Department of Economics
Working Paper No. 209

US Monetary Policy in a Globalized World

Jesus Crespo Cuaresma
Gernot Doppelhofer
Martin Feldkircher
Florian Huber

November 2015

WIRTSCHAFTS UNIVERSITÄT WIEN VIENNA UNIVERSITY OF ECONOMICS AND BUSINESS

EFMD EQUIS ACCREDITED

US Monetary Policy in a Globalized World*

Jesús Crespo Cuaresma[1,4,5,6], Gernot Doppelhofer[2], Martin Feldkircher[3], and Florian Huber[1,3]

[1]Vienna University of Economics and Business (WU)
[2]Norwegian School of Economics (NHH)
[3]Oesterreichische Nationalbank (OeNB)
[4] International Institute of Applied System Analysis (IIASA)
[5] Wittgenstein Center for Demography and Global Human Capital (IIASA,VID/OEAW,WU)
[6] Austrian Institute of Economic Research (WIFO)

Abstract

We analyze the interaction between monetary policy in the US and the global economy proposing a new class of Bayesian global vector autoregressive models that accounts for time-varying parameters and stochastic volatility (TVP-SV-GVAR). Our results suggest that US monetary policy responds to shocks to the global economy, in particular to global aggregate demand and monetary policy shocks. On the other hand, US-based contractionary monetary policy shocks lead to persistent international output contractions and a drop in global inflation rates, coupled with rising interest rates in advanced economies and a real depreciation of currencies with respect to the US dollar. We find considerable evidence for heterogeneity in the spillovers across countries, as well for changes in the transmission of monetary policy shocks over time.

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

JEL Codes: C30, E52, F41

*Corresponding author: Florian Huber, Oesterreichische Nationalbank (OeNB), Phone: +43-1-404 20-5208. E-mail: florian.huber@oenb.at. The opinions expressed in this paper are those of the authors and do not necessarily reflect the official viewpoint of the Oesterreichische Nationalbank or the Eurosystem. The authors thank Sylvia Frühwirth-Schnatter, Mike West, Gregor Kastner and Philipp Piribauer as well as participants at research seminars at the Norwegian School of Economics, the OeNB and the NBER-NSF 2015 time series workshop for helpful comments on an earlier draft of this paper The first two authors acknowledge support by funds of the Jubiläumsfonds of the Oesterreichische Nationalbank (project number 16244).

"... *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 (Kamin, 2010). 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 behaviour 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 policy directly in terms of the two major operational objectives of monetary policy, domestic inflation and economic growth. Indeed and 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 Taylor rule the US Fed sets monetary policy in response to developments in 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 its monetary 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.

In this paper we analyze both the effect of shocks to US monetary policy on global output and the feedback of fluctuations in the world economy on US monetary policy. Generally the empirical literature hitherto finds significant effects of US monetary policy on global output. Most studies assessing the effects of macroeconomic shocks in the US economy on the world either use 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 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, 2015). Employing this framework, Feldkircher and Huber (2015) 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.

The implications of increased globalization on the policy behaviour 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 (demand) shocks.

To address both of these questions, accounting for changes in the economic environment and the Fed reaction function seems essential. 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). 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).

In this contribution, we assess the dynamic relationship between US monetary policy and the world economy over time making use of a new class of global macroeconomic models. 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 model (TVP-SV-GVAR) is then 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 spillovers from US monetary policy to global output and analyze how the Fed reacts to external shocks. We address the changes in spillovers over time to judge whether the transmission from and to the US has significantly changed in the last decades.

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 global inflation rates together with (c) a rise in global interest rates in advanced economies, and (d) a relative real depreciation with respect to the US dollar. The estimated effects are in line with the empirical literature on the effects of shocks to monetary policy originated in the US on other economies (see Christiano et al., 1999). Second, we find evidence for important heterogeneity of the spillovers across economies, as well as for a changing transmission of monetary policy shocks at the global level over time. The global response to US monetary policy shocks became stronger over the last decade indicating an increase in financial globalization. We also find that it is important to allow for stochastic volatility in macroeconomic variables during crises and turbulent episodes. While many developed economies exhibit declining volatility until the mid-2000's (the so-called Great Moderation), we can observe a resurgence in volatility with the start of the Great Recession in 2007. Many emerging economies in Latin America and Asia experienced sharp changes in volatility due to economic crises. Last, monetary policy in the US responds to global macroeconomic shocks. In particular, US short-term interest rates respond to global demand (in the medium-term) and monetary policy shocks (in the short-term), while there is less evidence in the data for a reaction to global supply shocks. In general, the intensity of which US interest rates responded to global shocks has abated over time.

The paper is structured as follows. Section two presents the econometric framework including the necessary prior specifications and the Bayesian estimation strategy. Section three presents the results of the empirical study and section four 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, estimating a GVAR model consists of two distinct stages. In the first, we estimate a set of $N+1$ country-specific multivariate time series models, each of them including exogenous regressors that aim to capture cross-country linkages. In a 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. Explicitly allowing for variation of parameters over time is challenging from a computational point of view, so we rely on recent Bayesian techniques to estimate the model in an efficient manner.

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,

$$y^*_{ij,t} = \sum_{s=0}^{N} w_{ic} y_{cj,t} \text{ for } j = 1, \dots, k^*_i, \tag{2.1}$$

where w_{is} 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 or financial linkages. In line with the bulk of the literature on GVAR modelling, we assume that $\sum_{c=0}^{N} w_{ic} = 1$ and $w_{ii} = 0$.[1]

We depart from existing GVAR modelling efforts by specifying country-specific structural VAR models featuring exogenous regressors, time-varying parameters and stochastic volatility, so that

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

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 $\text{diag}(A_{i0,t}) = \iota_{k_i}$, where ι_{k_i} is a k_i-dimensional unit vector. 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}.

- $\varepsilon_{it} \sim \mathcal{N}(0, D_t)$ is a heteroskedastic vector error term with $D_t = \text{diag}(\lambda_{i0,t}, \dots, \lambda_{ik_i,t})$.[2]

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

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

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

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

[1] Hereby we assume that all variables and countries are linked together by the same set of weights.

[2] The assumption of a diagonal D_t simplifies the computational burden of model estimation enormously, since the k_i equations can be viewed as separate estimation problems and hence easily parallelised to achieve computational gains. See the Appendix for further details on the computational challenges involved in obtaining posterior distributions for model quantities.

allows us to rewrite equation (2.2) as

$$A_{i0,t} y_{it} = (I_{k_i} \otimes x'_{it}) \operatorname{vec}(\Psi_{it}) + \varepsilon_{it}. \tag{2.5}$$

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

$$a_{i0,t} = a_{i1,t-1} + \epsilon_{it}, \;\; \epsilon_{it} \sim \mathcal{N}(0, V_i) \tag{2.6}$$

where V_i is a (block-diagonal) variance-covariance matrix with $V_i = \operatorname{diag}(V_{i1}, \ldots, V_{ik_i})$. The block-diagonal nature stems from the fact that we estimate the model on a equation-by-equation basis, thus effectively disregarding the contemporaneous relationships between parameters in different equations. Likewise, we assume that the autoregressive coefficients in Ψ_{it} evolve according to

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

with $S_i = \operatorname{diag}(S_{i1}, \ldots, S_{ik_i})$ 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,

$$\log(\lambda_{il,t}) = \mu_{il} + \rho_{il}(\log(\lambda_{il,t}) - \mu_{il}) + \upsilon_{il,t}, \;\; \upsilon_{il,t} \sim \mathcal{N}(0, \varsigma^2_{il}), \tag{2.8}$$

where μ_{il} denotes the unconditional expectation of the log-volatility, ρ_{il} the corresponding persistence parameter and ς^2_{il} 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. On the other hand, the inclusion of weakly exogenous foreign variables accounts for cross-country linkages and enables us to investigate the stability properties of the model 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 $z_{it} = (y'_{it}, y^{*\prime}_{it})'$, we obtain

$$C_{it} z_{it} = \sum_{s=1}^{S} L_{is,t} z_{it-s} + \varepsilon_{it} \tag{2.9}$$

with $C_{it} = (A_{i0,t}, -\Lambda_{i0,t})$, $L_{is,t} = (B_{ip,t}, \Lambda_{iq,t})$ and $S = \max(P, Q)$. A global vector $y_t = (y'_{0t}, \ldots, y'_{Nt})'$ of dimension $k = \sum_{i=0}^{N} k_i$ and a corresponding country-specific linkage matrix W_i ($i = 1, \ldots, N$) 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,

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

Stacking the equations $N + 1$ times yields

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

where $G_t = ((C_{0s,t}W_0)', \ldots, (C_{Ns,t}W_N)')$, $F_{st} = ((L_{0s,t}W_0)', \ldots, (L_{Ns,t}W_N)')'$ and $e_t = (\varepsilon_{0t}', \ldots, \varepsilon_{Nt}')'$. The error term e_t is normally distributed with variance-covariance matrix $H_t = \text{diag}(D_{0l,t}, \ldots, D_{Nl,t})$. 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.

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). We impose a normally distributed prior on Ψ_{i0}, the initial state of Ψ_{it},

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

with $\underline{\Psi}_i$ a $k_i \times m_i$ prior mean matrix and $\underline{V}_{\Psi_i}$ 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 an inverted Wishart prior on the variance-covariance matrix S_i in equation (2.7), in line with the literature. Specifically, we assume that

$$S_{ir} \sim \mathcal{IW}(\underline{v}_i, \underline{S}_{ir}), \text{ for } r = 1, \ldots, k_i, \tag{2.13}$$

where $\underline{v}_i$ is the prior degrees of freedom and $\underline{S}_{ir}$ denotes a prior scale matrix. The normal prior on Ψ_{i0} and the set of inverted Wishart 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 zero. Second, the inverted Wishart prior can be set 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.

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

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

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 inverted Wishart priors on V_i

$$V_{ir} \sim \mathcal{IW}(\underline{m}_i, \underline{V}_{ir}), \text{ for } r = 1, \ldots, k_i, \tag{2.15}$$

where $\underline{m}_i$ denotes the prior degrees of freedom and $\underline{V}_{ir}$ is the prior scaling matrix.

Finally, we use the prior setup proposed in Kastner and Frühwirth-Schnatter (2014) and subsequently used in Huber (2014) on the coefficients of the log-volatility process in equation (2.8). A normal prior is imposed on μ_{il} $(l = 1, \ldots, k_i)$ with mean $\underline{\mu}_i$ and variance $\underline{V}_{\mu_i}$

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

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

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

which implies

$$\begin{aligned} E(\rho_{il}) &= \frac{2a_0}{a_0 + b_0} - 1 \\ \text{Var}(\rho_{il}) &= \frac{4a_0 b_0}{(a_0 + b_0)^2(a_0 + b_0 + 1)}. \end{aligned}$$

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 ς_{il}^2,

$$\varsigma_{il}^2 \sim \mathcal{G}\left(\frac{1}{2}, \frac{1}{2B_\sigma}\right). \tag{2.18}$$

This choice does not bind ς_{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, 2014). Following Lopes et al. (2013), we impose a Cholesky structure at the individual country level, which provides significant computational gains when sampling from the posterior distributions of interest.

