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Martin Feldkircher, Thomas Gruber, Florian Huber

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International effects of a compression of euro area yield curves

MARTIN FELDKIRCHER[1*], THOMAS GRUBER[1] and FLORIAN HUBER[2]

[1]*Oesterreichische Nationalbank (OeNB)*
[2]*Salzburg Centre of European Union Studies, University of Salzburg*

In this paper, we use a Bayesian global vector autoregressive model to analyze the macroeconomic effects of a flattening of euro area yield curves. Our findings indicate positive effects on real activity and prices, both within the euro area as well as in neighboring economies. Spillovers transmit through an exchange rate channel and a broad financial channel. We complement our analysis by conducting a portfolio optimization exercise. Our results show that multi-step-ahead forecasts conditional on the euro area yield curve shock improve Sharpe ratios relative to other investment strategies.

JEL CODES: C30, E52, F41, E32

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*Corresponding author: Martin Feldkircher, Foreign Research Division, Oesterreichische Nationalbank (OeNB). *Email*: martin.feldkircher@oenb.at. Any views expressed in this paper represent those of the authors only and do not necessarily coincide with those of the Oesterreichische Nationalbank or the Eurosystem. We would like to thank Geert Bekaert, Alessandro Galesi, Jennifer Gredler, Axel Jochem, Milan Nedeljković, Michael Pfarrhofer, Mária Širaňová, Stefan Voigt, participants of the 14th emerging markets workshop, Banco de España, Madrid, participants of the Jean Monnet workshop on Financial globalization and its spillovers, Cologne and Maastricht, participants of the VfS Annual Conference 2017, Vienna and two anonymous referees for helpful comments. The authors gratefully acknowledge financial support from the Austrian Science Fund (FWF): ZK 35) . This paper, subject to minor editorial changes, is forthcoming in the Journal of Banking & Finance.

1. INTRODUCTION

As a reaction to the global financial crisis, major central banks have considerably lowered their policy rates to stimulate economic growth and consumer price inflation. Since the room for conventional monetary policy quickly eroded, central banks increasingly adopted non-standard measures such as *Quantitative Easing* (QE). The literature that examines the macroeconomic effects of QE typically focuses on single-country models and domestic effects (see, among many others, Kapetanios et al., 2012; Baumeister and Benati, 2013; Gambacorta et al., 2014). Asset purchases by the central bank are supposed to bring inflation and inflation expectations in line with the central banks' target as well as stimulate economic growth. There are several channels through which QE can affect the real economy (see among others, Joyce et al., 2012; Fawley and Neely, 2013) and the empirical literature hitherto has reached a consensus that QE was effective in lowering yields (Gagnon et al., 2011; D'Amico and King, 2013; Krishnamurthy and Vissing-Jorgensen, 2011) as well as supporting economic growth (Lenza et al., 2010; Altavilla et al., 2016). How these effects are distributed across euro area countries, whether there are spillovers to neighboring economies, and whether this information can be used for asset management are the main questions we aim to answer in this study.

More specifically, this paper provides new evidence on the domestic and international effects of a decline in euro area term spreads. We focus on the term spread as a measure of QE since in a zero-interest rate environment, the most direct consequence of the central bank buying long-term securities is a flattening of the yield curve. This effect might be amplified by the duration channel (Krishnamurthy and Vissing-Jorgensen, 2011). That central banks indeed use unconventional measures to reduce interest rate spreads such as the term spread has been argued in Blinder (2012) and empirically demonstrated for the euro area by Altavilla et al. (2016) and Ambler and Rumler (2019).[1] Investigating macroeconomic effects of changes in the yield curve separates our work from others, who mostly focus on spillovers from unconventional monetary policy defined in a broader sense (see Horváth and Voslářová, 2017; Babecká Kucharčuková et al., 2016; Bluwstein and Canova, 2016; Moder, 2019). Also, new approaches to measure monetary policy in a broader sense focus on the yield curve as a vehicle through which monetary policy materializes (Inoue and Rossi, 2018).

The econometric framework adopted is a Bayesian global vector autoregressive (GVAR) model (Crespo Cuaresma et al., 2016; Feldkircher and Huber, 2016). The GVAR is based on estimating separate country-specific models that are then connected by specifying suitable link matrices which proxy economic ties and interactions between the countries. Using the GVAR hence enables us not only to accommodate cross-country heterogeneity of the receiving economies but also within the euro area. That euro area monetary policy causes heterogeneous effects among euro area countries is an important empirical fact that has been highlighted in work by Georgiadis (2015) and Burriel and Galesi (2018). For that purpose, we include data on single euro area countries while the ECB's reaction function is following a Taylor rule (Georgiadis, 2015) based on aggregate data. To identify the term spread shock, we use a novel combination of zero impact and sign restrictions on the cross-section that allows for the potential presence of heterogeneous intra-euro area effects.

Our main results may be summarized as follows. First, the term spread reduction leads to a significant increase in euro area output and consumer prices (see Bluwstein and Canova, 2016; Burriel and Galesi, 2018) as well as an increase in equity and house prices. Second, the euro area term spread shock triggers significant output and price spillovers to neighboring countries. These spillovers are

[1] For recent empirical applications that assess macroeconomic effects of QE by changes in the term spread, see European Central Bank (2017) and Bobeica and Jarociński (2019) for the euro area, and Chen et al. (2016) for the USA. The most comprehensive study that estimates full zero coupon yield curves is Wright (2011), who covers ten industrialized countries.

transmitted through both an exchange rate channel and a broad financial channel. More specifically, we find that local currencies strengthen vis-à-vis the euro, equity prices rise, and long-term rates decline, which spurs private credit growth. That long-term rates in CESEE economies decline, suggests that geographical closeness is an important consideration for international investors (besides other factors such as overall riskiness of assets). Last, we compute multi-step-ahead forecasts based on the identified structural shocks and find that using this information pays off in terms of Sharpe ratios, relative to a naive portfolio that weighs all assets equally and a portfolio that is constructed using unconditional forecasts.

The remainder of the paper is structured as follows. The next section summarizes the data and introduces the econometric framework. Section 3 lays out the strategy to jointly identify euro area term and risk spread shocks, while section 4 discusses the results. Finally, section 5 concludes.

2. DATA OVERVIEW AND ECONOMETRIC FRAMEWORK

In this section we turn to the description of our data set and econometric framework. We employ an extension of the traditional GVAR approach put forward by Pesaran et al. (2004), adopting a flexible stochastic volatility specification. The first subsection describes the global vector autoregressive model with stochastic volatility in fairly general terms. In the second subsection we briefly discuss the Bayesian prior setup adopted.

2.1. *The global vector autoregressive model with stochastic volatility*

The GVAR framework builds on a sequence of $N + 1$ country-specific VAR models that establish a relationship between domestic and international macroeconomic factors. To assess the international effects of recent financial market developments in the euro area, we use data on industrial production (y_{it}, index 2015=100), consumer prices (p_{it}, index 2010=100), short- and long-term interest rates ($i_{si,t}$ and $i_{li,t}$, 3-months and 10-year, respectively), stock prices (eq_{it}, index 2010=100), private credit (pc_{it}, index 2010=100), house prices (hp_{it}, index 2010=100) and the nominal exchange rate vis-à-vis the euro (er_{it}). These data are on a monthly frequency and span the period from 2000m10 to 2016m06. Short- and long-term interest rates are used to construct the term spread (tp_{it}) and the risk spread (rp_{it}), the latter being defined as euro area long-term yields over German government bond yields. All variables are in levels and, with the exception of interest rates, in logarithmic transform.[2]

These data are collected for euro area *core* (Austria, Belgium, Germany, Finland, France, Netherlands and Slovakia)[3] and *periphery* countries (Ireland, Italy, Portugal, Spain, Greece) on the one hand and neighboring countries from the *CESEE* region (Bulgaria, Croatia, the Czech Republic, Hungary, Poland, Romania and Slovenia) and other advanced non-euro area European countries (*other non-EA*, Denmark, Norway, the United Kingdom, Sweden and Switzerland) on the other hand. We further include data for Canada, China, Japan, Russia, Turkey and the USA to control for global factors. That

[2] Data on industrial production, consumer prices and house prices are de-seasonalized. Since industrial production data for Switzerland are only available on a quarterly basis over the time period we consider, we use the Chow Lin time disaggregation method (Chow and Lin, 1971) and the unemployment as a monthly time series to obtain monthly industrial production data.

[3] Due to their relatively small role in the asset purchase program, we exclude the Baltics, Cyprus and Malta. Moreover, Slovakia and Slovenia both adopted the euro over the sample period covered in this study. We assign Slovakia to the euro area core group and Slovenia to the set of CESEE countries, since for the latter data on long-term yields are not available, which prevents calculating spread shocks. Empirical results provided in section 4 are unaffected by the regional mapping of these two countries.

leaves us with a sample of good coverage of the euro area, non-euro area European countries and the G-8 industrialized advanced economies. The countries in our sample account for over 70% of global nominal output (averaged over the years 2010 to 2016) and reflect the most important trading partners of the euro area.

Euro area and non-euro area countries differ by the variables they include. For a typical euro area country i, we have a $k_j(=7)$-dimensional vector

$$\boldsymbol{x}_{it} = (tp_{it}, rp_{it}, y_{it}, p_{it}, eq_{it}, pc_{it}, hp_{it})'. \tag{1}$$

For a typical non-euro area country, the vector of endogenous variables is 8-dimensional and differs along two dimensions. First, we replace the term- and risk spread with the level of the short- and long-term interest rate, respectively. Second, we include the nominal exchange rate of a given country vis-á-vis the euro. For a complete list of variable coverage per country, see Tab. B.1 in the appendix.

We then assume that the dynamics of the k_i endogenous variables in country i are described by the following VARX($p = 2$, $q = 2$) model,

$$\boldsymbol{x}_{it} = \sum_{j=1}^{p=2} \boldsymbol{A}_{ij}\boldsymbol{x}_{it-j} + \sum_{s=0}^{q=2} \boldsymbol{B}_{is}\boldsymbol{x}^*_{it-s} + \boldsymbol{\varepsilon}_{it}, \tag{2}$$

with $\boldsymbol{A}_{ij}$ $(j = 1, \dots, p)$ being $k_i \times k_i$-dimensional coefficient matrices. $\boldsymbol{B}_{is}$, $(s = 0, \dots, q)$ are coefficient matrices of dimension $k_i \times k_i^*$ associated with the weakly-exogenous variables and $\boldsymbol{\varepsilon}_{it}$ is a normally distributed vector error term with zero mean and a time-varying variance-covariance matrix $\boldsymbol{\Sigma}_{it}$.

The weakly exogenous or international variables $\boldsymbol{x}^*_{it}$ serve as main channels for economic and financial spillovers across countries. These are computed as a simple weighted average of the other countries' endogenous variables:

$$\boldsymbol{x}^*_{it} = \sum_{j=0}^{N} w_{ij}\boldsymbol{x}_{jt}, \quad \text{for} \quad i \in \{0, \dots, N\}. \tag{3}$$

Here w_{ij} denotes a set of bilateral weights that reflect economic interactions between countries i and j, normalized to sum up to unity. For the sake of exposition, we assume here that all countries feature the same number of endogenous variables in $\boldsymbol{x}_{jt}$ – an assumption that we are going to relax in the empirical application. As is common in the literature using GVARs, w_{ij} are composed of bilateral trade flows.[4] Recently, other weights based on, e.g., financial flows have been proposed in the literature (see, e.g., Eickmeier and Ng, 2015). However, Feldkircher and Huber (2016) present a sensitivity analysis with respect to the choice of weights in Bayesian GVAR specifications and show that trade weights yield a reasonable model fit.

A convenient feature of the GVAR framework is that variable coverage can vary across countries. We include the respective foreign counterparts for all domestic variables $\boldsymbol{x}_{it}$ – with two exceptions. First, and to control for exchange rate movements in a broader sense, we include trade-weighted exchange rates in euro area countries although no domestic exchange rate (vis-à-vis the euro) exists.

