On the International Spillover Effects of Country-Specific Financial Sector Bailouts and Sovereign Risk Shocks*

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December 31, 2017

Abstract

We use sign-identified macroeconomic models to study the interaction of financial sector and sovereign credit risks in Europe. We show that structurally-identified national financial sector bailout shocks which transfer private sector risk onto the local sovereign generate negligible international spillovers. This is because the international benefits accruing from the reduction in financial sector risk are offset by the increase in sovereign risk. By contrast, structurally-identified sovereign risk shocks generate detrimental spillovers for the global financial sector and for international sovereign debt markets. We conclude that any bailout policy which calls the creditworthiness of the responsible sovereign into question is likely to exacerbate global credit risk.

JEL Codes: Financial sector bailouts; sovereign risk shocks; international spillovers; structural shocks; sign restrictions.

Keywords: C58, E61, F42, G01.

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We are grateful for illuminating discussions with Renée Fry-McKibbin, Stephen Kirchner, Leo Krippner, Yongcheol Shin, Artur Tarassow and Ben Wong. We acknowledge helpful comments from seminar participants at the University of Utah (March 2017) and the Reserve Bank of New Zealand (May 2017) and from participants at the Modelling Macroeconomic Shocks Workshop at the University of Tasmania (Hobart, October 2017) and the 2017 Sydney Banking and Financial Stability Conference (University of Sydney, December 2017). We use code from version 2.0 of Ambrogio Cesa-Bianchi's VAR Toolbox for MATLAB in the estimation of our sign-restricted models. We thank Raphael Brun-Aguerre at JP Morgan for providing data on the realised variance of the US 10-year Treasury Note. Greenwood-Nimmo gratefully acknowledges financial support from the Australian Research Council under grant number DE150100708. The usual disclaimer applies.

1 Introduction

Systemic banking crises are not rare events, with 147 recorded crises between 1970 and 2011 (Laeven and Valencia, 2013). To mitigate the economic consequences of systemic banking crises, national policymakers have employed a variety of measures to assist troubled financial institutions, including direct recapitalisation, debt guarantees and deposit guarantees. We shall refer to such measures collectively as *bailouts* throughout this paper. In the absence of an international resolution mechanism, financial sector bailouts have historically occurred on an idiosyncratic, unilateral basis. Much of the academic literature on financial sector bailouts reflects their country-specific nature and focuses on their domestic implications, including moral hazard costs and the negative impact on the fiscal position of the responsible sovereign (e.g. Mailath and Mester, 1994; Acharya and Yorulmazer, 2007; Acharya, Drechsler and Schnabl, 2014; Stângă, 2014). However, financial institutions typically have significant cross-border exposures and so unilateral bailouts may create substantial externalities, particularly if the fiscal burden of the bailout is so severe that it raises the possibility of a sovereign default. This leads to challenging issues of policy coordination between sovereign governments whose interests may diverge (e.g. Niepmann and Schmidt-Eisenlohr, 2013). Our contribution is to quantify the cross-border spillover effects arising from country-specific idiosyncratic financial sector bailouts and sovereign risk shocks in a panel of European core and peripheral economies using an ensemble of bilateral structural macroeconomic models.

Acharya et al. (2014) establish an important stylised feature of the recent financial sector bailouts in Europe which will play an important role in our modelling framework. Using credit default swap (CDS) spreads to measure the default risk of sovereign bonds and of corporate bonds issued by banks in a number of European countries[1], Acharya et al. (2014) demonstrate that bank bailouts are associated with a change in the comovement between sovereign risk and private sector credit risk. Prior to a bailout, there is no systematic comovement between sovereign risk and financial sector risk. At the time of a bailout, the sovereign absorbs a portion of the credit risk from the financial sector, leading to an inverse association between their respective credit spreads – financial sector credit risk falls while sovereign credit risk rises. Finally, after the bailout, sovereign and financial sector credit spreads start to move in the

[1]A CDS operates like an insurance contract in which a bondholder pays a premium to transfer the default risk of the bond onto the protection seller over a given period. CDS spreads are among the purest measures of credit risk because the CDS market is considered to be the leading forum for credit risk price discovery and certain segments on the CDS market are highly liquid (Blanco, Brennan and Marsh, 2005; Gyntelberg, Hördahl, Ters and Urban, 2013).

same direction as increased sovereign risk reduces the future value of the sovereign guarantee of the financial sector (in both its explicit and implicit forms) while simultaneously reducing the value of the financial sector's holdings of sovereign debt. As the financial sector becomes riskier, the likelihood that further sovereign intervention in the financial sector will be required rises, causing a further escalation of sovereign risk and so on in a mutually-reinforcing manner.

Stângă (2014) shows that the stylised facts documented by Acharya et al. (2014) form a basis for the sign-identification of financial sector bailout shocks and sovereign risk shocks in the context of a sign-restricted vector autoregressive (SRVAR) model. Stângă estimates a set of country-specific trivariate SRVAR models in the sovereign CDS spread, the financial sector CDS spread and the sovereign term spread and identifies three fundamental shocks: an expansionary business cycle shock, an adverse sovereign risk shock and a financial sector bailout shock. Her results indicate that country-specific bailouts in Europe have generally failed to achieve a sustained reduction in local financial sector credit risk, while adverse sovereign risk shocks have caused a sustained deterioration in local financial sector credit risk. The international spillover effects arising from these shocks are not addressed, however.

Our modelling framework differs from that of Stângă (2014) in several ways, the most important of which is that we estimate bilateral models to study bilateral (between-country) spillover effects rather than unilateral models to study unilateral (within-country) effects. The shift to a cross-country context necessitates careful handling of the sources of common variation, including variations in global macroeconomic conditions, investor risk appetite and funding liquidity (e.g. Chudik and Fratzscher, 2011; Erdem, Kalotychou, Remolona and Wu, 2016; Greenwood-Nimmo, Huang and Nguyen, 2017). Consequently, in addition to country-pair-specific information, our bilateral SRVAR models contain an array of global controls, including the equity and treasury variance risk premia defined by Bollerslev, Tauchen and Zhou (2009) and Mueller, Vedolin and Yen (2012), respectively, the TED spread and a selection of macrofinancial indicators for the US which, as the world's largest economy, is known to act as a dominant global unit (e.g. Pesaran and Chudik, 2013) and to exert a common influence on global credit spreads (e.g. Longstaff, Pan, Pedersen and Singleton, 2011; Augustin and Tédongap, 2016).

We use weekly data spanning the period January 2006 to July 2015 to estimate bilateral SRVAR models covering each of the ninety pairwise combinations of the following ten European countries: Austria, Belgium, France, Germany, Ireland, Italy, the Netherlands, Portugal, Spain and the UK. The scale of this exercise renders traditional impulse response analysis infeasible. For example, it would require $10 \times 9 \times 2 = 180$ impulse response functions to summarise the bilateral spillovers generated by a financial sector bailout shock occurring in each country onto sovereign and financial sector credit risk in every other country. Consequently, we adapt a visualisation technique from the empirical network literature associated with Diebold and Yilmaz (2009, 2014) and Alter and Beyer (2014) and use spillover tables to summarise the pair-wise spillovers among the countries in our sample. Our approach enjoys a substantial advantage within this literature because our analysis is based on structurally-identified idiosyncratic shocks rather than cross-sectionally correlated reduced form disturbances (as in Alter and Beyer, 2014 and Diebold and Yilmaz, 2014) or uncorrelated shocks obtained from an arbitrary triangular decomposition (as in Diebold and Yilmaz, 2009). To the best of our knowledge, ours is the first paper to construct structural spillover tables from a sign-identified system. Consequently, our technique represents a significantly addition to the literature on empirical network modelling.

Our first finding is that the point estimates of the spillover effect from a financial sector bailout shock occurring in country j onto financial sector credit risk in country $i \neq j$ are negative in 77 out of 90 cases. While this is suggestive of beneficial spillover effects arising from country-specific financial sector bailout shocks in Europe, the degree of uncertainty surrounding these estimates is such that the evidence is inconclusive — the 68% intervals obtained from the set of draws that satisfy our identifying sign restrictions includes zero in every case. This uncertainty reflects the nature of a financial sector bailout shock, where financial sector credit risk is not extinguished but merely transferred to the public sector. Consequently, if financial institutions in the ith country are exposed to both the jth financial sector and the jth sovereign, then the benefit accruing from the reduction in the credit risk of the jth financial sector brought about by the bailout will be offset to some extent by the accompanying increase in the credit risk of the jth sovereign. In addition, we find that financial sector bailout shocks do not generally spillover onto international sovereign credit spreads because the fiscal burden of a bailout programme is borne exclusively by the responsible sovereign. This finding also suggests the absence of information contagion effects, whereby the information revealed by the announcement of a bailout in country j leads investors to perceive an increased likelihood of a bailout in country i.

Next, given that the fiscal burden of a major bailout may raise serious concerns over the sustainability of sovereign debt, we consider the spillover effects arising from adverse sovereign risk shocks. It is well-established that an adverse sovereign risk shock affecting the jth sovereign is likely to exacerbate credit risk in the jth financial sector (e.g. Acharya et al., 2014; Acharya and Steffen, 2015; Fratzscher and Reith, 2015; Greenwood-Nimmo et al., 2017). This effect does not depend on home bias in the sovereign debt holdings of financial institutions but it is exacerbated by home bias and this was a major concern among European policymakers at the height of the debt crisis (European Systemic Risk Board, 2015). However, it follows from the analysis of Bolton and Jeanne (2011) that the diversification of sovereign debt portfolios directly exposes the financial sector to foreign sovereign risk shocks, thereby creating a direct channel by which sovereign risk shocks can generate international spillovers.[2] Moreover, even if financial institutions in country i are not directly exposed to debt issued by the jth sovereign, their dealings with financial institutions domiciled in country j may create indirect channels through which sovereign risk shocks can spread internationally.

Our results indicate that, due to these international linkages, sovereign risk shocks generate profound international spillovers onto foreign financial institutions. This is particularly true of sovereign risk shocks originating in the Eurozone core and of the larger peripheral economies, which suggests that the size and degree of development of the sovereign debt market is an important determinant of the degree to which sovereign risk shocks spillover onto global financial sector credit risk. Given the within-country linkage between financial sector risk and sovereign risk documented by Acharya et al. (2014), it follows that if a sovereign risk shock in country j drives up financial sector credit risk in country i, then sovereign risk in country i may also rise. It is therefore not surprising that we find broadly similar behaviour in the spillover of adverse idiosyncratic sovereign risk shocks onto foreign sovereigns, with large countries and those which are more strongly integrated into the Eurozone core exhibiting a greater potential to generate destabilising spillovers than smaller peripheral economies.

Our results sound a cautionary note regarding the unintended consequences of financial sector bailouts. Provided that the creditworthiness of the responsible sovereign is not called into question, then a bailout is likely to reduce financial sector credit risk locally and may also generate some beneficial international spillovers. However, if the fiscal burden of the bailout is sufficiently onerous that it raises the prospect that the sovereign may default, then the bailout is almost sure to fail in its local stabilisation objectives and poses a significant international financial contagion risk.

This does not mean that sovereigns should not attempt to stabilise ailing financial institutions during financial crises. Indeed, the losses associated with the disorderly progression

[2] Acemoglu, Ozdaglar and Tahbaz-Salehi (2015) have shown that diversification enhances financial stability if adverse shocks are small but it undermines financial stability in the presence of large adverse shocks.

of a systemic financial crisis are likely to prove severe, both locally and globally. Rather, our results suggest that a well-designed mechanism for the resolution of systemic banking crises should internalise the costs associated with elevated sovereign risk both for the local economy and for the global economy. Such an international resolution mechanism would also provide a means to coordinate bailouts across countries in order to maximise the potential for beneficial spillovers across borders. In addition, it would eliminate the uncertainty arising from the uncoordinated bailout actions of multiple sovereigns independently pursuing their own domestic objectives (Niepmann and Schmidt-Eisenlohr, 2013) as well as much of the moral hazard described by King (2015) in the context of the political bargaining surrounding bailout arrangements for multinational banks with bargaining power. We therefore view the European Commission's adoption of the Bank Recovery and Resolution Directive as a valuable framework for the coordination of future multilateral bailouts.