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

$$\begin{aligned}\text{vec}(\Psi_i^T) &= (\text{vec}(\Psi_{i1})', \ldots, \text{vec}(\Psi_{iT})')', \\ a_i^T &= (a_{i1}', \ldots, a_{iT}')', \\ \lambda_i^T &= (\lambda_{i1}, \ldots, \lambda_{iT})'.\end{aligned}$$

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).
- Conditional on $\text{vec}(\Psi_i^T)$ and a_i^T, the variance-covariance matrices in equation (2.6) and equation (2.7) can be sampled from inverted Wishart distributions with precision matrices given by $\overline{V}_{ir} = \underline{V}_{ir} + \sum_{t=1}^{T}(a_{it} - a_{it-1})(a_{it} - a_{it-1})'$ for equation (2.6) and $\overline{S}_{ir} = \underline{S}_{ir} + \sum_{t=1}^{T}(\text{vec}(\Psi_{it}) - \text{vec}(\Psi_{it-1}))(\text{vec}(\Psi_{it}) - \text{vec}(\Psi_{it-1}))'$ for equation (2.7), and posterior degrees of freedom $\overline{m}_i = \underline{m}_i + T$ and $\overline{v}_i = \underline{v}_i + T$.
- The history of log volatilities is sampled using the algorithm outlined in Kastner and Frühwirth-Schnatter (2014).[3]

3 The international dimension of US monetary policy

The following section introduces the data and the priors placed on the parameters of the model framework. Using this dynamic model we then investigate how a US monetary policy shock affects the world economy. Finally, we assess how the US Fed takes external shocks into account when conducting monetary policy. For the latter purpose we aim at disentangling direct effects from an increase in globalization from indirect effects that affect US monetary policy through spillovers to the domestic economy.

3.1 Data overview, model specification and prior implementation

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.

[Table 1 about here.]

The country-specific TVP-VAR-SV models include real GDP growth (Δy), the log-difference of the consumer price level (Δp), the log-difference of the real exchange rate (Δe) vis-á-vis the US dollar, short-term interest rates (i) and the term spread, constructed as the difference between short-term and long-term interest rates (s). Note

[3]Further details of the sampling algorithm by Kastner and Frühwirth-Schnatter (2014) can be found in the Appendix.

that not all variables are available for each of the countries we consider in this study. With the exception of long-term interest rates, that are used to calculate the term-spread, cross-country coverage of all variables is, however, above 80%.

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

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

We follow the bulk of the literature by including oil prices (*poil*) 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.[4] For a *typical* country i the set of *weakly exogenous* and global control variables comprises

$$\boldsymbol{x}^*_{it} = (\Delta y^*_{it}, \Delta p^*_{it}, i^*_{it}, s^*_{it}, \Delta poil^*)'. \tag{3.2}$$

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 trade weighted real exchange rate (Δe^*) is included as an additional control variable.

$$\begin{aligned} \boldsymbol{x}_{0t} &= (\Delta y_{0t}, \Delta p_{0t}, i_{0t}, s_{0t}, \Delta poil_t)', & (3.3)\\ \boldsymbol{x}^*_{0t} &= (\Delta y^*_{0t}, \Delta p^*_{0t}, \Delta e^*_{0t}, i^*_{0t}, s^*_{0t})'. & (3.4) \end{aligned}$$

For identification of the US monetary policy shock we will rely on recursive identification using exactly the same order of the variables as they appear in equation (3.3). This ordering has been proposed – among others – in Christiano et al. (2005) and implies that output, inflation and exchange rates react sluggishly, while the term spread is allowed to react instantaneously to a monetary policy shock. For all countries considered, we set the lag length of endogenous and weakly exogenous variables equal to one. Given the relatively short period spanned by our sample, the strong parametrization of the model and the quarterly frequency of the data, this seems to be a reasonable choice.[5]

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

[4] Note that recent contributions (Eickmeier and Ng, 2011; 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 (2015) for the results of a sensitivity analysis with respect to the choice of weights in Bayesian GVAR specifications.

[5] We 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.

hyperparameters. 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 $\underline{V}_{\Psi_i} = 10 \times I_{k_i m_i}$. Similarly we set $\text{vec}(\underline{a}_j) = 0$ 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 to be non-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 both $\underline{S}_{ir}$ and $\underline{V}_{ir}$ are diagonal matrices where the elements in the main diagonal equal 0.001 and the prior degrees of freedom equal to 40, respectively.[6] This choice is highly influential in practice and we have thus performed extensive robustness checks with respect to those hyperparameters. 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 which is quite loose. Given that we are interested in allowing the data to be as informative as possible 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 we evaluate is given by $(0.00001, 0.0001, 0.001, 0.005, 0.01)$ where we pick 0.001 as our reference value for both $\underline{S}_{ir}$ and $\underline{V}_{ir}$. Higher values typically lead to posterior draws which are excessively unstable, resulting in implausible impulse-responses.

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 is usually not informative about the autoregressive parameter on a latent factor, the corresponding 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 specify B_σ, which is set equal to unity.

We compute all relevant quantities by performing Monte Carlo integration by simply drawing 1,500 samples from a total chain of 30,000 draws, where the first 25,000 draws have been discarded. Standard diagnostic checks indicate convergence towards the stationary distribution, with inefficiency factors for the autoregressive coefficients and volatilities all well below 20 for all country models.

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

In this section we investigate the international responses of an unexpected monetary policy tightening in the US. The shock is normalized to a 50 basis points increase on impact throughout the sample period. The results are summarized in Figures 1

[6] This value implies a size of 40 observations for a hypothetical pre-sample period and, given the number of parameters to be estimated, guarantees a proper prior.

to 8. The plots show the posterior mean of the corresponding (cumulated) impulse response for selected countries, along with the cross-country means (in red) and the associated 25% and 75% credible sets (gray shaded regions) that 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 for three different horizons: after one quarter ($t = 1$), after eight quarters ($t = 8$) and after 12 quarters ($t = 12$).

Figures 1 and 2 show the output response for selected developed economies and Western European economies. The reaction of output in developed economies is very homogeneous, both across countries and over time. In all three different time horizons considered, the output reaction is negative and for most countries lies within the credible set. Taken at face value, this indicates a very persistent effect of spillovers generated from a US monetary policy shock on output and corroborates the findings by Feldkircher and Huber (2015), who use a linear, time-invariant version of the Bayesian GVAR model. While responses are in general very homogeneous, two countries stand out: on the one hand, Canada's response is even stronger than the domestic reaction of output in the US itself. On the other hand, the output response in Australia indicates that the country is relatively isolated in terms of spillovers from the US monetary policy shock. The response in selected Western European economies is also very homogeneous and – compared to the group of developed economies – smaller in magnitude. Spillovers from the monetary policy shock are very persistent from the mid-1990s onward.

There is much more time variation in responses when turning to emerging economies. In Asia, with the exception of China, all countries respond negatively on impact to the monetary tightening in the US. These effects, however, peter out in the long-run. Note also that credible sets for this group of countries are much wider than for developed and Western European economies. Some economies show a clear downward trend in responses, while for others spillover effects became less pronounced in the most recent period of the sample. While the response in China differs from its regional peers at the beginning of our sample, in the most recent part the Chinese response becomes more similar to that of India and Indonesia. Such detail about the changing characteristics of the spillovers of monetary policy shocks is only possible due to the time-varying nature of the specification of the parameter vector in our model and would be lost in a linear setting with fixed parameters. In Latin America, output responses are also very persistent, which is in line with findings of Feldkircher and Huber (2015).

[Figure 1 about here.]

[Figure 2 about here.]

Responses of inflation are depicted in Figures 3 and 4. On impact, the domestic response of inflation in the US is negative. Our model is thus rich enough to yield responses that do not create a price puzzle, which lends further confidence to the overall results. In the medium term, however, inflation in the US adjusts and becomes positive. However, note that even in this case and under the assumption that the sampling variability is similar across countries this implies that the responses are turning insignificant after the first few quarters. Again, Canada shows the most pronounced response to the monetary tightening in the US – even after three years, the effect is pronounced and negative. In general, responses of inflation show more time-variation compared to output responses. However, credible sets are also much wider, including zero responses throughout the sample period and for all three impulse response horizons. A similar picture arises in Western European economies: inflation responds negatively throughout the sample period and across countries. However, credible sets are large and responses in the medium-term hover around zero for most of the sample period. With respect to time variation, for most of the economies a downward shift in response has set in around 2005. In line with findings for output responses, reactions of inflation show much more variation over time for emerging as compared to advanced economies. In emerging Asia, the monetary tightening triggers negative reactions of inflation on impact, while responses in the medium term are accompanied by wide credible sets including the zero response. On impact, Thailand and Indonesia are the economies that appear to be most isolated from the shock. In the medium-term, the response in China is the strongest, with effects comparable to the lower credible bound throughout the sample period. Impact responses in Latin America tend to be more diverse. On the one hand, responses on impact are positive in Peru, Chile and Mexico, with the latter exceeding the upper threshold of the credible set throughout most of the sample period. On the other hand, responses are negative in Brazil and Argentina and the lower credible set throughout the sample period. In the medium-term, the response estimates indicate considerable time variation in the reaction of inflation, particularly at the beginning of the sample period, in which some countries experienced times of hyperinflation.

[Figure 3 about here.]

[Figure 4 about here.]

Figures 5 and 6 show the response of interest rates with respect to 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 (2015). While the response of US short-term rates has been fixed to 50 basis points on impact, spillovers for most other developed economies are weak. An exception is given by Canada, whose short-term rates rise strongly in response to the US monetary policy tightening. In the medium-term, interest rates in Canada and the US still increase, while responses of interest rates for the remaining countries are close to zero. Since responses are so diverse, the regional mean might not be informative in case of short-term interest rates. With the exception of Spain, impact responses in Western Europe are positive and credible sets are well above zero. For the case of Spain, responses are more similar to the rest of the group in the second part of the sample. In the medium-term, credible sets are wide throughout the region. Interest rate responses vary considerably over the period covered for most of the countries, emphasizing the importance to consider a time-varying parameter framework when assessing global linkages in the response to macroeconomic shocks. While in most advanced economies interest rates move in parallel with US short-term rates, interest rates in emerging economies tend to decrease in response to the contractionary US monetary policy shock. With the exception of Indonesia, all economies in emerging Asia respond negatively on impact and accompanying credible sets are relatively tight. After eight quarters, responses are still negative but barely significant. It should be noted that the variation of responses over time is pronounced. For example, responses in Korea are positive in the first part of the sample, after which they turn persistently negative. A similar picture arises for responses of short-term rates in Latin America. Here, on impact short-term rates decrease for all countries but Mexico and Chile in the most recent period of the sample. In the medium-term, interest rates still respond negatively, underlining the importance of the financial channel in transmitting external shocks from the US (Canova, 2005). Mexico and Chile deviate also in the medium term from its peers in Latin America. For both countries responses are strongly time varying and positive throughout the sample period. Other countries, like Argentina, also show pronounced variations of responses over time.

[Figure 5 about here.]

[Figure 6 about here.]

