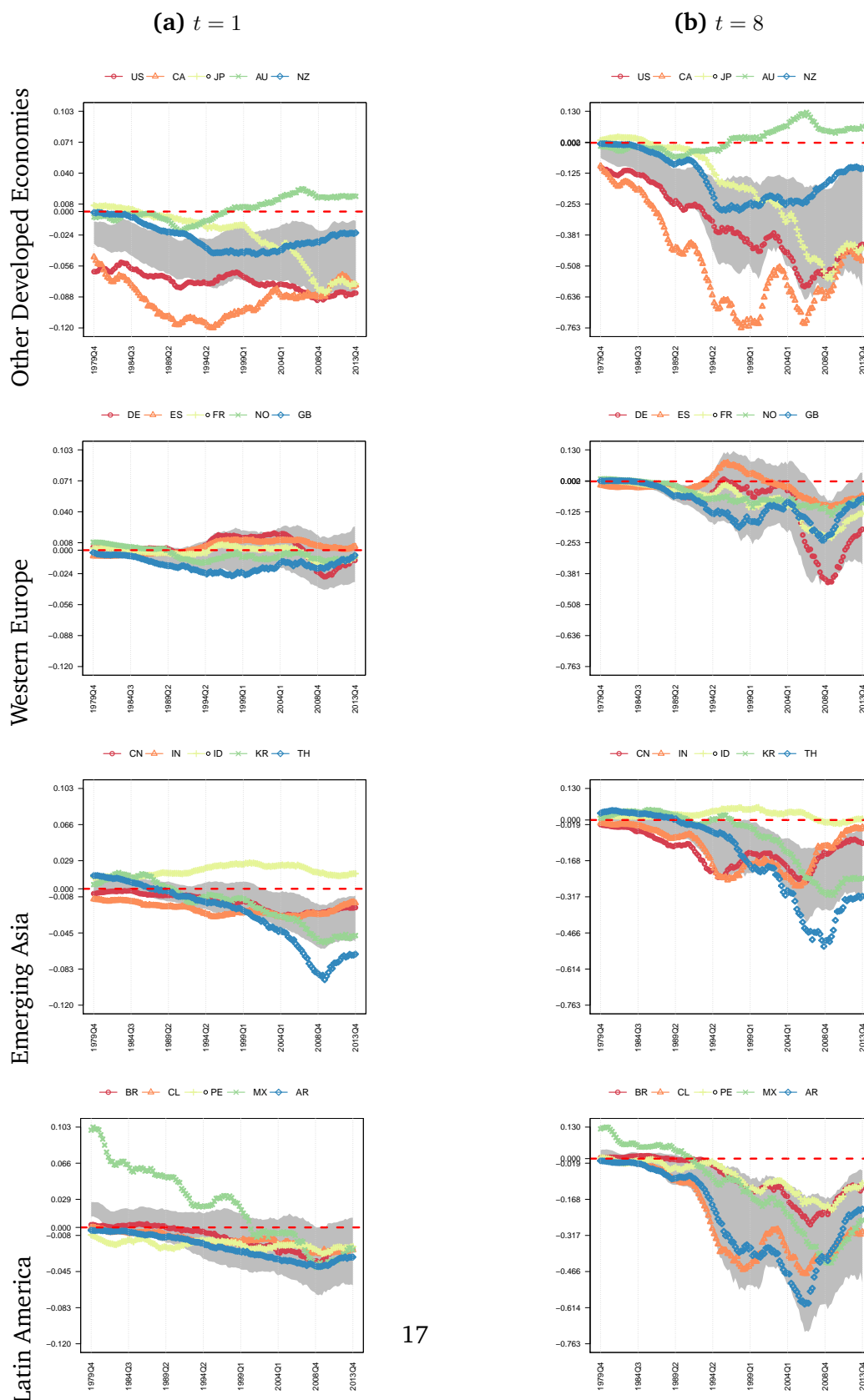
Figure 3 shows the (non-cumulative) response of interest rates following the monetary policy shock. Using a simpler specification than that employed here, comovements of interest rates have been identified as an important transmission channel of macroeconomic shocks in [Feldkircher and Huber \(2016\)](#). Indeed, in the short run, almost all countries follow the US rate hike. After eight quarters, the direct effect on domestic interest rates has practically disappeared in most countries. Countries that show a particularly pronounced behavior include Canada, Great Britain, India and Latin American economies. This implies that the interest rate reaction is large in countries that share strong economic ties with the US, as well as fast-growing emerging economies, that have been also hit in the past strongly when the Fed announced interest rate changes (see, e.g., the taper tantrum episode in mid-2013).

The importance of the financial channel in transmitting US shocks for Latin America has also been highlighted in [Canova \(2005\)](#). In line with results on international prices, most countries tend to show stronger medium-term responses during the mid-1990s to mid-2000s and comparably smaller responses in the most recent period of the sample.

Next, [Figure 4](#) shows the responses of the real exchange rate vis-à-vis the US dollar. As expected, responses tend to be positive on impact indicating a real appreciation of the US dollar as a consequence of the stronger increase in domestic interest rates. Naturally, advanced countries that are strongly linked to the US in economic terms respond more strongly to the monetary policy shock. These include Australia, Canada and New Zealand and, to a lesser extent, Great Britain. Among emerging economies, currencies that weaken against the US dollar include those of Korea, Brazil and Chile. The depreciation of the domestic currency is more pronounced in response to a hypothetical monetary policy shock hitting the economies in the 2000s than in the early part of our sample. Moreover, responses became smaller in the aftermath of the global financial crisis for all currencies.

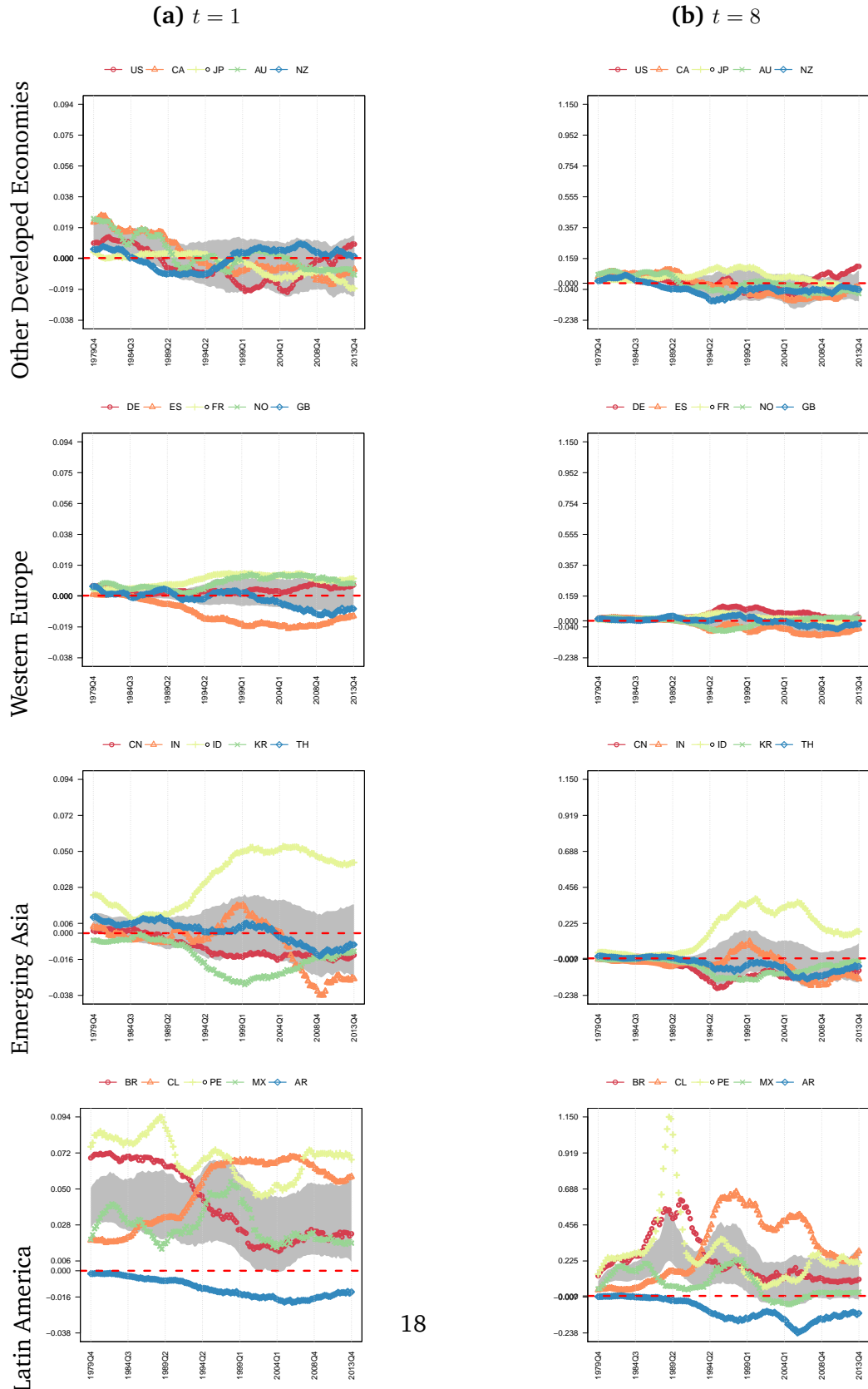
To complement the analysis, [Figures 5 and 6](#) depict the residual variance estimates for the variables discussed above. Our modeling framework provides also explicit inference on the dynamics of macroeconomic volatility. As an example, [Figure 5](#) plots the volatility of GDP growth, mean-standardized in order to facilitate cross-country comparison. A decline in the volatility of GDP growth in Western Europe and other developed economies can be observed until the middle of the 2000s, a development which is in line with the dampening of real fluctuations corresponding to the Great Moderation period. After 2007, a sharp increase in output growth volatility due to the outbreak of the global financial crisis can be observed, followed by a gradual return to lower volatility more recently.

