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Roberto De Santis, Michael Stein

Correlation changes between the risk-free rate and sovereign yields of euro area countries

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Abstract

We study correlations between the risk-free rate and sovereign yields of ten euro area countries using smooth transition conditional correlation GARCH (STCC-GARCH) specifications, controlling for credit risk in mean and variance equations and conditioning non-linearly to liquidity risk. Correlations are state-dependent and heterogeneous across jurisdictions. Using panel vector autoregression models, we identify the macro factors influencing the correlations: interbank credit risk, the Greek crisis, and break-up risk. We show that the European Central Bank's asset purchase programmes helped restore the pass-through relationship. We also make a methodological contribution by estimating all STCC-GARCH parameters at once and introducing an STCC-GARCHX.

Keywords: Monetary Policy, Government Bonds, Smooth Transition Models, Euro Area
JEL Codes: G12, G15

Non-technical Summary

During 2011-2012, the euro area's financial system was strongly affected by a surge in uncertainty associated with the risk of a break-up of the euro area, paired with an intensification of investors' concern regarding the sustainability of sovereign fiscal positions. The most critical period was in July 2012, when the long-term sovereign credit spreads of Italian and Spanish sovereign bonds vis-à-vis the risk-free rate, measured by the overnight indexed swap (OIS) with the same maturity, reached record highs (about 500-600 basis points).

The same spreads were about 200 basis points lower in March 2012 and these dynamics were associated with the risk of euro area break-up. Therefore, the Eurosystem announced the possibility to engage in outright monetary transactions (OMTs) in secondary sovereign bond markets. By mid-September 2012, the Italian and Spanish sovereign credit spreads fell by about 160-200 basis points compared to its peak in July and, subsequently, saw a steady decline. This anecdotal evidence suggests that sovereign bond markets can be driven by fear and panic and the self-fulfilling nature of these developments can lead a country into default.

The identification of a correlation breakdown between the risk free rate and the non-credit risk component of sovereign yields is paramount for policy-makers, given the benchmark role played by sovereign yields in the transmission mechanism of monetary policy. This question is even more important in a monetary union context, given that the single monetary policy instruments may result in being ineffective if financial markets are fragmented among member states.

Some might argue that during the sovereign debt crisis perceptions of expected macro-fundamentals changed and therefore the investors' reactions reflected in sovereign yields were fully rational. In other words, some would disagree with the monetary policy decision, as the dynamics of sovereign yields reflected changes in expectations about credit risk, which monetary policy should not aim at influencing.

We provide evidence that credit risk played an important role in both the conditional correlations and the volatility in most of the euro area sovereign debt markets, particularly those most affected by the sovereign debt crisis. However, even adjusting for credit risk, the conditional correlations between the sovereign yields adjusted for credit risk and the risk-free rate did not remain high and stable. On the contrary, they declined sharply during the hikes of the sovereign debt crisis in many euro area member states, suggesting that the transmission of the monetary policy was highly impaired. The breakdown in the conditional correlations is heterogeneous among country groupings, with the vulnerable countries - Ireland, Italy, Portugal and Spain - being affected most.

This result is obtained estimating the conditional correlations between the five-year risk-free rate and the sovereign yields adjusted for credit risk in the mean and variance equations,

conditional upon the degree of financial stress, which typically affects liquidity risk and uncertainty, of the largest ten euro area countries.

We also uncover the reasons behind the correlation breakdown. We report evidence that the counterparty risk in the banking system played an important role since the financial crisis started in August 2007. In addition, conditional correlations were affected by developments in Greek sovereign risk in 2010 and 2011 and by the perceived euro area's break-up risk from the end of 2011 onwards. The change in the trend began with the launch of outright monetary transactions by the ECB in the summer of 2012, which abated the euro break-up risk. Moreover, the financial market situation improved further from the beginning of 2014 onwards, when financial markets priced in a higher probability of additional expansionary monetary policy measures in the euro area. In particular, the econometric results suggest that the ECB public asset purchase program has been an important instrument to improve the conditional correlations in 2014 and 2015, particularly for the more vulnerable countries.

1. Introduction

During 2011-2012, the euro area's financial system was strongly affected by a surge in uncertainty associated with the risk of a break-up of the euro area, paired with an intensification of investors' concern regarding the sustainability of sovereign fiscal positions. The most critical period was in July 2012, when the long-term sovereign credit spreads of Italian and Spanish sovereign bonds vis-à-vis the risk-free rate, measured by the overnight indexed swap (OIS) with the same maturity, reached record highs (about 500-600 basis points).

The same spreads were about 200 basis points lower in March 2012 (see Figure 1) and these dynamics were associated with the risk of euro area break-up. Therefore, the Eurosystem announced the possibility to engage in outright monetary transactions (OMTs) in secondary sovereign bond markets. By mid-September 2012, the Italian and Spanish sovereign credit spreads fell by about 160-200 basis points compared to its peak in July and, subsequently, saw a steady decline. This anecdotal evidence suggests that sovereign bond markets can be driven by fear and panic (De Grauwe and Ji (2013)) and the self-fulfilling nature of these developments can lead a country into default.

[Insert Figure 1, here]

Sharp increases in sovereign spreads can lead to a massive misallocation of resources because sovereign yields are generally used as benchmark reference rates to price key interest rates, such as lending rates to households and corporations or corporate bond prices, and can impair the transmission of the Eurosystem's single monetary policy. Moreover, flight-to-safety and flight-to-liquidity can further fragment the sovereign bond market in a monetary union with countries considered to be safer and more liquid benefiting as the interest rates on their sovereign bonds decline.[3]

De Santis and Stein (2015) have shown that conditional correlations between sovereign yields and the risk-free rate are state-dependent using Smooth Transition Conditional Correlation GARCH (STCC-GARCH) methods, a class of multivariate GARCH models introduced by Silvennoinen and Teräsvirta (2005, 2015).[4] However, the dynamics of sovereign yields may reflect changes in expectations about credit risk driven by, for example, a nonlinear relationship with macro fundamentals (Boumparis et al. (2015)).

If the correlation breakdown between sovereign yields and the risk-free rate is due to

[3]Large sovereign spreads may also create incentives - particularly for banks - to initiate carry-trade activities, as investments in higher yielding sovereign assets can be financed with short-term debt, potentially leading to excess risk-taking and strengthening the bank and sovereign nexus (Acharya and Steffen (2015)).

[4]STCC-GARCH models have been used to study the correlation between stocks (Aslanidis et al. (2009), Silvennoinen and Teräsvirta (2005, 2009, 2015) and Chelley-Steeley et al. (2013)), stocks and bonds (Stein et al. (2013)), stocks and exchange rates (Lee et al. (2011)), and other asset classes (Silvennoinen and Thorp (2013) and Koch (2011)).

time-varying credit risk, then it is tolerable and desirable because an increase in credit risk premia can force policy-makers to undertake structural reforms, which otherwise would not be taken; if not, it is unwelcome. Hence, some would disagree with the monetary policies addressing the sovereign debt crisis, as the dynamics of sovereign yields reflected changes in expectations about credit risk, which monetary policy should not aim at influencing.

We provide evidence that credit risk played an important role in both the conditional correlations and the volatility in most of the euro area sovereign debt markets, particularly those most affected by the sovereign debt crisis. However, even adjusting for credit risk, the conditional correlations between the risk-free rate and (credit risk-adjusted) sovereign yields did not remain high and stable. On the contrary, they declined sharply during the hikes of the sovereign debt crisis in many euro area member states, suggesting that the transmission of the monetary policy was highly impaired. The breakdown in the conditional correlations is heterogeneous among country groupings, with the vulnerable countries - Ireland, Italy, Portugal and Spain - being affected most.

We expand the analysis relative to the existing literature along three dimensions. First, we control for the jurisdiction's credit risk and identify factors potentially affecting the dynamics of such correlations. Second, we introduce two methodological novelties: we maximize the likelihood function using a single step rather than by conditional maximum likelihood and, thus, estimate all parameters jointly, and control for credit risk in not only the mean but also the volatility equations, thereby introducing the STCC-GARCHX specification. Third, we cover a relatively large number of euro area countries.

