IWH Halle Institute for Economic Research
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Discussion Papers

No. 3
January 2023

Active Driver or Passive Victim – On the Role of International Monetary Policy Transmission

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ISSN 2194-2188

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Abstract

We provide new insights into determinants of international interest rates spillovers across seven advanced economies. To disentangle and quantify their respective importance, we identify country-specific structural monetary policy, demand, and supply equations in a Bayesian structural panel vector autoregressive model. We formulate prior beliefs on magnitudes and signs of contemporaneous structural coefficients (i.e., (semi-)elasticities), based on a standard theoretical multi-country open economy model from the literature. Our findings show that interest rate spillovers occur via an aggregated demand channel. Unexpected monetary tightening causes modest declines in most foreign interest rates, while demand and supply shocks result in increased foreign interest rates. Our results support that central banks respond to changes in the domestic macroeconomic environment induced by domestic or foreign shocks rather than directly reacting to foreign shocks. Spillovers are quantitatively stronger for shocks originating in economically large areas with strong trade linkages.

Keywords: informative priors, panel vector autoregressions, spillovers, structural vector autoregressions

JEL classification: C11, C30, E52, F42

1 Introduction

A large literature reports that interest rate movements spill over internationally. However, there is little evidence on the exact mechanisms behind this effect. On the one hand, interest rate changes in one country might spill over to other countries via exchange rate movements. Under such a mechanism, monetary policy would be an active driver of international interest rate developments. This might risk undesired real cross-border effects, if, for example, a tightening in the US causes other countries to raise interest rates as well even if macroeconomic conditions would caution against such a move. On the other hand, there might be real cross-border effects, for example because domestic aggregate demand shifts with foreign output. Even if monetary policy only reacts to domestic developments, interest rates in several countries might move in a seemingly coordinated fashion in this case – monetary policy would be a passive victim of foreign developments. Thus, potential spillovers via real channels would call for monetary policy to account for this "beggar-thy-neighbour" effect.[1] In this paper, we disentangle different determinants of international reactions in interest rates and quantify their respective importance.

We show that interest rate spillovers materialize mostly via an aggregated demand channel, whereby increased foreign output raises aggregate demand. Thus, contractionary monetary policy shocks cause modest negative reactions in the majority of foreign interest rates, while expansionary supply and demand shocks raise foreign interest rates considerably with a time lag. In comparison to spillovers via aggregate demand, the exchange rate channel is rather unimportant according to our results. Our findings support that monetary policy authorities in our sample mainly react endogenously to the domestic economic reactions (caused both by domestic and foreign shocks) rather than reacting directly to foreign shocks.

We derive our results from a Bayesian structural panel vector autoregressive (BSPVAR) model including Australia (AU), Canada (CA), the Euro area (EA), Japan (JP), South

[1] For example, Kim (2001), Maćkowiak (2007), Gambacorta et al. (2014), Bluwstein and Canova (2016), Dedola et al. (2017), Rogers et al. (2018), Crespo Cuaresma et al. (2019), De Santis and Zimic (2022) show spillovers in interest rates as results of foreign monetary policy shocks. Feldkircher and Huber (2016) and De Santis and Zimic (2022) provide evidence for interest rate spillovers driven by foreign supply and demand shocks. Furthermore, Hausman and Wongswan (2011), Glick and Leduc (2012), Bauer and Neely (2014), Neely (2015), and Fratzscher et al. (2018), among others, report cross-border *financial* spillover effects of US monetary policy shocks.

Korea (KO), the United Kingdom (UK), and the United States (US).[2] We use data on the output gap, inflation, shadow rates and real effective exchange rates between 1980:Q3 and 2019:Q4. In our setup, it is possible to jointly identify country-specific structural monetary policy, supply and demand equations, which allows us to quantify international endogenous and exogenous reactions of monetary policy. Specifically, we develop an open-economy version of the model of Baumeister and Hamilton (2018) and extend it to a multi-country framework. The Bayesian SVAR approach of Baumeister and Hamilton (2015, 2018) is notable in its reliance on informative priors on structural contemporaneous relationships; i.e., (semi-)elasticities of structural economic equations. To derive these priors, we rely on the open economy model of Lubik and Schorfheide (2007) and its closed economy version of Lubik and Schorfheide (2004) as main sources of information (the latter one is used in Baumeister and Hamilton, 2018).

Our paper provides new evidence on cross-border interest rate reactions to monetary policy, supply, and demand shocks and into international real effects of monetary policy shocks. We show differences in spillover effects of monetary policy shocks according to the origin country of the shock. The main transmitter countries are the EA and US while shocks from smaller countries result in virtually no spillovers. Furthermore, economies respond to unexpected monetary tightening in countries with close trade links stronger – spillover effects are strong for the EA to UK or the US to CA. Moreover, our findings highlight that international interest rate transmission seems to work mainly via an aggregated demand channel. To that end, we show that domestic contractionary monetary policy shocks induce considerable reductions in foreign output gaps and inflation, which causes an expansionary foreign monetary policy response. Likewise, expansionary domestic supply and demand shocks mainly cause foreign monetary policy reactions through their positive effect on foreign output gaps. We further quantify the importance of the drivers by means of forecast error variance decompositions. Across our sample, domestic shocks explain around two thirds of long-run forecast error variances in interest rates.

Our PVAR model generates direct insights into drivers and responses of all included countries and variables. This distinguishes our paper from a large part of the literature where many studies limit their analysis to setups with two countries (e.g., Kim and Roubini,

[2]Even though the EA is not a country but a group of countries, we still refer to it as such from now on for simplicity.

2000; Kim, 2001; Canova, 2005; Maćkowiak, 2007) or cross-country effects of a single country's structural shocks (e.g. Dees et al., 2007; Georgiadis, 2015; Feldkircher and Huber, 2016; Burriel and Galesi, 2018; Crespo Cuaresma et al., 2019). Similar to our paper, Gerko and Rey (2017), Rogers et al. (2018), and Liu et al. (2022) analyze international spillovers of monetary policy shocks of multiple countries, paying no attention though to cross-border effects of supply and demand shocks which we find quantitatively to be slightly more important than monetary policy shocks. In terms of results, we can compare our paper to a large literature on the international effects of US monetary policy shocks. Our finding of the importance of the aggregate demand channel implies a negative comovement of foreign interest rates in reaction to domestic monetary policy shocks. This is in line with Feldkircher and Huber (2016); Dedola et al. (2017); Crespo Cuaresma et al. (2019), but contrasts with findings of no (Gerko and Rey, 2017; Liu et al., 2022) or even positive comovement (Rogers et al., 2018; De Santis and Zimic, 2022).

In general, structural multi-country VAR models such as ours face an identification and estimation challenge. The identification challenge results from the presence of multiple country-specific structural shocks. The estimation challenge comes from the large number of free parameters as the model includes variables of all countries jointly. We tackle those challenges by working directly on the structural form of the VAR model. To facilitate this, we rely on two types of prior knowledge. First, consistent with the literature we use trade-weighted averages of foreign variables (instead of each foreign country separately) in each structural equation. This approach implies homogeneity restrictions on structural contemporaneous and autoregressive coefficients associated with foreign terms, and thereby effectively removes the curse-of-dimensionality that usually arises from extending the panel dimension of the model. While the restrictions are in the spirit of those imposed on the reduced form in global VAR (GVAR) models (see, e.g., Pesaran et al., 2004, 2009; Crespo Cuaresma et al., 2016), they are in our opinion easier to defend since homogeneity restrictions on structural model parameters do in general not imply similar restrictions on reduced form coefficients (and vice versa).

Second, we use prior information from theoretical open-economy models to identify the structural equations in our PVAR as country-specific open-economy Phillips curves, IS curves, monetary policy rules, and exchange rate equations. This prior knowledge implies economically meaningful exclusion restrictions on some contemporaneous structural coefficients. Moreover, it allows us to formulate informative prior distributions on remaining contemporaneous structural coefficients. Thereby, we incorporate *identification* uncertainty around restrictions accounting for a lack of conclusive theoretical evidence and avoid unjustifiable recursive structures (as used in, for example, Chen et al., 2016; Kucharukov et al., 2016; Bluwstein and Canova, 2016, where the order of countries matters) or block-exogeneity (as in Kim and Roubini, 2000; Kim, 2001; Canova, 2005; Maćkowiak, 2007). Compared to studies applying sign and/or magnitude restrictions in multi-country models (Gambacorta et al., 2014; Liu et al., 2022; De Santis and Zimic, 2022), we clearly acknowledge the uncertainty around those restrictions. To the best of our knowledge, we are the first who draw inference on a fully identified structural PVAR model (i.e., a model where all country-specific structural equations are identified) without imposing restrictions on the reduced form. Moreover, direct inference on the structural model implies that conditional on structural contemporaneous coefficients (i.e., elasticities and semi-elasticities), we can estimate the model equation-by-equation, which reduces *estimation* uncertainty.

2 Bayesian structural PVAR model

We use a structural PVAR model including Australia (AU), Canada (CA), the Euro Area (EA), Japan (JP), South Korea (KO), the United Kingdom (UK) and the United States (US); this set of countries represents around 54% of the world's economic activity as of 2019. We aim to model country-specific supply, demand, monetary policy, and exchange rate equations. For each country c, we include the output gap (y_{ct}) as a measure of economic output, year-on-year inflation rates (π_{ct}) and Krippner-shadow interest rates (r_{ct}) which capture both conventional and unconventional monetary policy actions (Krippner, 2013). Additionally, we include year-on-year growth rates of the real effective exchange rates (σ_{ct}). The real effective exchange rates are defined such that an increase in σ_{ct} indicates an increase in competitiveness. Our sample spans from 1980:Q3 to 2019:Q4 using quarterly observations.[3]

[3]An Online Appendix contains additional information. In particular, Online Appendix A explains the data in more detail. Two country selections deserve note. First, we rely on constructed data (provided by Eurostat, the ECB and Oxford Economics) for a counterfactual Euro Area between 1980 and 1999. Second, we exclude China. However, especially for the first part of our sample, there are issues with the availability and quality of Chinese data. As China is not included in the narrow real effective exchange

Stacking the $n_c = 4$ variables from each country $c \in \{1, \dots, C\}$, we formulate a structural panel VAR with $p = 4$ lags. We include a deterministic trend to counter the trending behavior of (mostly) shadow rates in our comparatively short sample. Thus, our baseline model contains a total of $n = Cn_c = 28$ endogenous variables captured in the $n \times 1$ vector $\mathbf{y}_t = (\mathbf{y}_{1t}, \dots, \mathbf{y}_{Ct})'$ with $\mathbf{y}_{ct} = (y_{ct}, \pi_{ct}, r_{ct}, \sigma_{ct})'$. The right-hand-side variables (p lags, a constant and trend) are collected in the $k \times 1$ vector $\mathbf{x}_{t-1}$, with $k = Cn_c p + 2 = 114$.[4] The model has the following compact formulation:

$$\mathbf{A}\mathbf{y}_t = \mathbf{B}\mathbf{x}_{t-1} + \mathbf{u}_t. \tag{1}$$

The $n \times n$-matrix $\mathbf{A}$ contains structural contemporaneous parameters and the $n \times k$-matrix $\mathbf{B}$ the structural lag coefficients. The $n \times 1$ vector of structural shocks $\mathbf{u}_t$ follows a normal distribution with mean zero and diagonal variance matrix $\mathbf{D}$. The independence of structural shocks implies that – conditional on $\mathbf{A}$ – the model can be estimated equation-by-equation. For later use, let $\mathbf{a}_i$, $\mathbf{b}_i$ and d_{ii} denote the structural coefficients (contemporaneous and lagged) as well as the variance of structural shocks in equation $i \in \{1, \dots, n\}$.