Figure 1: Output responses to a +100 basis point (bp) US monetary policy shock



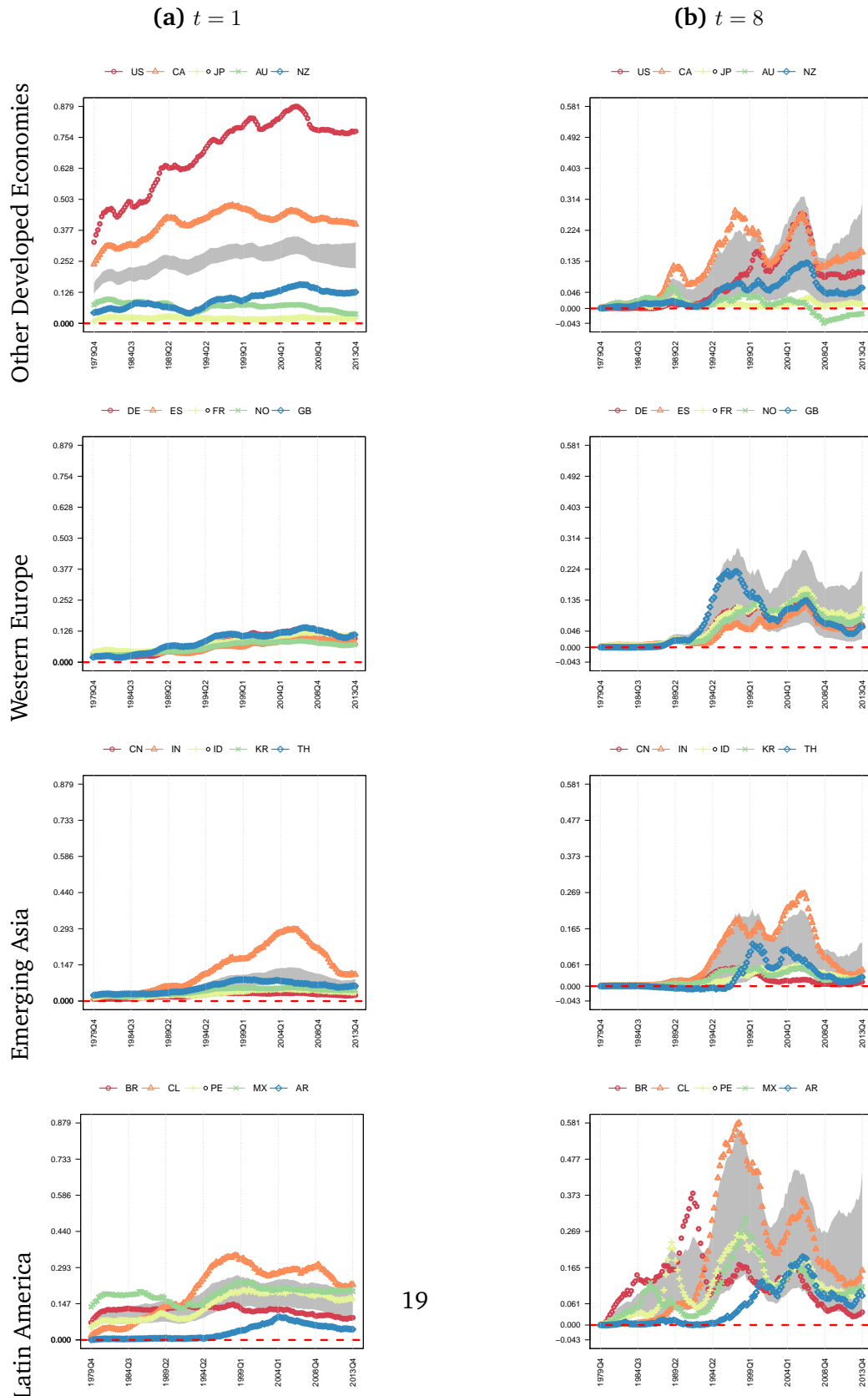
Notes: The plots show the posterior for selected countries along with 25% and 75% percentiles of the posterior distribution of the cross-country mean (gray shaded regions) after one and eight quarters. Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

Figure 2: Price responses to a +100 basis point (bp) US monetary policy shock



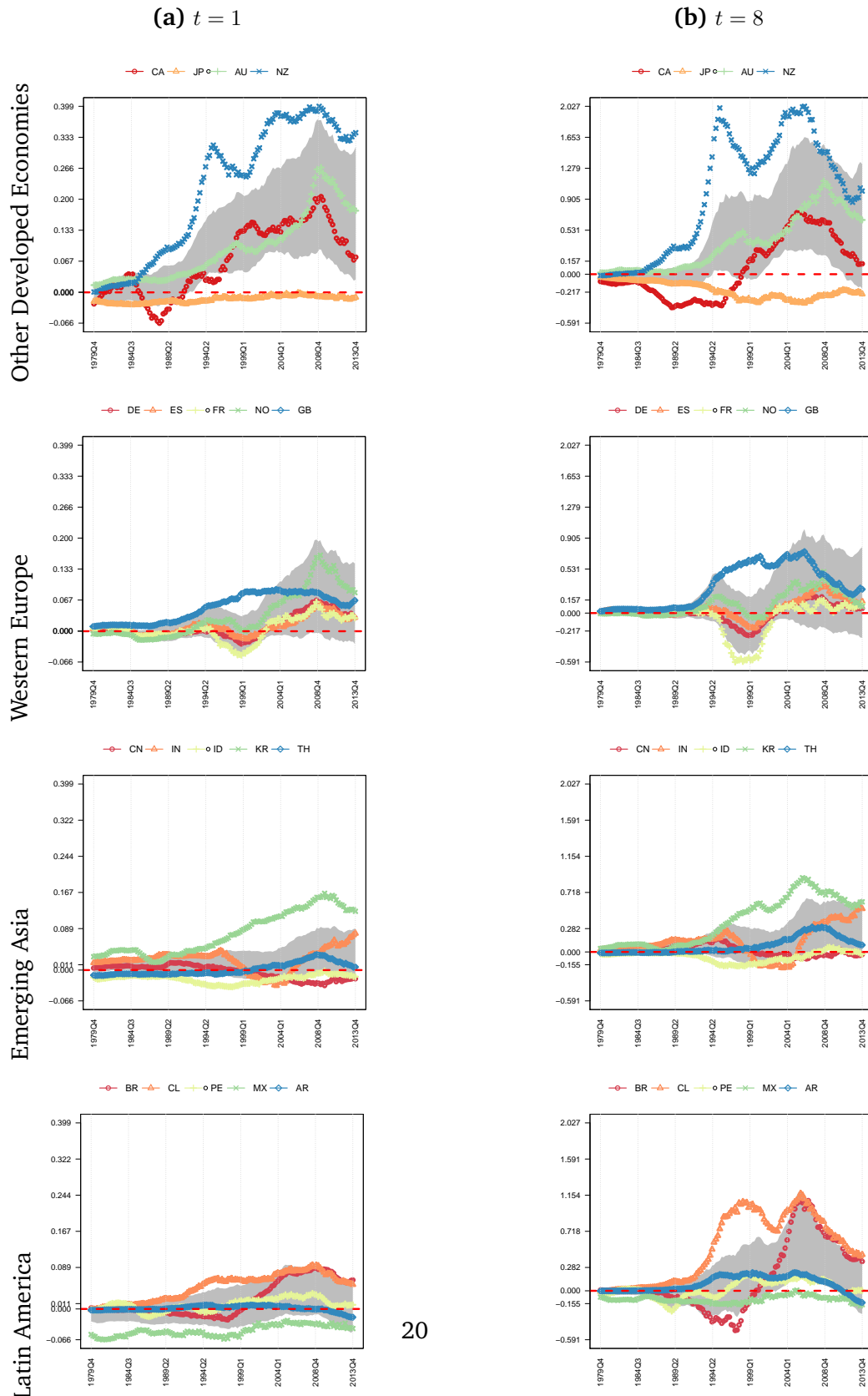
Notes: The plots show the posterior for selected countries along with 25% and 75% percentiles of the posterior distribution of the cross-country mean (gray shaded regions) after one and eight quarters. Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