Our paper contributes to two strands of literature. First, we build upon the empirical network modelling literature associated most notably with Diebold and Yilmaz (2009, 2014). Studies in this literature have analysed the topology of financial networks based on reduced form VAR models. While this approach has several appealing features including ease of implementation and interpretation, it does not provide a straightforward method to study the spillovers arising from structurally identified shocks or to conduct counterfactual analyses. Our technique addresses these limitations. Second, we add to the literature on the sovereign-financial credit risk nexus. Many studies in this literature have addressed the within-country domestic credit risk transfers associated with financial sector bailouts and the conditions under which destabilising feedback effects can emerge between financial sector and sovereign credit risk in a given country (e.g. Acharya et al., 2014; Fratzscher and Reith, 2015). However, little research has addressed the potential for international credit risk transfers, which we show to be substantial.

This paper proceeds as follows. Section 2 outlines our empirical framework and provides a detailed discussion of our identifying sign restrictions. Details of our dataset including data sources, data transformations and descriptive statistics may be found in Section 3. Our estimation results are presented in Section 4, while Section 5 concludes.

2 The Empirical Framework

Our analysis is based on an array of bilateral sign-identified structural macroeconomic models covering all pairwise combinations of the following $N = 10$ countries: Austria, Belgium, France, Germany, Ireland, Italy, the Netherlands, Portugal, Spain and the UK. Our sample spans the period January 2006 to July 2015 at weekly frequency for a total of $T = 498$ observations. During this time, each of our sample countries is a member of the European Union and all except the UK are members of the Eurozone. For the $\{i, j\}$th country-pair, we first estimate a reduced form VAR model and we then use the pure sign restrictions framework of Uhlig (2005) to identify three fundamental shocks occurring in country j — an expansionary business cycle shock, an adverse sovereign risk shock and a financial sector bailout shock. Our analysis focuses on the latter two shocks, with the primary role of the business cycle shock being to sharpen our identification scheme. These three shocks correspond to those considered by Stângă (2014) but in the more general context of a bilateral multivariate model as opposed to the unilateral trivariate SRVAR models that she considers.

2.1 Specification of the Bilateral VAR Models

We start by estimating a set of reduced form bilateral VAR models which exhaustively cover each of the of $10 \times 9 = 90$ pairwise combinations of countries for $i, j = 1, 2, \ldots, N$, $i \neq j$. For the ith country, we observe the $k \times 1$ vector of country-specific variables, $\boldsymbol{x}_{it}$:

$$\boldsymbol{x}_{it} = (g_{it}, b_{it}, s_{it}, l_{it})', \ i = 1, 2, \ldots, N \tag{1}$$

where time periods measured in weeks are indexed by $t = 1, 2, \ldots, T$, g_{it} is the sovereign credit spread, b_{it} is the aggregate financial sector credit spread constructed by Greenwood-Nimmo et al. (2017), s_{it} is the sovereign term spread and l_{it} is the spread between the 3-month interbank interest rate and the 3-month government bond yield.

The data used to estimate the bilateral VAR model for the $\{i, j\}$th country-pair naturally includes the vectors of country-specific variables for countries i and j, $\boldsymbol{x}_{it}$ and $\boldsymbol{x}_{jt}$. In addition, the model includes the relative yield on long-term government bonds issued by countries i and j, $r_{ij,t} = r_{it}/r_{jt}$. This country-pair-specific information will play an important role in our identification strategy. Lastly, to account for the influence of global conditions on both countries i and j, the $\{i, j\}$th bilateral VAR model includes a set of global controls. Adequately controlling for sources of common variation in the data for countries i and j is important if we are to interpret country-specific shocks as idiosyncratic in nature. Consequently, we include the following $k_i^* \times 1$ vector of observed global controls:

$$\boldsymbol{x}_t^* = \left(\boldsymbol{x}_{0t}', v_t^q, v_t^s\right)' \tag{2}$$

The vector $\boldsymbol{x}_{0t}$ is defined analogously to (1) and contains the sovereign and financial sector credit spreads, the term spread and the interbank–treasury spread for the US.[3] The last two variables in $\boldsymbol{x}_t^*$, v_t^q and v_t^s, denote the S&P 500 equity variance risk premium (VRP) and the US treasury VRP, respectively (see Bollerslev et al., 2009; Mueller et al., 2012). VRPs are among the foremost measures of investor risk appetite, variations in which have been identified as an important factor in the transmission of the GFC (e.g. Chudik and Fratzscher, 2011; Erdem et al., 2016). By including the equity and treasury VRPs separately, we are able to control for the risk appetite of equity and fixed income investors without restricting them to be equal.

Our use of US data to control for global conditions follows a long-established precedent that reflects the dominance of the US in the global economic and financial system. In the context of multi-country VAR models, the US has been shown to act as a dominant unit in the sense that it exerts a common influence on all other countries in the model (e.g. Chudik and Fratzscher, 2011; Pesaran and Chudik, 2013). Furthermore, controlling for US macroeconomic conditions is particularly important in the study of sovereign credit risk. Augustin and Tédongap (2016) demonstrate the existence of a time-varying premium in the sovereign CDS spread which reflects exposure to US macroeconomic risk, while Longstaff et al. (2011) show that sovereign CDS spreads in many countries respond more strongly to US stock and high yield markets than to domestic conditions.

With the above definitions in hand, the bilateral VAR model for the $\{i, j\}$th country-pair can be written as follows:

$$\boldsymbol{z}_{ij,t} = \boldsymbol{\alpha}_{ij} + \sum_{\ell=1}^{p} \boldsymbol{\phi}_{ij,\ell} \boldsymbol{z}_{ij,t-\ell} + \boldsymbol{e}_{ij,t} \tag{3}$$

where $\boldsymbol{z}_{ij,t} = \left(\boldsymbol{x}_{it}', r_{ij,t}, \boldsymbol{x}_{jt}', \boldsymbol{x}_t^{*\prime}\right)'$, Greek letters denote parameters to be estimated, $\boldsymbol{e}_{ij,t}$ denotes the vector of reduced form residuals with positive definite covariance matrix $\boldsymbol{\Omega}_{ij,t}$ and p is the VAR lag order, which is determined by minimisation of the Schwarz information criterion.

Two points are noteworthy in relation to (3). First, we model all variables endogenously as bidirectional feedback effects are likely to be prevalent among financial time series sampled at

[3]Note that, under our definition, the US interbank–treasury spread is precisely identical to the TED spread, which is a widely used indicator of global funding liquidity.

weekly frequency. Second, when countries i and j are both Eurozone members, it is infeasible to estimate (3) because $l_{it} \equiv l_{jt}$ by construction.[4] Our solution is to omit the interbank interest rate for the ith country when working with a Eurozone country-pair, a step which does not entail any loss of information from the model and which can be easily accommodated within our identification strategy. Consequently, the data vector $\boldsymbol{z}_{ij,t}$ is of dimension $K_{ij} \times 1$, where $K_{ij} = k_i + k_j + 1 + k^* = 15$ for non-Eurozone pairs and $K_{ij} = (k_i - 1) + k_j + 1 + k^* = 14$ for Eurozone pairs.

2.2 Structural Identification using Sign Restrictions

We are not directly interested in the reduced form innovations, $\boldsymbol{e}_{ij,t}$, but rather in a subset of three mutually independent fundamental shocks occurring in the jth country: (i) an expansionary business cycle shock; (ii) an adverse sovereign risk shock; and (iii) a financial sector bailout shock. In line with several related papers in the literature, our approach to identification is based on sign restrictions applied to the impulse responses derived from the VAR model (e.g. Chudik and Fratzscher, 2011; Stângă, 2014). Particularly in the context of low frequency financial data where contemporaneous feedback effects are prevalent, the use of sign restrictions confers several benefits relative to traditional identification strategies including the Wold-causal identification routine of Sims (1986), the use of short-run restrictions as in Blanchard and Watson (1986) and the use of long-run restrictions following Blanchard and Quah (1989). Most importantly from our perspective, sign-identification supports the agnostic identification of shocks where minimal structure is imposed on the response of the system to the identified shocks. Specifically, sign-identification makes use of inequality constraints which are inherently weaker than the point restrictions used in alternative identification schemes. Furthermore, unlike Wold causal identification, sign-identification does not invoke recursivity and is invariant to the ordering of the variables in the system.

The one-step ahead prediction error from the $\{i, j\}$th reduced form bilateral VAR model, $\boldsymbol{e}_{ij,t}$, is related to the set of K_{ij} fundamental shocks, $\boldsymbol{v}_{ij,t}$, as follows:

$$\boldsymbol{e}_{ij,t} = \boldsymbol{A}_{ij}\boldsymbol{v}_{ij,t} \tag{4}$$

[4] We use the 3-month Euribor to measure the rate of interest on interbank loans in the Eurozone and we compute the interbank spread for all Eurozone countries relative to to the 3-month German Bund yield, which we treat as approximately free from default risk. This is consistent with other recent papers in the literature (e.g. Pelizzon, Subrahmanyam, Tomio and Uno, 2016).

where $\boldsymbol{A}_{ij}$ is a $K_{ij} \times K_{ij}$ identifying matrix. As is common in the literature, we assume that there are the same number of fundamental shocks as there are reduced form disturbances and that the fundamental shocks are mutually uncorrelated with unit variance such that $E\left(\boldsymbol{v}_{ij,t}\boldsymbol{v}'_{ij,t}\right) = \boldsymbol{I}_{K_{ij}}$. Our interest is limited to the identification of three fundamental innovations, so we need only develop a partially identified model focusing on three column vectors from $\boldsymbol{A}_{ij}$, $\boldsymbol{a}^b_{ij}$, $\boldsymbol{a}^s_{ij}$ and $\boldsymbol{a}^f_{ij}$, which correspond to the business cycle, sovereign risk and financial sector bailout shocks, respectively. These three *impulse vectors* define the impact effect of the respective shock on each of the variables in the system.

For clarity of exposition, we outline the sign restrictions framework in the simple case of a single impulse vector — see Mountford and Uhlig (2009) for the generalisation to two or more impulse vectors. Following Uhlig (2005), one may define the identifying matrix $\boldsymbol{A}_{ij}$ by Cholesky decomposition such that $\boldsymbol{A}_{ij}\boldsymbol{A}'_{ij} = \boldsymbol{\Omega}_{ij}$ without loss of generality. Having defined $\boldsymbol{A}_{ij}$, Uhlig shows that if there exists a conformable vector of unit length, $\boldsymbol{m}_{ij}$, then the corresponding impulse vector is $\boldsymbol{a}_{ij} = \boldsymbol{A}_{ij}\boldsymbol{m}_{ij}$. Given the impulse vector $\boldsymbol{a}_{ij}$, the corresponding impulse response at the h-period horizon, $\boldsymbol{R}_{a,ij}(h)$, is defined as:

$$\boldsymbol{R}_{a,ij}(h) = \sum_{\ell=1}^{K_{ij}} m_{ij,\ell}\boldsymbol{R}_{ij,\ell}(h) \tag{5}$$

where $m_{ij,\ell}$ is the ℓth element of $\boldsymbol{m}_{ij}$ and $\boldsymbol{R}_{ij,\ell}(h)$ is a vector containing the h-period ahead impulse responses with respect to the ℓ-th shock in the Cholesky decomposition of $\boldsymbol{\Omega}_{ij}$.