To fully specify the contemporaneous relations among our included variables, we rely on structural relations derived in theoretical open-economy models such as Lubik and Schorfheide (2007). Specifically, we formulate an empirical open economy Phillips curve, IS curve, monetary policy rule, and an exchange rate equation for every country $c \in \{1, \dots, N\}$:

$$y_{ct} = \alpha^{c,\pi}\pi_{ct} + \alpha^{c,\sigma}\sigma_{ct} + \mathbf{b}'_{c,s}\mathbf{x}_{t-1} + u^s_{ct} \tag{AS}$$

$$y_{ct} = \beta^{c,r}r_{ct} + \beta^{c,\pi}\pi_{ct} + \beta^{c,\sigma}\sigma_{ct} + \beta^{c,y^*}y^*_{ct} + \mathbf{b}'_{c,d}\mathbf{x}_{t-1} + u^d_{ct} \tag{AD}$$

$$r_{ct} = (1 - \rho^c)\left(\psi^{c,\pi}\pi_{ct} + \psi^{c,y}y_{ct} + \psi^{c,\sigma}(\sigma_{ct} - \pi^*_{ct})\right) + \mathbf{b}'_{c,m}\mathbf{x}_{t-1} + u^m_{ct} \tag{MP}$$

$$\sigma_{ct} = \theta^{c,y}(y_{ct} - y^*_{ct}) + \theta^{c,\sigma^*}\sigma^*_{ct} + \mathbf{b}'_{c,\sigma}\mathbf{x}_{t-1} + u^\sigma_{ct}. \tag{ER}$$

Given four equations per country, and seven countries in our sample, we thus have the above described fully-identified system of 28 structural equations. The vector $\mathbf{b}_{c,j}$, $j \in$

rates of the BIS, trade weights are not distorted.

[4] In Online Appendix F, we show the main results for models with $p = 8$ lags and without a trend. While the main findings hold, excluding a trend leads to diverging posterior credibility sets visible for the EA, Figures F.31 – F.33. The results for 8 lags are robust (Figures F.28 – F.30).

$\{s, d, m, \sigma\}$, denotes the corresponding row of the matrix of autoregressive parameters, $\mathbf{B}$. We define the structural country-specific supply, demand, and monetary policy shocks $(u^s_{ct}, u^d_{ct}, u^m_{ct}, u^\sigma_{ct})$ as the structural error terms of the country-specific Phillips curve, IS curve, monetary policy rule, and exchange rate equation, respectively – hence, as exogenous shifts to those equations. The (semi-)elasticities of the structural equations are given by the $\alpha, \beta, \psi, \theta$-coefficients. Note that we aggregate foreign terms into a single variable and coefficient as in the term "$\beta^{c,y^*} y^*_{ct}$" in the aggregate demand equation. This aggregation is possible under homogeneity restrictions discussed in subsection 2.1. All aggregated foreign variables $(y^*_{ct}, \pi^*_{ct}, \sigma^*_{ct})$ are understood to be trade-weighted averages of country-specific terms.[5] To derive specific coefficients for each foreign country c^*, we only need to multiply the identified "aggregated" coefficients $(\beta^{c,y^*}, \psi^{c,\sigma}, \theta^{c,y}, \theta^{c,\sigma^*})$ with the corresponding trade weight w_{cc^*}.

As in Baumeister and Hamilton (2015), the joint prior distribution of the structural model parameters equals

$$p(\mathbf{A}, \mathbf{B}, \mathbf{D}) = p(\mathbf{A})p(\mathbf{D}|\mathbf{A})p(\mathbf{B}|\mathbf{A}, \mathbf{D}). \tag{2}$$

To facilitate direct inference on the structural VAR model, we rely on informative prior knowledge on the structural contemporaneous coefficients in $\mathbf{A}$. The approach of Baumeister and Hamilton (2015) offers two distinct advantages in this regard. First, using prior distributions on (semi-)elasticities in $\mathbf{A}$ transparently shows the full information set imposed by us. Second, we choose each of our priors with a specified degree of uncertainty (albeit never "uninformative"). Modeling the uncertainty around the prior knowledge is especially appealing in multi-country models because typically a larger number of identifying restrictions is necessary while at the same time precise prior information from theoretical models is lacking.

In this paper, we rely on two types of prior knowledge derived from theoretical open economy models. First, we formulate the above-mentioned homogeneity restrictions on all structural coefficients on foreign variables, introduced in section 2.1. Second, in section 2.2

[5]Since we use real effective exchange rates and not real exchange rates with respect to the US$ for a limited number of countries, the exchange rates included in our model do not form a closed system (Bussière et al., 2009; Feldkircher and Huber, 2016). That is, σ_{ct} cannot be derived only as a function of all remaining σ_{c^*t} for $c^* \neq c$. Thus we can treat σ_{ct} as a standard endogenous variable and include it in the set of foreign variables for all countries.

we explain how we use theoretical models and empirical estimates (a) to impose exclusion restrictions on some coefficients in **A** as well as (b) to formulate informative prior distributions on the remaining coefficients (those explicitly mentioned in equations (AS) to (ER)). In section 2.3, we discuss the prior distributions on **B** and **D**.

2.1 Homogeneity restrictions

We impose homogeneity restrictions on all structural contemporaneous and lag parameters in **A** and **B** that correspond to foreign variables. We assume that these coefficients are homogeneous across foreign countries (but varying across variables and lags). As an example, let us consider the relation between domestic aggregate demand and foreign output gaps. A standard open economy model does usually not differentiate between different foreign countries but combines a single elasticity, say β^{c,y^*}, with an aggregate measure of output gaps from the rest of the world. For the latter, we use the (trade-weighted) sum of foreign output gaps, $y^*_{ct} = \sum_{c^* \neq c} w_{cc^*} y_{c^*t}$. That is, our aggregate demand curve contains the term

$$\beta^{c,y^*} y^*_{ct} = \beta^{c,y^*} \sum_{c^* \neq c} w_{cc^*} y_{c^*t} = \sum_{c^* \neq c} \underbrace{(w_{cc^*} \beta^{c,y^*})}_{= \beta^{c,y_{c^*}}} y_{c^*t}. \tag{3}$$

In our baseline model, weights w_{cc^*} are the average of BIS bilateral trade weights used to calculate real effective exchange rates for a narrow "basket" of 27 economies.[6]

The assumption $\beta^{c,y_{c^*}} = w_{cc^*} \beta^{c,y^*}$ allows us to identify a single parameter β^{c,y^*}, distributed to each foreign country c^* according to its weight, instead of $C - 1$ separate foreign coefficients $\beta^{c,y_{c^*}}$. Thus, homogeneity assumptions effectively remove the curse of dimensionality arising from the panel dimension of the model. In our baseline model and without taking exclusion restrictions into account, it reduces the number of free parameters in every row of **A** from $n = 28$ to $2n_c = 8$ and the number of the lagged structural coefficients **B** from $k = 114$ to $2n_c p + 2 = 34$ coefficients.

While our homogeneity restrictions are comparable to those imposed in GVAR models, there are important differences. First, we argue that homogeneity restrictions are more convincingly put on the structural than the reduced form of the VAR. The two alternatives

[6] Bilateral weights are reported for three-year periods and are quite stable over time. We take the average over all three-year blocks in our sample period, and normalize bilateral weights in our restricted sample of seven countries to sum up to 1.

are not equivalent, since restrictions on $\mathbf{A}$ as described in Online Appendix C do not imply the same homogeneity restrictions on the inverse of $\mathbf{A}$. However, the latter would be a prerequisite for restrictions to hold on the reduced form as well. That is, while our structural model fulfills the homogeneity restrictions, the reduced-form counterpart does not. Second, the covariance matrix in GVAR models only captures the relations among one country and the weighted sum of all other countries. Thus, GVAR models commonly rely on generalized impulse response functions since identifying all structural shocks is infeasible (see, e.g., Dees et al., 2007; Feldkircher and Huber, 2016; Crespo Cuaresma et al., 2019). Opposed to this, we can investigate dynamic responses of each country's variable to all domestic and foreign shocks easily because we identify the parameters of the full model $\mathbf{A}, \mathbf{B}$. Third, conditionally on $\mathbf{A}$ the estimation of the remainder of the structural model can be done equation-by-equation due to the orthogonality assumption on the structural shocks, reducing the computational burden further.

Notwithstanding their advantages, the homogeneity restrictions in their chosen form may be a poor approximation of reality. As an extreme alternative, one could avoid restrictions altogether as long as the panel dimension of the model does not grow too large. To counter the increase in identification uncertainty, this would necessitate the use of much more informative priors on $\mathbf{A}$. Such an approach would be critical because – especially for countries other than the US – we usually do not have enough exogenous information to pin down elasticities precisely. For $\mathbf{B}$, the researcher could reduce the number of lags in order to deal with the larger estimation uncertainty arising from the lower degrees of freedom, with obvious disadvantages.[7]

Beyond this extreme alternative, one might have three concerns related to the use of time-constant trade weights as the aggregator of choice. First, global supply chains vary over time, with some countries rising to global importance while others become less central. We think that this argument is less of a concern in our sample which does not encompass China. Second, trade weights may be suboptimal if cross-border capital flows rather than trade are related to foreign interest rates (Feldkircher and Huber, 2016). Third, the US

[7]Alternative ways to guarantee the feasibility of the estimation in *reduced from* PVAR models are restrictions assuming homogeneity or no dependence across countries (Canova and Ciccarelli, 2013; Breitung, 2015), penalization estimators or Bayesian shrinkage priors specific to the nature of panel data (Koop and Korobilis, 2016; Korobilis, 2016; Billio et al., 2019; Camehl, 2022), or aggregating parameters in factor structures (Canova and Ciccarelli, 2004, 2009; Koop and Korobilis, 2019). While those techniques facilitate the estimation, identification of the structural model still remains as a challenge.

may have a special role, both as an originator and transmitter of international shocks – an argument underlying the often used block-exogeneity assumption in the literature. We deviate from trade weights along the lines of the second and third concern in robustness checks.

