Spillover effects of credit default risk in the euro area and the effects on the Euro: A GVAR approach*

Timo Bettendorf
Deutsche Bundesbank

Abstract

During the 2008 financial crisis, increasing risk and spillovers became a main concern for policy makers and banks. In addition, changes in sovereign and bank risk are believed to have had strong effects on world-wide exchange rates. This paper aims to analyze these dynamics empirically. We estimate a Global VAR (GVAR) model for nine EMU countries plus Japan, the United Kingdom as well as the United States and identify structural risk shocks using sign restrictions, which are based on a theoretical model by Acharya et al. (2014, JF). Our results indicate that spillover effects of general risk are much stronger than those of bailouts. Furthermore, we demonstrate that the Euro depreciates significantly against the Yen and US Dollar following general risk shocks in the euro area and only to a small extent following bailout shocks. The Pound Sterling is not affected by any of these shocks. The Euro variability is, from the EMU perspective, mainly driven by shocks stemming from large countries (e.g. Germany, France and Italy). However, shocks from third countries also play an important role.

Keywords: credit default swaps, bailouts, exchange rates, global var

JEL classification: C55, F31, H63.

*Contact address: Timo Bettendorf, Deutsche Bundesbank, Wilhelm-Epstein-Strasse 14, 60431 Frankfurt am Main, E-Mail: Timo.Bettendorf@bundesbank.de. The author is grateful to Angela Abbate, Sandra Eickmeier, Christoph Fischer, Ulrich Grosch, Miguel A. León-Ledesma, Irina Stanga as well as participants of the Bundesbank research seminar for valuable comments and suggestions. This paper represent the authors' personal opinion and does not necessarily reflect the views of the Deutsche Bundesbank or its staff.

1 Introduction

When the 2008 financial crisis began, credit default risk and contagion became a big concern for banks and policy makers. Some financial institutions were so distressed that they became a threat to the domestic and even global banking system. Therefore, some governments bailed out domestic banks by transferring risk from the bank balance sheets to public sector balance sheets. This, however, led to so-called feedback loops (see Acharya, Drechsler, and Schnabl (2014)) or doom loops (see Farhi and Tirole (2014)), as banks usually hold government bonds in their portfolios, which became more risky due to the risk transfer.

These measures even intensified the crisis in countries such as Ireland, leading to a sovereign debt crisis, which also threatened to affect foreign economies. This had far-reaching consequences, as exchange rates became more and more affected by sovereign risk (see Della Corte, Sarno, Schmeling, and Wagner (2014)). Our aim, is therefore, to investigate the international effects of increases in general credit default risk and bank bailouts in the EMU as well as their effects on the Euro. The main contribution of this research is to decompose credit default risk into general risk as well as risk related to bailouts in a global model and to quantify the international spillover effects of idiosyncratic shocks.

Spillover effects of credit default risk have been extensively discussed in the literature. They occur between sovereigns, between banks, but also between sovereigns and banks. This can happen via different channels (see BIS (2011), Prisker (2001) and De Bruyckere, Gerhardt, Schepens, and Vennet (2013)). Using the canonical model of contagion by Pesaran and Pick (2007), Metiu (2013) finds evidence for contagion in EMU bond yields. De Bruyckere et al. (2013) define contagion as excess correlation and find empirical evidence for contagion between banks and sovereigns in 2012. A shift in the sources of contagion is documented by Caceres, Guzzo, and Segoviano (2010). They find that the source of contagion shifted from countries such as Austria, the Netherlands and Ireland, which were severely affected by the financial crisis, to countries with high short-term refinancing risk and uncertain long-term fiscal sustainability. Fratzscher and Rieth (2015) show that there was a strong bank-sovereign nexus during the crisis and that policies undertaken to reduce risk led to negative feedback loops as suggested by Acharya et al. (2014). Studies more closely related to our research are Alter and Beyer (2014), Gray, Gross, Paredes, and Sydow (2013) and Gross and Kok (2013). Alter and Beyer (2014) measure credit risk spillovers across sovereign and bank CDS between October 2009 and July 2012 using a methodology based on Diebold and Yilmaz (2011).[1] Their results show increasing interdependencies between banks and sovereigns and reveal mitigating effects of policy measures such as the EFSF and LTROs. Gross and Kok (2013) confirm these findings in a Mixed-Cross-Section GVAR and show further that spillover potential was high for banks in 2008 and high for sovereigns in 2012. To investigate the effects of sovereign and bank risk on the real economy, Gray et al. (2013) employ a different risk measure, namely expected loss, and add a macro sphere to the Mixed-Cross-Section GVAR. Their results indicate a strong contraction in real activity following sovereign risk shocks. These studies rely on generalized impulse responses, which have the advantage that they are invariant to the ordering of the variables and can easily be implemented in a framework such as Diebold and Yilmaz (2011). However, they do not allow for a structural interpretation. Our identification strategy is based on an economic model and enables us to give the identified shocks an economic interpretation.

The literature on linkages between credit risk and exchange rates is scarce. Only Della Corte et al. (2014) provide an extensive analysis of the relationship between changes in sovereign CDS and exchange rates. Their findings suggest that an increase in a country's sovereign credit default risk translates into a depreciation of the domestic currency. They explain this behavior using a model in which agents ask for a risk premium for holding a specific currency. As sovereign credit risk is priced into this risk premium, the value of the currency falls with an increase in sovereign credit default risk *et vice versa*. However,

[1] Diebold and Yilmaz (2011) provide a measure of interdependence of asset returns which is based on variance decompositions.

this relationship can also be explained by the standard monetary model of exchange rate determination. If, for example, a negative business cycle shock reduces output, then the exchange rate will depreciate. At the same time, the negative business cycle shock is supposed to increase the sovereign (and bank) CDS spread (see Stanga (2014)) because economic conditions worsen.

Other studies focus on FX options and volatility, but not on currency returns. Hui and Fong (2015) show that in the long run currency options of the US Dollar, Yen, Swiss Franc and the Euro are driven by corresponding CDS. Carr and Wu (2007) demonstrate, using data for Mexico and Brazil, that CDS and currency option-implied volatilities are positively related.

Within this context, the literature often fails to differentiate between contagion and spillover effects. In line with Alter and Beyer (2014), we follow a definition by Allen and Gale (2000), who interpret contagion as excess spillover effects. Spillover effects can thus be seen as a necessary condition for contagion. Consequently, it is desirable to focus on spillover effects of credit default risk rather than contagion, which is difficult to capture without additional assumptions regarding *excess*.

This research contributes to two strands of literature on credit default risk. First, we contribute to the body of literature that is concerned with spillover effects and contagion within the nexus of sovereign and bank credit default risk by identifying country-specific (structural) general risk and bailout shocks within a GVAR model. The model enables us to quantify the domestic and international effects of these structural shocks on CDS premia. Second, we contribute to the literature that analyzes the effects of credit default risk on exchange rates, as the model provides us with evidence on how these structural risk shocks affect exchange rates. In particular, we may quantify the relative importance of country-specific shocks within the EMU on the Euro during the crisis.

Our results demonstrate that spillover effects of general risk shocks to sovereign and bank CDS spreads are stronger and more important than those of bailout shocks. Moreover, we observe heterogeneity in spillovers across different regions, as non-EMU countries are significantly less affected by shocks which originate in the EMU. Exchange rate effects of risk shocks are strong and depend on the economic size of the country in which the shock originates. However, third-country effects are strong as well.

The remainder of this paper is structured as follows. In Section (2) we present the data and methodology which we apply for our analysis. The results of this study are discussed in Section (3). Finally, Section (4) concludes.

2 Methodology

Estimating the effects of credit default risk on the Euro is not straightforward, as the euro area (EA) is an aggregate of countries in which credit default risk occurs at the national level and is unequally distributed. Considering aggregates of EA variables may thus provide evidence only from a very general perspective. However, by considering a large set of variables, standard VAR models in the spirit of Sims (1980) become subject to the so-called curse of dimensionality, as the number of regressors increases exponentially with the number of variables involved. Chudik and Pesaran (2014) discuss different methodologies which may help to circumvent this problem. Besides model reduction (see, for example, Doan, Litterman, and Sims (1984)) and data shrinkage (see, for example, Bernanke, Boivin, and Eliasz (2005)), they propose the Global VAR model, which shrinks the model and data using restrictions that may, for example, be derived from trade flows or financial flows. Within the GVAR, each economy is modeled by an individual VAR model in which country-specific shocks can be identified. Moreover, variables such as the Euro, which are affected by several countries simultaneously, can be modeled more consistently. The GVAR is thus best suited for our analysis of spillover effects of credit default risk and its effects on the Euro.

We employ a GVAR model (see Pesaran, Schuermann, and Weiner (2004), Dees, di Mauro, Pesaran, and Smith (2007) and Dees, Holly, Pesaran, and Smith (2007)) with

12 countries[2] (see Table (1)) to estimate the linkages between different risk shocks and exchange rates across countries $i = 1, ..., N$. Using the sign restriction approach proposed by Eickmeier and Ng (2015), we can identify country-specific structural shocks.

Table 1: Countries

Region	Countries
EMU	Germany, France, Belgium, Ireland, Netherlands, Portugal, Spain, Austria, Italy
Non-EMU	Japan, United Kingdom, United States

Note: The table separates the countries used in the model into different regions.

