Does BRRD mitigate the bank-sovereign risk nexus?*

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Abstract

We investigate the effectiveness of the Bank Recovery and Resolution Directive (BRRD) in mitigating the bank-sovereign nexus in Italy, Spain, Germany and France. Using CDS spreads to capture bank and sovereign credit risk in a DCC-MIDAS model, we document that the dynamic correlation between bank and sovereign credit risk has decreased significantly over the period in which the BRRD has been implemented. Bank-level results indicate that the decline in the interconnectedness between bank and sovereign risk is not driven by the banks' capital adequacy, size or holdings of domestic sovereign securities. If the BRRD bail-in framework is credible, exogenous bank risk shocks should not be transmitted to sovereign risk. In order to investigate this causality, we adopt an SVAR approach in which we use bank earnings announcements to identify exogenous bank risk shocks. Our results show that sovereign risk is no longer impacted by bank shocks in the BRRD regime in France and Germany, but the bank-sovereign nexus remains alive in Italy and Spain. These results suggest that the BRRD is least effective in countries where it is needed most.

JEL classification: C58; G28; G32.

Keywords: BRRD; Sovereign-bank nexus; CDS spread; Dynamic correlation; SVAR.

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1 Introduction

Is the BRRD bail-in regime credible? This is an important policy question since the Bank Recovery and Resolution Directive (BRRD) is a key component of the European resolution framework which is designed to organize the orderly unwinding of failing banks in such a way that a bail-in can be organized without involvement of governments. European leaders created the Banking Union in 2012 in an explicit attempt to break the bank-sovereign feedback loop that was threatening to break up the euro area. The Banking Union consists of three pillars: common bank supervision by the Single Supervisory Mechanism (SSM) of the ECB, a common resolution framework under the Single Resolution Mechanism (SRM) and common deposit insurance (EDIS). The first two pillars are in operation, but the deposit insurance is taking more time to complete. The SRM was introduced together with the BRRD, which provides the tools to be applied in bail-ins. The SRM applies to all banks under the supervision of the ECB and is administered by the Single Resolution Board (SRB), which treats systemic banks itself, while it can decide that non-systemic banks can be resolved under national law. As Farhi & Tirole (2018) have shown, centralized supervision at the euro area level can soften the bank-sovereign nexus. While there is broad agreement that bank supervision by the ECB works and that the harmonized banking supervision even had a positive valuation effect on banks (Loipersberger, 2018), there are persistent questions about the credibility of the resolution framework. The SRB has treated several cases according to the book since it became operational from 2016 onwards. The resolution of Banco Popular in 2017 was described by the financial press as a model to deal with failing banks (Financial Times, 8 June 2017). Nevertheless, there have also been instances in which sovereigns intervened to rescue stressed banks (e.g. Banca MPS). As a result, doubts remain about the willingness of countries to fully apply the bail-in rules. Moreover, although resolution may be a workable solution for individual bank distress cases, it remains unclear whether a generalized bail-in would work in a banking crisis affecting many banks simultaneously (Beck et al., 2020).

In terms of credibility of the new bail-in regime, several arguments can be developed. The fact that there is a unified European resolution framework and that it is administered by the SRB should give power to the new regimes and in any case makes it more difficult for national authorities to intervene. Actions speak louder than words (Schäfer et al., 2016), implying that effective bail-in cases administered by the SRB should reinforce the trust of financial markets in the application of bail-in. That is why various empirical papers, which we review later, conduct event studies around effective bail-in events. Doubt in the new bail-in regime can have various sources. Many banks still hold substantial amounts of their domestic sovereign debt, which could limit the weakening of the bank-sovereign nexus. Directives have to be transposed into national law, yet there was significant delay in the implementation of the BRRD across countries, which might result in legal uncertainties. The BRRD and the EU rules on state aid provide some degree of flexibility in the use of the bail-in tool, which might open the door for political influence. Also, under the BRRD, the SRB can discretionally decide to exclude, fully or partially, certain instruments from the bail-in, based on their maturities or holders. The possibility that bail-in cases would proceed heterogeneously across jurisdictions represents a potential source of uncertainty for market participants (Huertas, 2016).

In this paper we directly address the credibility issue by investigating the transmission of bank risk to sovereign risk. This is the channel of interest, since the hypothesis that the bail-in framework is credible implies that an increase in bank risk should not be transmitted to higher sovereign risk. A litmus test of the credibility of the resolution framework can be designed as follows. The first step is to analyze whether the conditional correlation between the risk of the bank and that of the sovereign decreases. If the bank-sovereign correlation decreases in the period in which the bail-in framework is operational, this would be consistent with the hypothesis that the doom loop has diminished in the BRRD era. If the bank-sovereign correlation does not decrease or even increases, this would suggest that markets still believe that bank bailouts are possible and that the doom loop is alive. The next step is to analyze the determinants of the individual bank-sovereign correlation, and thus to which degree it is driven by BRRD implementation events. Finally, we identify shocks to the risk of a bank that are unrelated to the risk of the sovereign in order to single out the channel of interest. More specifically, we identify events characterized by an exogenous increase in the riskiness of banks and we examine whether this increased bank risk is transmitted differently to sovereign risk after BRRD was implemented. The bank-specific events we consider are the quarterly earnings announcements by the banks, which are disclosed according to a predefined time schedule and thus exogenous to bank or sovereign stress. When the announcement is accompanied by an increase of the bank's CDS spread, we assume that the announcement is considered as bad news and hence increases the risk profile of the bank. Similarly, when the announcement is accompanied by a decrease of banks' CDS spread, the announcement is assumed to be positive news. We argue that this setup provides a genuine test of the credibility of the BRRD bail-in framework because we consider bank-specific events in which only bank risk is revealed without a contemporaneous effect on sovereign risk. We apply this analysis to the banks and their sovereign in two core euro area countries (Germany and France) and two periphery countries (Spain and Italy). Since the exposure of banks to their home sovereign is typically higher in the periphery countries, the doom loop is more pronounced there. If the BRRD framework is credible, it should thus benefit the more vulnerable countries most.

In the literature we find extensive research on the existence of the bank-sovereign feedback loop and the two-way risk spillovers between banks and their sovereign (De Bruyckere et al., 2013; Acharya et al., 2014; Fratzscher & Rieth, 2019). To isolate the bank-to-sovereign transmission or bailout channel, Böhm & Eichler (2020) use an instrumenting approach and report an economically meaningful and highly significant impact of bank sector distress on to sovereign distress. The transmission from the sovereign to banks, or sovereign-bond channel, is related to the fact that banks typically hold substantial amounts of domestic government bonds on their balance sheet because they are not subject to additional capital charges and because they can be used as collateral against central bank liquidity and for interbank operations (Allegret et al., 2017). In the European sovereign debt crisis, the interconnections between banks and sovereigns caused severe stress in the banking system and a new round of bailouts. Some even called it the diabolic bank-sovereign loop (Brunnermeier et al., 2016).

Since the introduction of the BRRD, research on the effect of the European resolution framework on the bank-sovereign nexus is receiving increasing attention. In general, theory is ambiguous on the effect of a bank resolution framework on the stability of banks. On the positive side, reducing the likelihood of bailouts and thus taxpayer support, allowing early intervention and providing tools for the orderly resolution of failing banks reduces moral hazard risk (Repullo, 2005; Farhi & Tirole, 2012). More specifically, the potential of a bail-in and ex ante knowledge on how losses will be distributed in case of bank failure may increase market discipline. They can also reduce incentives for banks to build up leverage in their balance sheets (Geanakoplos, 2010; Adrian & Shin, 2014). On the other hand, a rule-based resolution system can result in bank runs and contagion which would render the banking system less stable because of the direct interlinkages between banks and the possibility of a sudden reassessment of bank risk (Acharya & Yorulmazer, 2008; Eisert & Eufinger, 2019). According to this view, bailouts of failing banks can protect other banks from contagion and thus provide incentives to reduce risk-taking (Cordella & Yeyati, 2003; Dell'Ariccia & Ratnovski, 2019).

In order to assess the competing claims on the usefulness of resolutions regimes on the stability of the banking system, Beck et al. (2020) undertake an empirical analysis of bank resolution frameworks during systemic banking distress. Using a database on bank resolution regimes in 22 member countries of the Financial Stability Board, they assess the ability of bank resolution frameworks to deal with systemic banking fragility. They show that systemic risk, measured by CoVaR, increases more for banks in countries with more comprehensive bank resolution frameworks after negative system-wide shocks. Their results suggest that more comprehensive bank resolution may exacerbate the effect of system-wide shocks and should not be solely relied on in cases of systemic distress. However, they also find that bank systemic risk decreases more in countries with more comprehensive resolution frameworks after positive system-wide shocks such as Mario Draghi's 'whatever it takes' speech.

Turning to the case of the European resolution framework, Fiordelisi et al. (2020a) empirically examine the effect of changes in sovereign CDS spreads on bank CDS spreads in three periods surrounding the introduction of the BRRD. They report that the effect of sovereign CDS changes on banks' CDS diminishes in subperiods after the introduction of the BRRD, suggesting that the new bail-in regime decreased the interconnections between sovereigns and banks. However, while the authors argue that other regulatory reforms in the same period (e.g. Basel III) probably do not produce confounding effects, it might still be the case that the monetary policy of the ECB contributed significantly to a decrease of both sovereign risk and bank risk, as shown by Soenen & Vander Vennet (2020). Studies by Schäfer et al. (2016); Giuliana et al. (2019); Fiordelisi et al. (2020b) focus on investors' reactions to bail-in announcements and usually report a decrease in investors' bailout expectations, which suggests that bail-in has become more credible. Schäfer et al. (2016) estimate the stock price and CDS reaction to actual bail-ins (e.g. the Cyprus banks and the Portuguese Banco Espirito Santo) and to announcements related to the introduction of the BRRD. The authors find an increase in CDS spreads suggesting that the new regime reduces bail-out expectations. Comparable results for stock prices have been reported by Fiordelisi et al. (2020b) examining a broader set of regulatory announcements. Giuliana et al. (2019) considers the difference in yields between banks' bail-inable bonds and non-bail-in bonds. The main finding is that the spread between both types of bonds increases after events signaling an increased commitment by authorities to bail-in (again based on effective bail-in cases and regulatory announcements). While event studies provide evidence of the immediate reaction of investors to specific bail-in cases, the question is how long-lasting the effects are. Pancotto et al. (2019) conclude that the markets do not judge the BRRD as credible. They use a difference-in-differences approach with banks as the treated group and non-financial corporations as the control group. Instead of the expected widening of the gap between bank and sovereign CDS spreads in the BRRD period, the gap narrows, implying that bail-in is not credible. Interestingly, they report that the strongest credibility of the new regime is revealed in Italy. Neuberg et al. (2016) exploit a 2014 change in the definition of credit defaults swaps for European banks to show that the market price of protection against losses from government interventions exhibited a downward trend from 2014 to 2016, but this trend reversed, indicating a reversal of credibility of the bail-in instrument. These papers rely on the timing of the BRRD to conduct their analyses. Yet, determining when the markets consider the BRRD to be effective is far from easy. In fact, the European Commission revealed its plans for a new EU framework for crisis management in the banking sector in October 2010, it proposed rules for bank recovery and resolution in June 2012, and the European Parliament adopted the BRRD on 15 April 2014. Finally, the BRRD entered into legal force in all EU member states on 1 January 2015, while the Single Resolution Mechanism became fully operational on 1 January 2016. Hence, determining exact dates for event studies is difficult, also because such regulation is often anticipated by financial markets.

We contribute to this literature by investigating the longer-term credibility of the BRRD regime by analyzing the effect on the time-varying correlation between banks and their sovereign. Our analysis reveals that over the period from 2011 to 2020 the dynamic correlation between bank and sovereign risk has decreased dramatically for both the selected peripheral and core countries. Moreover, when we focus on cases in which the risk shock unambiguously originates in the banking sector, we demonstrate that in the core countries Germany and France, and to a lesser extent in Italy, this decrease in correlation is accompanied by a decrease of the bailout channel. The weakening of the spillover from bank to sovereign in the BRRD era, supports the hypothesis of a credible bail-in regime in those countries.

The paper unfolds in the following way. In section 2 and 3 we provide details on the data and methodology. In section 4 we discuss the results of our estimations. Section 5 concludes the paper and formulates some policy considerations.