2.2 Informative priors on structural contemporaneous parameters

Theoretical models offer convincing arguments that the structural equations are (contemporaneously) independent of some of the variables in our model. For example, the semi-elasticity of aggregate supply to domestic interest rates is usually assumed to be zero. Hence, we set some contemporaneous relations in the structural model in equations (AD) to (ER) to zero. We can interpret the zero restrictions as prior knowledge imposed by very informative prior distributions with zero variance and which are not updated by the data. The exclusion restrictions are uncontroversial, as they apply mostly to the role of foreign variables, or to domestic interest rates in the aggregate supply curve.

Despite the exclusion of many foreign terms, the country-specific blocks are not fully isolated from foreign contemporaneous developments (such as the US block under block-exogeneity assumptions) because of the potential role of foreign output gaps for aggregate demand and the exchange rate equation. Moreover, we allow monetary policy to react to movements in the nominal exchange rate. We show further below that results are robust if we deviate from some of the exclusion restrictions, notably on foreign terms in the supply equation and monetary policy rule.

In order to derive prior distributions on the remaining structural contemporaneous coefficients (the coefficients explicitly introduced in equations (AS) to (ER)), we mainly build on the New-Keynesian small open economy model of Lubik and Schorfheide (2007). This model is an extension of a closed economy model (Lubik and Schorfheide, 2004) that is used for similar purposes in Baumeister and Hamilton (2018). Based on theoretical insights, we formulate prior beliefs in the form of informative prior distributions and – in some cases – sign restrictions on the remaining structural parameters in **A**. Online Appendix B documents in detail the derivation of structural contemporaneous parameters.

Our prior beliefs on $(\alpha^{c,\pi}, \alpha^{c,\sigma}, \beta^{c,r}, \beta^{c,\pi}, \beta^{c,\sigma}, \beta^{c,y^*}, \psi^{c,\pi}, \psi^{c,y}, \psi^{c,\sigma}, \theta^{c,y}, \theta^{c,\sigma^*}, \rho^c)_{c=1}^{C}$ are sum-

Table 1: Prior on contemporaneous parameters for country c

parameter	distribution	prior mode	prior scale	restrictions
$\alpha^{c,\pi}$	$t(3)$	2	0.4	≥ 0
$\alpha^{c,\sigma}$	$t(3)$	-0.5	0.4	
$\beta^{c,r}$	$t(3)$	-1	0.4	≤ 0
$\beta^{c,\pi}$	$t(3)$	0.75	0.4	
$\beta^{c,\sigma}$	$t(3)$	0.2	0.4	
β^{c,y^*}	$t(3)$	0.5	0.4	
$\psi^{c,\pi}$	$t(3)$	1.5	0.4	≥ 0
$\psi^{c,y}$	$t(3)$	0.5	0.4	≥ 0
$\psi^{c,\sigma}$	$t(3)$	0	0.4	
$\theta^{c,y}$	$t(3)$	1	0.4	
ρ^{c}	*Beta*(2.6,2.6)	0.5	0.2	$0 \leq \rho^c \leq 1$
θ^{c,σ^*}	*Gen-Beta*(2.6,2.6,-1,0)	-0.5	0.2	$-1 \leq \theta^{c,\sigma^*} \leq 0$

marized in Table 1. We report distribution class, mode and scale, and potentially additional sign restrictions for a generic country c since we do not differentiate priors across countries. As in Baumeister and Hamilton (2018), we set a t-distribution with three degrees of freedom (and scale of 0.4) for the majority of parameters, allowing for larger tails compared to a normal distribution. We follow Baumeister and Hamilton (2018) in the prior specifications for the parameters of their closed-economy model, $(\alpha^{c,\pi}, \beta^{c,r}, \beta^{c,\pi}, \psi^{c,\pi}, \psi^{c,y}, \rho^c)_{c=1}^{C}$. We refer the reader to their paper for a detailed discussion on these prior beliefs. Notable among the coefficients is ρ^c, which captures potential interest rate smoothing by monetary authorities.

In an open economy, supply depends on exchange rate changes, measured by $\alpha^{c,\sigma}$. The parameter relates to the import share, intertemporal substitution elasticity, discount factor and the slope coefficient of the Phillips curve of the theoretical model.[8] We follow Lubik and Schorfheide (2007) and assume an import share centered around 20%. A larger share, as for example in Monacelli (2005), would be more suitable for small open economies, while an import share of zero reduces the model to the closed economy version. To avoid singularities in the theoretical model, intertemporal substitution elasticity are commonly restricted to be larger than zero and smaller than one (Lubik and Schorfheide, 2007; Justiniano and Preston, 2010). We follow Lubik and Schorfheide (2007) and settle for a mean of 0.5. As in Baumeister and Hamilton (2018) we assume a discount factor of zero and a slope coefficient

[8]See Online Appendix B for the exact relations of the coefficients in the SVAR model and the theoretical model.

of 0.25. Similarly, Lubik and Schorfheide (2004) set a prior for the slope coefficients allowing for a wide rage between 0 and 1. We substitute expectations with autoregressive processes of order one with an autoregressive parameter equal to 0.75. Based on these values we set the prior mode for $\alpha^{c,\sigma}$ to -0.5. Implicit in the formulation above are restrictions that foreign terms do not contemporaneously influence aggregate supply. We slacken this assumption in robustness checks and find no changes to results.

In the open economy IS curve, we set a positive prior mode of 0.5 for β^{c,y^*}. We follow standard open economy models where the Marshall-Learner condition implies a positive impact of foreign output on domestic output. We do not impose an extra sign restriction to acknowledge that the theoretical evidence on short-run dependencies is not conclusive (Krugman and Baldwin, 1987). Next, we assume a prior mode of 0.2 for $\beta^{c,\sigma}$, i.e. a slightly positive dependence of aggregate demand on competitiveness.

The monetary policy authority can set interest rates according to a generalized Taylor rule in line with the specification in Lubik and Schorfheide (2007), Adolfson et al. (2007), and Justiniano and Preston (2010). Since we include shadow rates in our model, we assume that we can model monetary policy behavior via such an interest rate rule for both conventional and unconventional actions. In an robustness analysis, we challenge this assumption by re-estimating our model based on data until 2007:Q3 only. Our main findings hold in the sub-sample, Figures F.37–F.39, supporting the view that an interest rate rule based on shadow rates captures conventional and unconventional policies. We apply the negative of $\psi^{c,\sigma}$ to foreign inflation to incorporate our prior belief that central banks (if at all) are influenced in their policy by nominal rather than real exchange rate fluctuations. Moreover, we set the prior mode of $\psi^{c,\sigma}$ to zero (as in Adolfson et al., 2007), as the countries in our sample are characterized by flexible exchange rate regimes. Justiniano and Preston (2010), instead, use a Gamma distribution with mean 0.25 as prior belief in a model including Australia, Canada, and New Zealand.

Equation (ER) determines exchange rates as a function of contemporaneous domestic and foreign output as well as foreign exchange rates. Given that purchasing power parity holds, σ_{ct} directly relates to the terms of trade. As in Lubik and Schorfheide (2007), we assume that the difference in domestic and foreign demand growth determines the terms of trades endogenously. That is, the terms of trades determine the relative market-clearing import and export prices at which growth in domestic and foreign demand balances out. We thus apply $\theta^{c,y}$ to the difference between domestic and foreign output gaps. A prior mode of one is based on the relation of import shares and intertemporal substitution elasticity.

By definition, real effective exchange rates are highly correlated, as a relative appreciation of one country implies a relative depreciation of (most of) its trading partners. If unaccounted for, this feature can lead to estimation difficulties and implausible estimates (Lubik and Schorfheide, 2007). To deal with this issue, we allow for contemporaneous relations of domestic to foreign exchange rates through θ^{c,σ^*}, for which we formulate a generalized Beta distribution between the boundaries -1 and 0.

Following Baumeister and Hamilton (2018), we impose prior beliefs on two types of impact responses to economic shocks. For these priors, we use asymmetric t-distributions with location parameter μ, scale parameter σ, degrees of freedom ν, and shape parameter λ, the latter one controlling the degree of asymmetry. First, we assume for every country that the output response to a contractionary monetary policy shock is smaller than the interest rate response ($\mu = -0.3; \sigma = 0.5; \nu = 3; \lambda = -2$). Second, we assume a positive output response to aggregate supply shocks ($\mu = 0.3; \sigma = 0.5; \nu = 3; \lambda = 2$). Finally, we impose a fairly uninformative prior distribution on $\det(\mathbf{A})$ to avoid sign changing of the determinate ($\mu = 2; \sigma = 40; \nu = 3; \lambda = 10$).

2.3 Remaining priors, and posterior inference

The remaining prior distributions on $\mathbf{B}$ and $\mathbf{D}$ are standard choices in the literature. We assume orthogonality of the structural shocks, which implies a diagonal matrix $\mathbf{D}$. For the variances of individual structural shocks (i.e., conditional on $\mathbf{A}$), we follow Baumeister and Hamilton (2015, 2018) and use inverse-gamma prior distributions.

The prior distribution of $\mathbf{B}$ is conditional on $\mathbf{A}$ and $\mathbf{D}$ (Baumeister and Hamilton, 2018). We combine a Minnesota prior – an assumption that each data series follows an AR(1)-process with coefficient 0.75, with prior confidence increasing with the lag length – with the prior structure of the model imposed by $\mathbf{A}$ and $\mathbf{D}$. We use fairly uninformative priors on the constant terms and deterministic trend using a tightness of 100. Also in line with Baumeister and Hamilton (2018), we account for interest rate smoothing in the monetary policy equation by adding an additional prior on the lagged coefficients of the monetary policy equation of country c. This prior is a multivariate normal distribution with mean zero with the exception of the coefficient on $r_{c,t-1}$, for which we set a mean of ρ^c, and variance matrix $0.1 I_k$.

We generate draws from the posterior distributions based on the Metropolis-Hastings algorithm of Baumeister and Hamilton (2015). For details, we refer to their paper and Online Appendix D, which gives a sketch of the algorithm, standard convergence statistics and the full posterior distributions. After a pre-sampling used to obtain suitable starting values and search directions for draws of $\mathbf{A}$, we run the algorithm on 100 parallel chains. From each chain, we obtain 300,000 draws and keep every 100th draw after a burn-in of 200,000 draws.