Credit risk can be proxied with the credit default swap (CDS) spread (Duffie (1999), Pan and Singleton (2008), Beber et al. (2009), Longstaff et al. (2011)). However, the CDS market is also influenced by flight to liquidity and, more generally, aggregate market uncertainty (Ang and Longstaff (2013)). Therefore, as a proxy of credit risk, we consider the CDS component that is orthogonal to (i) the bid-ask spread associated with the sovereign yield, (ii) the implied volatility of option prices on the EURO STOXX 50 index (VSTOXX) and (iii) the spread between the Kreditanstalt für Wiederaufbau (KfW) bond and the German Bund. The latter is used as a proxy for flight to liquidity, because both bonds are guaranteed by the German government and, therefore, carry the same default risk (De Santis (2014), Monfort and Renne (2014), Ejsing et al. (2015)).[5]

All empirical analyses are conducted using variables in first-differences with the exception of the signal that enters the transition function of the STCC-GARCHX model. In sum,

[5] Any difference between agency and government bond yields should reflect international investors' preference for assets with the lowest liquidity risk (Longstaff (2004)).

sovereign yields are regressed against the component of CDS spreads, which are orthogonal to measures of liquidity and aggregate risks, and the resulting correlation with the risk-free rate is non-linearly conditioned to liquidity risk and/or aggregate risk. The non-linear relationship is statistically tested and can be motivated by the possibility that correlations may depend upon the degree of financial stress, which typically affects liquidity and aggregate risks.

Our findings support the view of time-varying, state-dependent, and heterogeneous correlations between the risk-free rate and (credit risk-adjusted) sovereign yields, particularly after the collapse of Lehman Brothers and during spring 2011 and autumn 2012, when the sovereign debt crisis spread to Italy and Spain.

To identify the key macro factors underlying the pass-through breakdown in the sovereign debt markets, we use impulse responses based on panel vector autoregression (PVAR) models, which includes the time-varying conditional correlations, a global financial factor (US VIX), and three risk measures characterizing the financial crisis in the euro area. The results suggest that the estimated changes in the correlations are driven by (i) the counterparty risk in the banking system over the entire sample period, (ii) the Greek sovereign risk, particularly between 2010 and 2011, and (iii) the redenomination risk from the end of 2011.

Moreover, the results indicate that the OMT and ECB' Public Sector Asset Purchase Programme (PSPP) helped improve the conditional correlations between the risk-free rate and sovereign yields, thereby partly re-establishing a cardinal monetary policy transmission mechanism. In particular, the econometric results suggest that the ECB PSPP has been an important instrument to improve the conditional correlations in 2014 and 2015, particularly for the more vulnerable countries.

Our study contributes to various strands of the literature. First, and methodologically, we enhance the smooth transition method of Silvennoinen and Teräsvirta (2015) by maximizing the likelihood function in one step rather than by conditional maximum likelihood and controlling for credit risk in the GARCH equation. Second, a vast literature study the determinants of sovereign yields during the sovereign debt crisis. Countries with poor fundamentals saw their yields dramatically increase,[6] a pattern that has been identified as "wake-up call" contagion. In this study, we control for credit risk and condition non-linearly the correlations to the "wake-up call" variables, such as liquidity premia and uncertainty measures. Third, the literature investigates the impact of unconventional monetary policy

[6]See for example Arghyrou and Kontonikas (2012), Bernoth and Erdogan (2012), Bernoth et al. (2012), Beirne and Fratzscher (2013), Giordano et al. (2013), D'Agostino and Ehrmann (2014), De Santis (2014) and Dergiades et al. (2014))

on asset prices.[7] We investigate the impact of ECB policy programs on the correlations between (credit risk-adjusted) sovereign yields of several euro area countries and the euro area risk-free rate.

This study is structured as follows. Section 2 summarizes the methods. Section 3 describes the data and the indicators. Section 4 and section 5 discuss and interpret the results. Section 6 provides robustness checks. Section 7 concludes.

2. Standard methods and innovations

We estimate the conditional correlations of sovereign bond markets and the risk-free rate using a STCC-GARCH model that we augment with an exogenous regressor in the variance equation (STCC-GARCHX model). The results based on the standard dynamic conditional correlation (DCC)-GARCH model introduced by Engle (2002) are used as a benchmark. Both DCC-GARCH and STCC-GARCH models require the definition of the conditional mean equation, the conditional variance and the conditional correlation matrices. The major differences between the two models are the specification of the correlation matrix and the estimation procedure. The conditional correlation matrix in the DCC-GARCH models is defined with a dynamic structure, whereas the conditional correlations in the STCC-GARCH models change smoothly between two "extreme" correlation matrices. The DCC-GARCH models are estimated using a two-step approach with conditional correlations computed after estimating the univariate conditional variances. The STCC-GARCH models are typically estimated using conditional maximum likelihood with an iterative procedure for the parameters governing conditional variances, conditional correlations and transition parameters (Silvennoinen and Teräsvirta (2015)). This iterative approach for the three sets of parameters ensures convergence.

In this study, we introduce two methodological novelties: the inclusion of exogenous factors in the variance equation (STCC-GARCHX model), and the one-step joint estimation of all GARCH, correlation and transition parameters, rather than using conditional maximum likelihood. In the next sub-sections, we detail the setup for both the mean and variance equations.

[7] Studies focusing on US government bond yields are Doh (2010), Gagnon et al. (2011), Krishnamurthy and Vissing-Jorgensen (2011), Meaning and Zhu (2011), D'Amico et al. (2012), D'Amico and King (2013) and Li and Wei (2013). Studies focusing on UK government bond yields are Meier (2009), Joyce et al. (2011), Joyce and Tong (2012), Meaning and Zhu (2011), Breedon et al. (2012), Christensen and Rudebusch (2012) and McLaren et al. (2014). Studies focusing on euro area government bond yields are Altavilla et al. (2015); Szczerbowicz (2015) and De Santis (2016).

2.1. Mean equation

Theory can help us select the indicators signaling a change in correlations between the risk-free rate and sovereign yields. Typically, the nominal sovereign long-term rate with maturity L in country c, $r^L_{c,t}$ can be disaggregated in the following main components:

$$r^L_{c,t} = (i^{MP}_t + E^{MP}_t(i_t) + ... + E^{MP}_{t+L-1}(i_t))/L + cp^L_{c,t} + lp^L_{c,t} + gp_t \tag{1}$$

where the first component in parentheses is the average of the expected monetary policy rates, $(i^{MP}_t + E^{MP}_t(i_t) + ... + E^{MP}_{t+L-1}(i_t))/L$ common to all euro area countries; the second component is the credit risk premium for sovereigns in country c, $cp^L_{c,t}$; the third component is the liquidity premium for sovereigns in country c, $lp^L_{c,t}$; and the fourth component is aggregate risk premium, gp_t.

In general, sovereign yields depend only on credit risk and risk-free rate: $lp^L_{c,t} = 0$ and $gp_t = 0$. However, $lp^L_{c,t}$ and gp_t are time-varying and their dynamics may affect the correlation between the risk-free rate and $r^L_{c,t} - cp^L_{c,t}$, particularly during financial stress.