Figures 7 and 8 show the responses of the real exchange rate vis-á-vis the US dollar. As expected, responses are positive on impact indicating a real appreciation of the US dollar as a consequence of the interest rate increase. Also consistent with our findings so far, Canada is the most affected economy while Australia's currency seems resilient to the US interest rate shock. This patterns also holds in the medium-term. Strikingly, the effect of the monetary policy shock on real exchange rates appears increasing over time for all countries but Australia. This implies that interest rate increases now tend to have stronger appreciating effects on the dollar than at the beginning of our sample period. Qualitatively a similar picture arises when we consider developed economies in Western Europe. On impact, all currencies weaken against the US dollar. These effects are most pronounced for the United Kingdom, which historically shares strong trade and financial linkages with the US. In the medium run, credible sets widen. However, the effects of monetary policy on exchange rates also become more pronounced over the period considered and are especially well estimated from the mid 2000s onward. In emerging Asia, currencies weaken against the US dollar on impact, after eight quarters and after 12 quarters. Over all three forecast horizons, effects are most pronounced for Korea, while the other countries show a very homogeneous response. In line with results for advanced economies, responses increase gradually with the sample period, indicating a stronger sensitivity of the currencies in the most recent period of the sample. Also Latin American currencies weaken against the US dollar on impact and credible sets are tight. In the medium term credible sets widen and include the zero response for all currencies considered. Similar to the results for the other regions, currencies tend to react more strongly in response to a US monetary policy shock in the most recent period of the sample. However, and in contrast to responses of currencies in emerging Europe, in Latin America reactions seem to be less gradual. Around the year 2000, responses of the real exchange rate started to rise significantly, while in the most recent period responses even declined.

[Figure 7 about here.]

[Figure 8 about here.]

The time-varying nature of the volatility of macroeconomic variables is integrated in the model specification and thus taken into account when deriving the posterior quantities presented above. Our modelling framework provides also explicit inference on the dynamics of macroeconomic volatility. As an example,[7] Figure 9 plots the volatility of GDP growth for four regional groups and a selection of countries within each group: (a) Western Europe (France, Germany, Great Britain, Norway and Spain), (b) Asia (China, India, Indonesia, Korea and Thailand), (c) Latin America (Argentina, Brazil, Chile, Mexico, and Peru), and (d) other developed economies (Australia, Canada, Japan, New Zealand and the US). Figure 9 shows the cross-country median of the standardized volatility estimate per group (in red), which can be interpreted as the corresponding volatility of a typical country in the region under consideration. A decline in the volatility of GDP growth in Western Europe and other developed economies can be observed until the middle of the 2000's, a development which is in line with the dampening of real fluctuations corresponding to the Great Moderation period. After 2007, we see a sharp increase in output growth volatility due to the outbreak of the global financial crisis, followed by a gradual return to lower volatility more recently. 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.

[Figure 9 about here.]

Our set of results concerning the changing international dimension of US monetary policy emphasize the importance of considering variation of responses and uncertainty over time. Relying on specifications with fixed parameters may lead to biased and unreliable inference on the spillovers that US based monetary policy shocks have on the global economy.

[7] Due to space constraints, we present here only the results for GDP growth volatility dynamics. Results for the rest of the macroeconomic variables in our TVP-SV-GVAR model are available from the authors upon request.

3.3 Does the US Fed respond to global shocks?

Typically, the reaction function of the US Fed is modeled as a linear function of purely domestic quantities (Christiano et al., 1999). For instance, reaction functions based on Taylor rules (Taylor, 1993) assume that the Fed sets the policy rate according to a simple linear function of inflation expectations, the output gap and possibly the effective exchange rate (Taylor, 2002; Clarida et al., 1998). By establishing such rule-based behavior, it is theoretically ruled out that the Fed reacts to international economic developments beyond the spillovers which are directly reflected in domestic US macroeconomic variables. In addition, the assumption of a linear monetary policy reaction function implies that the central bank conducts monetary policy with elasticities which are independent from the prevailing state of the economy, i.e., the reaction function is the same in boom and bust phases.

To provide further insights on the theoretical relationship between the preferences of the central banks and global macroeconomic developments, we use a simplified variant of the model outlined in Rudebusch and Svensson (1999) and Bernanke et al. (2005). We assume that the economy is described by the following set of equations

$$\Delta p_t = \delta \Delta p_{t-1} + \kappa(y_{t-1} - \overline{y}_{t-1}) + \mathfrak{s}_t, \tag{3.5}$$

$$y_t = \phi y_{t-1} - \psi(i_{t-1} - \Delta p_{t-1}) + \theta \chi_t + d_t, \tag{3.6}$$

$$\chi_t = \omega y_t^* - \alpha y_t, \tag{3.7}$$

where Δp_t and y_t denote the rate of inflation and current output, $\overline{y}_t$ denotes potential output and $\mathfrak{s}_t$ is a serially correlated cost-push shock. Furthermore, i_t denotes the policy rate controlled by the monetary policy authority and y_t^* denotes foreign output. Let us assume that the central bank sets the policy rate according to

$$i_t = \beta \Delta p_t + \gamma(y_t - \overline{y}_t) + \epsilon_t. \tag{3.8}$$

Here, ϵ_t denotes a zero mean monetary policy shock with constant variance. Plugging equation (3.7) into equation (3.6) and solving for y_t yields

$$y_t = \frac{1}{1+\alpha}\left[\phi y_{t-1} - \psi(i_{t-1} - \Delta p_{t-1}) + \theta\omega y_t^* + d_t\right]. \tag{3.9}$$

By substituting equation (3.9) in equation (3.8), we relate foreign output with the policy rate through an augmented Taylor rule equation,

$$i_t = \beta\pi_t + \tilde{\psi} y_{t-1} - \tilde{\psi}(i_{t-1} - \Delta p_{t-1}) + \tilde{\omega} y_t^* + \tilde{\gamma} d_t - \gamma \overline{y}_t + \epsilon_t, \tag{3.10}$$

with $\tilde{\psi} = \frac{\gamma\psi}{1+\alpha}, \tilde{\psi} = \frac{\gamma\psi}{1+\alpha}, \tilde{\omega} = \frac{\gamma\theta\omega}{1+\alpha}$ and $\tilde{\gamma} = \frac{\gamma}{1+\alpha}$. Assuming $\alpha = 0$ and $\omega = 0$ leads to the model of Bernanke et al. (2005). Note that equation (3.10) implies that $\frac{\partial i_t}{\partial y_t^*} = \tilde{\omega}$, which is greater than zero if $\gamma, \theta, \omega, \alpha > 0$, a set of assumptions which is routinely assumed to hold in this type of model. If $\tilde{\omega} > 0$, the central bank increases its policy rate as a response to a positive international output shock. This shows that even if the central bank does not explicitly consider international output in its reaction function, there are indirect channels that lead to global developments playing a role in domestic monetary policy. It is straightforward to show that the model given by equation (3.5) - equation (3.10) is a restricted variant of equation (2.2), where only weakly exogenous output is included and the parameters are assumed to be constant over time.[8]

In practice, the structural parameters embodied in the coefficients of equation (3.10) can be thought of as changing over time and we can relax the assumption that international output is the only variable affecting domestic monetary policy by assuming that the policy instrument is set according to

$$i_t = \mathfrak{f}_t(\Omega_t) + \epsilon_t, \tag{3.11}$$

where $\mathfrak{f}_t(\Omega_t)$ is a non-linear function of the information set of the central bank up to time t (Ω_t). Relating equation (3.11) to the GVAR model outlined in Section 2 implies that Ω_t now may include information on international output, interest rates, prices, exchange rates and term spreads. This allows us to investigate the behavior of US monetary policy when facing shocks to the aforementioned international quantities.

To assess the international dimension of US monetary policy we perform a set of simple counterfactual exercises by estimating the response of US interest rates to three distinct structural shocks:

1. a one standard deviation global aggregate demand shock,
2. a one standard deviation global supply shock and
3. a one standard deviation monetary policy shock.

We identify the structural shocks by imposing zero-impact restrictions in the spirit of the identifying assumptions employed for the US-based structural shocks. We assume that output, inflation and exchange rates react sluggishly, while the term spread is allowed to react instantaneously to a monetary policy shock. Since the number of restrictions that have to be imposed is large, the simple recursive scheme imposed in the TVP-SV-GVAR proves to be a convenient way to retrieve the structural form of the model.[9]

[Table 2 about here.]

[8]It is straightforward to extend the theoretical framework and incorporate further international macroeconomic variables such as international price movements or changes in foreign interest rates. For the sake of brevity and since the model presented here is purely exemplary, we exclusively consider output as a foreign variable.

[9]We performed several robustness checks concerning the shock identification scheme. For instance, assuming different orderings of the variables in the vector of endogenous covariates leads to results that are qualitatively similar. In addition, we also computed generalized impulse response functions (Pesaran and Shin, 1998). The responses obtained are comparable to the ones obtained from structural identification schemes.

Figure 10 depicts the posterior median of the response of the US short-term interest rate with respect to the different shocks described above. In addition, Table 2 presents the posterior estimates of the responses of the US short-term interest rates, averaged across different periods corresponding to distinct US monetary policy regimes. Several findings are worth emphasizing.

First, Figure 10(a) displays US short-term interest rate responses to a global aggregate demand shock. From the second row of 2 the following picture arises: In the first regime, labeled the *Volcker regime*, following the global demand shock short-term interest rates increased by around 4.6 basis points after one year. The interest rate responses become weaker in subsequent periods, increasing by around 4.3 basis points in the *Greenspan regime*, while the median response reached 3.9 basis points during Bernanke's chairmanship. Interestingly, medium-run responses in the final part of our sample turn insignificant, suggesting that under Ben Bernanke's leadership the behavior of the Fed might be less driven by global demand shocks. Since the final period covers the recent financial crisis, which originated in the US and subsequently engulfed the rest of the world, such a result appears intuitive. The most recent period summarized in Table 2 is extraordinary due to the depth of the recession that it covers, with the Fed adopting unconventional monetary policy measures. These measures have been designed to foster economic growth exclusively in the US, without actively reacting to international economic developments. By contrast, the responses prior to 2006 provide some evidence that the Fed acted as if it closely monitored international output, reacting significantly to such developments within four quarters.

Similarly to the global demand shock, the responses to a global aggregate supply shock shown in Figure 10(b) tend to display only a minor degree of time variation. In addition, the rows related to the aggregate supply shock in Table 2 suggest that the responses of the short-term interest rate are not statistically significant for all periods under scrutiny. This implies that while the Fed is actively monitoring global output movements, it tends to put less weight on international price developments within its reaction function.

By contrast, responses to a global monetary policy shock shown in Figure 10(c) display a more pronounced degree of time variation, especially within the first half of the sample. While short-term interest rates tend to increase on impact, responses after one year tend to be rather negative, varying significantly across time. The somewhat stronger response in the beginning of the sample marks the first half of Paul Volcker's term as Fed chairman, a period where monetary policy started to react aggressively to domestic inflationary developments. Within that regime, the Fed responded to global monetary policy shocks by increasing interest rates by around 3.6 basis points. For the second part of the sample, interest rate responses declined marginally, reaching around 2.2 basis points while in the final part, responses dropped to around 2.0 basis points. Note that the degree of uncertainty surrounding our point estimates was highest within the period marked by the reign of Paul Volcker, whereas the dispersion subsequently declined under Alan Greenspans' and Ben Bernankes' chairmanships. Since policy responses after one year suggest that the Fed is actually lowering interest rates, our findings imply that two mechanisms tend to influence policy making. First, a falling external value of the US dollar would lead to improving terms of trade that would, in general, be beneficial for output growth. Second, a global monetary policy shock would most likely imply that external demand falls. Thus, while our identification scheme rules out contemporaneous reactions of output growth to monetary policy shocks, demand effects materialize after one year. In this case, the Fed reacts by lowering interest rates after one year, thus providing further stimulus to counterbalance the detrimental effects of a fall in external demand.