3 Results

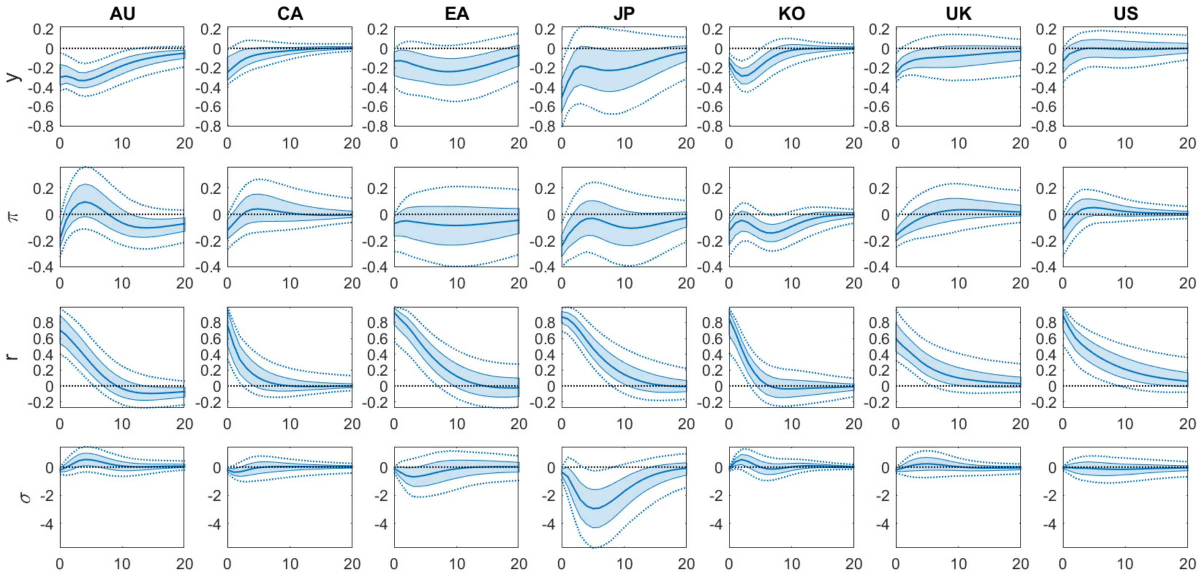
3.1 Domestic responses

Figure 1 shows median impulse responses of country-specific variables in rows to country-specific one unit contractionary monetary policy shocks (in columns) for 20 quarters together with the 68% and 95% posterior credibility sets (shaded area and dotted lines, respectively). Unexpected tightening in monetary policy raises interest rates, lowers output gaps and inflation domestically. Exchange rates react differently across countries, for some exchange rates increase after a few quarters (AU, CA, KP, UK), while for others they decrease (EA, JP) or barely move (US).

The domestic interest rate reaction peaks on impact, but is smaller than the initial shock size. This attenuating macroeconomic feedback effect, which arises from the direct endogenous response of monetary policy to lower output gaps and inflation caused by the contractionary monetary policy shock, is particularly strong for the UK. Incidentally, this is the country where monetary policy smooths interest rates very little and endogenously reacts very strongly to output gaps and inflation, see Figure E.4. After impact, the posterior credibility bands associated with the impulse responses do not contain zero for several quarters, in part because of the significant degree of interest rate smoothing in all countries, as indicated by the posterior distributions of ρ^c.

Domestic supply shocks lead to an expansion in output, a sharp drop in inflation, and a

Figure 1: Impulse responses of domestic variables to country-specific monetary policy shocks



NOTES: The solid lines in the figure show median impulse responses of country-specific variables (in rows) to country-specific monetary policy shocks (in columns) over 20 quarters. The shaded areas (dotted lines) show the 68% (95%) posterior credibility sets. The shocks have size of one unit (i.e., one percentage point).

delayed increase in exchange rates and cause initially a small decrease or no effect in interest rates, see diagonal elements in Figures E.10 to E.13. For most economies a re-bouncing effect is visible after a few quarters. Demand shocks increase domestic interest rates on impact with the peak effect for the countries within one year, see diagonal elements in Figures E.14 to E.17. The shocks raise output gap and inflation. Exchange rates decrease with a trough after on average around a year. The associated posterior credibility sets for the domestic responses exclude zero for several quarters.

3.2 International effects of monetary policy shocks

Our model allows (contemporaneously) for international interest rate transmissions via exchange rates and via aggregate demand (see also Jones et al., 2022).[9] In case of exchange rate targeting, foreign central banks should increase their interest rates in reaction to a contractionary domestic monetary policy shock in an effort to keep their exchange rates stable. Thus, we might expect a positive comovement of international interest rates after

[9]In robustness checks, we allow for additional transmission channels via additional structural contemporaneous parameters, and via additional variables. We discuss the corresponding results further below.

monetary policy shocks, especially for countries in which the exchange rate coefficient in the monetary policy rule, $\psi^{c,\sigma}$, is positive. Transmission channels via aggregate demand, on the other hand, would imply a negative comovement, at least in a textbook model: as contractionary monetary policy shocks lower domestic output gaps, net foreign exports and therefore foreign aggregate demand falls. This "beggar-thy-neighbour"-effect should result in endogenous expansionary monetary policy by the foreign central bank.

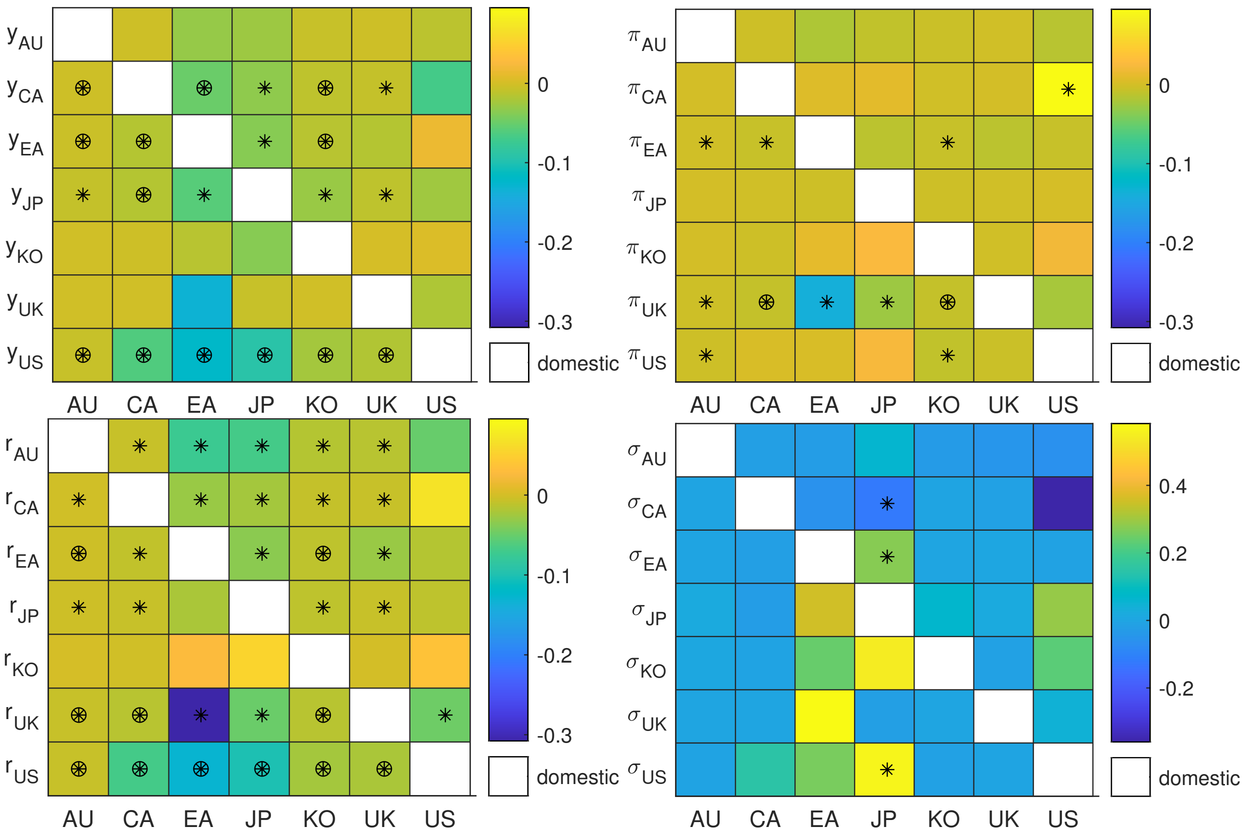
Just like theoretical predictions, the empirical findings on signs of monetary policy spillovers in the literature are somewhat mixed. The majority of papers, focusing on different sets of (advanced) economies and different shock origins, finds mostly negative comovements (Feldkircher and Huber, 2016; Dedola et al., 2017; Crespo Cuaresma et al., 2019). However, especially with respect to US monetary policy shocks, there exists conflicting evidence in the sense that there could be either no comovement (Liu et al., 2022; Degasperi et al., 2021) or even positive comovement (Rogers et al., 2018; Ha, 2021; Ilzetzki and Jin, 2021; De Santis and Zimic, 2022), especially when the "receiving" country is a small open economy with strong trade connections to the US like Canada.

Our results fit to the above literature. In short, we find – if anything – mostly negative international comovement of interest rates after monetary policy shocks, with the exception of US shocks. Moreover, the strength of spillovers depends on the size of the origin economy and the strength of trade relations. Figure 2 shows in one subplot for each of the four macroeconomic variables the median impulse responses of country-specific variables (in rows) to country-specific contractionary monetary policy shocks (in columns) at horizon four. We add an asterisk (with circle) if zero is not contained in the 68% (95%) posterior credibility sets, respectively. The color scales are aligned for the response of output gap, inflation, and interest rates but not for exchange rate responses.[10]

International transmissions of monetary policy shocks to foreign interest rates (the lower left subplot in Figure 2) are modest in size, as indicated by the dominance of olive colored cells. Three regularities stand out. First, shocks from economically large currency areas (US, EA, JP) or those with strong trade relations (for example, CA to US) lead to stronger reactions. Monetary policy shocks from smaller countries or those with weak trade relations result in virtually no international interest rate response. Second and most importantly, the median responses of foreign interest rates at horizon four are in many cases negative.

[10]Figures E.6 to E.9 in the Appendix show the full impulse responses over all horizons.