We model the risk-free rate with a constant and an autoregressive term:

$$\triangle r_{i,t} = \beta_{i0} + \beta_{i1} \triangle r_{i,t-1} + \varepsilon_{i,t}, \tag{2}$$

Instead the mean equation for sovereign yields controls for credit risk. The literature suggests the use of CDS spreads; however, they are not exogenous to liquidity risk and aggregate market uncertainty (Ang and Longstaff (2013)). Thus, to disentangle the pure credit risk component, CDS spreads are first regressed against measures of liquidity and aggregate risks in an auxiliary regression and, then, the residuals are used in the mean equation for changes in the sovereign yields:

$$\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1} \triangle BA_{i,t} + \phi_{i2} \triangle KfW_t + \phi_{i3} \triangle VSTOXX_t + \xi_{i,t}, \tag{3}$$

$$\triangle r_{i,t} = \beta_{i0} + \beta_{i1} \triangle r_{i,t-1} + \beta_{i2} \hat{\xi}_{i,t} + \varepsilon_{i,t}, \tag{4}$$

where $BA_{i,t}$ is the bid-ask spread of the sovereign yield i at time t, KfW_t is the KfW-Bund spread and $VSTOXX_t$ is the euro area's implied volatility. $\beta_{i1} BA_{i,t}$, $\beta_{i2} KfW_t$ and $\beta_{i3} VSTOXX_t$ measure sovereign-specific liquidity premia and expected changes in the euro area macroeconomic outlook as perceived by the market. Therefore, the residuals $\hat{\xi}_{i,t}$ from equation (3) represent the pure sovereign-specific credit risk and $\hat{\varepsilon}_{i,t}$ obtained from equation (4) captures the changes in sovereign yields orthogonal to credit risk. This implies that important changes in the correlations between $\varepsilon_{i,t}$ and the OIS rate might require policy attention. It is important to stress that the OLS estimates provide the correct value for β_{i2} because we use the residuals from a supplementary regression and not the predictors (see Model 4 in Pagan (1984)).

Differences in the specification of the mean equations in multivariate GARCH studies are often seen playing a marginal role.[8] In the present, controlling for credit risk in the mean equations is important, particularly for countries heavily affected during the sovereign debt crisis. Introducing the CDS spreads in the mean equation is one answer to the view that developments in sovereign yields were driven by investors' re-appraisal of credit risk.

2.2. Variance equations and correlation matrices

We employ bivariate GARCH models for the variances in each sovereign yield i and the OIS rate. In both the DCC-GARCH and STCC-GARCH(X) models, the univariate conditional variances D_t are coupled with a conditional correlation matrix R_t to obtain the conditional variance-covariance matrix H_t:

$$H_t = D_t R_t D_t, \; with \; D_t = (h_{i,t}^{1/2}, h_{OIS,t}^{1/2}). \tag{5}$$

D_t is a 2x2 matrix with $h_{i,t}$ denoting the conditional variance of the sovereign yield i and $h_{OIS,t}$ indicating the conditional variance of the OIS rate.

The variance of the DCC-GARCH model is defined as a GARCH(p,q) process

$$h_{i,t} = a_{i0} + \sum_{j=1}^{q} a_{ij}\varepsilon_{i,t-j}^2 + \sum_{l=1}^{p} b_{il} h_{i,t-l}, \; \varepsilon_{i,t} = h_{i,t}^{1/2} z_{i,t} \; and \; \varepsilon_{i,t} \mid \psi_{i,t-1} \sim N(0, h_{i,t}), \tag{6}$$

where $\varepsilon_{i,t}$ are the errors in the mean equations and $z_{i,t}$ is a standard normal variate, as the standardized errors are assumed to be normally distributed gven the information set $\psi_{i,t-1}$. The variance equation for the OIS rate is identical to equation (6).

However, the dynamics of such correlations may depend on the degree of financial stress, which typically characterizes $lp_{c,t}^L$ and gp_t. In other words, $lp_{c,t}^L$ and gp_t can be related non-linearly to standard factors s_t used in the literature, such as the bid-ask spreads, the KfW-Bund spread and stock market-implied volatility. The STCC-GARCH model allows for a non-linear relationship in the correlations between the returns of two assets.

In particular, Silvennoinen and Teräsvirta (2015) define a logistic transition function G bounded between zero and one, governed by the difference between the transition variable

[8]Chelley-Steeley et al. (2013), for example, add the Fama-French factors to the mean equation of an STCC-GARCH model; however, they do not find any significant difference in the results.

s_t and the endogenously determined threshold c,[9] with two extreme states of correlation represented by the correlation matrices R_1 and R_2:

$$R_t = (1 - G_t) \cdot R_1 + G_t \cdot R_2, \tag{7}$$

$$G_t(\gamma, c, s_t) = (1 + \exp\{-\gamma(s_t - c)\})^{-1}, \gamma > 0. \tag{8}$$

Time-varying conditional correlations change smoothly between the extreme states. How quickly transitions occur depends on γ, the speed of transition parameter.[10] Conditional correlations become constant if $s_t = c$ for all t, $R_1 = R_2$ or $\gamma = 0$. Otherwise, they range between the two extreme states with the non-linear specification being preferred if γ is statistically different from zero.

Moreover, we consider the possibility that both the conditional mean and variance may be affected by credit risk. Therefore, we enhance the variance equation with an exogenous factor $\hat{\xi}^2_{i,t}$ for the first time in the literature (STCC-GARCHX model):

$$h_{i,t} = a_{i0} + \sum_{j=1}^{q} a_{ij} \varepsilon^2_{i,t-j} + \sum_{l=1}^{p} b_{il} h_{i,t-l} + \sum_{k=1}^{r} \delta_{ik} \hat{\xi}^2_{i,t-k}. \tag{9}$$

The variance equation of the risk-free rate does not depend on credit risk and, therefore, the OIS rate is always modeled with the standard GARCH(p,q) specification.

We assume that $p = q = r = 1$. Therefore, the parameter set to be estimated is: $\theta = \{a_{i0}, a_{i1}, b_{i1}, \delta_{i1}, a_{OIS0}, a_{OIS1}, b_{OIS1}, \gamma, c, \rho_1, \rho_2\}$, where ρ_1 and ρ_2 are the off-diagonal values of the correlation matrices R_1 and R_2 in our bivariate case.

2.3. Joint parameter estimation

Silvennoinen and Teräsvirta (2015) suggested that STCC-GARCH models can be estimated with conditional maximum likelihood, separating three sets of parameters: univariate GARCH, correlation and transition parameters. The suggested iterative approach has the advantage of allowing feedback effects between volatility, correlation and transition parameters, which ensures convergence and, in general, smaller standard errors. This estimation procedure is termed “simultaneous” to indicate the feedback loop between the parameter subsets.

In this study, we jointly estimate all parameters in a single step with the advantages of: (i) never encountering problems when calculating the standard errors of the parameter

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[9] Berben and Jansen (2005) independently developed a time-varying STCC (TV-STCC)-GARCH in the same year when the STCC-GARCH model was introduced by Silvennoinen and Teräsvirta (2005), with the transition variable s_t being a time trend in the Berben and Jansen (2005) specification.

[10] To eliminate scale effects, we follow the common practice of standardizing the transition parameter γ with the standard deviation of the transition variable.

estimates, with all matrices being positive semi-definite, (ii) not having to fix the speed of the transition parameter γ for the standard error calculations as in previous studies[11], (iii) not needing to change a simple GARCH(1,1) to a GJR-GARCH specification as in Silvennoinen and Teräsvirta (2015), and (iv) not obtaining integrated GARCH results as those in De Santis and Stein (2015). The joint estimation also eradicates the sensitivity to initial parameter values and the danger of local minima.

The joint parameter estimation is conducted as follows: first, we calculate rolling correlations over the full sample to identify the best initial guesses of extreme correlations, and run the univariate GARCH estimations to identify the best starting values of the variance equations; second, we use a large set of parameter combinations for the transition parameters c and γ to find the initial values that provide the highest values for the likelihood function.[12] Using this set of initial values for all parameters, we jointly maximize the likelihood to obtain the parameter set $\theta = \{a_{i0}, a_{i1}, b_{i1}, \delta_{i1}, a_{OIS0}, a_{OIS1}, b_{OIS1}, \gamma, c, \rho_1, \rho_2\}$:

$$l_t(\theta) = -log(2\pi) - \frac{1}{2}(log\,(h_{i,t}) + log\,(h_{OIS,t})) - \frac{1}{2}log \mid R_t \mid - \frac{1}{2} z_t' R_{t-1}^{-1} z_t. \tag{10}$$

3. Data

The sovereign bond yield that we use has a five-year maturity for two main reasons: first, the aggregate demand is typically affected by long-term interest rates and, therefore, the correlation between long-term sovereign yields and the risk-free rate is useful to monitor; and second, the market for CDS spreads used to measure the price of credit risk is more liquid at five-year maturity.