2.1 Data and model specification

For our analysis, we use weekly data for the time period 2008-2011, which provides us with 157 observations. This is the main period in which governments bailed out financially distressed banks.[3] We rely on 5-year sovereign and bank credit default swaps denominated in USD from Markit. For sovereigns, we consider CDS spreads with CR (full restructuring) Doc Clauses, which are the sovereign CDS derivatives with the highest liquidity. For banks, we use CDS with the Doc Clause preferred by the corresponding region. Therefore, EU bank CDS have an MM (modified-modified restructuring) Doc Clause, Japanese banks a CR Doc Clause and US banks an XR (no restructuring) Doc Clause (see Packer and Zhu (2005) for further information on contractual terms of CDS). We build proxy variables for country-specific bank risk by computing unweighted averages of bank CDS spreads (see Table (7) in the Appendix). Nominal exchange rate data for the Euro refer to reference rates from the ECB statistical data warehouse. Exchange rates are expressed in quantity quotation from a euro-area perspective, meaning that an increase in the exchange rate

[2]The choice of countries is limited by data availability. Therefore, we focus on those EMU countries for which data are available as well as the three countries providing the most important currencies. Due to the Greek default, CDS data on Greece are not available for the full sample. A workaround would be to use expected loss data as Gray et al. (2013). This, however, would prevent us from studying the important relationship between *observable* CDS data and exchange rates.

[3]After 2011, the volatility of the CDS time series increases dramatically, leading to a strong increase in the volatility of the shocks. This would overshadow the effects of the policy measures during the banking crisis.

corresponds to an appreciation of the Euro. In order to minimize possible biases caused by different trading hours across countries, data in daily frequency are converted into weekly frequency (end of period). To account for global risk appetite we introduce the log difference of the S&P500 Volatility Index (VIX) as a global variable (see Heinz and Sun (2014)), which is endogenous in the US model, but exogenous in other countries. For the empirical study we transform the data by taking the first differences (denoted by the first difference operator, Δ) of the log CDS series and the first differences of the log Euro exchange rates as well as the VIX, which is the standard approach in this literature (see, for example, Yalin Gündüz and Orcun Kaya (2014), Alter and Beyer (2014), Gross and Kok (2013) or Gray et al. (2013)).

Each country i in the GVAR is represented by a VAR model that consists of a constant ($a_{i,0}$), a vector of domestic and endogenous variables ($x_{i,t}$) as well as a vector of foreign and weakly exogenous variables ($x^*_{i,t}$)

$$x_{i,t} = a_{i,0} + \sum_{l=1}^{p_i} \Phi_{i,l} x_{i,t-l} + \sum_{m=0}^{q_i} \Lambda_{i,m} x^*_{i,t-m} + u_{i,t}. \tag{1}$$

We discuss the GVAR for a VARX(1,1) specification. Specifically, $a_{i,0}$ denotes a $k_i \times 1$ column vector of constants. k_i stands for the rank of the corresponding vector or matrix relating to the model of country i. We set the lag orders p_i and q_i according to the SBC for all countries $i = 1..N$ (see Table (2)).[4] $\Phi_{i,j}$ as well as $\Lambda_{i,k}$ are coefficient matrices of the vectors of domestic variables and foreign variables with dimension $k_i \times k_i$. $u_{it} = iid(0, \Sigma_{ui})$ denotes a vector of serially uncorrelated residuals and is of $k_i \times 1$ dimension.

In all EMU countries, the vector of domestic variables ($x_{i,t}$) which are all endogenous contains the sovereign ($CDS^{Sovereign}$) and bank (CDS^{Bank}) CDS spreads. The weighted averages of foreign variables enter the model through the vector of foreign (weakly exogenous) variables ($x^*_{i,t}$). Since EMU countries have no country-specific endogenous exchange rates, the aggregate of weighted foreign (non-EMU) exchange rates, which can be inter-

[4]We rely on the SBC criterion as the AIC is, in some cases (e.g. Ireland), allowing for too many lags, leading to a cyclical pattern in the impulse response functions. The SBC leads to a more parsimonious model while having no significant effects on the serial correlation of the error term.

Table 2: Lag orders

Country	p_i	q_i
Germany	1	1
France	1	1
Belgium	2	1
Ireland	1	1
Netherlands	1	1
Portugal	1	1
Spain	1	1
Austria	1	1
Italy	1	1
Japan	1	1
United Kingdom	1	1
United States	1	1

Note: Table shows the lag orders used in the country-specific VARX(p_i,q_i) models.

preted as the effective Euro ($\Delta s^*_{i,t}$), enters EMU models as exogenous variable as well. The VIX as measure of global risk appetite is also treated as exogenous:

$$x_{i,t} = \begin{bmatrix} \Delta CDS^{Sovereign}_{i,t} \\ \Delta CDS^{Bank}_{i,t} \end{bmatrix}, \quad x^*_{i,t} = \begin{bmatrix} \Delta CDS^{Sovereign^*}_{i,t} \\ \Delta CDS^{Bank^*}_{i,t} \\ \Delta s^*_{i,t} \\ \Delta vix_t \end{bmatrix}$$

In the United Kingdom and Japan, the first differences of log domestic sovereign and bank CDS spreads as well as the log exchange rate are endogenous. Here, the exchange rates are endogenous because the UK and Japan are assumed to be the main drivers of the EUR/GBP and EUR/JPY exchange rates, respectively. Therefore, only the first differences of the foreign sovereign and bank CDS spreads as well as the VIX appear in the vector of foreign variables:

$$x_{i,t} = \begin{bmatrix} \Delta CDS^{Sovereign}_{i,t} \\ \Delta CDS^{Bank}_{i,t} \\ \Delta s_{i,t} \end{bmatrix}, \quad x^*_{i,t} = \begin{bmatrix} \Delta CDS^{Sovereign^*}_{i,t} \\ \Delta CDS^{Bank^*}_{i,t} \\ \Delta vix_t \end{bmatrix}$$

In the US model, the first differences of log domestic sovereign and bank CDS spreads as well as the log exchange rate and the VIX are endogenous. As the US is the dominant financial center in the world, the VIX, a measure of US stock market volatility and proxy for global risk appetite, is endogenous here. Only the first differences of the foreign sovereign and bank CDS spreads appear in the vector of foreign variables:

$$x_{i,t} = \begin{bmatrix} \Delta CDS_{i,t}^{Sovereign} \\ \Delta CDS_{i,t}^{Bank} \\ \Delta s_{i,t} \\ \Delta vix_t \end{bmatrix}, \quad x_{i,t}^{*} = \begin{bmatrix} \Delta CDS_{i,t}^{Sovereign^*} \\ \Delta CDS_{i,t}^{Bank^*} \end{bmatrix}$$

The foreign aggregates are computed as

$$\begin{aligned} \Delta CDS_{i,t}^{Sovereign^*} &= \sum_{j=1}^{N} \omega_{i,j} \Delta CDS_{j,t}^{Sovereign}, \\ \Delta CDS_{i,t}^{Bank^*} &= \sum_{j=1}^{N} \omega_{i,j} \Delta CDS_{j,t}^{Bank}, \\ \Delta s_{i,t}^{*} &= \sum_{j=1}^{N} \omega_{i,j} \Delta s_{j,t}, \end{aligned}$$

where $\omega_{i,j}$ denotes the weight of country $j's$ variables in the aggregate of foreign variables in country i. Note here that $\omega_{i,i} = 0$ and $\sum_{j=1}^{N} \omega_{i,j} = 1$. In the GVAR literature, the weights $\omega_{i,j}$ are usually trade weights (see Dees et al. (2007) or Dees et al. (2007)). Alternatively, weights could be derived from the geographic distance (see Vansteenkiste (2007)), sectoral input-output tables (see Hiebert and Vansteenkiste (2010)), distances from fiscal fundamentals (see Favero (2013)), portfolio positions, FDI as well as banking claims (see Eickmeier and Ng (2015)). Gross and Kok (2013) take a different path and use weights that minimize the squared residuals of the unit-specific VAR models. Since our study focuses on dynamics of measures that are related to portfolio assets, we compute weights from the IMF Coordinated Portfolio Investment Survey (CPIS)[5] data and average

[5]The CPIS survey contains information on the bilateral holdings of portfolio investment securities for

the portfolio positions over the years 2008-2011 (see Table (3)) so that the full estimation period is covered.[6] The columns in the table sum up to 1 and represent the shares of foreign portfolio assets held by investors in the corresponding country. This implies, for example, that spillovers from the United States are strong in Japan, as the share of US assets in Japanese portfolios is relatively high (0.55).