2 Data and sample selection

We construct a dataset containing daily CDS spreads of banks and sovereigns retrieved from IHS Markit and daily market variables obtained from Refinitiv. We include all banks that have a CDS spread quotation at least 25% of the time over the sample period.

The application of the selection criterium results in a sample of 7 banks for Italy, 6 banks for Spain, 6 banks for France and 5 banks for Germany during the period of 2011-2020. These banks represent a large share of the banking sector in each country respectively. The sample period covers the pre- and post-crisis era and includes both the great financial crisis, the sovereign debt crisis and the onset of the corona crisis.

2.1 Bank and sovereign default risk indicator

We capture bank and sovereign default risk by their CDS spreads on 5-year senior bonds because they are a market-based, unbiased measure of default risk (Altavilla et al., 2018).

CDS spreads have three main advantages compared to bank- and sovereign bond spreads. First, CDS spreads provide timelier market-based pricing (Blanco et al., 2005). Second, using CDS spreads avoids the difficulty in dealing with time to maturity as in the case of using interest rate spreads (of which the zero coupon bonds would be preferred). Third, bond spreads include inflation expectations and demand/supply for lending conditions as well as default risk. Since we explicitly want to capture default risk, we focus on CDS spreads (De Bruyckere et al., 2013).

The top panel of Figure 1 displays the evolution of the CDS spreads for the Italian banks in our sample. We can discern four periods of heightened CDS spreads. First, the global financial crisis and the sovereign debt crisis led to significant increases in perceived bank and sovereign default risk. Second, during 2016-2017 doubts concerning the viability of bank business models translated to elevated CDS spreads. Third, in 2018-2019 there is a new surge in the CDS spreads of Italian banks, in this instance coinciding with the Italian general re-election and subsequent negotiations. Finally, In the first half of 2020 we notice elevated levels due to the onset of the corona pandemic. The bottom panel shows the equally-weighted average CDS spread of the Italian banks and the sovereign CDS spread. Noteworthy is that the sovereign CDS spread is typically lower than the bank CDS spreads. Market perceived Italian sovereign credit risk was particularly elevated during 2011-2012, a period characterized by the sovereign debt crisis in the euro area.

In Figure 2 we give a similar overview for the CDS spreads of Spanish banks and sovereign. Similar to the Italian banks, financial institutions experienced a first surge in market perceived credit risk with the onset of the great financial crisis and a second surge during the sovereign debt crisis. In contrast to the credit risk of Italian banks, Spanish CDS spreads decreased after the sovereign debt crisis and only increased mildly during 2016-2017. The onset of the corona pandemic gave rise to a modest surge, but stagnates at a much lower level than during the sovereign debt crisis. Similar to Italian sovereign credit risk, the Spanish CDS spreads reached a peak during the sovereign debt crisis, amid concerns on the debt burden of peripheral countries of the euro area to which Spain belongs.

Comparing the perceived credit risk of banks and sovereigns in peripheral countries to those in the core euro area, we notice clear differences. Two countries that belong to the latter are France and Germany. Figure 3 and 4 display the evolution of the CDS spreads of their banks and sovereign, respectively. A noticeable difference in the evolution of credit risk in both countries is that the levels of the CDS spreads are significantly lower than their peripheral counterparts, both for the banking sector and the sovereign. The CDS spread of the German sovereign in particular is only a fraction of the credit risk levels for Italy, Spain and even France.

In the remainder of the paper, we evaluate on the one hand the correlation between the CDS spreads of the bottom panel, i.e. the equally-weighted average bank CDS spread and the sovereign CDS spread. This allows us to estimate the correlation of the entire banking sector and sovereign credit risk for each country individually. On the other hand we estimate the correlation for each bank and respective sovereign separately, which we include in a panel setup to identify potential drivers of the interconnectedness.

2.2 Market variables

Credit risk of banks and sovereigns, and consequently their correlation may be affected by prevailing financial market conditions. Therefore we include a series of market variables when we estimate the SVAR on top of the CDS spreads, enriching the system of variables that endogenously affect each other. For each country we include a stock market index, captured by the MSCI Italy, Spain, France and Germany. We also add the volatility of these indices, estimated with a GARCH(1,1) model. Finally, we include the stock index of the broader European market, which we capture by including the STOXX Europe 600 index.

The top panel of Figure 5 shows the evolution of each country's MSCI index and the STOXX Europe 600. Three observations stand out in this figure. First, the global financial crisis impacted the stock market similarly in each country, but the consequent recovery is observably lower in Italy and also remains structurally lower for the remainder of the sample period. Second, the sovereign debt crisis is less visible in the stock market indices when compared to the impact it had on the CDS spreads. Finally, we distinguish the onset of the corona pandemic in the early stages of 2020, followed by a quick recovery in Germany and to a lesser extent in France. It is clear that the growth of the market indices is higher for the core euro area countries (Germany and France) in comparison with the peripheral countries (Spain and Italy). The bottom panel displays the evolution of the volatility of each country-specific stock index. The volatility of each country's stock index evolves quite similarly, with observable episodes of increased volatility during the global financial crisis, the sovereign debt crisis and the corona pandemic.

The descriptive statistics for the CDS spreads of banks and sovereigns, and market variables are reported in Table 1.

3 Methodology

Since we hypothesize that the bank-sovereign risk nexus evolves dynamically and is affected by the implementation of the BRRD, our empirical analysis proceeds in three steps. First, we describe the dynamic correlation of bank and sovereign credit risk with a short- and long-run component specification, using the DCC-MIDAS approach, on both a country level and a bank level. Second, we analyse whether or not the implementation of BRRD regulation was effective in reducing this correlation. We adopt a bank level panel approach, in which we estimate the effect of the BRRD directly on the bank-sovereign correlation. Additionally, we indirectly estimate the impact of BRRD, by examining the reaction of bank- and sovereign credit risk to an exogenous bank shock in an SVAR both in a pre-BRRD and BRRD regime, using the market reaction of bank credit risk on banks' earnings announcement days to identify the bank shock. This allows us to identify whether a shock in the banking sector transmits to sovereign credit risk. Both the estimation of the dynamic correlation and the SVAR are conducted on a country level, whereas the panel estimation is analyzed on a bank level.

3.1 Correlation of bank and sovereign credit risk

Since we hypothesize that bank and sovereign default risk are interconnected due to e.g. banks' home bias and implicit government guarantees, we assume a positive level of correlation between bank and sovereign credit risk. This positive relation could be susceptible to short-term deviations and long-term structural shocks. Therefore we hypothesize that the correlation of bank and sovereign default risk evolves dynamically and adopt a DCC-MIDAS model to be able to estimate the dynamic evolution of the correlation on the one hand, and to disentangle the short- and long-term components of the correlation on the other hand.

DCC-MIDAS is a natural extension of the DCC-GARCH model of Engle (2002), in which the long-run correlation is assumed to be constant. In the DCC-MIDAS the shocks to the correlation are mean-reverting around a long-run component. The DCC-MIDAS is thus a combination of the Engle (2002) DCC model and the Engle et al. (2006) and Engle & Rangel (2008) GARCH with a short- and long-run specification. In the latter case, two components of volatility are disentangled, one regarding short-term fluctuations and the other concerning a secular component, which is driven by realized volatilities computed over a longer time period, e.g. monthly, quarterly or annually (Engle et al., 2006). Using a MIDAS weighting scheme Ghysels et al. (2005), the realized volatilities are converted to a slowly moving secular component (Colacito et al., 2011). This specification can also be applied to correlations where the daily dynamics, which follow a DCC scheme, evolve around a long-run secular component, estimated using realized correlations through a MIDAS weighting scheme.

The estimation of a DCC-MIDAS model proceeds in two steps. In a first step the univariate conditional volatility is estimated using the GARCH-MIDAS model for both the bank and sovereign CDS spreads, yielding the univariate standardized residuals. The second step consists of the estimation of the parameters of the DCC-MIDAS model, using the standardized residuals from step one.

3.1.1 Step 1: GARCH-MIDAS

In the first step, we estimate the univariate conditional variance of both the bank and sovereign CDS spreads, allowing a short-run and a long-run secular component. The latter is estimated using exclusively financial market variables. The short-run is estimated by a GARCH(1,1) specification on the squared daily returns, whereas the long-term component is based on realized volatilities computed over a monthly basis, similar to Engle et al. (2006).

The 2×1 vector $\boldsymbol{r_t}$ contains the log return of bank and sovereign CDS spreads at day t during time period τ. In this paper, the time period chosen for τ is a month. However, choosing a different time period such as a quarter or bi-annual period has no major effect on the results. We thus assume that for each CDS spread $i = \{\textit{bank, sovereign}\}$ the univariate volatilities follow a GARCH-MIDAS(1,1) process:

$$r_{i,t} = \mu_i + \varepsilon_{i,t} \tag{1}$$

$$\varepsilon_{i,t} = \sqrt{m_{i,\tau}\sigma_{i,t}^2}\nu_t \qquad \text{with } \nu_t \sim N(0,1) \tag{2}$$

$$\sigma_{i,t}^2 = (1 - \alpha_i - \beta_i) + \alpha_i \frac{\varepsilon_{t-1}^2}{m_{i,\tau}} + \beta_i \sigma_{t-1}^2 \tag{3}$$

where μ_t shows the demeaned return, obtained from the conditional mean equation and ν_t represents the standardized residuals. The volatility consists of two components, a short-run component ($\sigma_{i,t}$) which follows a GARCH(1,1) process and varies daily and a MIDAS component ($m_{i,\tau}$) which is constant during each time period τ. Following Colacito et al. (2011), the latter component is a weighted sum of K lags of realized variances (RV) over a longer horizon. The realized variance is estimated by summing over the squared daily returns of time period τ:

$$m_{i,\tau} = \bar{m}_i + \theta_i \sum_{j=1}^{K} \phi_l(\omega_i) RV_{i,\tau-j} \tag{4}$$

$$RV_{i,\tau} = \sum_{i=1}^{N_\tau} r_{i,t,\tau}^2 \tag{5}$$

where N_τ contains the number of days in time-period τ. The weights given to each realized variance are Beta weights defined as:

$$\phi_l(\omega_i) = \frac{\left(1 - \frac{l}{K}\right)^{\omega_i - 1}}{\sum_{j=1}^{K}\left(1 - \frac{j}{K}\right)^{\omega_i - 1}} \tag{6}$$

where we choose K, the number of months over which to lag in the estimation of the monthly realized variances. The model is estimated using a quasi-maximum likelihood estimator where the parameters $(\mu_i, \alpha_i, \beta_i, \bar{m}_i, \theta_i, \omega_i)$ are optimized in such a way that the standardized residuals are most likely to represent a standard normal distribution.

3.1.2 Step 2: DCC-MIDAS

Similar to the GARCH-MIDAS specification to disentangle volatilities into a short- and long-run component, we adopt the same intuition when estimating the dynamic correlation between two time series, using the standardized residuals from the GARCH-MIDAS process $\varepsilon_{i,t}$. The 2×1 vector $\boldsymbol{r_t}$ contains the log return of bank and sovereign CDS spreads at day t. From Engle (2002), we assume the returns follow a stochastic i.i.d. process:

$$\boldsymbol{r_t} \sim N(\mu, H_t) \tag{7}$$

$$H_t = D_t R_t D_t \qquad \text{with } D_t = \text{diag}\{\sqrt{\sigma_t}\} \tag{8}$$

where μ is the vector of unconditional means, H_t is the conditional covariance matrix, D_t contains the conditional standard deviations on the diagonal axis and R_t is a correlation matrix containing the conditional correlations of the standardized residuals ε_t. The conditional variance matrix is thus determined by the product of the correlation and the univariate standard deviations of the bank and the sovereign.

$$[H_t]_{i,j} = \sqrt{\sigma_{it}\sigma_{jt}}\rho_{ijt} \qquad \text{where } i, j = \{\mathit{bank},\ \mathit{sovereign}\} \tag{9}$$

The conditional correlation must be specified such that the conditional covariance matrix H_t is positive definite. Therefore the conditional correlation matrix must also be positive definite. Moreover, the absolute value of the elements in the correlation matrix R_t have to be equal to or less than one by definition. To ensure that both requirements hold, a matrix Q_t is constructed of which the elements $(q_{i,j,t})$ are computed as follows:

$$q_{i,j,t} = (1 - a - b)\bar{\rho}_{i,j,t} + a\nu_{i,t-1}\nu_{j,t-1} + bq_{i,j,t-1} \tag{10}$$

$$\bar{\rho}_{i,j,t} = \sum_{l=1}^{K^{i,j}} \phi_l(\omega_{i,j})c_{i,j,t-l} \tag{11}$$

$$c_{i,j,t} = \frac{\sum_{k=t-N_{i,j}}^{t} \nu_{i,k}\nu_{j,k}}{\sqrt{\sum_{k=t-N_{i,j}}^{t} \nu_{i,k}^2}\sqrt{\sum_{k=t-N_{i,j}}^{t} \nu_{j,k}^2}} \tag{12}$$

From which correlations are computed as:

$$\rho_{i,j,t} = \frac{q_{i,j,t}}{\sqrt{q_{i,i,t}}\sqrt{q_{j,j,t}}} \tag{13}$$

The parameters a and b are scalars, and have similar interpretations as the ARCH and GARCH parameters in univariate GARCH models. The values of $q_{i,j,t}$ are normalized to a correlation by dividing it by the square root of $q_{i,i,t}$ and $q_{j,j,t}$. The long-run correlation, $\bar{\rho}_{i,j,t}$ is estimated as the weighted sum of monthly correlations, $c_{i,j,t-l}$. Similar to the estimation of the long-run component of the univariate volatilities, we use weights from a Beta distribution, captured by ϕ_l. To ensure that H_t is a positive definite matrix the parameters a and b should be positive and their sum should be less than 1. Additionally, the unconditional covariance is used as starting value for both $\nu_{i,0}\nu_{j,0}$ and $q_{i,j,0}$. The parameters of the DCC model are obtained using the quasi-maximum likelihood (QML) algorithm of Bollerslev & Wooldridge (1992)[1].