As a proxy for the risk-free rate, we employ the euro area OIS rate with the same maturity. According to the ECB (2014), the OIS rate has a very low perceived credit risk and, during the crisis period, it was much less sensitive to flight-to-liquidity flows than the euro area AAA yield curve and German Bund yields.[13]

[11] Silvennoinen and Teräsvirta (2005) fix the value at the general upper limit of 100 for standardized γ above which the likelihood function is merely insensitive to changes in the parameter, with this numerical detail not being the focus in Silvennoinen and Teräsvirta (2015). De Santis and Stein (2015) fix it at the estimated value for the error calculation.

[12] In several tests with both actual and simulated data, the results are consistent for all estimation runs if we use more than 10,000 combinations in the grid search for γ and c of the transition function. In this study, we use 10 million combinations for the sake of robustness and set the upper and lower bounds in the grid to $\gamma \in [0, 100]$ and $c \in [min(s), max(s)]$ with s_t being the signal.

[13] The euro over-night index average (EONIA) swap index or OIS rate is a fixed-floating rate interest rate swap where the floating rate is indexed to the EONIA rate at which banks provide loans to each other for a day's duration. Banks may qualify for the EONIA Swap Index Panel if they: 1) are active players in the euro derivative markets, in the euro area or worldwide, and have the ability to transact large volumes in EONIA swaps, even under turbulent market conditions, 2) have a high credit rating, and exhibit high standards of

The price of liquidity risk is proxied by the sovereign bid-ask spread and the KfW-Bund spread, and that of aggregate risk is proxied by the VSTOXX, which is highly correlated with the VIX (see Figure 2). These measures are all well-known in the literature except for the KfW-Bund spread, which has only recently drawn attention. The construction of such a spread is documented in De Santis (2014), Monfort and Renne (2014)) and Ejsing et al. (2015). The key point is that the KfW bond and the German Bund are characterized by the same credit risk, with both being guaranteed by the German state. Any difference between agency and government bond yields should thus reflect liquidity premia, as international investors prefer to hold very liquid assets, such as the Bund, which can be easily dismissed in large quantities if required. This positive spread is also due to the fact that the portfolio composition of mutual funds with a low-risk profile includes the German Bund but not the KfW bond.

[Insert Figure 2, here]

The sample period under investigation covers the ten-year period April 2005 - March 2015; the frequency of the sample is daily; and the country coverage includes Austria, Belgium, Finland, France, Germany, Ireland, Italy, the Netherlands, Portugal and Spain. A summary of the descriptive statistics is presented in Table 1.

[Insert Table 1, here]

The various stages of the sovereign debt crisis in the euro area are clearly described by the developments of the sovereign yields and CDS spreads, which are also obtained from Bloomberg (see Figure 1). All the benchmark sovereign yields and OIS rates tightly co-moved up to September 14, 2008. With the intensification of the financial crisis in September 2008, following the collapse of Lehman Brothers, the sovereign spreads of countries with a weak fiscal space increased. CDS and KfW-Bund spreads followed similar developments (see Figure 2).

The sovereign debt crisis spread alarmingly to Italy and Spain in 2011 and 2012. Only after the "whatever it takes speech" by Mario Draghi, the sovereign credit and bid-ask spreads as well as the KfW-Bund spread began a steady decline. The VIX and VSTOXX also reverted their trend, although they were already fluctuating much below the levels previously recorded.

ethical behavior and enjoy an excellent reputation, and 3) disclose all relevant information requested by the Steering Committee. At present, 25 prime banks constitute the EONIA Swap Index Panel. These selected banks are obliged to quote the EONIA swap index for the complete range of maturities in a timely manner on every business day and with an accuracy of three decimal places. The independent Steering Committee, which consists of 10 members, closely monitors all market developments and ensures, by reviewing panel banks' contributions on a regular basis, strict compliance with the code of conduct. It also has the right to request information and remove or appoint panel banks.

To extract a "pure" credit risk component from CDS spreads, we regress each CDS spread on the bid-ask spread, KfW-Bund spread and euro area VSTOXX, as defined in equation (3). These variables all have the correct positive sign and can explain around 10-15% of the variation in the CDS spreads (see Table 2). De Santis (2014) identifies the KfW-Bund spread as a euro area common risk factor, which captures the portfolio shift because of a higher appetite for the German Bund, thereby affecting all the euro area sovereign yields. The results in Table 2 suggest that the CDS market has also been highly influenced by flight-to-liquidity considerations.

[Insert Table 2, here]

The CDS spreads and the orthogonalized measure of credit risk, obtained by cumulating the residuals of equation (3), are depicted in Figure 1. In the cases of Italy and Spain, during the hike of the crises, the CDS spreads amounted to about 600 basis points, while the adjusted measures reached 400 basis points. The 200 basis points difference can be attributed to the liquidity premia and aggregate uncertainty. In low-yielding countries, such as Germany, Finland and the Netherlands, this difference is smaller - about 20-30 basis points - although in relative terms it still amounts to about 20-30% of the traded CDS spreads.

4. Empirical results

4.1. Mean equation

The OLS estimates of the mean equations (2) and (4) are summarized in Table 3. The autoregressive component of the changes in sovereign yields is statistically significant in nine out of ten countries in a simple AR(1) model, but the explanatory power does not exceed 5% for any assets.

As expected, a large proportion of the variance in the sovereign yields can be explained by credit risk, as the adjusted R^2 for Spain, Portugal, Italy and Ireland substantially increases to 25-45% when estimating equation (4). The coefficient for credit risk is positive, except in the case of Germany. The eyeballing of Figure 1 indicates a clearly negative relationship, which is particularly evident for Germany during the hike in the sovereign debt crisis, because of to flight-to-safety and flight-to-liquidity motives. The coefficient for Germany is negative, but the adjusted R^2 is 1%, which means that the dynamics of German sovereign spreads are not affected by such a variable.

In sum, the credit risk in the Dutch, Finnish and German sovereign bond markets plays a negligible role and the sovereign yields of these countries are treated by investors as merely risk-free rates. This implies that their correlations with the risk-free rate should be relatively less affected during the sovereign debt crisis.[14]

[Insert Table 3, here]

4.2. Conditional correlations and transition parameters

When using the sovereign bid-ask spreads or the stock market-implied volatilities as transition variables, the conditional correlations change rather quickly. Frequent switches are typical of standard regime-switching models, and these are not helpful in making key decisions. In particular, when using stock market-implied volatilities, the dynamic correlations do not provide a clearly consistent signal of regime changes that is in line with market narratives and expectations, and this reflects the volatile nature of the signal. Therefore, to save space, the results reported here refer only to those models that use the KfW-Bund spread as a signal of a potential regime change.[15]

First, we compare the results obtained by estimating the STCC-GARCHX model using three estimation methods: the two-step, iterative and joint estimation approaches. The two-step approach is similar to the approach used to estimate the DCC-GARCH models, where a feedback loop between the subset of the parameters is not existent. The iterative approach is common in the literature. The joint estimation approach proposed in this study maximizes the likelihood function after an intensive grid search of initial parameters.

The results presented in Table 4 are summarized for France, Germany and Italy as case studies reflecting groups of countries that were affected differently during the sovereign debt crisis. Most of the estimated coefficients are similar across the various approaches, but at times, these differences can be large. Most importantly, there is a clear improvement in the likelihood value from the two-step and the iterative approaches to the joint estimation. The same results also apply to other countries. Therefore, we can safely conclude that our suggested method provides an improvement relative to the existing literature.