Estimation proceeds as follows. First, one draws from a Normal-Wishart posterior for $(\boldsymbol{\phi}_{ij}, \boldsymbol{\Omega}_{ij})$ and computes the Cholesky factor, $\boldsymbol{A}_{ij}$. Next, for each of these parameter draws, one draws repeatedly from a uniform distribution over the unit sphere for $\boldsymbol{m}_{ij}$. Each time, one computes the impulse vector $\boldsymbol{a}_{ij}$. The impulse responses derived in this way are then evaluated against the sign restrictions — the draw is discarded if the sign restrictions are violated and retained otherwise. The process is repeated until the desired number of draws — in our case 1,000 — has been retained. Analysis and inference is then conducted using the set of retained draws. Under Uhlig's pure sign restrictions approach, all impulse responses that satisfy the sign restrictions are considered to be equally valid and are therefore given equal weight. In this case, the set of retained impulse response functions are summarised via a measure of their central tendency such as their median or the median target defined by Fry and Pagan (2011) as well as selected percentiles of their empirical distribution. Alternatively, one could employ penalised combinations of the retained impulse response functions. However, as noted by Fry and Pagan (2011), the use of a penalty function relies on additional non-sign information, which amounts to the introduction of additional modelling assumptions beyond those which are explicitly laid out in the formulation of the sign restrictions. We therefore adopt the pure sign restrictions approach due to its transparency.

2.3 Identification Strategy

Table 1 summarises our sign restrictions, which extend those used by Stângă (2014) to the bilateral case and which owe an intellectual debt to Acharya et al. (2014). For the $\{i, j\}$th bilateral model, we identify three shocks originating in country j: (i) an expansionary business cycle shock; (ii) an adverse sovereign risk shock; and (iii) a financial sector bailout shock. We impose all of the sign restrictions in Table 1 both in the week of impact and in the following week. Although Fry and Pagan (2011) show that increasing the duration over which sign restrictions bind does not guarantee tighter identification, in our case we find that imposing restrictions for two weeks rather than one yields narrower intervals and richer dynamics. Nonetheless, our results are qualitatively similar under either setting — a detailed comparison is available on request. As our restrictions bind for two weeks, we select the lag order for each bilateral model from the set $p \in \{2, 3\}$. This ensures that the duration over which the restrictions bind does not exceed the VAR lag length and also excludes the possibility of excessively high lag orders given the high-dimensional nature of our model and the relatively short span of data available for estimation.

	Domestic Economy					Foreign Economy				Global Economy					
	g_i	b_i	s_i	l_i	r_i/r_j	g_j	b_j	s_j	l_j	g_0	b_0	s_0	l_0	v^q	v^s
Bus. Cycle					−	−	−	+							
Sov. Risk					−	+	+	+							
Fin. Bailout					−	+	−	+	−						

Notes: Non-negative and non-positive sign restrictions are denoted '+' and '-', respectively. All restrictions hold only in the week of impact and in the following week. Note that the domestic interbank–treasury spread, l_i, is omitted from models where both countries i and j belong to the Eurozone without loss of generality.

Table 1: Identifying Sign Restrictions

Although our primary interest is in the identification of sovereign risk and financial sector bailout shocks, the inclusion of the business cycle shock serves to sharpen our identification scheme. To see this, note that the business cycle and sovereign risk shocks separate two cases in which sovereign and financial sector credit risk share a common sign. By identifying both of these shocks together, we introduce a large number of additional relevant restrictions which will improve the probability of recovering well-identified and meaningful fundamental shocks (see Canova and Paustian, 2011, for a related discussion). Note that we avoid imposing restrictions on country i when identifying shocks arising from country j to ensure that our sign restrictions do not mechanically determine the pattern of international spillover activity. In addition, we do not impose any restrictions on the global controls. In our baseline model, we allow all of the global controls to react freely to the identified shocks. Given that the economies that we study are small relative to the global economy, we do not expect country-specific shocks to exert a strong influence on any of our global controls. We test this assumption through a number of exclusion exercises in which key global channels of shock transmission are manually shut down, as detailed in subsection 2.4. The intuition underlying each of our identified shocks is summarised below.

2.3.1 An Expansionary Business Cycle Shock

Our expansionary business cycle shock represents an improvement in economic fundamentals in country j. This improves the jth sovereign's fiscal position and its ability to service debt, lowering its credit risk. Likewise, financial sector credit risk in country j declines as a result of improved earnings prospects in the private sector. As long-term interest rates in country j adjust to reflect the higher expected growth rate, the term spread increases. In addition, the increase in long-term yields in country j reduces the relative long-term yield r_i/r_j under the maintained assumption that there is no change in the long-term yield in country i.

2.3.2 An Adverse Sovereign Risk Shock

An adverse sovereign risk shock in country j raises the credit risk of the jth sovereign by definition. As noted by Acharya et al. (2014), this increases the credit risk of the jth financial sector for two reasons: (i) it reduces the value of domestic sovereign debt held by the financial sector and this will typically account for a substantial proportion of a financial institution's sovereign debt portfolio due to home bias; and (ii) the reduction in the sovereign's borrowing capacity diminishes the future value of the sovereign's guarantee of the financial sector. The increase in sovereign risk is reflected in a rising default risk premium on debt issued by the jth sovereign. This increases the term spread in country j and reduces the relative long-term yield

r_i/r_j assuming that the long-term bond yield in country i is unaffected.

2.3.3 A Financial Sector Bailout Shock

A financial sector bailout in country j transfers credit risk from the jth financial sector to the jth sovereign as described by Acharya et al. (2014). This raises the default risk premium on sovereign debt, increasing the yield spread in country j and reducing the relative long-term yield r_i/r_j. This is consistent with Acharya et al. (2014)'s observation that the fiscal burden of a financial sector bailout in the jth country is borne by the jth sovereign and that this burden may be such that the conditions under which the sovereign can raise funds are no longer as advantageous as they were prior to the bailout. By contrast, the reduction in financial sector credit risk relaxes the interbank money market, narrowing the interbank–treasury spread in country j.

2.4 Excluding Global Responses

As mentioned above, our baseline results are obtained under the assumption that the global variables are free to respond to idiosyncratic shocks arising in country j. This allows country-specific shocks to propagate internationally through a number of channels, including direct bilateral spillovers between countries i and j as well as indirect effects arising due to variations in global investor risk appetite (reflected in the two VRPs) and variations in global liquidity (reflected in the TED spread), for example. In practice, there is little reason to believe that these global variables should respond strongly to idiosyncratic shocks occurring in the small open economies that we study. We analyse this contention via the following exclusion exercises:

(i) we shut down the response of the VRPs to the identified shocks, thereby ensuring that shock propagation does not occur via variations in global investor risk appetite;

(ii) we shut down the response of the TED spread to the identified shocks, which ensures that variations in global liquidity do not play a role in the propagation of the shocks;

(iii) we combine cases (i) and (ii) and simultaneously shut down the responses of the VRPs and the TED spread to the identified shocks; and

(iv) we shut down the responses of all six global variables to the identified shocks, thereby ensuring that variations in global risk appetite, global liquidity conditions and US macro-financial conditions do not play a role in shock propagation.

If our identification scheme is successful in identifying idiosyncratic country-specific shocks, then shutting these global effects down should not materially affect the inferences derived from our model. Consequently, these four exclusion exercises represent valuable tests of the performance of our identification scheme.

3 Construction and Properties of the Dataset

In this section we provide details of the country-specific, country-pair-specific and global variables used to estimate our bilateral SRVAR models. Recall that our analysis spans the period January 2006 to July 2015 at weekly frequency and focuses on ten European countries — Austria, Belgium, France, Germany, Ireland, Italy, the Netherlands, Portugal, Spain and the UK — with additional data for the US being used to control for global conditions. Although it would have been desirable to include Greece in our sample given that it experienced a profound sovereign debt crisis, the absence of reliable data on the Greek sovereign credit spread in 2012 and 2013 precludes this possibility.[5]

3.1 Sovereign Credit Spreads

We measure the sovereign credit risk for the ith country using the five-year sovereign CDS spread. In line with the CDS market conventions outlined by Bai and Wei (2017), we work with US dollar denominated CDS in all cases except for the US, where we employ Euro denominated CDS. In addition, following Bai and Wei, we use CDS contracts with a complete restructuring clause for each sovereign. The CDS data is sourced from *Markit* at daily frequency and is converted to weekly frequency by taking the Wednesday observation of each week.[6] By defining a week from Thursday to Wednesday, we avoid any potential weekend effects. In practice, the use of a period average relative to simply working with the Wednesday observation of each week

[5]The five-year Greek sovereign CDS spread exceeds 10,000bp on 15-Feb-2012 and is not reported by *Markit* over the period 09-Mar-2012 to 06-Jun-2013. During this time, Greek sovereign CDS contracts traded on a points upfront basis to ensure that protection sellers would not be obliged to pay out following a credit event without having received an income stream from the contract. Consequently there is a large gap in the data. We attempted to estimate our bilateral SRVAR models including Greece using a reduced sample period ending in February 2012 but unfortunately there was too little data to obtain reliable results. Therefore, we are obliged to exclude Greece from our analysis.

[6]Stângă (2014) also converts daily CDS spread data to weekly frequency. Sampling at weekly rather than daily frequency suppresses much of the noise in the data and tends to increase the retention rate across draws in the sign-identification algorithm, thereby significantly reducing the computational time. Mitigating the computational burden is critical here as we must evaluate $10 \times 9 = 90$ bilateral SRVAR models for our benchmark case and for each of our four exclusion exercises, for a total of 450 models.

has little effect on our results (a detailed comparison is available on request).[7]

3.2 Financial Sector Credit Spreads

We measure financial sector credit risk in the ith country using the aggregate financial sector CDS spread originally constructed by Greenwood-Nimmo et al. (2017). The index for the ith country is computed as an equally weighted average of the single name CDS spreads for locally-domiciled financial firms which satisfy a set of selection criteria inspired by those used by Acharya et al. (2014). Specifically, with few exceptions, firms must: (i) be listed in the Markit CDS database and have USD denominated five year CDS spread data which accords with the corporate CDS market conventions documented by Bai and Wei (2017) and which covers at least 10% of the sample period originally considered by Greenwood-Nimmo et al. (2017), which closely maps onto our sample period but at daily frequency; (ii) be classified by Markit as *financials*; (iii) be classified as either banking or insurance firms in Bureau van Dijk's *Osiris* database; (iv) be identified by Markit as operating in the ith country; (v) hold assets of USD10bn or more in at least one year between 2006 and 2015; and (vi) have publicly traded equity.[8] Overall, data for 137 financial institutions is used to construct the aggregate financial sector CDS spreads that we use in estimation.

3.3 Term Spreads

In keeping with Stângă (2014), our models include the term spread as a barometer of macroeconomic conditions in the ith country. The term spread is invaluable in this regard as it offers a parsimonious means to capture information concerning macroeconomic fundamentals, liquidity premia, inflation and output growth expectations and the stance of monetary policy. We compute the term spread as the spread between the ten-year and the three-month government bond yields. Details of the series used to compute the term spread for each country may be found in the Data Appendix.

[7] We have also experimented with the use of monthly data defined using end-of-month values. Using first-order SRVAR models and imposing our sign restrictions on impact only reveals that our results are robust to the change of frequency. However, the dynamics obtained using monthly data are less rich than with weekly data.

[8] Greenwood-Nimmo et al. (2017) make a small number of exceptions to these selection criteria. For example, even though they do not appear in *Osiris*, ABN Amro and Fortis are manually added to the sample on account of their regional importance during the GFC. Likewise, even though it does not have publicly traded equity, Raiffeisen Zentralbank is included in the composite Austrian financial sector CDS spread as otherwise the index would be based on data for a single firm. Unlike Acharya et al. (2014), the financial sector indices constructed by Greenwood-Nimmo et al. (2017) include CDS data for non-bank financial institutions (e.g. insurance firms) as well as several institutions which became state-owned as a result of the crisis, such as the Irish Bank Resolution Corporation. For a comprehensive discussion of the construction and properties of the financial sector CDS spreads, the reader is referred to Greenwood-Nimmo et al. (2017).