Table 3: Weight matrix: IMF Coordinated Portfolio Investment Survey

Countries	DE	FR	BE	IE	NL	PT	ES	AT	IT	JP	UK	US
DE	0.00	0.15	0.10	0.09	0.18	0.09	0.12	0.32	0.15	0.10	0.11	0.11
FR	0.18	0.00	0.24	0.08	0.15	0.11	0.19	0.13	0.22	0.09	0.10	0.13
BE	0.02	0.05	0.00	0.01	0.02	0.02	0.03	0.02	0.02	0.01	0.01	0.01
IE	0.09	0.06	0.08	0.00	0.05	0.23	0.07	0.06	0.13	0.03	0.09	0.05
NL	0.14	0.15	0.20	0.05	0.00	0.09	0.14	0.10	0.12	0.05	0.09	0.09
PT	0.02	0.02	0.02	0.04	0.01	0.00	0.05	0.01	0.01	0.00	0.00	0.00
ES	0.11	0.11	0.08	0.03	0.05	0.15	0.00	0.05	0.07	0.02	0.03	0.04
AT	0.06	0.03	0.03	0.01	0.03	0.01	0.02	0.00	0.03	0.01	0.01	0.01
IT	0.12	0.15	0.09	0.08	0.06	0.12	0.19	0.10	0.00	0.03	0.06	0.03
JP	0.01	0.06	0.00	0.03	0.02	0.00	0.00	0.01	0.01	0.00	0.08	0.18
UK	0.12	0.12	0.07	0.23	0.11	0.09	0.09	0.09	0.11	0.10	0.00	0.36
US	0.12	0.10	0.09	0.36	0.30	0.08	0.10	0.11	0.15	0.55	0.41	0.00

Note: Table shows lenders in columns and debtors in rows. Sum of each column equals 1.

We solve the GVAR by transforming (1) to

$$A_{i,0} z_{i,t} = a_{i,0} + A_{i,1} z_{i,t-1} + u_{i,t}, \tag{2}$$

where

$$z_{i,t} = (x_{i,t}, x^*_{i,t})', \quad A_{i,0} = (I_{k_i}, -\Lambda_{i,0}), \quad A_{i,1} = (\Phi_{i,1}, \Lambda_{i,1}).$$

Now, the vector $z_{i,t}$ may be written as $W_i x_t$, the product of a suitable weight matrix, W_i, and the *global* vector $x_t = (x'_{1,t}, x'_{2,t}, ..., x'_{N,t})$. This leads to the GVAR equation

$$A_{i,0} W_i x_t = a_{i,0} + A_{i,1} W_i x_{t-1} + u_t, \tag{3}$$

a large number of countries.

[6] We also estimated the model using identical weights for all countries and did not observe major differences in the results based on estimations using the CPIS weights.

which can be simplified to

$$G_0 x_t = b_0 + G_1 x_{t-1} + c_t, \tag{4}$$

where

$$b_0 = \begin{pmatrix} a_{1,0} \\ a_{2,0} \\ \vdots \\ a_{N,0} \end{pmatrix}, \quad b_1 = \begin{pmatrix} a_{1,1} \\ a_{2,1} \\ \vdots \\ a_{N,1} \end{pmatrix}, \quad c_t = \begin{pmatrix} u_{1,t} \\ u_{2,t} \\ \vdots \\ u_{N,t} \end{pmatrix}$$

and

$$G_0 = \begin{pmatrix} A_{1,0}W_1 \\ A_{2,0}W_2 \\ \vdots \\ A_{N,0}W_N \end{pmatrix}, \quad G_1 = \begin{pmatrix} A_{1,1}W_1 \\ A_{2,1}W_2 \\ \vdots \\ A_{N,1}W_N \end{pmatrix}.$$

Premultiplying (4) by G_0^{-1} yields

$$x_t = f_0 + F_1 x_{t-1} + \epsilon_t, \tag{5}$$

where

$$f_0 = G_0^{-1} b_0, F_1 = G_0^{-1} G_1, \epsilon_t = G_0^{-1} c_t.$$

2.2 Shock identification

To simulate economic shocks in a vector autoregressive model, one can employ different methodologies. The GVAR literature often relies on generalized impulse responses (see Pesaran and Shin (1998)), which are invariant to the ordering of the variables in the system. These estimated impulse responses can be thought of as the statistically most likely response following a shock to a variable in the system. Consequently, it is not possible to give the impulse responses a structural interpretation. In order to do so, the shocks have to be orthogonal. Shocks can, for example, be orthogonalized by a Cholesky scheme (see Sims (1980)), which yields a triangular variance-covariance matrix. This, however, requires the assumption that some shocks have no contemporaneous effect on certain variables. In a system of financial variables, this assumption is unlikely to hold, because financial variables usually respond quickly to changes in other variables. This problem can be circumvented by imposing restrictions on either the sign of the conditional correlation (see Hau and Rey (2004)) or the sign of the impulse responses directly (see Uhlig (2005)). The advantages of the latter approaches are that restrictions can be derived from economic models and that no zero restrictions on the correlation structure of shocks are required.

Since the underlying model by Acharya et al. (2014) provides us with information about the expected signs of impulse responses, we follow the latter approach and identify the shocks by imposing sign restrictions on impulse response functions of country-specific models in the GVAR (see Eickmeier and Ng (2015)). For the estimation of our model, we use a modified version of the GVAR Toolbox by Smith and Galesi (2011).

2.2.1 Theory

In our model, we identify two different shocks using sign restrictions (see Table (4)). The sign restrictions have previously been employed by Stanga (2014) for single country models in a different setting. According to the model by Acharya et al. (2014), the philosophy behind the bailout shock is that a bailout is characterized by a risk transfer from the bank balance sheets to the public sector balance sheets, leading to a decline in the CDS spreads of banks and an increase in the CDS spread of the sovereign.[7]

[7] What this identification strategy does not capture are potential unidirectional responses following a bailout shock. If a bailout improves economic agents' expectations regarding future economic developments so that sovereign CDS spreads decrease together with bank CDS spreads, then this would be captured as a general risk shock. However, we show in section (2.3) that the identified shocks appear at

By contrast, bank risk shocks as well as sovereign risk shocks are expected to raise sovereign CDS spreads as well as bank CDS spreads. While a bank risk shock affects the sovereign CDS spread by reinforcing the expectation of future financial distress and economic contraction, a sovereign risk shock affects the bank CDS via sovereign bonds in bank portfolios. Therefore, we do not disentangle sovereign risk from bank risk, but rather consider a general risk shock, which is characterized by a contemporaneous increase in bank CDS and sovereign CDS. General risk shocks also capture most macro shocks such as business cycle shocks, for example.

The advantage of this identification strategy is that it is very agnostic. Bailouts do not necessarily affect markets on the announcement date. In certain cases, the expectation of a bailout may affect the markets more than the bailout itself (see Stanga (2014)).

We identify these shocks for Germany, France and Italy, the three largest EMU economies in terms of GDP, as well as Ireland, an EMU country which was severely hit by the banking crisis and where bailouts were relatively large. An additional reason for this specific selection is that in some of the other countries, no bailouts took place during the sample period.

Since EMU models have only two endogenous variables, the identification of two shocks is straightforward as it requires fewer assumptions than Stanga (2014), who identifies three shocks, forcing the imposition of additional restrictions on the time horizon of impulse responses. Having the same number of variables in each model is also important for our identification strategy, which we explain in the following section.

Table 4: Shock profiles

Shock	$\Delta\ CDS^{Sovereign}$	$\Delta\ CDS^{Bank}$
Bailout	+	-
General risk	+	+

Note: Table reports the required sign restriction profiles for each shock.

times of bailout announcements. Therefore, we are confident that bailout shocks are correctly identified.

2.2.2 Technical implementation

In order to identify the country-specific structural shocks, we use the reduced form residuals from each country model i to compute lower triangular Cholesky matrices P_i, which are then used to derive the matrix

$$P = \begin{pmatrix} P_0 & 0 & \dots & \dots & 0 \\ 0 & \ddots & & & \vdots \\ \vdots & & P_i & & \vdots \\ \vdots & & & \ddots & 0 \\ 0 & \dots & \dots & 0 & P_N \end{pmatrix}.$$

Given the P matrix and the vector moving average representation of the GVAR, ϕ^h, we obtain the impulse responses $\psi^h = \phi^h G_0^{-1} P$. Here, h denotes the horizon.

The traditional approach in the sign restrictions literature (see, for example, Eickmeier and Ng (2015), Georgiadis (2015)) is to draw random country-specific $k_i \times k_i$ orthonormal matrices. QR decompositions of these matrices then provide unique weight matrices Q_i which satisfy $Q_i Q_i' = I$. The next step is to test if the impulse response functions $\Psi_i^h = (\psi_i^h Q_i')'$ satisfy the sign restrictions

$$\Psi_i^0 = \begin{bmatrix} \geq 0 & \leq 0 \\ \geq 0 & \geq 0 \end{bmatrix}.$$

Having obtained a set of Q_i matrices which produce impulse response functions that satisfy the sign restrictions, it is possible to plot location parameters or confidence bands. However, Fry and Pagan (2007) criticize that these impulse responses may relate to different data-generating processes. They propose to select the Q_i, which produces impulse responses that are as close as possible to the median over all variables, which we call M_i. This so-called median target can then be used for bootstraps and forecast error variance decompositions in the traditional way.

However, by performing this computation for every country in a Global VAR individually, it is likely that M_i differs across countries. We therefore propose a novel approach by rotating Q_i and testing if it yields impulse responses that satisfy the sign restrictions in every single country under investigation. For this approach it is important to note that all country models must be modeled using exactly the same variables. Afterwards, we follow Fry and Pagan (2007) and select the Q_i, which produces impulse responses which are as close as possible to the median over all variables and, according to our novel approach for the GVAR model also over all countries. This procedure provides us with one unique M, which can be applied to every single country. The advantage of having one unique M is that the same weight is applied to every country, making the impulse responses easier to compare.

