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Monetary policy, bank bailouts and the sovereign-bank risk nexus in the euro area¹

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Abstract

The paper analyses the empirical relationship between bank credit risk and sovereign credit risk in the euro area, using a system of simultaneous equations identified through heteroskedasticity. We first confirm a two-way causality between both risks, which amplifies initial credit risk shocks. We also document significant credit risk spillovers between sovereigns and banks in the periphery and the core countries. The paper then focuses on the impact of ECB non-standard monetary policy and bank bailout policies. We show that bailouts have reduced both risks. Monetary policy lowered in most but not all cases bank and sovereign risk.

Keywords: Credit risk, banks, sovereigns, monetary policy, bank bailout, heteroskedasticity, spillovers.

JEL No.: E52, G10, E60.

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1. Introduction

A key feature of the European crisis has been the ever closer link between sovereign and bank risk in the euro area. A major concern of policymakers was the feedback loop, in which adverse shocks to banks led to an increase in sovereign credit risk, and deterioration in sovereign risk, in turn, raised the credit risk of banks. Indeed, the correlation between credit default swap (CDS) spreads of banks and sovereigns rose from 0.1 in 2007 to 0.8 in 2013 (see Figure 1). The feedback loop has been blamed for deepening the recession, making an escape from the crisis more difficult.

Policymakers have adopted two principal approaches to dealing with mounting credit risks during the global financial and the subsequent European crisis. On the one hand, national governments implemented bank rescue policies, by providing capital injections, debt guarantees or deposit guarantees. On the other hand, monetary policy played a central role, by providing liquidity to banks on a massive scale and by intervening in sovereign debt markets through outright purchases or by giving an implicit guarantee against a speculative run.

Have these policies reduced sovereign and bank risk? The academic literature has analysed different channels through which these two policy approaches function. Concerning government policies for banks, there is a growing theoretical literature about how such policies affect bank risk and sovereign risk (see Allen et al., 2013; Cooper and Nikolov, 2013; Acharya, Drechsler and Schnabl, 2014; Leonello, 2017; Farhi and Tirole, forthcoming). Guarantees and capital injections can prevent bank runs, reduce uncertainty, and improve the stability of the financial system. However, if the bailouts are large and imply a challenge to the sustainability of public debt, then such a transfer of risk may actually worsen the outlook for the economy, and lead to sovereign and bank risk moving in lockstep.

There is also an extensive literature on the impact of unconventional monetary policy on financial markets during the crises. In particular, there is an expanding literature on the effects of the Federal Reserve's quantitative easing policies (see Gagnon et al., 2011; Fratzscher, Lo Duca and Straub, 2018) and of the European Central Bank's (ECB) policies (see Manganelli, 2012; Ghysels et al., 2016). The controversy in this literature is to what extent the unconventional policies have reduced risks to banks or to sovereigns, but these risks have mostly been analysed in isolation, without studying the mutual feedback effects. Finally, there is the possibility of a transmission of credit risk shocks across countries (see Popov and Van Horen, 2014; Breckenfelder and Schwaab, 2015; Caporin et al., 2017).

This paper starts with a general analysis of the causality between bank risk and sovereign risk, and whether there is evidence for cross-country spillovers within the euro area. It then provides an empirical test of the effect of bank rescue policies and non-standard policies of the ECB on the nexus between sovereigns and banks. The primary interest is whether such policies have reduced credit risks, or whether they

have intensified the crisis. It then goes into detail of this link, by investigating whether these policies have functioned differently for healthy versus crisis countries.

Using daily data for the period 2004-13, we employ a set of simultaneous equation models. To identify the structural shocks to the endogenous variables, we exploit the heteroskedasticity in the data, following Sentana and Fiorentini (2001) and Rigobon (2003). In short, this identification strategy exploits that changes in the volatility of the structural shocks contain additional information on the relation between the endogenous variables. For example, in a period of large bank risk shocks we learn more about the response of sovereign risk to bank risk as the covariance between both risks increases. During these periods, bank risk shocks are more likely to occur and can be used as a ‘probabilistic instrument’ (see Rigobon, 2003) to trace out the response of sovereign risk. To identify the effects of non-standard monetary policy and bank bailout announcements, we follow the ‘event study’ literature (see Kuttner, 2001; Gürkaynak, Sack and Swansson, 2005) and use dummy variables.

We find evidence for a two-way causality between sovereign risk and bank risk, both at the euro area level and in many member states individually. The bidirectional effect generates a strong feedback loop which amplifies initial credit risk shocks substantially. Shocks to sovereign or bank risk also lead to a drop in equity prices and an increase in credit risk of non-financial institutions. If one interprets these variables as reflecting the prospects of the real economy, then the implication is that the feedback loop between banks and sovereigns extends to the real economy. In terms of economic significance, variance decompositions show that shocks to bank and sovereign risk explain a sizable magnitude of the variability of the other variables.

The empirical findings further suggest that there are strong cross-country spillovers of sovereign and bank risk. The spillovers stem from shocks to sovereign risk in the euro area periphery to sovereigns in the core, from shocks to banks in the periphery to banks in the core, as well as vice versa. What is striking is that the impact of credit risk in the periphery on the core countries has been similar in magnitude or even larger than the effect in the opposite direction. This is important since the core countries constitute the much larger share of the euro area and its underlying sovereign debt market.

Finally, the empirical analysis shows that ECB non-standard policies and bank bailouts by national governments reduced in most but not all cases credit risks of banks and sovereigns. They also exerted a large and significant impact on equity returns and credit risk of non-financial firms. Announcements about the Securities Market Programme (SMP), consisting of government purchases, and Outright Monetary Transactions (OMT), implying the conditional commitment to purchase public debt, lowered sovereign and bank CDS spreads significantly. Similarly, bank bailout announcements by fiscal policy makers led to a strong decline in sovereign and bank spreads. For Longer-Term Refinancing Operations (LTROs), the results reveal an interesting yet intuitive difference between the announcement and the actual

implementation. While the former was taken as a disappointment by financial markets, with spreads increasing, the latter reduced sovereign and bank risk.

The paper contributes to a growing line of research on the sovereign-bank risk nexus. Several authors use time-series analysis and quantify the size of risk spillovers or risk transfers between both sectors (see Ejsing and Lemke, 2011; Alter and Schöler, 2012; Bicu and Candelon, 2013; De Bruyckere et al., 2013; Alter and Beyer, 2014). These papers do not aim at identifying causal effects, which we do using the heteroskedasticity of the data. This is a first main contribution of the paper.

Other papers use panel data analysis to study the importance of the financial sector (risk) for sovereign risk (see Mody and Sandri, 2012; Dieckmann and Plank, 2012; and Kallestrup, Lando and Murgoci, 2016) or the relevance of sovereign risk for bank risk (see Acharya, Drechsler and Schnabl, 2014). Both sets of papers provide evidence that the credit risk of one sector is important for risk of the other. Our results, based on a different methodology and data, confirm these findings. Using panel regressions with monthly data, Dieckmann and Plank (2012) show that the relative market capitalisation of the domestic financial sector and its asset growth prior to the global financial crisis covary with CDS spreads of western European sovereigns. We corroborate the link between financial sector risk and sovereign risk using CDS spreads as a measure of bank credit risks. Based on panel estimation with daily data, Archarya, Drechsler and Schnabl (2014) provide evidence that a 10 percent increase in sovereign CDS translates into a 0.9 percent increase in bank CDS. We also document significant risk spillovers in that reverse direction. However, our benchmark specification for the euro area suggests stronger effects: a 100 basis points increase in sovereign CDS leads to a 32 basis points increase in bank CDS.

Furthermore, our paper complements existing evidence by providing two novel insights into the relation between sovereign and bank risk. First, in the aforementioned papers the two directions of risk transmission have been analysed in isolation. Our econometric methodology instead allows us to analyse them jointly and disentangle the directions of causality. This means that we can for the first time directly compare the relative strength of the effect running in either direction, holding everything else constant. Our results reveal that the impact of sovereign risk is more than twice as large as the reverse effect. This suggests that sovereign risk has been the more important driver of the sovereign-bank risk nexus in the euro area during the global and European crisis and seems to be more important for the pricing of bank CDS than vice versa.

Second, our methodological approach allows us to quantify the amplification of sovereign and bank risk shocks due to the direct and contemporaneous feedback between both risks (and potentially other asset prices). Policy makers and market participants are well aware of the existence of such a feedback, in which exogenous increases in sovereign risk, say, push up bank risk, and the latter, in turn, raises sovereign risk, and so forth. But little is known about the strength and scope of the

resulting amplification of initial credit risk shocks.¹ Our estimates suggest that the amplification is strong. Original shocks to sovereign or bank risk are accentuated by up to 30 percent. Finally, our model also highlights through which other transmission channels initial credit risk shocks dissipate. The estimates suggest that shock propagation and amplification extends to the real economy, as measured by equity returns and the term spread.

The paper also relates to a rapidly evolving literature on unconventional monetary policy. Kozicki, Santor and Suchanek (2011) or Stone, Fujita and Ishi (2011) provide reviews for the US and UK. Regarding the euro area, many studies focus on the impact of ECB policy on macroeconomic variables such as monetary aggregates, credit volumes, inflation, or output (see Giannone et al, 2011; De Pooter, Martin and Pruitt, 2012). Only few papers concentrate on credit risks. Lucas, Schwaab and Xin (2014) propose an empirical framework to measure joint and conditional probabilities of credit events and investigate the impact of several non-standard ECB policies on them. Pelizzon et al. (2016) analyse the dynamic relationship between credit risk and liquidity in government bond markets and show that the relation is affected by ECB policy. Eser and Schwaab (2016) have access to daily data on government bond purchases by the Eurosystem and study their effects on sovereign yields and CDS. They find that the SMP led to a significant reduction in sovereign risk. Finally, Krishnamurthy, Nagel and Vissing (2018) use a Kalman-filter to decompose changes in sovereign yields around ECB policy events into changes in different yield components. They find that lower sovereign default risk is important to explain the decline in sovereign yields of stressed euro area countries.

By and large, we confirm the results of these papers. Our study extends the analysis of the financial impact of ECB non-standard policy along several dimensions. First, we estimate the dynamic effects of the policy measures and document that they had in many instances persistent effects on CDS spreads. Accounting for persistence is important because our estimates imply larger effects than the previous studies which focus on the immediate market impact. Second, our analysis also includes bank credit risk. This is relevant for policy makers because several ECB measures were targeted specifically at alleviating strains in the financial sector.

Finally, within the same model, we study the effects of bank bailout announcement. This facilitates a comparison of the dynamic asset market effects of the two principle policy approaches to the crises. The results suggest that bailout announcements reduced bank risk more than monetary announcements.

¹ Previous papers have estimated the overall effect of sovereign on bank risk, or vice versa. This impact already incorporates all direct contemporaneous feedback between both risks (and potentially other asset prices) and does not provide an estimate of the amplification of initial credit risk shocks.

2. Conceptual Framework

In this section, we first discuss potential transmission channels between bank risk, sovereign risk, and other asset prices. Then, we outline how non-standard monetary policy can affect asset prices.