Figure 3: Short-term interest rate responses to a +100 basis point (bp) US monetary policy shock



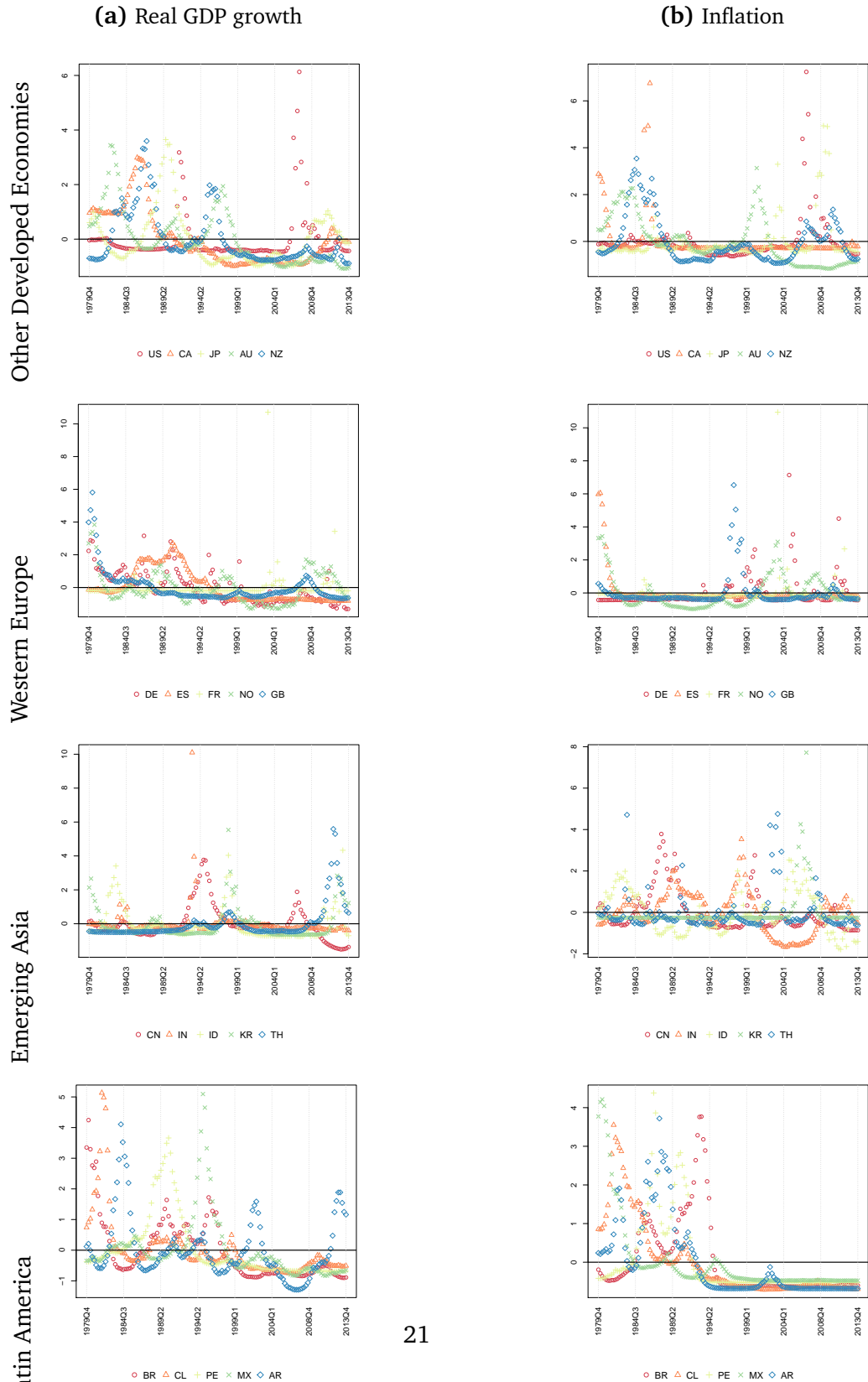
Notes: The plots show the posterior for selected countries along with 25% and 75% percentiles of the posterior distribution of the cross-country mean (gray shaded regions) after one and eight quarters. Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

Figure 4: Real exchange rate responses to a +100 basis point (bp) US monetary policy shock



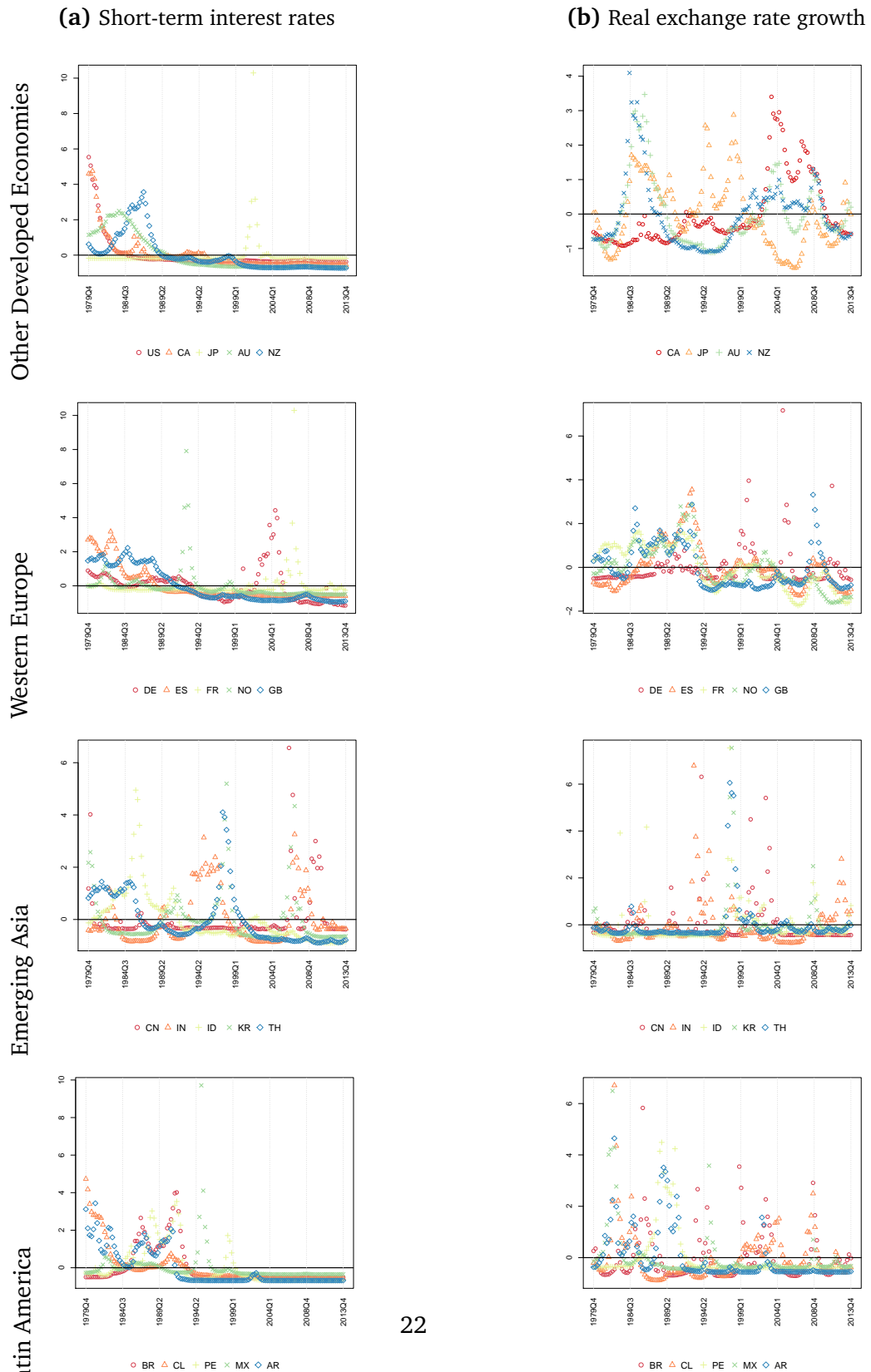
Notes: The plots show the posterior for selected countries along with 25% and 75% percentiles of the posterior distribution of the cross-country mean (gray shaded regions) after one and eight quarters. Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