[Figure 10 about here.]

4 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 vectorautoregression augmented 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 taking explicitly into account that the extent of spillovers might have changed over time. Finally, we ask the reverse question: Does the US Fed respond to international shocks and if yes have these responses changed over time. This part of the analysis is carried out by simulating three global structural shocks and investigate the subsequent response of the US policy rate.

First, 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. This result is in line with Feldkircher et al. (2015). Following the response of the US, global inflation rates tend to decrease. Short-term interest rates, by contrast, follow their US counterparts only in advanced economies, while in emerging economies a lot of short-term interest rates decrease after the policy shock. Naturally, the US tightening causes a nominal appreciation of the US dollar. This appreciation, however, carries also over in real terms. The estimated effects are in line with the empirical literature on the effects of shocks to monetary policy originated in the US on other economies (see Christiano et al., 1999). These results describe global trends in our sample. We observe, however, a great deal of cross-country heterogeneity regarding the spillovers. Countries that are more strongly affected from US monetary policy shocks comprise Canada and Great Britain considering advanced economies. Emerging economies show a larger degree of heterogeneity and there is no single country that always reacts strongly to the US shock. Depending on the variable under considerations, strong responses are reported for China, Korea, Argentina and Mexico among others.

Second, we find evidence for a changing transmission of monetary policy shocks at the global level over time. The global response to US monetary policy shocks became stronger over the last decade indicating an increase in financial globalization. We also find that it is important to allow for stochastic volatility in macroeconomic variables during crises and turbulent episodes. Many developed economies show decreasing volatility until the mid-2000's – a period dubbed the Great Moderation. With the onset of the Global Financial Crisis, we can observe a resurgence of volatility. Also, emerging economies in Latin America and Asia experienced sharp changes in volatility due to economic crises.

Last, we find evidence that monetary policy in the US responds to global macroeconomic shocks. More specifically, US short-term rates react to global aggregate demand and monetary policy shocks. By contrast, we do not find evidence for US monetary policy reacting to global supply shocks. Taken at face value this implies that global interest movements and changes in global real activity do exert influence on how interest rates are set in the US, while international price movements seem to play a minor role. Moreover, US rates tend to react to global demand in the medium-term, for which we find significant responses throughout our sample period, while responses are only tightly estimated at the short-term horizons regarding to global monetary policy shocks. This might imply that, while the Fed is willing to adjust short-term interest rates in response to fluctuations of international interest rates in the short-run, global interest rates do not determine the Fed's long-term interest rate target. Last, the intensity of which US interest rates responded to global demand and monetary supply shocks has abated over time. This finding might be driven by the fact that the most recent period of our sample covers the global financial crisis and its aftermath – a period that is characterized by strong differences in the monetary policy stance with some central banks, including the Fed, engaging in new forms of monetary easing but to a varying degree.

References

Bernanke BS, Boivin J and Eliasz P (2005) Measuring the Effects of Monetary Policy: A Factor-Augmented Vector Autoregressive (FAVAR) Approach. *The Quarterly Journal of Economics* 120, 387–422

Bloom N (2009) The Impact of Uncertainty Shocks. *Econometrica* 77(3), 623–685

Canova F (2005) The transmission of US shocks to Latin America. *Journal of Applied Econometrics* 20(2), 229–251

Carter CK and Kohn R (1994) On Gibbs sampling for state space models. *Biometrika* 81(3), 541–553

Christiano LJ, Eichenbaum M and Evans CL (1999) Monetary policy shocks: What have we learned and to what end? *Handbook of Macroeconomics* 1, 65–148

Christiano LJ, Eichenbaum M and Evans CL (2005) Nominal rigidities and the dynamic effects of a shock to monetary policy. *Journal of Political Economy* 113(1), 1–45

Clarida R, Galı J and Gertler M (1998) Monetary policy rules in practice: some international evidence. *European Economic Review* 42(6), 1033–1067

Dees S, di Mauro F, Pesaran HM and Smith LV (2007a) Exploring the international linkages of the euro area: a global VAR analysis. *Journal of Applied Econometrics* 22(1), 1–38

Dees S, Holly S, Pesaran HM and Smith VL (2007b) Long Run Macroeconomic Relations in the Global Economy. *Economics - The Open-Access, Open-Assessment E-Journal* 1(3), 1–20

Dees S, Pesaran H, Smith V and Smith RP (2010) Supply, demand and monetary policy shocks in a multi-country New Keynesian Model. Working Paper Series 1239, European Central Bank

Dovern J and van Roye B (2014) International transmission and business-cycle effects of financial stress. *Journal of Financial Stability* 13(0), 1 – 17

Ehrmann M and Fratzscher M (2009) Global Financial Transmission of Monetary Policy Shocks. *Oxford Bulletin of Economics and Statistics* 71(6), 739–759

Eickmeier S and Ng T (2011) How do credit supply shocks propagate internationally? A GVAR approach. Discussion Paper Series 1: Economic Studies 2011,27, Deutsche Bundesbank, Research Centre

Feldkircher M and Huber F (2015) The international transmission of US shocks – Evidence from Bayesian global vector autoregressions. *European Economic Review* forthcoming

Feldkircher M, Moder I and Huber F (2015) Towards a New Normal – How Different Paths of US Monetary Policy affect the World Economy. *Economic Notes* Oesterreichische Nationalbank, mimeo

Fernández-Villaverde J, Guerró-Quintana P, Rubio-Ramírez JF and Uribe M (2011) Risk Matters: The Real Effects of Volatility Shocks. *American Economic Review* 101(6), 2530–61

Frühwirth-Schnatter S (1994) Data augmentation and dynamic linear models. *Journal of Time Series Analysis* 15(2), 183–202

Huber F (2014) Density Forecasting using Bayesian Global Vector Autoregressions with Common Stochastic Volatility Oesterreichische Nationalbank, mimeo

Kamin SB (2010) Financial Globalization and Monetary Policy. International Finance Discussion Papers 1002, Board of Governors of the Federal Reserve System

Kastner G and Frühwirth-Schnatter S (2014) Ancillarity-sufficiency interweaving strategy (ASIS) for boosting {MCMC} estimation of stochastic volatility models. *Computational Statistics and Data Analysis* 76, 408 – 423

Kim S (2001) International transmission of U.S. monetary policy shocks: Evidence from VAR's. *Journal of Monetary Economics* 48(2), 339–372

Lopes HF, McCulloch R and Tsay R (2013) Cholesky Stochastic Volatility Models for High-Dimensional Time Series. Technical report, University of Chicago Booth School of Business, mimeo

Omori Y, Chib S, Shephard N and Nakajima J (2007) Stochastic volatility with leverage: Fast and efficient likelihood inference. *Journal of Econometrics* 140(2), 425–449

Orphanides A (2003) Historical monetary policy analysis and the Taylor rule. *Journal of Monetary Economics* 50(5), 983–1022

Pesaran MH, Schuermann T and Weiner SM (2004) Modeling Regional Interdependencies Using a Global Error-Correcting Macroeconometric Model. *Journal of Business and Economic Statistics* 22, 129–162

Pesaran MH and Shin Y (1998) Generalized impulse response analysis in linear multivariate models. *Economics Letters* 58(1), 17–29

Primiceri GE (2005) Time varying structural vector autoregressions and monetary policy. *The Review of Economic Studies* 72(3), 821–852

R Development Core Team (2011) *R: A Language and Environment for Statistical Computing*. R Foundation for Statistical Computing, Vienna, Austria

Rodgers DP (1985) Improvements in multiprocessor system design. In *ACM SIGARCH Computer Architecture News*, volume 13. IEEE Computer Society Press, 225–231

Rudebusch G and Svensson LE (1999) Policy rules for inflation targeting. In *Monetary policy rules*. University of Chicago Press, 203–262

Rue H (2001) Fast sampling of Gaussian Markov random fields. *Journal of the Royal Statistical Society: Series B (Statistical Methodology)* 63(2), 325–338

Sims CA and Zha T (2006) Were There Regime Switches in U.S. Monetary Policy? *American Economic Review* 96(1), 54–81

Taylor JB (1993) Discretion versus policy rules in practice. *Carnegie-Rochester Conference Series on Public Policy* 39(1), 195–214

Taylor JB (2002) The Monetary Transmission Mechanism and the Evaluation of Monetary Policy Rules. In Loayza N, Schmidt-Hebbel K, Editor) NLS and (Series KSH, eds., *Monetary Policy: Rules and Transmission Mechanisms*, volume 4 of *Central Banking, Analysis, and Economic Policies Book Series*, chapter 2. Central Bank of Chile, 021–046

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)				

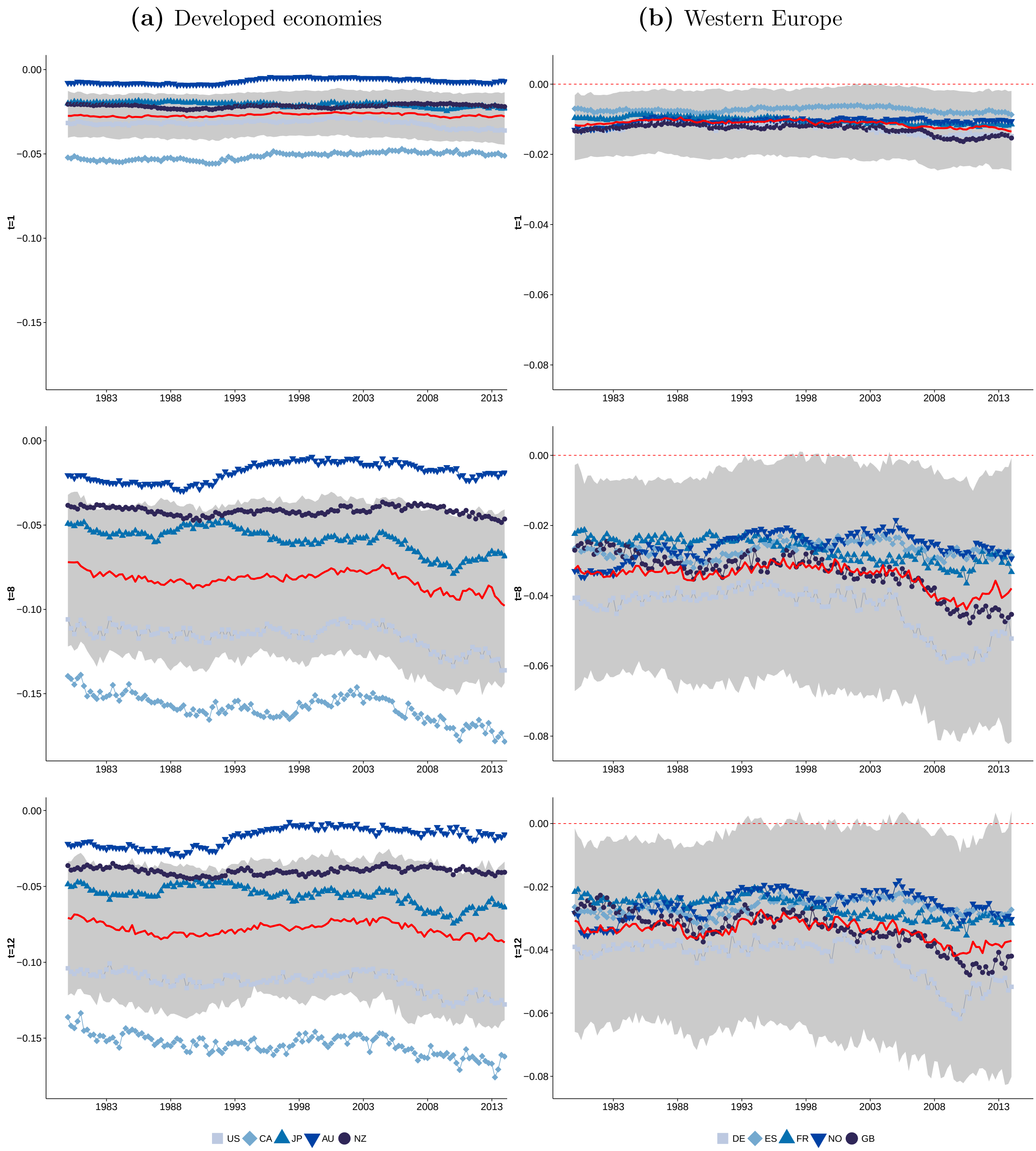
Notes: ISO-2 country codes in parentheses. Empirical results shown for countries in bold.