3.4 Interbank–Treasury Spreads

We measure interbank liquidity in the ith country using the spread between the three-month interbank interest rate and the appropriate three-month government bond yield. In the case of the US, this measure is identical to the TED spread, which has been widely used as a measure of funding liquidity (e.g. Greenwood-Nimmo, Nguyen and Rafferty, 2016a). For the Eurozone countries, we employ the Euribor-DeTBill spread used by Pelizzon et al. (2016), for example, which is defined as the spread between the three-month Euribor and the three-month German Bund yield.[9] Details of the data used to compute the interbank–treasury spreads may be found in the Data Appendix.

3.5 Relative Long-Term Bond Yields

We use the relative yield on ten-year government bonds issued by sovereigns i and j, r_i/r_j, to measure changes in the relative debt-servicing cost that each sovereign faces. Details of the data used to construct the relative long-term bond yields are provided in the Data Appendix.

3.6 Variance Risk Premia

We use the equity and treasury variance risk premia to control for variations in investor risk appetite. In line with the original definition put forward by Bollerslev et al. (2009), we define the VRP as $v_t = IV_t - E\left[RV_{t+1}^{(22)}\right]$, where IV_t is the implied variance and $E\left[RV_{t+1}^{(22)}\right]$ is the one-month-ahead forecast of the realised variance. For the equity VRP, IV_t is defined as the de-annualised squared VIX index and $RV_t^{(22)}$ is the monthly realised variance of the S&P 500 index measured as the sum of squared five-minute intraday returns over a period of 22 trading days (approximately one month). Meanwhile, for the treasury VRP, IV_t is the defined as the de-annualised squared TYVIX index and $RV_t^{(22)}$ is the monthly realised variance of the US 10 year treasury note. The VIX and TYVIX data are obtained from the Chicago Board Options Exchange. The daily realised variance of the S&P 500 is sourced from the Oxford Man Institute's Realized Library (Heber, Lunde, Shephard and Sheppard, 2009, ver. 0.2). Lastly, the monthly realised variance of the US 10 year treasury note was constructed from a daily realised variance series provided by JP Morgan. In line with the forecast evaluation results presented by Bekaert

[9] In principle, we could define a unique interbank spread for each Eurozone country by using the three-month domestic government bond yield in each case rather than the German Bund yield. However, we elect to use the common Euribor-DeTBill spread due to the approximately default-risk-free nature of German government debt, which itself is reflected in the historically low default risk of German bonds relative to the debt of other Eurozone sovereigns.

and Hoerova (2014), we use Corsi's (2009) heterogeneous autoregressive model augmented with the square of the VIX to construct the realised variance forecasts required to compute the VRPs.

3.7 Properties of the Data

Table 2 reports the mean, median and standard deviation for each country-specific and global variable in the dataset. The countries in our sample naturally divide into two groups according to the level and volatility of their data. As one may expect, countries in the Eurozone periphery — Italy, Ireland, Portugal and Spain — display higher average credit spreads with substantially greater volatility than the core countries. It is generally the case that the median of both the sovereign and financial sector credit spreads is lower than the mean, indicating a right skew in the credit spread data, with this effect being most pronounced in those countries that experienced the deepest crises. The term spreads of the peripheral economies display marked volatility that stands in constrast to the bond market convergence observed among the other European sovereigns. The interbank spread is identical for every Eurozone country, by definition. Finally, the rightmost column of the table relates to the long-term bond yield in each country. These series are used to construct the relative long-term bond yield for each country-pair. Again, the excess volatility of the Eurozone periphery is easily seen.

— Insert Table 2 here —

4 Estimation Results

4.1 An Illustration: Spillovers between France and Germany

To establish a frame of reference for the multi-country analysis that follows, we begin with a descriptive analysis of the the spillover effects associated with financial sector bailout shocks and adverse sovereign risk shocks among a single country-pair via traditional impulse response analysis. For this exercise, we focus on France and Germany, two of the largest Eurozone economies, both of which host major financial centers. The purpose of this exercise is to demonstrate the nature of the impulse responses obtained from our bilateral models in a familiar format before introducing the spillover tables that will form the basis of our multilateral analysis. Consequently, we shall defer the detailed discussion of the observed spillover effects to Sections 4.2 and 4.3, below.

Figure 1 presents a selection of impulse response functions obtained from two bilateral models: the $\{DE, FR\}$ model where we identify structural shocks occurring in France and the $\{FR, DE\}$ model where we identify structural shocks originating in Germany. Sixteen impulse response functions are reported in the form of a table which is separated into quadrants to enhance the clarity of presentation. Each column of the table relates to a separate structural shock, while each row relates to a separate variable which is affected by the shock. The upper (lower) left quadrant shows the response of financial sector (sovereign) credit spreads to financial sector bailout shocks, while the upper (lower) right quadrant shows the response of financial sector (sovereign) credit spreads to sovereign risk shocks.

— Insert Figure 1 here —

In each panel of Figure 1, the median impulse response is shown as a solid red line. Although our discussion will focus on the median impulse response, we also report the median target impulse response of Fry and Pagan (2011) as a dashed black line for comparison. In most cases, the median and median target impulse responses are qualitatively similar so we do not discuss them separately. In addition, we plot the central 68% (90%) interval of the set of retained impulse responses as a dark (pale) gray band. It is well-known in the sign restrictions literature that these intervals should not be interpreted as confidence intervals. Rather they reflect the extent of variation in the set of impulse response functions obtained from rotations of the underlying VAR model which satisfy the sign restrictions. That is, they reflect the range of variation over different candidate structural models. Following the common practice in the sign restrictions literature, our discussion of the range of model uncertainty focuses on the 68% intervals (see Uhlig, 2005, for example), although we include the 90% intervals in Figure 1 for completeness.

First, consider the spillovers arising from financial sector bailout shocks. We begin with a one standard deviation financial sector bailout shock occurring in France, the effects of which are reported in the first column of Figure 1. Our identifying restrictions ensure that the French financial sector credit spread does not rise in the first two weeks while the French sovereign credit spread does not fall during this time. A one standard deviation bailout shock is relatively small in the case of France, with French financial sector credit risk falling by approximately 1.50bp on impact and French sovereign credit risk increasing by approximately 0.75bp on impact. The magnitudes of these impact impulse responses are comparable to those obtained by Stângă (2014) on the basis of a single-country trivariate SRVAR model. The domestic effects of the French bailout shock dissipate relatively quickly, with the 68% intervals around the French financial sector and sovereign credit risk impulse responses encompassing zero after approximately one month. This is again consistent with Stângă's (2014) observation that financial sector bailouts in Europe generally failed to achieve sustained reductions in financial sector credit risk. However, we find that the French bailout shock spills over onto the German financial sector in delayed fashion, reducing the credit spread by approximately 1.5bp after 8 weeks. By contrast, there is little evidence of any spillover onto the German sovereign. The second column of Figure 1 reveals that much the same picture is obtained in the case of a German financial sector bailout, the effects of which are qualitatively similar to the French case examined above, with the exception that the German bailout shock does not attenuate French financial sector credit risk in an economically significant way.

Now consider the spillovers arising from adverse sovereign risk shocks. In the case of a one standard deviation adverse sovereign risk shock occurring in France, our sign restrictions imply that the French sovereign and financial sector credit spreads both respond non-negatively on impact and in the following week. This effect is sustained at around 2bp in both cases and, importantly, it also generates a concomitant deterioration in both German financial sector credit risk and German sovereign credit risk. The effect of an adverse German sovereign risk shock is qualitatively similar, with German sovereign credit risk rising by approximately 1bp and German financial sector credit risk rising by around 2bp in a sustained fashion after some mild overshooting on impact. The German sovereign risk shock elevates French financial sector credit risk by a similar amount and the French sovereign credit spread also increases although this effect is subject to some uncertainty, as reflected in the width of the 68% interval.

Next, in order to assess whether the bilateral spillover effects documented above are merely a reflection of the influence of common factors which jointly affect credit risk in France and Germany, Figure 2 shows the median impulse responses from each of the four exclusion exercises described in Section 2.4 overlaid on our baseline results from Figure 1 (to minimise clutter we suppress the median target impulse responses). Note that each country that we model can be considered 'small' in the sense that idiosyncratic shocks particular to that country are unlikely to generate sizable fluctuations in global macrofinancial conditions, including global liquidity and global risk appetite. Consequently, if our identification scheme is successful in identifying idiosyncratic country-specific shocks, then the results obtained from each of the exclusion exercises should be similar to our baseline results. Figure 2 therefore provides strong support for the validity of our identifying sign restrictions, as the median impulse responses from each exclusion exercise closely track the median impulse responses from our baseline setting and never breach the associated 68% intervals.

— Insert Figure 2 here —

Traditional impulse response analysis of this type provides a detailed illustration of the spillover dynamics but this very detail is the reason that it rapidly becomes infeasible as the dimensionality of the system increases. To illustrate, note that if we were to extend Figure 1 to cover all ten countries in our model, then each quadrant would contain 100 panels and the figure as a whole would contain 400 impulse response functions. The need for a compact visualisation is apparent. To this end, we take inspiration from a fast-growing body of work that employs VAR models in the analysis of economic and financial networks (e.g. Diebold and Yilmaz, 2009, 2014; Alter and Beyer, 2014), as detailed below.

4.2 Spillovers Arising from Financial Sector Bailouts

Studies in the VAR-based network literature construct *spillover tables* by cross-tabulating pair-wise combinations of impulse response functions or forecast error variance decompositions evaluated at a given horizon (see Diebold and Yilmaz, 2009, for an influential early contribution to the field). We adopt a similar approach, although we report results for horizons $h = 0$ (the impact spillover) and $h = 3$ (the spillover after approximately one month, which is reported in rounded parentheses) in the same table in order to convey not just the strength of bilateral spillovers at a fixed horizon but also a sense of their evolution across horizons. The impact spillover is of inherent interest because it measures the strength of contemporaneous credit risk transmission between countries. In principle, we could select any longer horizon to convey information regarding spillover dynamics — we use the one-month horizon because the impulse response functions derived from each bilateral model have typically converged to their long-run values by this time. We will present four spillover tables in total, one corresponding to each quadrant of Figure 1.

The impulse response analysis in Section 4.1 draws attention to an issue that must be handled carefully in the construction of the spillover tables — the size of the shocks will differ across countries. As an illustration, note that a one standard deviation financial sector bailout shock occurring in Germany reduces domestic financial sector credit risk by 1.67bp and raises domestic sovereign risk by 0.47bp. The equivalent values for a Portuguese bailout shock are -5.06bp and 6.11bp and for Irish bailout shock they are -14.07bp and 5.20bp. To facilitate direct cross-country comparisons requires that we normalise the scale of the shocks affecting each country. To this end, we normalise the ith financial sector bailout shock such that it reduces financial sector credit risk in country i by 50bp on impact. Likewise, we normalise the ith sovereign risk shock such that it raises sovereign risk in country i by 50bp on impact. Non-normalised versions of the spillover tables are available on request.

We start with Table 3, which corresponds to the upper left quadrant of Figure 1 and which measures the bilateral spillovers from idiosyncratic financial sector bailout shocks onto country-specific measures of financial sector credit risk. The $\{i, j\}$th cell of Table 3 reports the normalised median impulse response of the financial sector credit spread for country i to a financial sector bailout shock occurring in country j. Consequently, the leftmost column of the table reports the effect of an Austrian bailout shock on financial sector credit risk in each country, while the uppermost row of the table reports the response of Austrian financial sector credit risk to bailout shocks occurring in each country in the system. Given that we do not estimate any unilateral models — i.e. there is no $\{i, i\}$th model — the values reported in the $\{i, i\}$th cell of Table 3 are median impulse responses obtained by pooling the retained draws from each of the nine bilateral models where structural shocks occurring in country i are identified — that is, each of the models appearing in off-diagonal positions in the ith column of Table 3.