Figure 5: Stochastic volatility over time



Notes: the plots depict the posterior mean of standardized volatility across regions over the estimation sample. Results based on 1,500 posterior draws from a total chain of 30,000 draws

Figure 6: Stochastic volatility over time



Notes: the plots depict the posterior mean of standardized volatility across regions over the estimation sample. Results based on 1,500 posterior draws from a total chain of 30,000 draws

Economies in Latin America and Asia witnessed episodes of increased volatility of GDP growth also during crises in the 1980's and 1990's, respectively. In some emerging economies (Thailand, Korea and Argentina) volatility following the global financial crisis increased sharply. Volatility spikes in other variables occur more frequently in emerging Asia and Latin American economies, while they are less frequent in advanced economies. The timing of the spikes also differs. For example, high volatility attached to inflation spikes is a common phenomenon in the early 1980s in Latin America, when some countries witnessed periods of hyperinflation. On the other hand, residual variance increases sharply in advanced countries around the period of the global financial crisis, which was marked by deflationary pressures. Naturally, among all variables considered, volatility of real exchange rates exhibit the highest variability for all regions, including advanced economies. Overall, our model framework correctly identifies periods of heightened uncertainty such as the Asian and the global financial crisis.

Summing up, we find that a US monetary tightening tends to decrease international output with effects which are visible even after eight quarters, while prices tend to decrease in the short-term but adjust quickly thereafter. International interest rates tend to follow the US rate hike and most currencies weaken against the US dollar in response to the tightening. In the medium-term, responses tend to be more pronounced for those countries that share strong economic links with the US and for emerging economies. We also find relevant time variation in international spillovers. With the exception of output, most responses tend to be large during the mid-1990s to mid-2000s, while effects tend to get smaller in the most recent period of the sample. Output responses appear more pronounced during the period of the global financial crisis and less so afterwards. The residual variance component of our model correctly identifies known periods of heightened uncertainty in the past.

4.2 Why are some countries more affected by international spillovers than others?

So far we have established that spillovers from US monetary are significant, time-varying and different across countries. In this section we assess explicitly the drivers of differences in the size of international spillovers across economies.

To investigate the determinants of cross-country differences of spillovers in a systematic fashion, it proves useful to recall the uncovered interest (UIP) rate parity condition in its approximate form, which states that, under a no-arbitrage condition, domestic interest rates should equal foreign interest rates bar the expected appreciation of the domestic currency. It is easy to see that an unexpected increase in US short term rates should thus have a direct impact on either the nominal exchange rate or domestic interest rates, thereby affecting the domestic macroeconomy. Other factors such as country risk premia, that affect differences in

investment sentiment across countries of the world, are also expected to affect the degree of international spillovers affecting a particular economy in a given point of time.

We aim at explaining differences in the strength of international spillovers using linear panel regressions of the form:

$$z_{it} = \alpha_t + \gamma_i + \beta_s X_{si,t} + u_{it}, \quad (4.1)$$

with z_{it} denoting absolute cumulative spillovers averaged per year, α_t and γ_i are time and country fixed effects, respectively. $X_{si,t}$ is a matrix containing s explanatory variables from the pool of potential covariates and u_{it} is a normally distributed error term with fixed variance. In order to account for specification uncertainty and ensure the robustness of the results concerning the determinants of spillovers, we make use Bayesian model averaging (BMA) techniques to perform inference in this setting.⁹

We collect data on a series of potential covariates explaining the intensity of spillovers, including information on factors related to exchange rate stability, macroeconomic and fiscal vulnerabilities, financial depth and stability, as well as financial and trade openness. A detailed description of the data used can be found in [Table B.1](#) of the appendix. These variables are then related to the absolute value of the cumulative international impulse responses from [section 4](#). A similar econometric exercise using a linear, time-invariant version of the GVAR has been recently provided in [Georgiadis \(2015\)](#). Since the data are on annual rather than quarterly frequency, we take yearly averages of the absolute cumulative impulse responses. Due to data availability, we have to limit the time span for the panel regressions to 1995-2011 and drop Great Britain, Sweden, Chile, Indonesia and Saudi Arabia from the set of countries. That leaves us with a balanced panel of 17 yearly observations for 31 countries and 26 potential explanatory variables.

We focus on variables that receive a posterior inclusion probability which is larger than 0.5 (in bold) and whose parameter estimates have a ratio of posterior mean to posterior standard deviation above 1.3 in absolute terms. These measures ensure that the variable is an important determinant of international spillovers and its effect is precisely estimated (see [Barbieri and Berger, 2004](#); [Masanjala and Papageorgiou, 2008](#)).

The results provided in [Table 2](#) suggest that there is no single determinant that explains spillovers equally well for all variables. Nevertheless there are some general patterns that emerge from the data. First, measures of trade and financial openness tend to be related to the extent of spillovers. International spillovers are stronger for countries with either a high

⁹We use the benchmark prior of [Fernández et al. \(2001\)](#) on the parameters and a binomial beta prior on the model space ([Ley and Steel, 2009](#)).