Table 2: Posterior distribution of US short-term interest rates responses to three global structural shocks

		"Volcker" regime 1979 - 1987			"Greenspan" regime 1987 - 2006			"Bernanke" regime 2006 - 2013		
		$\text{Low}_{0.25}$	Median	$\text{High}_{0.75}$	$\text{Low}_{0.25}$	Median	$\text{High}_{0.75}$	$\text{Low}_{0.25}$	Median	$\text{High}_{0.75}$
AD	$t=0$	-3.18	2.38	8.95	-4.02	1.40	7.60	-3.67	1.45	7.56
	$t=4$	0.20	4.62	9.66	0.42	4.30	9.08	-0.07	3.89	8.82
	$t=8$	-1.45	0.20	2.57	-1.51	0.16	2.39	-1.72	0.35	2.74
AS	$t=0$	-4.67	1.16	7.42	-3.85	0.69	6.09	-3.31	0.58	5.64
	$t=4$	-4.31	2.59	10.92	-3.28	2.22	9.61	-3.02	1.71	8.39
	$t=8$	-4.06	-0.68	3.64	-3.44	-0.23	3.79	-3.05	-0.02	3.98
MP	$t=0$	0.78	3.65	6.79	0.34	2.44	4.78	0.93	2.09	3.39
	$t=4$	-8.93	-3.53	1.61	-6.79	-2.87	1.05	-2.31	-0.22	1.78
	$t=8$	-6.04	-1.96	2.08	-4.52	-1.34	1.53	-1.48	-0.17	1.20

Notes: The table presents the posterior distribution of the impulse responses associated with a global aggregate demand, supply and monetary policy shock (one standard error). Results are based on 1,500 posterior draws from a total chain of 30,000 iterations. Responses in basis points.

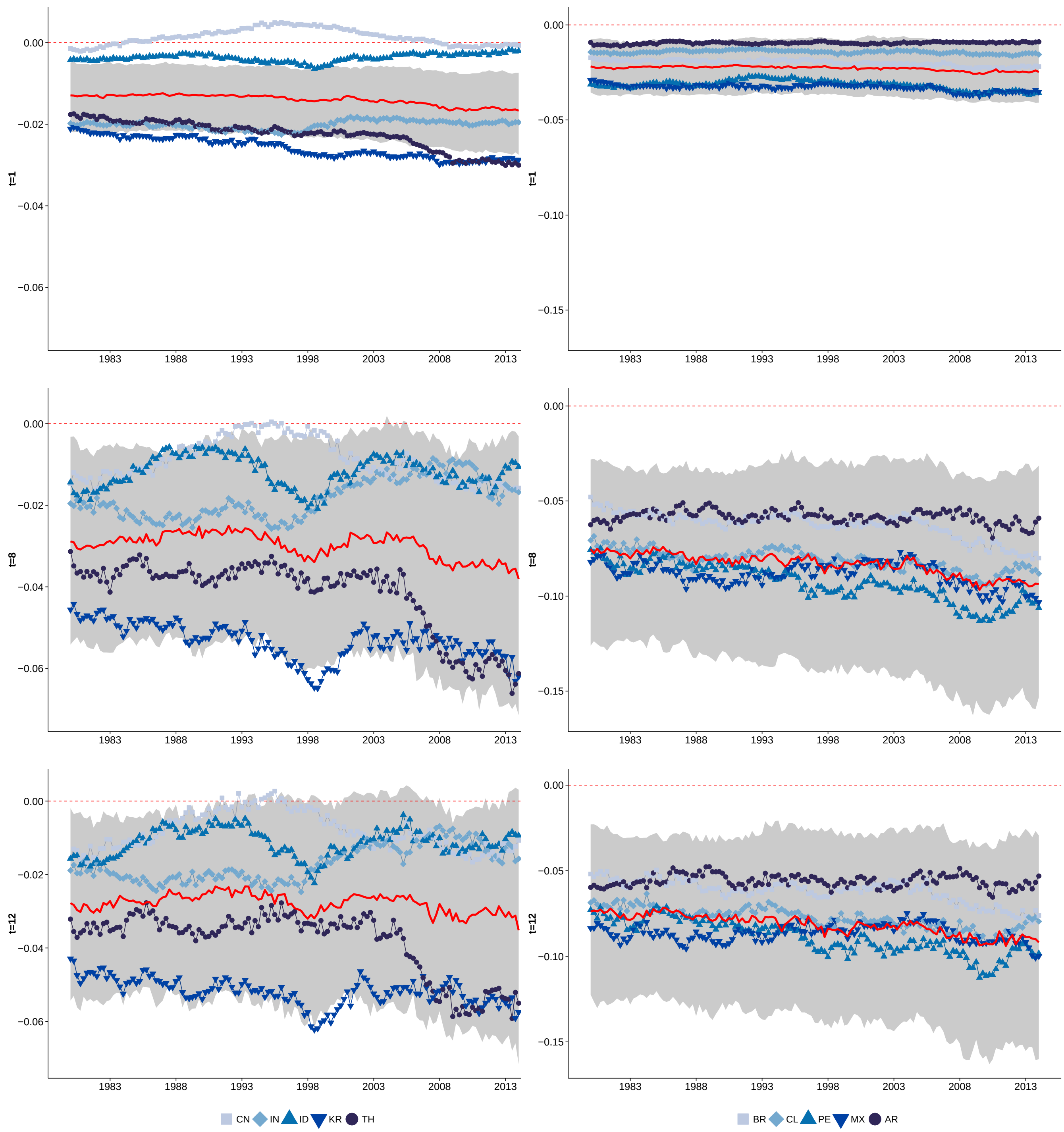
Figure 1: Output responses to a +50 basis point (bp) monetary policy shock in the US



Notes: The plots show the posterior for selected countries along with the cross-country means (in red) and associated 25% and 75% credible sets (gray shaded regions)for three distinct horizons, namely after one quarter ($t = 1$), after 8 quarters ($t = 8$) and after 12 quarters ($t = 12$). Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

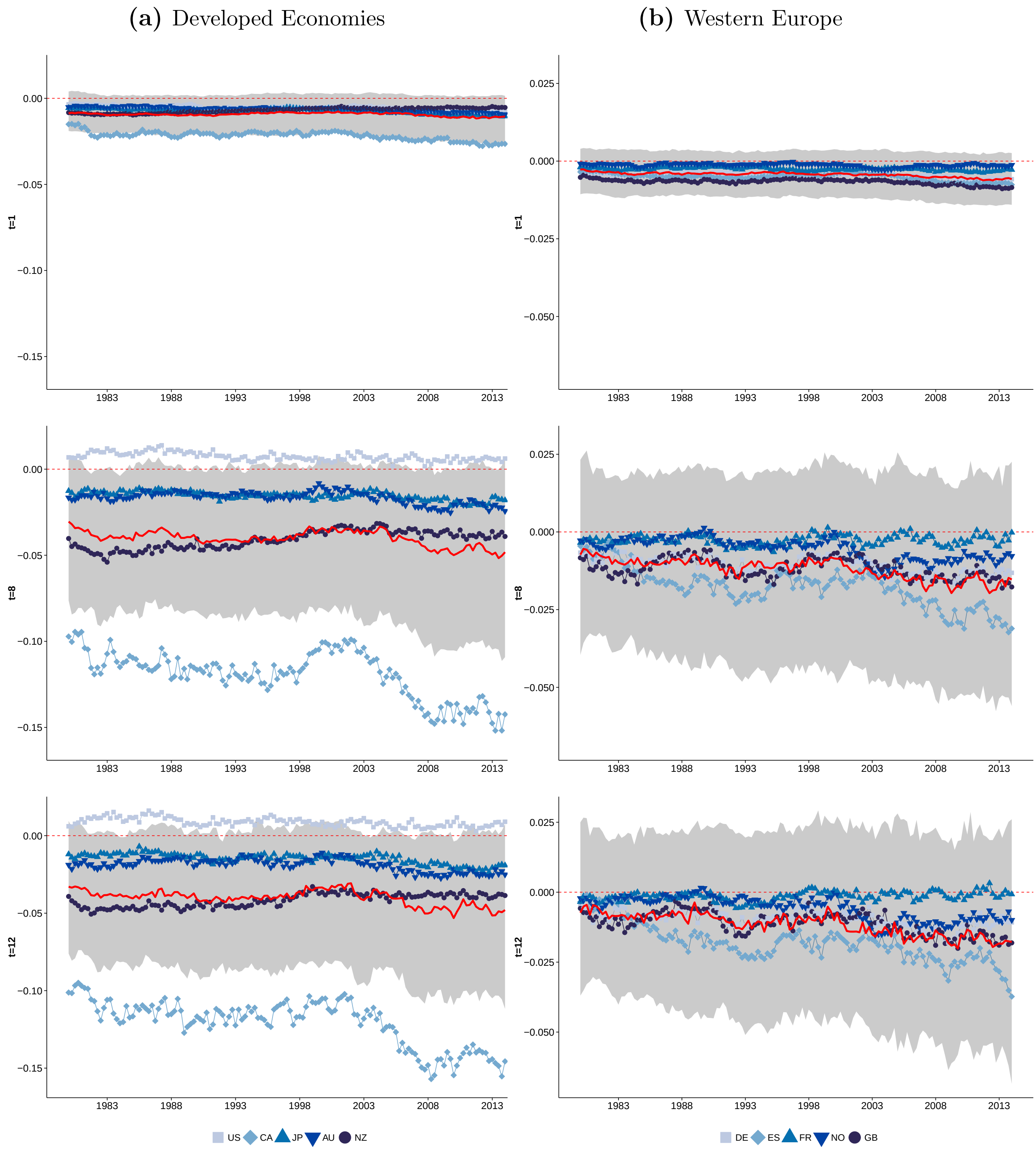
Figure 2: Output responses to a +50 basis point (bp) monetary policy shock in the US