3.2 Bank-specific variables, BRRD and bank-sovereign correlation

For different types of individual bank-sovereign correlations, we estimate whether or not a series of bank-specific variables, and various stages in the implementation of the BRRD are relevant in determining the change in correlation. Hence, the following model is estimated:

$$CORR_{i,t} = \alpha_i + \lambda_t + \beta_k \sum_{k=1}^{K} BSV_{k,i,t} + \gamma BRRD_{2015,t} + \theta BRRD_{2016,t} + \varepsilon_{i,t} \tag{14}$$

where $CORR_{i,t}$ represents either the quarterly estimated correlation of changes in bank and sovereign credit risk, or the long-term component of the dynamic correlation estimated in the DCC-MIDAS model. Both estimates of the correlation between bank and sovereign credit risk thus have a quarterly frequency t. The k^{th} fundamental of bank i is contained in the vector $BSV_{k,i}$. We capture the effect of the implementation of the BRRD by adding a dummy variable which starts either in Q1/2015 or Q1/2016. The former captures the quarters that the BRRD came into full effect (i.e. the 1st of January 2015), while the latter

[1]We explain the optimization procedure in more detail in Appendix A

coincides with the quarter when the SRM became fully operational. The model controls for unobserved heterogeneity at the bank level by including bank fixed effects (α_i). We also include quarterly fixed effects to filter out the effects of common shocks (λ_t).

3.3 Narrative SVAR

The estimation of the dynamic correlation allows us to identify whether the relation between sovereign and bank credit risk changes over time, a higher level of correlation representing a higher level of interconnectedness between bank and sovereign credit risk. However, the dynamic correlation does not allow us to identify which transmission channels are causing these changes.

The objective of the BRRD is to avoid government support for banks in times of distress. If the BRRD is credible, then it should decrease the link between bank and sovereign credit risk, specifically after an exogenous bank shock. As the BRRD does not regulate the degree of sovereign exposures we assume it has no impact on the transmission of exogenous sovereign shocks. Therefore, we focus on the transmission of banks shocks and estimate their impact on bank and sovereign credit risk.

We examine the reaction of bank and sovereign credit risk on a series of exogenous bank shocks in a SVAR model, using the market reaction of bank CDS spreads on earnings announcements in a narrative identification approach. We estimate the SVAR in 2 regimes: a regime prior to the BRRD (2011-2013) and a regime after BRRD was fully implemented (2018-2020). The BRRD, first proposed in 2014, was finally written into law in 2016, which makes it unfavorable to use a specific cut-off point to separate both regimes. Additionally, in 2017 BRRD regulation was applied twice with the bail-in of Banco Popular Espanol and MPS. Therefore, we omit observations between 2014-2017, capturing both the implementation period of BRRD and its first application. From 2018 we assume that market expectations have adjusted to the environment of BRRD regulation. To identify a set of exogenous bank shocks, we use earnings releases, which contain information on the earnings and profitability of the bank during the previous quarter. The announcement day itself is usually set a number of weeks after the fiscal end of the quarter and is not influenced by sovereign or market conditions. We identify the shock as the change of the CDS spread during the day that the earnings were announced. We assume that markets have a priori expectations with respect to the earnings of the bank and the deviation from these expectations is captured by the change of banks' CDS spreads. To reduce the possibility of confounding events, we control for macroeconomic (e.g. a monetary policy announcement of the ECB which had a clear effect on financial markets) and sovereign events (e.g. news related to a political crisis, or an election result) that potentially contribute to CDS spread changes. We examine headlines of the Financial Times newspaper on each of the announcement days in the sample. Announcement days where we identify potential confounding events are omitted from the sample.

The SVAR is estimated on a number of financial variables which we assume that are potentially impacted by exogenous bank shocks and are endogenously related to one another. The included variables are the CDS spread of both the banks and the sovereign, a stock and volatility index of the country and a broader European stock and credit risk index for which we use the STOXX Europe 600 and Itraxx index respectively. The SVAR is thus:

$$Y_t = A(L)Y_{t-1} + R\varepsilon_t \tag{15}$$

where Y_t is an N-dimensional vector of the endogenous variables. The narrative bank shock is zero on non announcement days and has the value of the change in bank CDS spreads on earnings announcements. The orthogonal structural innovations (ε_t) is an N-dimensional vector with mean zero, and $A(L)$ and R are $N \times N$ time-invariant parameter matrices. The reduced-form residuals corresponding to this structural model are given by the relationship $\eta = R\varepsilon$. Following the narrative approach, applied in Ramey (2011), we assume that the bank shock is exogenous to the other variables in the SVAR. We identify the impact matrix R using a Cholesky decomposition, where the bank shock is exogenous to all other shocks. We make no assumptions on the order of the other shocks. However, this is not a problem as the underlying order of the other shocks has no impact on the transmission of the bank shock to the variables in the system. The transmission of the other shocks are not within the scope of this paper and are discarded.

4 Results

4.1 Dynamic correlation of bank and sovereign credit risk

The estimation of the dynamic correlation of bank and sovereign credit risk proceeds in two steps. The first step requires the estimation of the conditional volatilities of the individual bank and sovereign CDS spread. In Table 2 we analyze the parameters of the GARCH-MIDAS models, estimated for bank and sovereign credit risk and for each country separately. The model is estimated with a monthly frequency on the MIDAS component, and we set the number of lagged RV's to 3 years, which means that K in Equation 6 is set to 36. The volatilities of bank and sovereign CDS spreads are mainly driven by the parameters α and β. The former captures the impact that shocks have on the volatility, whilst the latter is a measure of the persistency of new shocks to the volatility. Shocks from Italian and Spanish banks are more persistent than those of French and German banks. With regards to the sovereign volatilities, it is noteworthy that the Italian sovereign has the highest persistency. The long-term component in GARCH-MIDAS is identified from the parameters m, θ and ω. The value m captures the long-run mean of the long term component, we note that this is higher for the volatility of sovereign CDS spreads than banks' CDS spreads. The degree to which the lagged realized variances (RV) are allowed to impact the long term component is displayed in the value of θ. A lower θ value allows for a lower impact of RV's on the long-term component. Finally, higher values for the parameter ω give a higher weight to the most recent RV's. The parameters are optimized through a quasi ML estimator.

In Figure 6 we display the results of the GARCH-MIDAS models graphically. The green curve shows the short-run daily component which is a GARCH With regard to the long term component of the volatilities of both banks and the sovereign are visibly higher during the sovereign debt crisis and the corona pandemic. The peaks in CDS volatility are driven by common macroeconomic events such as the onset of the corona pandemic, as well as country specific events, e.g. the Italian general election in 2018 led to an increase in the volatility of both Italian bank and sovereign credit risk.

We capture the dynamic correlation between bank and sovereign credit risk through the estimation of a DCC-MIDAS model on the standardized residuals obtained from the

GARCH-MIDAS model. Similar to the GARCH-MIDAS model, we obtain a short- and long-run component, the former captured by a DCC-GARCH process and the latter through RV's. We show the estimated parameters of the DCC-MIDAS model in Table 3 for each country. Similar to the parameters of the GARCH-MIDAS model, α captures the degree to which short-term deviations from the long-run correlation are possible and β captures the persistence of these temporary deviations. Finally, ω is the factor that determines how lagged RV's influence the long-run component of the dynamic correlation. A higher value for ω gives a higher weight on recent RV's, and vice versa a lower ω shifts the weight on lagged RV's more equally. These parameters are optimized with a quasi ML estimator.

In Figure 7 we show the results of the DCC-MIDAS model graphically. The blue curve represents the long-run component of the dynamic correlation around which the short-run component deviates, displayed by the green curve. The short-run component is thus captured by a mean reverting DCC model, and where the mean is captured by the MIDAS specification. Similar to the estimation of the conditional volatility, the sample period starts in 2010-2011. The estimation of the long-run component is accomplished by the RV's of 36 months in total, once for the estimation of the GARCH-MIDAS and also for the estimation of the DCC-MIDAS. Therefore, the first observation for the long-run component is obtained after 72 months, which is in 2010-2011.

The results indicate that correlation between bank and sovereign credit risk, captured by CDS spreads, has decreased over the sample period. For the peripheral countries (Italy and Spain) this decrease starts in 2016, which coincides with the implementation of BRRD. The correlation in Germany decreases from 2013 onward. Finally, the French correlation also decreases, but without a clear trend and higher volatility. From the DCC-MIDAS analysis it is clear that the correlation between bank and sovereign credit risk has decreased between 2011-2020.

4.2 Bank-specific variables, BRRD and bank-sovereign correlation

In Table 6 we analyze the drivers of bank-sovereign correlation of the individual banks in our sample [2]. The results show the impact of bank-specific variables and BRRD dummy variables on either the quarterly estimated correlation of bank and sovereign CDS spread changes or the long-run component of the bank-sovereign correlation estimated in a DCC-MIDAS model. The bank-specific variables are categorized into capital adequacy, asset structure, income diversification, asset quality, profitability and funding mix. We saturate each model with bank and quarterly fixed effects and we cluster the standard errors at the bank level.

We find that the unweighted equity/asset ratio (CAP) remains insignificant in each specification, indicating that within banks, changes in the capital ratio are not perceived as reducing the bank-sovereign correlation. On the contrary, the long-run component of the dynamic correlation seems to be positively associated with an increase in the risk-weighted capital ratio (CET1), suggesting that risk-weighted capital ratios are potentially biased measures of buffers against unexpected losses.

In terms of balance sheet structure, a higher LTA ratio implies a higher bank-sovereign correlation, suggesting that banks primarily focused on lending are perceived as having a higher interconnectedness with their sovereign than banks with a more diversified assets structure.

With respect to the implementation of the BRRD, we observe that the dummy variable that captures the implementation of Q1/2015 is negatively associated with the bank-sovereign correlation and highly significant, with decreases in the correlation ranging from 0.12-0.21. The additional effect of the implementation of BRRD in Q1/2016 is insignificant, suggesting that their was no additional level shift in the bank-sovereign correlation post-2015.

[2]The estimation results of the bank level bank-sovereign correlations are obtainable on demand for the interested reader.