[Insert Table 4, here]

The estimates for DCC-GARCH and STCC-GARCHX are summarized in Tables 5 and 6. The parameters are strongly statistically significant, pointing to the usefulness of comparing the linear and the non-linear model specifications. As for the STCC-GARCHX model, the estimated speed of transition, γ, is relatively moderate and statistically significant in six out of ten cases, which justifies the use of the chosen model particularly for Belgium, France, Germany, Italy, Portugal and Spain.

[14]The constant in the mean equation is significant for the Finnish regression in the two models because the sample period for Finland begins only at the end of 2008; thus the constant captures the declining trend in the Finnish sovereign yield.

[15]All other calculations and results not presented here can be provided by the authors upon request.

The parameter estimates are similar for some country groups: (i) Italy, Spain and Portugal; (ii) Austria and Belgium; and (iii) Finland, Netherlands and France; and separately (iv) Ireland and Germany. This implies that the conditional correlations are expected to be highly heterogeneous across jurisdictions.

Finally, credit risk in the GARCH equation is statistically significant in many cases, such as Austria, Belgium, Ireland, Italy, the Netherlands and Spain, providing support for the use of the STCC-GARCHX specification, especially for countries most affected by credit risk considerations.

The analysis suggests that the threshold to a crisis regime is reached when the five-year KfW-Bund spread is above around 10 basis points for Finland, Germany and Portugal; 20 basis points for Italy and Spain; 30 basis points for Ireland; and 40 basis points for Belgium and France. The threshold for Austria and the Netherlands is not statistically significant (see Table 6). Therefore, as a rule of thumb, many sovereign debt markets move to a crisis regime if the five-year KfW-Bund spread is above 20-40 basis points, and this was the case during 2008-2013.

[Insert Tables 5-6, here]

In the non-crisis regime, the "extreme" correlations are above 90% in all the countries except Ireland (60%) and Finland (80%). In the crisis regime, the correlations are below 50% for all countries except Germany (80%).

Figure 3 plots the conditional correlations obtained using the DCC-GARCH and STCC-GARCHX specifications. The comparison of the conditional correlations suggests that those obtained with the STCC-GARCHX models are less volatile and never negative, a feature that fits well with the theoretical argument that the risk-free rate is always positively correlated with sovereign yields. This is not the case for the DCC-GARCH models, which show negative conditional correlations between 2010 and 2012 for most of the vulnerable countries, including Ireland, Italy, Portugal and Spain.

In general, despite controlling for credit risk, we can safely say that the pass-through from the risk-free rate to sovereign yields of many euro area countries declined during the 2008-2014 crisis period relative to the previous periods, particularly for sovereigns that suffered the most, such as Ireland, Italy, Portugal and Spain.

[Insert Figure 3, here]

The conditional correlations show similar patterns for three country-pairs: Italy and Spain, Austria and Belgium, and France and the Netherlands. The conditional correlation between the Bund and the OIS rate ranged from 80% to 90% during the entire sample period (2005-2015), regardless of the developments in the signal. This suggests that the Bund yield behaves like a risk-free rate. The correlations of Finland and the Netherlands are slightly more volatile, reaching a trough at 50% in 2009 and 2012 during the hikes of the financial crisis.

In sum, the correlations between the risk-free rate and the (credit risk-adjusted) sovereign yields are state-dependent, contingent on the degree of financial stress and flight-to-liquidity motives, and heterogeneous across jurisdictions.

Figure 4 plots the conditional variances that are well-behaved, with the peaks reached after the Lehman collapse in 2008 and the exacerbation of the sovereign debt crisis in 2011 and 2012.[16] The dynamics of the volatilities between the two models are identical for all countries.

[Insert Figure 4, here]

4.3. Correlation dynamics

Conditional correlations began moving out of the non-crisis regime as early as in August 2007, when the interbank credit crisis began. By summer 2008, the correlations were close to 50% in many countries, suggesting that the risk-free rate and credit risk were no longer key drivers of sovereign yields in the euro area. Conversely, the conditional correlations estimated using the DCC-GARCH models report a change only after the Lehman Brothers collapse in some countries, and a clear and persistent correlation breakdown only during the sovereign debt crisis.

After the Lehman Brothers collapse, the correlations estimated using the STCC-GARCHX model further declined as the financial crisis developed and investors began repricing risk. The situation began to improve during spring 2009 following the announcement of stringent fiscal stabilization measures by the Irish government on February 22, 2009. It can be argued that the improvement was instead a result of receding global uncertainty. However, the STCC-GARCHX models with VIX and VSTOXX as transition variables (not shown) do not support this argument. On the other hand, the correlations slightly increased during 2009 following the sharp decline in the KfW-Bund spread, but then declined after the disclosure of Greece's severe fiscal problems in October 2009.

The plateau at low levels was reached in 2011 and 2012 in all countries (except Germany) in a full crisis-regime mode (see Figure 3). The volatility reached its highest point in all countries, including Germany either immediately after the Lehman collapse (Finland,

[16] The GARCH parameters and their sums point to a persistence effect in the volatility of sovereign yields and OIS rates, although none of the models come close to an integrated GARCH, with the STCC-GARCHX showing less overall persistence than the DCC-GARCH models. Hillebrand and Medeiros (2009) provide an extensive discussion of long-range dependence and structural change in a realized volatility framework. Amado and Teräsvirta (2014) examine the short- and long-run properties of the conditional correlations in a multivariate GARCH framework with a non-stationary component in the variance equations.

Germany, the Netherlands) or before the speech of Mario Draghi in London on July 26, 2012 (all other countries). Following the launch of OMTs in summer 2012, the correlations began rising and the volatilities decreasing, reflecting the abatement of the euro area break-up risk.

The market situation further improved from the beginning of 2014, as market participants expected additional expansionary monetary policy measures. A gradual improvement in the financial conditions is also reflected in the KfW-Bund spread, which provides a timely signal of a breakdown or improvement in the conditional correlations between sovereign yields and the risk-free rate due to the flight-to-liquidity phenomena that have characterized the euro area sovereign debt market during the financial crisis and captured by this indicator.

5. Interpreting the results

5.1. Reasons for the correlation breakdown

Can we explain the reasons underlyign the breakdown in the transmission mechanism from the risk-free rate to sovereign yields?

Three key episodes have characterized the financial crisis in Europe. The first is associated with the interbank lending crisis. In August 2007, interbank lending collapsed because of problems associated with asset-backed securities. The comparison between unsecured (EURIBOR) and secured (OIS) market rates heightened the perceptions of counterparty risk in the banking system. The EURIBOR-OIS spread rose sharply to above 60 basis points (see the top plot on the left in Figure 5). The same indicator ballooned to 200 basis points with the bankruptcy of Lehman Brothers. Hence, the three-month EURIBOR-OIS spread is a good proxy to measure the strains on the interbank lending market.

[Insert Figure 5, here]

Following the disclosure of Greece's severe fiscal problems in October 2009, sovereign spreads sharply increased in most of the euro area countries, causing the European monetary union to face its largest ever challenge. Therefore, the second episode is associated with the sovereign risk in Greece (see the middle plot on the left in Figure 5).

The third episode is associated with the widespread contagion of the Italian and Spanish sovereign debt markets and the risk of the euro area break-up, labelled "the redenomination risk", namely the compensation demanded by market participants for the risk that a euro asset could be redenominated into a devalued new legacy currency. De Santis (2015) employs the quanto CDS of Italy and Spain relative to the quanto CDS in Germany as a measure of the redenomination risk. The quanto CDS, namely the difference between the CDS quotes in US dollars and euros, was expected to be positive for Italian and Spanish sovereign contracts during the first half of 2012, since being paid-off in euro-denominated credit event protection seemed less attractive. Should, for example, Italy declare a credit event, 10 million euro worth of protection on an Italian credit event would be worth much less after the event given the devaluation of the euro or, if the euro were to no longer exist, as a result of the devaluation of the "new" currency. This led to a major drop in the demand for EUR-denominated Italian (and other European) protection. The use of countries' quanto CDS relative to Germany rests on the idea that such a spread would be close to zero if the break-up risks of the euro area were minor (see the bottom plot on the left of Figure 5). The proposed measure for currency redenomination risk peaked for both Italy and Spain just before the speech made by ECB President Draghi on July 26, 2012. Thereafter, it declined and, since the end of 2012, has remained contained.