Figure 2: Impulse responses at horizon four of foreign variables to country-specific monetary policy shocks

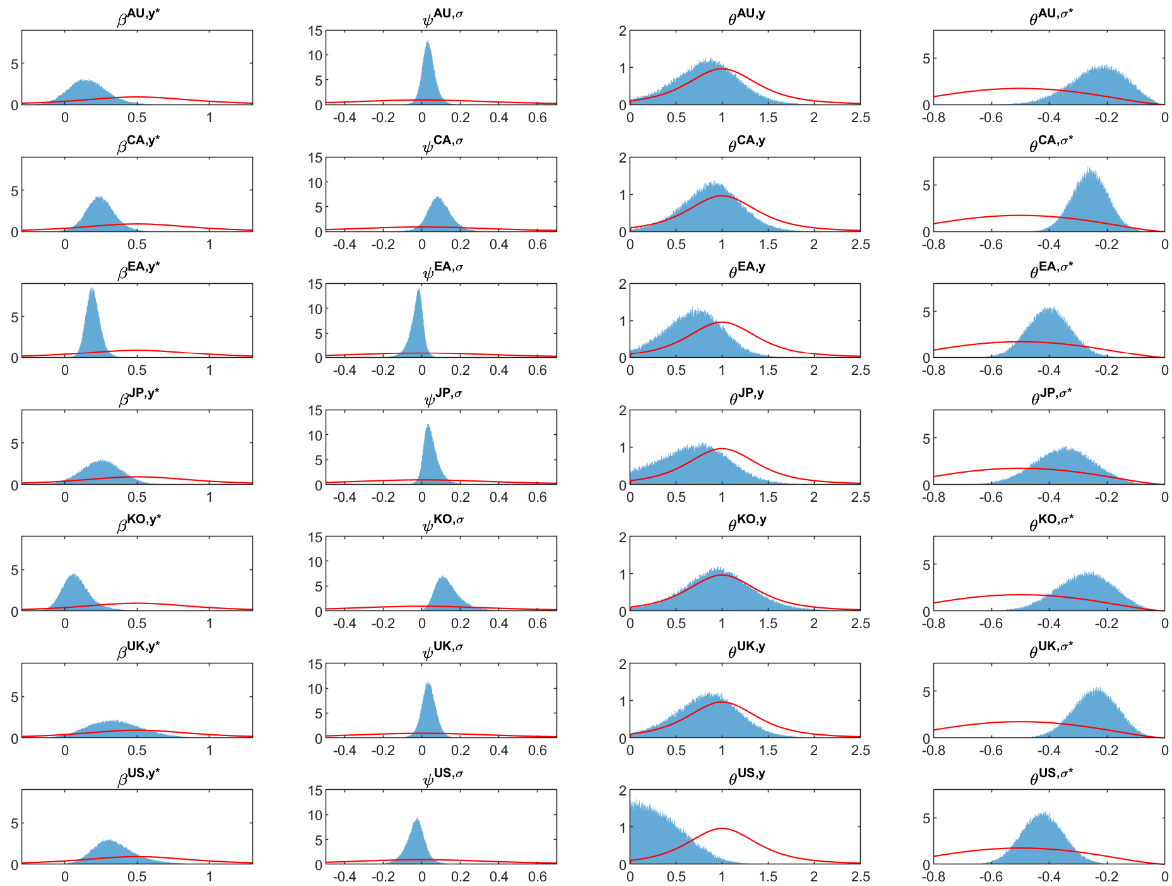


NOTES: Heat map of median impulse responses of country-specific variables (in rows) to country-specific monetary policy shocks (in columns) at horizon four. The asterisk (asterisk with circle) indicates that zero is not contained in the 68% (95%) posterior credibility sets. The color scales for the response of output gap, inflation, and interest rates are aligned.

Third, the majority of 68% credibility sets do not contain zero, even when responses are quantitatively close to zero.

These regularities can be jointly explained by the strength and duration of domestic output gap reactions together with real spillovers across borders via foreign aggregate demand. The two subplots in the first row of Figure 2 display a negative response of foreign output and inflation in reaction to a contractionary domestic monetary policy shock for the majority of country pairs (see also Georgiadis, 2015; Feldkircher and Huber, 2016; Dedola et al., 2017; Crespo Cuaresma et al., 2019). The transmission seems to work mostly via output gaps rather than inflation, as the posterior credibility sets of the former usually exclude zero, while they mostly include zero for the latter. Moreover, the lower right subplot of the figure shows that foreign real effective exchange rates react weakly and the 68% credibility sets contain zero. That is, the cross-border transmission of monetary policy shocks is dominated by the aggregate demand channel rather than the exchange rate

Figure 3: Selection of structural contemporaneous parameters in AD, MP, and EX equations



NOTES: The solid red lines in the Figure show the prior distribution of the structural contemporaneous parameters (columns: parameters; rows: countries). The histograms show posterior distributions.

transmission channel.

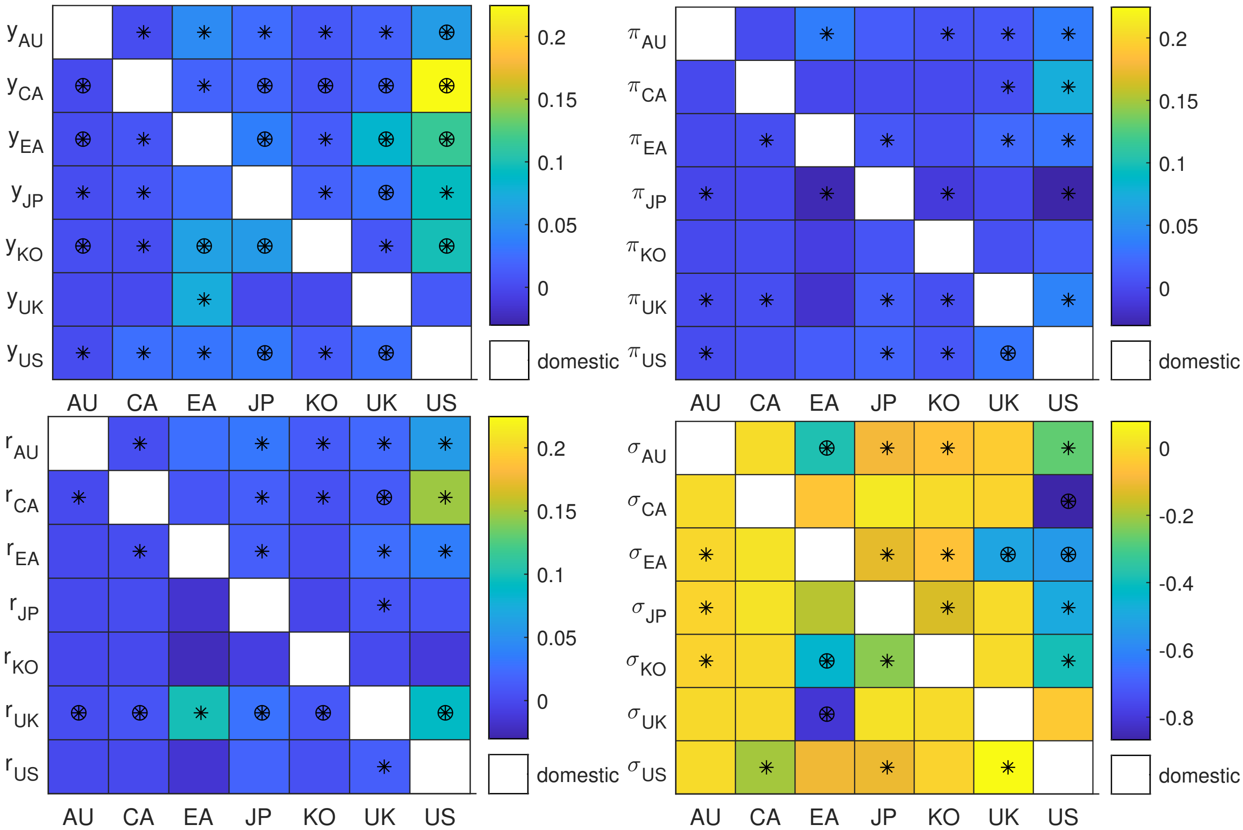
There are two notable exceptions. First, foreign interest rate, for the most part, react only weakly to US monetary policy shocks, with zero included in nearly all credibility sets. The reason lies in the particularly short-lived reaction of US output to contractionary monetary policy shocks, which has already died out at horizon four. As for the positive Canadian interest rate response, we see that this is matched by an increase in inflation after a contractionary US monetary policy shock. Second, Korean interest rates do not react to any foreign interest rate shock. The fact that the same holds for the Korean output gap and inflation as well points towards the possibility that the aggregate demand channel may be absent for Korea.

To investigate these findings further, we turn to the posterior distributions of the elasticities of foreign variables (blue histograms in Figure 3) together with their prior distributions (solid red lines). The aggregate demand channel predicts a positive effect of higher foreign output gaps on domestic inflation. The identified coefficient β^{c,y^*}, first column in Figure 3, mirrors this prediction: it is overwhelmingly positive for nearly all countries. Moreover, the spillover is much smaller than one-to-one, as posterior distributions are always smaller than one. Korea is the exception with a particularly weak aggregate demand channel, with a lower median coefficient than the other countries and only 80% of the posterior mass being positive. The exchange rate channel is captured by the endogenous reaction of domestic monetary policy to (nominal) exchange rate fluctuations, $\psi^{c,\sigma}$, second column in Figure 3. This coefficient is essentially zero for the EA and US, but has more than 80% positive posterior mass for all other countries in our sample. Nonetheless, we find the exchange rate channel to be comparably unimportant because of the weak reaction of real effective exchange rates to monetary policy shocks (last row of Figure 1 and lower right subplot of Figure 2). As the positive coefficient $\theta^{c,y}$ describes the elasticity of real effective exchange rates with respect to output gap differences, an increase in domestic (foreign) output gaps triggers a simultaneous increase (decrease) in real effective exchange rates, i.e. higher (lower) domestic competitiveness. This effect is mostly lower than one to one as the posterior distributions, especially in the US, shift towards zero relative to the prior, third column in Figure 3. The contemporaneous relations between real effective exchange rates, captured by the coefficients θ^{c,σ^*}, fourth column in Figure 3, show some degree of negative dependency.

While a dominant transmitter role of the US is stressed in the literature (such as Miranda-Agrippino and Rey, 2020; Rey, 2016; De Santis and Zimic, 2022), our results give a more diverse picture, as they highlight the global effects of EA shocks and within regional clusters as well. Albeit modest, these reactions to foreign monetary policy shocks provide evidence against block-exogeneity assumptions in VAR models. Hence, our findings relate to studies reporting similarities in spillover effects and mutual reactions caused by unconventional monetary policy shocks of the Fed and ECB (such as Curcuru et al., 2018; Miranda-Agrippino and Nenova, 2022; Jarociński, 2022). However, it should be noted that a generalization of our results towards other currency areas might not be possible. Especially for the case of emerging economies, one would need to pay extra attention to the role of exchange rate arrangements and capital controls.