2.1 Credit risks, the term spread, and equity returns

The literature discusses several potential links between bank risk and sovereign risk. We start with a discussion of the two main transmission channels of bank risk shocks to sovereign risk. According to a credit supply channel, banks reduce the supply of credit if they incur unexpected losses. This depresses investment, economic growth, the tax base and ultimately increases sovereign risk. Second, there exists a bailout channel (see Allen et al., 2013; Cooper and Nikolov, 2013; Acharya, Drechsler, Schnabl, 2014, Leonello, 2017; Farhi and Tirole, forthcoming). If sovereigns implicitly or explicitly guarantee banks' liabilities, bank risk shocks also increase sovereign risk. We thus expect a positive effect of bank risk shocks on sovereign risk.

Regarding the reverse direction, the literature discusses three main channels how sovereign risk shocks can be transmitted to bank risk. The first functions through prices of government bonds. If sovereign risk increases, bond prices fall and banks incur portfolio losses. In addition, bank funding conditions deteriorate as the value of (bond) collateral used in refinancing operations declines. Angeloni and Wolff (2012), Kallestrup, Lando and Murgoci (2013), and De Bruyckere et al. (2013) provide evidence for these mechanisms in the euro area. A second channel is based on rating ceilings according to which private entities cannot be rated higher than their sovereign. A downgrade of the sovereign then triggers a down rating of domestic banks. This increases bank risk since many investors are legally constrained regarding the rating structure of their portfolios. Arezki, Candelon and Sy (2011) find evidence for this channel. Finally, if sovereigns guarantee banks' liabilities, there is also a positive transmission of sovereign risk shocks to bank risk. These considerations imply that we also expect a positive effect of sovereign risk shocks on bank risk.

Taken together, if we find evidence of positive shock transmission in both directions, a feedback loop between bank risk and sovereign risk can emerge. The feedback can amplify initial shocks to both bank and sovereign risk and a main goal of the analysis is to quantify this amplification effect.

Concerning other potential drivers of bank risk and sovereign risk, shocks to credit risk of the non-financial corporate sector are likely to play an important role. These shocks are expected to increase both bank and sovereign risk as the number of non-performing loans rises and the tax base declines. Ejlsing and Lemke (2011) show that a common credit risk factor accounts for a large part of the variation of sovereign and bank CDS spreads. For sovereign risk, Gerlach and Schulz (2010) find that the first principal component explains three fourths of the variance. Regarding bank risk,

Acharya, Drechsler and Schnabl (2014) identify the iTraxx Europe CDS index as important. Guided by these findings, we include non-financial credit risk into the model.

In addition, the term spread, defined as the difference between the ten-year and three-month rate on government bonds, consists of two components which are important for the evolution of sovereign and bank risk. The first component reflects business cycle expectations. In recessions, public deficits widen and the number of non-performing loans grows. Hence, this component implies a negative co-movement between the term spread and credit risk of sovereigns and banks. The second component reflects sovereign risk perception in the bond market and implies a positive co-movement. Palladini and Portes (2011) suggest that the European sovereign CDS market moves ahead of its cash market while Fontana and Scheicher (2010) find that price discovery can run in both directions. Overall, we expect the business cycle component to dominate in non-stressed countries and the credit component to prevail in stressed countries.

Finally, Longstaff et al. (2011) and Dieckmann and Plank (2012) show that stock returns are one important determinant of sovereign CDS spreads. As equity markets reflect real growth expectations and risk appetite of investors, we expect that shocks to equity returns lower credit risk in the sovereign, banking, and non-financial sector.

2.2 Non-standard monetary policy and credit risk

In this subsection, we discuss potential transmission channels of important ECB policies to sovereign risk and bank risk. We consider LTROs with maturity 6-12 and 12-36 months, respectively, SMP, and OMT. LTROs provide liquidity to banks. Their objective is to restore the functioning of impaired interbank markets. Lower funding risk reduces the probability that liquidity turns into default risk. Hence, we expect LTROs to reduce bank credit risk. If this translates into higher credit supply, output and tax revenues, they are likely to also lower sovereign risk. Moreover, banks could use the additional funds to buy government debt, also reducing sovereign risk (see Acharya and Steffen, 2015).

SMP and OMT are supposed to support the depth and liquidity of secondary government bond markets through purchases of government securities. There are two main transmission channels. First, through a portfolio channel, ECB purchases lower the outstanding amount of debt securities. If assets are imperfect substitutes, bond prices increase (see Tobin, 1958; or Vayanos and Vila, 2009). Second, large potential demand for bonds by the ECB lowers liquidity premia (see De Pooter, Martin and Pruitt 2012). Both channels, by lowering yields, reduce public financing costs and hence credit risk. To the extent that banks hold government debt, or that sovereigns guarantee for banks liabilities, they are expected to also reduce bank risk.

3. Empirical Methodology and Data

In this section, we present our baseline econometric model, the data, the identification strategy, and the estimation procedure.

3.1 Model specification and data

Our baseline specification is a system of five endogenous variables which includes credit default risk of sovereigns, banks, and non-financial corporations, the term spread, and an equity market index in the vector of endogenous variables y_t :

$$Ay_t = \tilde{c} + \tilde{A}_1 y_{t-1} + \Lambda + \tilde{A}_p y_{t-p} + \tilde{\Gamma}_0 x_t + \Lambda + \tilde{\Gamma}_q x_{t-q} + \varepsilon_t. \quad (1)$$

Here, x_t is a vector of exogenous variables, consisting of policy measures and other control variables, \tilde{c} a vector of constants, \tilde{A}_i with $i = 1, \dots, p$ and $\tilde{\Gamma}_j$ with $j = 0, \dots, q$ are coefficient matrices, and ε_t a vector of structural shocks with diagonal variance matrix $\Sigma_\varepsilon = E(\varepsilon_t \varepsilon_t')$. The focus of the paper is on A and $\tilde{\Gamma}_j$. These matrices contain the contemporaneous effects of structural shocks on the endogenous variables and the effects of the policy measures, respectively. The identification strategy regarding the policy effects is based on the ‘event study’ approach. The identifying assumptions are detailed in Section 5. In short, we focus only on the most important monetary policy and bank bailout announcements in the sample and assume that they did not occur in response to contemporaneous market prices but in response to generally worsening financial and economic conditions, and can thus be treated as exogenous with respect to the other daily variables in the system.

We collect daily data until 31 July 2013. An online appendix provides details on all variables, definitions and sources as well as summary statistics. This appendix also contains all robustness tests and additional material that we mention throughout the paper. To measure credit risk, we use CDS spreads on senior unsecured debt, modified-modified restructuring, mid spread, with maturity of five years. This is the most liquid maturity segment. We collect sovereign spreads for Austria, Belgium, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal and Spain. The country sample is restricted by the joint availability of CDS data for all three credit sectors in a given country. For several countries, like Belgium or Greece, there is only one non-sovereign entity for which data is available. For the non-financial sector, we use all CDS data available in Datastream. This yields data for 56 entities.

We compute unweighted averages for each sector for the euro area and at the country level. We use unweighted indices in analogy to the main CDS indices used in practice - the family of iTraxx indices. Our main results are insensitive to using GDP-weighted averages, to excluding one country at a time, or to using alternative sector CDS indices for banks and nonfinancial companies. While for some countries CDS data for all three sectors are available from 2004 onwards, data for Ireland become available

only after 2006. This restricts the analysis for the euro area level to the years 2006-2013. Moreover, we exclude Greece from euro area averages since its sovereign CDS data end after the default in 2011. Concerning equity markets, we employ the Euro Stoxx 50 return index for the euro area and benchmark local return indices when looking at individual countries.

For the monetary policy and bank bailout policies we use dummy variables to indicate days when the ECB or national policy makers made important announcements. For monetary policy, we differentiate between announcements regarding LTROs with 6-12 months maturity, LTROs with 36 months, SMP, and OMT. For LTROs, we also estimate the implementation effects, which are measured in billion euros. We do not aim at estimating the implementation effects of SMP since public data on purchases are available only at the weekly frequency. Concerning bailout announcements, we consider three types: debt guarantees, deposit guarantees, and capital injections.

The model also contains several control variables. To capture uncertainty in euro area financial markets, we employ the VDax New. This index measures option-implied volatility in the German stock market. To control for money market conditions, we use the difference between the three month Euribor and the corresponding OIS spread. Both exogenous variables could potentially also be treated as endogenous. We include them exogenously, however, as the system with five endogenous variables is already relatively large and challenging to estimate. Moreover, we are more interested in the relation between the CDS market and the two main other asset classes, stocks and bonds, which better reflect the prospects of the real economy. Our results are similar when we either treat both variables as endogenous or drop them from the model.

To correct for macroeconomic news shocks, we include the unexpected component of 12 economic indicators for the euro area. Finally, we estimate the model in first differences of CDS spreads, the term spread, the VDax New and the interbank risk premium, and in first log differences of the stock market indices.

3.2 Identification strategy

For estimation of the structural parameters, we pre-multiply equation (1) by A^{-1} and define $c \equiv A^{-1}\tilde{c}$, $A_i \equiv A^{-1}\tilde{A}_i$, $\Gamma_j \equiv A^{-1}\tilde{\Gamma}_j$, which yields

$$y_t = c + A_1 y_{t-1} + \Lambda + A_p y_{t-p} + \Gamma_0 x_t + \Lambda + \Gamma_q x_{t-q} + u_t. \quad (2)$$

Here, u_t is a vector of reduced-form residuals. It is related to the structural shocks according to $u_t = A^{-1}\varepsilon_t$. The matrices c , A_i , Γ_j and Σ_u of model (2) can be estimated consistently by ordinary least squares. To recover the structural parameters, we need to identify the impact matrix A . From (1) and (2), $\Sigma_u = A^{-1}\Sigma_\varepsilon(A^{-1})'$. In this system, the number of unknowns is larger than the number of independent equations. As additional information we therefore exploit the heteroskedasticity in the data.

For illustration of this approach, we consider a bivariate system without constants, lags, or exogenous variables. If there are two regimes in the variances of the structural shocks, low (L) and high (H), the system reads

$$\begin{pmatrix} \omega_{11}^L & \omega_{12}^L \\ \cdot & \omega_{22}^L \end{pmatrix} = \begin{pmatrix} 1 & -a_{12} \\ -a_{21} & 1 \end{pmatrix}^{-1} \begin{pmatrix} \sigma_{11}^L & 0 \\ 0 & \sigma_{22}^L \end{pmatrix} \begin{pmatrix} 1 & -a_{12} \\ -a_{21} & 1 \end{pmatrix}^{-1'} \\ \begin{pmatrix} \omega_{11}^H & \omega_{12}^H \\ \cdot & \omega_{22}^H \end{pmatrix} = \begin{pmatrix} 1 & -a_{12} \\ -a_{21} & 1 \end{pmatrix}^{-1} \begin{pmatrix} \sigma_{11}^H & 0 \\ 0 & \sigma_{22}^H \end{pmatrix} \begin{pmatrix} 1 & -a_{12} \\ -a_{21} & 1 \end{pmatrix}^{-1'}.$$

While an additional volatility regime adds two new parameters to be estimated, the variances of the structural shocks, it also provides three new moments. Now, there are as many independent equations as unknowns and the system can be solved.

Identification relies on several important assumptions. First, the different types of structural shocks are uncorrelated. Second, the structural shocks are uncorrelated over time. Third, the ratio of the shock variances changes significantly across regimes. Fourth, A is constant across regimes.

The first assumption is standard in structural vector autoregressions (VAR). Moreover, it is likely to hold in our setup as we control for common effects through macroeconomic news and other exogenous variables which can affect the endogenous variables simultaneously. To make the second assumption likely to hold, we include five lags of the endogenous variables although information criteria suggest less lags. This relatively high number of lags ensures that the reduced-form residuals are largely free of autocorrelation. For all variables except for stock returns, Portmanteau tests for lags 1 to 1-5 do not reject the null hypothesis of no autocorrelation at the 10% level.