To formally address these potential explanations, we make use of impulse response functions (IRFs) generated by a multivariate PVAR, which takes the following form:

$$A_0^G Y_{i,t}^G = \phi_i^G + \sum_{j=1}^{p} A_j^G Y_{i,t-j}^G + \eta_{i,t}^G, \tag{11}$$

where G is the group of pooled countries, $Y_{i,t}^G$ is the 5×1 vector of variables observed at time t for country i, ϕ_i^G is the country-fixed effect that allows for cross-country heterogeneity, p is the lag length equal to two according to the Akaike information criterion (AIC) and $\eta_{i,t}$ is a 5×1 vector of innovations, defined as being uncorrelated among factors but interdependent across i within G. A_0^G is the impact matrix. The dynamics of the system are assumed to be heterogeneous among G, through $\sum_{j=1}^{p} A_j^G$. Therefore, the intercepts and the variance of the shocks are unit-specific, while the slopes are heterogeneous across groups. These assumptions allow for ample cross-sectional heterogeneity.

The variables include the US VIX to control for global factors, the estimated conditional correlations that measure the pass-through from the risk-free rate to the sovereign yield, the three-month EURIBOR-OIS spread as a proxy for interbank-credit risk, the 10-year Greek sovereign yield-OIS spread as a proxy for Greek sovereign risk, and the average quanto CDS spread of Italy and Spain relative to Germany as a proxy for redenomination risk.

Restrictions must be imposed on A_0^G to uniquely recover the structural form. The identification restrictions imposed on A_0^G are recursive, that is they are equivalent to a Cholesky factor of the variance-covariance matrix of the reduced-form white noise innovations with the three factors ordered last. This assumption is rather conservative, as we assume that all common contemporaneous innovations are generated by the US VIX and the conditional correlations.

As for the ordering of the factors, we follow the time-line of episodes previously discussed, with the EURIBOR-OIS spread being ordered third, and the Greek sovereign yield-OIS spread fourth, followed by the relative quanto CDS spread. Hence, the shock to the

EURIBOR-OIS spread is interpreted as a shock to the interbank lending market, that to the Greek sovereign spread is considered as a shock from re-pricing the Greek sovereign debt, and the relative quanto CDS is a shock from re-pricing the break-up risk.

First, we present the results pooling all euro area countries and assuming cross-sectional slope homogeneity, but with the intercepts and the variance of the shocks remaining unit-specific. The IRFs for the STCC-GARCHX model estimated for the entire sample 2005-2015 are reported in Figure 5. They indicate that the estimated correlations respond negatively to all three factors. In particular, the responses peak after 250 days, if the shock is generated within the interbank market, after 100 days if the euro is questioned by the market, and after few days if the shock originated in Greece.

Given the role played by the Greek crisis in 2010-2011 and the euro area's break-up risk since 2011, we compare the results of the models estimated during the crisis period of 2007-2015, but separated into two sub-samples for the last quarter of 2011, when the market questioned the irreversibility of the euro.

The results reported in Figure 6 suggest that the shock generated within the interbank credit market affected the transmission channel from the risk-free rate to sovereign yields during both sub-periods, with the largest impact in the first period, while the Greek sovereign debt crisis affected such a pass-through in the first phase of the crisis with a large and prolonged impact, and the redenomination risk affected the dynamics of the conditional correlations in the second phase.

[Insert Figure 6, here]

Second, we present the results while pooling the countries into two groups: the more vulnerable countries (Ireland, Italy, Portugal and Spain), characterized by higher sovereign and CDS spreads and larger correlation breakdown, and all other euro area countries. The results are presented separately for the two sample periods in Figures 7 and 8. There are large similarities between the two groups, which confirm that the correlation breakdown initiated with tensions in the interbank market continued with the re-pricing of the Greek sovereign debt in the first phase of the crisis and ended with the re-pricing of the euro area break-up risk in the second phase. As one expected, the impact of the Greek sovereign risk and redenomination risk is larger for the more vulnerable countries.

Third, we investigate the results while pooling countries as per the similarities in the parameter estimates: (i) Italy, Spain and Portugal; (ii) Austria and Belgium and (iii) Finland, Netherlands and France. The IRFs (not shown) are similar to those of the vulnerable countries in the first group and of other countries in the second and third groups.

In sum, these results are robust to (i) further group disaggregation, (ii) methods of estimation of conditional correlations, (iii) alternative ordering schemes to identify shocks, and (iv) the assumption that $\phi_i = \phi$.

[Insert Figures 7 and 8, here]

5.2. Impact of ECB asset purchase programmes on correlations

After sovereign and CDS spreads reached their peak, Mario Draghi pledged to do "whatever it takes" to address the break-up risk on July, 26 2012. On August 2, 2012, he announced the discussion of the Governing Council to address the severe malfunctioning in the price formation process in the bond markets of euro area countries.[17] The actual transactions were never executed, but the upward shift of the correlations following July, 26 2012 is evident in Figure 3.

The quantification of the impact of policy measures on asset prices is often based on price changes observed within a narrow time window surrounding the policy announcement. We can capture the OMT effect using a dummy variable, which takes the value of one for two consecutive days after the 26 July and 2 August.

Furthermore, the ECB launched its quantitative easing (QE) programme in January 2015. Through the PSPP, the ECB aimed to purchase 1.14 trillion euro of public and private sector securities from March 2015 to September 2016 representing 11% of nominal GDP, a program further expanded in December 2015. The monetary policy announcement in January 2015 was already implicitly communicated to the market in 2014. Therefore, we can assess whether the program helped restore the conditional correlations between sovereign yields and the risk-free rate to higher values in 2014 and 2015, as is evident in Figure 3.

However, the identification of the PSPP impact for the euro area is more challenging, because the monetary policy announcement on January 22, 2015, followed by the details of the new program and the initial purchases in March 2015, was implicitly communicated to the market in autumn 2014 and many believed that the Jackson Hole speech by Mario Draghi on August 22, 2015 already raised such expectations, as worries about rising deflationary risks coupled with negative news about the real economy were communicated. This speech is relevant against the background that, on April 24, 2014, Mario Draghi stated that the worsening of the medium-term inflation outlook would provide a reason for broad-based asset purchases.[18]

[17]In August 2012, as pointed out by Mario Draghi, *"the Governing Council of the Eurosystem extensively discussed the policy options to address the severe malfunctioning in the price formation process in the bond markets of euro area countries. Exceptionally high risk premia are observed in government bond prices in several countries and financial fragmentation hinders the effective working of monetary policy. Risk premia that are related to fears of the reversibility of the euro are unacceptable, and they need to be addressed in a fundamental manner. The euro is irreversible."*

[18]See `https://www.ecb.europa.eu/press/key/date/2014/html/sp140424.en.html`.

After the Jackson Hall speech, stocks rose, the euro fell and bond yields dropped as the comments fanned speculation that the ECB was heading for further monetary stimulus.[19] This suggests that event studies based on the official announcement on January 22, 2015 would not be satisfactory to identify the PSPP shock for the euro area.

To estimate the incremental impact of the expected monetary policy intervention, we examine the impact of market news about the ECB asset purchases. The PSPP for the euro area is identified using the number of references to the program in news stories recorted on Bloomberg, with the underlying intuition being that the more intense the discussion about the program, the greater the expectation from the euro area PSPP. In particular, the news variable is defined as the sum of all Bloomberg news jointly containing the following keywords: "Draghi, and QE or quantitative easing, and sovereign, and euro area". To be assured about the suggested identification, a similar series containing the following keywords was constructed: "Bernanke or Yellen, and QE or quantitative easing, and US". The correlation between the EA and US QE news is statistically non-significant.