Table 2: Empirical determinants of international spillovers

	Real GDP growth		Inflation		Real exchange rate change		Short-term interest rate		Term spread	
	PIP	Post Mean	PIP	Post Mean	PIP	Post Mean	PIP	Post Mean	PIP	Post Mean
Exchange rate	0.0311	0.0000	0.0181	0.0000	0.3078	0.0000	0.0331	0.0000	0.0219	0.0000
ER Volatility	0.1697	-0.0191	1.0000	-0.5593*	0.0643	-0.0117	1.0000	-0.5821*	0.0203	-0.0014
FX Exposure	0.0256	0.0000	0.0455	0.0005	0.6979	0.0414*	0.0812	0.0004	1.0000	0.2518*
FX Reserves	0.0228	-0.0001	0.2006	0.0062	0.1863	-0.0128	0.0628	-0.0004	0.9546	0.2101*
Min Deviation	0.0419	0.0012	0.0306	-0.0005	0.0196	0.0007	0.9559	-0.0694*	0.0198	0.0005
Max Deviation	0.0461	0.0004	0.1991	0.0152	0.0423	-0.0019	0.9985	0.0972*	0.0191	0.0002
Asset Exposure	0.0458	-0.0002	0.2072	0.0012	0.0663	0.0006	0.5920	-0.0021	0.0358	-0.0007
Zero Change	0.0683	0.0000	0.0311	0.0000	0.0465	0.0000	0.0672	0.0000	0.0300	0.0000
Base Exchange Rate	0.0267	0.0000	0.0187	0.0000	0.0457	0.0000	0.1393	0.0000	0.0213	0.0000
Range	0.7588	-0.0175*	1.0000	0.1016*	0.7704	0.0368*	1.0000	0.0481*	0.0191	-0.0001
Current Account	0.8143	0.0585*	0.0547	-0.0019	0.4049	-0.0470	0.0341	0.0001	0.0269	-0.0021
Fiscal Deficit	0.5312	0.0005	0.3954	-0.0003	0.2650	0.0004	0.9611	-0.0004*	0.2079	-0.0006
Government Debt	0.9121	0.0002*	0.0232	0.0000	0.0177	0.0000	0.9090	0.0001*	0.0459	0.0000
Gross Savings	0.2829	-0.0002	1.0000	0.0013*	0.0695	-0.0001	0.9904	0.0004*	0.9999	-0.0084*
Bank Credit to Deposits	0.1952	0.0000	0.0950	0.0000	0.0935	0.0000	0.0907	0.0000	1.0000	-0.0009*
Deposit Money	0.0678	0.0000	1.0000	-0.0004*	0.0638	0.0000	0.6039	-0.0001	0.9373	0.0007*
Financial Deposits	0.0450	0.0000	0.8740	0.0002*	1.0000	0.0013*	0.1376	0.0000	0.9709	-0.0014*
Liquid Liabilities	0.0637	0.0000	0.0569	0.0000	0.9833	-0.0008*	0.1066	0.0000	0.0238	0.0000
Private Credit	0.9517	-0.0002*	0.9989	0.0004*	0.1205	0.0000	0.7169	0.0001	0.0375	0.0000
FDI Assets	0.0347	-0.0001	0.0314	0.0001	0.1548	0.0032	0.4674	0.0032	0.0249	-0.0003
Capital Restrictions (inflow)	0.8696	-0.0165*	0.0570	-0.0005	0.9631	-0.0454*	0.1788	0.0008	0.1264	0.0100
Capital Restrictions (outflow)	0.0496	-0.0004	0.0205	0.0000	0.0204	0.0001	0.0376	0.0000	0.1175	-0.0073
Portfolio Assets	0.0491	-0.0004	0.0728	0.0009	0.1688	0.0057	0.2892	0.0026	0.0403	0.0018
Portfolio Liability	0.9997	-0.0347*	0.0301	0.0002	0.0152	0.0000	0.1545	-0.0007	0.9982	0.1452*
Foreign Liabilities	0.1872	0.0006	0.0494	0.0001	0.0309	0.0001	0.1495	0.0002	0.0734	-0.0021
Foreign Assets	0.7272	0.0033	0.1411	0.0004	0.0423	0.0001	0.5176	0.0013	0.9406	-0.0404*
Trade Openness	0.9999	0.0003*	0.5359	0.0001	0.0189	0.0000	0.0304	0.0000	0.0362	0.0000

Notes: The table reports results of five separate pooled panel estimations (including country and time fixed effects) of spillovers to international output, inflation, the real exchange rate, short- and long-term interest rates. Results obtained by Bayesian model averaging using 2 million draws as burn-ins and 3 million posterior draws, a binomial beta prior on the model space and Zellner's g prior on the coefficients. Variables with posterior inclusion probabilities (PIP) above 0.5 in bold, posterior means with asterisk denote standardized coefficients (posterior mean / posterior deviation) that exceed 1.3 in absolute terms. All calculations were carried out with the R package BMS described in [Zeugner and Feldkircher \(2015\)](#).

share of exports and imports in percent of GDP or with few restrictions on their capital account. Implementing restrictions on capital (in-)flows might thus serve as a way to shield the domestic economy from international shocks. Evidence for other variables related to financial openness depend on the spillover variable under consideration. For example, countries with a high share of foreign assets to GDP experience stronger spillovers on domestic output, whereas a high share of portfolio liabilities to GDP seem to act as a buffer to this effect. The reverse holds true for spillovers to the term spread. That said, specifications which include variables from the group of financial openness measures receive large posterior mass in the model space, implying that spillovers are shaped by these measures. Second, measures of exchange rate volatility of the domestic currency tend to be related to spillovers; higher volatility implies smaller spillovers. This is a direct consequence of the UIP, which predicts that adjustment through exchange rates should lead to a less complete pass-through of international interest rates. Also a high share of foreign exchange reserves or more generally foreign exchange exposure of the economy translates into higher spillovers from changes in US monetary policy. Third, macroeconomic vulnerabilities amplify spillovers. A high current account deficit or limited fiscal space, for instance, tend to be related to stronger spillovers. Also, a high share of gross savings – which is frequently a characteristic of oil or gas exporters, as well as emerging economies – tends to be related with larger spillovers. Last, evidence for variables related to financial stability and financial depth is mixed. As pointed out in [Georgiadis \(2015\)](#), the effect of deep financial markets on spillovers is theoretically ambiguous. On the one hand, in deep financial markets the credit channel might play a more important role to transmit monetary policy shocks. On the other hand, when financial deepening is high, competitiveness might lead to a more efficient financial system, mitigating financial accelerator effects. These ambiguous theoretical considerations are mirrored in our estimation results, which indicate different impacts depending on the spillover under consideration.

Summing up, we find that the size of international spillovers appear robustly related to how strongly that country is integrated with the world economy in terms of trade, the volatility of its exchange rate, the degree of restrictions on its capital account, the share of foreign exchange reserves and foreign currency exposure, and macroeconomic vulnerabilities including the current account deficit and the fiscal space left to adjust to shocks. The fact that our modeling strategy is able to exploit within-country variation in spillovers and accounts explicitly for model uncertainty lends particular confidence to these results.

4.3 Does the US Fed respond to international shocks?

In this section we investigate if and how US interest rates react to foreign shocks. We perform a set of simple counterfactual exercises by estimating the response of US interest rates to four distinct regional shocks, namely an increase of 100bp in short-term interest rates, a rise in inflation (by 1 percentage point), a deceleration of real GDP growth (by 1 percentage point) and a 1 percent real appreciation of the US dollar. These shocks are assumed to happen simultaneously in either Western Europe, Asia and Latin America. We do not compute a regional shock for the group of "other developed economies" since this group is rather heterogeneous and from an economic perspective it seems unlikely that these economies are hit by similar regional shocks.

The impact of these shocks on US short-term interest rates is quantified by generalized impulse response functions (GIRFs) as proposed in [Pesaran and Shin \(1998\)](#). GIRFs are appealing since they are insensitive to the ordering of the variables in the system, while the shocks in general remain (weakly) correlated, which strictly speaking prohibits a structural and economic interpretation. In practice, however, residual correlation is weak, especially when using a GVAR approach, since the weakly exogenous variables absorb a lot of the existing correlation.

The results are depicted in [Table 3](#). The table presents the posterior estimates of the responses of the US short-term interest rates, averaged across different periods corresponding to the mandates of the Fed's chairmen Volcker, Bernanke and Greenspan.

Table 3: Posterior distribution of US short-term interest rates responses to four regional shocks (GIRF).

		(a) Western Europe								
		"Volcker" regime 1979 - 1987			"Greenspan" regime 1987 - 2006			"Bernanke" regime 2006 - 2013		
Shock to		Low _{0.25}	Median	High _{0.75}	Low _{0.25}	Median	High _{0.75}	Low _{0.25}	Median	High _{0.75}
<i>i</i>	<i>t</i> = 1	6.7	41.9	95.5	22.7	64.8	132.5	32.6	94.7	171.9
	<i>t</i> = 8	-465.8	-294.6	-160.0	-503.6	-327.5	-182.2	-567.9	-362.1	-189.8
Δp	<i>t</i> = 1	12.9	43.9	79.6	33.3	73.7	120.6	5.6	35.9	65.1
	<i>t</i> = 8	-168.7	-54.2	78.1	-228.8	-85.7	50.8	-182.4	-76.8	12.7
Δy	<i>t</i> = 1	-165.8	-122.0	-79.0	-134.1	-97.1	-64.3	-126.6	-91.1	-58.2
	<i>t</i> = 8	-214.5	-107.2	-38.4	-156.0	-79.0	-20.7	-138.8	-62.6	-15.8
<i>e</i>	<i>t</i> = 1	-0.5	4.0	8.2	-0.0	3.9	7.3	-0.0	4.0	7.4
	<i>t</i> = 8	-9.8	-2.2	5.8	-8.0	-1.6	4.4	-6.8	-0.4	5.7