4.3 SVAR

In Figures 8-11 we report the transmission channel of exogenous bank shocks on the CDS spread of the bank and sovereign, the MSCI index and its volatility of that country, and the STOXX Europe 600 and Itraxx indices, which capture the European economic conditions. We measure these transmissions in 2 regimes: a regime where the transmission is not yet affected by BRRD regulation (pre-BRRD) and a regime where the transmission should fully incorporate the new dynamic due to the implementation of the BRRD directive (BRRD). For this reason the first regime is estimated during 2010-2013, and the second regime during 2018-2020. Over the period 2014-2017 we assume that BRRD is either being negotiated and implemented, or that the market is unsure as to how the implementation will effectively take place. After the resolution of two major banks, Banco Popular Espanol and Banca MPS in 2017, we assume that market participants have adjusted to the BRRD directive and that the transmission of exogenous bank shocks from 2018 onwards therefore fully captures the new regulation. The transmission channel is estimated in an SVAR, where we identify an exogenous bank shock by including a variable that captures the change in bank CDS spreads on earnings releases. We omit those days where we find confounding effects with respect to the macro economy or the sovereign, by analyzing FT newspapers. This shock is added to the SVAR and put first in a Cholesky decomposition. In the results we show the impulse responses stemming from the exogenous bank shock. The impulse responses are calibrated such that a bank shock increases the bank CDS spread with 2 basis points, which allows for a straightforward interpretation of the results. An exogenous bank shock is thus calibrated to increase bank credit risk.

The results in Figure 8 show the transmission of exogenous bank shocks in Italy. The results indicate that prior to the BRRD, bank shocks lead to a persistent increase in bank credit risk that remains significant over 40 days after the shock. This is accompanied by an increase in sovereign credit risk, suggesting the interconnectedness of banks and sovereigns. This increase is lower and less persistent than the reaction of bank credit risk. With respect to the control variables, we notice that the volatility of the MSCI Italy increases marginally for a brief period and the MSCI and STOXX Europe 600 indices are negatively affected by a bank shock. This is unsurprising as the public banks in our sample are constituents in both indices. Finally, we note that there is a short lasted positive increase of the credit risk in the European economy after an Italian bank shock of 2 basis points. After BRRD implementation, we note that bank shocks are less persistent on the increase in bank credit risk. Sovereign credit risk increases still increases after exogenous bank shocks. However, the impact is lower than pre-BRRD, suggesting that in the BRRD regime, the transmission of bank shocks to the sovereign has decreased. The volatility no longer increases after a bank shock, suggesting that this type of bank shocks (i.e. after earnings releases) no longer increase uncertainty on Italian financial markets. The other control variables have similar responses to the bank shocks in comparison with the pre-BRRD regime.

In Figure 9 we analyze the transmission of bank shocks in Spain. Bank shocks increase bank CDS spreads. However, in Spain, we note that the transmission of exogenous bank shocks, based on earnings releases, lead to a more persistent transmission into bank credit risk after the implementation of the BRRD. Prior to the BRRD, we find no evidence that these shocks affected sovereign credit risk. However, during the BRRD regime, we notice that sovereign credit risk is positively affected by bank shocks. Therefore, this result suggests that despite the implementation of the BRRD, the interconnectedness between banks and sovereigns has increased. The control variables suggest that after the implementation of the BRRD, Europe wide credit risk, captured by the Itraxx, also increases after exogenous bank shocks in Spain.

Figure 10 displays the impulse response functions of exogenous bank shocks in Germany. Similarly to Italy and Spain, we find that these bank shocks significantly increase bank credit risk. This transmission is more persistent prior to the implementation of the BRRD. In Germany we find no evidence of a strong interconnectedness between banks and sovereigns after the occurrence of a bank shock which captures the surprise component of an earnings release. Sovereign credit risk increases slightly, but the increase is more limited than that of the bank and the effects are short-lived. However, during the BRRD regime, we note that the effect of bank shocks on the sovereign decreases further. With respect to the control variables, we find similar transmissions as in Italy and Spain. Volatility tends to increase, and the stock market indices of Germany and Europe decrease after a bank shock. However, these effects are less pronounced after the implementation of BRRD. Europe wide credit risk increases after bank shocks, but the transmission is less persistent during the BRRD regime.

Finally, in Figure 11 we estimate the transmission of bank shocks in France. The exogenous bank shocks tend to increase credit risk of banks, however the transmission is less persistent compared to the credit risk in Italy and Germany. With respect to the transmission of bank shocks to the sovereign, the transmission shows a delayed reaction, however only temporary. This result suggests that banks and the sovereign in France are not as interconnected as in e.g. Italy. After the implementation of BRRD, we see a further reduction of this interconnnectedness. Sovereign credit risk no longer increases after a bank shock. Prior to the BRRD we show that the volatility increases, with a delay, and that the MSCI and STOXX Europe 600 index decrease, while we also note an increase in Europe wide credit risk. These transmissions are all subdued after the implementation of BRRD.

The results in Figures 8-11 confirm that the transmission of exogenous bank shocks, captured by the surprise element in earnings releases, did in fact change after the implementation of BRRD. Sovereign credit risk has a lower and less persistent reaction to these shocks after BRRD was implemented, which suggests that the bank-sovereign risk nexus has diminished after BRRD, with the exception of Spain. This confirms the possibility of country-level heterogeneity with respect to the effectiveness of BRRD regulation.

5 Conclusion

Since 2008 European banks and sovereigns have witnessed a stressful decade, with a banking crisis and a sovereign debt crisis, during which sovereigns performed bailouts of banks to safeguard the financial sector. In order to avoid this doom loop between banks and sovereigns, the European Banking Union was established, consisting of three pillars: a Single Supervisory Mechanism (with the ECB in the lead), a unified recovery and resolution framework regulated by the Bank Recovery and Resolution Directive, and a European deposit insurance mechanism.

The objective of this paper is to assess the effectiveness of the second pillar of the banking union, i.e. the BRRD regulation. We perform this analysis for four countries in the euro area, two peripheral countries (Italy, Spain) and two core countries (France and Germany). Theoretically, if the bail-in framework introduced by the BRRD directive is considered by market participants as credible, the probability of bank bailouts should decrease and we should observe a significant decrease in the interconnectedness between banks and their sovereign. Unanticipated shocks to the credit risk of banks should not, or to a lower extent, be transmitted to the credit risk of the sovereign. This is our motivation to empirically investigate whether or not the implementation of BRRD has diminished the interconnectedness of bank and sovereign default risk, proxied by bank and sovereign CDS spreads.

We implement a novel approach to estimate the interconnectedness of bank and sovereign credit risk, using a DCC-MIDAS model from which we obtain the dynamic correlation of bank and sovereign default risk. Our results show that the correlation between bank and sovereign credit risk has diminished for all countries over the period 2011-2020. This decrease is most pronounced in Germany, where the correlation shows a clear downward trend from 2013 onwards. In the peripheral countries, Italy and Spain, the dynamic correlation also exhibits a downward trend, which sets in after 2016. In France, the correlation decreases, although it is more susceptible to positive and negative shocks. Hence, a main finding is that the dynamic correlation between bank and sovereign credit risk has decreased significantly over the period in which the BRRD has been negotiated and implemented.

Next, we investigate four potential drivers of the interconnection between bank and sovereign credit risk: the BRRD implementation versus three bank-specific features (capital adequacy, size and the banks' exposures to their sovereign). Capital adequacy, which captures the buffer of banks against unexpected losses, and size, as a proxy for the too-big-to-fail status of the banks, are insignificant. Also higher exposure to the home sovereign, through sovereign bond holdings, is not associated with a higher bank-sovereign correlation. Hence, we do not find that the bank-sovereign nexus is affected by bank-specific features that are linked to their bail-out probability. The strong significance of the dummy variable capturing the period after January 1st 2015, i.e. the effective entry into force of the BRRD, confirms that the BRRD has a significant downward impact on the dynamic correlation between bank and sovereign credit risk.

If the BRRD bail-in framework is judged by investors to be credible, exogenous bank credit risk shocks should no longer affect sovereign risk in the BRRD regime. To investigate this line of causality, we adopt an SVAR approach to estimate whether or not the transmission of exogenous bank shocks to sovereign credit risk has changed after the implementation of the BRRD. We estimate the SVAR in both the pre-BRRD and BRRD regimes. We identify a series of exogenous bank shocks captured by the change of the CDS spread of banks on earnings announcement days. To control for confounding effects, we omit days on which the Financial Times reports major economic events or news related to the domestic sovereign. To control for the state of the domestic and European economy as well as the domestic volatility and economywide credit risk, we add the country-specific MSCI index, the STOXX Europe 600, the volatility of the MSCI index and the Itraxx. Our results confirm that the bank-sovereign feedback loop diminishes after the implementation of the BRRD, although not in each country. In Italy, the results show that an exogenous bank shock affects the sovereign CDS also in the BRRD regime, although the change in sovereign credit risk is lower in magnitude compared to the pre-BRRD era. In Spain, an exogenous bank shock even provokes a significant increase in sovereign credit risk in the BRRD regime compared with the pre-BRRD period. Hence, investors judge that the bank-sovereign nexus remains alive in the two peripheral euro area countries in the period during which the BRRD is in force. For Germany we do find that the response of sovereign credit risk to an exogenous bank shock is lower after the introduction of the BRRD. Similarly, in France there is no impact of bank risk shocks on sovereign credit risk in the BRRD era. Hence, the bail-in framework is judged as credible in the core countries of the euro area.

We plan two additional analyses to investigate the robustness of the results. One is to make a distinction between negative and positive bank earnings announcement surprises to investigate whether there is asymmetry in the response of sovereign credit CDS spreads. The other is to use a series of actual bail-in events as shocks and analyze their impact on the dynamic bank-sovereign correlation.

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Tables

Table 1: Descriptive statistics. CDS spread data and market data are obtained from Markit and Refinitiv respectively.

Variable	Mean	SD	P1	P50	P99
CDS senior Germany	87.21	52.01	8.35	82.44	257.13
CDS senior Italy	160.25	128.69	9.07	136.84	579.32
CDS senior Spain	163.24	156.67	9.03	114.31	673.66
CDS senior France	83.90	69.55	6.25	67.70	309.43
CDS sovereign Germany	16.52	16.08	1.59	10.05	69.05
CDS sovereign Italy	113.87	94.73	6.10	97.01	448.90
CDS sovereign Spain	97.72	99.61	2.49	64.14	425.04
CDS sovereign France	32.59	31.93	1.57	22.98	145.44
MSCI Italy	99.96	21.29	60.34	99.85	148.71
MSCI Spain	155.65	28.40	98.76	156.12	203.55
MSCI France	167.54	50.97	88.04	156.92	284.21
MSCI Germany	198.80	64.02	98.37	185.29	309.24
Eurostoxx 600	173.05	51.17	84.54	161.89	276.86
Volatility MSCI Germany	1.21	0.57	0.63	1.06	3.77
Volatility MSCI Italy	1.41	0.71	0.63	1.25	4.28
Volatility MSCI France	1.21	0.62	0.57	1.04	3.92
Volatility MSCI Spain	1.38	0.68	0.68	1.19	4.46
Correlation Germany	0.22	0.11	0.00	0.19	0.45
Correlation Italy	0.60	0.13	0.28	0.63	0.82
Correlation Spain	0.48	0.14	0.07	0.49	0.77
Correlation France	0.34	0.14	0.00	0.33	0.69

Table 2: Estimation results of the GARCH-MIDAS model

Estimation	mu	alpha	beta	m	theta	omega
Italy banks	−0.11	0.16	0.76	4.28	0.02	3.72
Spain banks	−0.07	0.17	0.80	7.21	0.01	0.54
Germany banks	−0.04	0.22	0.68	6.89	0.01	9.78
France banks	−0.06	0.22	0.66	6.68	0.02	3.80
Italy sovereign	−0.05	0.16	0.81	25.54	0.00	9.63
Spain sovereign	−0.06	0.19	0.71	16.79	0.01	−0.13
Germany sovereign	0.04	0.19	0.69	42.74	0.00	6.56
France sovereign	−0.06	0.20	0.72	11.61	0.06	0.47

Table 3: Estimation results of the DCC-MIDAS model

Estimation	alpha	beta	omega
Italy	0.06	0.72	6.07
Spain	0.07	0.81	2.59
Germany	0.03	0.79	0.88
France	0.12	0.59	0.19

Table 4: Overview of the earnings announcement days for each country during the pre-BRRD bail-in regime (2011-2013). Events shows the total number of announcement days for each country. Next, the second column shows the number of these days where we found confounding sovereign and/or macroeconomic announcements, obtained from the Financial Times. Through the reaction of the bank CDS spread we determine whether the event was perceived as positive (CDS decrease) or negative (CDS increase), which are shown in columns 3 and 4.

Country	Events	Omitted	Positive	Negative
Italy	53	21	18	14
Spain	57	30	17	10
France	51	23	18	10
Germany	36	10	17	9

Table 5: Overview of the earnings announcement days for each country during the post-BRRD bail-in regime (2018-2020). Events shows the total number of announcement days for each country. Next, the second column shows the number of these days where we found confounding sovereign and/or macroeconomic announcements, obtained from the Financial Times. Through the reaction of the bank CDS spread we determine whether the event was perceived as positive (CDS decrease) or negative (CDS increase), which are shown in columns 3 and 4.