(a) Emerging Asia **(b)** Latin America



Notes: The plots show the posterior for selected countries along with the cross-country means (in red) and associated 25% and 75% credible sets (gray shaded regions)for three distinct horizons, namely after one quarter ($t = 1$), after 8 quarters ($t = 8$) and after 12 quarters ($t = 12$). Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

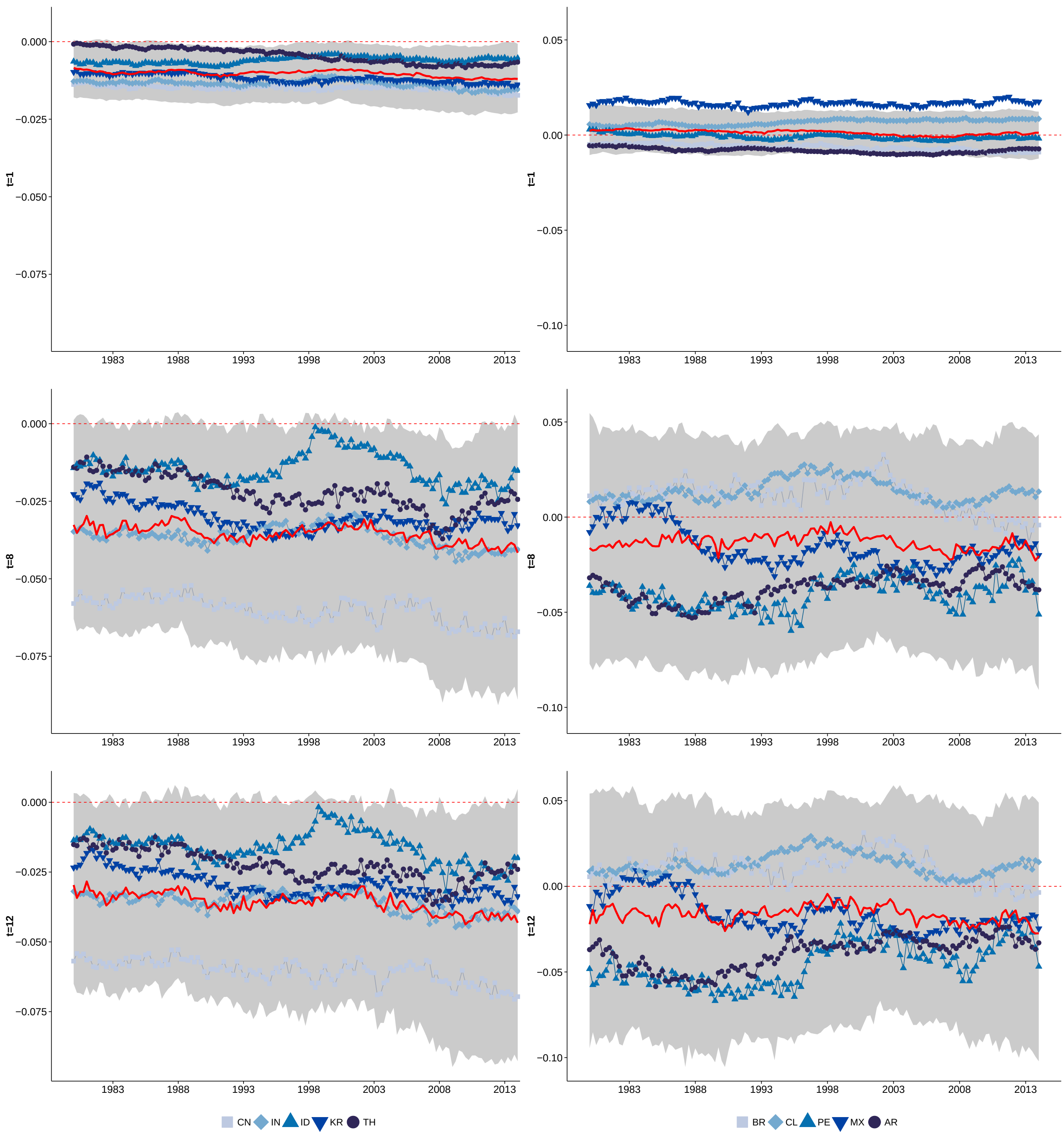
Figure 3: Inflation responses to a +50 basis point (bp) monetary policy shock in the US



Notes: The plots show the posterior for selected countries along with the cross-country means (in red) and associated 25% and 75% credible sets (gray shaded regions)for three distinct horizons, namely after one quarter ($t = 1$), after eight quarters ($t = 8$) and after 12 quarters ($t = 12$). Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

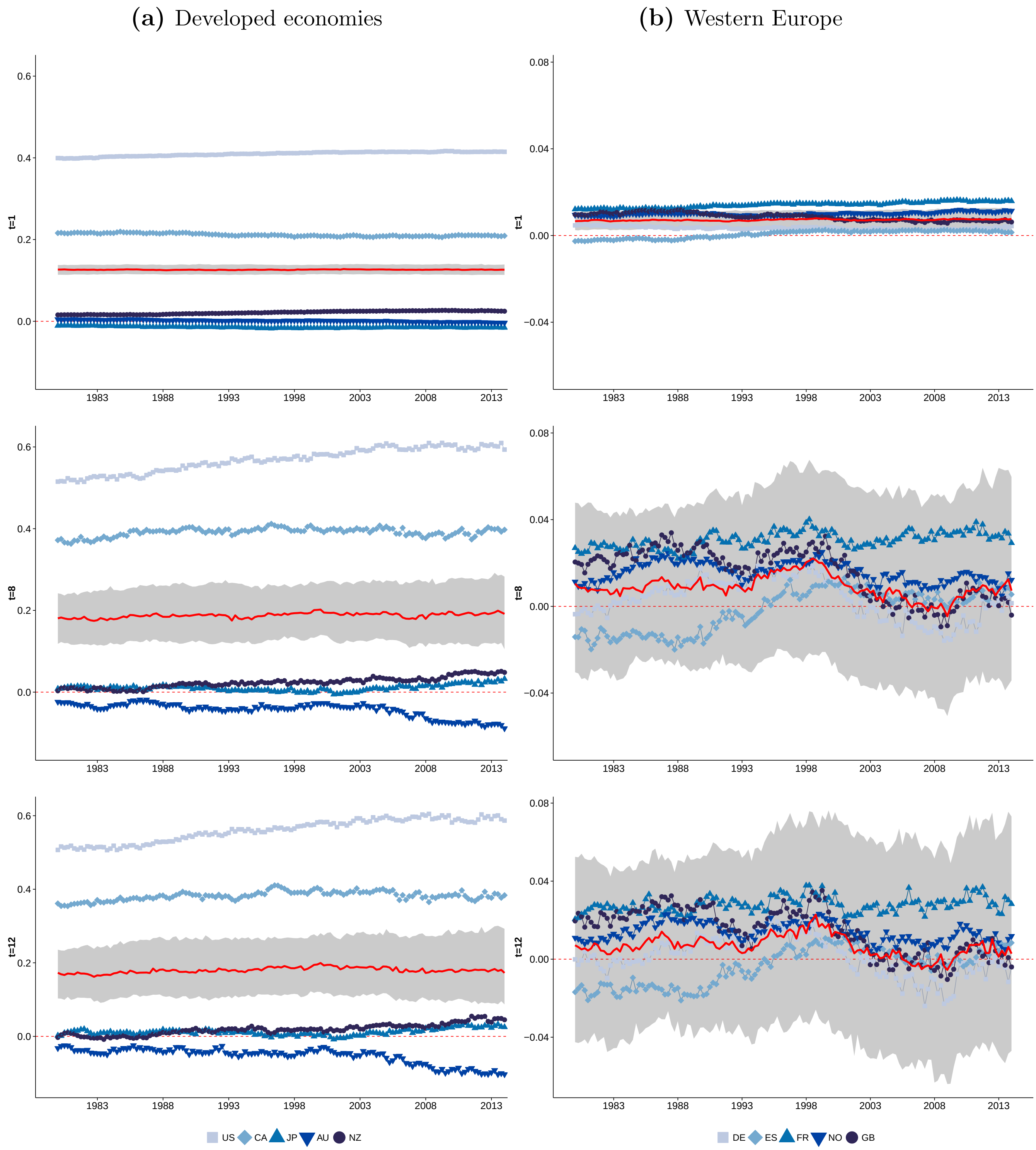
Figure 4: Inflation responses to a +50 basis point (bp) monetary policy shock in the US

(a) Emerging Asia **(b)** Latin America



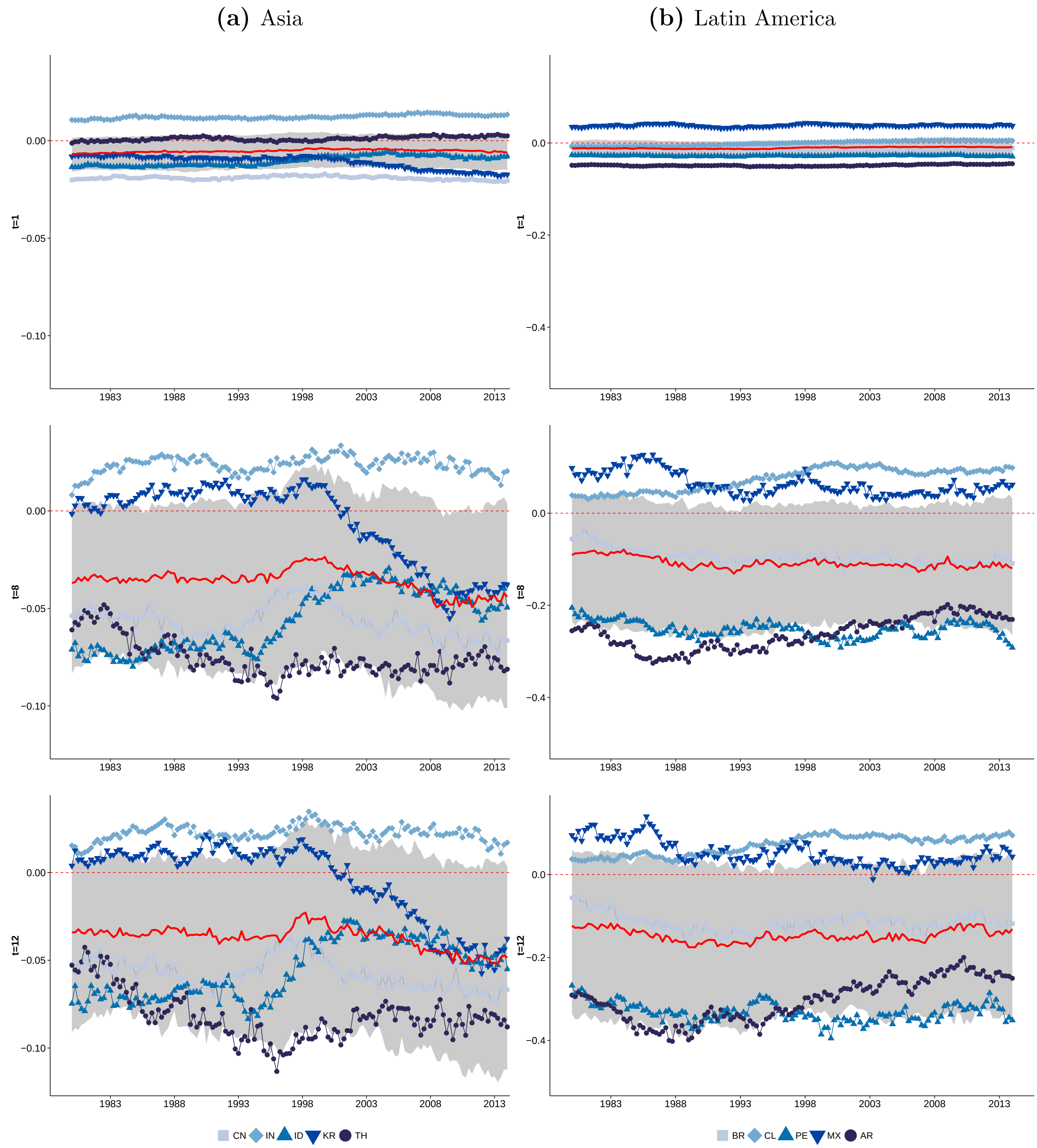
Notes: The plots show the posterior for selected countries along with the cross-country means (in red) and associated 25% and 75% credible sets (gray shaded regions)for three distinct horizons, namely after one quarter ($t = 1$), after eight quarters ($t = 8$) and after 12 quarters ($t = 12$). Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