		(b) Asia								
		"Volcker" regime 1979 - 1987			"Greenspan" regime 1987 - 2006			"Bernanke" regime 2006 - 2013		
Shock to		Low _{0.25}	Median	High _{0.75}	Low _{0.25}	Median	High _{0.75}	Low _{0.25}	Median	High _{0.75}
<i>i</i>	<i>t</i> = 1	-7.2	-3.0	1.2	-7.0	-2.2	3.1	-25.9	-13.2	-4.9
	<i>t</i> = 8	-43.2	-25.9	-13.2	-59.3	-36.1	-20.2	-80.0	-46.3	-26.3
Δp	<i>t</i> = 1	20.0	42.1	66.0	18.0	36.3	59.2	17.3	32.8	52.8
	<i>t</i> = 8	-106.2	-54.8	-3.4	-107.8	-59.0	-9.4	-85.6	-43.8	-6.1
Δy	<i>t</i> = 1	-120.6	-87.9	-58.5	-132.2	-95.8	-59.4	-125.9	-91.6	-56.2
	<i>t</i> = 8	-174.6	-106.6	-59.2	-209.4	-132.2	-70.7	-210.3	-127.8	-67.4
<i>e</i>	<i>t</i> = 1	-10.5	-5.0	-0.1	-10.1	-5.0	-0.5	-12.9	-6.5	-1.8
	<i>t</i> = 8	-23.0	-9.4	0.6	-16.2	-5.0	4.7	-11.2	-0.4	10.8

		(c) Latin America								
		"Volcker" regime 1979 - 1987			"Greenspan" regime 1987 - 2006			"Bernanke" regime 2006 - 2013		
Shock to		Low _{0.25}	Median	High _{0.75}	Low _{0.25}	Median	High _{0.75}	Low _{0.25}	Median	High _{0.75}
<i>i</i>	<i>t</i> = 1	-2.4	-0.9	0.4	-4.3	-2.0	0.3	-19.5	-9.1	-0.6
	<i>t</i> = 8	-0.8	3.3	8.6	-6.6	-0.6	5.7	-34.2	-11.0	5.3
Δp	<i>t</i> = 1	-9.8	3.8	19.0	-3.8	9.1	22.0	-3.8	15.0	34.9
	<i>t</i> = 8	-40.9	-1.9	25.5	-34.0	2.6	28.9	-45.4	0.9	40.4
Δy	<i>t</i> = 1	-60.1	-44.5	-28.7	-67.4	-50.1	-32.1	-78.1	-56.5	-36.8
	<i>t</i> = 8	-91.0	-56.5	-30.9	-99.6	-63.3	-33.6	-111.3	-67.7	-34.6
ER	<i>t</i> = 1	-2.8	0.8	5.1	-3.6	0.3	3.8	-1.5	2.1	6.5
	<i>t</i> = 8	-14.7	-6.4	2.1	-19.2	-9.2	-1.9	-22.3	-11.2	-4.7

Notes: The table presents the posterior distribution of generalized impulse response functions (GIRFs) to an increase in regional interest rates (*i*), inflation (Δp), a decrease in output growth (Δy) and a depreciation of regional currencies against the US dollar (*e*). Responses are based on 1,500 posterior draws from a total chain of 30,000 iterations and in basis points. Responses for which credible sets do not include a zero value in bold.

Several findings are worth emphasizing. First, looking at the median responses to an increase in regional interest rates, we find that short-term rates in the US tend to increase in the short-run when the shock originates in Western Europe and tend to decrease when the shock comes from emerging economies in Asia and Latin America. Eight quarters ahead, the response of US short-term rates is negative to all three regional shocks. That is, if policy rates are raised abroad the Fed tends to cut domestic rates, probably to stabilize output and to

compensate for the shortfall in foreign demand. This is in contrast to our results on spillovers generated by US monetary policy shocks, where international short-term rates tend to follow the US rate hike even in the medium-term. Concentrating on the variation of responses over time, the effects on interest rates have increased on the back of stronger financial globalization. With the exception of those corresponding to shocks in Latin America, effects are tightly estimated for almost all periods. Wide credible sets for Latin American shocks might be driven by the fact that many countries in that region have historically resorted to fixed exchange rate regimes vis-a-vis the US dollar and thus would rather follow than lead US rate changes.

Considering an increase in foreign inflation next, the posterior median indicates that domestic interest rates in the US tend to increase in the short run when shocks originate in Western Europe or Asia. After eight quarters, US rates decrease, enhancing domestic demand. Effects are well estimated in the short-run when shocks originate in Western Europe and even in the medium-term when the shock origin is Asia. Responses to inflation shocks from Latin America are accompanied by wide credible sets for all the time periods considered.

Next we consider responses to decreases in foreign output growth. Here, estimates based on the posterior median point to an immediate rate decrease and rather persistent effects. Responses are well estimated for all three time periods and regardless of where the shock originates.

Last, we assess how US rates respond to a regional strengthening of the US dollar. Here, the immediate response of US short-term rates depends strongly on where the shock comes from. If the US dollar appreciates against a basket of Western European and Latin American economies, short-term interest rates tick up in the short-run. By contrast, if the shock originates in Asia, the Fed responds by lowering short-term rates. In the medium term, short-term rates decrease in response to all three shocks but credible sets are wide for all regions with the exception of Latin America in the two most recent periods of our sample. Here, with the exception of the Volcker regime, interest rates medium-term responses are tightly estimated and negative.

Summing up, we find significant responses of US short-term interest rates to an increase in foreign interest rates as well as to a decrease in foreign output regardless of the region of shock origin. The cut in domestic interest rates drives up economic growth thereby compensating for the fall in foreign demand. Responses to other shocks depend on their origin. The responses which are estimated best correspond to shocks from Asia, which includes China. Here, US rates also respond to an exchange rate shock in the short-run and to a shock to inflation in the medium-term.

5 Closing remarks

This paper analyzes the interlinkages of US monetary policy and the global economy. For that purpose we develop a time-varying parameter global vector autoregression with stochastic volatility (TVP-SV-GVAR). We use this framework to assess spillovers originating from disturbances to US monetary policy on a country-by-country basis taking explicitly into account that the extent of spillovers might have changed over time. We further assess the drivers of these spillovers and whether and how US interest rates respond to international shocks.

We find 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 is more persistent than transitory, a result which is in line with [Feldkircher and Huber \(2016\)](#). Following the response of the US, global inflation rates tend to decrease immediately and adjust quickly in the medium term. Also global short-term interest rates tend to follow their US counterparts increasing in response to the US rate hike. Naturally, the US tightening causes a nominal appreciation of the US dollar. This appreciation, however, carries also over in real terms.

These results describe global trends in our sample. We find, however, significant evidence for a changing international transmission of monetary policy shocks over time. More specifically, most responses are more pronounced during the mid-1990s to mid-2000s, while effects tend to be smaller in the most recent period of the sample. This holds true for effects on international inflation, short-term interest rates and exchange rates. By contrast, the international effects of US monetary policy on output are strongly shaped by the economic developments during the most recent part of our sample. Effects are most pronounced during the global financial crisis, when uncertainty was high and a boost to stimulate the economy might have been most effective. In the aftermath of the crisis, international spillover effects are comparably smaller but still more pronounced than in the early part of our sample, when financial globalization was less developed.

Third, we examine in a systematic fashion what drives the extent to which a country is more affected than another. We find that countries that are heavily integrated with the world economy via trade, economies with limited fiscal space, a pronounced current account deficit, a small degree of exchange rate volatility are most strongly affected by changes in US monetary policy. Also spillovers to countries with a high degree of FX reserves (which are often denominated in US dollars) or foreign exchange component in the external balance tend to be stronger. Either reducing these vulnerabilities or having capital account restrictions in place can mitigate the impact of external shocks on the domestic economy.

Last, our framework allows also to investigate whether due to an increase in (financial) globalization US monetary policy responds to foreign regional macroeconomic shocks. Depending on the nature of the foreign shock we find significant responses of US short-term rates. More specifically, if foreign policy rates are raised or foreign output growth decelerates, US rates decrease in response. This boosts economic growth in the US and compensates for the shortfall in foreign demand. The effects of these foreign shocks on US interest rates have increased over time. We do not find such compelling evidence in response to a foreign inflation shock or an unexpected weakening of foreign currencies against the US dollar. An exception to this are shocks that originate from the Asian region which includes China. Here, almost all shocks trigger a significant response on US short-term interest rates emphasizing the important role this region plays for the US economy.