The third assumption can be tested after estimation by formally evaluating the inferred relative changes in volatility. While theoretically two regimes can be enough for just-identification, in practice larger systems tend to require more regimes as there are more shocks to be disentangled. More regimes enhance the likelihood of finding a regime for each shock where that shock changes significantly in volatility vis-à-vis the other shocks. For this reasons, we define five volatility regimes (see below). While we show that already four regimes are sufficient for identification, having one additional regime has the advantage that the fourth identifying assumption, the constancy of A , becomes overidentifying. It can then be tested by comparing the impact effects estimated on different subsets of volatility regimes, as in Rigobon and Sack (2003).

Two prominent alternative identification strategies exist. First, we could use zero restrictions on A , arguing for delayed responses of some endogenous variables to others. This seems too restrictive, however, as financial markets are likely to respond to each other in nearly continuous time. Second, we could use sign restrictions. They allow for contemporaneous effects among all variables. However, neither theory nor

empirical evidence gives unambiguous predictions for the signs of several key parameters and it is our aim to determine the signs empirically. On the other hand, a well-known drawback of our agnostic identification approach is that the structural shocks are more difficult to interpret as they are not based on a priori assumptions. We address this issue below. There, we provide a detailed analysis of the identified shock series and show that they correspond well with our interpretation.

3.3 Volatility regimes, estimation, and identification tests

Before estimation, we need to determine the volatility regimes. We use a narrative approach, following Rigobon (2003). The global financial and the European debt crisis provide a natural framework for this methodology as they are characterised by strong, persistent and datable increases in volatility. Using media reports and following previous studies (see BIS, 2009; Alter and Schöler, 2012; Mody and Sandri, 2012; Alter and Beyer, 2014), we divide the sample into five regimes.

The first regime runs until 1 July 2007 when the subprime crisis emerged in the US. Regime 2 captures the period when these tensions were transmitted to Europe. The subprime crisis culminated in the bankruptcy of Lehman Brothers on 15 September 2008, which marks the beginning of regime 3, with subsequent major stress in western financial markets. Thereafter, a denotable European dimension of the banking crisis emerged with large scale state support for the banking sector. Bailouts occurred in nearly all euro area member states, including liability guarantees, capital injections, and deposit insurances worth several trillion euro. Regime 4 begins with the bailout of Greece on 2 May 2010. It contains the bailouts of Ireland and Portugal, as well as the haircut on Greek sovereign debt. The last regime starts with the announcement of OMT by the ECB on 26 July 2012, which marks the beginning of the end of the crisis.

We use two types of structural break tests to assess whether these choices are supported by the data. First, we employ supremum Wald tests for a break in the level of the squared reduced-form residuals of the baseline five-variable VAR model at an unknown break date, using a trimming rate of 15%. The Wald statistics peak at the start of regime 3 and 4, respectively. Due to trimming, the statistics cannot be used to see whether there are more regime changes at the start or the end of the sample. But this can be done ex post (also for regimes 3 and 4) by testing the constancy of the reduced-form covariance matrix. We follow Lanne and Lütkepohl (2008) and perform likelihood ratio tests on the null hypothesis that every two neighbouring regimes have the same covariance matrix, and on the joint null that all regimes have the same covariance matrix. Table 1 shows that the assumption of homoscedasticity is clearly rejected. The data prefer a model with heteroskedasticity.

As a robustness test, we study whether the main results are sensitive to using two alternative, purely statistical approaches to defining volatility regimes. We also verify that the main results are not affected by including dummies for days of the week that control for ‘institutional heteroskedasticity’. Finally, it should be noted that

identification through heteroskedasticity yields consistent estimates even if the regimes are misspecified (see Rigobon, 2003).

Changes in volatility identify the system only up to a row permutation, however. To pin down the permutation that reflects the true underlying economic model, we need to impose one sign restriction, independent of the number of volatility regimes ≥ 2 . We assume that a positive shock to equity returns lowers credit risk of non-financial corporates. We choose this assumption because higher expected profits in the overall corporate sector are likely to lower credit risk of non-financials. Moreover, the constraint does not restrict the signs of the bidirectional effect between sovereign risk and bank risk, and it is not binding in the estimation.

Following Ehrmann et al. (2011), we estimate the parameters of A by minimising

$$\|g'g\| = \sqrt{\text{tr}(g'g)} = \sqrt{\text{vec}(g)' \text{vec}(g)}, \quad \text{with } g = \sum_{i=1}^5 A \Sigma_{u,i} A' - \Sigma_{\varepsilon,i}$$

where $\Sigma_{\varepsilon,i}$ is the variance of structural shocks and $\Sigma_{u,i}$ is the covariance matrix of the estimated reduced-form residuals for regime $i=1, \dots, 5$. We perform 1000 bootstrap replications. In each replication, we use the regime-specific covariance matrices to generate new data from which we obtain estimates using the minimisation procedure. We judge the significance by computing the share of estimates beyond zero.

As mentioned above, one assumption for identification is that the estimated variance-ratios of the uncorrelated structural shocks are sufficiently distinct across regimes (see Herwartz and Lütkepohl, 2014). This can be assessed after estimation in two ways. First, Table 2 contains the ratios of all shock pairs for all volatility regimes. It shows that for each pair there is a regime where the standard error of the ratio does not overlap with the standard error of any of the other ratios. In regime 4, for example, the volatility of sovereign risk shocks increases overproportionally relative to the volatility of the other shocks. Similarly, one can find regimes for bank risk and the other shocks with significant relative volatility shifts. This shows also that the regime definition not only allows disentangling sovereign and bank risk shocks from the other shocks, but also the latter from each other. The volatility ratio of term spread and stock market shocks, for example, is significantly lower in regime 3. Second, we use a linear Wald test for each shock pair to assess whether the variance-ratios are the same across regimes. Table 3 shows that the null hypothesis of no changes in volatility is strongly rejected by the data in all cases. The model is identified.

4. Credit Risk Propagation in the Euro Area

We first analyse the euro area as one block before we look at single countries. Then, we explore the relation between the euro area core and periphery. Throughout, we standardise the endogenous variables prior to estimation by subtracting the mean and

dividing by the standard deviation. This facilitates a direct comparison of the effects across variables and countries. Moreover, it reduces the computational challenges of the minimisation procedure as the parameters are of similar order of magnitude.

4.1 Analysis of the euro area as a single entity

We start the analysis at the euro area level. The upper part of Table 4 shows the estimated *direct* effects of a structural shock of one (in columns) on the endogenous variables (in rows). All other variables are kept constant within period and there is no contemporaneous feedback among the endogenous variables. The direct effects correspond to the elements in A of the structural model $Ay_t = \dots + \varepsilon_t$ in (1). We reverse the signs of the off-diagonal entries for ease of interpretation. The lower part of the table shows the *overall* effects. They take into account all contemporaneous feedback among the endogenous variables, are contained in A^{-1} , and refer to model (1) pre-multiplied by A^{-1} : $y_t = \dots + A^{-1}\varepsilon_t$. Bootstrap p -values are below point estimates.

Regarding shock transmission between sovereign risk and bank risk, both the direct and the overall effects are positive. Moreover, they are statistically and economically significant. A unit bank risk shock increases sovereign risk *directly* by 0.18 standard deviations. Reversely, the effect of sovereign risk shocks is even larger at 0.44. Mutual feedback between sovereign and bank risk (and other asset prices) amplifies both types of shocks substantially within the same day. The amplification can be quantified by comparing the direct and the overall effects. The feedback loop is strong: both types of shocks are amplified by about 30%. Together, the estimates suggest a significantly positive two-way causality between sovereign and bank risk that generates an important contemporaneous feedback loop.

Compared to Archarya, Drechsler and Schnabl (2014), our estimates, based on a different methodology and sample, imply a larger impact of sovereign risk on bank risk. The authors estimate that a 10 percent increase in sovereign CDS translates into a 0.9 percent increase in bank CDS. Our specification suggests that a 100 basis points increase in sovereign CDS leads to a 32 basis points increase in bank CDS.

All other types of shocks except term spread shocks affect both sovereign and bank risk significantly and the signs correspond to our priors. Reversely, shocks to sovereign and bank risk increase non-financial credit risk and reduce equity returns. The effects of both shocks are similar in magnitude. Both shocks also significantly affect the term spread. Intuitively, the effect of sovereign shocks is stronger.

As indicated above, if we have more than two volatility regimes and the additional regimes contain significant relative changes in volatility, the system is overidentified. Wald test as those in Section 3.3 indeed show that the system is identified when using any four out of the five regimes. The estimates based on these subsets should be statistically indistinguishable from the baseline results using all five regimes, if the assumption of constant impact effects across all five regimes holds (see Rigobon and

Sack, 2003). We can test whether this is the case. We first focus on the two main parameters of interest, the effects between sovereign and bank risk. Figure 2 shows the bootstrap distribution of the difference between the alternative estimates and the baseline estimate, for the five possible subsets. In nine of the ten cases the hypothesis of constant parameters is not rejected at the 5% level. The difference is only borderline significant at that level in one case, but not at the 1% level.

We next test the constancy of the full A -matrix. For this, we estimate the model on data corresponding to regimes 1-3 and 4-5 separately and compare the difference in estimates across subsamples. Figure 3 shows that in 18 out of 20 estimated parameters there is no indication of instability, and in particular not for the main parameters of interest. We thus conclude that the identifying assumption of no change in impact effects across volatility regimes is consistent with the data.

To see whether the results in Table 4 are sensitive to using alternative volatility regimes, we employ two purely statistical approaches to defining the regimes. The first approach follows Rigobon and Sack (2003) and Ehrmann, Fratzscher and Rigobon (2011). For each of the five equations we classify the residuals into a high volatility regime when the 8-day rolling standard deviation of a given residual exceeds a threshold of 1, and those of the others do not. Additionally, we specify a low (high) volatility regime where all residuals are below (above) the threshold, yielding seven regimes. Table 5 shows that the two-way effects between bank risk and sovereign risk increase, as well as the amplification effects. The second approach is based on the estimated variances of variable-specific GARCH(1,1) models.² The online appendix contains further details about both approaches and shows that the second also yields similar results.

Returning to the estimates from the narrative regime definition, we perform the following counterfactual exercise to understand what drives the positive correlation between sovereign and bank spreads shown in Figure 1. First, we compute historical decompositions of both spreads. They yield the contribution of each structural shock to the daily evolution of these variables. Then, we calculate the evolution of both variables assuming that only one type of shock materialised and compute the correlation between both series over rolling windows of 200 days. Finally, we compare the shock implied correlations to the actual correlation to assess the importance of each shock for the evolution of the latter.³

Figure 4 contains the results. The shock implied correlations can be grouped in two. The first group (upper panel) contains the correlations implied by shocks to bank risk, to non-financial risk, and to stock returns. Bank risk shocks explain the high

² We are thankful to an anonymous referee for suggesting the use of GARCH models.

³ We neglect the correlation implied by term spread shocks since they are insignificant in explaining bank risk and sovereign risk according to the overall effects (see Table 4).

frequency movements of the actual correlation. Non-financial risk shocks and stock market shocks account for its level.