2.3 Shock analysis

As stated in section (2.2.1), we perform the identification procedure for four countries, namely Germany, France, Italy and Ireland.[8]

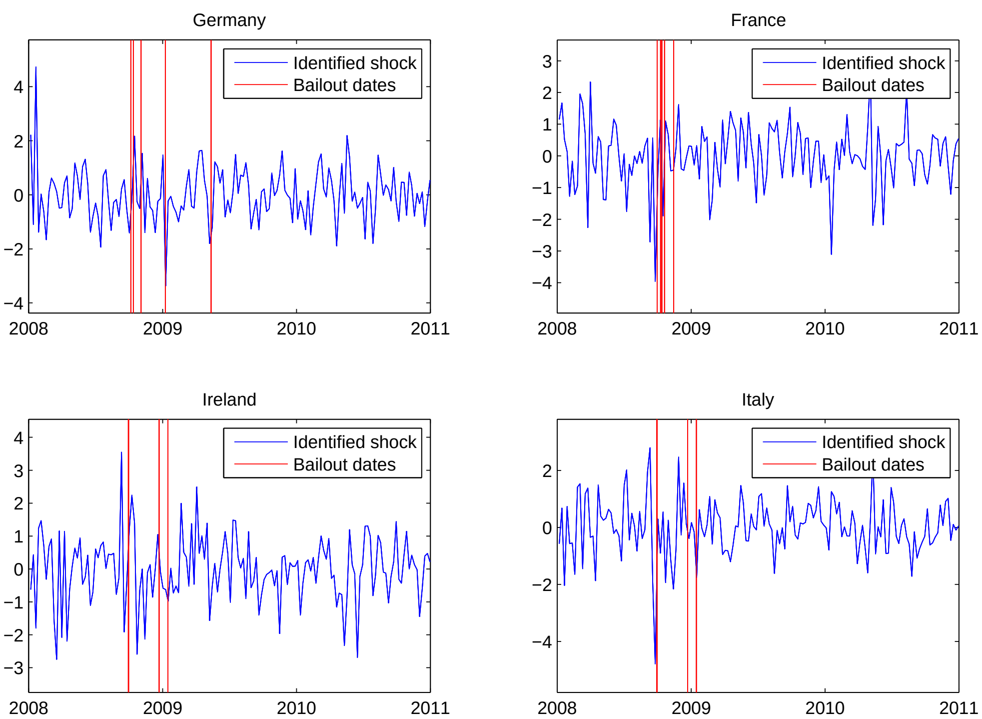
Our identification strategy with a block-diagonal P-matrix relies on the assumption that the cross-section correlation of residuals is low. Although the matrix G_0 captures the contemporaneous effects across countries, there may still be cross-section correlation left. In such a case, the interpretation of identified shocks as country-specific would be problematic. However, the average absolute pairwise cross-section correlation over all variables is 0.08. The absolute average pairwise cross-section correlation between identified shocks is 0.14 for general risk shocks, as well as for bailout shocks. We are, therefore, comfortable with interpreting the identified shocks as country-specific.

Since shocks can be interpreted as country-specific, we may now investigate if the identified structural shocks correspond to actual bailout dates. We have already highlighted that the applied identification strategy is very agnostic in the way we identify bailout shocks as shocks that drive sovereign CDS spreads up and bank CDS spreads down. The

[8]For a detailed list of bailout dates, we refer the reader to Stanga (2014).

advantage of this strategy is that we identify a shock exactly when market participants are shocked. This is not necessarily when the announcement is made. However, it should be close by. Therefore, we compare the country-specific shocks with the actual announcement dates. Figure (1) demonstrates that the bailout dates correspond to negative peaks in the corresponding bailout shock series. Particularly in the shock series of France, Ireland and Italy, we observe large negative peaks shortly before the actual announcement, indicating that market participants anticipated the bailout and that markets had already reacted before the announcement. This is in line with findings by Bernal, Oosterlinck, and Szafarz (2010) and Dam and Koetter (2012), who show that markets reacted before the announcement. Moreover, it shows that the applied agnostic identification is — despite its drawbacks — a reasonable choice, as it may capture shocks better than an event study based on announcement dates.

Figure 1: Structural bailout shocks and bailout dates



Note: Figure shows identified structural bailout shocks (blue) together with the actual bailout dates (red lines).

3 Results

In order to investigate the responses of variables to identified shocks, we compute impulse response functions and discuss the impact of bailout and general risk shocks in terms of statistical significance based on a bootstrap procedure with 1,000 runs. Then we employ a forecast error variance decomposition to analyze by how much the variability of variables is explained by specific shocks.

3.1 Impulse response analysis

Figures (2)-(7), which concern impulse response functions, show the 90% confidence intervals of cumulated impulse response functions five weeks following the shocks. We present the cumulated effects for the five weeks following the shocks, as this is the period during which the effects generally have their maximum impact (see Figures (8)-(11)). Bailout and general risk shocks are normalized in the way that the sovereign CDS spread increases by one standard deviation.

3.1.1 Spillover effects of structural risk shocks

First, we investigate the international propagation of bailout shocks. Figure (2) displays spillover effects of German, French, Irish and Italian bailout shocks to CDS spreads in all modeled countries. The shocks always affect domestic CDS spreads the most, which is intuitive, as the shock is idiosyncratic. However, spillover effects of the German shock are relatively strong, particularly for Austria. The strong effect in Austria may be related to the relatively high share of German assets in Austrian portfolios as documented in the IMF Coordinated Portfolio Investment Survey (see Table (3)). For Italy, we observe a very similar pattern. But in France and Ireland, spillover effects to sovereign CDS spreads are negligible. Interestingly, we observe for all countries that spillover effects to

EMU countries appear to be stronger than those to non-EMU countries, which could imply that market participants anticipate a common euro-area-wide rescue scheme for stressed countries. In the following section, we will discuss this heterogeneity in depth by testing if this difference is statistically significant.

Spillover effects of bailout shocks to bank CDS spreads appear to be generally stronger. They are significant in all countries following French and Irish shocks. However, for both countries we observe negative effects to Japan, and for the Irish shock also to the US. An explanation for this pattern would be that the bailouts within the EMU lowered the credit default risk of foreign (non-EMU) banks and thus the necessity of bailouts in non-EMU countries. At the same time market participants may have anticipated a common euro-area-wide rescue scheme. German shocks have significant spillover effects in France, the Netherlands, the United Kingdom and the United States. Italian shocks have no statistically significant effects at all in foreign countries. This reflects strong linkages between banks in the euro-area interbank market.

Second, we analyze the spillover effects of general risk shocks on sovereign CDS spreads (see Figure (5)). Overall, we observe statistically significant spillover effects in all countries following shocks from all countries under investigation. The international effects of general risk shocks on bank CDS spreads are very similar. These findings are supported by Alter and Beyer (2014), Gray et al. (2013), and Gross and Kok (2013). However, Figure (6) also shows that spillover effects are significant in all countries, while non-EMU countries appear to be less affected. In the following section, we will test if this new insight is statistically significant.

3.1.2 Heterogeneity in spillover effects

In the previous section, we observed heterogeneity in spillovers across countries. Specifically, we observed that the spillover effects to sovereign CDS spreads of the identified shocks appeared to be much stronger in EMU countries than in non-EMU countries. To test if these differences are statistically significant, we compute the differences of averaged impulse responses for different control groups. For every test we normalize the shock in the way that the variable under investigation shows a positive response. Then, we compute the difference between the average EMU country response and the average non-EMU country response for every bootstrap. This procedure enables us to compute confidence intervals for these differenced impulse response functions.

Figure (8) displays these confidence bands for a general risk shock to sovereign CDS spreads. Given the normalization, a significantly positive response implies that sovereign CDS spreads of EMU countries respond, on average, significantly more strongly to a general risk shock than non-EMU countries. We find evidence for this effect following shocks originating in all four countries. However, the effect becomes insignificant after approximately three weeks. The effects of general risk shocks to bank CDS spreads are slightly different. For Germany, France and Italy, Figure (9) reveals that non-EMU countries respond more strongly on impact. However, after about three to four weeks, the impulse response functions become significantly positive for all shocks.

Bailout shocks have slightly different effects than general risk shocks. While our measure is low on impact, indicating no major differences between responses of sovereign CDS spreads in EMU and non-EMU countries, it is significantly positive in the long run (see Figure (10)). Except for Germany, where our measure is only significant over the first four weeks following the shock, we find evidence for permanently stronger effects on bank CDS spreads in EMU countries following shocks originating in France, Ireland and Italy (see Figure (11)).

Overall, we observe that the differences are statistically significant for every single country. Consequently, spillover effects of EMU shocks to other EMU countries are (on average) significantly larger than spillover effects to non-EMU countries. This result may imply that market participants anticipated EMU-wide bailout actions and may also reflect strong relationships between banks in the EMU.