We use the country-panel VAR described by the system of equation (11) and include as exogenous regressors the OMT dummy, the ECB and FED Bloomberg news (i.e. VARX). The results reported in Table 7 suggest that the ECB policies increased the conditional correlations and the coefficients are strongly statistically significant. For brevity, only the coefficients for OMT and PSPP shocks are shown.

In addition, to control for contemporaneous developments among the regressors, we regress the first difference of the conditional correlations estimated using either with DCC-GARCH or STCC-GARCHX against the OMT dummy, the ECB and FED Bloomberg news, the lagged endogenous variable and the following regressors in first difference with their respective lag also used in the VAR: the US VIX, the three-month EURIBOR-OIS spread, the 10-year Greek sovereign-yield OIS spread, and the average quanto CDS spread of Italy and Spain relative to Germany. The OLS econometric results confirm that the OMT and the PSPP have been important instruments to restore or improve the correlation between the OIS rate and sovereign yields (see Table 7).

The country group split is very informative. While the OMT announcement has posi-

[19]For example, on August 25, 2014, Simon Kennedy and Alessandro Speciale in a Bloomberg article entitled "Draghi pushes ECB closer to QE as deflation risks rise" reported the following: The 22 August speech "was a major event and marked a turning point in ECB rhetoric," said Philippe Gudin, chief European economist at Barclays Plc in Paris. "We think the recent economic developments have increased the chance of outright QE as the next step." The article also stated that in the previous week Citigroup Inc.'s economists predicted that the ECB would unveil a QE program in December valued at 1 trillion euro ($1.3 trillion), split between public and private assets and aimed at reducing borrowing costs and increasing liquidity. See `http://www.bloomberg.com/news/articles/2014-08-24/draghi-pushes-ecb-closer-to-qe-as-deflation-risks-rise`.

tively affected the correlations of the less vulnerable countries in the Eurosystem (i.e. the coefficients of the more vulnerable countries have a similar magnitude but are not statistically significant), the PSPP has been successful in affecting the correlations of the more vulnerable countries. The estimated coefficients for the more vulnerable country's group are three times those of all other countries and are all statistically significant.

[Insert Table 7, here]

6. Robustness

Our results are robust to alternative univariate volatility specifications, such as an asymmetric GARCH of the Glosten-Jagannathan-Runkle type (Glosten et al. (1993)) used by Silvennoinen and Teräsvirta (2015).

We also consider a specification with two transition variables and four regimes of correlation between which the process may vary, using the double STCC (DSTCC) GARCH model of Silvennoinen and Teräsvirta (2009). The two transition variables include those used for the auxiliary regression in equation (3) and the KfW-Bund spread. Except for changes in the transition speeds, the conditional variances and correlation patterns remain similar to those already presented in Figure 3. Moreover, the threshold of one transition variable and the correlation of at least one additional regime are non-significant in these robustness checks. In other words, the combination of transition variables does not strengthen the estimation of the correlations obtained using the suggested STCC-GARCHX models.

It could be argued that an STCC-GARCH(X) model is redundant for the German Bund. However, the LM-type test of constant conditional correlation (CCC) under the null against the alternative of the STCC-GARCH model, proposed by Silvennoinen and Teräsvirta (2015),[20] still suggests the rejection of the null hypothesis for the German Bund. Further, the CCC-GARCH correlation of about 82% for Germany is close to the means of the conditional correlations obtained using the STCC-GARCHX model. This evidence further supports our choice of adopting a regime-dependent correlation model which is stable and reliable.

Finally, we investigate the role of non-linearity in the mean equations. the results for the conditional correlations remain invariant.

[20] Silvennoinen and Teräsvirta (2005) and Silvennoinen and Teräsvirta (2015) base their specification on an earlier test proposed by Tse (2000), Engle and Sheppard (2001) and Bera and Kim (2002).

7. Conclusions

The identification of a correlation breakdown between the risk-free rate and the non-credit risk component of sovereign yields is paramount for policy-makers, given the benchmark role played by sovereign yields in the transmission mechanism of monetary policy. This question is even more important in a monetary union context, given that the single monetary policy instruments may result in being ineffective if financial markets are fragmented among member states.

To address this issue, we estimate the conditional correlations between the five-year risk-free rate and the sovereign yields adjusted for the credit risk in the mean and variance equations, conditional on time-varying liquidity premia and uncertainty measures, of the largest ten euro area countries. All the volatility, correlation and transition parameters of the newly introduced STCC-GARCHX models are jointly estimated using one single step, rather than by conditional maximum likelihood.

We provide evidence that the one-step joint estimation outperforms the respective estimation in terms of likelihood values. In addition, we show that the credit risk plays an important role in both the conditional correlations and volatility in most euro area sovereign debt markets, particularly those most affected by the sovereign debt crisis. Nevertheless, the conditional correlations between the sovereign yields adjusted for credit risk and the risk-free rate sharply declined during the hikes in the sovereign debt crisis in many euro area member states. The breakdown in the conditional correlations is heterogeneous among country groupings, with the more vulnerable countries (Ireland, Italy, Portugal and Spain) being affected the most.

We also discuss the reasons underlying the correlation breakdown. We found evidence that the counterparty risk in the banking system played an important role throughout the sample period. In addition, the conditional correlations were affected by developments in Greek sovereign risk in 2010 and 2011 and the perceived euro area's break-up risk since he end of 2011. The change in trend began with the launch of outright monetary transactions by the ECB in summer 2012, which abated the euro break-up risk. Moreover, the financial market situation further improved from the beginning of 2014, when financial markets priced in a higher probability of additional expansionary monetary policy measures in the euro area. In particular, the econometric results suggest that the ECB public asset purchase programme has been instrumental in improving the conditional correlations in 2014 and 2015, particularly for the more vulnerable countries.

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Table 1: Descriptive Statistics

This table shows the descriptive statistics of the data set. The country coverage includes Austria (AUT), Belgium (BEL), Finland (FIN), France (FRA), Germany (GER), Ireland (IRE), Italy (ITA), the Netherlands (NED), Portugal (POR) ad Spain (ESP). The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability.

Panel A: Sovereign Yield										
	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP
Mean	240.35	267	237.03	166.05	210.77	401.06	343.95	225.62	518.22	345.14
Min	-6.9	-5.8	1.65	-8.55	-12.2	22.5	43.4	-7.15	81	45.65
Max	487.35	540.35	492.25	486.73	476.15	1761.3	770.35	487.95	2174.75	749.8
Panel B: Change Sovereign Yield										
	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP
Mean	-0.1	-0.11	-0.1	-0.23	-0.11	-0.09	-0.09	-0.1	-0.07	-0.08
Min	-38.05	-53.1	-29.25	-31.33	-21.6	-160.2	-95.65	-22.5	-221.5	-99.3
Max	37.4	45.9	33.65	34.33	29.7	165.2	70.2	24.7	289.9	51.75
Panel C: CDS Spread										
	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP
Mean	53.99	74.26	56.48	35.49	29.29	203.57	146.28	38.21	283.23	146.2
Min	0.5	1	0.5	6.5	0.6	1.5	5.3	1	3.4	1.05
Max	273	403.01	247.08	93.92	120.59	1286.91	595.68	136.21	1762.1	636.68
Panel D: Change CDS Spread										
	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP
Mean	0.01	0.02	0.02	0.01	0.01	0.02	0.04	0.01	0.05	0.04
Min	-27.35	-56.64	-29.7	-9.9	-14.32	-178.69	-76.36	-14.22	-192.17	-79.21
Max	41.8	36.67	22.82	12.8	10.97	119.18	72.15	23.7	174.99	54.09
Panel E: Bid-Ask Spread										
	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP
Mean	2.5	1.83	1.37	1.86	0.49	11.3	2.11	1.15	17.98	2.67
Min	0.3	0.2	0.2	0.4	0	0.2	0.2	0.2	0.5	0.4
Max	14.5	10.3	6	4.35	1.5	89.2	18.1	3.4	157.1	27.6
Panel F: Change Bid-Ask Spread										
	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP
Mean	0	0	0	0	0	0	0	0	0	0
Min	-6.8	-3.8	-2	-2.75	-1.3	-48.6	-12	-1.8	-47	-15.6
Max	4.3	6.1	2.3	2.75	1	53.7	11	1.6	57.2	20.2