But these shocks alone cannot explain the large shift in the actual correlation. This shift is driven by sovereign risk shocks and own shocks (see lower panel). Own shocks refer to sovereign risk being driven by sovereign shocks only and bank risk by bank shocks only. These shocks drive up the implied correlation by about 0.5-1.0. Two scenarios help to understand the low level of the correlation implied by own shocks before 2010. In the first, the implied correlation is low if mainly bank shocks occur. This was the case when the US subprime crisis spilled over to European interbank markets, while sovereigns stood sheltered apart. In the second scenario, the implied correlation is low if shocks to both sovereign risk and bank risk occur but if they tend to go in opposite directions. This happened after the collapse of Lehman Brothers when most countries implemented major bailout packages for their domestic banking sectors. These bailouts came as negative shocks to bank risk that tended to push up sovereign risk.

To understand the high level of the correlation implied by own shocks after 2010, we consider a third scenario. Here, sovereign risk and bank risk are positively correlated if shocks to both variables occur and if they move in the same direction. This scenario seems particularly relevant when the European debt crisis gathered pace. One illustrative example is the introduction of the European Financial Stability Facility on 10 May 2010. Finally, these scenarios are mirrored in the correlation implied by sovereign shocks, which increases as their importance grows over the sample period.

4.2 Overall effects in individual euro area member countries

As the member states were affected by and responded to the crises in different ways, we next assess whether there is heterogeneity in the relationship between sovereign risk and bank risk across countries. For this, we estimate the model using country-specific data for all five endogenous variables, leaving the exogenous variables unchanged. Table 6 shows the results, focusing on the overall effects between sovereign risk and bank risk.⁴ The first column repeats the results for the euro area for comparison. In Austria, Belgium, Italy, Spain, and France there are significantly positive effects in both directions. Throughout the effect of sovereign risk is larger than vice versa. For Austria and Belgium the impact of sovereign shocks is smaller than for the euro area, whereas it is of similar magnitude in France, Italy, and Spain.

⁴ By and large, the signs of the other coefficients in the A and A^{-1} matrices are as expected. For example, credit risks across all three credit sectors tend to be positively related, while the effects between the term spread and other asset prices vary across country. In periphery countries, the credit component mostly dominates, inducing a positive relation with credit markets and a negative with the equity market, whereas in many core countries the business cycle part prevails, implying opposite co-movements.

In the second group of countries, there are no or only unidirectional effects. With the exception of the Netherlands, the effect of sovereign risk on bank risk is always significant. For Greece and Portugal this finding is consistent with the notion that credit risk in these countries originated mainly in the public sector.

Two underlying features help understand some of the cross-country heterogeneity in the estimates. It seems that larger bailout packages for the banking sector imply tighter links between banks and sovereigns. Bank support was relatively large in Spain, Belgium, and France. Exceptions are the Netherlands, with large bank support but only a small uni-directional effect, and Italy, with small bank support and strong ties between sectors. For Ireland, Greece, and Portugal there seems to be a second explanation. All three sovereigns were bailed out. This seems to have broken the link between domestic banks and sovereigns as bank credit risks on the balance sheets of these sovereigns was forwarded to other sovereigns. In light of these findings, the parameter estimates for the bidirectional effects between sovereigns and banks from the previous subsection are potentially underestimated, given that they are based on the full sample and including CDS spreads of sovereigns that were bailed out.

4.3 Credit risk spillovers between euro area core and periphery

We now investigate whether there are credit risk spillovers across countries. We resort to the specification for the euro area level but split the sovereign and bank risk variable each into core and periphery.⁵ This yields a system of seven endogenous variables. Table 7 shows the overall effects. Most of the cross-effects in the sovereign-bank risk nexus are significant and positive. Within each country group, the effect of sovereign risk on bank risk is always stronger than vice versa.

Regarding cross-country spillovers, we find tight links both within each credit sector and across sectors. Spillovers of sovereign risk and bank risk are similar in magnitude, or even larger from the periphery to their core counterparts than the other way around. This finding surprises at first sight given the relative importance of the underlying economies and debt markets. It can be rationalized by the high credit exposure of core banks to periphery sovereigns. Across sectors, credit risk spillovers are stronger from sovereigns to banks.

To rank the strength of the effects of the shocks, we sum columnwise over the top four rows. The largest spillovers come from core sovereign risk (2.68), followed by periphery sovereign risk (2.44), periphery bank risk (2.06), and core bank risk (1.21). The large effects of core sovereign risk are consistent with their large exposure to periphery sovereign risk, due to sovereign bailouts, and their high exposure to banking risk, due to bank bailouts.

⁵ We classify as core Austria, Belgium, France, Germany, and Netherlands and as periphery Ireland, Italy, Portugal, and Spain.

To quantify the average economic significance of all types of structural shocks, we compute one day ahead forecast error variance decompositions.⁶ They yield the percentage contribution of each shock to the variance of the endogenous variables. Since the decompositions are regime-specific, we calculate a weighted average over all regimes for clarity, using the number of observations per regime as weights.

The results in Table 8 mirror the estimated overall effects. First, sovereign risk is more important in explaining bank risk than vice versa. It explains 27% and 19% of the variability in periphery and core bank risk, respectively. In contrast, bank risk shocks contribute only up to 9% to sovereign risk variability. Second, sovereign risk shocks are especially important in the periphery, where they explain 8% of the variance of bank risk. Third, periphery shocks contribute similarly to the variance of core variables as core shocks to periphery variable variance. Fourth, other than own shocks play a role for bank risk variability, whereas for sovereign risk variability their contributions are negligible. The last observation is in line with the finding in Table 7 of large effects of core sovereign risk. Due to multiple types of guarantees, shocks to core sovereign risk are important for the evolution of other variables.

We close the section by returning to the issue of the interpretation of the structural shocks. We address it in several ways. First, use the variance decompositions to see whether the shocks that are named after the equation they appear in explain a large fraction of the variance of that variable. This is motivated by the idea that one would expect sovereign risk shocks, say, to be an important driver of sovereign risk. The estimates show that this is the case for all shocks.

Second, we look at the contribution of the shocks to their own variables over time using historical compositions. Figure 5 shows that in particular sovereign risk shocks (grey line) help to explain the evolution of actual sovereign risk (black line) during the European debt crisis, both in the core and in the periphery. In the latter country group the increase in the cumulated shock series occurs around the bailouts of Greece, Portugal, Ireland, and Spain. Similarly, bank risk shocks contribute especially to bank risk in core countries during the period that one would expect a priori, namely, during and following the global financial crisis. Shocks to non-financial credit risk and the stock market are also important in explaining the dynamics of these variables during this period.

Third, following Herwartz and Lütkepohl (2014), we check whether the shocks show higher volatility during those times that we would expect, given our interpretation of shocks. Fourth, we interpret selected single shocks for illustration. The last two approaches are deferred to the online appendix. All in all, these analyses support our shock interpretations.

⁶ Given that there is essentially no persistence in the differenced data, the one-day ahead variance decompositions are close to the long-run forecast error variance decompositions.

5. Monetary Policy and Bank Bailout Announcements

In this section, we first analyse the effects of specific ECB measures on euro area asset prices, before we estimate the effects of bank bailout announcements on credit risks. Our identification strategy is based on the ‘event study’ approach (see Kuttner, 2001). Generally, the approach uses high frequency data on financial variables directly or indirectly related to a policy indicator. The idea is that before a policy announcement these variables incorporate the expected endogenous response of policy to economic conditions. Changes in the variables from before to after the announcement then reflect the unexpected component of the announcement and are interpreted as exogenous (see Gürkaynak, Sack and Swansson, 2005). This implies that our estimates do not allow quantifying the overall effect of these policies, which contains also the endogenous and expected response of policy to the state of the economy or the financial system, and thus provide a lower bound of the effectiveness.

The identifying assumption is that the monetary and bank bailout announcements which we consider are unlikely to have occurred in response to daily market conditions. To ensure that this is the case, we focus on the most important announcements contained in our sample. The specific events are listed below in the respective subsections. They were a reaction to generally worsening financial and economic conditions. Such decisions are typically preceded by lengthy internal discussion and preparation and not taken on a day-by-day basis.

To trace out the full dynamic implications of the announcements, we compute cumulative impulse responses of the endogenous variables over a horizon of three trading weeks from the corresponding lags of the event dummies and taking into account the autocorrelation structure of the model. Market participants may have needed time to understand the new information, in particular given that the measures we consider were unprecedented in the euro area at that time. As the window size grows, the effect of the announcements may be estimated less precisely if other information approaches the market. To mitigate this risk, we control for macroeconomic surprises (see Section 3.1).

5.1 Monetary policy, credit risks, and other asset prices

We first analyse the effects of monetary policy at the euro area level. Then, we look at individual countries, focusing on the response of sovereign and bank risk. We estimate the announcement effects through four daily indicator variables taking value one on one or more event days. In line with other papers studying the effects of ECB non-standard policy (Altavilla, Giannone and Lenza, 2016; Acharya et al., 2017; Krishnamurthy, Nagel and Vissing-Jorgensen, 2018), we use the following event days: LTRO with maturity up to 12 months (28 Mar., 4 Sep., and 15 Oct. 2008; 7 May 2009; 4 Aug. and 6 Oct. 2011), three-year LTRO (8 Dec. 2011), SMP (10 May 2010 and 8 Aug. 2011), OMT (26 Jul., 2 Aug., and 6 Sep. 2012). As some of the indicators

contain more than one event, they estimate the average effect of the considered announcements regarding a specific policy.⁷

For OMT, the first day corresponds to the “whatever it takes” speech of ECB President Mario Draghi in London. The other two events are the official announcement and subsequent clarification of the details of OMT during ECB press conferences. For the SMP, the first event day coincides with the creation of the European Financial Stability Fund (EFSF). To separate the effects of SMP announcements from that policy, we create an EFSF dummy (and corresponding lags) equal to one on the first SMP day and on three other important event days related to the EFSF, which are not overlapping with the second SMP event.⁸ The inclusion of this EFSF dummy into the regression allows disentangling the effects of SMP from EFSF announcements.

5.1.1 Monetary policy and euro area asset prices

We return to the five-variable system for the euro area en bloc. Figure 6 contains the cumulative responses of the endogenous variables (in rows) to ECB announcements (in columns), along with their 90 percent confidence bands.⁹ Many policy measures have strong effects on euro area asset prices, driving sovereign and bank risk in the same direction. This underscores the tight interdependence between both risks.

OMT announcements have the largest and most persistent effect on sovereign and bank risk, with CDS spreads declining significantly by about 40 and 20 basis points, respectively. Non-financial credit risk and the term spread also drop significantly, although the effects are more short-lived and turn borderline insignificant after one week. Equity prices increase by 5 percent. SMP announcements have strong effects on the term spread, reducing it persistently by 50 basis points. This is consistent with the main aim of this policy of mitigating tensions in sovereign debt markets. The effects on sovereign and non-financial CDS is also significantly negative during the first days, but die out afterwards. The impact on the other variables is insignificant.

In contrast, the announcement of three-year LTROs actually leads to a significant increase in credit risks. Sovereign and bank spreads rise by up to 60 basis point, before levelling out at 30 basis points. Non-financial spreads increase by 10 basis points. Mirroring these developments, stock prices drop by more than 5 percent. Interestingly, the effect on the term spread is first positive but then turns negative after

⁷ An alternative would be to estimate the effects of the events individually and then compute the average impact for a given policy type.