Country	Events	Omitted	Positive	Negative
Italy	28	12	10	6
Spain	53	20	15	18
France	53	13	20	20
Germany	31	9	13	9

Table 6: Baseline regression results. This table shows the results explaining the drivers of the bank-sovereign correlation using bank-specific variables, and BRRD implementation variables (BRRD1 = 1 from 1/1/2015; BRRD2 = 1 from 1/1/2016). The correlation is either estimated by the quarterly correlation between bank- and sovereign CDS spreads, or the long-run component of a DCC-MIDAS model between bank- and sovereign CDS spreads. All estimations use bank- and quarterly fixed effects to control for unobserved heterogeneity and common macroeconomic shocks. Standard errors in parentheses are clustered at the country level. *, ** and *** represent significance a the 10%, 5% and 1% percent level, respectively.

	Correlation Quarterly		Correlation DCC-MIDAS	
CAP	−0.012		0.004	
	(0.021)		(0.013)	
CET1		0.004		0.020**
		(0.011)		(0.008)
SIZE	−0.144	−0.109	0.048	0.065
	(0.139)	(0.125)	(0.123)	(0.097)
LTA	0.011***	0.011***	0.008***	0.007***
	(0.002)	(0.002)	(0.002)	(0.002)
NPL	−0.007	−0.007	0.001	0.003
	(0.007)	(0.007)	(0.002)	(0.002)
ROA	0.041	0.032	0.022	0.010
	(0.028)	(0.028)	(0.024)	(0.017)
DEP	−0.001	−0.001	−0.004	−0.003
	(0.004)	(0.004)	(0.003)	(0.003)
HOME BIAS	−0.002	−0.002	−0.003	−0.004
	(0.005)	(0.006)	(0.003)	(0.002)
BRRD 1	−0.186***	−0.210***	−0.124***	−0.146***
	(0.031)	(0.038)	(0.021)	(0.029)
BRRD 2	−0.021	−0.037	−0.025	−0.087*
	(0.054)	(0.060)	(0.039)	(0.047)
Bank fixed effects	Yes	Yes	Yes	Yes
Quarterly fixed effects	Yes	Yes	Yes	Yes
R^2	0.480	0.481	0.620	0.642
No. of banks	15	15	15	15
No. of obs	387	383	387	383