3.1.3 Effects of structural risk shocks on the Euro

We find that general risk shocks originating in any country under investigation translate into a significant depreciation of the Euro-Yen and Euro-US Dollar exchange rates (see Figure (7)) in line with findings by Della Corte et al. (2014). The effects on the Yen are slightly stronger than those on the US Dollar. However, we do not find any significant effects on the Pound Sterling.

Bailouts have much less pronounced effects on the Euro (see Figure (4)). However, the Euro depreciates significantly against the US Dollar following bailout shocks originating in any country. German and Italian shocks seem to have the strongest effects on the Euro-Dollar exchange rate. Only the Irish bailout shock has a slightly significant effect on the Euro-Sterling exchange rate. This effect can be explained by the strong economic ties between both countries.

Our results suggest that for the Euro-Yen and Euro-Dollar exchange rates, general risk shocks are of particular importance. Bailout shocks mainly affect the Euro-Dollar exchange rate, but to a much smaller extent than general risk shocks. However, the effect may be large because the variance of the corresponding variables is large. Therefore, we decompose the forecast error variance in the following section.

3.2 Generalized forecast error variance decomposition

We now analyze the forecast error variance of Euro exchange rate pairs explained by risk shocks originating in the euro area. To this end, we will first focus on the international role of total EMU shocks and take a look at country-specific EMU shocks in a second stage. Since our identified shocks are only orthogonal within the country of origin, but weakly correlated across countries, we rely on a generalized forecast error variance composition (see Pesaran and Shin (1998)). In this case, the forecast error variance of variables does not add up to exactly 1 (100%).[9] In order to achieve comparable variance shares, we scale the shares so that they do add up to 1.

[9] In fact, the explained forecast error variance adds up to values very close to 1 (e.g. approx. 0.97-0.99).

Table (5) displays the shares of explained exchange rate variability by aggregate EMU shocks and non-EMU shocks. Overall, EMU shocks explain approximately 25% of the Euro-Dollar variability, 21% of the Euro-Yen variability, but only 4% of the Euro-Sterling variability. The table also shows that bilateral exchange rates may be substantially driven by shocks originating in third countries. In particular, a large fraction of the Euro-Yen variability (59%) is explained by foreign shocks. But also for the Euro-Sterling and Euro-Dollar exchange rates, third-country shocks account for approximately 12% and 22% of the exchange rate variabilities, respectively. This is surprising, as the exchange rate is supposed to reflect the relative price between two economies. Our finding, however, relates to the so-called *exchange-rate disconnect puzzle* (see Obstfeld and Rogoff (2000)). It states that the ability of domestic macroeconomic fundamentals to forecast bilateral exchange rates is often far below what one would expect according to traditional exchange rate models (see Engel, Mark, and West (2007)). In this context, Engel, Mark, and West (2015), and Hodrick and Vassalou (2002) show that factors of third-country variables or multi-country models may improve bilateral exchange rate forecasts. As possible reasons for the importance of third countries, Berg and Mark (2015) identify monetary policy rules taking exchange rates or interest rates of third countries into account, as well as differences in price stickiness across countries.

Table 5: GFEVD: International role of shocks

Exchange rate	EMU	Japan	UK	US	Sum
Euro-Yen	0.21	0.19	0.12	0.47	1.00
Euro-Sterling	0.04	0.01	0.84	0.11	1.00
Euro-Dollar	0.25	0.07	0.15	0.53	1.00

Note: Table reports the exchange-rate-specific forecast error variance explained by shocks originating in different regions and countries respectively.

We can now analyze the importance of the identified country-specific shocks. Table (6) displays the shares of explained forecast error variance by country-specific general risk and bailout shocks for the Euro-Yen, Euro-Sterling and Euro-Dollar exchange rates. The total share of the identified shocks is in all cases very high, meaning that the identified shocks in the four EMU countries are the most important contributors to the Euro variability within the EMU. An unambiguous result is that the identified bailout shocks are not important drivers of exchange rate fluctuations. However, general risk shocks are of importance. Particularly Germany and France, the two (in terms of GDP) largest economies, play a role in this context. General risk shocks from both countries explain at least 5% of the Euro-Yen and Euro-Dollar variability. Although EMU shocks generally do not explain much of the Euro-Sterling variability, the shocks stemming from the latter two countries are relatively important. Ireland and Italy are slightly less important than Germany and France, although Ireland appears to be slightly more important than Italy for all exchange rate pairs.

The variance decompositions confirm our previous findings that shocks from countries with a high GDP are important drivers of the Euro. However, Ireland, one of the smallest, but most distressed countries in our sample, plays an important role as well.[10]

Table 6: GFEVD: Importance of country-specific shocks

Country	Shock	Euro-Yen	Euro-Sterling	Euro-Dollar
Germany	General risk	0.05	0.01	0.05
	Bailout	0.00	0.00	0.00
France	General risk	0.05	0.01	0.06
	Bailout	0.00	0.00	0.00
Ireland	General risk	0.04	0.01	0.04
	Bailout	0.00	0.00	0.00
Italy	General risk	0.03	0.00	0.03
	Bailout	0.00	0.00	0.00
	Total	0.17	0.03	0.18

Note: Table reports the exchange-rate-specific forecast error variances explained by specific shocks originating in EMU countries.

[10]It is likely that Greece, for example, is also an important contributor to the Euro exchange rate variability. However, due to the Greek default, data on CDS spreads are not available over the whole sample period.

4 Conclusion

During the 2008 financial crisis, risk spillovers became a major concern for policy makers and banks. The changes in sovereign and bank risk are believed to have influenced world-wide exchange rates. We investigate these dynamics empirically and estimate a Global VAR model for nine EMU countries and Japan, the United Kingdom as well as the United States in which we identify general risk shocks as well as bailout shocks.

The results from our model suggest that spillover effects of general risk shocks to sovereign and bank CDS spreads are strong and significant. For bailouts, however, spillovers are much less pronounced and mostly insignificant. In this context we observe that spillovers are significantly stronger within the EMU, indicating that risk spills over strongly across EMU sovereigns and banks.

Regarding the effects on the Euro, our results indicate that general risk shocks cause significant depreciations of the Euro-Yen and Euro-Dollar exchange rates, while bailout shocks affect only the Euro-Dollar exchange rate significantly. In this context, economically (in terms of GDP) large and also distressed countries appear to be the most important drivers. Interestingly, the global model reveals that shocks stemming from third countries are important sources of exchange rate fluctuations as well.

Consequently, risk spillovers are an important factor for country-specific risk, particularly within the EMU. In this way, country-specific bailouts during the financial crisis helped — to a certain extent — to mitigate euro-area-wide credit default risk. According to our estimates, the risk increases also put relatively strong pressure on the Euro exchange rate. However, not all exchange rate pairs were affected equally.

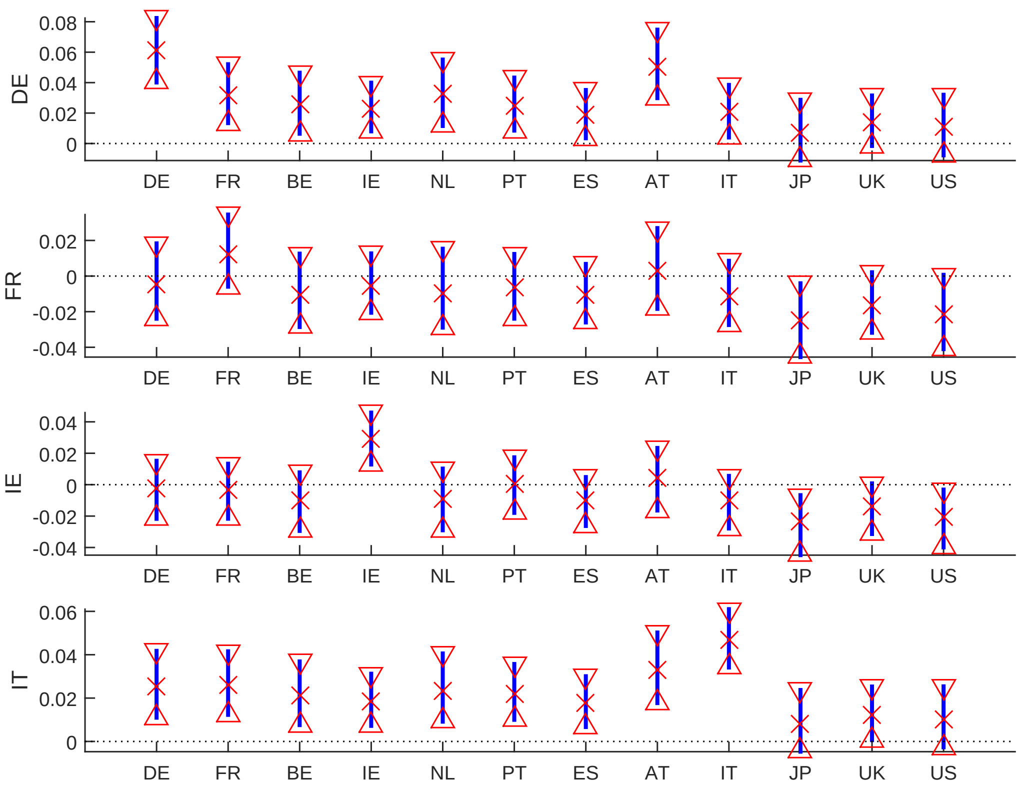
Appendix: Tables and Figures

Table 7: Composition of Bank CDS series

Country	Bank	Weight
Austria	Raiffaisen	1
Belgium	KBC Bank	1
Germany	Deutsche Bank	0.33
	Commerzbank	0.33
	Bayerische Landesbank	0.33
Spain	Banco Bilbao	0.5
	Santander	0.5
France	BNP Paribas	0.33
	Crédit Agricole	0.33
	Sociéte Générale	0.33
Italy	Banca Monte dei Paschi di Siena	0.33
	Intensa Sanpaolo	0.33
	Banco Popolare	0.33
Ireland	Bank of Ireland	1
Portugal	Banco Espirito Santo	0.5
	Banco Comercial Portugues	0.5
Netherlands	ING	0.5
	Rabobank	0.5
United Kingdom	Barclays	0.5
	Royal Bank of Scotland	0.5
Japan	Nomura	0.25
	Norinchukin	0.25
	Mizuho Bank	0.25
	Mitsui Sumitomo	0.25
United States	Bank of America	0.25
	JP Morgan Chase & Co	0.25
	Citigroup	0.25
	Goldman Sachs	0.25