— Insert Table 3 here —

To assist the reader, if the impact impulse response is negative (positive) then the cell is shaded green (red), with the depth of shading indicating the relative magnitude of the spillover effect. Consequently, green-shaded cells denote beneficial spillovers and red-shaded cells denote detrimental spillovers on impact. Furthermore, in each case, if the 68% interval does not include zero, then the associated spillover effect is printed in bold face. Finally, the value shown in square parentheses in each cell is a count of the number of times that the inferences drawn from our baseline model on impact and at the one-month horizon are upheld in each of the four exclusion exercises described in Section 2.4. The maximum score for each exclusion exercise is two points — one for the impact horizon and another for the one-month horizon — and therefore the overall maximum score is eight points.[10]

[10] To illustrate, consider the $\{AT, BE\}$ cell in Table 3. We start by classifying the impact spillover effect in

The impact spillovers reported along the prime diagonal of Table 3 are all negative by virtue of our identifying sign restrictions and are scaled to -50bp due to our normalisation. Our main interest is in the off-diagonal elements of the table, which capture bilateral spillovers in a freely-estimated fashion. The large majority of these spillover effects are negative, which suggests that a financial sector bailout shock occurring in country j may act to attenuate financial sector credit risk in country $i \neq j$. This is true both on impact and at the one-month horizon, although the spillover effects dissipate over time. However, the 68% intervals associated with these spillover effects are generally wide, indicating considerable uncertainty. Indeed, none of the off-diagonal elements of Table 3 is printed in bold face, which indicates that the 68% interval includes zero in every case. This is a robust finding which is upheld in the overwhelming majority of our exclusion exercises — the lowest value of the robustness count shown in square parentheses is six out of a possible eight points. Consequently, although the fact that the median impulse responses are overwhelmingly negative is suggestive of beneficial spillover effects arising from country-specific financial sector bailout shocks in Europe, the evidence is inconclusive.

The uncertainty surrounding the international spillover effects of bailouts is a direct consequence of the nature of a financial sector bailout shock. Acharya et al. (2014) note that a financial sector bailout occurring in the jth country amounts to a transfer of credit risk from the jth financial sector onto the jth sovereign. This risk transfer is explicitly embodied in our identifying sign-restrictions. The key point is that credit risk is not extinguished in this process, merely transferred. If financial institutions in country $i \neq j$ are exposed to both the jth financial sector and the jth sovereign, then it follows that two countervailing forces act on the counterparty risk faced by the ith financial sector — downward pressure due to the reduction in foreign financial sector risk and upward pressure due to the increase in foreign sovereign risk. These effects of these two opposing forces offset one-another, at least in part, thereby contributing to the observed uncertainty.

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our baseline model into one of three cases: (a) the 68% interval includes zero; (b) the spillover effect is negative and the 68% interval does not include zero; and (c) the spillover effect is positive and the 68% interval does not include zero. Consequently, we allocate the impact spillover effect in the $\{AT, BE\}$ cell to case (a). We then repeat this exercise for each of our four exclusion exercises. For each exclusion exercise where the impact spillover effect is also allocated to case (a), we score one point; we score zero points if it is allocated either to case (b) or case (c). Consequently, if the inference regarding the impact spillover effect drawn in our baseline model is never upheld in any exclusion exercise, we score zero points; if it is upheld in one of the four exclusion exercises, we score one point and so on up to a maximum of four points. We then repeat the entire exercise for the one-month spillover effect. Finally, we sum the points obtained on impact and at the one-month horizon together and report this value in square parentheses in each cell of the table. A value of $n \in \{0, 1, 2, \ldots, 8\}$ indicates that the proportion of cases in which our baseline results are upheld is $n/8$. For the $\{AT, BE\}$ cell, the score of 8 therefore indicates that our results at both the impact and one-month horizons are maintained throughout all of our exclusion exercises.

This discussion offers a natural segue into the analysis of the international implications of financial sector bailouts for sovereign credit risk, which are addressed in Table 4. In this case, our identifying sign restrictions impose a positive sign on impact along the prime diagonal of the table. Ex ante, one can identify mechanisms which may generate both negative and positive spillovers onto international sovereign risk. For example, sovereign risk in country $i \neq j$ may rise via an information-contagion effect if the information revelation associated with the bailout in country j leads investors to perceive an increased likelihood of a similar bailout in the ith country. Alternatively, sovereign risk in country i may fall if investors perceive that the bailout in country j reduces the risks faced by the ith financial sector sufficiently that the ith sovereign is now less likely to be required to bail-out its financial sector. However, this latter effect is likely to be weak given the uncertainty surrounding the spillovers of financial sector bailouts onto foreign financial sector risk documented in Table 3. Overall, our prior expectation is that a financial sector bailout shock occurring in country j should have little effect on sovereign risk in country $i \neq j$ because the fiscal burden of the bailout is borne solely by the jth sovereign.

Our results are consistent with the latter interpretation. In our baseline setting and throughout all of our exclusion excerises, the only case where the 68% interval surrounding the median impulse response does not include zero is the spillover from the Belgian bailout shock onto the German sovereign at the one-month horizon. In this case, we find that a bailout shock which raises Belgian sovereign risk by more than 35bp in a sustained fashion raises German sovereign risk by as much as 14bp. The strength of this spillover effect may reflect the fact that the Belgian sovereign intervened to protect several major financial institutions which were collectively very large relative to the size of the Belgian economy. In addition, Belgium was the country in the European core which came closest to defaulting on its debt. The issuance of Staatsbons helped the sovereign to raise funds and averted a default but the eruption of an acute sovereign debt crisis in the European core may have led investors to believe that intervention from Germany — as the strongest economy in the Eurozone — would be required to restore stability.

— Insert Table 4 here —

Overall, our results indicate that the effects of financial sector bailout shocks are largely localised. The effect on foreign sovereigns is negligible because the fiscal burden of a bailout is borne exclusively by the responsible sovereign and we find no evidence of information contagion. Meanwhile, because bailout shocks reduce local financial sector credit risk while simultaneously increasing local sovereign credit risk, the overall effect on foreign financial institutions which are exposed to both the bailed-out financial sector and the responsible sovereign is ambiguous. We now shift our attention to the spillovers arising from idiosyncratic sovereign risk shocks. In this case, our ex ante belief is that the mutually-reinforcing comovement of financial sector and sovereign credit risk is likely to generate substantial international spillover effects. This view is supported by a growing literature which shows that localised sovereign risk shocks may exacerbate foreign financial sector and sovereign credit risk (e.g. Arghyrou and Kontonikas, 2012; Alter and Beyer, 2014; Bostanci and Yilmaz, 2015; Greenwood-Nimmo, Nguyen and Shin, 2016b; Greenwood-Nimmo et al., 2017).

4.3 Spillovers Arising from Sovereign Risk Shocks

Table 5 shows the response of financial sector credit spreads to country-specific adverse idiosyncratic sovereign risk shocks. Note that domestic financial sector credit risk deteriorates on impact in response to an adverse domestic sovereign risk shock by virtue of our identification routine. This occurs because an increase in sovereign credit risk in the jth country reduces the value of locally-domiciled financial institutions' holdings of sovereign debt and also reduces the future value of the jth sovereign's guarantee of its financial sector (e.g. Acharya et al., 2014). If financial institutions attach excess weight to the debt of their home sovereign — as many do — then the link between sovereign risk and the credit risk of the *local* financial sector is strengthened.[11] As the European sovereign debt crisis deepened, regulators became increasingly concerned about the intensification of the sovereign–financial credit risk nexus brought about by sovereign debt home bias (e.g. European Systemic Risk Board, 2015).

— Insert Table 5 here —

As shown by Bolton and Jeanne (2011), however, the diversification of sovereign debt portfolios does not offer a simple solution as it exposes the *foreign* financial sector to a troubled sovereign, exacerbating contagion risk. In addition, the ith financial sector is indirectly exposed to the sovereign risk shock occurring in country j due to the network of cross-border claims that exists between financial institutions — that is, financial firms in country j are exposed to financial firms in country i which are, in turn, directly affected by the ith sovereign risk shock.

[11] There is a puzzling aspect to home bias because, although sovereign debt is regulated favourably relative to many other asset classes — for example, over our sample period, European banks have enjoyed a zero risk weight exemption on European sovereign debt — *home country* sovereign debt receives no special treatment. Consequently, several theories of home bias have been advanced, including moral suasion by governments wishing to obtain rollover finance on advantageous terms (De Marco and Macchiavelli, 2016; Ongena, Popov and Van Horen, 2016) and strategic behaviour by financial institutions wishing to affect the sovereign's choice set in the event of a bailout (Gaballo and Zetlin-Jones, 2016).

With these direct and indirect exposures in mind, it is not surprising that Table 5 reveals profound detrimental spillovers of sovereign risk shocks onto foreign financial institutions. As in the case of financial sector bailout shocks above, these findings are remarkably robust throughout each of our exclusion exercises, with the robustness count rarely dropping below seven out of a possible eight points. It is interesting to note that these international spillover effects are strong relative to the within-country effects reported on the prime diagonal of the table, particularly for sovereign risk shocks occurring in the European core which, in our model, is comprised of Austria, Belgium, France, Germany and the Netherlands. This is especially evident in the case of Germany, where a sovereign risk shock which raises the German sovereign credit spread by 50bp and the German financial sector credit spread by 131.57bp raises financial sector credit spreads elsewhere in the Eurozone core by 160–192bp and in the Eurozone periphery by 245–575bp. The heightened sensitivity of pan-European financial institutions to German sovereign risk reflects the safe haven status of German Bunds for investors seeking access to European sovereign debt markets, as reflected in the evidence of flight-to-safety favouring German Bunds (e.g. Fratzscher and Reith, 2015). Consequently, a deterioration in the quality of German sovereign debt undermines a primary source of safe Euro-denominated assets.

Interestingly, we find that sovereign risk shocks occurring throughout the Eurozone periphery transmit to peripheral financial institutions but, among the peripheral countries, it is only Italian and Spanish sovereign risk shocks which strongly affect financial credit risk in the Eurozone core. The difference between Italian and Spanish sovereign risk shocks on the one hand and Irish and Portuguese sovereign risk shocks on the other hand suggests that the size of the economy and the size of the sovereign debt market are key factors influencing the propagation of peripheral sovereign risk shocks through the financial sector.

Table 6 reveals broadly similar behaviour in the spillover of adverse idiosyncratic sovereign risk shocks onto foreign sovereigns. The impact effects on the prime diagonal of the table are positive by construction and are normalised to 50bp. These within-country effects are relatively sustained, with the one-month impulse response taking values no lower than 38bp in any case. The majority of bilateral spillover effects are also positive and, among the Eurozone core countries, the evidence of detrimental international spillover effects is often sufficiently strong that the 68% intervals exclude zero.

— Insert Table 6 here —

These strong bilateral spillovers between sovereigns may result either from direct sovereign-to-sovereign risk linkages (due to implicit or explicit risk-sharing among European sovereigns, for example) or from indirect linkages created by the network of cross-border exposures of financial institutions discussed above. In either case, our results suggest that the contagion potential of core sovereigns exceeds that of peripheral sovereigns because of the relative sizes of the corresponding economies and debt markets. When viewed in this way, it is fortunate that large sovereign risk shocks have so far been confined to the Eurozone periphery and have not arisen in the core.