Figures

Figure 1: Time series of the CDS spreads of Italian banks and the sovereign

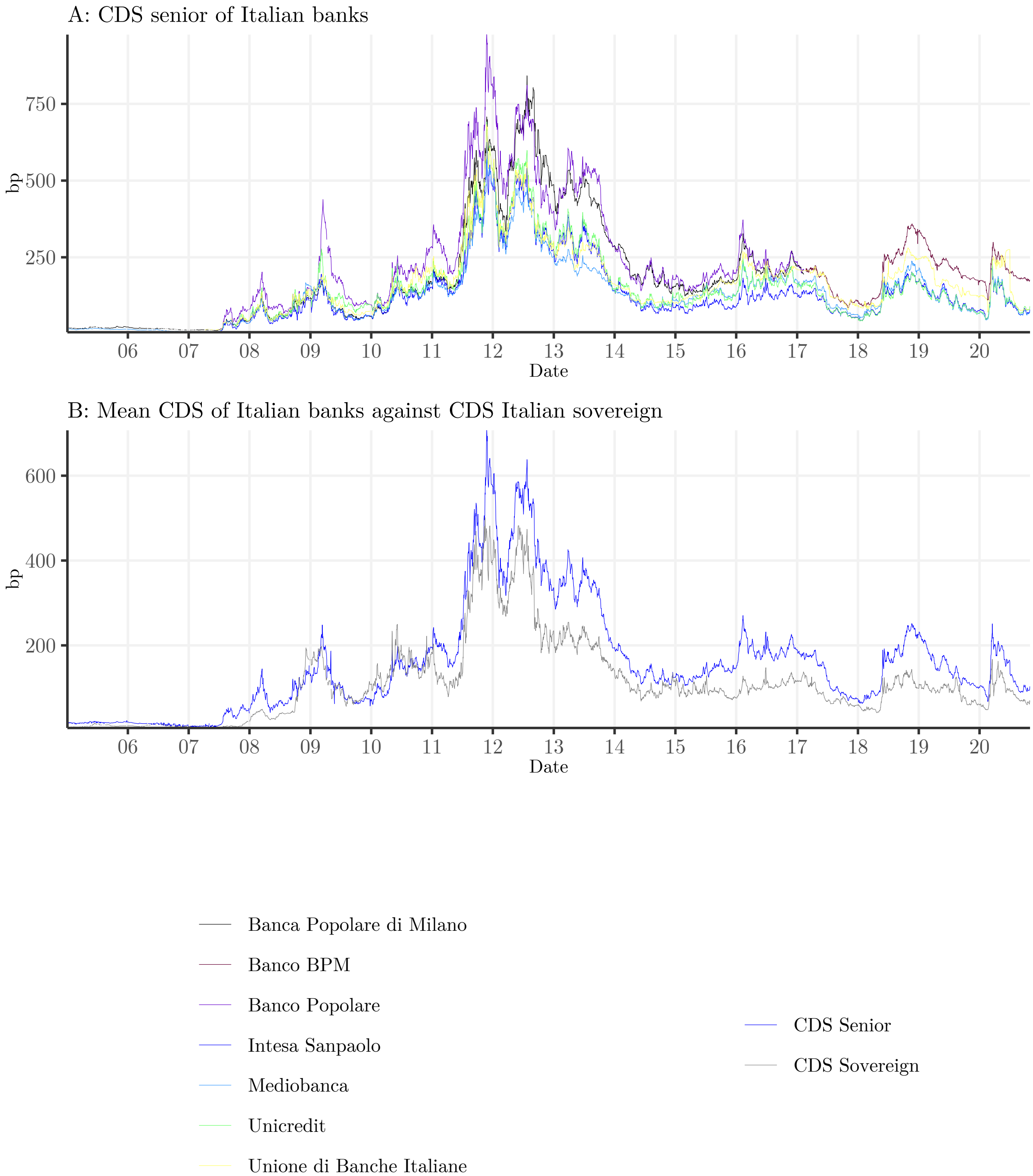


Figure 2: Time series of the CDS spreads of Spanish banks and the sovereign

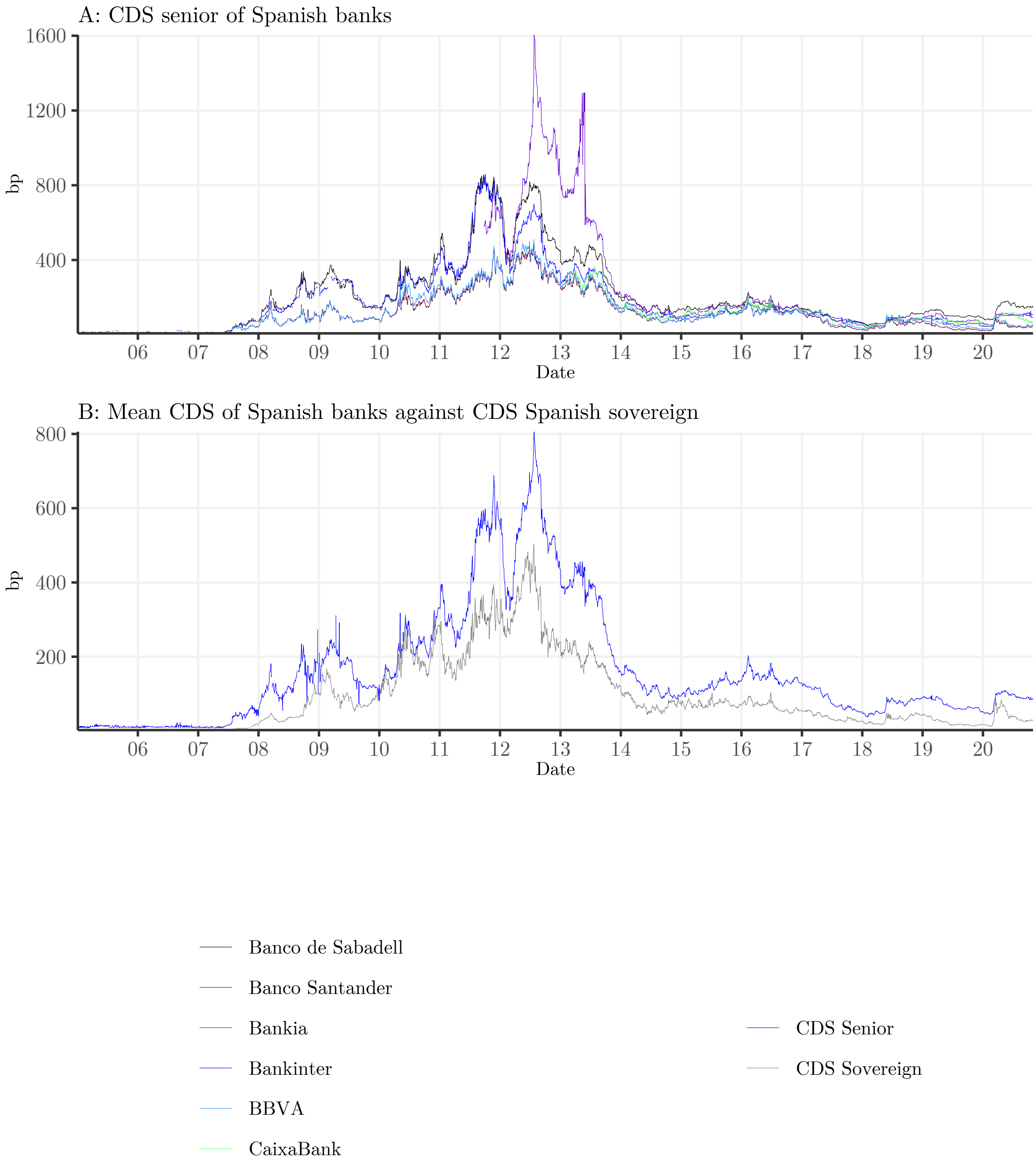


Figure 3: Time series of the CDS spreads of French banks and the sovereign

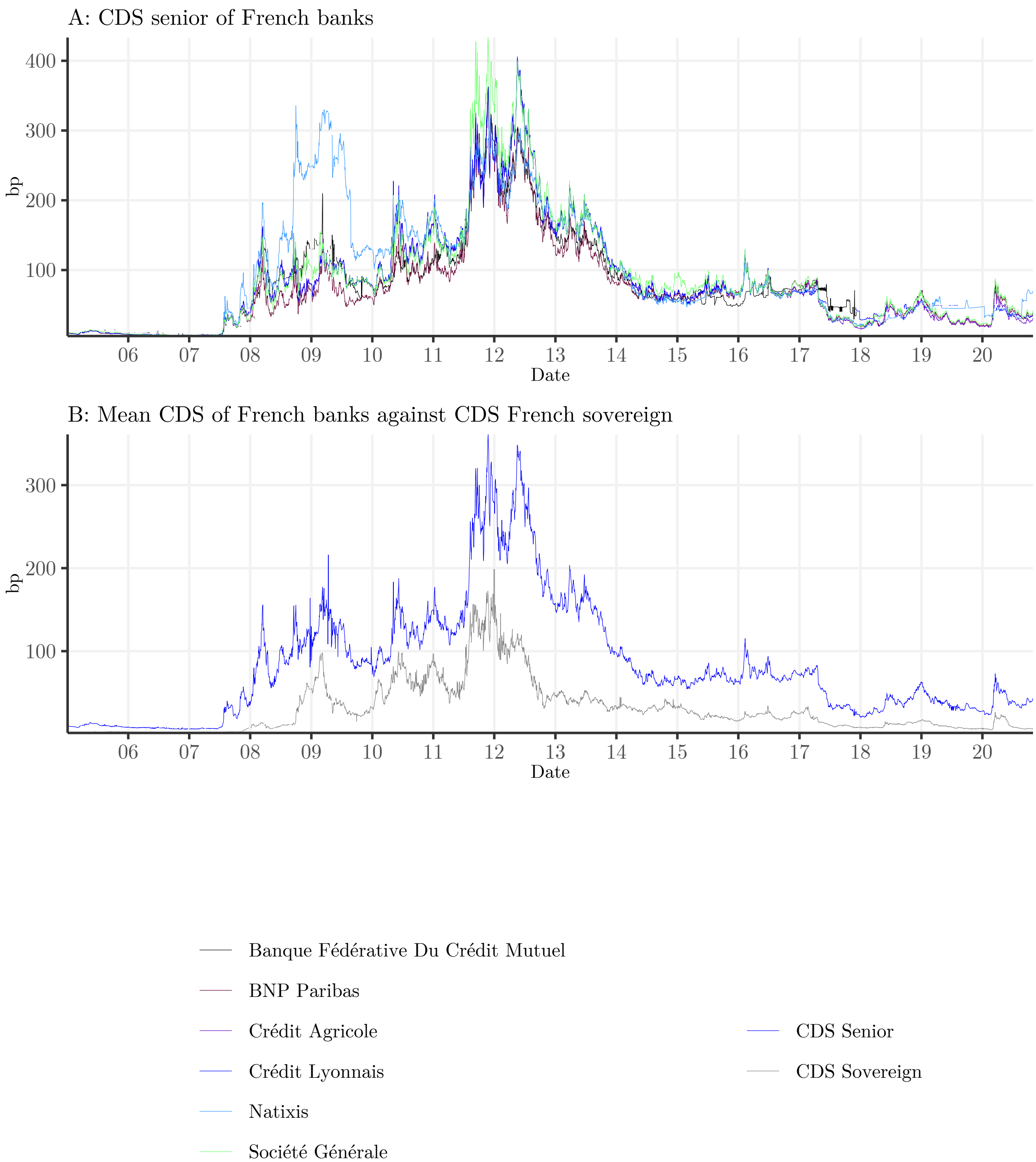


Figure 4: Time series of the CDS spreads of German banks and the sovereign

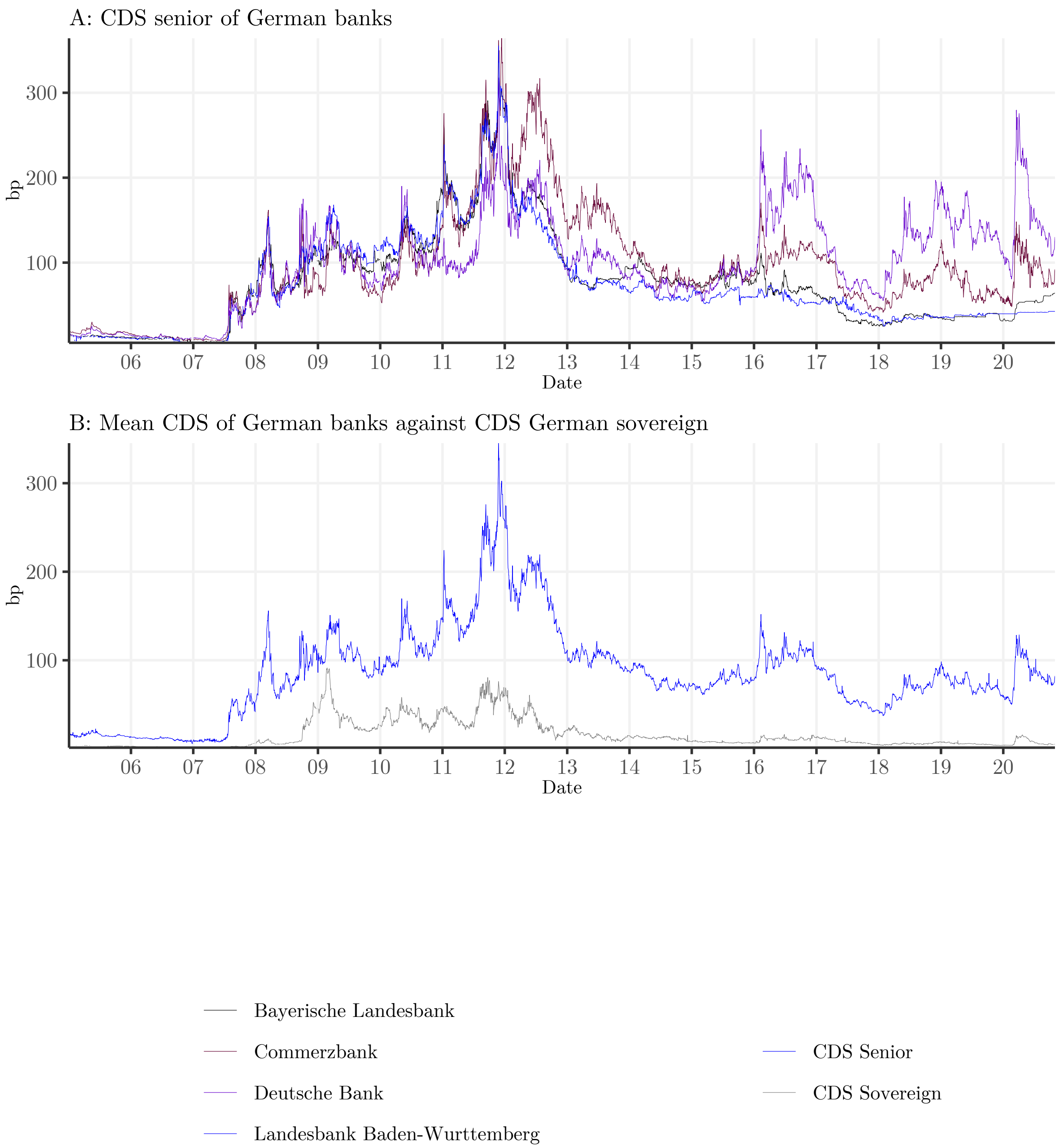


Figure 5: Panel A displays the time series of MSCI indices of Italy, Spain, Germany and France, and the STOXX Europe 600 (index base = 31/12/2004). Panel B contains the time series of the conditional volatilities of the respective MSCI indices, estimated with a GARCH(1,1) model.

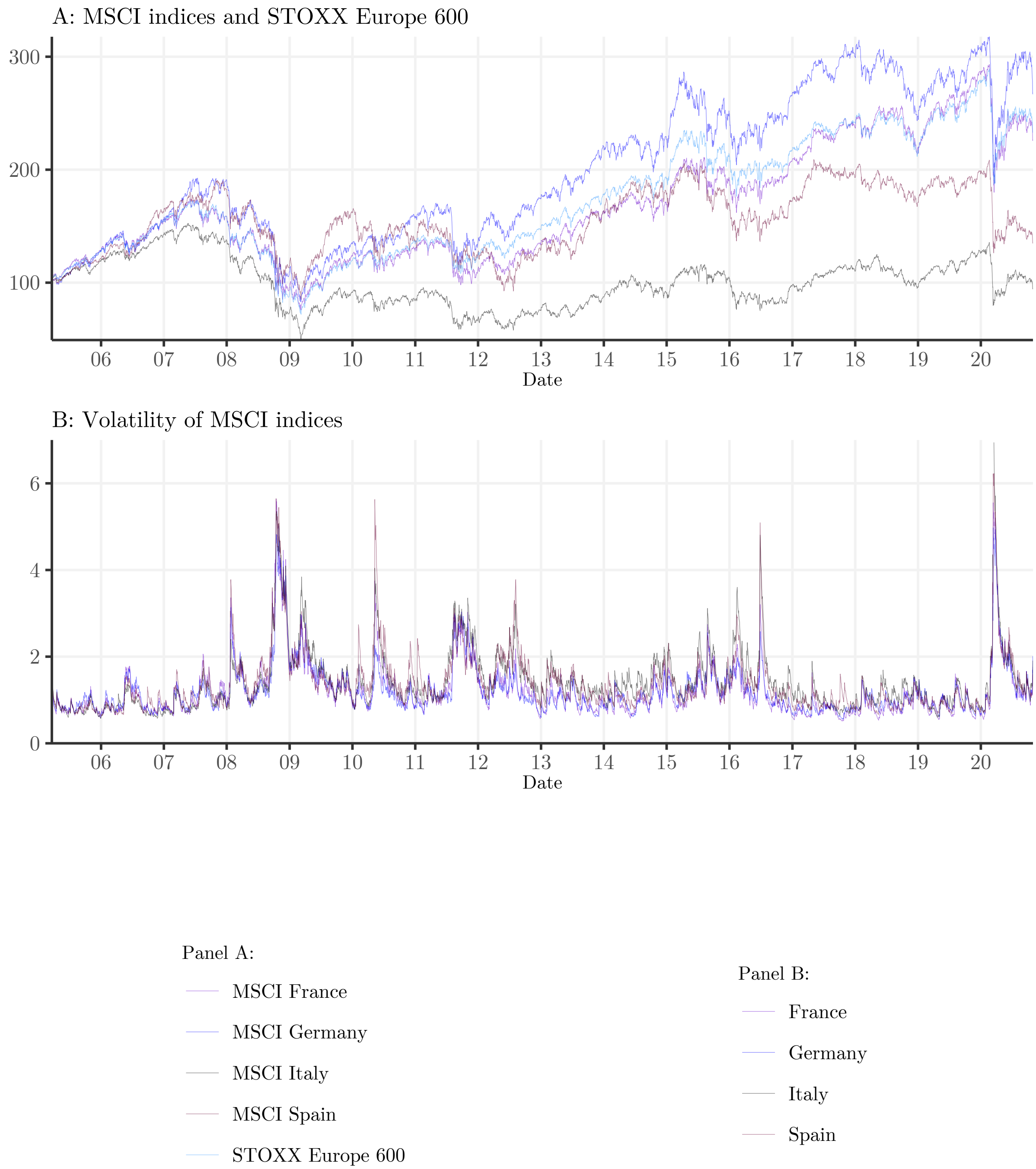


Figure 6: Time series of the dynamic conditional volatility of both the bank and sovereign CDS spreads estimated through the GARCH-MIDAS model.

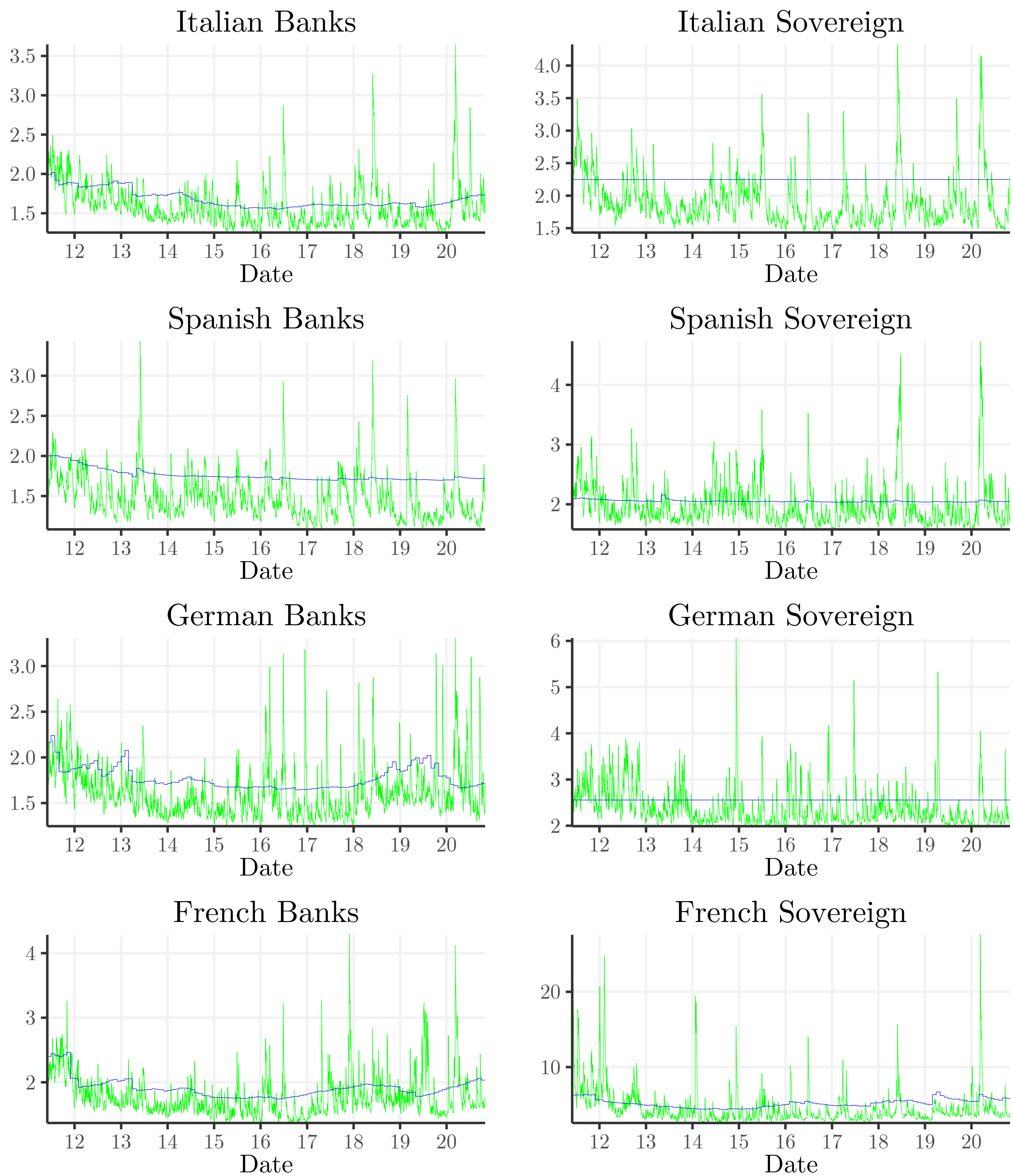


Figure 7: Time series of the dynamic conditional correlation between the log changes in the senior bank CDS and sovereign CDS spreads, estimated through the DCC-MIDAS model.