⁸ Following Szczerbowicz (2015), the dates are 10 May 2010 (creation of the EFSF), 12 Mar. 2011 (authorization of EFSF to intervene in primary sovereign debt markets), 26 Mar. 2012 (agreement to increase overall lending ceiling of EFSF from 500 to 700 billion euro), and 29 Jun. 2012 (allowance for EFSF to recapitalize Spanish banks directly and to purchase Italian sovereign debt).

⁹ We use a parametric bootstrap with 1000 replications which takes into account the heteroskedasticity.

a week. This finding is consistent with the argument that banks loaded up on domestic sovereign risk (see Archarya and Steffen, 2015). Finally, the effects of announcements regarding LTRO with 6-12 month maturity are insignificant, apart from a small and short negative impact on sovereign spreads.

The positive effects of three-year LTROs on spreads indicate that markets took the announcement of this measure as a disappointment. Indeed, financial press and market commentary around the event suggest that investors were expecting a reactivation of SMP. Consistent with this interpretation, the increase of sovereign spreads is stronger than that of bank spreads. Moreover, most market participants probably underestimated the final size of the LTROs, which exceeded one trillion euro.

Note that the finding of an increase in CDS spreads does not conflict with but is at the heart of our identification approach. It is precisely the fact that financial markets expect and price a certain policy response to economic conditions which controls for the endogenous part of policy. In other words, the daily dummy variable whose effects we trace out captures the unexpected component of monetary policy.

The implementation effects of LTROs, on the other hand, suggest that these programmes also reduced credit risks. Table 9 shows that the effect on bank risk is significantly negative for both maturities.¹⁰ The allotment of one billion euro lowers bank CDS spreads by 0.014 basis points. Given a programme size of roughly one trillion euro each, the cumulative implementation effect is –28 basis points. The implementation effect of three-year LTROs on sovereign risk is also significant. Netting the implementation with the announcement effects of LTROs gives a small positive impact on sovereign and bank risk on average.

5.1.2 Monetary policy effects on country CDS spreads

Next we explore the effect of monetary policy on sovereign and bank CDS spreads of individual euro area countries. To focus the discussion, we concentrate on the cumulative effects after five days for the announcement effects. Table 10 shows the results for sovereigns. OMT announcements lower spreads in all countries, and mostly significantly. The effects are strongest in Spain and Italy with –96 and –70 basis points, respectively. SMP announcements have smaller impacts and the point estimates are less significant. Again, the largest effects are found for Italy and Spain. For Portugal the point estimate is larger but barely misses significance. Together, the findings complement existing evidence from the cash bond market. Krishnamurthy et al. (2018) show that SMP and OMT reduce government bonds spreads mainly by lowering sovereign credit risk. Our estimates tend to be larger, probably because we take into account the persistence of the effects.

¹⁰ We include one lag based on the significance of the individual lag coefficients.

On the other hand, the three-year LTRO announcements have significantly positive effects on sovereign spreads of all countries (except of Greece). The largest increases are in countries that were covered by the SMP programme. The announcement effects of shorter LTROs are insignificant. The implementation effects of both liquidity schemes are mostly negative. The effects are in particular strong for Italy and Spain, in line with Acharya and Steffen (2015) showing that banks in these two countries engaged in carry trades, using the liquidity to buy domestic government bonds.

Table 11 shows that the announcement effects of ECB policy on bank spreads are qualitatively similar, but tend to be smaller. The implementation effects are similar as well, but stronger – as expected. In particular three-year LTROs reduce bank spreads significantly in most member states.

5.2 The effects of bank bailout announcements on euro area credit risks

We return to the analysis of bank bailouts. In Section 4 we analysed their effects implicitly, pooling them with other sources of exogenous variation in bank risk into ‘bank risk shocks’. These shocks contained both announcement and implementation effects of bank bailouts. We now focus on the announcement effects, modelling them explicitly. We neglect the implementation effects as the various bailout schemes differed substantially across countries, which prevents a meaningful comparison. It is also more difficult to deal with potential problems of endogeneity as the schemes have been implemented over months or even years.

Following CGFS and BIS (2010), we consider three types of bailout announcements by federal governments: (a) debt guarantees, (b) deposit guarantees, and (c) capital injections. The announcement dates across all countries are 29 and 30 Sep., 06, 07, 08, 09, 10, 13, and 14, Oct., and 15 Dec. 2008. From those, we compute the average number of announcements per bailout type and day at the euro area level. This yields three indicator variables bounded between 0 and 1. An analysis at the country level is infeasible because most member states announced more than one bailout type on the same day and the composition of jointly announced bailout types differs across countries. We use dummy variables instead of measures that aim at capturing the magnitude of the programmes as most announcements were general in nature and, at the time of the announcement, did not specify implied volumes.

We resort to the five-variable specification for the euro area and include the announcement variables as additional exogenous variable, one at a time. Including all bailout dummies jointly did not yield meaningful results due to problems of multicollinearity. Figure 7 shows the cumulative responses of sovereign risk (left column) and bank risk (right column) to the different types of announcements (in rows). The responses are negative throughout and mostly statistically significant. In particular the effects on bank risk are strong and persistent. Following the announcement of a capital injection or a deposit guarantee, bank spreads decline by up to 200 basis points, depending on the horizon considered. The negative impact of

debt guarantee announcements is also sizeable, exceeding 100 basis points after about two weeks. Interestingly, the announcements also reduce sovereign spreads in most cases, although the effects are smaller and less significant.

6. Conclusions

Using a system of simultaneous equations that exploits the heteroskedasticity of the underlying daily financial market data for identification, the analysis shows that sovereign risk shocks had an important impact on bank risk in the euro area. While also shocks to bank risk had an adverse impact on sovereign risk within the own country, sovereign risk overall appears to have been the more important driver of the sovereign-bank risk nexus. Additionally, we find evidence for large spillovers of sovereign risk and bank risk in the euro area between core and periphery countries. Overall, the paper confirms a significant two-way causality between bank risk and sovereign risk, and shows that the ensuing feedback loop amplifies initial credit risk shocks by 30%.

The core analysis of the paper focuses on whether and how bank bailout policies and monetary policy by the ECB have affected this sovereign-bank risk nexus. The announcements of bank bailout policies, such as capital injections into ailing banks and debt and deposit guarantees, exerted a significant negative impact on both sovereign risk and bank risk. As to monetary policy, our results show that announcements about Outright Monetary Transactions had the largest effects on credit risks, as they reduced CDS spreads of both sovereigns and banks significantly. The Securities Market Programme also led to improvements in credit markets, but the effects were more limited. In contrast, Longer-Term Refinancing Operations had overall limited effects on the sovereign-bank nexus, and in some instances even led to an increase in risk.

All in all, these findings suggest that both rescue policies of the banking system and monetary policy exerted a significant impact on sovereign and bank risk as well as on risks and prospects of the real economy. The empirical estimates also indicate that not all policy announcements lowered risks.

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Figures

Figure 1: Rolling correlation between sovereign credit risk and bank credit risk in the euro area

The figure shows the 200 days rolling correlation between (differenced) average euro area bank and sovereign credit default swap spreads (five year, senior) over the sample period.

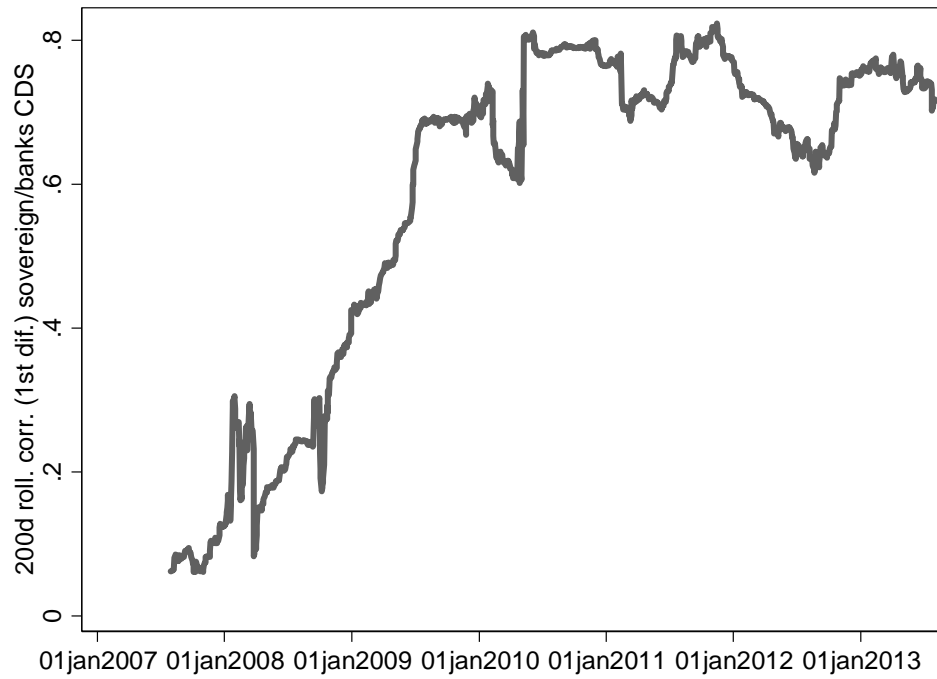


Figure 2: Test of parameter constancy of effects between bank risk and sovereign risk

The figure shows the distribution based on 1000 bootstrap replications of the difference between estimates when using volatility regimes 1-5 and when excluding one regime at a time. The upper panel contain the results for the effect of bank risk shocks on sovereign risk, parameter $A(1,2)$ of the matrix A of the structural model, and the lower panel those for the impact of sovereign risk on bank risk, parameter $A(2,1)$. Dashed and dotted lines mark out 5% and 1% probability mass, respectively.

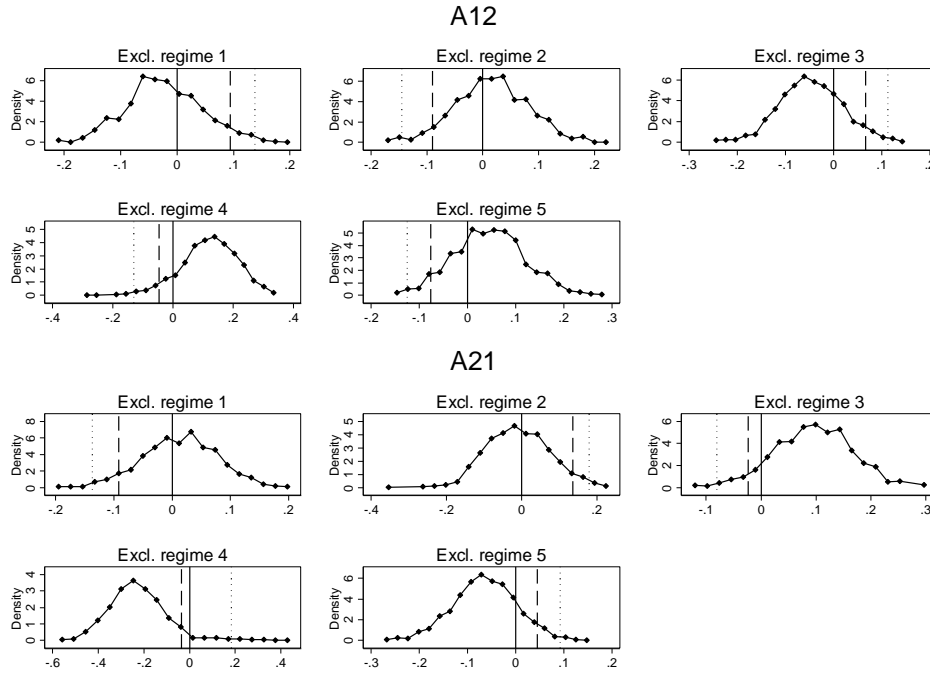


Figure 3: Test of overidentifying restrictions

The figure shows the distribution based on 1000 bootstrap replications of the difference between estimates of the elements of the A -matrix when estimating the model on data corresponding to regimes 1-3 and regimes 4-5. The rows correspond to the rows of the A -matrix excluding the main diagonal (consisting of ones). Dashed and dotted lines mark out 5% and 1% probability mass, respectively.