The second exception relates to how we model monetary policy in the euro area. Following Georgiadis (2015) we introduce an "ECB" country model where monetary policy is governed by

[4] More precisely, we use annual data from the World Input Output Database (WIOD), averaged over the period from from 2000 to 2014. The data were retrieved from http://www.wiod.org/home and are available only up until 2014. For a detailed description see Timmer et al. (2015).

a simple Taylor rule. More specifically, the euro area short-term interest rate (i_s^{EA}) is regressed on purchasing power parity (PPP)-weighted averages of output and consumer prices of euro area countries. In a second step, i_s^{EA} enters then into all (also non-euro area) country models as a weakly exogenous variable. In this sense, the treatment of domestic interest rates is not symmetric among euro area countries on the one hand and the rest of the countries included in the analysis on the other hand. Last, we include oil prices ($poil_t$, Brent, in US dollar) as a global control variable. Following the bulk of the literature, oil prices are assumed to be endogenously determined within the US country model (see, e.g., Pesaran et al., 2004).

For the specification of the error term, we deviate from the standard approach that assumes a homoskedastic variance. Following Cogley and Sargent (2005) we can decompose $\boldsymbol{\Sigma}_{it}$ as follows

$$\boldsymbol{\Sigma}_{it} = \boldsymbol{U}_i \boldsymbol{H}_{it} \boldsymbol{U}_i', \tag{4}$$

where $\boldsymbol{U}_i$ is a $k_i \times k_i$-dimensional lower triangular matrix with unit diagonal and off-diagonal elements denoted by $u_{ij,n}$ $(j = 2, \ldots, k_i; n = 1, \ldots, j-1)$ and $\boldsymbol{H}_{it}$ is a diagonal matrix with $\boldsymbol{H}_{it} = \text{diag}(e^{h_{i1,t}}, \ldots, e^{h_{ik_i,t}})$. We assume that the log-volatilities $h_{ij,t}$ follow an AR(1) process,

$$h_{ij,t} = \mu_{ij} + \rho_{ij}(h_{ij,t-1} - \mu_{ij}) + \kappa_{ij,t}. \tag{5}$$

Hereby, we let μ_{ij} denote the (unconditional) mean of the log-volatility, ρ_{ij} is the persistence parameter, while $\kappa_{ij,t}$ denotes a white noise error with variance ς_{ij}^2. We include stochastic volatility in the VAR framework for three reasons: First, the time period under study is rather volatile since many countries experienced a boom period (up until 2006) and a severe bust period (2008/09) followed by a diverse recovery. Hence accounting for time variation might improve the fit of the model. This can be done in general by either letting coefficients in the model drift, or by allowing residual variances to change over time. Several studies (Primiceri, 2005; Sims and Zha, 2006) find rather limited evidence in favor of time-varying parameters but recognize the importance to control for heteroscedasticity.[5]

Second, we emphasize that our specification nests the homoscedastic case. More specifically, if ς_{ij}^2 equals zero, the log-volatility process remains constant at its unconditional mean, implying that heteroscedasticity is effectively ruled out. In our Bayesian setup, we introduce a suitable shrinkage prior to allow the data to speak about the degree of heteroscedasticity over time (see Appendix A for more information). Third, as will be described in Section 3, we focus on a joint shock to multiple countries. This implies that the shock composition matters for the impact estimates of the impulse response functions. Hence, incorrect inference about the underlying volatilities could potentially translate into biased impulse responses.

The sequence of $N+1$ country models can be combined to yield a global VAR model,

$$\boldsymbol{G}\boldsymbol{x}_t = \sum_{n=1}^{p^*} \boldsymbol{F}_n \boldsymbol{x}_{t-n} + \boldsymbol{\varepsilon}_t. \tag{6}$$

Hereby, we let $\boldsymbol{x}_t = (\boldsymbol{x}_{0t}', \ldots, \boldsymbol{x}_{Nt}')'$ denote a $k = \sum_{j=0}^{N} k_j$-dimensional vector that collects all endogenous variables in the system, $\boldsymbol{G}$ is a $k \times k$ matrix of contemporaneous coefficients that are a function of the $\boldsymbol{B}_{i0}$ matrices and the weights in w_{ij} and $p^* = \max(p, q) = 2$. Moreover, $\boldsymbol{F}_n$ are $k \times k$ matrices of autoregressive coefficients that are driven by the weights and the estimates of $\boldsymbol{A}_{ij}$ for all countries and $\boldsymbol{\varepsilon}_t = (\boldsymbol{\varepsilon}_{0t}', \ldots, \boldsymbol{\varepsilon}_{Nt}')'$ is a k-dimensional vector white noise process with a block-diagonal variance-covariance matrix $\boldsymbol{\Sigma}_t = \text{diag}(\boldsymbol{\Sigma}_{0t}, \ldots, \boldsymbol{\Sigma}_{Nt})$.

[5] See Huber (2016) for predictive evidence in favor of stochastic volatility specifications within large dimensional models.

Multiplying with $\boldsymbol{G}^{-1}$ from the left yields the reduced-form GVAR model that closely resembles a standard VAR model with parametric restrictions imposed through the weights w_{ij},

$$\boldsymbol{x}_t = \sum_{n=1}^{p^*} \boldsymbol{\psi}_n \boldsymbol{x}_{t-n} + \boldsymbol{v}_t. \tag{7}$$

The reduced-form VAR coefficients are given by $\boldsymbol{\psi}_n = \boldsymbol{G}^{-1}\boldsymbol{F}_n$ and $\boldsymbol{v}_t = \boldsymbol{G}^{-1}\boldsymbol{\varepsilon}_t$ is a k-dimensional vector of white noise errors with variance given by $\boldsymbol{\Omega}_t = \boldsymbol{G}^{-1}\boldsymbol{\Sigma}_t(\boldsymbol{G}^{-1})'$.

2.2. *Stochastic variable selection in GVAR models*

While the GVAR modeling approach imposes parsimony by restricting the coefficients related to other countries' endogenous variables to be driven by economic weights (see Eq. (3)), the remaining number of parameters in Eq. (2) is still typically higher than the number of available observations. This calls for Bayesian shrinkage priors that effectively deal with this problem by shrinking the parameter space toward some stylized prior model. In this section we outline the main ideas behind the stochastic search variable selection (SSVS) prior (see George and McCulloch, 1993; George et al., 2008). The full prior setup as well as information on the Markov chain Monte Carlo algorithm is given in Appendix A.

To set the stage, we collect all regression coefficients in a K_i-dimensional vector $\boldsymbol{c}_i = \text{vec}\{(\boldsymbol{A}_{i1}, \dots, \boldsymbol{A}_{ip}, \boldsymbol{B}_{i0}, \dots, \boldsymbol{B}_{iq})'\}$, with $K_i = k_i(pk_i + qk_i^*)$. We follow Feldkircher and Huber (2016) and specify a stochastic search variable selection (SSVS) prior in the spirit of George and McCulloch (1993) and George et al. (2008) on each element of $\boldsymbol{c}_i$,

$$c_{ij}|\delta_{ij} \sim \mathcal{N}(0, \tau^2_{ij,0})\delta_{ij} + \sim \mathcal{N}(0, \tau^2_{ij,1})(1-\delta_{ij}) \quad \text{for } j = 1, \dots, K_i. \tag{8}$$

Hereby we assume that the prior on c_{ij} depends on a Bernoulli distributed random variable δ_{ij} that selects the prior variance $\tau^2_{ij,0} \gg \tau^2_{ij,1}$ with $\tau_{ij,1}$ set close to zero.

Thus, if δ_{ij} equals unity, we choose the first Gaussian distribution with mean equal to zero and a rather large variance $\tau^2_{ij,0}$. This case introduces little prior information on c_{ij}, implying that the posterior is strongly driven by the information contained in the likelihood function and little shrinkage is applied to c_{ij}. This effectively translates into a point estimate of c_{ij} near the maximum likelihood estimate. By contrast, if δ_{ij} equals zero, the Gaussian prior adopted features a tiny prior variance, strongly pushing the corresponding posterior distribution of c_{ij} toward zero.[6] For δ_{ij} we use a Bernoulli distributed prior with prior inclusion probability of 0.5 for all i and j. This implies that we view all variables as equally likely to appear in the corresponding country-specific VAR model.

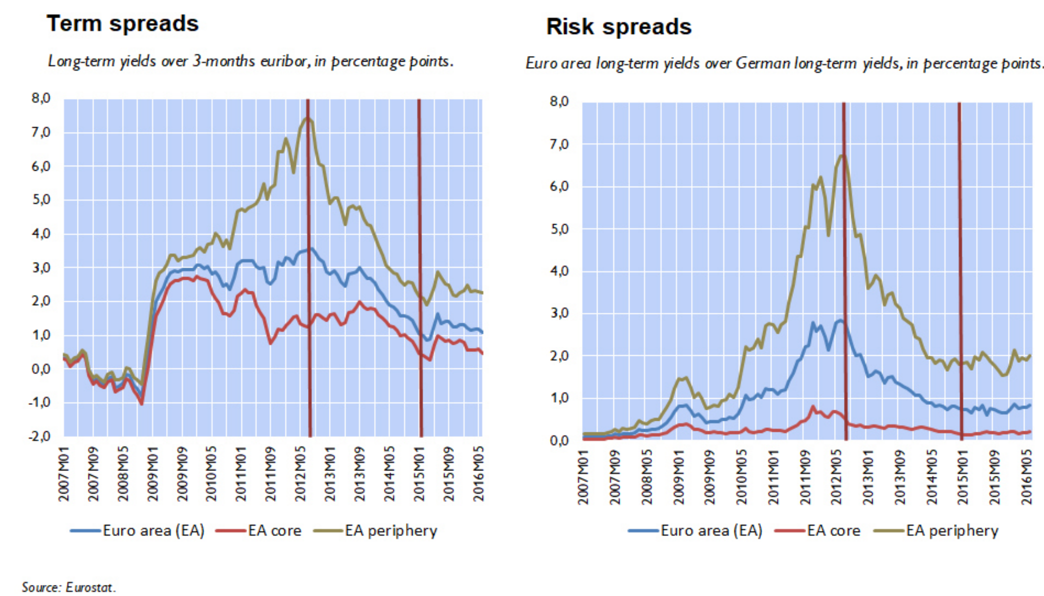
The presence of δ_{ij} implies that the proposed model specification is a hierarchical model, which in turn suggests that conditional on c_{ij}, the corresponding posterior density of δ_{ij} is independent from the data, i.e., $p(\delta_{ij}|c_{ij})$. It is then easy to show that the posterior distribution of δ_{ij} is a Bernoulli distribution with probability that $\delta_{ij} = 1$ given by

$$p\left(\delta_{ij} = 1 \middle| c_{ij}\right) = \frac{\mathcal{N}(c_{ij}|0, \tau^2_{ij,0})}{\mathcal{N}(c_{ij}|0, \tau^2_{ij,0}) + \mathcal{N}(c_{ij}|0, \tau^2_{ij,1})}. \tag{9}$$

We let $\mathcal{N}(c_{ij}|0, \tau^2_{ij,0})$ and $\mathcal{N}(c_{ij}|0, \tau^2_{ij,1})$ denote the density of a Gaussian distribution evaluated at c_{ij}. If c_{ij} is close to zero, the spike mixture component would be more probable, implying that

[6] For both Gaussian components in Eq. (8), the prior mean of the first own lag of a given variable/equation is specified to equal unity to mimic features of the Minnesota prior.

Fig. 1: Evolution of term and risk spreads in the euro area



Notes: The left panel of the plot shows the evolution of term spreads and the right panel that of risk spreads. Risk spreads are defined as the spread of 10-year government bonds over German 10-year government bond yields; euro area (EA) core and periphery countries as defined in Section 2. Purchasing power parities are used to calculate regional aggregates. Vertical bars refer to the "Whatever it takes" speech (July 2012) and the launch of the extended asset purchase program (March 2015).

$p\left(\delta_{ij} = 1|c_{ij}\right)$ is close to zero. By contrast, if c_{ij} is large, the spike component of the prior would be highly improbable, leading to a value of $p\left(\delta_{ij} = 1\middle| c_{ij}\right)$ near one. In what follows, we also apply this prior specification to obtain a parsimonious representation of the variance-covariance matrix within a given country. This allows us to shrink covariance parameters within a given country to zero, if necessary.