3.3 International effects of supply and demand shocks

Figure 4: Impulse responses at horizon four of foreign variables to country-specific supply shocks



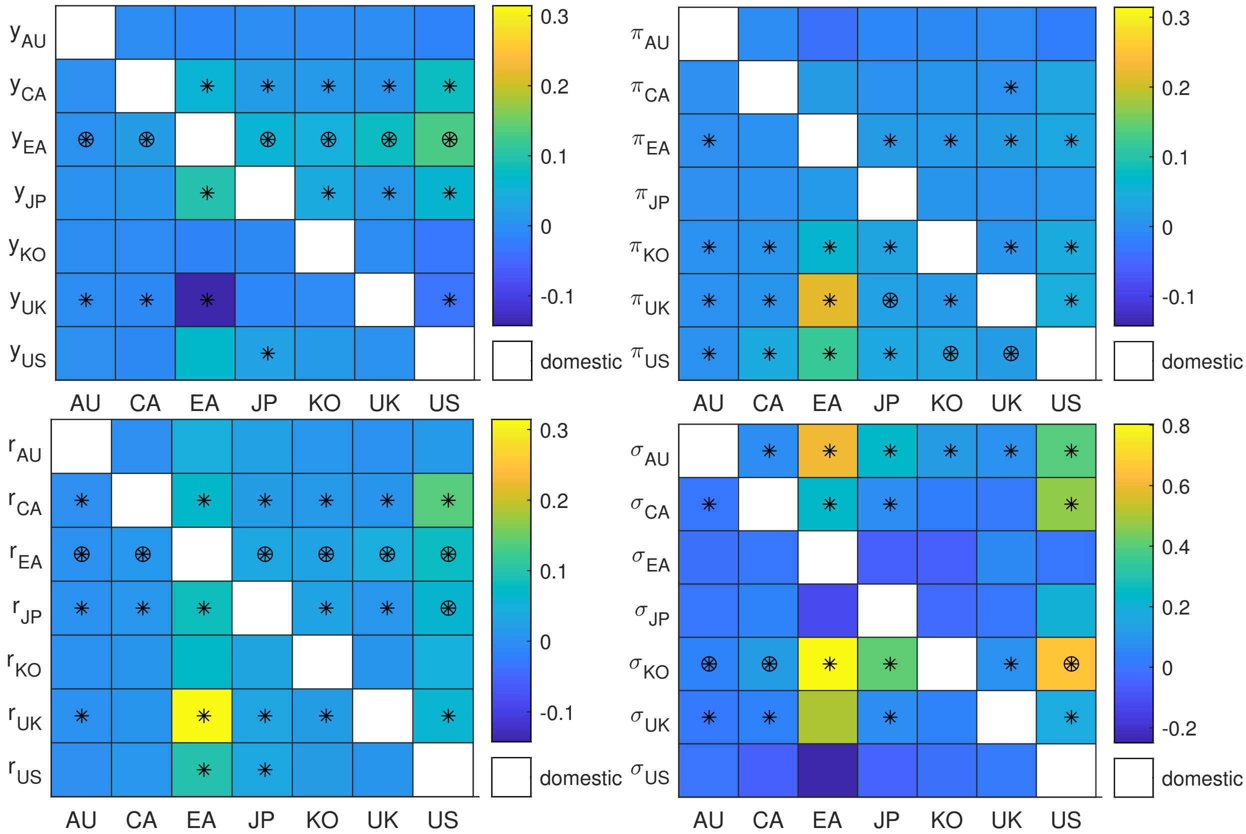
NOTES: Heat map of median impulse responses of country-specific variables (in rows) to country-specific supply shocks (in columns) at horizon four. The asterisk (asterisk with circle) show the 68% (95%) posterior credibility sets. The color scales for the response of output gap, inflation, and interest rates are aligned.

Next, we discuss the international transmission of shocks to aggregate supply and demand. Figures 4 and 5 show the median foreign responses to supply and demand shocks, respectively, at horizon four.[11] The color scales, showing the size of the response, are aligned for the response of output gap, inflation, and interest rates but not for exchange rate responses. In both cases and for many shock-country pairs, we observe that foreign output gaps expand, albeit not to the same degree as domestic output. Inflation and interest rates also go up.

These findings are consistent with those reported by Feldkircher and Huber (2016) for US shocks. As with monetary policy shocks, they can be explained by the aggregate demand channel. The increase of domestic output in reaction to the two shocks shifts

[11]The full impulse response functions, including the domestic responses, is reported in Figures E.10 to E.17 in Online Appendix E.

Figure 5: Impulse responses at horizon four of foreign variables to country-specific demand shocks

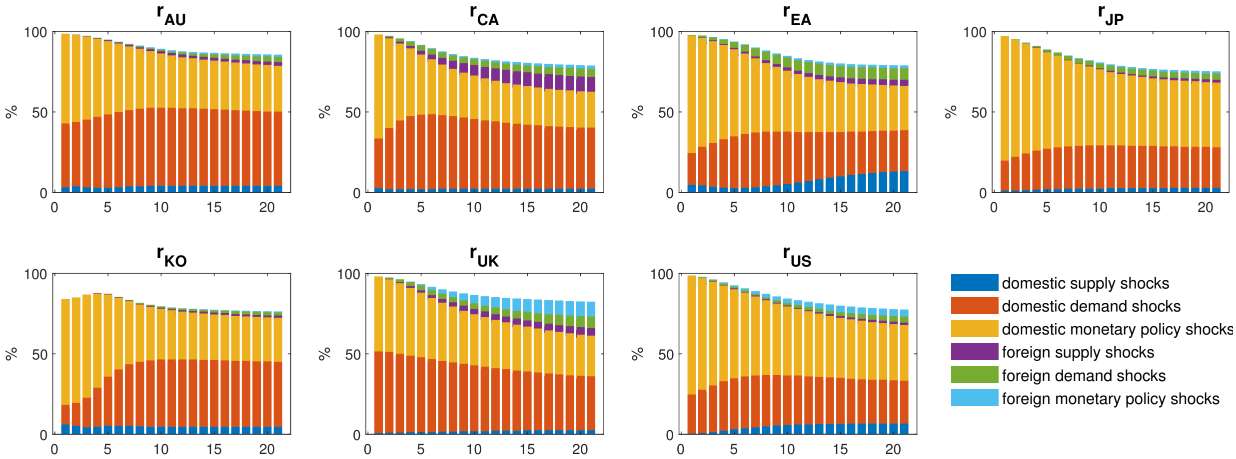


NOTES: Heat map of median impulse responses of country-specific variables (in rows) to country-specific demand shocks (in columns) at horizon four. The asterisk (asterisk with circle) show the 68% (95%) posterior credibility sets. The color scales for the response of output gap, inflation, and interest rates are aligned.

foreign aggregate demand outwards, resulting in the observed economic expansion and increase in inflation. The outward shift in aggregate demand is muted by the endogenous foreign monetary policy reaction (the increase in interest rates). For many shock-country pairs, the 68% or even 95% credibility sets of these responses, denoted by an asterisk and asterisk with circle, do not contain zero. Especially the credibility of the output response of the Euro Area is high. With respect to the relative strength of responses across shocks (the colors in the Figures), we see the importance of country sizes and trade weights. US and Euro Area shocks in general create the strongest responses. Moreover, the Euro Area economy is, for example, affected stronger by shocks from the United Kingdom than by Japanese shocks.

Last, we observe that foreign real effective exchange rate growth is negatively affected by supply shocks and positively by demand shocks. As before, shocks from the US and EA usually create stronger reactions. One interesting result is that demand shocks only

Figure 6: Forecast error variance decomposition of interest rates



NOTES: The figure shows the forecast error variance decomposition of country-specific interest rates (in subplots) to domestic and aggregated foreign supply, demand, and monetary policy shocks over 20 quarters.

cause credible responses for smaller countries (AU, CA, KO and UK). The main reason for these findings might be competitiveness. After four quarters, domestic real effective exchange rate growth is positive after a supply shock, indicating a relative increase in competitiveness, and negative after a demand shock. Correspondingly, foreign countries need to loose (relative) competitiveness after a supply shock, and vice versa after a demand shock.

3.4 Decomposition of variables to domestic and foreign shocks

Next, we evaluate the importance of foreign shocks relative to domestic shocks in explaining output gaps, inflation, interest rates, and growth in exchange rates for each country. Figure 6 shows the forecast error variance decomposition for interest rates to domestic and aggregated foreign supply, demand, and monetary policy shocks over 20 quarters. The decompositions of the remaining variables are in Figures E.18 to E.20.

We find that variation in interest rates can be mostly explained by domestic monetary policy and demand shocks. The contribution of monetary policy shocks decreases over time, while demand shocks become more important. This result is consistent with previous findings by De Santis and Zimic (2022) and Baumeister and Hamilton (2018), although the latter find a higher contribution of demand shocks than we do. This can partially be explained by the sample (monetary policy shocks become less important once we restrict our estimation to end with the financial crisis), and by the fact that they only estimate the model for the US without accounting for international transmission channels. Domestic supply shocks play a minor role, picking up importance for interest rates of EA and US over the horizons. In total, domestic shocks explain – for all countries and horizons, between 60% and 100% of the variation in interest rates.

The relative importance of domestic vs. foreign shocks is similar for output and inflation (see Figures E.18 and E.19), although the relative importance of domestic supply, demand and monetary policy shocks differs across variables and countries. As extreme examples, we find that supply shocks are most relevant for UK output gap and inflation. In the Euro Area, supply shocks explain variations in inflation and demand shocks around 50% of the long-run variation in output gaps.

Overall, the three foreign shocks have increasing contributions over the horizons, albeit to a different degree across countries. They seem to matter mostly for Canada, the EA and UK. Beyond these differences, the relative contribution of different foreign shocks is diverse for interest rates across countries. They vary with foreign supply shocks (in Canada), with foreign demand shocks (EA) and with foreign monetary policy shocks (UK). For output gaps, the relative contributions are more homogeneous: foreign monetary policy shocks are irrelevant, while supply and demand shocks are roughly equally relevant. Inflation seems to be somewhat affected by foreign shocks only in Canada and UK. As for exchange rates, we observe that the variation, especially at short horizons, is nearly completely driven by exchange rate shocks (which are not depicted explicitly).