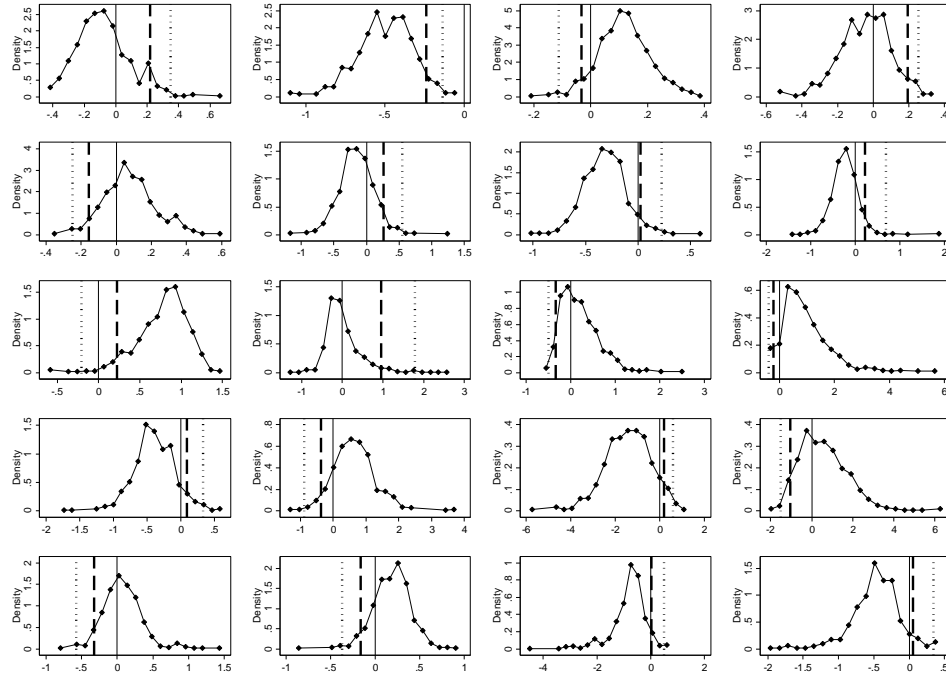


Figure 4: Actual and implied correlation between euro area sovereign risk and bank risk driven by different types of shocks

The figure shows the actual 200 days rolling correlation between euro area sovereign and bank CDS spreads and the shock implied correlations between these two variables. The latter are the counterfactual correlation between sovereign and bank CDS spreads in the euro area that would be observed if only one particular type of shock had materialised.

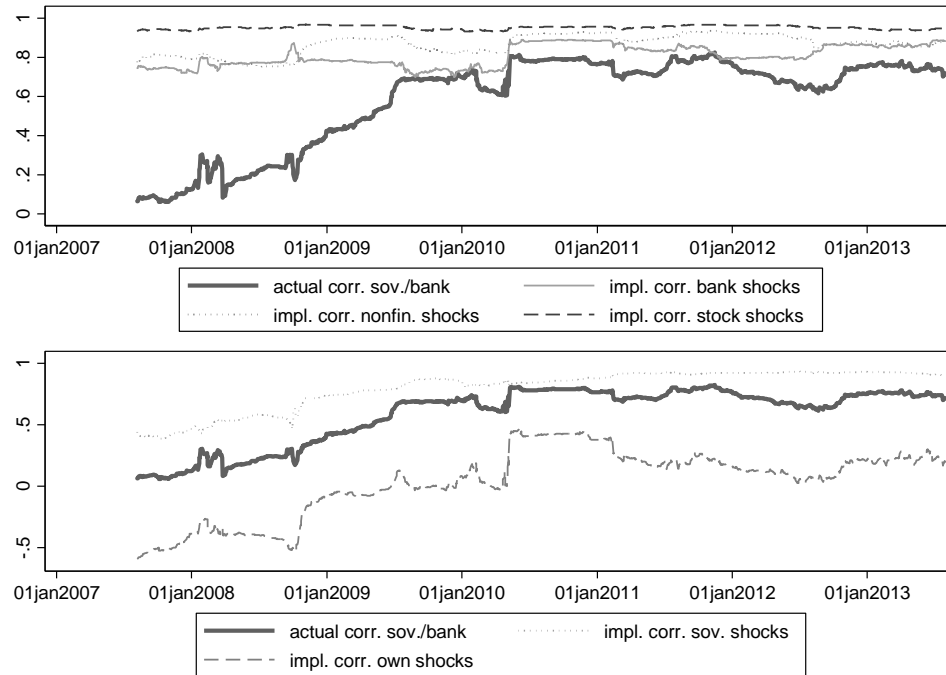


Figure 5: Cumulated effects of own shocks on endogenous variables

The figure shows the cumulative change of the endogenous variables of a seven-variable SVAR model for the euro area (black lines) and the cumulative effect of the structural shocks to the respective equations. The sample period is 24 October 2006 until 31 July 2013.

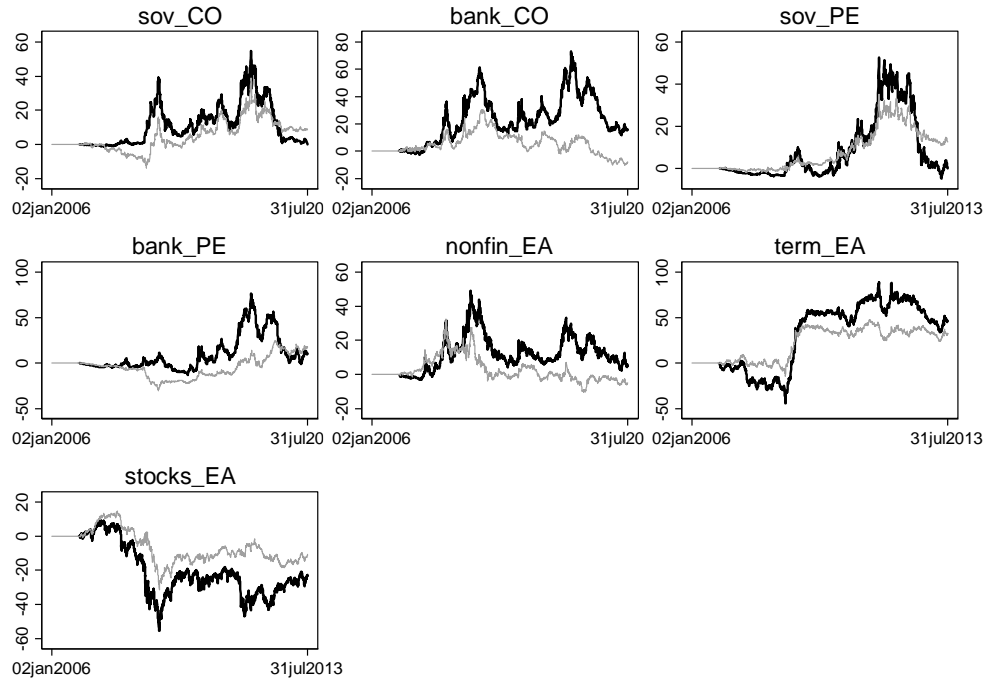


Figure 6: Impulse responses of euro area variables to monetary policy announcements

The figure shows the cumulative responses of the endogenous variables (in rows) to the different types of monetary policy announcements by the ECB (in columns) over a horizon of 15 trading days. 90% confidence bands based on 1000 parametric bootstrap replications. We consider the following announcement dates: LTROs with maturity up to 12 months (28 March, 4 September, and 15 October 2008; 7 May 2009; 4 August and 6 October 2011), three-year LTRO (8 December 2011), SMP (10 May 2010 and 8 August 2011), OMT (26 July, 2 August, and 6 September 2012).

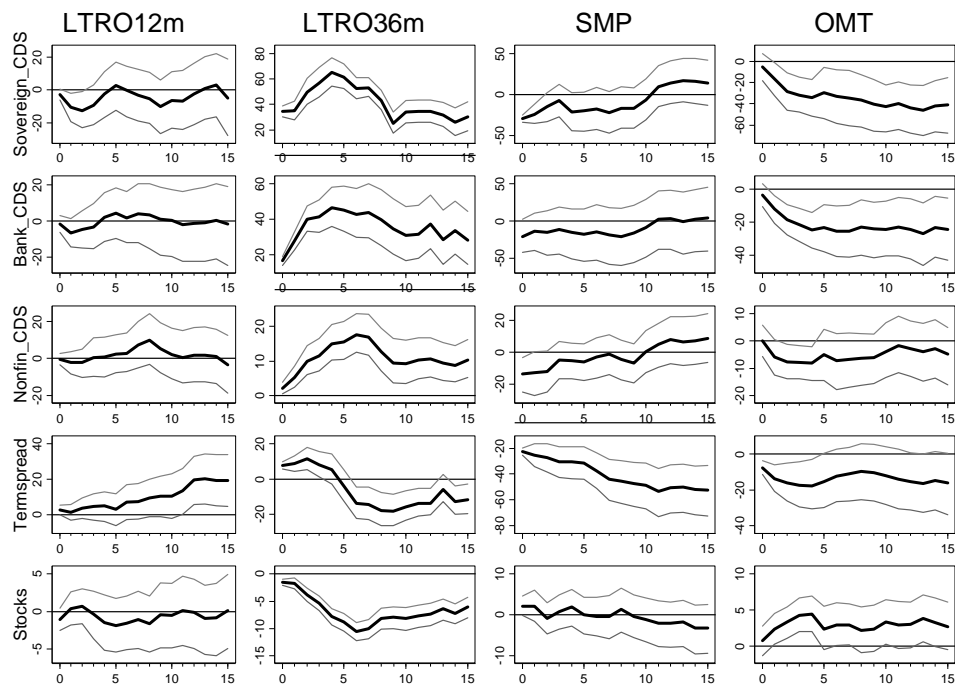
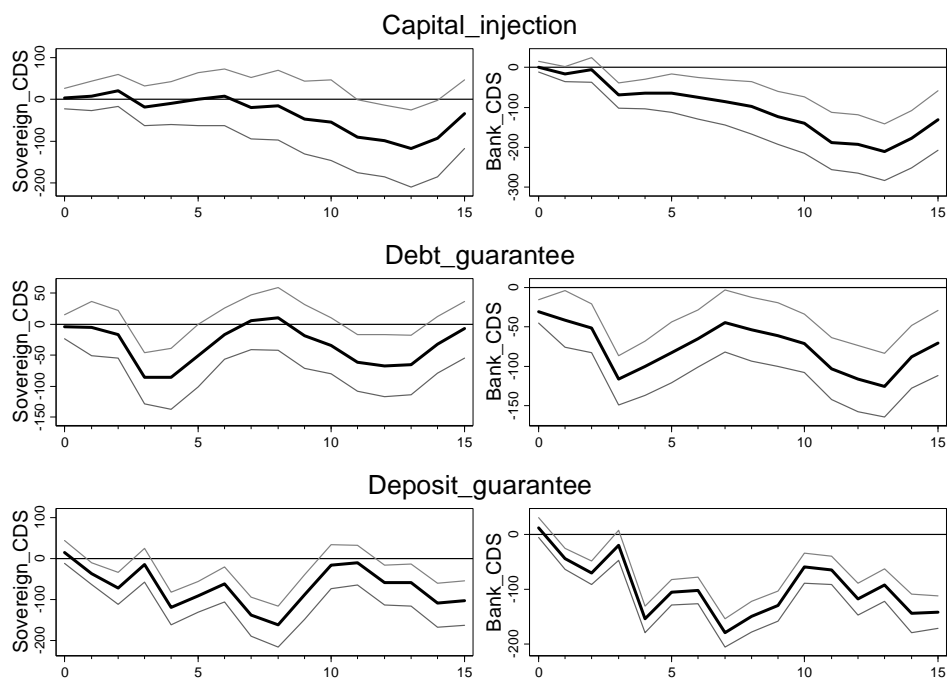