3. IDENTIFICATION

Following Baumeister and Benati (2013), we assume that large-scale purchases of longer-term securities result in a compression of the yield curve in the euro area. To illustrate this, Fig. 1, left panel shows the dynamics of term spreads for the euro area aggregate as well as PPP-weighted averages of euro area core and periphery countries.

Term spreads increased notably during the global financial crisis and around 2011 in the wake of the sovereign debt crisis. While the first increase in term spreads has been similar across countries, the second increase had a much stronger impact on periphery countries than on core economies, indicating a divergence of euro area long-term yields.

An important event that significantly impacted term spreads has been the "Whatever it takes" speech of ECB president Mario Draghi, delivered at the Global Investment Conference in London.[7] This speech was aimed at containing a divergence of long-term yields within the euro area. To measure

[7] See Acharya et al. (2017) for an empirical assessment of the macroeconomic effects of the "Whatever it takes" speech.

this, we define a risk spread as the spread between long-term yields in a euro area country over German long term yields (perceived as risk free), depicted on the right-hand side in Fig. 1. In July 2012, risk spreads declined sharply, while they have been rather unaffected during the period when the asset purchase program was launched.

Since the decline in term spreads in July 2012 was unrelated to actual purchases of longer-term securities, it could be argued that focusing solely on the term spread as a vehicle through which asset purchases affect the economy could contaminate the results. In the empirical exercise that follows, we hence complement the term spread analysis by also assessing risk spreads and identify both shocks jointly within a single empirical model. Doing this effectively allows us to purge the dynamic responses to a joint shock to euro area term spreads from a risk spread shock that was purely driven by a single verbal intervention.

Both shocks are empirically identified by imposing a mixture of zero and sign restrictions on the impulse response functions. While a similar identification approach was proposed in Mumtaz and Surico (2009) in an international context and by using a factor augmented VAR model applied to US data, it is novel to the GVAR literature. In a GVAR context, imposing $k(k-1)/2$ identifying restrictions is usually avoided by identifying the structural shocks of interest locally (Dees et al., 2007). By contrast, our identification approach is based on considering the structural form of the global model,

$$\boldsymbol{\Lambda} \boldsymbol{x}_t = \sum_{n=1}^{p^*} \tilde{\boldsymbol{\psi}}_n \boldsymbol{x}_{t-n} + \boldsymbol{u}_t, \tag{10}$$

where $\boldsymbol{\Lambda}$ is a $k \times k$ matrix of coefficients that determines the contemporaneous relationships between the elements in $\boldsymbol{x}_t$ and $\tilde{\boldsymbol{\psi}}_n = \boldsymbol{\Lambda}\boldsymbol{\psi}_n$ is a structural coefficient matrix associated with the nth lag of $\boldsymbol{x}_t$ while $\boldsymbol{u}_t$ is a vector of structural shocks with variance given by $\boldsymbol{H}_t$. In what follows we assume that $\boldsymbol{\Lambda} = \boldsymbol{U}^{-1}\boldsymbol{G}$ with $\boldsymbol{U}^{-1} = \text{diag}(\boldsymbol{U}_0^{-1}, \ldots, \boldsymbol{U}_N^{-1})$ being a block-diagonal matrix.

We introduce zero impact restrictions by appropriately re-ordering the equations in Eq. (10) such that all quantities that are assumed to react sluggishly to euro area financial developments (i.e., real output, inflation, and total credit) *and* do not belong to the euro area are ordered first. Moreover, let $\boldsymbol{x}_t^{\text{RA}}$ denote an r_{RA}-dimensional vector that stacks these variables. The next equation is the one that determines euro area short-term interest rates (i_{st}^{EA}), followed by a block of variables that consists of euro area macroeconomic (i.e., real and financial market) quantities collected in an r_{EA}-dimensional vector $\boldsymbol{x}_t^{\text{EA}}$. The zero impact restriction on i_{st}^{EA} is necessary to disentangle the reduction in term spreads driven by a decrease in long-term yields from that of a conventional monetary policy shock. Regarding the latter, Benati and Goodhart (2008) note that a conventional monetary policy tightening also causes a flattening of the yield curve since long-term yields typically do not increase as much as the policy rate. Finally, we assume that financial market quantities outside the euro area are fast moving and react instantaneously to shocks (exchange rates, short- and long-term interest rates, equity prices, house prices and oil prices). Variables belonging to this last block are collected in an r_{FI}-dimensional vector denoted as $\boldsymbol{x}_t^{\text{FI}}$.

Zero impact restrictions on spillovers and the euro area short-term interest rate are complemented with sign restrictions for the euro area. Conditional on the ordering of variables, we can rewrite Eq. (10) as

$$\boldsymbol{\Lambda}^+ \begin{pmatrix} \boldsymbol{x}_t^{\text{RA}} \\ i_{st}^{\text{EA}} \\ \boldsymbol{x}_t^{\text{EA}} \\ \boldsymbol{x}_t^{\text{FI}} \end{pmatrix} = \sum_{n=1}^{p^*} \tilde{\boldsymbol{\psi}}_n^+ \begin{pmatrix} \boldsymbol{x}_{t-n}^{\text{RA}} \\ i_{st-n}^{EA} \\ \boldsymbol{x}_{t-n}^{\text{EA}} \\ \boldsymbol{x}_{t-n}^{\text{FI}} \end{pmatrix} + \boldsymbol{u}_t^+, \tag{11}$$

whereby $\boldsymbol{\Lambda}^+, \tilde{\boldsymbol{\psi}}_n^+(n = 1, \ldots, p^*)$, and $\boldsymbol{u}_t^+$ are the coefficient matrices and reduced-form shocks with rows reshuffled in consistency with the timing restrictions described. Within the euro area block $\boldsymbol{x}_t^{\mathrm{EA}}$, we introduce a set of sign restrictions outlined in Tab. 1. These restrictions are imposed by using a $k \times k$-dimensional block-diagonal rotation matrix $\boldsymbol{R} = \mathrm{diag}(\boldsymbol{I}_{r_{RA}+1}, \hat{\boldsymbol{R}}, \boldsymbol{I}_{r_{FI}})$ with $\boldsymbol{RR}' = \boldsymbol{I}_k$ and $\hat{\boldsymbol{R}}$ denoting an orthonormal rotation matrix. This rotation matrix implies that we only rotate the shocks of euro area countries to introduce the sign restrictions while using the timing restrictions outlined above for all remaining quantities.[8]

The rotation matrices are obtained following the algorithm proposed in Arias et al. (2015).[9] The zero impact and sign restrictions we use are outlined in Tab. 1, with ↑ / ↓ denoting an increase/decrease, "med"/"all" indicating that restrictions have to hold at least for half of the euro area countries/all euro area countries and 1,4 the time horizon until which the restrictions are binding.

Table 1: Zero impact & sign restrictions.

Country	Variable	Term spread shock	Risk spread shock
ECB model:	i_s^{EA}	0_1	0_1
EA countries:	$tp_{\{EA\backslash DE\}}$	$\downarrow_{all,4}$	–
	tp_{DE}	$\downarrow_4$	0_1
	$rp_{\{EA\}}$	–	$\downarrow_{all,1}$
	$y_{\{EA\}}$	$\uparrow_{med,4}$	–
	$p_{\{EA\}}$	$\uparrow_{med,4}$	–
	$eq_{\{EA\}}$	$\uparrow_{med,4}$	–
Non-EA countries	$y_{\{non-EA\}}$	$0_{all,1}$	$0_{all,1}$
	$p_{\{non-EA\}}$	$0_{all,1}$	$0_{all,1}$
	$pc_{\{non-EA\}}$	$0_{all,1}$	$0_{all,1}$
	$eq_{\{non-EA\}}$	–	–
	$hp_{\{non-EA\}}$	–	–
	$er_{\{non-EA\}}$	–	–
	$i_{s\{non-EA\}}$	–	–
	$i_{l\{non-EA\}}$	–	–

Notes: The restrictions are imposed on impact and the subsequent three periods. The subscript "med" indicates that at least half of the countries in the considered country group have to fulfill the sign restrictions, "all" means that restrictions have to hold simultaneously for all considered countries. The numbers indicate how long restrictions are binding (1 or 4 months). "EA" refers to the euro area (core and periphery) and "non-EA" to non-euro area countries.

The term spread shock is calibrated to yield a simple average decrease of 100bp in the euro area. The cross-country composition of the shock depends on the estimated error volatilities which are tilted toward periphery countries.[10] The compression of the yield curve – as a consequence of quantitative

[8] Alternatively, one could introduce a $k \times k$-dimensional rotation matrix that rotates all shocks in the system. This, however, increases estimation uncertainty and the computational complexity involved.

[9] In principle, the fact that we rely on the Cholesky decomposition of $\boldsymbol{\Omega}_t$ indicates that the corresponding IRFs are not invariant with respect to different orderings of the elements in $\boldsymbol{x}_t^{\mathrm{RA}}$ and $\boldsymbol{x}_{t-n}^{\mathrm{FI}}$. Based on 30 random permutations of the elements in $\boldsymbol{x}_t^{\mathrm{RA}}$ and $\boldsymbol{x}_t^{\mathrm{FI}}$ we show that the ordering has empirically little influence on the results.

[10] More specifically, impact responses lie in the range of -43bp to -57bp for DE, SK, AT, NL, FI, -68bp to -100bp for FR, BE, ES, IT and -143bp to -267bp for IE, PT, and GR.

easing – should increase economic activity (y_t) and prices (p_t), while short-term interest rates in the euro area do not deviate from zero. These restrictions basically follow Baumeister and Benati (2013).

In addition, we include a restriction for equity prices. Empirical evidence for the reaction of stock markets to monetary policy-induced interest rate changes has been provided in a range of contributions (see, among others, Thorbecke, 1997; Rigobon and Sack, 2004; Bernanke and Kuttner, 2005). Kuttner and Mosser (2002) postulate that after a monetary loosening, investors who search for yields adjust their portfolio from (now less profitable) fixed income assets to the equity market (eq_t). Bekaert et al. (2013) more generally find that monetary easing lowers the risk awareness of investors. In our setting, the term spread decrease should trigger a reduction in risk awareness, with investors diversifying into more risky (i.e., equity) markets as a result. The response of private credit remains ambiguous a priori. On the one hand, heightened economic activity coupled with lower long-term yields should boost demand for private credit. On the other hand, a reduction in term spreads typically induces a decrease in net interest rate margins and the incentive for banks to lend money.[11] Consequently we leave the response of private credit (pc_t) unrestricted. In addition, and without further prior knowledge, we do not restrict house price responses (hp_t).

To further sharpen the analysis and to separate the effects of a term spread shock from the induced reactions of the "Whatever it takes" speech, we also identify a risk spread shock that is orthogonal to the term spread shock. In analogy to the term spread shock, the risk spread shock is implemented by an average decrease of euro area risk spreads by 100bp. Following a similar logic as before, long-term interest rates in Germany should be zero on impact to rule out the case where risk spreads decrease due to an increase in German long-term yields. This is implemented by putting a zero restriction on the German term spread, which together with the zero restriction on euro area short-term interest rates, implies a zero restriction on German long-term yields. The zero impact restriction on the German term spread also separates the risk spread from the term spread shock (which assumes a reduction in German term spreads).[12] Further timing restrictions are again imposed on real activity, prices and private credit in non-euro area countries.

Finally, two points of our identification strategy deserve further discussion.[13] First, the sign restrictions are only imposed to hold for all euro area economies jointly for the variable we shock (i.e., term and risk spreads). To account for intra-euro area heterogeneity, the remaining restrictions are binding only for the median of euro area impulse response functions. In a similar framework, Georgiadis (2015) and Burriel and Galesi (2018) put restrictions on simple averages of euro area quantities, which could lead to a situation with one country driving the results. Alternatively, one could place restrictions on PPP-weighted averages of euro area single countries' impulse response functions. This would ensure that restrictions of large countries are fulfilled and hence that the euro area aggregate follows the outlined assumptions. In our application, these modifications of the restrictions do not alter the results qualitatively. Second, since we employ stochastic volatility, we would technically have different impulse response functions for each point in the sample. For the sake of brevity, we show impulse responses based on a shock composition using the sample means

[11] Recently, Wieladek and Pascual (2016) have shown that the credit easing channel plays a role in the transmission of asset purchases to the economy. This suggests that banks lend out money even against the background of decreasing margins. The reason is that interest on reserves is low or negative and banks still have an incentive to lend more, competing with their peers in the pass-through of savings (Wieladek and Pascual, 2016).