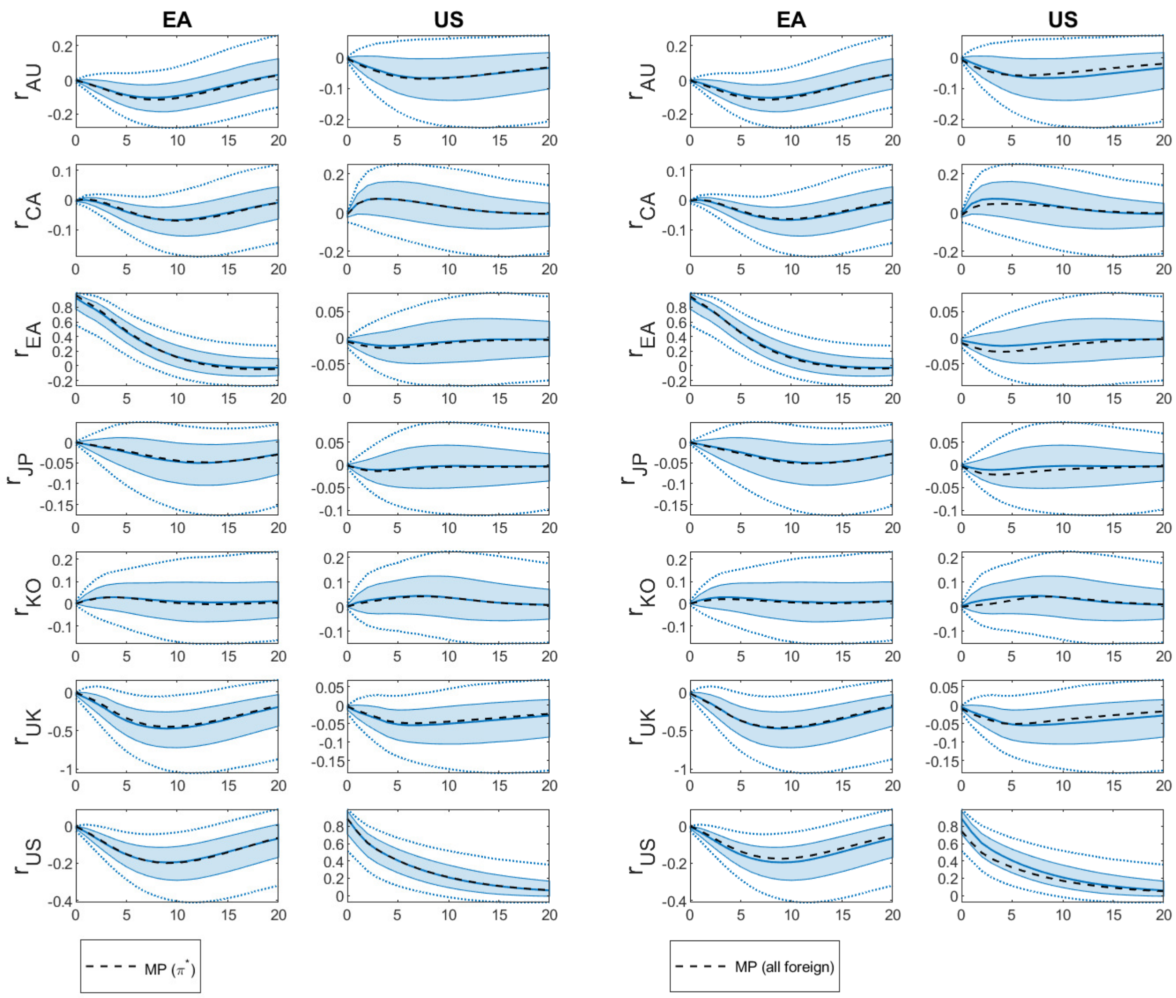
4 Further analysis

4.1 Transmission channels of monetary policy shocks

In the previous section, we argued that monetary policy shocks in our sample are mostly transmitted via aggregate demand channels. Additionally, we allow for an exchange rate channel, which we find to be rather unimportant. However, instead of the aggregate demand channel, domestic monetary policy could directly react to changes in foreign output and inflation. We check this in two robustness checks where we extend the monetary policy rules to contain coefficients on (a) foreign inflation only and (b) foreign output gaps and foreign inflation. Figure 7 plots interest rate responses to EA and US monetary policy shocks for model (a) and (b) in dashed lines along with the posterior credibility sets of the baseline model. We focus on EA and US shocks because we previously found that these two source countries generated the largest foreign responses. Generalizing the monetary policy equation does not alter the impulse responses notably. The reason for this is that the posterior distributions of the additional contemporaneous coefficients are firmly centered around zero.

Figure 7: Impulse responses of interest rates to EA/US monetary policy shocks under alternative monetary policy rules

(a) MP including π^* **(b)** MP including π^* and y^*



NOTES: Figure shows country-specific impulse responses of interest rates to monetary policy shocks in the EA and US over 20 quarters. The posterior mean responses (black lines) of the different models are plotted together with the 68% (shaded areas) and 95% (blue dotted lines) credibility set of our baseline model. The shocks have size of one unit.

The two channels described above should be visible in the responses of trade-related variables. To analyze whether this is the case, we augment our model by total trade, exports or imports, respectively.[12] Figure 8, panel (a), shows the impulse response functions of total trade (% of GDP) in reaction to contractionary monetary policy shocks in the Euro Area and US. We observe that the median responses are often negative (albeit with credibility sets including zero). Exports and imports show very similar responses. This supports the aggregate demand channel, whereby trade should decrease when domestic and foreign output gaps go down. Under the exchange rate channel, however, we would expect a more nuanced reaction where domestic imports increase and exports fall after an inflow of capital causes exchange rate appreciations.

In addition to the channels described above, monetary policy shocks – in particular with the start of unconventional measures – can transmit internationally via financial channels, such as the wealth and credit channel (for a description of different financial channels see, for example, Glick and Leduc, 2012; Bauer and Neely, 2014; Neely, 2015; Fratzscher et al., 2018). The central bank's actions can alter asset prices by stimulating or dampening the demand for assets. Changes in asset prices impact the wealth of households and companies leading to adjusted spending (wealth channel). The monetary authority actions can affect the availability of credit in the market which in turn alters spending and investments (credit channel).

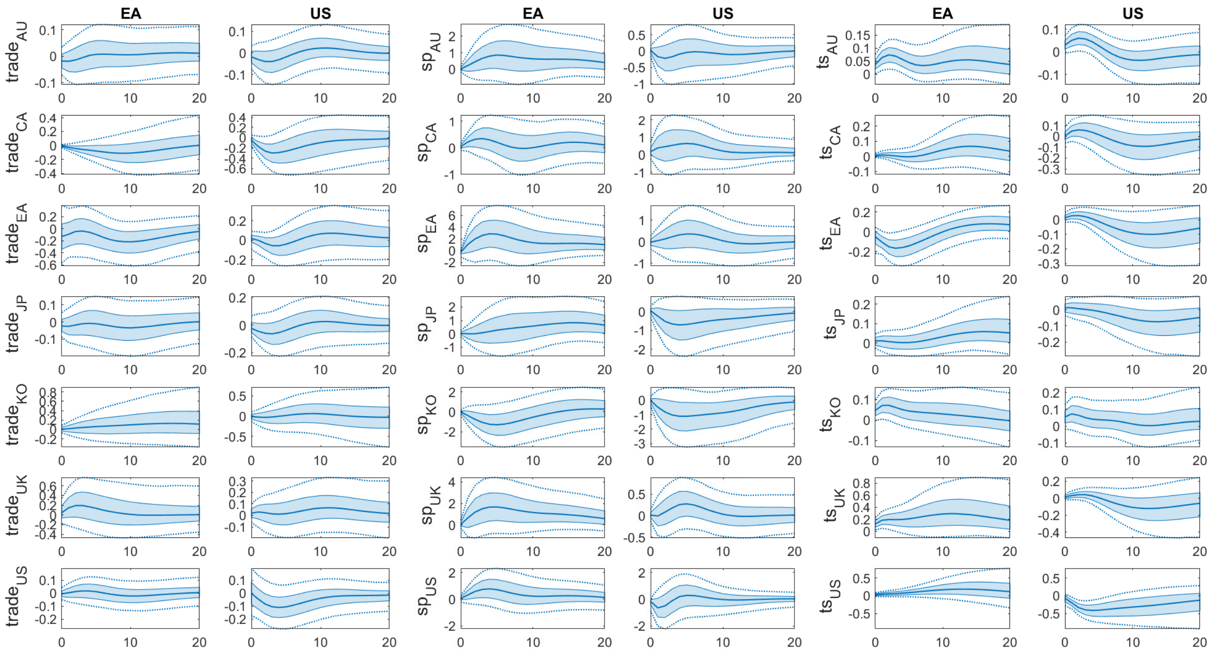
Similar to the trade extensions above, we augment our model by growth in stock prices (wealth channel) or term spreads (credit channel), measured as 5-year yields minus shadow interest rates.[13] Figure 8 shows in (b) and (c) the impulse responses of growth in stock prices and term spreads to EA and US monetary policy shocks. In line with the moderate interest rate responses, international reactions of growth in stock prices and term spreads are modest. Hence, we do not find clear evidence for an international wealth and credit

[12]This implies a large number of additional contemporaneous relations in the existing equations, and a set of additional (residual) equations. We include the additional trade variable (domestic and foreign) contemporaneously in all baseline equations. In the additional residual equations, we allow all variables to contribute contemporaneously to the development of the trade variable, with the exception of foreign exchange rates and foreign trade variables. We set relatively wide priors on the additional contemporaneous parameters, Student t priors with mode zero and scale one, indicating that we lack clear theoretical arguments for these parameters. We also add prior beliefs that contractionary monetary policy raises domestic interest rates, lowers domestic output gaps and inflation, and decreases competitiveness, by setting asymmetric t distributions with $\mu = 0$, $\sigma = 1$, $v = 3$, and $\lambda = 20$ on the impact effects of monetary policy shocks.

[13]In addition to the restrictions in the trade extensions, we also restrict the impact response of the stock prices or term spreads to a domestic monetary policy shock to be negative.

Figure 8: Impulse responses of additional variables to EA/US monetary policy shocks

(a) total trade (b) stock prices (c) term spreads



NOTES: Figure shows country-specific impulse responses of (a) total trade, (b) stock prices and (c) term spreads to monetary policy shocks in the EA and US over 20 quarters, obtained from models extended to five variables per country. The posterior mean responses (solid lines) are plotted together with the 68% (shaded areas) and 95% (blue dotted lines) credibility sets. The shocks have size of one unit.

channel. Domestically, our results show that the credit channel is active since domestic term spreads of the EA and US respond to the respective monetary policy shocks. Overall, including the channel variables does not alter the impulse responses of interest rates substantially (Figures F.21 to F.26).

4.2 Additional robustness analysis

We run a series of robustness checks to validate our results. In the following, we discuss the use of alternative choices for prior distributions on contemporaneous structural coefficients, alternative sample specifications and alternative foreign country weights. Visual presentations of these results are given together with the baseline results, similar to Figure 7.

4.2.1 Alternative prior distributions

Similarly to the robustness checks with additional coefficients in the monetary policy rule, we check robustness to alternative specifications of aggregate supply and the exchange rate equations.

To that end, we (1) add foreign output gaps in the supply equation, restricting the coefficient to be positive. Thus, we allow that domestic output contemporaneously reacts to changes in foreign output hypothesizing that an increase in foreign output raises domestic output, for example due to common technology shocks. Second, we (2) adjust the exchange rate equation such that interest rate differential enter. Hence, we model directly the uncovered interest rates parity such that differences in interest rates between the countries should equalize relative changes in exchange rates, potentially strengthening the exchange rate channel for the transmission of monetary policy shocks. As before, the responses of interest rates to monetary policy shocks of the EA and US, shown in Figure F.27 for model (1) in dashed lines and (2) in dashed-dotted lines, align with those of the baseline model.