Figure 5: Short-term interest rate responses to a +50 basis point (bp) monetary policy shock in the US



Notes: The plots show the posterior for selected countries along with the cross-country means (in red) and associated 25% and 75% credible sets (gray shaded regions)for three distinct horizons, namely after one quarter ($t = 1$), after eight quarters ($t = 8$) and after 12 quarters ($t = 12$). Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

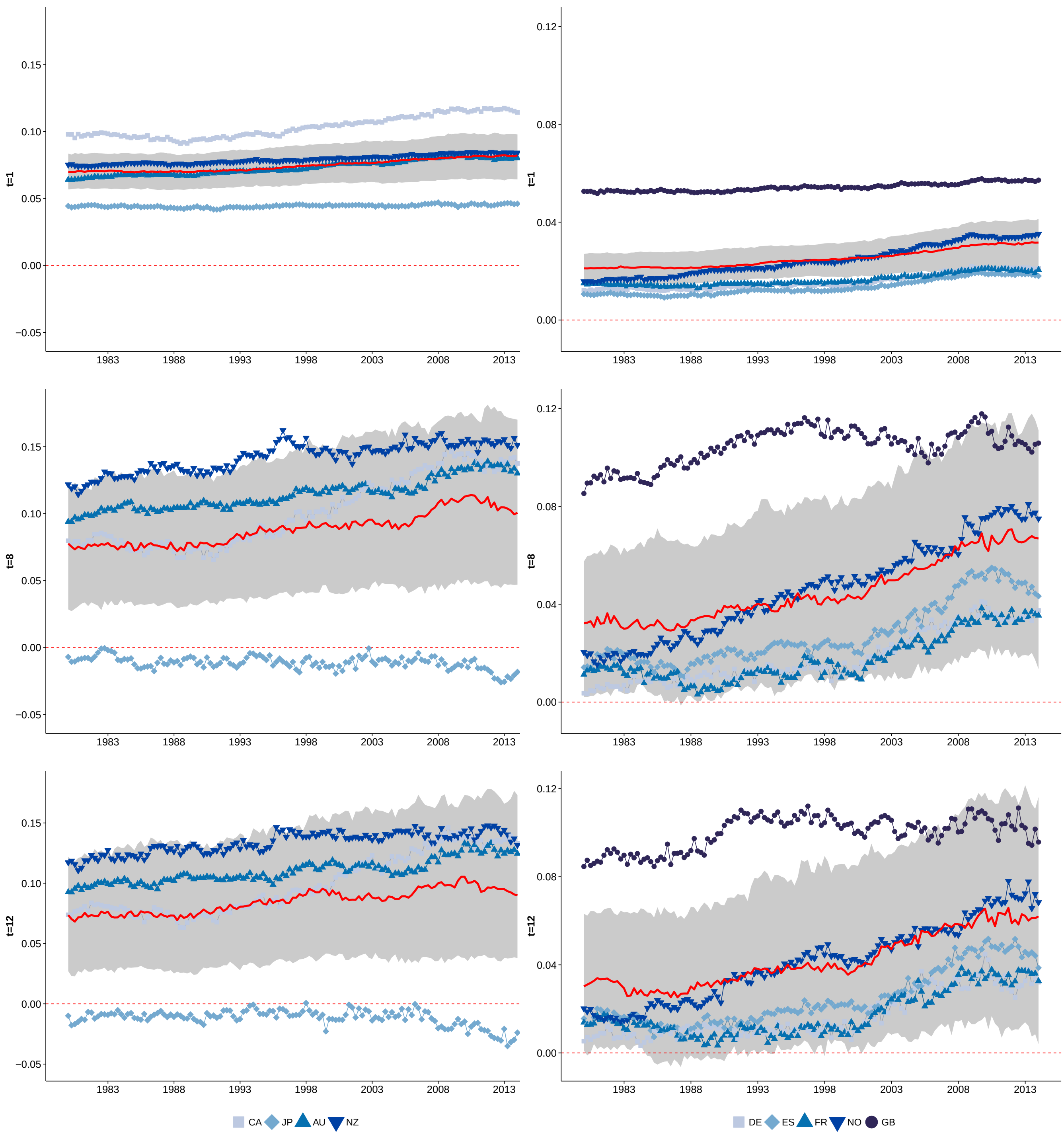
Figure 6: Short-term interest rate responses to a +50 basis point (bp) monetary policy shock in the US



Notes: The plots show the posterior for selected countries along with the cross-country means (in red) and associated 25% and 75% credible sets (gray shaded regions)for three distinct horizons, namely after one quarter ($t = 1$), after eight quarters ($t = 8$) and after 12 quarters ($t = 12$). Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

Figure 7: Real exchange rate responses to a +50 basis point (bp) monetary policy shock in the US

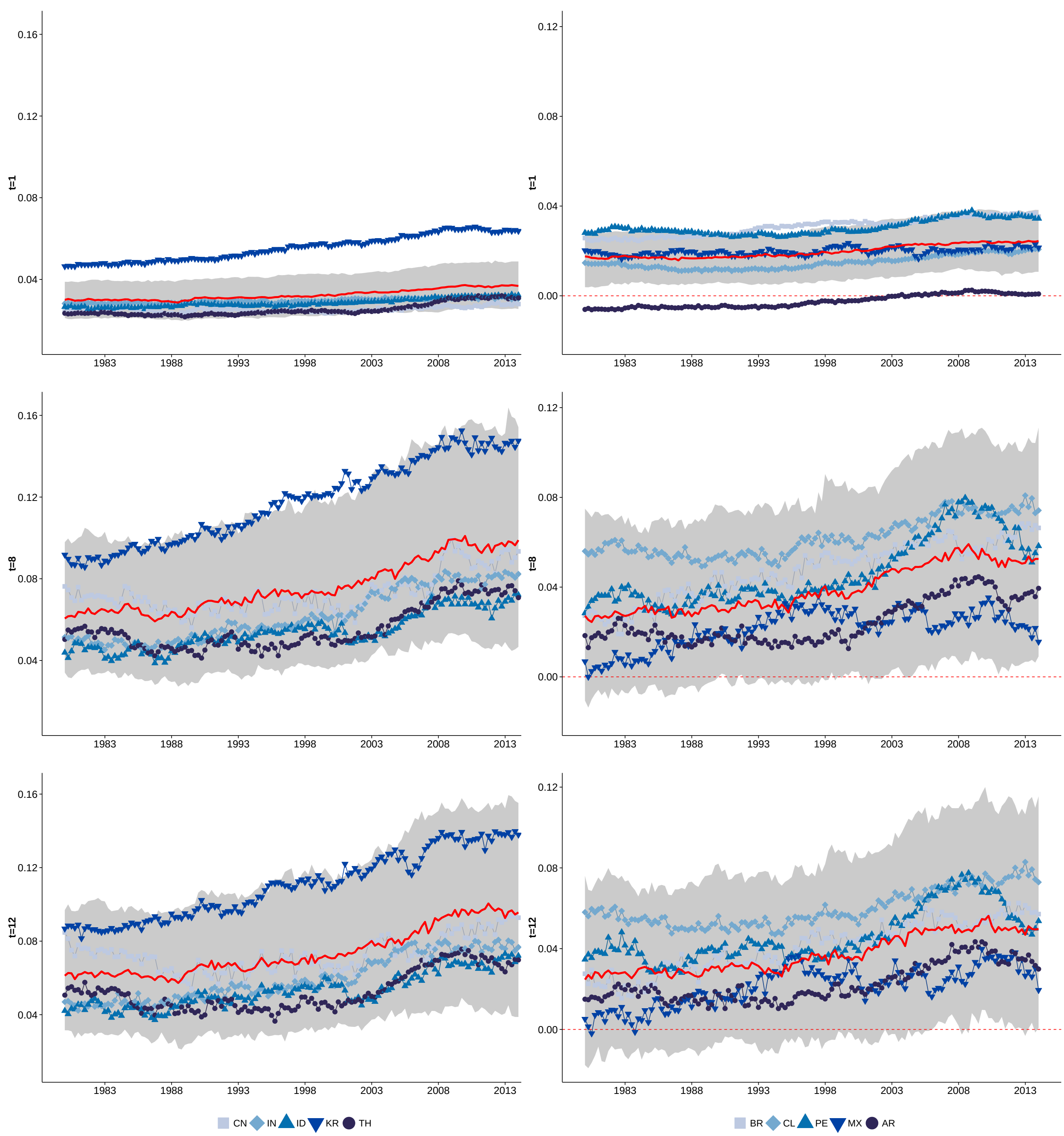
(a) Developed economies **(b)** Western Europe



Notes: The plots show the posterior for selected countries along with the cross-country means (in red) and associated 25% and 75% credible sets (gray shaded regions)for three distinct horizons, namely after one quarter (t = 1), after eight quarters (t = 8) and after 12 quarters (t = 12). Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

Figure 8: Real exchange rate responses to a +50 basis point (bp) monetary policy shock in the US