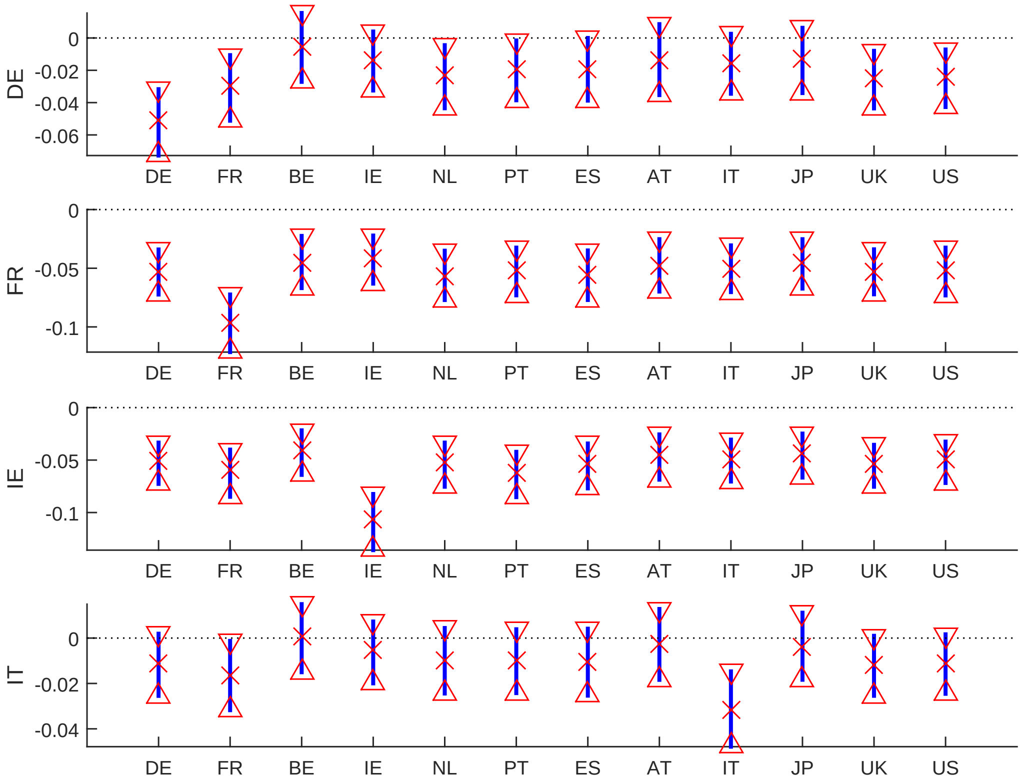
Note: Table reports the banks and bank-specific weights that have been used to compute the country-specific bank CDS series.

Figure 2: Effects of bailout shocks on sovereign CDS spreads



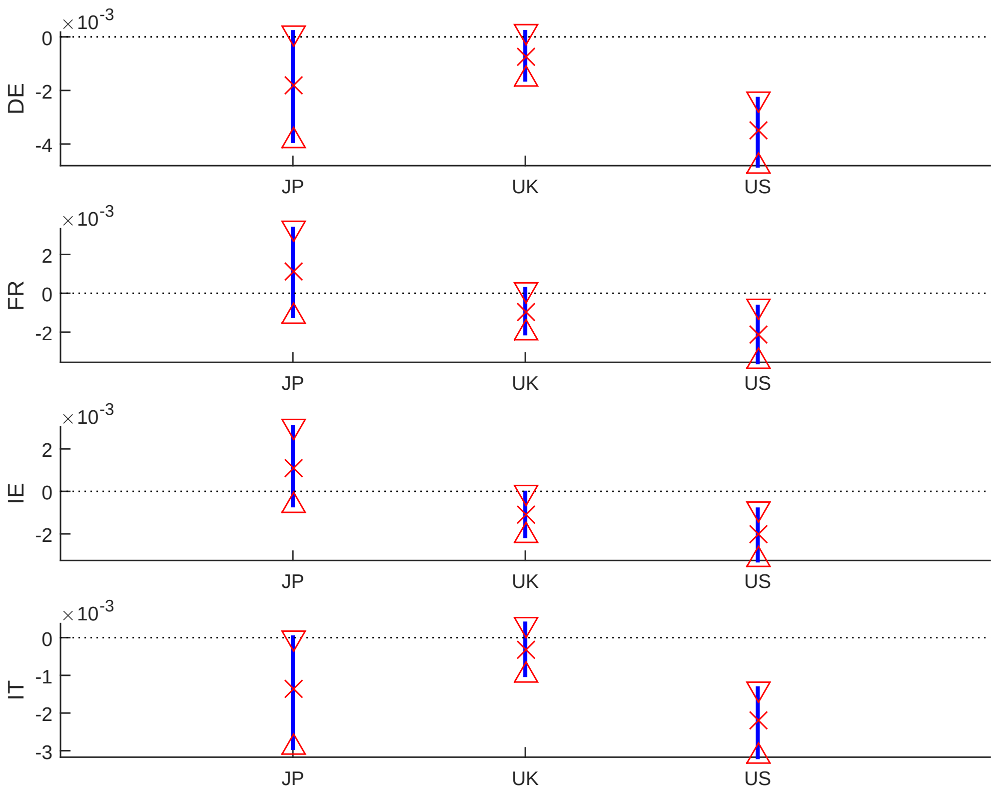
Note: Figures show 90% confidence bands (blue lines) and medians (red crosses) of bootstrapped IRFs (1,000 runs) 5 weeks following the shock.

Figure 3: Effects of bailout shocks on bank CDS spreads



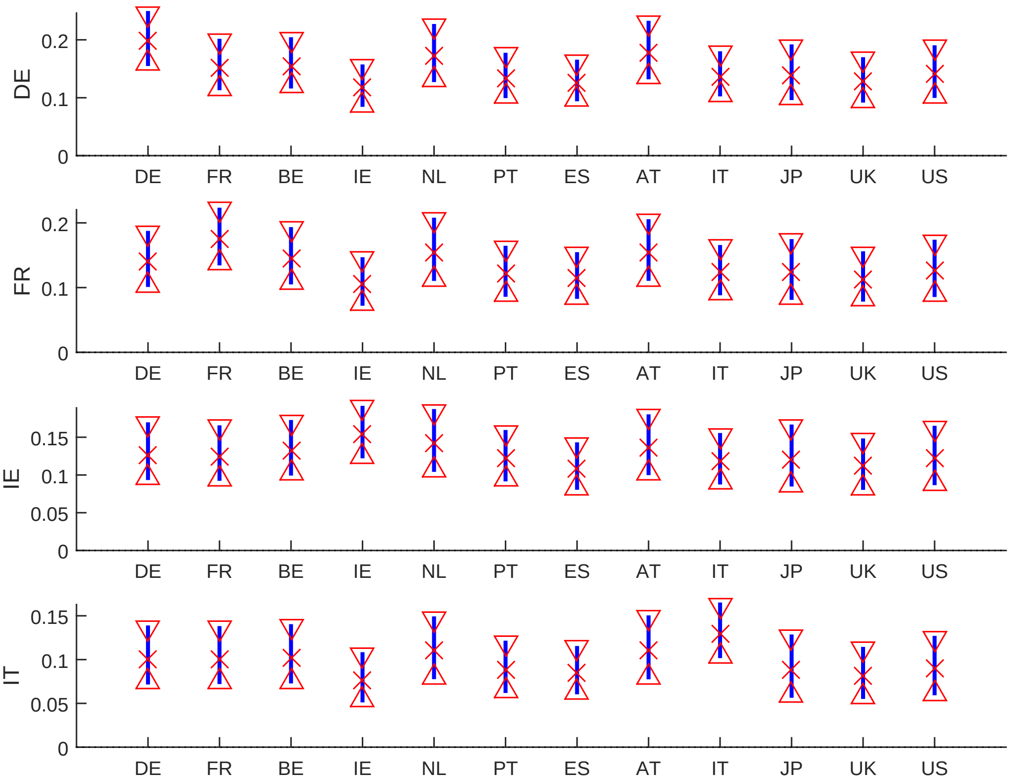
Note: Figures show 90% confidence bands (blue lines) and medians (red crosses) of bootstrapped IRFs (1,000 runs) 5 weeks following the shock.