[12] We achieve this by introducing an additional zero restriction using the methods outlined in Arias et al. (2015).

[13] Also note that the proposed implementation of zero and sign restrictions differs from the local treatment of sign restriction in the GVAR framework as proposed in Eickmeier and Ng (2015) and applied among others in Fadejeva et al. (2017) and Feldkircher and Huber (2016). See the working paper version of this article for a detailed robustness exercise in this regard.

of estimated volatilities in the term/risk spread equation. We cross-check our results when using volatilities corresponding to the last data point in our sample in subsection 4.3.

4. EMPIRICAL RESULTS

In this section we examine the domestic and international effects of a simultaneous compression of euro area single countries' yield curves, controlling for developments of euro area risk spreads.

4.1. *Effects of a compression of the yield curve*

For the sake of illustration we aggregate country-specific impulse response functions by purchasing power parities into four regions of interest, namely the euro area *core* and *periphery*, neighboring countries from the *CESEE* region and other advanced non-euro European countries, *other non-EA*. The regions are defined as in section 2 and single country responses are available from the authors upon request.

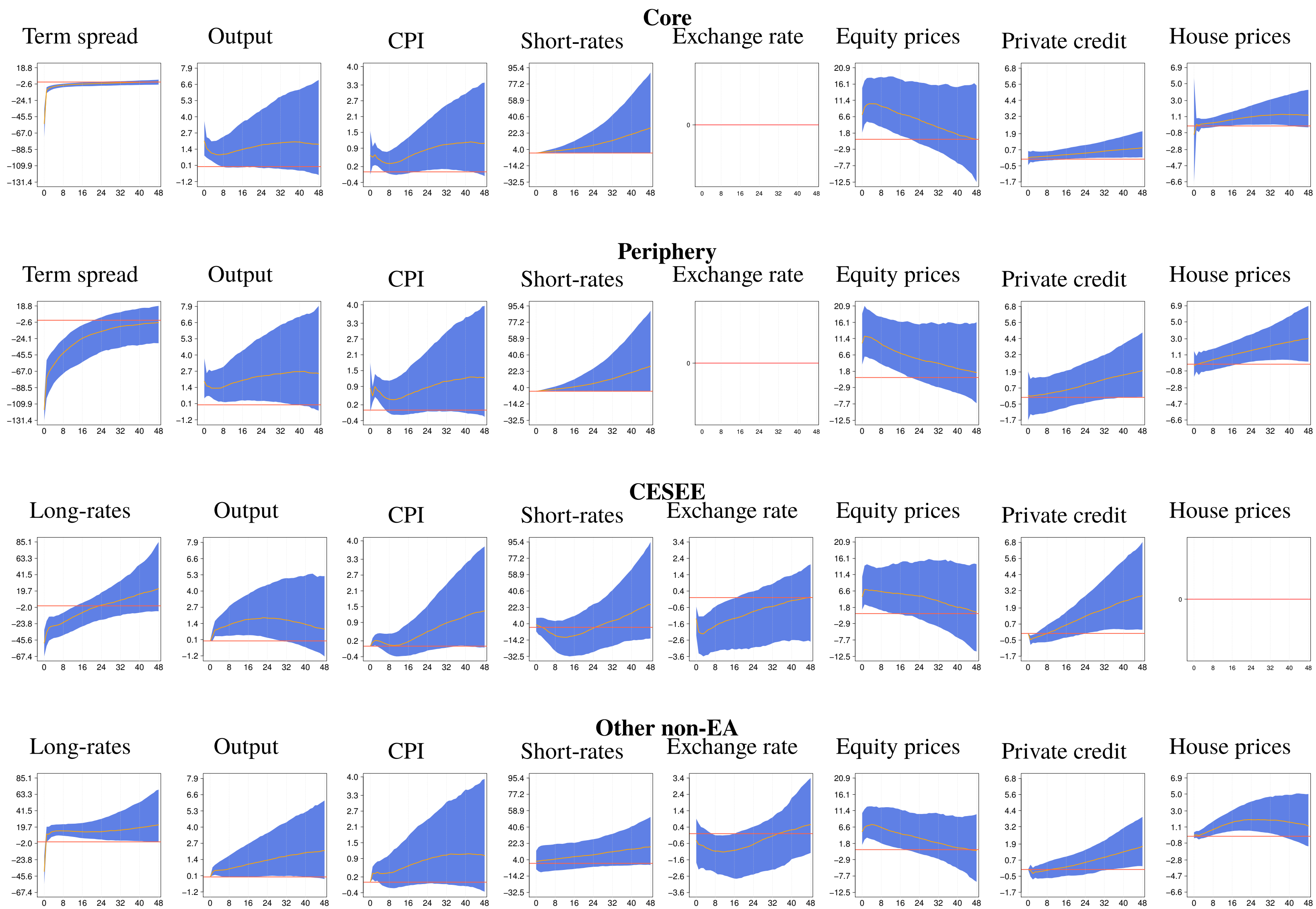
Results for the term spread shock are depicted in Fig. 2 which shows the posterior median (orange line) along with 68% credible intervals (blue shaded area). For completeness, Fig. B.1 in the appendix presents the full posterior distribution of the regional impulse responses to gauge uncertainty beyond the 68% credible intervals.

Considering the response of term spreads in the euro area reveals a negative immediate reaction that gradually fades out after around 16 months. In terms of cross-country differences, we find that short-run effects are more pronounced for euro area periphery compared to core economies. This, to some degree, reflects the shock composition that is more tilted toward peripheral countries.

The reduction in term spreads, caused by a decline in long-term yields, increases real activity. While the increase is enforced by the sign restrictions (which have to hold for the first 4 periods), estimated magnitudes of the effects are determined by the model and the data. The same holds true for medium- and long-run shapes of the impulse responses. Despite the larger immediate reduction in periphery countries' term spreads relative to those in core countries, short-run effects on output appear to be similar across both country groups. Estimated impact effects of about two percent are comparable to the findings reported in Baumeister and Benati (2013) for the USA and Kapetanios et al. (2012) for the UK. In general, output effects are more front loaded for euro area core (significant up to 8 months) compared to periphery countries (significant up to 32 months). Long-run output effects in periphery countries amount to 2.5%, which is about 0.6 percentage points, higher compared to effects in core countries. That intra-euro area responses differ corroborates findings of Georgiadis (2015), Altavilla et al. (2016) and Burriel and Galesi (2018) who show that the transmission of monetary policy in the euro area is heterogeneous. In the context of spillovers, this could imply that models which ignore intra-euro area heterogeneity for reasons of simplicity (see e.g., Dees et al., 2007; Hájek and Horváth, 2018) could provide misleading results, since not all neighboring countries have equally strong economic and financial links with the same euro area countries. In terms of international output effects, we find that responses in CESEE and other non-euro area European countries show a rather similar pattern to responses in euro area periphery countries. Comparing the size of output spillovers to domestic euro area effects, reveals somewhat smaller long-run effects (by about 1 percentage point).

Next, as economic activity picks up, consumer prices increase across all regions. In both euro area core and periphery countries, prices rise by 0.5% in the short-run (up to 8 months), while long-run effects are precisely estimated only for core countries. This result is in line with findings of Wieladek and Pascual (2016), who demonstrate a stronger effect on prices in selected euro area core compared

Fig. 2: Term spread shock - regional results



Notes: The figure shows impulse responses to a simultaneous 100bp reduction in the euro area term spread. The dark blue shaded area denotes the 68% credible set and the orange line the posterior median. Exchange rate refers to nominal exchange rates vis-à-vis the euro (decrease implying a depreciation of the euro). Regional figures are aggregated using purchasing power parities. Responses of risk spreads and the German term spread are available from the authors upon request.

to periphery countries. We also find evidence for international price spillovers. In CESEE and other non-euro area European economies, effects are significant in the short-run. In terms of magnitudes, estimated short-run responses are slightly weaker than in the euro area (by 0.1 to 0.3 percentage points). In general and with the exception of euro area core countries, we find rather short-lived effects on consumer prices. This might be driven by the link between asset prices and inflation, as recently analyzed by de Haan and van den End (2018). They find that the transmission of financial developments to inflation can be quite long and that overall effects of quantitative easing on inflation can be uncertain, both in timing and direction. Bluwstein and Canova (2016) who assess spillovers of a broad measure of unconventional monetary policy find stronger effects on prices than output. The finding is reversed, however, when solely focusing on sovereign bond purchases, which is the focus of this study.

Next, we look at responses of short-term interest rates to gauge how international central banks react to the euro area monetary policy shock. While the impact response of short-term interest rates in the euro area is by construction zero, the longer-run behavior is determined by the assumed Taylor rule. As output and prices in the euro area pick up, short-term rates increase significantly. They also increase in CESEE and other non-euro area European countries, but here credible intervals are wide. This is especially true for CESEE countries, which points toward a considerable degree of heterogeneity of monetary policy regimes within that region.

The decline in euro area term spreads could also bring down international long-term yields (Bluwstein and Canova, 2016). Couré (2017) provides stylized facts that indicate a re-allocation of capital after the launch of the asset purchase program toward more advanced economies and currencies perceived as safe havens and less so to international emerging markets such as the BRICs. Looking at the responses of international long-term rates, we find a significant and tightly estimated decline for both CESEE and other non-euro area European economies. In other non-euro area European economies, long-term rates adjust quickly and persistently after the initial drop. That long-term rates decline significantly in comparably less developed CESEE economies suggests that geographical closeness is a further important factor (besides riskiness of assets) that international investors consider. Our results hence refine the general conclusions of Couré (2017).

Another important transmission channel for small open economies is through adjustments in exchange rates. Looking at Fig. 2, we find that monetary easing in the euro area strengthens currencies of CESEE economies and other non-euro area European countries vis-à-vis the euro. Our results hence corroborate recent findings of the literature using event study approaches and high frequency data (Rogers et al., 2014; Ciarlone and Colabella, 2016; Fratzscher et al., 2016; Georgiadis and Gräb, 2016). CESEE currencies appreciate particularly strongly against the euro and responses are tightly estimated. This could be driven by a rise in capital flows toward the region (Ciarlone and Colabella, 2016). Our results on local currency movements are also in line with Chen and Tsang (2013), who show that if relative yield curves flatten, the currency of the home country (i.e., the euro area) depreciates. In fact, albeit long-term yields in both the CESEE region and other non-euro area European countries decline, yield curves relative to the euro area remain steeper.

We now turn to spillovers to other financial variables, namely real equity prices, private credit and house prices. We expect equity prices to rise due to Keynesian effects that should boost consumption and growth (Nickel and Vansteenkiste, 2013) on the one hand and through an adjustment of investors who search for yields into equity markets (Kuttner and Mosser, 2002) on the other hand. In fact, and in parallel with the reduction of term spreads, equity prices increase in the euro area and abroad. That international equity prices rise in the short-run, is in line with findings reported in the event-study literature that uses high frequency data (Fratzscher et al., 2016; Ciarlone and Colabella, 2016;

Georgiadis and Gräb, 2016). The findings provided in Fig. 2, however, show that increases in international equity prices are in fact rather persistent (up to 16 months).

Responses of private credit are significant, positive and persistent across all regions. For CESEE economies, we find especially pronounced effects that even outpace those of euro area countries (by about 1 to 2 percentage points). This result is in line with Fadejeva et al. (2017) who analyze spillovers to credit in response to a range of macroeconomic shocks. The strong response of credit might be explained by the particularly volatile business cycle movements in these countries during a large part of the time period under study. Another explanation might be given by considering the recent work of Bruno and Shin (2015), who demonstrate that a monetary policy loosening in a center country (e.g., the euro area) from which global banks obtain funding can accelerate cross-border banking flows to the countries where these banks operate (e.g., CESEE).