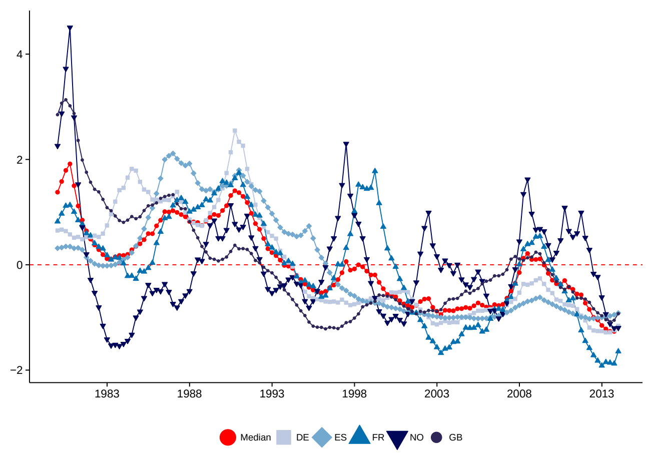
(a) Asia **(b)** Latin America



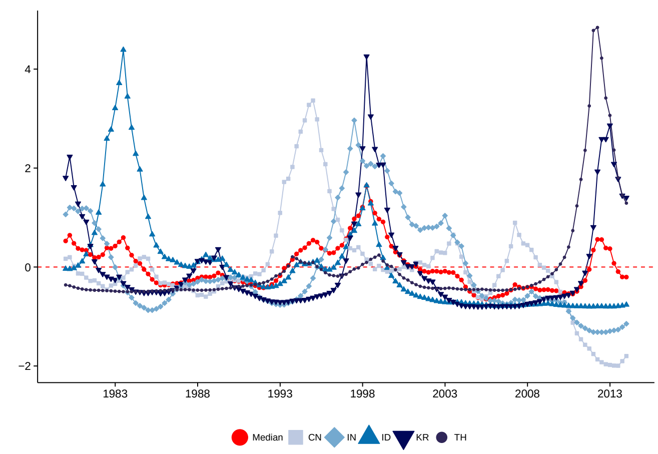
Notes: The plots show the posterior for selected countries along with the cross-country means (in red) and associated 25% and 75% credible sets (gray shaded regions). Responses are based on 1,500 posterior draws from a total chain of 30,000 posterior draws, shown for three distinct horizons, namely after one quarter ($t = 1$), after eight quarters ($t = 8$) and after 12 quarters ($t = 12$).

Figure 9: Volatility of GDP growth

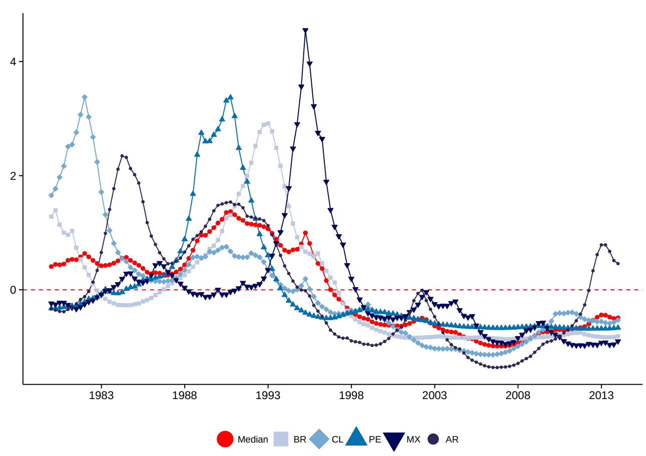
(a) Western Europe



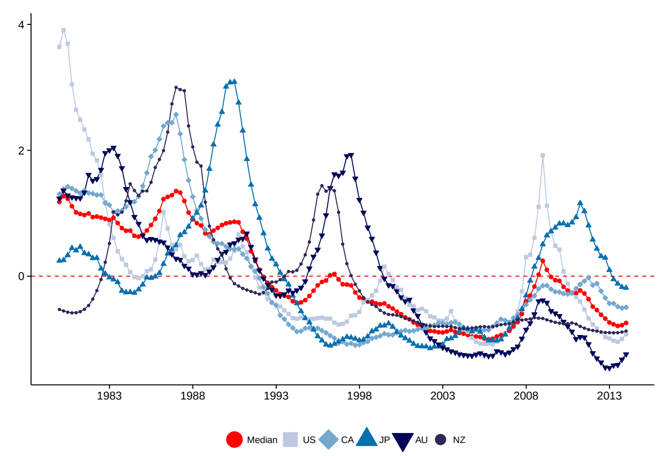
(b) Asia



(c) Latin America

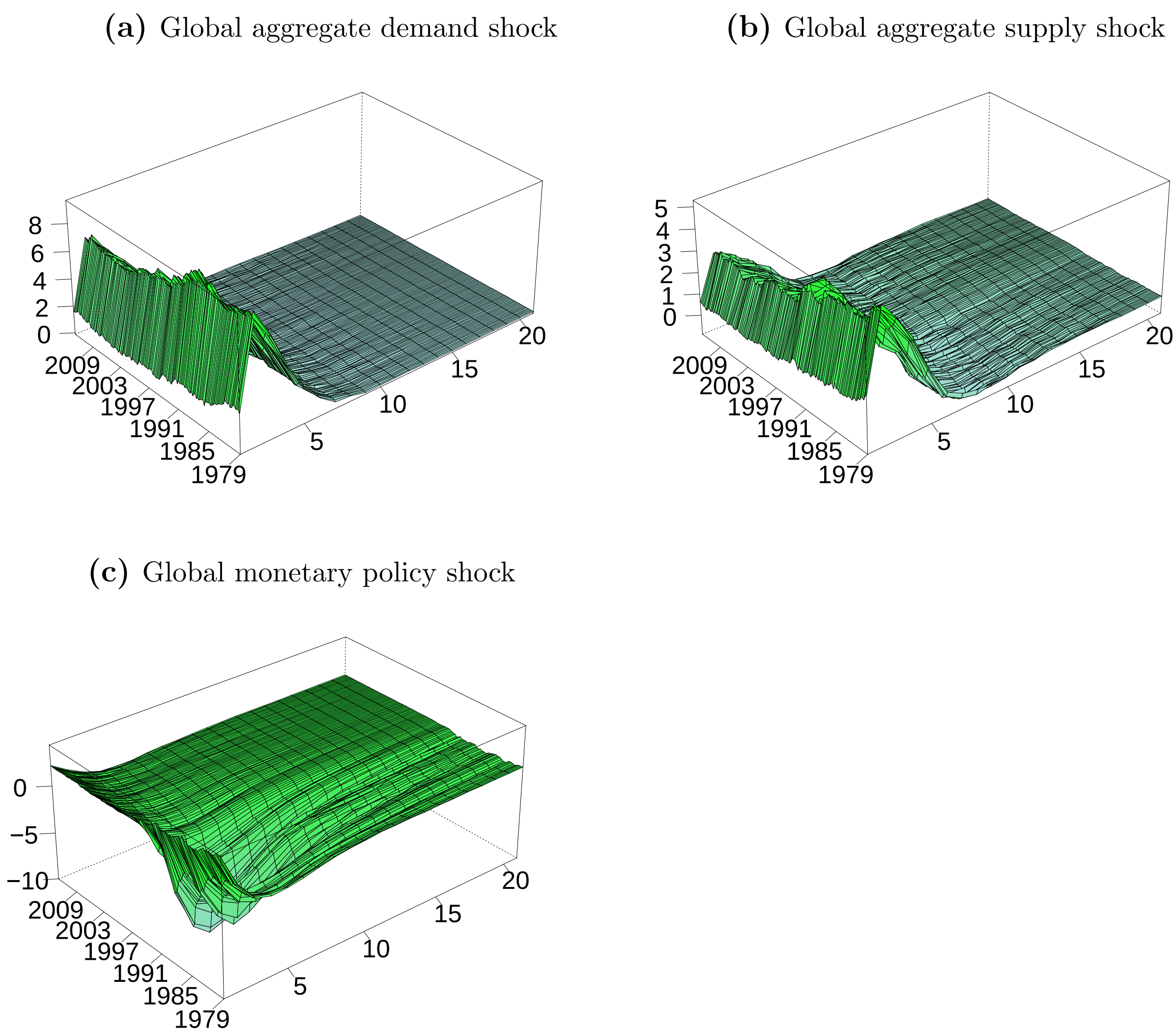


(d) Developed economies



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 10: US short-term interest rate responses to global shocks



Notes: The plots depict the posterior mean response (in basis points) of the short-term interest rates with respect to a global aggregate demand, supply and monetary policy shock. Responses are based on 1,500 posterior draws from a total chain of 30,000 draws.

Appendix A

A.1 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 (2014), 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

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

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 (2014) consider $\lambda_{ij,t}$ in its centered parametrization given in equation (2.8) and in its non-centered form given by

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

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

Let us consider the jth equation of equation (A.1). Squaring and taking logs yields

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

Since $\ln(u_{i,t}^2) \sim \log \chi^2(1)$, 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 $r_{ij,t} \in \{1, \ldots, 10\}$. $m_{r_{i,t}}$ 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,

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

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 (2014). 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 $\ln(\lambda_{ij,-1}) = (\ln(\lambda_{ij,2}), \ldots, \ln(\lambda_{ij,T}))'$ in terms of a multivariate normal distribution

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

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

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

In this expression, the posterior moments are given by

$$\Omega_{\lambda_{ij}} = \begin{pmatrix} \frac{1}{s^2_{r_{ij,2}}} + \frac{1}{\varsigma^2_{ij}} & \frac{-\rho_{ij}}{\varsigma^2_{ij}} & 0 & \cdots & 0 \\ -\frac{\rho_i}{\varsigma^2_{ij}} & \frac{1}{s^2_{r_{i,3}}} + \frac{1+\rho_{ij}}{\varsigma^2_i} & -\frac{\rho_{ij}}{\varsigma^2_{ij}} & \ddots & \vdots \\ 0 & -\frac{\rho_{ij}}{\varsigma^2_{ij}} & \ddots & \ddots & 0 \\ \vdots & \ddots & \ddots & \frac{1}{s^2_{r_{ij,T-1}}} + \frac{1+\rho_{ij}}{\varsigma^2_{ij}} & \frac{-\rho_{ij}}{\varsigma^2_{ij}} \\ 0 & \dots & 0 & -\frac{\rho_{ij}}{\varsigma^2_{ij}} & \frac{1}{s^2_{r_{ij,T}}} + \frac{1}{\varsigma^2_{ij}} \end{pmatrix} \tag{A.7}$$

and

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

The moments for the non-centered case are given by $\tilde{\Omega}_i = \varsigma^2_{ij}\Omega_{h_{ij}}$ and $\tilde{c}_{ij} = \varsigma^2_{ij}c_{ij}$. 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 ς^2_{ij} 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 $\tilde{y}^2_{ij,t} - \ln(\lambda_{ij,t}) = \tilde{u}_{ij,t}$ with $\tilde{u}_{ij,t} \sim \mathcal{N}(m_{r_{ij,t}}, s^2_{r_{ij,t}})$. Posterior probabilities for each $r_{ij,t}$ are then given by

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

where $p(r_{ij,t} = c)$ is the weight associated with the cth 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 (2014).[10]

A.2 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.

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 c processors. In this case, the maximum speedup gained by parallelization is given by

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

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 *embarassingly 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 c equals k, computation time almost boils down to that required for estimating the equation with the maximum number of parameters.

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