5 Concluding Remarks

In this paper, we use an ensemble of bilateral sign-identified macroeconomic models to study the international spillover effects arising from idiosyncratic financial sector bailout shocks and sovereign risk shocks among a panel of core and peripheral European countries. Our approach exploits several stylised features of the joint evolution of financial sector and sovereign credit spreads during the European bank bailouts of 2008–9 documented by Acharya et al. (2014). In this way, we are able to identify both financial sector bailout shocks and adverse sovereign risk shocks. The former transfers credit risk from the financial sector onto the sovereign, which generates an inverse comovement between the financial sector and sovereign credit spreads. The latter, by contrast, causes both sovereign and financial sector credit spreads to increase as a deterioration of sovereign credit risk reduces the value of financial sector holdings of sovereign debt while simultaneously reducing the value of the guarantee of the financial sector extended by the sovereign.

Our results indicate that a financial sector bailout shock occurring in the jth country may generate beneficial spillovers onto the ith financial sector but that this effect is subject to considerable uncertainty. In a setting where the ith financial sector holds a diversified portfolio of foreign claims, this uncertainty reflects the offsetting effect of a reduction in financial sector credit risk in country j coupled with an increase in sovereign risk in country j. Meanwhile, given that the fiscal burden of the jth bailout is borne exclusively by the jth sovereign, we find little evidence that financial sector bailout shocks spillover onto foreign sovereign risk. This also suggests that financial sector bailouts do not generate information contagion effects in the sense that we find no evidence that a bailout shock occurring in country j leads investors to perceive an increased likelihood of a bailout in country i.

By contrast, an adverse sovereign risk shock in the jth country causes sovereign and financial sector credit risk in country j to comove in a mutually-reinforcing manner. As a result, there is no offsetting effect and so adverse sovereign risk shocks generate detrimental spillovers affecting international financial institutions and foreign sovereigns. In addition, we find that the contagion potential of core sovereigns exceeds that of peripheral sovereigns because of the relative sizes of the corresponding economies and debt markets.

Overall, our results sound a cautionary note regarding financial sector bailouts. Provided that the fiscal sustainability of the responsible sovereign is not called into question, then a bailout is likely to reduce financial sector credit risk locally and may also generate some beneficial international spillovers. However, if the fiscal burden of a bailout is such that the risk of sovereign default rises substantially, then a bailout is likely to fail in its local stabilisation objective and also poses a significant contagion risk. Given the potential spillovers associated with financial sector bailout programmes — especially if they undermine the stability of the responsible sovereign — the current norm whereby bailouts are enacted unilaterally is likely to lead to inferior economic outcomes relative to a multilateral resolution mechanism that appropriately prices these externalities within the global economy. Such an international resolution mechanism would also provide a means to coordinate bailouts across countries in order to maximise the potential for beneficial spillovers across borders. In addition, it would eliminate the uncertainty arising from the uncoordinated bailout actions of multiple sovereigns independently pursuing their own domestic objectives (Niepmann and Schmidt-Eisenlohr, 2013) as well as much of the moral hazard described by King (2015) in the context of the political bargaining surrounding bailout arrangements for multinational banks with bargaining power. We therefore view the European Commission's adoption of the Bank Recovery and Resolution Directive as a valuable framework for the coordination of future multilateral bailouts.

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Data Appendix

Sovereign Term Spreads

Where data is available, we compute the term spread using the yield to redemption on the 10-year benchmark bond and the 3-month treasury bill. In some cases, data limitations oblige us to substitute the 3-month yield with an appropriate zero coupon yield. The data sources for each country are listed below. Daily data is converted to weekly frequency as described in the main text.

	10-year yield		3-month yield	
	Source	Series ID	Source	Series ID
Austria	Datastream	S06676	Bloomberg	F90803M†
Belgium	Datastream	TRBG10T	Bloomberg	GBGT3MO
France	Datastream	TRFR10T	Bloomberg	GTFRF3M
Germany	Datastream	BDBRYLD	Bloomberg	GETB1
Ireland	Datastream	TRIE10T	Bloomberg	F91803M†
Italy	Datastream	TRIT10T	Bloomberg	GBOTG3M
Netherlands	Datastream	TRNL10T	Bloomberg	GTBN3M
Portugal	Datastream	TRPT10T	Bloomberg	GSPT3M
Spain	Datastream	TRES10T	Bloomberg	GSPG3M
UK	Datastream	UKMBRYD	Datastream	S02162
US	Datastream	USBD10Y	Bloomberg	USGB090Y

† denotes cases where we use zero coupon yields instead of 3-month yield to redemption.

Interbank-Treasury Spreads

We compute the interbank-treasury spread for three cases: the Eurozone (3-month Euribor minus 3-month German bond yield); the UK (3-month LIBOR minus 3-month UK bond yield); and the US (3-month USD-LIBOR minus 3-month US Treasury bill yield). Data sources are reported below. Daily data is converted to weekly frequency as described in the main text.

	3-month interbank		3-month yield	
	Source	Series ID	Source	Series ID
Eurozone	Datastream	Y03728	Datastream	S3098Q
UK	Datastream	S97086	Datastream	S02162
US	Datastream	S97074	Datastream	S02553

Relative Long-Term Bond Yields

The relative long-term bond yield for each country-pair is computed using the same 10-year bond yield data used to compute the term spread. Daily data is converted to weekly frequency as described in the main text.

	Sov. CDS Spread			Fin. CDS Spread			Term Spread			Interbank Spread			LT Bond Yield		
	Mean	Med.	S.D.	Mean	Med.	S.D.	Mean	Med.	S.D.	Mean	Med.	S.D.	Mean	Med.	S.D.
Austria	56.71	39.24	55.44	140.82	143.35	87.36	154.00	173.20	91.82	46.83	36.60	39.82	301.51	339.45	115.45
Belgium	78.38	49.05	79.87	126.95	117.73	86.82	200.21	220.10	113.67	46.83	36.60	39.82	332.38	372.60	112.16
France	57.72	47.00	54.56	108.75	95.42	74.27	174.19	204.60	100.50	46.83	36.60	39.82	306.11	340.60	105.68
Germany	30.52	24.01	26.52	115.60	119.54	66.58	149.41	151.89	84.23	46.83	36.60	39.82	264.95	294.69	120.45
Ireland	211.63	129.18	247.95	439.74	261.85	490.77	232.48	185.85	185.89	46.83	36.60	39.82	487.61	448.75	227.42
Italy	154.80	125.70	137.80	186.38	146.14	144.53	257.62	307.95	142.56	46.83	36.60	39.82	423.90	432.40	103.89
Netherlands	38.50	33.92	32.88	137.98	142.48	80.25	167.24	174.35	94.36	46.83	36.60	39.82	290.34	326.20	116.52
Portugal	298.73	170.03	351.60	345.83	247.19	347.87	349.67	339.20	252.95	46.83	36.60	39.82	573.81	456.60	282.81
Spain	154.34	101.29	144.04	228.51	177.24	183.69	259.12	310.70	157.84	46.83	36.60	39.82	422.30	420.25	112.32
UK	41.53	35.70	33.49	127.49	123.39	97.16	152.55	172.47	137.34	40.58	23.57	40.08	334.80	340.91	112.27
US	26.77	28.61	18.82	184.57	165.86	145.31	199.01	221.01	115.35	51.43	30.08	54.91	314.37	296.90	103.34

NOTES: Values are reported in basis points. The data is sampled at monthly frequency over the period January 2006–July 2015.

	Variance Risk Premia		
	Mean	Med.	S.D.
Equity	1521.13	918.52	3508.89
Treasury	-164.05	-200.22	169.59

NOTES: Values are reported in variance units.

Table 2: Descriptive Statistics

	AT	BE	FR	DE	IE	IT	NL	PT	ES	UK
AT	**−50.00** [6] (−33.22)	−15.25 [8] (−19.15)	−36.31 [8] (−35.37)	−15.56 [8] (−11.30)	−1.49 [8] (−4.51)	−33.46 [8] (−29.57)	−12.04 [8] (−11.80)	−6.62 [8] (−12.26)	−11.29 [8] (−15.43)	−29.97 [8] (−17.01)
BE	−41.91 [8] (−35.31)	**−50.00** [8] (−28.84)	−47.94 [8] (−40.62)	−45.41 [8] (−45.69)	−2.55 [8] (−4.88)	−39.79 [7] (−36.49)	−37.57 [8] (−30.75)	−12.84 [8] (−11.90)	−21.69 [8] (−19.68)	−26.21 [8] (−39.27)
FR	−26.21 [8] (−12.89)	10.34 [8] (12.81)	**−50.00** [8] (**−29.16**)	−39.64 [8] (−28.49)	0.47 [8] (0.18)	−24.60 [8] (−12.91)	−30.79 [8] (−14.88)	−10.83 [8] (−6.31)	−13.73 [8] (−7.05)	−36.86 [8] (−30.16)
DE	−20.21 [8] (−3.08)	5.07 [8] (2.14)	−33.69 [8] (−26.01)	**−50.00** [8] (**−37.70**)	0.67 [8] (1.00)	−16.27 [8] (−7.64)	−16.53 [8] (−10.06)	−4.52 [8] (−4.58)	−0.20 [8] (3.00)	−14.55 [8] (−14.26)
IE	−160.08 [8] (−92.57)	−44.94 [8] (24.38)	−187.10 [8] (−104.92)	−174.13 [8] (−76.28)	**−50.00** [8] (−9.58)	−82.95 [8] (−114.16)	−122.91 [8] (−27.92)	−25.38 [8] (−49.56)	−12.69 [8] (−9.69)	−110.67 [8] (1.32)
IT	−16.97 [8] (−13.03)	11.92 [8] (0.41)	−52.55 [8] (−44.27)	−44.72 [8] (−64.12)	1.78 [8] (1.52)	**−50.00** [8] (**−25.54**)	−27.11 [8] (−29.52)	−12.56 [8] (−16.68)	−22.15 [8] (−7.75)	−67.46 [8] (−59.90)
NL	−34.44 [8] (−20.63)	−13.59 [8] (−8.07)	−45.34 [8] (−35.34)	−42.46 [8] (−37.36)	−0.91 [8] (−1.40)	−24.64 [8] (−16.11)	**−50.00** [7] (−30.97)	−8.22 [8] (−6.77)	−5.66 [8] (−2.70)	−17.84 [8] (−18.84)
PT	−69.84 [8] (−39.68)	10.44 [8] (36.00)	−92.01 [8] (−58.93)	−102.69 [8] (−102.59)	1.95 [8] (3.29)	−84.13 [8] (−71.82)	−40.50 [8] (−4.53)	**−50.00** [8] (**−42.14**)	−24.24 [8] (−19.43)	−34.72 [8] (−21.53)
ES	−37.47 [8] (−8.86)	15.88 [8] (33.31)	−55.50 [8] (−31.63)	−48.30 [8] (−39.69)	−1.32 [8] (0.12)	−37.36 [8] (−17.26)	−25.70 [8] (7.39)	−18.73 [8] (−14.69)	**−50.00** [7] (−15.77)	−48.06 [8] (−32.51)
UK	−7.97 [8] (−4.30)	−1.98 [8] (−16.31)	−35.96 [8] (−21.54)	−9.17 [8] (3.82)	1.44 [8] (0.75)	−11.87 [8] (−14.33)	3.45 [8] (19.46)	0.35 [8] (2.48)	1.60 [8] (8.04)	**−50.00** [8] (**−49.25**)

NOTES: The $\{i,j\}$th cell of the table records the median impulse response of financial sector credit risk in country i to an idiosyncratic financial sector bailout shock occurring in country j. The value printed without parentheses is the impact impulse response ($h = 0$ weeks), while the value printed underneath in rounded parentheses is the one-month impulse response ($h = 3$ weeks after the week of impact). The $\{i,i\}$th cell of the table summarises the median impulse response of financial sector credit risk in country i to an idiosyncratic financial sector bailout shock occurring in country i having pooled all of the retained draws from each of the nine bilateral models in the ith column of the table (i.e. all models where shocks originating in country i are identified). To assist the reader, where the impact impulse response is negative (positive), the cell is shaded green (red), with the depth of shading indicating the relative magnitude of the impact impulse response. In addition, if the $(16\%, 84\%)$ interval associated with the median impulse response on impact or at $h = 3$ weeks does not include zero, then the value is printed in bold face. The value in square parentheses is a count of the number of times that the inference drawn from our baseline model is upheld in each of our four exclusion exercises. Values are reported in basis points.