Finally, we analyze the impact of the term spread shock on international house prices. As the yield curve flattens, investors might re-allocate capital into other asset class such as the housing sector, thus driving up real estate prices (Iacoviello, 2005). That this is indeed the case is corroborated by our results. More specifically, we find positive and significant effects for euro area core countries (with a peak effect of about 1.4%), periphery countries (peak effect of about 3%) as well as for other non-euro area European countries (peak effect of about 2%). For CESEE economies, no data on house prices are available for the period under study.

Summing up, we find that a reduction in euro area term spreads positively and persistently affects industrial production in the euro area and abroad. International price spillovers are also positive, but effects more short-lived. These findings are in line with Horváth and Voslářová (2017) for CESEE and Gambacorta et al. (2014) for euro area core countries. Spillovers transmit via both the exchange rate and the financial channel as we see a depreciation of the euro vis-á-vis these currencies on the one hand and a rise in financial variables such as equity prices, private credit and – for non-euro area European countries – house prices on the other hand. Also, long-term yields in non-euro area countries decline significantly in responses to the reduction in euro area term spreads.

4.2. *A simple portfolio analysis*

This section complements the structural analysis by conducting a portfolio analysis. For that purpose, we compute h-step-ahead forecasts based on the structural shocks identified in section 3. These scenario-based forecast distributions are then used to compute a global minimum-variance portfolio for each forecast horizon (for an approach that uses Bayesian predictive distributions to guide portfolio choice, see, Jorion, 1985). In a second step, we benchmark the corresponding portfolio returns against returns arising from a minimum variance portfolio based on using the unconditional h-step-ahead predictive distribution and a naive portfolio that uses equal weights for each asset class. The naive portfolio is a hard-to-beat benchmark since estimation errors associated with data-based approaches potentially hamper overall performance (for a discussion, see DeMiguel et al., 2007).

To set the stage, let $p(\boldsymbol{x}_{t+h}|\boldsymbol{u}_t = \boldsymbol{\xi}, \mathcal{I}_{t-1})$ denote the h-step-ahead predictive density, $\boldsymbol{\xi}$ is a joint structural shock that yields a 100 basis point average decline in euro area term spreads, with remaining elements set equal to zero, and $\mathcal{I}_{t-1}$ represents the information set up to time $t-1$. Similarly, $p(\boldsymbol{x}_{t+h}|\boldsymbol{u}_t = \boldsymbol{0}, \mathcal{I}_{t-1})$ denotes the unconditional h-step-ahead predictive distribution with all shocks set equal to zero. Taking the conditional expectations of both predictive densities then yields the h-step-ahead impulse response function (Koop et al., 1996; Pesaran and Shin, 1998),

$$\text{IRF}(h, \boldsymbol{\xi}, \mathcal{I}_{t-1}) = E(\boldsymbol{x}_{t+h}|\boldsymbol{u}_t = \boldsymbol{\xi}, \mathcal{I}_{t-1}) - E(\boldsymbol{x}_{t+h}|\boldsymbol{u}_t = \boldsymbol{0}, \mathcal{I}_{t-1}), \tag{12}$$

which shows that one could easily compute the full predictive distribution *conditional* on the structural shocks at time t (henceforth labeled the 'conditional' scenario). In order not to mix information sets, we re-estimate the GVAR model using data until December 2014 – one month before the ECB announced its asset purchase program. After obtaining the structural impulse responses, we use Eq. (12) to compute $p(\boldsymbol{x}_{t+h}|\boldsymbol{u}_t = \boldsymbol{\xi}, \mathcal{I}_{t-1})$ using Monte Carlo integration for $h = 1, 2, \ldots, 24$ months ahead and calculate annualized Sharpe ratios.

Since we need returns of assets, our analysis is limited to financial quantities that enter the model in (log) prices, namely equity prices, exchange rates, and house prices. For these variables, the corresponding impulse responses can be interpreted as forecasts of returns in response to a term spread shock. Unfortunately, this rules out the inclusion of bond prices since these enter the model in yields. Moreover, bond-specific characteristics play a crucial role for portfolio optimization which could potentially distort our findings (for a discussion, see Bredendiek et al., 2016).

Let z_{t+h} be a vector of returns associated with m assets. The vector z_{t+h} includes the h-step-ahead returns of the three asset classes in response to a term spread shock. In what follows, $\boldsymbol{\mathcal{V}}_{t+h}^{(j)}$ denotes the posterior predictive variance-covariance matrix of the returns, for $j \in \{\text{conditional, unconditional}\}$ indicating the scenario adopted. The investor then selects a set of weights $\boldsymbol{\nu}_{t+h}^{(j)}$ such that the expected portfolio variance is minimized. Formally,

$$\begin{aligned} \underset{\boldsymbol{\nu}_{t+h}^{(j)}}{\text{minimize}} \quad & (\boldsymbol{\nu}_{t+h}^{(j)})' \boldsymbol{\mathcal{V}}_{t+h}^{(j)} \boldsymbol{\nu}_{t+h}^{(j)} \\ \text{subject to} \quad & \mathbf{1}' \boldsymbol{\nu}_{t+h}^{(j)} = 1, \end{aligned} \tag{13}$$

with $\mathbf{1}$ being an m-dimensional vector of ones. This simple approach is sufficient to adequately describe changes in the weights with respect to declines in euro area term spreads.[14] It is worth emphasizing that we consider optimal weights across forecast horizons, all obtained by solving the static optimization problem in the spirit of Markowitz (1952). In principle, one could also solve a dynamic optimization problem to obtain an optimal portfolio allocation across assets (see, inter alia, Han, 2005; Benzoni et al., 2007; Bansal and Kiku, 2011; Gârleanu and Pedersen, 2013).

Table 2 shows (annualized) Sharpe ratios, mean returns, and standard deviations across portfolio strategies. Comparing the three scenarios reveals that using information based on the dynamic responses translates into higher portfolio returns, a lower portfolio variance, and higher Sharpe ratios when benchmarked against the competing strategies. The bad performance of the portfolio based on the unconditional predictive distribution can be traced back to the fact that it seems to be backward looking, extrapolating historical price movements and correlations across asset classes. This translates into negative annualized returns and a high portfolio variance. In contrast, using information based on the structural impulse responses potentially improves the probability of observing a turning point in asset returns. Notice, however, that the naive portfolio strategy yields returns comparable to the more complex conditional scenario but with a markedly higher portfolio volatility. We conjecture that the better performance of our conditional scenario stems from using additional information on the effects of the shocks without inflating estimation noise.

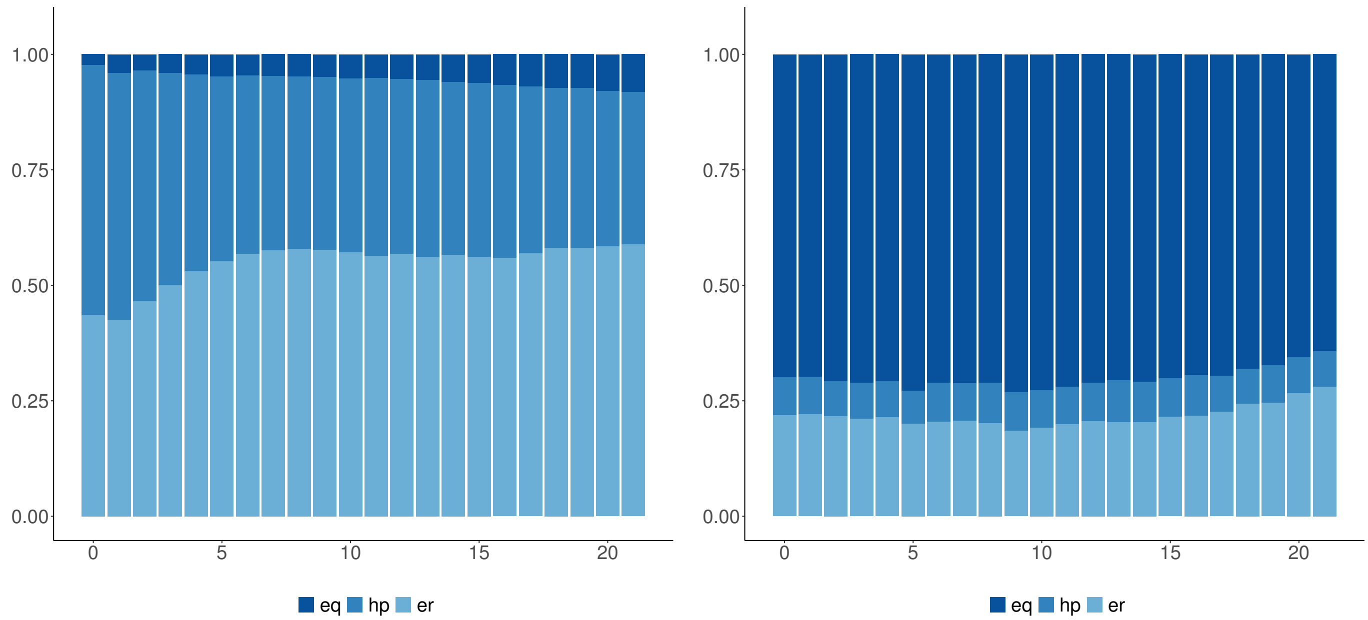
To analyze the causes of the dismal performance of the portfolio strategy based on the unconditional forecast, we consider differences in optimal weights across asset classes (see Fig. 3(a)) and countries (see Fig. 3(b)). These numbers are obtained by summing optimal weights across assets in Fig. 3(a) and across countries in Fig. 3(b).

[14]Notice that we set transaction costs equal to zero. For a recent study that deals with portfolio optimization under transaction costs, see Hautsch and Voigt (2019).

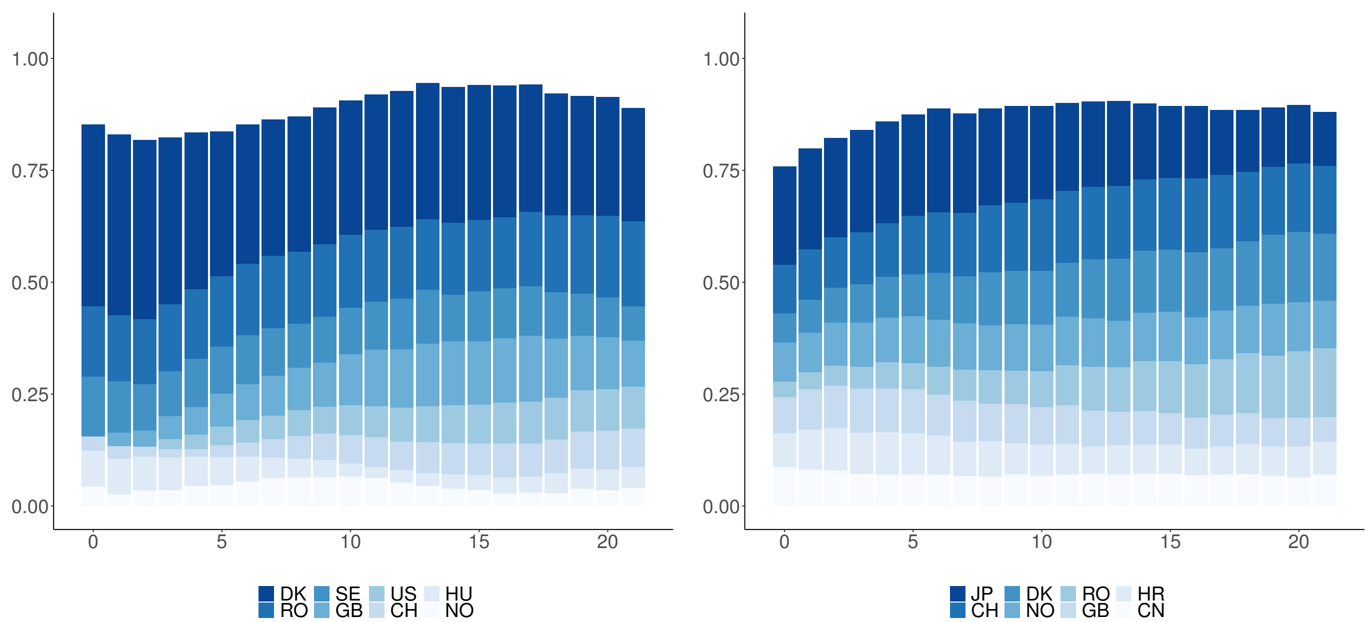
Fig. 3: Optimal portfolio weights across asset classes and countries

Conditional scenario **Unconditional scenario**

((a)) Asset classes



((b)) Countries (top 8)



Notes: This figure depicts optimal portfolio weights associated with different asset classes and countries (in terms of the top 8 countries according to the absolute weight received) between the global minimum variance portfolio based on the conditional (left column) and unconditional (right column) predictive distribution.