Figure 4: Effects of bailout shocks on the Euro



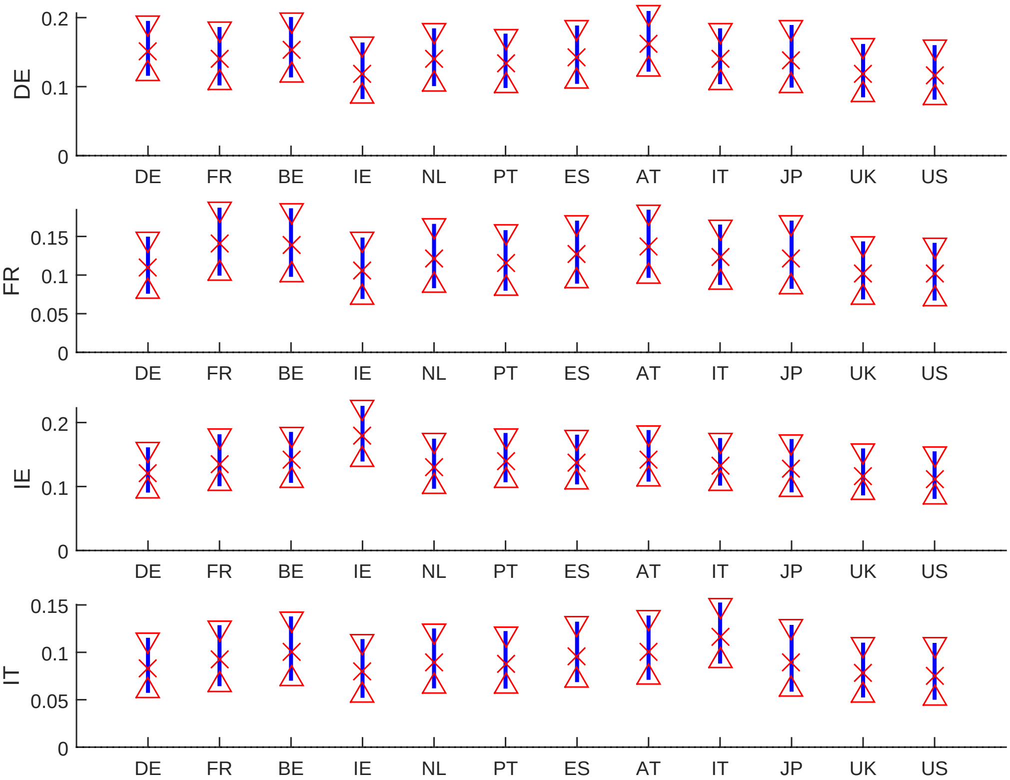
Note: Figures show 90% confidence bands (blue lines) and medians (red crosses) of bootstrapped IRFs (1,000 runs) 5 weeks following the shock.

Figure 5: Effects of general risk shocks on sovereign CDS spreads



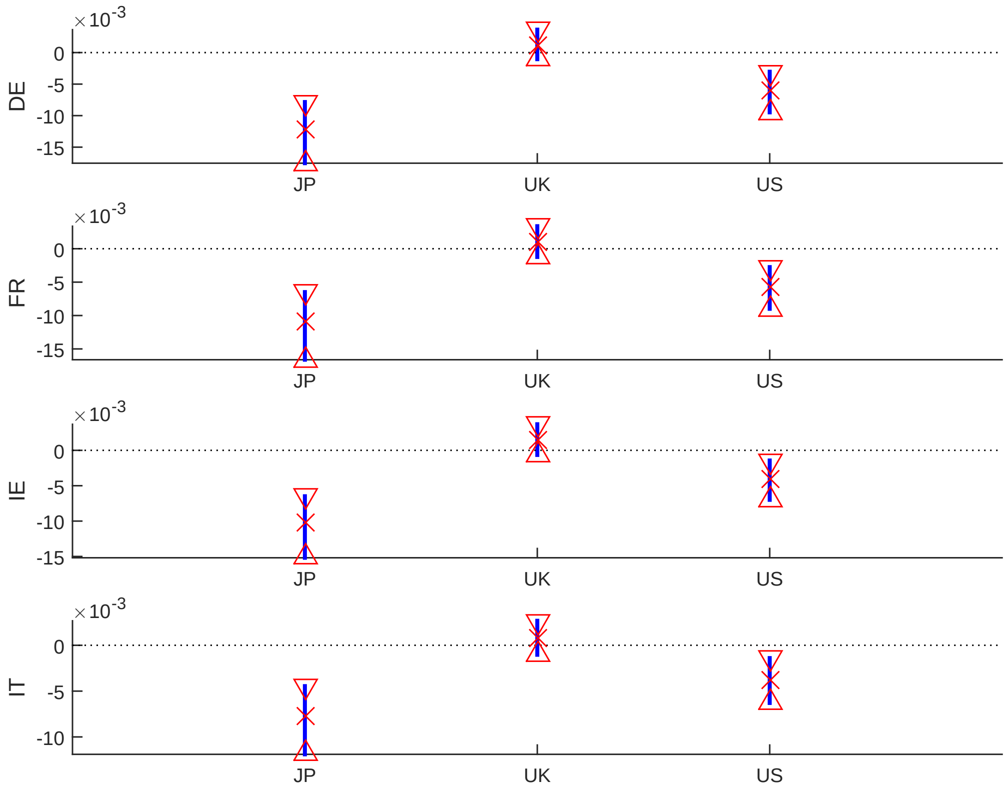
Note: Figures show 90% confidence bands (blue lines) and medians (red crosses) of bootstrapped IRFs (1,000 runs) 5 weeks following the shock.

Figure 6: Effects of general risk shocks on bank CDS spreads



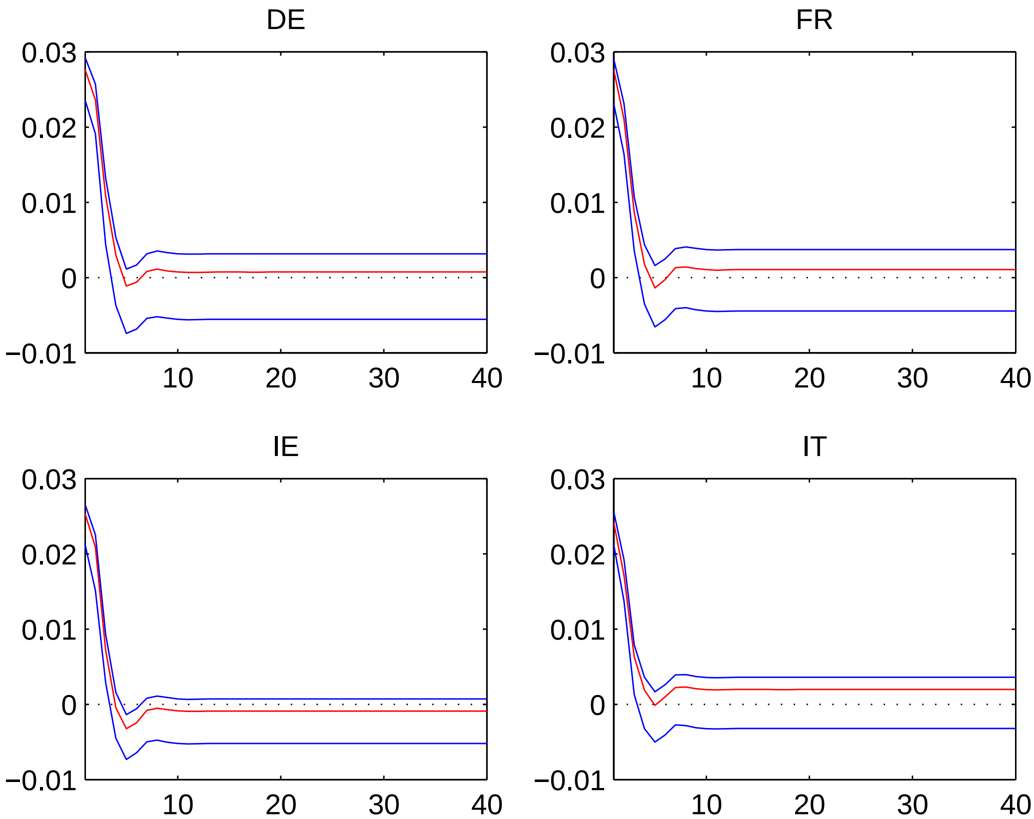
Note: Figures show 90% confidence bands (blue lines) and medians (red crosses) of bootstrapped IRFs (1,000 runs) 5 weeks following the shock.