Table 3: The Response of Financial Sector Credit Risk to Financial Sector Bailouts, on Impact and after One Month

	AT	BE	FR	DE	IE	IT	NL	PT	ES	UK
AT	**25.08** [4] (18.87)	10.96 [8] (8.32)	7.56 [8] (12.82)	20.39 [8] (7.62)	−0.63 [8] (−1.82)	−6.14 [8] (−1.90)	17.76 [8] (7.49)	−7.69 [8] (−9.56)	−3.98 [8] (−2.77)	−4.17 [8] (11.63)
BE	5.58 [8] (8.08)	**36.53** [8] (**36.34**)	6.64 [8] (9.34)	5.71 [8] (−7.18)	−1.57 [8] (−1.26)	−4.11 [8] (6.38)	2.00 [8] (2.32)	−11.44 [8] (−9.91)	−10.64 [8] (−1.99)	−7.08 [8] (−3.51)
FR	7.17 [8] (10.98)	19.71 [6] (19.77)	**18.88** [8] (**18.51**)	6.32 [8] (1.36)	−0.61 [8] (0.35)	−0.23 [8] (7.31)	4.90 [8] (5.54)	−8.09 [8] (−4.69)	−5.08 [8] (2.53)	−2.73 [8] (−1.44)
DE	4.58 [8] (9.84)	10.66 [8] (**14.02**)	8.23 [6] (11.60)	**14.78** [8] (8.21)	0.12 [8] (0.81)	2.02 [8] (6.63)	8.26 [8] (10.32)	−2.33 [8] (−1.37)	−0.12 [8] (1.70)	6.72 [8] (12.18)
IE	−44.37 [8] (−37.56)	−16.03 [8] (−13.98)	−67.76 [8] (−65.19)	−36.82 [8] (−78.14)	**18.13** [8] (**19.68**)	−27.28 [8] (−28.43)	−31.63 [8] (−49.71)	−13.40 [8] (−5.92)	−4.61 [8] (−0.21)	−44.23 [8] (−22.85)
IT	8.91 [8] (−8.63)	34.44 [8] (24.21)	−15.04 [8] (−8.78)	−21.22 [8] (−37.94)	3.20 [8] (2.85)	**26.46** [8] (**28.97**)	−11.57 [8] (−30.85)	−9.85 [8] (−8.72)	2.77 [8] (2.65)	−33.26 [8] (−27.50)
NL	10.80 [8] (12.26)	13.70 [8] (12.05)	12.79 [8] (12.70)	11.64 [8] (2.81)	0.85 [8] (1.00)	0.85 [8] (2.70)	**17.85** [8] (**14.69**)	−3.92 [8] (−1.45)	−0.26 [8] (1.04)	2.28 [8] (6.67)
PT	−48.30 [8] (−41.07)	−33.56 [8] (−33.92)	−107.97 [8] (−123.93)	−75.70 [8] (−105.88)	7.55 [8] (7.27)	−55.04 [8] (−58.05)	−72.53 [8] (−71.56)	**59.27** [8] (**53.58**)	−12.13 [8] (−17.70)	−60.90 [8] (−89.58)
ES	6.01 [8] (17.69)	34.23 [8] (41.17)	−6.03 [8] (8.89)	−17.54 [8] (−6.98)	0.21 [8] (5.12)	13.29 [8] (25.42)	−0.52 [8] (3.78)	−6.87 [8] (2.95)	**24.62** [8] (**37.45**)	−19.24 [8] (−0.21)
UK	8.50 [8] (9.82)	4.23 [8] (0.84)	1.14 [8] (3.14)	15.34 [8] (7.27)	0.70 [8] (0.38)	0.63 [8] (−0.35)	14.54 [8] (9.50)	−2.18 [8] (−2.14)	2.31 [8] (1.26)	**16.24** [8] (**13.28**)

NOTES: The $\{i,j\}$th cell of the table records the median impulse response of sovereign credit risk in country i to an idiosyncratic financial sector bailout shock occurring in country j. The value printed without parentheses is the impact impulse response ($h = 0$ weeks), while the value printed underneath in rounded parentheses is the one-month impulse response ($h = 3$ weeks after the week of impact). The $\{i,i\}$th cell of the table summarises the median impulse response of sovereign credit risk in country i to an idiosyncratic financial sector bailout shock occurring in country i having pooled all of the retained draws from each of the nine bilateral models in the ith column of the table (i.e. all models where shocks originating in country i are identified). To assist the reader, where the impact impulse response is negative (positive), the cell is shaded green (red), with the depth of shading indicating the relative magnitude of the impact impulse response. In addition, if the $(16\%, 84\%)$ interval associated with the median impulse response on impact or at $h = 3$ weeks does not include zero, then the value is printed in bold face. The value in square parentheses is a count of the number of times that the inference drawn from our baseline model is upheld in each of our four exclusion exercises. Values are reported in basis points.

Table 4: The Response of Sovereign Credit Risk to Financial Sector Bailouts, on Impact and after One Month

	AT	BE	FR	DE	IE	IT	NL	PT	ES	UK
AT	**77.85** [8] (**83.27**)	**59.79** [8] (52.95)	**76.30** [8] (57.25)	**159.96** [8] (90.35)	8.04 [8] (5.53)	15.75 [8] (14.41)	**144.51** [5] (116.50)	4.44 [8] (3.70)	20.94 [8] (15.41)	**137.31** [8] (**147.67**)
BE	**73.31** [8] (**75.88**)	**59.00** [8] (**63.90**)	**96.17** [8] (**92.11**)	**191.66** [8] (**155.82**)	13.25 [8] (10.85)	**23.41** [8] (**24.23**)	**180.96** [8] (**146.46**)	8.97 [7] (9.04)	**27.19** [8] (**25.34**)	**117.19** [8] (**107.16**)
FR	**59.97** [7] (**47.17**)	**64.72** [8] (**52.62**)	**82.69** [8] (**65.10**)	**181.93** [8] (**106.32**)	11.96 [8] (9.66)	**27.16** [8] (**23.60**)	**128.32** [8] (**95.56**)	8.83 [8] (6.93)	**27.69** [8] (**25.50**)	70.65 [8] (38.71)
DE	**43.65** [8] (**35.20**)	**37.58** [8] (**31.18**)	**57.61** [8] (**45.71**)	**131.57** [8] (**93.07**)	5.25 [8] (4.94)	**17.60** [4] (**17.08**)	**104.29** [8] (**85.65**)	3.70 [8] (3.37)	**18.44** [7] (**16.85**)	**80.93** [8] (**82.54**)
IE	223.27 [8] (117.98)	222.40 [8] (128.23)	260.34 [8] (108.73)	575.37 [8] (348.15)	**144.26** [8] (**114.97**)	87.75 [8] (42.55)	439.23 [8] (249.43)	**67.56** [8] (33.80)	101.47 [8] (47.68)	495.67 [8] (354.53)
IT	76.88 [8] (51.92)	**82.53** [8] (64.69)	**132.02** [8] (**111.04**)	**245.27** [5] (173.18)	22.53 [8] (16.80)	**46.92** [8] (**47.48**)	**162.18** [8] (137.64)	**17.64** [8] (12.52)	**48.07** [8] (**48.48**)	68.28 [8] (42.38)
NL	**64.14** [8] (**66.40**)	**54.06** [8] (**53.27**)	**75.13** [8] (**67.53**)	**166.90** [8] (**139.71**)	9.70 [8] (9.28)	**22.14** [8] (**21.92**)	**126.54** [8] (**120.33**)	3.83 [8] (3.42)	**22.78** [7] (**20.81**)	**111.33** [8] (**121.04**)
PT	89.97 [8] (81.28)	104.97 [8] (88.98)	156.92 [8] (151.08)	**356.77** [8] (246.33)	**49.56** [8] (**59.32**)	**60.17** [8] (52.40)	190.64 [8] (160.10)	**39.58** [8] (**40.14**)	**77.84** [8] (**74.58**)	146.95 [8] (112.52)
ES	90.66 [8] (65.88)	**92.15** [8] (62.92)	**141.09** [8] (**118.68**)	**282.81** [5] (206.97)	**29.45** [8] (22.37)	**54.69** [8] (**47.14**)	**184.13** [8] (115.97)	**19.32** [8] (9.89)	**53.00** [8] (**61.52**)	93.94 [8] (27.31)
UK	**68.39** [8] (**72.45**)	**50.64** [5] (45.37)	**67.62** [8] (57.48)	**144.60** [5] (115.57)	16.12 [8] (16.03)	18.66 [8] (16.38)	**126.47** [8] (**134.26**)	6.32 [8] (3.31)	21.97 [4] (18.35)	**106.37** [8] (**162.15**)

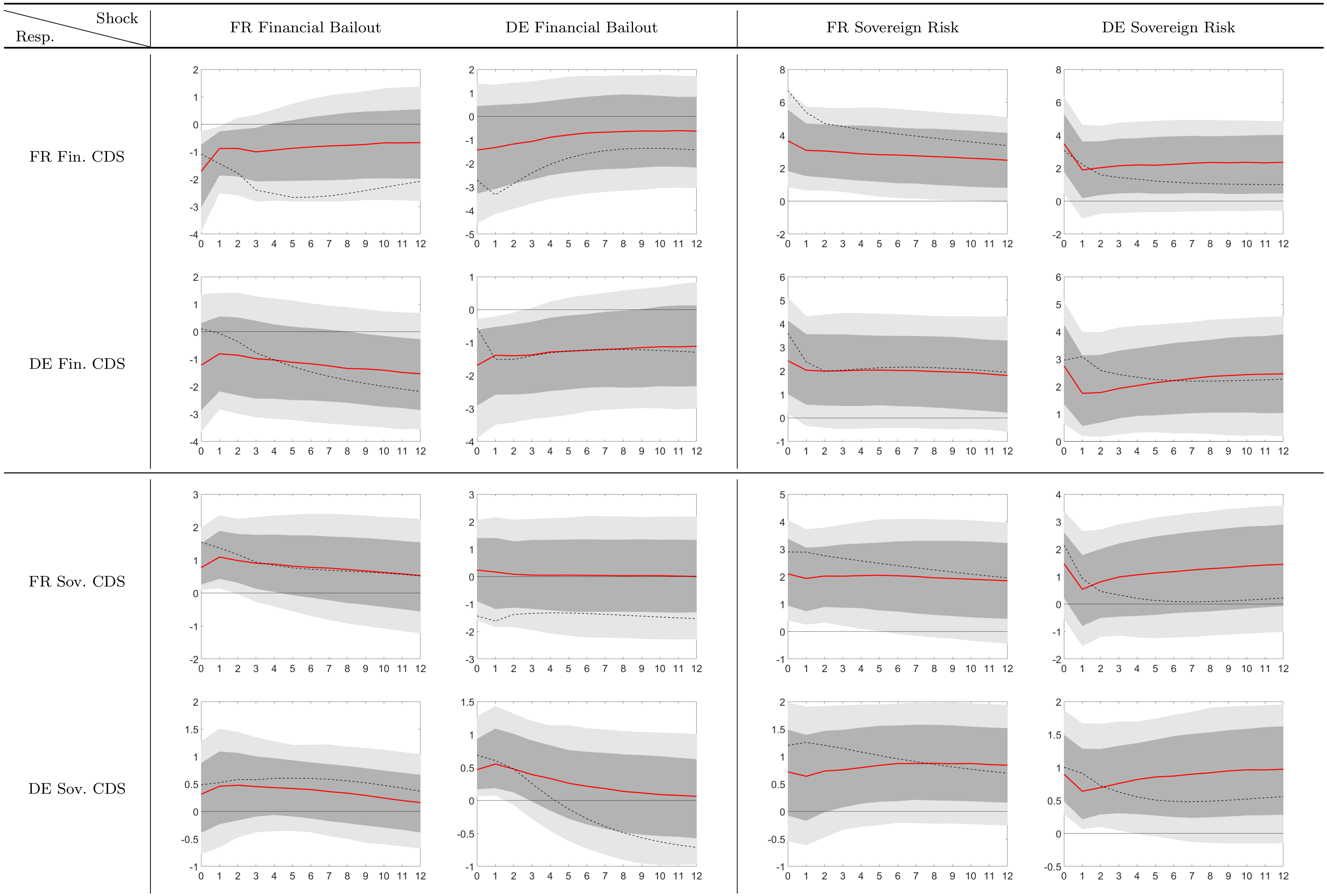
Notes: The $\{i,j\}$th cell of the table records the median impulse response of financial sector credit risk in country i to an idiosyncratic adverse sovereign risk shock occurring in country j. The value printed without parentheses is the impact impulse response ($h = 0$ weeks), while the value printed underneath in rounded parentheses is the one-month impulse response ($h = 3$ weeks after the week of impact). The $\{i,i\}$th cell of the table summarises the median impulse response of financial sector credit risk in country i to an idiosyncratic adverse sovereign risk shock occurring in country i having pooled all of the retained draws from each of the nine bilateral models in the ith column of the table (i.e. all models where shocks originating in country i are identified). To assist the reader, where the impact impulse response is negative (positive), the cell is shaded green (red), with the depth of shading indicating the relative magnitude of the impact impulse response. In addition, if the $(16\%, 84\%)$ interval associated with the median impulse response on impact or at $h = 3$ weeks does not include zero, then the value is printed in bold face. The value in square parentheses is a count of the number of times that the inference drawn from our baseline model is upheld in each of our four exclusion exercises. Values are reported in basis points.