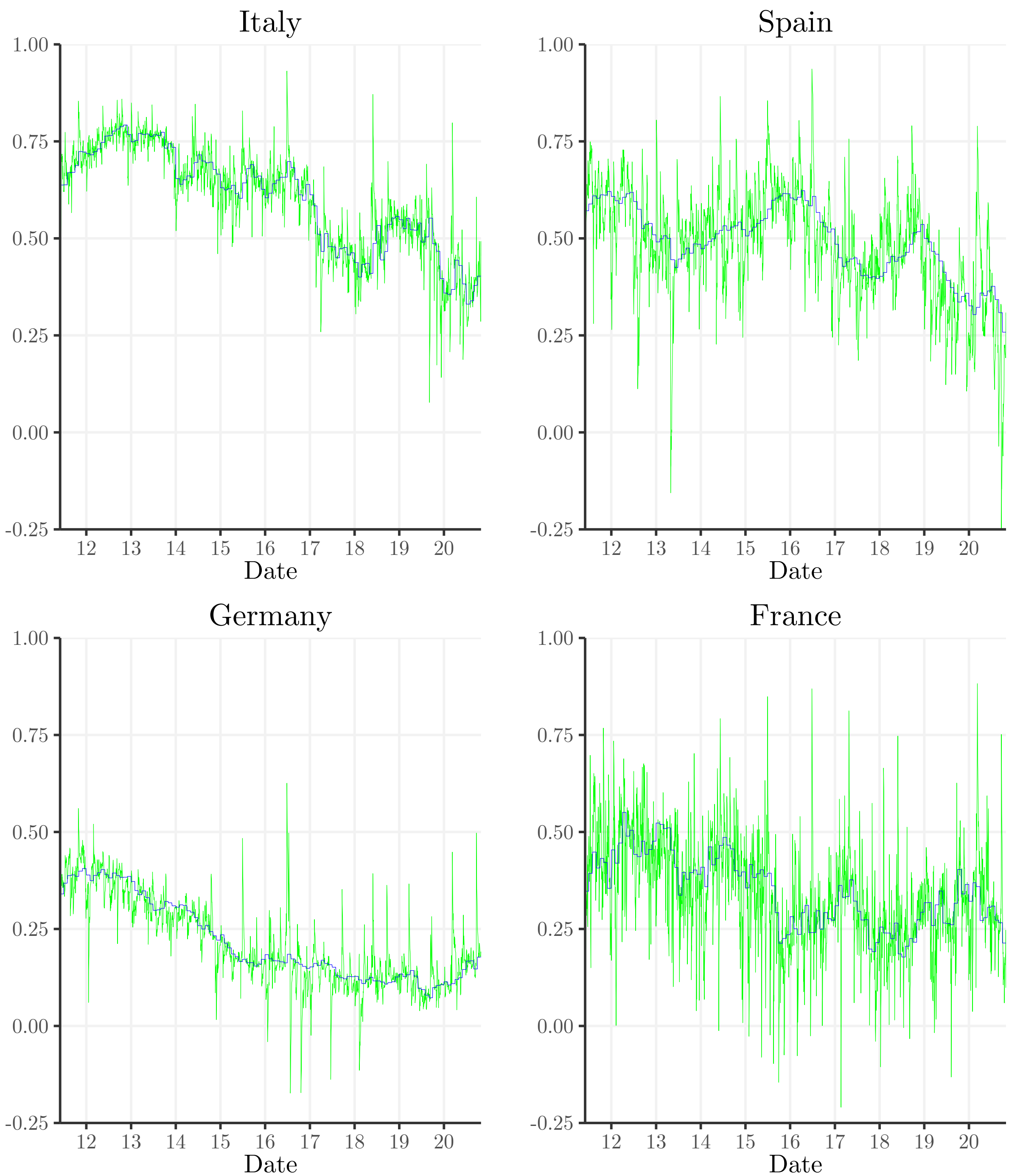


Figure 8: Impulse response function of an exogenous bank shock on the CDS spreads of Italian banks and sovereign bonds, the volatility of the MSCI Italy and the index itself, and the Itraxx, identified using the narrative approach. Panel A shows the impulse response function is estimated from the pre-BRRD period, whilst Panel B shows the impulse responses of the post-BRRD regime.

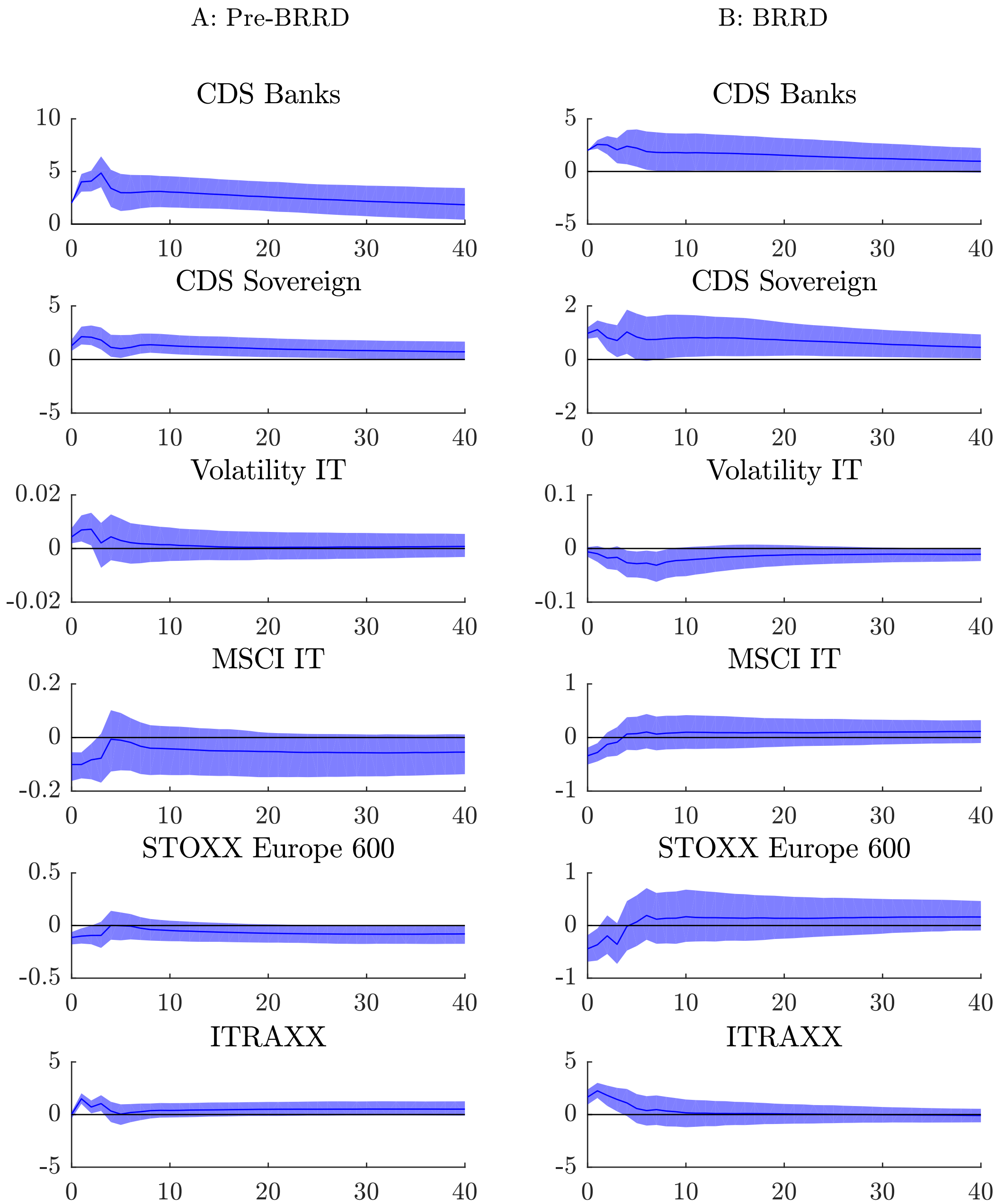


Figure 9: Impulse response function of an exogenous bank shock on the CDS spreads of Spanish banks and sovereign bonds, the volatility of the MSCI Spain and the index itself, and the Itraxx, identified using the narrative approach. Panel A shows the impulse response function is estimated from the pre-BRRD period, whilst Panel B shows the impulse responses of the post-BRRD regime.

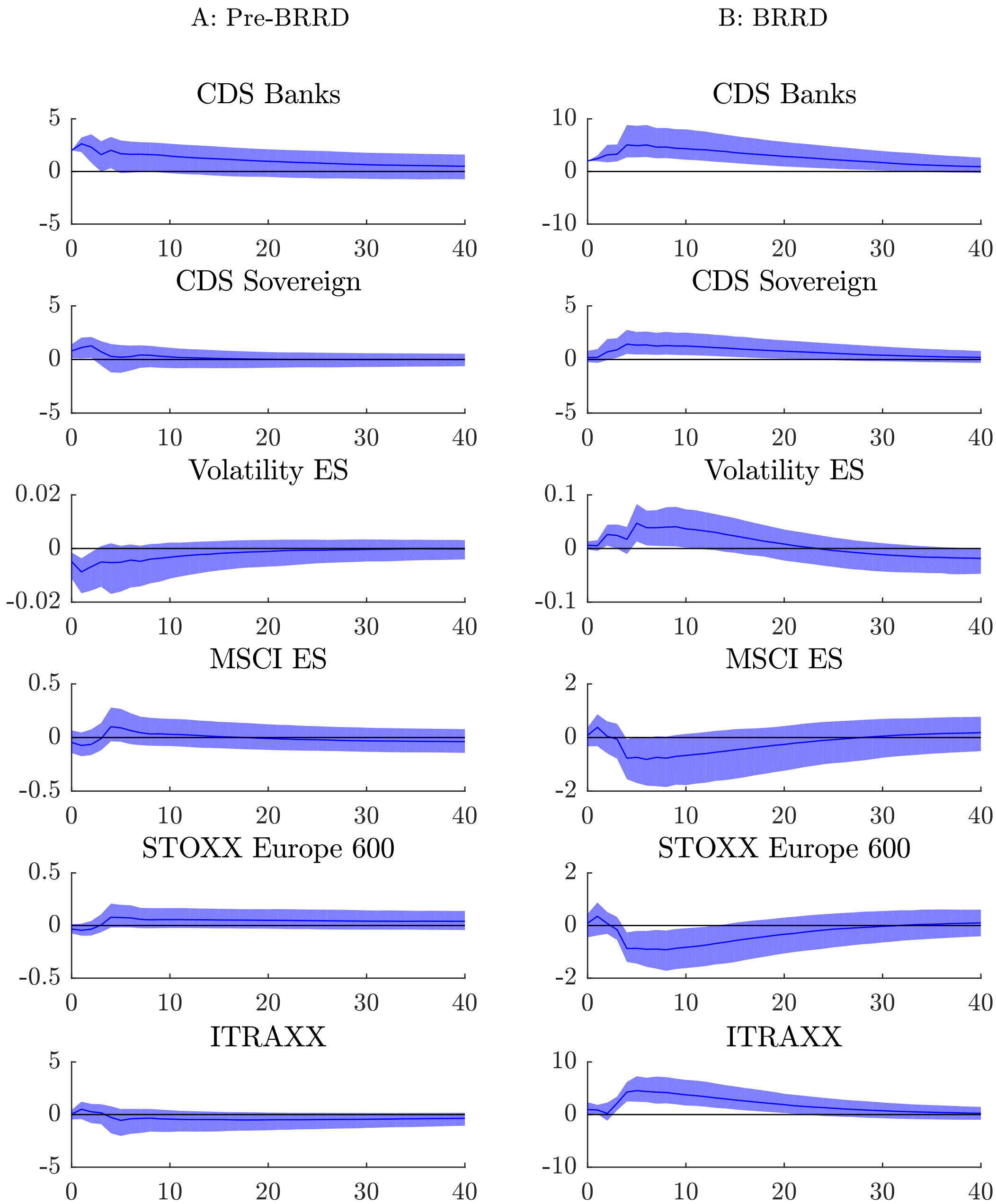


Figure 10: Impulse response function of an exogenous bank shock on the CDS spreads of German banks and sovereign bonds, the volatility of the MSCI Germany and the index itself, and the Itraxx, identified using the narrative approach. Panel A shows the impulse response function is estimated from the pre-BRRD period, whilst Panel B shows the impulse responses of the post-BRRD regime.