Figure 7: Response of euro area sovereign and bank credit risk to bank bailout announcements

The figure shows the cumulative responses of the euro area sovereign risk (left column) and bank credit risk (right column) to the different types of bank bailout announcements (in rows) by national government over a horizon of 15 trading days. 90% confidence bands based on 1000 parametric bootstrap replications. The announcement dates are 29 Sep. 2008, 30 Sep. 2008, 06 Oct. 2008, 07 Oct. 2008, 08 Oct. 2008, 09 Oct. 2008, 10 Oct. 2008, 13 Oct. 2008, 14 Oct. 2008, and 15 Dec. 2008.



Tables

Table 1: Tests for constancy of reduced-form covariance matrix

The table shows the results of likelihood ratio tests on the null hypothesis that all regimes have the same reduced-form covariance matrix and that each neighbouring regimes have the same reduced-form covariance matrix. Bootstrap p-values are in parentheses.

H_0	all $\Sigma_{u,i}$ equal	$\Sigma_{u,1} = \Sigma_{u,2}$	$\Sigma_{u,2} = \Sigma_{u,3}$	$\Sigma_{u,3} = \Sigma_{u,4}$	$\Sigma_{u,4} = \Sigma_{u,5}$
LR statistic	2746.1	494.9	426.7	634.3	393.2
p -value	0.0000	0.0000	0.0000	0.0000	0.0000

Table 2: Variance-ratios of structural shocks by volatility regime

The table shows the estimated ratios of the variances of the structural shocks for the euro area model by volatility regime. The shocks are numbered after the equations they appear in. The ordering of variables is sovereign CDS, bank CDS, non-financial CDS, term spread, Euro Stoxx 50. Bootstrap standard errors based on 1000 replications are below point estimates. The sample is 24 October 2006 until 31 July 2013.

Shock k	Volatility Regime				
	1	2	3	4	5
	<u>Variance-ratio sovereign credit risk shock/shock k</u>				
Bank risk	0.21 (0.03)	0.41 (0.04)	0.76 (0.08)	1.85 (0.18)	0.85 (0.14)
Nofin. risk	0.37 (0.06)	0.06 (0.01)	0.30 (0.03)	4.25 (0.43)	1.27 (0.17)
Term spread	0.11 (0.02)	0.11 (0.01)	0.28 (0.03)	1.51 (0.15)	0.39 (0.06)
Stock market	0.27 (0.04)	0.28 (0.04)	0.31 (0.03)	4.64 (0.42)	1.23 (0.17)
	<u>Variance-ratio bank credit risk shock/shock k</u>				
Nofin. risk	1.78 (0.33)	0.14 (0.02)	0.40 (0.04)	2.29 (0.22)	1.49 (0.25)
Term spread	0.55 (0.09)	0.27 (0.03)	0.38 (0.04)	0.82 (0.08)	0.46 (0.09)
Stock market	1.31 (0.20)	0.68 (0.08)	0.41 (0.04)	2.51 (0.22)	1.44 (0.22)
	<u>Variance-ratio non-financial credit risk shock/shock k</u>				
Term spread	0.31 (0.05)	1.90 (0.26)	0.93 (0.10)	0.36 (0.03)	0.31 (0.05)
Stock market	0.74 (0.11)	4.78 (0.61)	1.03 (0.10)	1.09 (0.10)	0.97 (0.13)
	<u>Variance-ratio term spread shock/shock k</u>				
Stock market	2.40 (0.36)	2.51 (0.34)	1.10 (0.12)	3.07 (0.28)	3.11 (0.44)

Table 3: Identification tests

The table shows the Wald statistics and associated p -values of linear Wald tests on the joint null hypothesis that the estimated variance ratios of two structural shocks are the same across volatility regimes, for each pair of structural shocks. The tests are based on 1000 bootstrap replications.

Shock to	Shock to Sovereign risk	Bank risk	Nonfin. risk	Term spread
	<u>Wald statistic</u>			
Bank risk	94.75	.	.	.
Nonfin. Risk	163.31	168.10	.	.
Term spread	95.37	48.74	80.26	.
Stock market	117.28	106.93	46.62	71.62
	<u>p-value</u>			
Bank risk	0.0000	.	.	.
Nonfin. Risk	0.0000	0.0000	.	.
Term spread	0.0000	0.0000	0.0000	.
Stock market	0.0000	0.0000	0.0000	0.0000

Table 4: Direct and overall effects among endogenous variables at the euro area level

The table shows the estimated direct effects and overall effects of a unit structural shock on the endogenous variables, based on a five-variable SVAR for the euro area. The direct effects are contained in the matrix A of the structural model $Ay_t = \tilde{c} + \tilde{A}y_{t-1} + \Lambda + \varepsilon_t$. The overall effects are contained in A^{-1} and refer to the structural model pre-multiplied by A^{-1} : $y_t = c + A_1y_{t-1} + \Lambda + A^{-1}\varepsilon_t$. The overall effects capture all contemporaneous propagation among variables. Impulse variables are in columns, response variables are in rows. The sample period is 24 October 2006 until 31 July 2013. The number of observations is 1761. ***, **, * denotes significance at the 1%, 5%, 10% level, respectively. P-values based on 1000 bootstrap replications are below point estimates.

	Impulse				
Response	Sovereign risk	Bank risk	Nonfin. risk	Term spread	Stock market
<u>Direct effects</u>					
Sovereign risk	1	0.18***	0.01	0.01	-0.06**
<i>p-value</i>	.	(0.00)	(0.33)	(0.29)	(0.02)
Bank risk	0.44***	1	0.24***	-0.02	-0.04
<i>p-value</i>	(0.00)	.	(0.00)	(0.38)	(0.14)
Nonfin. risk	0.05**	0.08**	1	0.14***	-0.39***
<i>p-value</i>	(0.02)	(0.01)	.	(0.00)	(0.00)
Term spread	0.39***	0.22**	-0.32***	1	0.09
<i>p-value</i>	(0.00)	(0.02)	(0.00)	.	(0.14)
Stock market	-0.08***	-0.16***	-0.08***	0.02	1
<i>p-value</i>	(0.00)	(0.00)	(0.00)	(0.35)	.
<u>Overall effects</u>					
Sovereign risk	1.12***	0.23***	0.07***	0.02	-0.10***
<i>p-value</i>	(0.00)	(0.00)	(0.00)	(0.24)	(0.00)
Bank risk	0.55***	1.16***	0.30***	0.02	-0.19***
<i>p-value</i>	(0.00)	(0.00)	(0.00)	(0.33)	(0.00)
Nonfin. risk	0.24***	0.22***	1.05***	0.14***	-0.42***
<i>p-value</i>	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Term spread	0.46***	0.25**	-0.26***	0.97***	0.14**
<i>p-value</i>	(0.00)	(0.01)	(0.00)	(0.00)	(0.04)
Stock market	-0.19***	-0.22***	-0.15***	0.01	1.08***
<i>p-value</i>	(0.00)	(0.00)	(0.00)	(0.46)	(0.00)

Table 5: Effects among endogenous variables at the euro area level using statistical approach

The table shows the estimated direct effects and overall effects of a unit structural shock on the endogenous variables, based on a five-variable SVAR for the euro area and using a statistical approach to define volatility regimes. The direct effects are in the matrix A of the structural model $Ay_t = \Lambda + \varepsilon_t$. The overall effects are contained in A^{-1} and refer to the structural model pre-multiplied by A^{-1} : $y_t = \Lambda + A^{-1}\varepsilon_t$. The overall effects capture all contemporaneous propagation among variables. Impulse variables are in columns, response variables are in rows. The sample period is 24 October 2006 until 31 July 2013. The number of observations is 1761. ***, **, * denotes significance at the 1%, 5%, 10% level, respectively. P-values based on 1000 bootstrap replications are below point estimates.

Response	Impulse				
	Sovereign risk	Bank risk	Nonfin. risk	Term spread	Stock market
<u>Direct effects</u>					
Sovereign risk	1	0.37***	-0.04*	0.07*	0.00
<i>p-value</i>	.	(0.00)	(0.09)	(0.05)	(0.46)
Bank risk	0.42***	1	0.19***	0.03	-0.08
<i>p-value</i>	(0.00)	.	(0.00)	(0.21)	(0.10)
Nonfin. risk	0.00	0.15**	1	0.19***	-0.35***
<i>p-value</i>	(0.47)	(0.04)	.	(0.00)	(0.00)
Term spread	0.40***	0.18	-0.21**	1	0.13
<i>p-value</i>	(0.00)	(0.12)	(0.01)	.	(0.21)
Stock market	-0.15**	-0.17**	-0.09**	-0.04	1
<i>p-value</i>	(0.01)	(0.01)	(0.03)	(0.36)	.
<u>Overall effects</u>					
Sovereign risk	1.25***	0.49***	0.02	0.11**	-0.04
<i>p-value</i>	(0.00)	(0.00)	(0.23)	(0.04)	(0.35)
Bank risk	0.62***	1.32***	0.21***	0.13**	-0.17*
<i>p-value</i>	(0.00)	(0.00)	(0.00)	(0.03)	(0.07)
Nonfin. risk	0.32***	0.38***	1.04***	0.25***	-0.37***
<i>p-value</i>	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Term spread	0.50***	0.31**	-0.19**	1.00***	0.17
<i>p-value</i>	(0.00)	(0.03)	(0.01)	(0.00)	(0.19)
Stock market	-0.35***	-0.35***	-0.13***	-0.10	1.06***
<i>p-value</i>	(0.00)	(0.00)	(0.00)	(0.11)	(0.00)

Table 6: Overall effects between sovereign risk and bank risk in euro area member states

The tables show the contemporaneous overall effects between sovereign credit risk shocks and bank credit risk shocks in the euro area and in member states, based on five-variable structural VARs for the euro area and the individual countries, identified through heteroskedasticity. The numbers correspond to the respective elements in the A^{-1} matrices. P-values are below point estimates based on 1000 bootstrap replications.