Table 5: The Response of Financial Sector Credit Risk to Sovereign Risk Shocks, on Impact and after One Month

	AT	BE	FR	DE	IE	IT	NL	PT	ES	UK
AT	**50.00** [8] (**45.51**)	33.11 [5] (32.49)	38.53 [8] (33.98)	66.52 [5] (39.98)	5.59 [8] (4.58)	11.16 [8] (11.64)	64.52 [8] (52.09)	2.43 [8] (1.96)	9.32 [8] (7.12)	51.23 [8] (55.62)
BE	26.67 [8] (17.26)	**50.00** [8] (**45.15**)	**47.63** [8] (28.54)	93.54 [7] (47.06)	8.52 [8] (9.18)	15.31 [8] (15.94)	50.53 [8] (32.13)	4.06 [8] (4.77)	15.20 [8] (18.02)	7.70 [8] (4.41)
FR	26.05 [8] (20.97)	**30.15** [8] (**28.12**)	**50.00** [8] (**39.84**)	**73.83** [8] (47.30)	5.55 [8] (6.46)	13.72 [4] (**13.63**)	49.83 [8] (40.87)	4.06 [8] (4.51)	12.29 [4] (**13.09**)	9.48 [8] (2.42)
DE	**16.05** [8] (**17.36**)	14.06 [8] (**16.74**)	16.42 [8] (**17.90**)	**50.00** [8] (**38.13**)	3.48 [8] (3.93)	5.35 [7] (**6.91**)	**32.27** [6] (**34.41**)	1.97 [8] (2.42)	5.49 [8] (6.68)	22.48 [8] (24.40)
IE	15.97 [8] (−11.54)	16.16 [8] (−5.24)	0.91 [8] (−32.69)	112.06 [8] (8.30)	**50.00** [8] (**59.07**)	20.26 [8] (10.10)	33.09 [8] (−33.35)	19.43 [8] (18.48)	22.55 [8] (6.26)	6.66 [8] (−58.04)
IT	34.91 [8] (7.05)	51.55 [8] (37.04)	**85.17** [7] (**67.92**)	162.52 [8] (89.79)	9.78 [8] (7.96)	**50.00** [8] (**45.64**)	82.08 [8] (43.55)	9.27 [8] (8.85)	31.48 [8] (26.84)	−30.23 [8] (−55.65)
NL	**20.41** [7] (**19.87**)	**18.38** [8] (**18.98**)	**21.76** [8] (19.12)	**45.33** [8] (37.56)	3.91 [8] (4.44)	7.98 [2] (8.11)	**50.00** [8] (**44.71**)	1.74 [8] (1.60)	7.04 [8] (7.70)	29.29 [8] (34.29)
PT	−13.01 [8] (−19.94)	−2.04 [8] (10.75)	10.62 [8] (0.99)	8.40 [8] (−92.42)	9.10 [8] (12.69)	4.96 [8] (10.51)	−54.54 [8] (−71.31)	**50.00** [8] (**44.18**)	28.21 [8] (26.79)	−90.36 [8] (−128.82)
ES	48.08 [8] (24.97)	48.49 [8] (28.11)	**87.83** [8] (57.54)	164.33 [8] (97.88)	12.73 [8] (8.83)	**35.99** [8] (**28.84**)	76.60 [8] (43.53)	10.53 [8] (9.14)	**50.00** [8] (**46.22**)	−18.93 [8] (−38.42)
UK	17.55 [8] (13.38)	15.55 [8] (14.75)	15.75 [8] (12.11)	42.80 [4] (37.84)	5.67 [8] (4.14)	5.50 [8] (4.53)	**38.09** [7] (32.87)	1.93 [8] (1.88)	5.89 [8] (4.82)	**50.00** [8] (**47.32**)

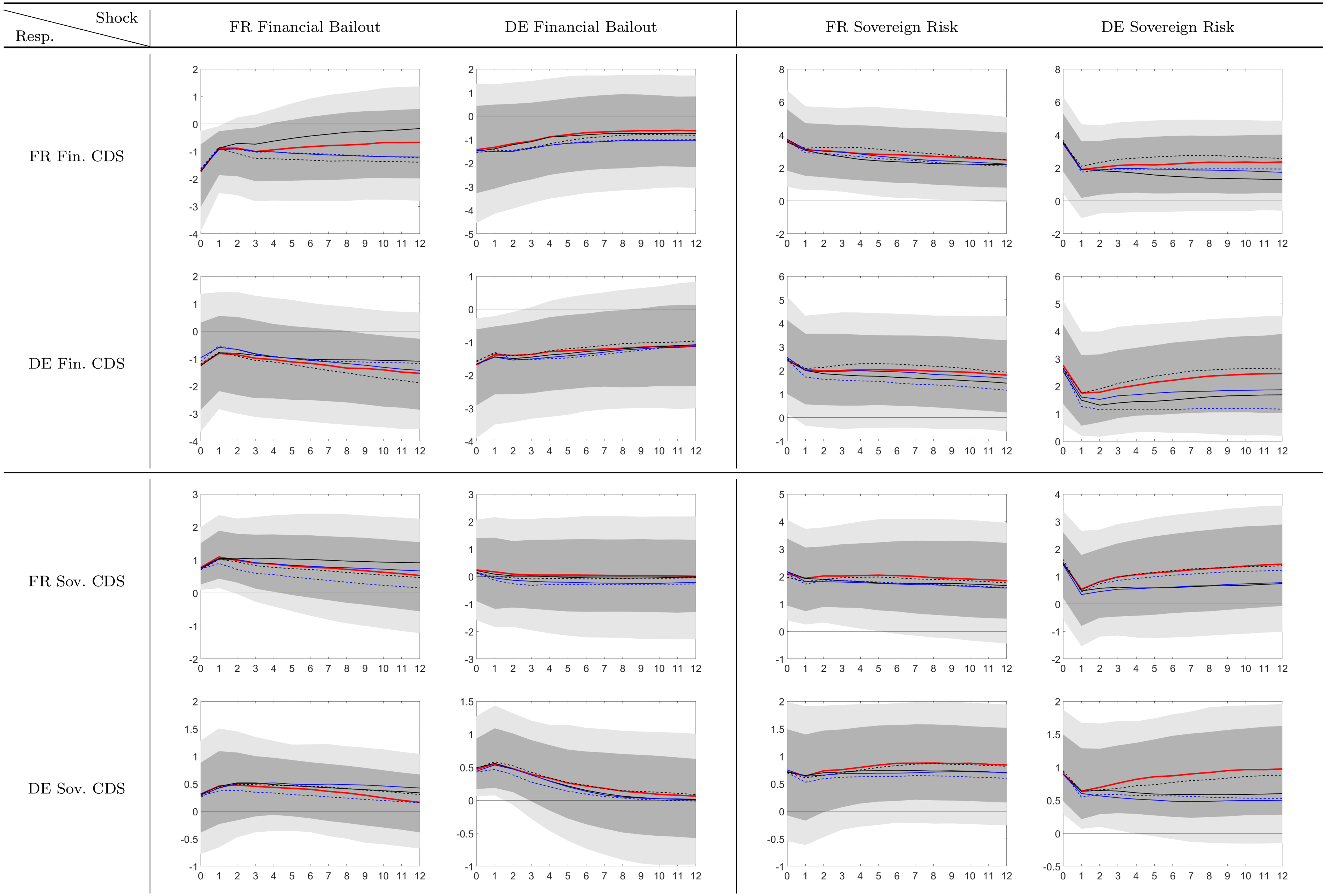
NOTES: The $\{i, j\}$th cell of the table records the median impulse response of sovereign credit risk in country i to an idiosyncratic adverse sovereign risk shock occurring in country j. The value printed without parentheses is the impact impulse response ($h = 0$ weeks), while the value printed underneath in rounded parentheses is the one-month impulse response ($h = 3$ weeks after the week of impact). The $\{i, i\}$th cell of the table summarises the median impulse response of sovereign credit risk in country i to an idiosyncratic adverse sovereign risk shock occurring in country i having pooled all of the retained draws from each of the nine bilateral models in the ith column of the table (i.e. all models where shocks originating in country i are identified). To assist the reader, where the impact impulse response is negative (positive), the cell is shaded green (red), with the depth of shading indicating the relative magnitude of the impact impulse response. In addition, if the $(16\%, 84\%)$ interval associated with the median impulse response on impact or at $h = 3$ weeks does not include zero, then the value is printed in bold face. The value in square parentheses is a count of the number of times that the inference drawn from our baseline model is upheld in each of our four exclusion exercises. Values are reported in basis points.

Table 6: The Response of Sovereign Credit Risk to Sovereign Risk Shocks, on Impact and after One Month



Notes: The red line in each figure is the median of the set of retained impulse responses which satisfy the identifying sign restrictions. The fine dashed black line is the median target impulse response, which we obtain as the impulse response function derived from the single draw of the sign restrictions algorithm that is closest to the median impulse response for all three of our identified shocks. The dark (light) gray band marks the central 68% (90%) of the distribution of retained impulse responses at each horizon. The vertical axis measures the response of the named variable in basis points while the horizontal axis records the time horizon in weeks, where the week in which the shock occurs is denoted week 0.

Figure 1: Pairwise Spillovers between France and Germany



Notes: The red line in each figure is the median of the set of retained impulse responses which satisfy the identifying sign restrictions. The dark (light) gray band marks the central 68% (90%) of the distribution of retained impulse responses at each horizon. The other four lines show the median impulse responses obtained from our four exclusion exercises as follows: black solid line – exercise (i); black dashed line – exercise (ii); blue solid line – exercise (iii); and blue dashed line – exercise (iv). The vertical axis measures the response of the named variable in basis points while the horizontal axis records the time horizon in weeks, where the week in which the shock occurs is denoted week 0.

Figure 2: Pairwise Spillovers between France and Germany: Comparison of the Benchmark Results with Exclusion Exercises (i)–(iv)