Table 2: Annualized Sharpe ratios, mean returns ('mean'), and standard deviations ('S.D.') across investment strategies

	Sharpe ratio	Mean	S.D.
Conditional	1.50	2.44	1.62
Unconditional	-0.21	-1.65	7.77
Naive	0.56	2.24	4.00

The upper figure clearly suggests that the stronger performance in terms of Sharpe ratios stems from two sources. First, using conditional forecasts alters the variance-covariance structure of returns vis-á-vis using unconditional predictions such that more weight is placed on housing and exchange rates. Second, the weights associated with equities appear to be much smaller as compared to the portfolio based on unconditional forecasts. In light of the rather weak performance of equity markets in selected countries in our sample, this partly explains the weak performance of this investment strategy.

Considering which countries receive the largest weights shows that, under the conditional scenario, assets in Denmark, Sweden and the USA dominate the optimal portfolio. These countries are often regarded as safe havens. Weights of the unconditional scenario show a more diverse cross-country distribution with largest weights attached to assets in Japan and Denmark on the one hand followed by developing European countries such as Romania and Croatia.

4.3. *Robustness analysis*

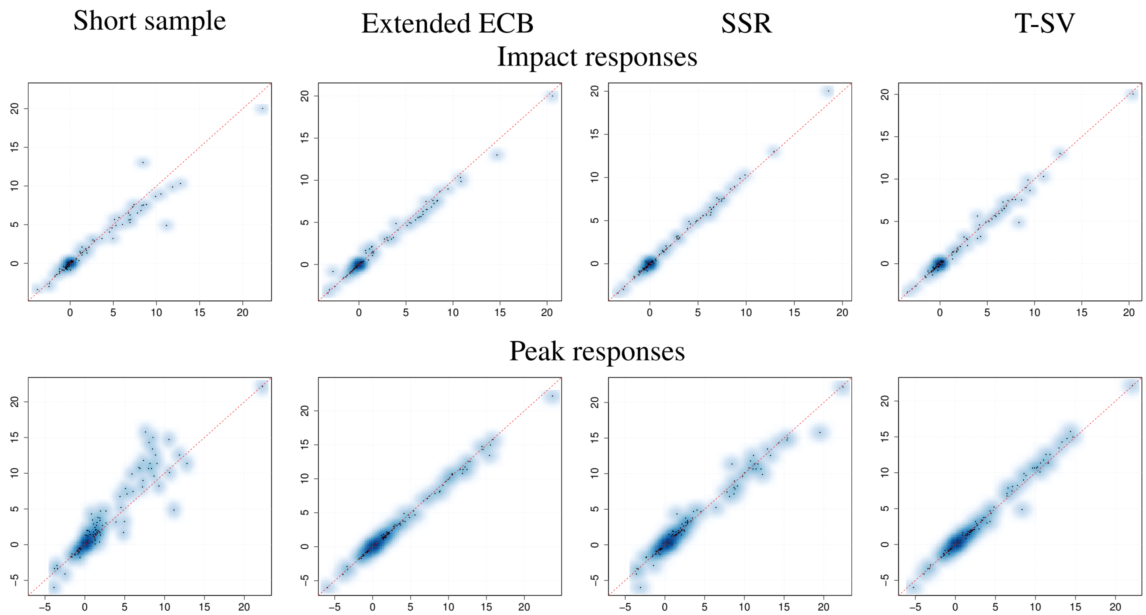
In this section, we carry out a range of robustness checks to assess the sensitivity of our results. Our sample period covers both normal times and the period of the zero lower bound. Hence, one could question the appropriateness of assuming that the ECB's monetary policy is well described by a simple Taylor rule, or more generally the validity of our identification approach across both sub-samples. We address these questions by carrying out three robustness exercises, namely changing the estimation period, modifying the identification of the term spread shock and replacing euro area short-term interest rates by a shadow rate.[15] A fourth robustness check deals with sensitivity of our results with respect to the estimated time-varying volatilities that influence the cross-country shock composition. For all robustness checks and to reduce the computational burden, we focus solely on the term spread shock and use the identification provided in the left-hand side column of Table 1 (i.e., without controlling for the risk spread shock).

First, we assess the sub-sample sensitivity of our results. Following Bluwstein and Canova (2016), Boeckx et al. (2017) and Burriel and Galesi (2018), we restrict the sample to match the period over which the ECB has intensively used its balance sheet as a policy tool. More specifically, we re-estimate the model over the period from 2007m01 to 2016m06.

Second, we enlarge the ECB country model by including the ECB's total assets and the Composite Indicator of Systemic Stress (CISS) described in Holló et al. (2012). Adding these variables to the

[15]The assessment of conventional or unconventional monetary policy shocks using models that are estimated and identified during both normal and zero lower bound periods, is frequently criticized. In turn, recent empirical approaches use time-varying parameter models to address the issue. For example, Crespo Cuaresma et al. (2019) use a time-varying parameter GVAR model and shadow rates as a policy instrument to examine spillovers from a US monetary policy shock. Kimura and Nakajima (2016) us a latent threshold approach that allows the central bank to switch between conventional and unconventional monetary policy measures.

Fig. 4: Comparing impact and peak effects for alternative estimation and identification set-ups



Notes: The picture shows scatter plots of posterior median impact and peak responses, stacked over all variables and countries. On the y-axis, responses are shown for the -100bp euro area term spread shock described in section 2. On the x-axis responses refer to a -100bp euro area term spread shock based on (I) a shortened sample (short sample), (II) an extended ECB country model (extended ECB), (III) shadow rates instead of short rates as a policy measure for the euro area, Japan, the UK and the USA (SSR) and (IV) shocks based on volatilities that refer to the last data point in our sample (T-SV). Responses of the term spread are in percentage points to facilitate comparability with responses of other variables.

ECB country model now changes the simple Taylor rule regression to a broader VAR framework with three endogenous variables, namely short-term interest rates, total assets and the CISS indicator. In addition to the sign restrictions provided in Tab. 1, we restrict changes in the ECB's total assets to increase. To rule out the systematic positive reaction of the ECB to an increase in financial stress, we further place a negative restriction on the CISS indicator. These two additional restrictions have been proposed in the context of unconventional monetary policy shocks and euro area data (Gambacorta et al., 2014; Boeckx et al., 2017). Both restrictions have to hold on impact and the subsequent three periods.

Third, we substitute so-called shadow interest rates (SSR) for short-term interest rates in major economies where actual interest rates remained at the zero lower bound for an extended period of time. These are the euro area (ECB country model), Japan, the UK, and the USA. Shadow interest rates are derived from term structure models and mirror actual rates during normal times, while they become negative during periods when the zero lower bound is binding. We use the shadow rates of Krippner (2013), which are publicly available from the webpage of the Reserve Bank of New Zealand for all four countries that witnessed periods of policy rates reaching zero in our sample.

Finally, as pointed out in subsection 2.2, the estimates of volatility can alter the shock composition across euro area countries. In case of a single shock, including stochastic volatility would simply re-scale the impulse responses. Since we are interested in a multiple euro area shock, changes in the volatility can alter the shock composition and thus affect not only the scale of estimated impulse responses but also their dynamics. Hence we cross-check our benchmark estimations that are based on averaged volatilities over the sample period with estimates referring to the last data point in our sample (2016m06).

To gauge the robustness of our baseline results for the term spread shock, we investigate scatter plots of impact and peak responses provided in Fig. 4. Points clustered around the 45 degree line indicate similar responses.

Impact responses based on the short sample, provided in the left panel of Fig. 4, are very similar to those of the benchmark results provided in Section 4. Looking at the peak responses, however, reveals stronger responses for some observations using the whole sample period. This is indicated by observations above the 45 degree line. Taken at face value, this would imply that results are overestimated using the whole sample period. However, one should bear in mind that the scatter plots relate to posterior median responses and that estimation uncertainty is considerably larger for estimation results using the short sample. Next, we investigate the robustness of our baseline results when including balance sheet data for the ECB country model. In this case both impact and peak responses tend to be very similar as indicated by observations being mostly clustered around the 45 degree line (see case "extended ECB" in Fig. 4). The same holds true when including shadow rates instead of actual short rates for countries that faced the zero lower bound constraint over our sample period. Here we also find results that are very similar to our baseline estimations (see case "SSR" in Fig. 4). Finally, we examine the sensitivity of our results with respect to estimated volatilities again finding very similar impact and peak responses compared to our baseline estimation (see case "T-SV" in Fig. 4).

5. CLOSING REMARKS

Since the global financial crisis, the ECB has implemented a range of non-standard policies.The most recent program included buying large amounts of securities issued by euro area governments, agencies and EU institutions, asset-backed securities, and covered bonds with the goal to decrease longer-term yields and spur inflation. In this paper, we examine the effects of a 100bp reduction in euro area term spreads using a Bayesian GVAR model and an identification strategy that is novel to the GVAR literature. Specifically, we combine sign restrictions with zero impact restrictions across the panel of countries considered. To disentangle the effects of unconventional monetary policy measures from purely verbal interventions, we moreover identify a risk spread shock.

Looking at *intra-euro area effects of the term spread shock* first, our results point to positive effects on output and prices (Gambacorta et al., 2014), a depreciation of the euro (Burriel and Galesi, 2018) and a rise in financial variables, such as equity and house prices as well as private credit. We also find evidence for the heterogeneity of intra-euro area effects, confirming results of Georgiadis (2015) and Burriel and Galesi (2018). For example, the effect on prices tends to be more precisely estimated and larger in euro area core economies than in euro area periphery countries (Wieladek and Pascual, 2016).

Our main results are on the *international effects* of a reduction in euro area term spreads. Here, we find evidence for significantly positive and persistent spillovers to industrial production in CESEE and other non-euro area European countries. International effects on consumer prices are also positive and tightly estimated in the short-run. This implies that euro area financial developments have significant macroeconomic consequences for these regions.

Spillovers *transmit* through both the exchange rate channel and a broad-based financial channel. Local currencies appreciate against the euro and responses are especially precisely estimated for CESEE economies, reflecting the importance of this channel for the regions' export-oriented growth model. In terms of financial variables, we find significant evidence of a decline in long-term yields in neighboring European economies as a consequence of a reduction in euro area term spreads. That long-term yields decline not only in less risky, advanced economies but also more broadly in CESEE economies, is a novel finding and generalizes conclusions of Couré (2017). Moreover, the easing of financing conditions triggers an increase in private credit. Effects are particularly pronounced in CESEE economies, reflecting the regions' strong financial linkages with the euro area. On top of that, we find evidence for significant rises in stock market prices throughout the regions. The overall positive effect on equity prices coupled with positive spillovers to GDP suggest that wealth effects are crucial in providing stimulus to the economy (see also Nickel and Vansteenkiste, 2013; Eller et al., 2017, for the case of a fiscal shock).

Last, we construct minimum variance portfolios for multi-step-ahead forecasts conditional on information from the term spread shock and compare these to alternative investment strategies. Our findings show that using structural information to guide portfolio choice pays off in terms of annualized Sharpe ratios. This holds true for both a naive benchmark that uses equal weights over assets as well as a constructed portfolio using the unconditional forecasts of the model. We find that these differences mainly stem from considerably smaller weights associated to international equity markets.