Figure 7: Effects of general risk shocks on the Euro



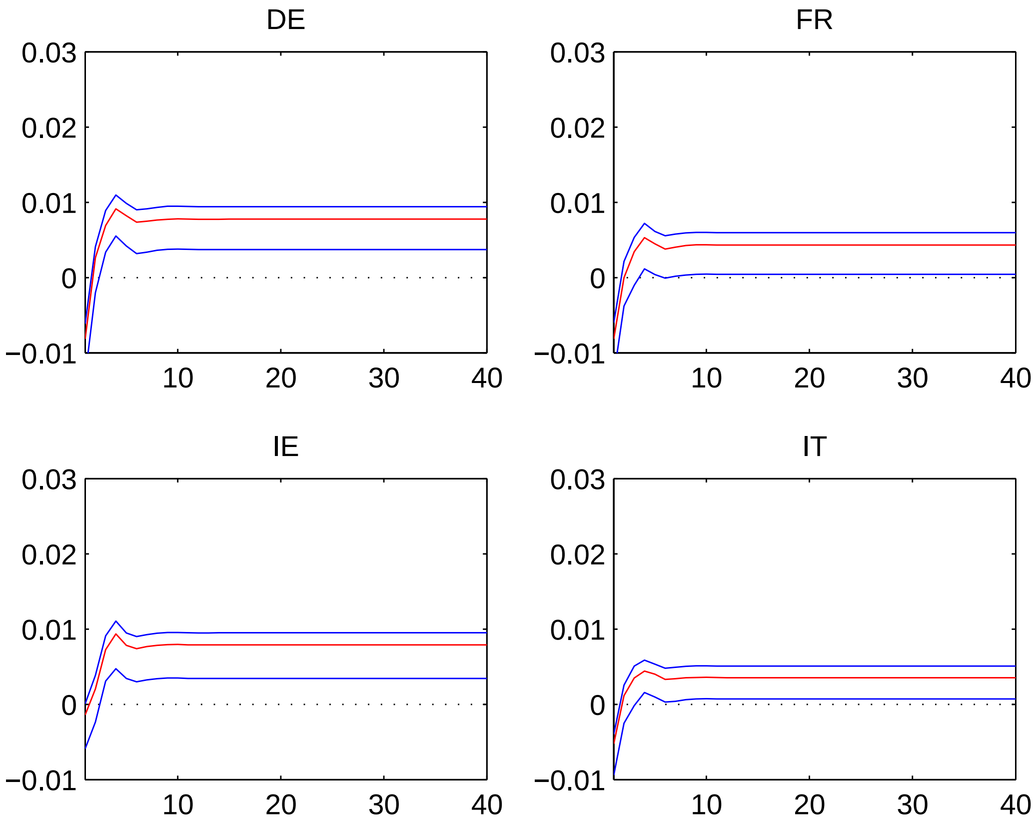
Note: Figures show 90% confidence bands (blue lines) and medians (red crosses) of bootstrapped IRFs (1,000 runs) 5 weeks following the shock.

Figure 8: Differences in averaged sovereign CDS spread impulse responses of EMU and non-EMU countries to a general risk shock



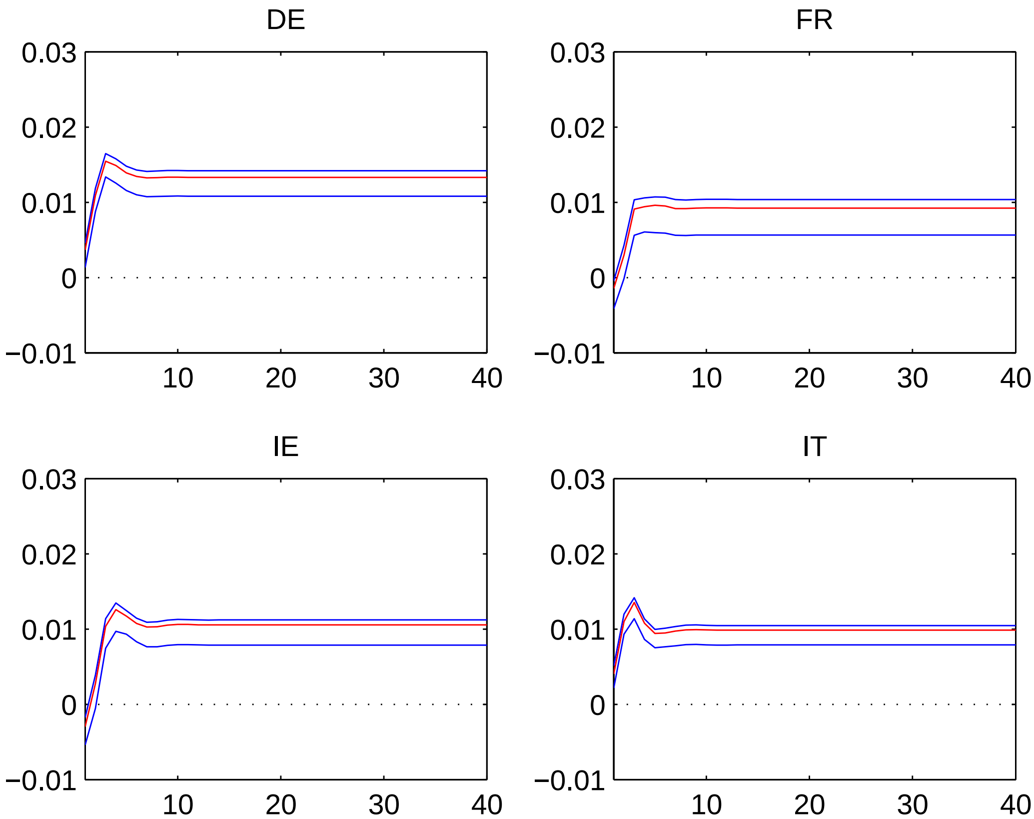
Note: Figures show 90% confidence bands (blue lines) and medians (red lines) of differences in bootstrapped IRFs (1,000 runs) between averages of EMU and non-EMU countries.

Figure 9: Differences in averaged bank CDS spread impulse responses of EMU and non-EMU countries to a general risk shock



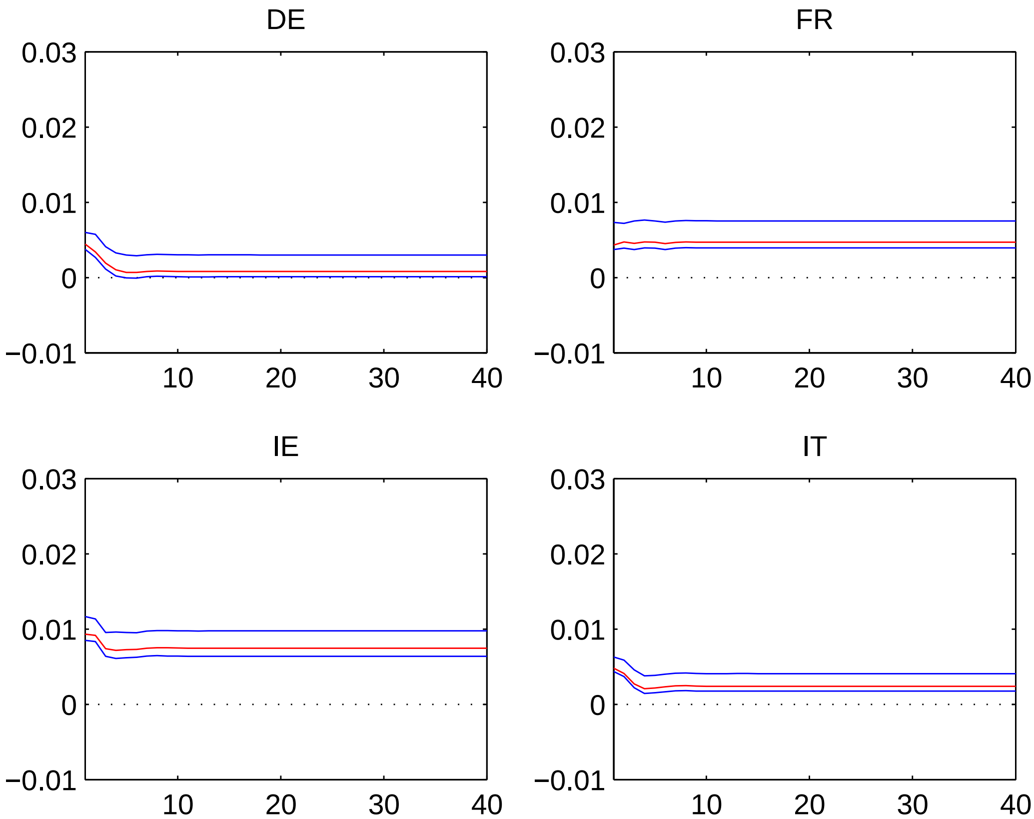
Note: Figures show 90% confidence bands (blue lines) and medians (red lines) of differences in bootstrapped IRFs (1,000 runs) between averages of EMU and non-EMU countries.

Figure 10: Differences in averaged sovereign CDS spread impulse responses of EMU and non-EMU countries to a bailout shock



Note: Figures show 90% confidence bands (blue lines) and medians (red lines) of differences in bootstrapped IRFs (1,000 runs) between averages of EMU and non-EMU countries.

Figure 11: Differences in averaged bank CDS spread impulse responses of EMU and non-EMU countries to a bailout shock



Note: Figures show 90% confidence bands (blue lines) and medians (red lines) of differences in bootstrapped IRFs (1,000 runs) between averages of EMU and non-EMU countries.

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