4.2.2 Alternative samples

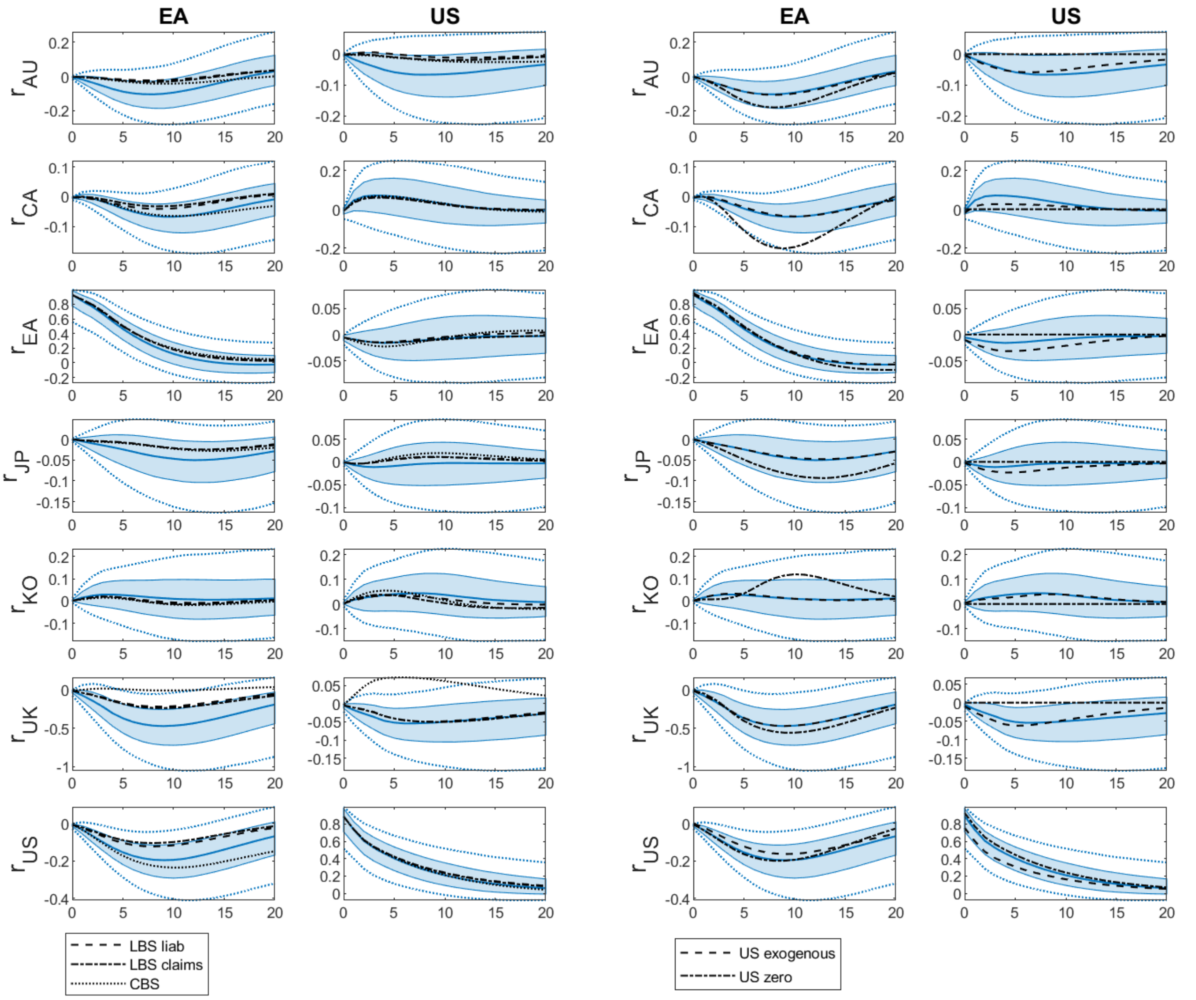
Furthermore, we show that our main findings hold for two sub-samples: (1) data start in 1999:Q1, figures F.34–F.36, and (2) data end in 2007:Q2, figures F.37–F.39. We use the first to assess whether including the EA as an aggregate before the introduction of the Euro alters the results. We find that results are broadly robust, with responses to EA shocks being a bit dampened. The second sub-sample allows us to study whether the switch to unconventional monetary policy measures impacts our main findings. In this case, the response to EA shocks even increases, while other reactions remain very similar.

4.2.3 Alternative weights

In general, financial linkages across countries might be of higher importance than trade linkages for international spillovers in interest rates. Therefore, we consider three alternative banking weights for foreign interest rates while the remaining foreign variables enter trade-weighted (Feldkircher and Huber, 2016, use similar weights). We construct weights based on BIS data using (1) total liabilities of the locational banking statistics, (2) total claims of the locational banking statistics, and (3) the consolidated banking statistics. The locational banking statistics include outstanding claims and liabilities of worldwide operating banks located in a reporting country. The consolidated banking statistics measure claims of banking groups whose headquarter are located in the reporting country.

Figure 9: Impulse responses of interest rates to EA/US monetary policy shocks using alternative weights

(a) Bank weights (b) US weights



NOTES: Figure shows country-specific impulse responses of interest rates to monetary policy shocks in the EA and US over 20 quarters. The posterior mean responses (black lines) of the different models are plotted together with the 68% (shaded areas) and 95% (blue dotted lines) credibility set of our baseline model. The shocks have size of one unit.

The results are summarized in Figure 9 which shows the impulse responses of interest rates to monetary policy shocks in the EA and US. The first two columns give the posterior mean responses of the models with different weights; (1) dashed, (2) dashed-dotted and (3) dotted lines together with the credibility set of our baseline model. While the majority of the responses do not change considerably, UK interest rates react notably less to a

EA monetary policy shock, and positively to a US shock (the latter one when using the consolidated banking statistics). The connections measured via banking weights between the UK and EA are around half the size than the bilateral trade link. Still, nearly all responses are within the credibility set of the baseline model and differences between the models using the three banking weights are negligible.

Next, we assess the role of the US as a catalyst for international spillovers. To that end, we compare our baseline model, first, with a specification in which the US is treated mainly as a transmitter but not receiver country (*US exogenous*). We remove the contemporaneous foreign terms in the US equations except those in the exchange rate equation. Autoregressive terms still include foreign terms as well. This comparison comes close to the exogeneity assumption for the United States, which has been used in a large part of the literature to investigate the spillover of US shocks. Second, we shut off transmissions of US shocks (*US zero*). Thus, spillovers cannot happen from or via the US.[14] To model this, we set the weights on US variables in the equations of all other countries to zero. This robustness check allows us to check the sensitivity of Euro Area transmissions to the inclusion of the US in the model.

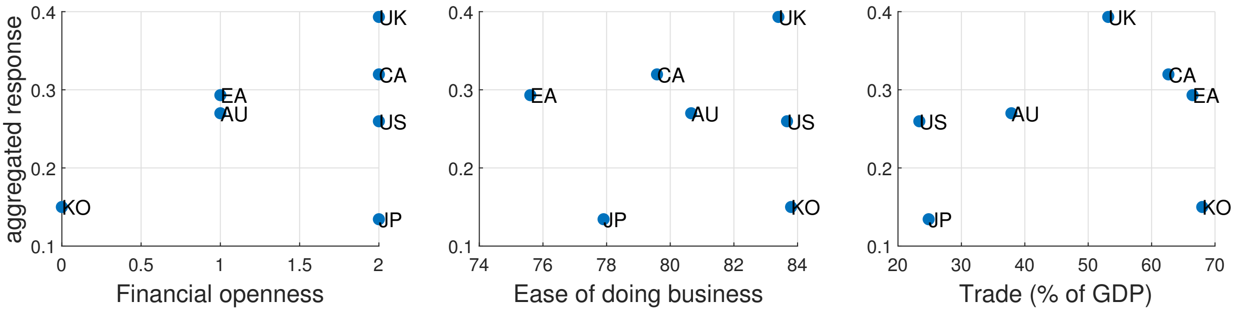
The last two columns of Figure 9 plot the median impulse responses of the two models; given by the dashed (*US exogenous*) and dashed-dotted lines (*US zero*). Overall, the reactions do not change considerably compared to the baseline model. Shutting off transmission of US shocks increases partly the effects of EA shocks, for example to AU and CA (i.e. countries that are fairly strongly connected to the US).

4.3 Country characteristics and the sensitivity of interest rates to foreign shocks

We further take a look whether stronger reaction to foreign shocks can be associated with specific country characteristics. In order to do that, we calculate absolute responses of interest rates at horizon four to standardized foreign monetary policy, supply, and demand shocks and sum those responses to capture the total strength of transmissions from other countries. Figure 10 shows country characteristics on the horizontal axis together with our

[14]The US still influence the identification of the model because of the orthogonality condition on all shocks.

Figure 10: Country characteristics and response of interest rates to foreign shocks at horizon four



NOTES: The figure shows country characteristics and summed absolute responses of interest rates to foreign supply, demand, and monetary policy shocks at horizon four.

transmission strength measure on the vertical axis. All country characteristics are averages over the years 1980 to 2019. The first subplot relates financial openness to the aggregated responses. We take the Chinn and Ito (2006)-index, where higher numbers indicate more open economies, as a measure for financial openness.[15] More financially open economies could be more affected by spillovers. We find a small positive correlation of 0.41. That is in line with Georgiadis (2016) and Degasperi et al. (2021) who report a small positive effect of financial openness of advanced economies on spillovers from US monetary policy shocks. In the second subplot, we consider the ease of doing business index from the World Bank. It measures the regulatory conditions to conduct a business in a country between 0 (worst) and 100 (best). We find no clear correlation (correlation of 0.08).

The third subplot includes total trade as percent of gross domestic product (World Bank data). We see these as measures of trade openness of a country. Trade integration could increase exposure to spillovers when foreign demand is an essential part in aggregated demand of an economy. On the other hand, trade integration can dampen spillovers when an expenditure-switching effect is dominant (Georgiadis, 2016). Dedola et al. (2017) report no evidence of a systematic relationship between trade openness to the US and reaction to US monetary policy shocks for their set of 36 countries. Aggregating over all shocks, we find instead a small positive relation between trade openness and stronger responses to foreign shocks (correlation of 0.26).

[15] To obtain a value for the EA we calculated the average of 18 member states (Luxembourg is not included).

5 Conclusions

In this paper, we find evidence for international spillovers in interest rates driven by country-specific demand, supply and monetary policy shocks. To draw inference on structural equations, we extend the approach of Baumeister and Hamilton (2018) to a multi-country framework. The approach has two advantages over the literature. First, by using informative priors on structural contemporaneous coefficients we achieve full identification of the model equations, allowing us to differentiate between aggregate supply, aggregate demand, a monetary policy rule and an exchange-rate equation. Second, we are the first to show how homogeneity restrictions on *structural* foreign coefficients can remove the curse of dimensionality which usually arises from the extending the panel dimension. Our restrictions – for the same variable, equation and lag, foreign coefficients are equal up to a constant weight – are already applied in some of the literature relying on estimation of the reduced form model. However, we argue that the restrictions are much more convincing if imposed on the structural form of the model.

We show that most international interest rate transmissions arise through an aggregate demand channel. Thus, an exogenous shift in the behavior of monetary authorities leads to a decline in foreign interest rates. Demand and supply shocks cause considerable, and positively correlated, responses in interest rates internationally. The strongest international reactions are found for shocks originating in the EA and US, which are the two largest trading partners for the majority of the countries in our sample. Compared to domestic shocks, foreign shocks are overall of lower importance. We show that our main findings hold when changing prior beliefs on contemporaneous relations and the model set-up.

Acknowledgements

We thank Christiane Baumeister, Peter Tillmann, Roland Winkler and participants of the IWH Macroeconomics Reading Group for useful comments and suggestions. We are also grateful to Helmut Lütkepohl and Marek Jarocinski for their constructive remarks on an earlier version of this paper.

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ISSN 2194-2188

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