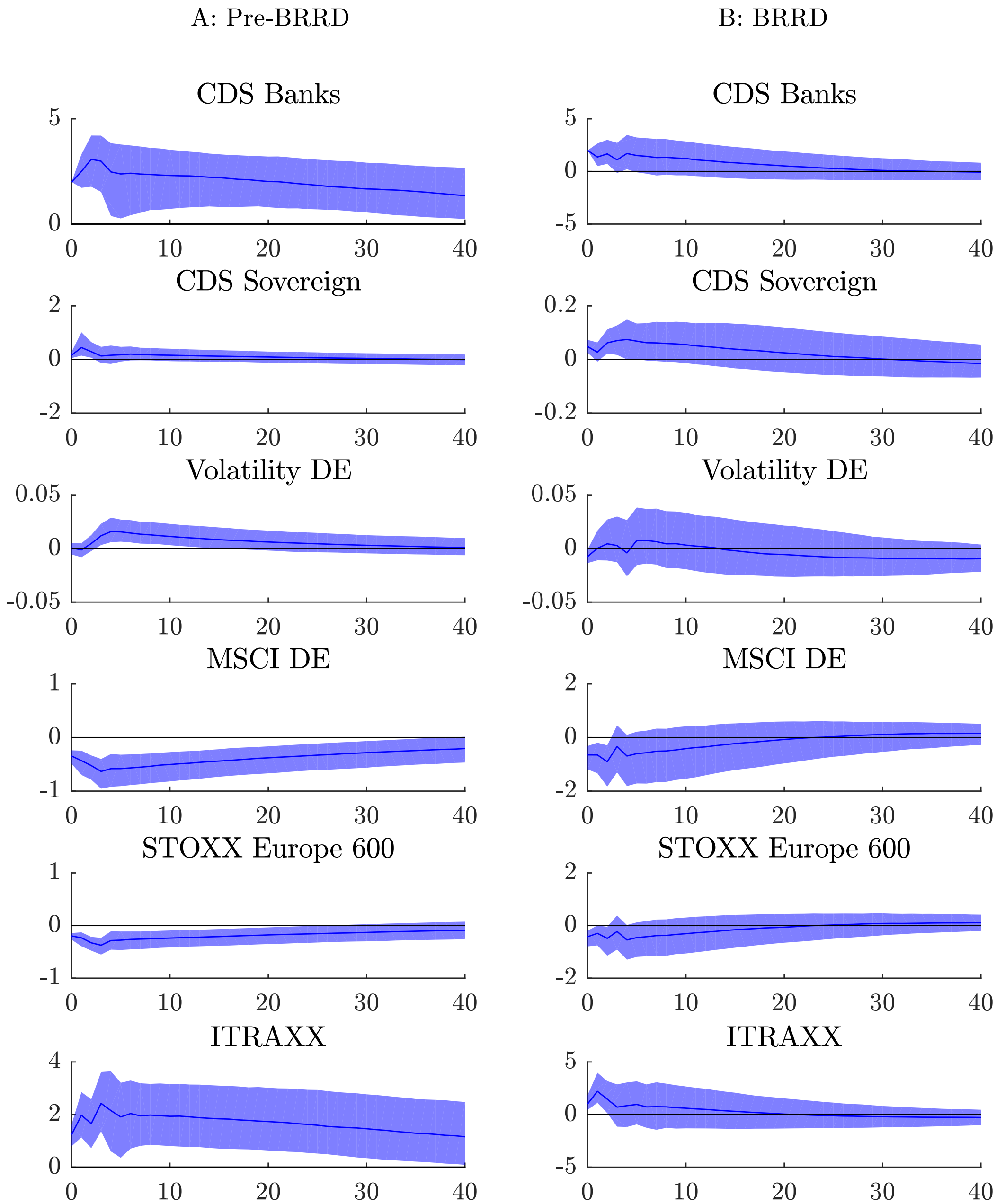
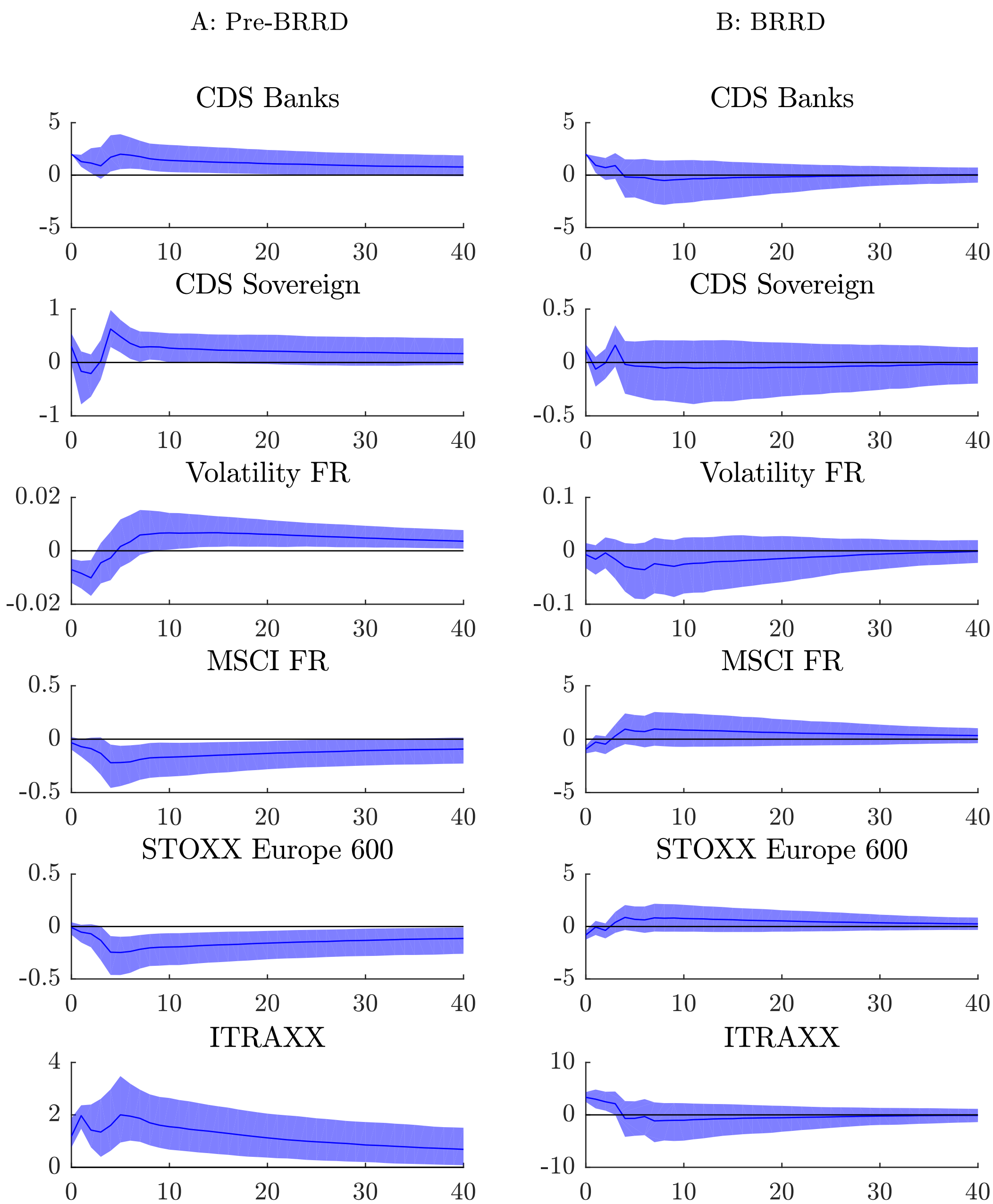


Figure 11: Impulse response function of an exogenous bank shock on the CDS spreads of French banks and sovereign bonds, the volatility of the MSCI France and the index itself, and the Itraxx, identified using the narrative approach. Panel A shows the impulse response function is estimated from the pre-BRRD period, whilst Panel B shows the impulse responses of the post-BRRD regime.



Appendix

A DCC-MIDAS quasi-maximum likelihood estimator

The likelihood is maximized of the assumption that the demeaned returns μ_t are normally distributed with covariance matrix H_t. This yields a log likelihood estimator which can be maximized over the parameters of the model. First, the joint distribution of the demeaned log returns of the CDS spreads is

$$f(\mu_t) = \prod_{t=1}^{T} \frac{1}{\sqrt{2\pi}} |H_t|^{-1/2} \exp\left(-\frac{1}{2}\,\mu_t' H_t^{-1} \mu_t\right) \tag{16}$$

assuming that the standardized residuals (ν_t) are mean zero and unit variance normally distributed.

$$\begin{aligned} L &= -\frac{T}{2}\log(2\pi) - \frac{1}{2}\sum_{t=1}^{T}\log|H_t| - \frac{1}{2}\sum_{t=1}^{T}\mu_t' H_t^{-1}\mu_t \qquad \text{with } H_t = D_t R_t D_t & (17)\\ &= -\frac{T}{2}\log(2\pi) - \frac{1}{2}\sum_{t=1}^{T}\log|D_t R_t D_t| - \frac{1}{2}\sum_{t=1}^{T}\mu_t' D_t^{-1} R_t^{-1} D_t^{-1}\mu_t & (18)\\ &= -\frac{1}{2}\sum_{t=1}^{T}\left(T\log(2\pi) + 2\log|D_t| + \mu_t' D_t^{-1} D_t^{-1}\mu_t - \nu_t'\nu_t + \log|R_t| + \nu_t' R_t^{-1}\nu_t\right) & (19) \end{aligned}$$

Estimating the log likelihood in Equation 19 is difficult, certainly instances where the covariance matrix is large. Therefore the DCC-model can be estimated in two stages, as in Engle (2002). The log likelihood function can be decomposed into a part that is driven by the volatility of the individual time series, and into a part that is driven by the correlation between the time series. Each part is optimised individually. The volatility (L_v) and correlation (L_c) terms are expressed as

$$\begin{aligned} L_v &= -\frac{1}{2}\sum_{t=1}^{T}\left(T\log(2\pi) + \log|D_t|^2 + \mu_t' D_t^{-2}\mu_t\right) & (20)\\ L_c &= -\frac{1}{2}\sum_{t=1}^{T}\left(-\nu_t'\nu_t + \log|R_t| + \nu_t' R_t^{-1}\nu_t\right) & (21) \end{aligned}$$

The estimation procedure, proposed by Engle (2002), estimates the correlation in two stages. First, the conditional volatility of each time series individually is optimised, through maximizing the log likelihood of the volatility. This is equivalent to maximizing the likelihood in Equation 20. The estimated volatilities are used in the second stage to obtain the standardized residuals (ν_t) after which the likelihood of the correlation term is maximized. The result of the maximization procedure is the dynamic conditional correlation matrix which is most likely, given that the standardized residuals have a zero mean and unit variance normal distribution.