Country	EMU	AT	BE	ES	IT	FR
Bank risk on sovereign risk	0.23***	0.05***	0.07**	0.14**	0.38***	0.05*
<i>p-value</i>	(0.00)	(0.00)	(0.01)	(0.03)	(0.00)	(0.05)
Sovereign risk on bank risk	0.55***	0.15***	0.15***	0.51***	0.60***	0.48***
<i>p-value</i>	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Observations	1764	2477	2142	2470	1765	2074

Country	DE	GR	IE	NL	PT
Bank risk on sovereign risk	-0.02	0.01	0.03	0.03*	0.09
<i>p-value</i>	(0.15)	(0.39)	(0.33)	(0.08)	(0.19)
Sovereign risk on bank risk	0.25***	0.10**	0.08*	-0.02	0.21**
<i>p-value</i>	(0.00)	(0.02)	(0.05)	(0.30)	(0.04)
Observations	1571	1466	2454	1765	2003

Table 7: Overall effects among euro area core and periphery sovereign risk and bank risk

The table shows the estimated overall effects of structural shocks of one standard deviation on the endogenous variables. They are extracted from the A^{-1} -matrix of a seven-variable structural VAR, containing sovereign and bank credit risk in the euro area core and periphery and other asset prices. Impulse variables are in columns, response variables are in rows. The sample period is 24 October 2006 until 31 July 2013. The number of observations is 1761. ***, **, * denotes significance at the 1%, 5%, 10% level, respectively. P-values based on 1000 bootstrap replications are below point estimates.

<u>Response</u>	<u>Impulse</u>						
	Sovereign risk core	Bank risk core	Sovereign risk periphery	Bank risk periphery	Nonfin. risk EMU	Term spread EMU	Stock market EMU
Sovereign risk core	1.15*** (0.00)	0.00 (0.41)	0.43*** (0.00)	0.03 (0.19)	0.08*** (0.00)	-0.04 (0.10)	-0.10** (0.02)
Bank risk core	0.49*** (0.00)	1.05*** (0.00)	0.31*** (0.00)	0.55*** (0.00)	0.43*** (0.00)	-0.12*** (0.00)	-0.27*** (0.00)
Sov. risk periphery	0.48*** (0.00)	0.02 (0.13)	1.25*** (0.00)	0.32*** (0.00)	0.04*** (0.00)	-0.01 (0.31)	-0.07** (0.01)
Bank risk periphery	0.56*** (0.00)	0.14*** (0.00)	0.45*** (0.00)	1.16*** (0.00)	0.21*** (0.00)	-0.08** (0.01)	-0.13*** (0.00)
Nonfin. risk EMU	0.24*** (0.00)	-0.01 (0.44)	0.22*** (0.00)	0.24*** (0.00)	1.03*** (0.00)	0.09*** (0.00)	-0.41*** (0.00)
Term spread EMU	0.31*** (0.00)	0.14*** (0.00)	0.44*** (0.00)	0.40*** (0.00)	-0.23*** (0.00)	0.93*** (0.00)	0.12* (0.06)
Stocks EMU	-0.21*** (0.00)	-0.01 (0.36)	-0.14*** (0.00)	-0.22*** (0.00)	-0.14*** (0.00)	0.05 (0.11)	1.07*** (0.00)

Table 8: Variance decomposition for euro area core and periphery specification

The table shows the weighted average one day ahead forecast error variance decomposition, based on the seven-variable euro area core-periphery structural VAR. The weights are the number of observations per regime. The sample period is 24 October 2006 until 31 July 2013.

<i>Variable</i>	<i>Shock</i>						
	Sovereign risk core	Bank risk core	Sov. risk periphery	Bank risk periphery	Nonfin. risk EMU	Term spread	Stock market
Sovereign risk core	0.84	0.00	0.10	0.00	0.03	0.00	0.02
Bank risk core	0.14	0.49	0.05	0.16	0.13	0.00	0.02
Sov. risk periphery	0.21	0.00	0.67	0.09	0.01	0.00	0.01
Bank risk periphery	0.19	0.04	0.08	0.56	0.09	0.01	0.02
Nonfin. risk EMU	0.08	0.00	0.06	0.07	0.63	0.01	0.14
Term spread EMU	0.03	0.01	0.05	0.05	0.07	0.76	0.04
Stocks EMU	0.09	0.00	0.05	0.07	0.07	0.00	0.72

Table 9: Implementation effects of monetary policy on credit risks and other asset prices in the euro area

The table shows the implementation effects of LTROs on the endogenous variables, based on the five-variable structural VAR specification for the euro area level. In columns (1)-(4), the dependent variables are in basis points, in column (5) the dependent variable is in percentage points. Statistical significance is indicated with ***, **, * for the 1%, 5%, and 10% level, respectively, and refers to testing whether the sum of the contemporaneous effect and its first lag is equal to zero. P-values are below point estimates.

	(1) Sovereign risk	(2) Bank risk	(3) Non-fin. risk	(4) Term spread	(5) Stock market
LTRO12m	-0.025	-0.014*	-0.011**	0.007	0.002**
<i>p-value</i>	(0.120)	(0.055)	(0.032)	(0.149)	(0.036)
LTRO36m	-0.008**	-0.014***	0.003	0.004*	0.000
<i>p-value</i>	(0.026)	(0.000)	(0.125)	(0.063)	(0.403)
R squared	0.26	0.34	0.38	0.25	0.61
N	1761	1761	1761	1761	1761

Table 10: Announcement and implementation effects of monetary policy on sovereign credit risk of euro area member states

The table shows the effects of non-standard monetary policy measures by the ECB on *sovereign CDS spreads* of euro area countries, based on the sovereign risk equation of the corresponding country-specific SVAR. The dependent variable is in basis points. The upper part of the table shows the cumulative announcement effects of specific monetary policy measures. The lower part shows the cumulative implementation effects. Statistical significance is indicated with ***, **, * for the 1%, 5%, and 10% level, respectively, and refers to testing whether the sum of lags 0-5 for announcement and 0-1 for implementation effects is equal to zero. We consider the following announcement dates: LTROs with maturity up to 12 months (28 March, 4 September, and 15 October 2008; 7 May 2009; 4 August and 6 October 2011), three-year LTRO (8 December 2011), SMP (10 May 2010 and 8 August 2011), OMT (26 July, 2 August, and 6 September 2012).

	EMU	AT	BE	DE	FR	NL	ES	IE	GR	IT	PT
<u>Cumulative announcement effects (in basis points)</u>											
LTRO12m	0.6	2.9	10.0	-0.4	2.0	-1.8	-2.7	-0.1	19.7	-2.5	2.3
<i>p-value</i>	(0.942)	(0.624)	(0.122)	(0.914)	(0.679)	(0.736)	(0.832)	(0.997)	(0.845)	(0.838)	(0.927)
LTRO36m	58.7***	30.2***	46.0***	9.6***	55.5***	25.6***	70.6***	79.4***	516.2	125.8***	116.6***
<i>p-value</i>	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.385)	(0.000)	(0.000)
SMP	-27.5*	2.8	-0.1	0.0	11.2	-4.6	-49.0*	2.4	-98.0	-49.4**	-79.1
<i>p-value</i>	(0.054)	(0.823)	(0.995)	(0.992)	(0.313)	(0.494)	(0.058)	(0.932)	(0.624)	(0.022)	(0.108)
OMT	-36.6***	-7.7	-17.3	-3.1	-13.5***	-10.5***	-96.3***	-51.1**	.	-69.5***	-47.8
<i>p-value</i>	(0.005)	(0.173)	(0.138)	(0.376)	(0.001)	(0.001)	(0.000)	(0.022)	.	(0.004)	(0.251)
<u>Cumulative implementation effects (in basis points)</u>											
LTRO12m	-0.025	-0.028***	-0.016	-0.008*	-0.010	-0.018***	-0.024	-0.049**	-0.468	-0.034*	-0.051
<i>p-value</i>	(0.117)	(0.000)	(0.199)	(0.083)	(0.347)	(0.000)	(0.188)	(0.034)	(0.317)	(0.073)	(0.235)
LTRO36m	-0.013	-0.007***	-0.027**	-0.007***	-0.017**	0.009	-0.031**	0.010	-2.458***	-0.078**	0.009
<i>p-value</i>	(0.403)	(0.007)	(0.023)	(0.000)	(0.012)	(0.630)	(0.021)	(0.580)	(0.002)	(0.010)	(0.917)
R squared	0.26	0.16	0.22	0.15	0.20	0.31	0.28	0.24	0.30	0.29	0.20
N	1761	2481	2139	2467	2071	1762	1762	1463	1580	2451	2000

Table 11: Announcement and implementation effects of monetary policy on bank credit risk of euro area member states

The table shows the effects of non-standard monetary policy measures by the ECB on *bank CDS spreads* of euro area countries, based on the sovereign risk equation of the corresponding country-specific SVAR. The dependent variable is in basis points. The upper part of the table shows the cumulative announcement effects of specific monetary policy measures. The lower part shows the cumulative implementation effects. Statistical significance is indicated with ***, **, * for the 1%, 5%, and 10% level, respectively, and refers to testing whether the sum of lags 0-5 for announcement and 0-1 for implementation effects is equal to zero. We consider the following announcement dates: LTROs with maturity up to 12 months (28 March, 4 September, and 15 October 2008; 7 May 2009; 4 August and 6 October 2011), three-year LTRO (8 December 2011), SMP (10 May 2010 and 8 August 2011), OMT (26 July, 2 August, and 6 September 2012).

	EMU	AT	BE	DE	FR	NL	ES	IE	GR	IT	PT
Cumulative announcement effects (in basis points)											
LTRO12m	3.6	-10.1	31.5	-1.5	5.4	-7.2	21.5	-8.5	-14.9	1.1	-1.4
<i>p-value</i>	(0.606)	(0.467)	(0.118)	(0.808)	(0.604)	(0.308)	(0.117)	(0.742)	(0.324)	(0.896)	(0.915)
LTRO36m	35.9***	11.6**	62.8***	24.0***	47.8***	10.2***	34.9***	-101.3***	112.6**	40.6***	37.0**
<i>p-value</i>	(0.000)	(0.011)	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.037)	(0.000)	(0.047)
SMP	-20.4	2.3	-6.2	-9.8	-7.0	-0.1	-31.8	153.5*	-40.2	-18.0	-39.8
<i>p-value</i>	(0.204)	(0.915)	(0.635)	(0.212)	(0.776)	(0.997)	(0.105)	(0.082)	(0.180)	(0.239)	(0.291)
OMT	-27.5***	-36.0**	-13.1**	-2.9	-18.4***	-12.9***	-24.8*	32.2**	.	-53.1***	-57.5**
<i>p-value</i>	(0.000)	(0.039)	(0.041)	(0.313)	(0.000)	(0.004)	(0.079)	(0.023)	.	(0.000)	(0.020)
Cumulative implementation effects (in basis points)											
LTRO12m	-0.014**	-0.028**	-0.005	-0.008	-0.020**	-0.012**	-0.020*	-0.025	0.020	-0.013	-0.005
<i>p-value</i>	(0.049)	(0.042)	(0.745)	(0.177)	(0.033)	(0.030)	(0.077)	(0.233)	(0.683)	(0.254)	(0.536)
LTRO36m	-0.021***	-0.028***	-0.047***	-0.018***	-0.025***	-0.011*	-0.018**	-0.011	-0.174**	-0.031***	-0.034
<i>p-value</i>	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.075)	(0.011)	(0.586)	(0.010)	(0.002)	(0.174)
R squared	0.34	0.18	0.22	0.16	0.29	0.15	0.29	0.11	0.28	0.29	0.23
N	1761	2481	2139	2467	2071	1762	1762	1463	1581	2451	2000