REFERENCES

Acharya VV, Eisert T, Eufinger C and Hirsch CW (2017) Whatever it Takes: The Real Effects of Unconventional Monetary Policy. Working Paper 152, SAFE Working Paper [7]

Altavilla C, Giannone D and Lenza M (2016) The Financial and Macroeconomic Effects of OMT Announcements. *International Journal of Central Banking* 12(3), 29–57 [2, 11]

Ambler S and Rumler F (2019) The effectiveness of unconventional monetary policy announcements in the euro area: An event and econometric study. *Journal of International Money and Finance* 94, 48–61 [2]

Arias J, Caldara D and Rubio-Ramirez J (2015) The systematic component of monetary policy in svars: an agnostic identification procedure. Technical Report 1131, FRB International Finance Discussion Paper [9, 10]

Babecká Kucharčuková O, Claeys P and Vašiček B (2016) Spillover of the ECB's monetary policy outside the euro area: How different is conventional from unconventional policy? *Journal of Policy Modeling* 38(2), 199–225 [2]

Bansal R and Kiku D (2011) Cointegration and long-run asset allocation. *Journal of Business & Economic Statistics* 29(1), 161–173 [15]

Baumeister C and Benati L (2013) Unconventional Monetary Policy and the Great Recession: Estimating the Macroeconomic Effects of a Spread Compression at the Zero Lower Bound. *International Journal of Central Banking* 9(2), 165–212 [2, 7, 10, 11]

Bekaert G, Hoerova M and Lo Duca M (2013) Risk, uncertainty and monetary policy. *Journal of Monetary Economics* 60, 771–788 [10]

Benati L and Goodhart C (2008) Investigating time-variation in the marginal predictive power of the yield spread. *Journal of Economic Dynamics and Control* 32(4), 1236–1272 [8]

Benzoni L, Collin-Dufresne P and Goldstein RS (2007) Portfolio choice over the life-cycle when the stock and labor markets are cointegrated. *The Journal of Finance* 62(5), 2123–2167 [15]

Bernanke BS and Kuttner KN (2005) What explains the stock market's reaction to federal reserve policy? *Journal of Finance* 60(3), 1221–1257 [10]

Blinder AS (2012) Revisiting Monetary Policy in a Low-Inflation and Low-Utilization Environment. *Journal of Money, Credit and Banking* 44(s1), 141–146 [2]

Bluwstein K and Canova F (2016) Beggar-Thy-Neighbor? The International Effects of ECB Unconventional Monetary Policy Measures. *International Journal of Central Banking* 12(3), 69–120 [2, 13, 17]

Bobeica E and Jarociński M (2019) Missing Disinflation and Missing Inflation: A VAR Perspective. *International Journal of Central Banking* 15(1), 199–232 [2]

Boeckx J, Dossche M and Peersman G (2017) Effectiveness and Transmission of the ECB's Balance Sheet Policies. *International Journal of Central Banking* 13(1), 297–333 [17, 18]

Bredendiek M, Ottonello G and Valkanov RI (2016) Corporate Bond Portfolios and Macroeoconomic Conditions. *Available at SSRN* [15]

Bruno V and Shin HS (2015) Capital flows and the risk-taking channel of monetary policy. *Journal of Monetary Economics* 71, 119–132 [14]

Burriel P and Galesi A (2018) Uncovering the heterogeneous effects of ecb unconventional monetary policies across euro area countries. *European Economic Review* 101, 201–229 [2, 10, 11, 17, 19]

Carriero A, Clark TE, Marcellino M et al. (2019) Large Vector Autoregressions with Asymmetric Priors. *Journal of Econometrics* forthcoming [24]

Chen Q, Filardo A, He D and Zhu F (2016) Financial crisis, US unconventional monetary policy and international spillovers. *Journal of International Money and Finance* 67(C), 62–81 [2]

Chen Y and Tsang KP (2013) What Does the Yield Curve Tell Us about Exchange Rate Predictability? *The Review of Economics and Statistics* 95(1), 185–205 [13]

Chow G and Lin A (1971) Best Linear Unbiased Interpolation, Distribution, and Extrapolation of Time Series By Related Series. *Review of Economics and Statistics* 53, 372–375 [3]

Ciarlone A and Colabella A (2016) Spillovers of the ECB's non-standard monetary policy into CESEE economies. *Ensayos sobre política económica* 34(81), 175–190 [13]

Cogley T, Morozov S and Sargent TJ (2005) Bayesian fan charts for UK inflation: Forecasting and sources of uncertainty in an evolving monetary system. *Journal of Economic Dynamics and Control* 29(11), 1893–1925 [24]

Cogley T and Sargent TJ (2005) Drifts and volatilities: monetary policies and outcomes in the post WWII US. *Review of Economic dynamics* 8(2), 262–302 [5]

Couré B (2017) The international dimension of the ECB's asset purchase programme. Speech at the Foreign Exchange Contact Group meeting, 11 July 2017 [13, 20]

Crespo Cuaresma J, Doppelhofer G, Feldkircher M and Huber F (2019) Spillovers from US monetary policy: evidence from a time varying parameter global vector auto-regressive model. *Journal of the Royal Statistical Society: Series A*, forthcoming [17]

Crespo Cuaresma J, Feldkircher M and Huber F (2016) Forecasting with Global Vector Autoregressive Models: a Bayesian Approach. *Journal of Applied Econometrics* Vol. 31(7), 1371–1391 [2]

D'Amico S and King TB (2013) Flow and stock effects of large-scale treasury purchases: Evidence on the importance of local supply. *Journal of Financial Economics* 108(2), 425–448 [2]

de Haan L and van den End JW (2018) The signalling content of asset prices for inflation: Implications for quantitative easing. *Economic Systems* 42(1), 45–63 [13]

Dees S, di Mauro F, Pesaran HM and Smith LV (2007) Exploring the international linkages of the euro area: a global VAR analysis. *Journal of Applied Econometrics* 22(1) [8, 11]

DeMiguel V, Garlappi L and Uppal R (2007) Optimal versus naive diversification: How inefficient is the 1/N portfolio strategy? *The Review of Financial Studies* 22(5), 1915–1953 [14]

Eickmeier S and Ng T (2015) How do US credit supply shocks propagate internationally? A GVAR approach. *European Economic Review* 74, 128 – 145 [4, 10]

Eller M, Feldkircher M and Huber F (2017) How would a fiscal shock in Germany affect other European countries? Evidence from a Bayesian global VAR model with sign restrictions. *Focus on European Economic Integration* Q1, 1–30 [20]

European Central Bank (2017) Low inflation in the euro area: causes and consequences. Technical Report 181, Occasional Paper Series [2]

Fadejeva L, Feldkircher M and Reininger T (2017) International Spillovers from Euro Area and US Credit and Demand Shocks: A focus on Emerging Europe. *Journal of International Money and Finance* 70, 1–25 [10, 14]

Fawley BW and Neely CJ (2013) Four stories of quantitative easing. *Review* , 51–88 [2]

Feldkircher M and Huber F (2016) The international transmission of US shocks – Evidence from Bayesian global vector autoregressions. *European Economic Review* 81(C), 167–188 [2, 4, 6, 10, 24]

Fratzscher M, Duca ML and Straub R (2016) ECB Unconventional Monetary Policy: Market Impact and International Spillovers. *IMF Economic Review* 64(1), 36–74 [13]

Gagnon J, Raskin M, Remache J and Sack B (2011) The Financial Market Effects of the Federal Reserve's Large-Scale Asset Purchases. *International Journal of Central Banking* 7(1), 3–43 [2]

Gambacorta L, Hofmann B and Peersman G (2014) The Effectiveness of Unconventional Monetary Policy at the Zero Lower Bound: A Cross-Country Analysis. *Journal of Money, Credit and Banking* 46(4), 615–642 [2, 14, 18, 19]

Gârleanu N and Pedersen LH (2013) Dynamic trading with predictable returns and transaction costs. *The Journal of Finance* 68(6), 2309–2340 [15]

George EI and McCulloch R (1993) Variable selection via Gibbs sampling. *Journal of the American Statistical Association* 88, 881–889 [6]

George EI, Sun D and Ni S (2008) Bayesian stochastic search for VAR model restrictions. *Journal of Econo-*

metrics 142(1), 553–580 [6, 24]

Georgiadis G (2015) Examining asymmetries in the transmission of monetary policy in the euro area: Evidence from a mixed cross-section global VAR model. *European Economic Review* 75(C), 195–215 [2, 4, 10, 11, 19]

Georgiadis G and Gräb J (2016) Global financial market impact of the announcement of the ECB's asset purchase programme. *Journal of Financial Stability* 26(C), 257–265 [13, 14]

Hájek J and Horváth R (2018) International spillovers of (un)conventional monetary policy: The effect of the ECB and the US Fed on non-euro EU countries. *Economic Systems* 42, 91–105 [11]

Han Y (2005) Asset allocation with a high dimensional latent factor stochastic volatility model. *The Review of Financial Studies* 19(1), 237–271 [15]

Hautsch N and Voigt S (2019) Large-scale portfolio allocation under transaction costs and model uncertainty. *Journal of Econometrics* forthcoming [15]

Holló D, Kremer M and Lo Duca M (2012) CISS - a composite indicator of systemic stress in the financial system. Working Paper Series 1426, European Central Bank [17]

Horváth R and Voslářová K (2017) International spillovers of ECB's unconventional monetary policy: the effect on Central Europe. *Applied Economics* 49(24), 2352–2364 [2, 14]

Huber F (2016) Density forecasting using Bayesian global vector autoregressions with stochastic volatility. *International Journal of Forecasting* 32(3), 818–837 [5]

Iacoviello M (2005) House prices, borrowing constraints, and monetary policy in the business cycle. *The American economic review* 95(3), 739–764 [14]

Inoue A and Rossi B (2018) The Effects of Conventional and Unconventional Monetary Policy: A New Identification Procedure. Technical report, Barcelona Graduate School of Economics [2]

Jorion P (1985) International portfolio diversification with estimation risk. *Journal of Business* , 259–278 [14]

Joyce M, Miles D, Scott A and Vayanos D (2012) Quantitative Easing and Unconventional Monetary Policy – an Introduction. *The Economic Journal* 122(564), F271–F288 [2]

Kapetanios G, Mumtaz H, Stevens I and Theodoridis K (2012) Assessing the economy-wide effects of quantitative easing. *The Economic Journal* 122(564), F316–F347 [2, 11]

Kastner G (2016) Dealing with Stochastic Volatility in Time Series Using the R Package stochvol. *Journal of Statistical Software* 69(1), 1–30 [24]

Kastner G and Frühwirth-Schnatter S (2014) Ancillarity-sufficiency interweaving strategy (ASIS) for boosting MCMC estimation of stochastic volatility models. *Computational Statistics & Data Analysis* 76, 408–423 [24]

Kimura T and Nakajima J (2016) Identifying conventional and unconventional monetary policy shocks: A latent threshold approach. *The B.E. Journal of Macroeconomics* 16, 277–300 [17]

Koop G, Pesaran MH and Potter SM (1996) Impulse response analysis in nonlinear multivariate models. *Journal of Econometrics* 74(1), 119–147 [14]

Krippner L (2013) Measuring the stance of monetary policy in zero lower bound environments. *Economics Letters* 118(1), 135–138. [18]

Krishnamurthy A and Vissing-Jorgensen A (2011) The Effects of Quantitative Easing on Interest Rates: Channels and Implications for Policy. *Brookings Papers on Economic Activity* 43(2 (Fall)), 215–287 [2]

Kuttner KN and Mosser PC (2002) The monetary transmission mechanism: some answers and further questions. *Economic Policy Review* , 15–26 [10, 13]

Lenza M, Pill H and Reichlin L (2010) Monetary policy in exceptional times. *Economic Policy* 25, 295–339 [2]

Markowitz H (1952) The utility of wealth. *Journal of Political Economy* 60(2), 151–158 [15]

Moder I (2019) Spillovers from the ECB's non-standard monetary policy measures on Southeastern Europe. *International Journal of Central Banking* forthcoming [2]

Mumtaz H and Surico P (2009) The Transmission of International Shocks: A Factor-Augmented VAR Approach. *Journal of Money, Credit and Banking* 41(s1), 71–100 [8]

Nickel C and Vansteenkiste I (2013) The international spillover of fiscal spending on financial variables. In Di Mauro F and Pesaran MH, eds., *The GVAR Handbook: Structure and Applications of a Macro Model of the Global Economy for Policy Analysis*. Oxford University Press, 182–192 [13, 20]