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Appendix A Additional technical information

A.1 Convergence properties of the MCMC algorithm

As noted in Section 3, our MCMC algorithm is repeated 30,000 times, with the first 15,000 draws being discarded as burn-in. Inspection of a range of diagnostic checks indicate that the Markov chain converged to its stationary distribution. We consider the procedure proposed in Geweke (1992) to assess whether two non-overlapping parts of the Markov chain (in our case the first 10% and the final 50%) come from the same statistical distribution. For the vast majority of parameters, the null hypothesis is confirmed at the 5% significance level.

In addition, inefficiency factors tend to be remarkably low, ranging from five to 20 for most parameters of the model. Values of inefficiency factors below 30 are typically considered to be satisfactory. These favorable convergence properties of our algorithm are not surprising given the fact that our model is a relatively simple TVP-SV-VAR with additional exogenous variables.

The trace plots of selected coefficients in some countries confirm the findings described above. We typically see well behaved distributions with low autocorrelation, again emphasizing the good properties of our sampler.

A.2 Proof of Equation 2.19

In what follows we prove the result for v_{ij}^2 . The proof for s_{ij}^2 is very similar. Before proceeding to the proof, it is worth noting that the density of a generalized inverse Gaussian is proportional to $x^{\alpha-1} \exp\{-\frac{1}{2}(\mathfrak{c}/x + \mathfrak{s}x)\}$.

Proof. Note that conditional on $a_{ij}^T = (a_{ij,1}, \dots, a_{ij,T})'$, the conditional posterior of v_{ij}^2 is independent of the data. The conditional posterior is proportional to the likelihood times the Gamma prior,

$$\begin{aligned} p(v_{ij}^2 | a_{ij}^T) &\propto (v_{ij}^2)^{-\frac{T}{2}} \times \exp\left\{-\frac{1}{2} \left(\frac{\sum_{t=1}^T (a_{ij,t} - a_{ij,t-1})^2}{v_{ij}^2} \right)\right\} \times (v_{ij}^2)^{\frac{1}{2}-1} \times \exp\left(-\frac{1}{2B_v} v_{ij}^2\right) \\ &\propto (v_{ij}^2)^{(\frac{1}{2}-\frac{T}{2})-1} \exp\left\{-\frac{1}{2} \left(\frac{\sum_{t=1}^T (a_{ij,t} - a_{ij,t-1})^2}{v_{ij}^2} + \frac{1}{2B_v} v_{ij}^2 \right)\right\}, \end{aligned}$$

which is the kernel of a generalized inverse Gaussian distribution with $\alpha = 1/2 - T/2$, $\mathfrak{c} = \sum_{t=1}^T (a_{ij,t} - a_{ij,t-1})^2$ and $\mathfrak{s} = 1/(2B_v)$. ■

A.3 Sampling from the posterior of the log volatilities

This appendix provides a brief overview of the MCMC algorithm put forward in Kastner and Frühwirth-Schnatter (2013), which is used as one of the required steps to sample from the posterior distribution of the parameters of our TVP-SV-GVAR model. We start by rewriting equation (2.5) as

$$A_{i0,t}^{-1} y_{it} - (I_{k_i} \otimes x'_{it}) \text{vec}(\Psi_{it}) = \tilde{y}_{it} = D_{it}^{\frac{1}{2}} u_{i,t}. \quad (\text{A.1})$$

Here $u_{i,t} \sim \mathcal{N}(0, I_{k_i})$ and $D_{it} = (D_{it}^{\frac{1}{2}})' D_{it}^{\frac{1}{2}}$. [Kastner and Frühwirth-Schnatter \(2013\)](#) consider $\lambda_{ij,t}$ in its centered parametrization given in [equation \(2.8\)](#) and in its non-centered form given by

$$\ln(\tilde{\lambda}_{ij,t}) = \rho_{ij} \ln(\tilde{\lambda}_{ij,t-1}) + \nu_{ij,t} \text{ for } j = 1, \dots, k_i, \quad (\text{A.2})$$

where $\nu_{ij,t}$ is a standard normal error term.

Let us consider the j th equation of [equation \(A.1\)](#). Squaring and taking logs yields

$$\tilde{y}_{ij,t}^2 = \ln(\lambda_{ij,t}) + \ln(u_{ij,t}^2) \text{ for } j = 1, \dots, k_i. \quad (\text{A.3})$$

Since $\ln(u_{ij,t}^2) \sim \log \chi^2(1)$, we follow [Omori et al. \(2007\)](#) and use a mixture of normal distributions to design the sampling procedure. This renders [equation \(A.3\)](#) conditionally Gaussian, i.e., $\ln(u_{ij,t}^2 | r_{j,t}) \sim \mathcal{N}(m_{r_{ij,t}}, s_{r_{ij,t}}^2)$. The indicators controlling the mixture components prevailing at time t are labeled as $r_{ij,t} \in \{1, \dots, 10\}$. $m_{r_{ij,t}}$ and $s_{r_{ij,t}}^2$ denote the mean and variance of the corresponding mixture normal component, respectively.

Conditional on $r_{ij,t}$, we can rewrite [equation \(A.3\)](#) as a (conditionally) Gaussian linear state space model,

$$\tilde{y}_{ij,t}^2 = m_{r_{ij,t}} + \lambda_{ij,t} + \zeta_{ij,t}, \quad (\text{A.4})$$

where $\zeta_{ij,t} \sim \mathcal{N}(0, s_{r_{ij,t}}^2)$.

We simulate the history of log volatilities and the parameters of the state equation according to the following algorithm outlined in [Kastner and Frühwirth-Schnatter \(2013\)](#). The algorithm proceeds as follows:

1. *Sample* $\ln(\lambda_{ij,-1}) | r_{ij}, \mu_{ij}, \rho_{ij}, \varsigma_{ij}, \Psi_{it}, A_{i0,t}$ *or* $\ln(\tilde{\lambda}_{ij,-1}) | r_{ij}, \rho_{ij}, \zeta_{ij}, \Psi_{it}, A_{i0,t}$ *all without a loop (AWOL)*. In the spirit of [Rue \(2001\)](#), it is possible to state $\ln(\lambda_{ij,-1}) = (\ln(\lambda_{ij,2}), \dots, \ln(\lambda_{ij,T}))'$ in terms of a multivariate normal distribution

$$\ln(\lambda_{ij,-1}) \sim \mathcal{N}(\Omega_{\lambda_{ij}}^{-1} c_i, \Omega_{\lambda_{ij}}^{-1}). \quad (\text{A.5})$$

In a similar fashion, the distribution of the full state vector $\tilde{\lambda}_{ij,-1} = (\tilde{\lambda}_{ij,2}, \dots, \tilde{\lambda}_{ij,T})$ is given by

$$\ln(\tilde{\lambda}_{ij,-1}) \sim \mathcal{N}(\tilde{\Omega}_{\lambda_{ij}}^{-1} \tilde{c}_i, \tilde{\Omega}_{\lambda_{ij}}^{-1}). \quad (\text{A.6})$$

In this expression, the posterior moments are given by

$$\Omega_{\lambda_{ij}} = \begin{pmatrix} \frac{1}{s_{r_{ij,2}}^2} + \frac{1}{\varsigma_{ij}^2} & \frac{-\rho_{ij}}{\varsigma_{ij}^2} & 0 & \dots & 0 \\ -\frac{\rho_{ij}}{\varsigma_{ij}^2} & \frac{1}{s_{r_{ij,3}}^2} + \frac{1+\rho_{ij}}{\varsigma_{ij}^2} & \frac{-\rho_{ij}}{\varsigma_{ij}^2} & \ddots & \vdots \\ 0 & \frac{-\rho_{ij}}{\varsigma_{ij}^2} & \ddots & \ddots & 0 \\ \vdots & \ddots & \ddots & \frac{1}{s_{r_{ij,T-1}}^2} + \frac{1+\rho_{ij}}{\varsigma_{ij}^2} & \frac{-\rho_{ij}}{\varsigma_{ij}^2} \\ 0 & \dots & 0 & \frac{-\rho_{ij}}{\varsigma_{ij}^2} & \frac{1}{s_{r_{ij,T}}^2} + \frac{1}{\varsigma_{ij}^2} \end{pmatrix} \quad (\text{A.7})$$

and

$$c_{ij} = \begin{pmatrix} \frac{1}{s_{r_{ij},2}^2} (\tilde{y}_{ij,2}^2 - m_{r_{ij},2}) + \frac{\mu_{ij}(1-\rho_{ij})}{\zeta_{ij}^2} \\ \vdots \\ \frac{1}{s_{r_{ij},T}^2} (\tilde{y}_{ij,T}^2 - m_{r_{ij},T}) + \frac{\mu_{ij}(1-\rho_{ij})}{\zeta_{ij}^2} \end{pmatrix}. \quad (\text{A.8})$$

The moments for the non-centered case are given by $\tilde{\Omega}_i = \zeta_{ij}^2 \Omega_{h_{ij}}$ and $\tilde{c}_{ij} = \zeta_{ij}^2 c_{ij}$. The initial states of $\ln(\lambda_{ij,1})$ and $\ln(\tilde{\lambda}_{ij,1})$ are obtained from the respective stationary distributions.

2. *Sample the parameters of the state equations for both parameterizations.* Due to the lack of conjugacy of the prior setup outlined in the main body, we combine Gibbs steps with Metropolis Hastings (MH) steps. We employ simple MH steps for the parameters of the state equations in (2.8) and (A.3). In the centered parametrization case, we sample μ_{ij} and ρ_{ij} jointly using a Gibbs step and ζ_{ij}^2 is updated through a simple MH step. For the non-centered parametrization, ρ_{ij} is sampled by means of a MH step and the remaining parameters are obtained by Gibbs sampling.
3. *Sample the mixture indicators through inverse transform sampling.* Finally, the indicators controlling the mixture distributions employed are obtained by inverse transform sampling in both cases. This step can be implemented by noting that $\tilde{y}_{ij,t}^2 - \ln(\lambda_{ij,t}) = \tilde{u}_{ij,t}$ with $\tilde{u}_{ij,t} \sim \mathcal{N}(m_{r_{ij,t}}, s_{r_{ij,t}}^2)$. Posterior probabilities for each $r_{ij,t}$ are then given by

$$p(r_{ij,t} = c | \bullet) \propto p(r_{ij,t} = c) \frac{1}{s_{ij,k}} \exp\left(-\frac{(\tilde{u}_{ij,t} - m_{ij,k})}{2s_{ij,t}^2}\right), \quad (\text{A.9})$$

where $p(r_{ij,t} = c)$ is the weight associated with the c th component.

In the implementation of the present algorithm we simply draw the parameters under both parametrization and, depending on the relationship between the innovation variances of equation (A.1) and equation (2.8), we decide ex-post whether we should discard draws obtained from the centered parametrization or keep them. This constitutes the interweaving part of the algorithm. For further details we refer the reader to [Kastner and Frühwirth-Schnatter \(2013\)](#).¹⁰

A.4 Computational aspects of posterior inference in the TVP-SV-GVAR model

Since our sampling scheme treats countries and equations as isolated estimation problems, parallel computing can be exploited to carry out posterior inference in the TVP-GVAR model. Such a modeling strategy proves to be an efficient means of estimating high-dimensional GVARs with drifting parameters, while imposing parametric restrictions only on the international linkages that take place through the weakly exogenous variables.

¹⁰The steps described here are implemented using the [stochvol](#) package in R, a language and environment for statistical computing ([R Development Core Team, 2011](#)).

The combination of the Cholesky structure in [equation \(2.2\)](#) and the presence of the weakly exogenous variables permit equation-by-equation and country-by-country estimation. This constitutes an estimation strategy that relies heavily on parallel computation to obtain parameter estimates for [equation \(2.11\)](#). The first strategy views the GVAR model as a system of k unrelated regression models, which can be spread across ϱ processors. In this case, the maximum speedup gained by parallelization is given by

$$\text{Maximum Speedup} = \frac{1}{\frac{f}{\varrho} + (1 - f)}. \quad (\text{A.10})$$

Here, f denotes the fraction of the problem which can be parallelized. Equation [\(A.10\)](#) is known as Amdahl's law ([Rodgers, 1985](#)) in computer science. If f equals unity the task at hand is called *embarrassingly parallel*, making it perfectly suitable for parallel computing. In the GVAR setting, f is close to unity after taking into account the costs of distributing the information across the different processing units. In addition, it is worth emphasizing that since we impose a triangular structure on the model and the number of endogenous variables per country model differs (note that in general, $k_i \neq k_j \forall j, i$), the number of parameters differs from equation to equation. The maximum computation time is bounded by the time required to estimate the equation with the maximum number of parameters. If the number of CPU cores ϱ equals k , computation time almost boils down to that required for estimating the equation with the maximum number of parameters.

Appendix B Data appendix

Table B.1: Potential determinants of spillover size

Variable	Description	Source
Exchange rate stability		
Exchange rate	Yearly average of the month-end exchange rate against the US dollar.	Shambaugh exchange rate regime classification (see "KS").
ER Volatility	Annual standard deviation of the monthly percentage change in the exchange rate against the base country. In our data set and according to the "base" definition of the Shambaugh data set, in 53% of years and countries the base country is the USA, followed by Germany / euro area (39%), Australia (3%), Malaysia (3%) and oil exporting countries (2%).	Shambaugh exchange rate regime classification (see "KS").
FX Exposure	Aggregate measure of foreign-currency exposure which is relevant in capturing the sensitivity of a country's external balance sheet to a uniform movement of its domestic currency against all foreign currencies	See Bénétrix et al. (2015) .
FX Reserves	Foreign exchange reserves minus gold in % of GDP.	IMF's IFS database.
Min Deviation	The largest downwards (i.e., appreciation) move in the exchange rate in a given month against the base currency.	Shambaugh exchange rate regime classification (see "KS").
Max Deviation	The largest upwards (i.e., depreciation) move in the exchange rate in a given month against the base currency.	Shambaugh exchange rate regime classification (see "KS").
Asset Exposure	The quantitative exposure of assets and liabilities to a uniform shift in the value of the domestic currency against all foreign currencies	See Bénétrix et al. (2015) .
Zero Change	Number of months with no change in the exchange rate against the base country.	Shambaugh exchange rate regime classification (see "KS").
Base Exchange Rate	Yearly average of the month-end exchange rate against base currency.	Shambaugh exchange rate regime classification (see "KS").
Range	Range of domestic currency to base currency movements (in %).	Shambaugh exchange rate regime classification (see "KS").
Macroeconomic vulnerabilities		
Current Account	Current account in % of GDP.	IMF's IFS or WEO database.
Fiscal Deficit	Fiscal deficit in % of GDP.	IMF's WEO database.
Government Debt	Government debt, in % of GDP.	IMF's WEO database.
Gross Savings	Gross savings in % of GDP.	IMF's WEO database.
Financial depth and stability		
Bank Credit to Deposits	Bank credit to bank deposits in % (financial stability).	World Banks' global financial development database, see also "BDL".
Deposit Money	Deposit money bank assets to GDP in % (financial depth).	World Banks' global financial development database, see also "BDL".
Financial Deposits	Financial systems deposits to GDP in % (financial depth).	World Banks' global financial development database, see also "BDL".
Liquid Liabilities	Ratio of liquid liabilities (broad money) to GDP (financial stability).	World Banks' global financial development database, see also "BDL".
Private Credit	Private credit by deposit money banks and other financial institutions to GDP in % (financial depth).	World Banks' global financial development database, see also "BDL".
FDI Assets	FDI asset side, in % of GDP.	External Wealth Mark II database, see also "LM".
Financial and trade openness		
Portfolio Assets	Portfolio equity, asset side (in % of GDP).	External Wealth Mark II database, see also "LM".
Portfolio Liability	Portfolio equity, liability side (in % of GDP).	External Wealth Mark II database, see also "LM".
Foreign Liabilities	Total foreign liabilities in % of GDP.	External Wealth Mark II database, see also "LM".
Foreign Assets	Total foreign assets in % of GDP.	External Wealth Mark II database, see also "LM".
Trade Openness	Exports + Imports in % of GDP.	World Banks' WDI database.
Capital Restrictions (inflow)	Overall capital account inflow restrictions index.	See Fernández et al. (2015) .
Capital Restrictions (outflow)	Overall capital account outflow restrictions index.	See Fernández et al. (2015) .

Notes: "KS" denotes [Klein and Shambaugh \(2008\)](#), "BDL" [Beck et al. \(2009\)](#) and "LM" [Lane and Milesi-Ferretti \(2007\)](#).