Pesaran MH, Schuermann T and Weiner SM (2004) Modeling Regional Interdependencies Using a Global Error-Correcting Macroeconometric Model. *Journal of Business and Economic Statistics, American Statistical Association* 22, 129–162 [3, 5]

Pesaran MH and Shin Y (1998) Generalized impulse response analysis in linear multivariate models. *Economics Letters* 58(1), 17–29 [14]

Primiceri GE (2005) Time varying structural vector autoregressions and monetary policy. *The Review of Economic Studies* 72(3), 821–852 [5]

Rigobon R and Sack B (2004) The impact of monetary policy on asset prices. *Journal of Monetary Economics* 51(8), 1553–1575 [10]

Rogers JH, Scotti C and Wright JH (2014) Evaluating Asset-Market Effects of Unconventional Monetary Policy:

A Multi-Country Review. *Economic Policy* 29(80), 3–50 [13]
Sims CA and Zha T (2006) Were there regime switches in US monetary policy? *The American Economic Review* 96(1), 54–81 [5]
Thorbecke W (1997) On stock market returns and monetary Policy. *Journal of Finance* 52(2), 635–654 [10]
Timmer MP, Dietzenbacher E, Los B, Stehrer R and de Vries GJ (2015) An Illustrated User Guide to the World Input-Output Database: the Case of Global Automotive Production. *Review of International Economics* 23, 575–605 [4]
Wieladek T and Pascual AIG (2016) The European Central Bank's QE: A New Hope. CESifo Working Paper Series 5946, CESifo Group Munich [10, 11, 19]
Wright JH (2011) Term Premia and Inflation Uncertainty: Empirical Evidence from an International Panel Dataset. *American Economic Review* 101(4), 1514–34 [2]

A. PRIOR SETUP AND POSTERIOR INFERENCE

This section provides additional details on the priors for the remaining parameters of the model. For the binary indicators that control the mixture to use we impose a Bernoulli prior with

$$\delta_{ij} \sim Bernoulli(p_{ij}). \tag{14}$$

We set $p_{ij} = Prob(\delta_{ij} = 1) = 1/2$ for all i, j. This implies that a priori, all variables are equally likely to enter Eq. (2).

Similarly to the prior on the regression coefficients we impose a SSVS prior on the off-diagonal elements of U_i,

$$u_{ij,n}|\kappa_{ij,n} \sim \mathcal{N}(0, \varphi^2_{ij,n0})\kappa_{ij,n} + \mathcal{N}(0, \varphi^2_{ij,n1})(1 - \kappa_{ij,n}), \tag{15}$$

where $\kappa^2_{i,jn}$ is, again, a Bernoulli distributed random quantity that selects the mixture Gaussian component and $\varphi^2_{ij,n0}, \varphi^2_{ij,n1}$ are prior scalings such that $\varphi^2_{ij,n0} \gg \varphi^2_{ij,n1}$.

Since prior information on inclusion/exclusion of a given covariance parameter is rather scarce, we again adopt a Bernoulli prior with prior inclusion probability set to $q_{ij,n} = Prob(\kappa_{ij,n} = 1) = 1/2$,

$$\kappa_{ij,n} \sim Bernoulli(q_{ij,n}). \tag{16}$$

We follow Kastner and Frühwirth-Schnatter (2014) and impose a normally distributed prior on $\mu_{ij} \sim \mathcal{N}(0, 10^2)$, a Beta distributed prior on $\frac{\rho_{ij}+1}{2} \sim \mathcal{B}(25, 5)$ and a Gamma prior on $\varsigma^2_{ij} \sim \mathcal{G}(1/2, 1/2)$.

As mentioned in Section 2.2, our model collapses to a homoscedastic GVAR model if ς^2_{ij} equals zero. The Gamma prior on ς^2_{ij} is equivalent to imposing a normally distributed prior on $\pm\varsigma_{ij}$,

$$\varsigma^2_{ij} \sim \mathcal{G}(1/2, 1/2) \Leftrightarrow \pm\varsigma_{ij} \sim \mathcal{N}(0, 1). \tag{17}$$

This prior centers ς_{ij} on zero, if necessary and thus softly shrinks the model toward a homoscedastic specification.

Posterior simulation is carried out by sampling from the $N + 1$ country-specific posterior distributions in parallel. The MCMC algorithm is standard in the literature for VAR models. Specifically, we sample C_i on an equation-by-equation basis (for details, see Carriero et al., 2019) from an multivariate normal distribution. The free elements of U_i can be simulated by noting that the system can be rewritten as a set of k_i univariate regression models with standard normally distributed errors (see Cogley et al., 2005). The log-volatilities and the parameters of the state equation Eq. (5) are simulated by means of the algorithm stipulated in Kastner and Frühwirth-Schnatter (2014) and implemented in the R package stochvol (Kastner, 2016). The indicator variables δ_{ij} and $\kappa_{ij,n}$ are sampled from their Bernoulli distributed conditional posterior distributions.For further information on the specific posterior moments, see Feldkircher and Huber (2016).

Finally, we specify the remaining hyperparameters for the prior. More specifically, following George et al. (2008) we set $\tau^2_{ij,0} = 3\hat{\sigma}^2_{ij}$ and $\tau^2_{ij,1} = 0.1\hat{\sigma}^2_{ij}$, where $\hat{\sigma}^2_{ij}$ are the OLS variances associated with c_{ij}. For the covariance parameters, we simply specify $\varphi^2_{ij,n0} = 3$ and $\varphi^2_{ij,n1} = 0.1$ for all i, j, n. For μ_j we set $v_\mu = 10^2$, leading to a rather uninformative prior on the level of the log-volatility. Finally, for the persistence parameter we set $a_0 = 25$ and $b_0 = 5$, placing significant mass on high persistence regions. We execute the MCMC algorithm for each country using parallel computing and 120,000 iterations with the first 60,000 being discarded as burn-ins. Due to storage limits we use a thinning interval to select 9,000 out of the 60,000 posterior draws. From these, we sort out unstable posterior draws which are characterized by large eigenvalues of the companion form of the global model which leads to 7,085 posterior draws upon which the impulse response analysis in Section 4 is based.

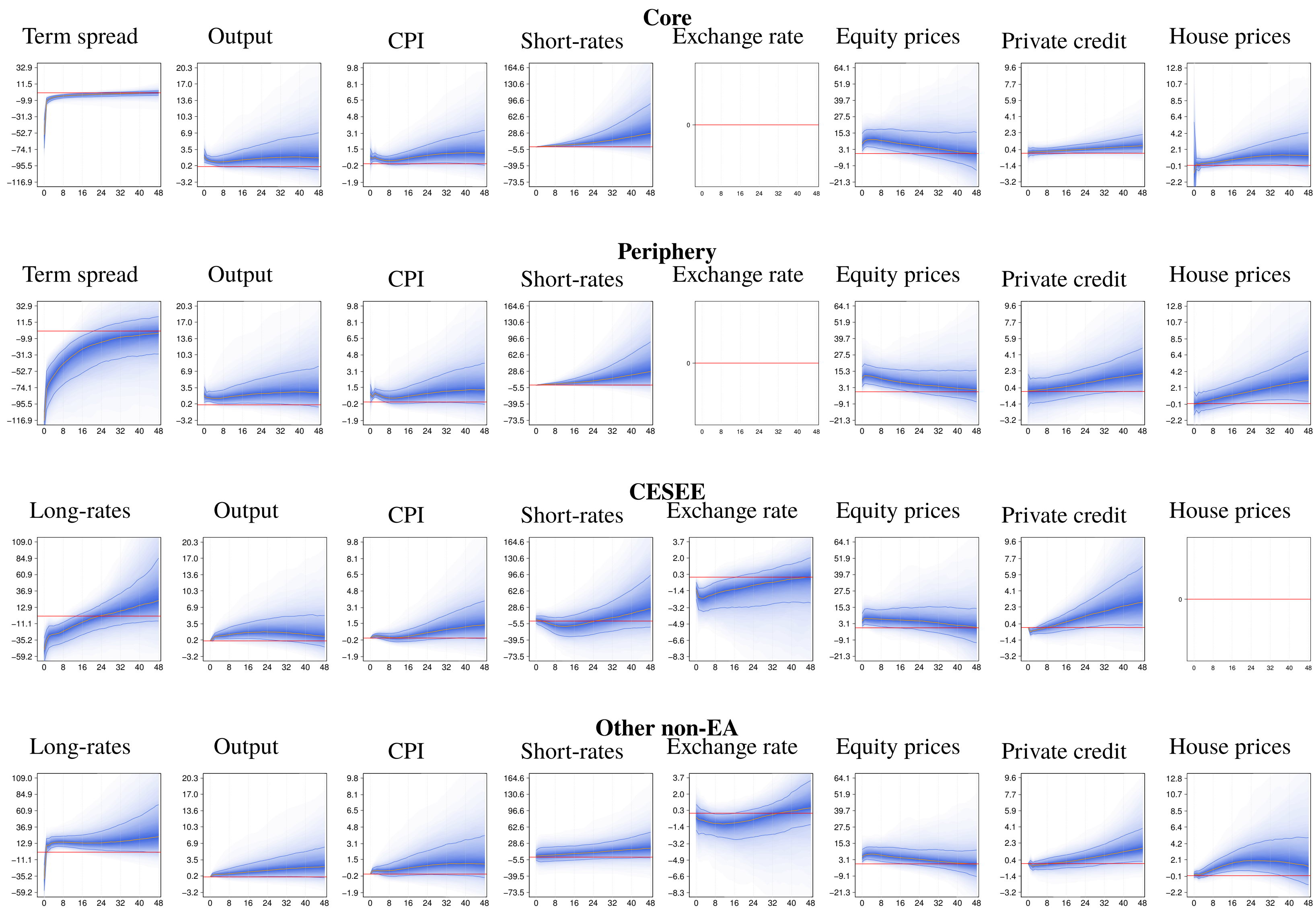
B. ADDITIONAL RESULTS

Table B.1: Variable coverage per country.

Country	Variables								
AT	y	p	tp	tp	pc	eq	hp		
BE	y	p	tp	tp	pc	eq	hp		
DE	y	p	tp		pc	eq	hp		
ES	y	p	tp	tp	pc	eq	hp		
FI	y	p	tp	tp	pc	eq	hp		
FR	y	p	tp	tp	pc	eq	hp		
IE	y	p	tp	tp	pc	eq	hp		
IT	y	p	tp	tp	pc	eq	hp		
NL	y	p	tp	tp	pc	eq	hp		
PT	y	p	tp	tp	pc	eq			
GR	y	p	tp	tp	pc	eq			
SK	y	p	tp	tp	pc	eq			
BG	y	p	i_s	i_l		pc	eq		
CZ	y	p	i_s	i_l	er	pc	eq		
HR	y	p	i_s		er	pc	eq		
HU	y	p	i_s	i_l	er	pc	eq		
PL	y	p	i_s	i_l	er	pc	eq		
RO	y	p	i_s		er	pc	eq		
SI	y	p	i_s			pc	eq		
CH	y	p	i_s	i_l	er	pc	eq	hp	
DK	y	p	i_s	i_l	er	pc	eq	hp	
GB	y	p	i_s	i_l	er	pc	eq	hp	
NO	y	p	i_s	i_l	er	pc	eq	hp	
SE	y	p	i_s	i_l	er	pc	eq	hp	
CA	y	p	i_s	i_l	er		eq	hp	
CN	y	p	i_s	i_l	er	pc	eq		
JP	y	p	i_s	i_l	er	pc	eq	hp	
RU	y	p	i_s	i_l	er	pc	eq	hp	
TR	y	p	i_s		er	pc	eq		
US	y	p	i_s	i_l	er	pc	eq	hp	$poil$

Notes: Endogenous variables for each country model. Country codes refer to ISO alpha-2. i_s^{EA} included in the ECB country model.

Fig. B.1: Term spread shock - regional results



Notes: The figure shows the full posterior distribution of impulse responses to a simultaneous 100bp reduction in the euro area term spread. The dark blue lines denote the 68%, credible sets and the orange line the posterior median. Exchange rate refers to nominal exchange rates vis-à-vis the euro (decrease implying a depreciation of the euro). Regional figures are aggregated using purchasing power parities. Responses of risk spreads and the German term spread are available from the authors upon request.