Panel G: Economic Variables						
	Mean	Min	Max	Mean Change	Min Change	Max Change
OIS	2.32	-0.02	4.81	-0.12	-22.6	21.45
KfW-Bund Spread	30.62	0.67	94.05	0	-18	18.1
VSTOXX	20.87	8.68	79.28	0	-15.52	18.23

Table 2: Auxiliary Regressions

This table shows the coefficient estimates of the regressions $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1} \triangle BA_{i,t} + \phi_{i2} \triangle KfW_t + \phi_{i3} \triangle VSTOXX_t + \xi_{i,t}$ (see equation (3)), where $CDS_{i,t}$ denotes the CDS spread of a sovereign bond i at time t at five-year maturity, $BA_{i,t}$ the bid-ask spread of a sovereign bond i at time t at five-year maturity, $VSTOXX_t$ the implied volatility of option prices on EURO STOXX 50 at time t. The country coverage includes Austria (AUT), Belgium (BEL), Finland (FIN), France (FRA), Germany (GER), Ireland (IRE), Italy (ITA), the Netherlands (NED), Portugal (POR) ad Spain (ESP). The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. HAC-robust standard errors are reported in parenthesis. ***, ** and * indicate significant coefficients at the 1%, 5% and 10% levels, respectively.

	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP
Constant	0.01	0.0159	0.008	0.0153	0.0048	0.0187	0.0396	0.0077	0.0515	0.035
	(0.0809)	(0.088)	(0.0399)	(0.0604)	(0.0334)	(0.2611)	(0.1569)	(0.0441)	(0.3711)	(0.1606)
Bid-Ask Spread	0.3091	0.0725	0.348**	0.3802	0.3619	0.1292	0.5167*	0.2705	0.139	0.4447***
	(0.257)	(0.4073)	(0.1734)	(0.4662)	(0.364)	(0.1018)	(0.3185)	(0.2335)	(0.149)	(0.1825)
KfW-Bund Spread	0.3591***	0.5061***	0.1348***	0.3336***	0.1651***	1.2282***	1.1376***	0.1955***	1.8289***	1.0899***
	(0.0636)	(0.0905)	(0.0226)	(0.0544)	(0.0283)	(0.2149)	(0.169)	(0.0358)	(0.3567)	(0.1585)
VSTOXX	0.3295***	0.4534***	0.0917***	0.3243***	0.1614***	0.9098***	0.8661***	0.1783***	1.1656***	0.7838***
	(0.0691)	(0.1108)	(0.0273)	(0.0684)	(0.0286)	(0.2187)	(0.187)	(0.0383)	(0.3384)	(0.1622)
$adj.R^2$	0.0997	0.1189	0.0964	0.1319	0.1119	0.0924	0.1625	0.0911	0.0859	0.1459

Table 3: Mean Equations

This table shows the coefficient estimates of the mean equations for two different models. Panel A contains the estimation results of a simple AR(1) model, $\triangle r_{i,t} = \beta_{i0} + \beta_{i1}\triangle r_{i,t-1} + \varepsilon_{i,t}$ (see equation (2)), where $r_{i,t}$ denotes the first difference of the OIS rate or the sovereign bond yield i at time t at five-year maturity. Panel B provides the OLS estimates of the mean equation $\triangle r_{i,t} = \beta_{i0} + \beta_{i1}\triangle r_{i,t-1} + \beta_{i2}\hat{\xi}_{i,t} + \varepsilon_{i,t}$ (see equation (4)), where $\hat{\xi}_{i,t}$ denotes credit risk obtained as the residuals of the following regression $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1}\triangle BA_{i,t} + \phi_{i2}\triangle KfW_t + \phi_{i3}\triangle VSTOXX_t + \xi_{i,t}$, (see equation (3)). The country coverage includes Austria (AUT), Belgium (BEL), Finland (FIN), France (FRA), Germany (GER), Ireland (IRE), Italy (ITA), the Netherlands (NED), Portugal (POR) ad Spain (ESP). The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. HAC-robust standard errors are reported in parenthesis. ***, ** and * indicate significant coefficients at the 1%, 5% and 10% levels, respectively.

Panel A: AR(1) process

	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP	OIS
Constant	-0.0923	-0.0839	-0.2222**	-0.0946	-0.1074	-0.0798	-0.0732	-0.0938	-0.0544	-0.0686	-0.1136*
	(0.0978)	(0.1199)	(0.0995)	(0.096)	(0.0945)	(0.4758)	(0.1639)	(0.0956)	(0.1602)	(0.1414)	(0.0847)
AR(1)	0.1076***	0.2122***	0.0359	0.0568**	0.032*	0.1339***	0.1448***	0.0531**	0.2176***	0.181***	-0.04*
	(0.0279)	(0.0439)	(0.0329)	(0.0273)	(0.0246)	(0.0535)	(0.0444)	(0.0244)	(0.0675)	(0.0363)	(0.025)
$adj.R^2$	0.0116	0.0451	0.0021	0.0032	0.0011	0.0179	0.021	0.0029	0.0473	0.0328	0.0017

Panel B: The role of credit risk

	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP
Constant	-0.0914	-0.0881	-0.2234**	-0.0943	-0.1086	-0.0882	-0.0794	-0.0935	-0.0622	-0.0764
	(0.0969)	(0.1047)	(0.0988)	(0.0922)	(0.0901)	(0.3383)	(0.138)	(0.099)	(213.1832)	(0.131)
AR(1)	0.1158***	0.1687***	0.0303	0.0605**	0.0216	0.0385	0.0682**	0.0559**	0.0959	0.0839***
	(0.0272)	(0.0359)	(0.0328)	(0.0264)	(0.0237)	(0.0421)	(0.035)	(0.0245)	(0.4156)	(0.027)
Credit risk	0.2061***	0.4842***	-0.1172*	0.3075***	-0.2981***	0.6363***	0.6758***	0.088	0.727***	0.7298***
	(0.0508)	(0.0473)	(0.0846)	(0.0638)	(0.0919)	(0.0665)	(0.0437)	(0.0768)	(0.1347)	(0.0413)
$adj.R^2$	0.0323	0.1817	0.0039	0.0351	0.01	0.2747	0.378	0.0043	0.4147	0.4356

Table 4: Smooth Transition Conditional Correlation GARCHX (STCC-GARCHX): Comparison among Estimation Methods

This table shows the STCC-GARCHX estimates for sovereign yields of France, Germany and Italy and the OIS rate at 5-year maturity using three different estimation methods. The comparison is based on the model $\triangle r_{i,t} = \beta_{i0} + \beta_{i1}\triangle r_{i,t-1} + \beta_{i2}\hat{\xi}_{i,t} + \varepsilon_{i,t}$ for the conditional mean (see equation (4)). $\hat{\xi}_{i,t}$ denotes credit risk obtained as the residuals of the following regression $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1}\triangle BA_{i,t} + \phi_{i2}\triangle KfW_t + \phi_{i3}\triangle VSTOXX_t + \xi_{i,t}$ (see equation (3)). To obtain the conditional variance-covariance matrix $H_t = D_t R_t D_t$ with $D_t = (h^{1/2}_{i,t}, h^{1/2}_{OIS,t})$, the conditional variance is estimated as a GARCHX(p,q,r) model with $p = q = r = 1$ of the form $h_{it} = a_{i0} + \sum_{j=1}^{q} a_{ij}\varepsilon^2_{i,t-j} + \sum_{l=1}^{p} b_{il}h_{i,t-l} + \sum_{k=1}^{r}\delta_{ik}\hat{\xi}^2_{i,t-k}$, $\varepsilon_{i,t} = h^{1/2}_{i,t}z_{i,t}$ *and* $\varepsilon_{i,t} \mid \psi_{i,t-1} \sim N(0, h_{i,t})$ (see equation (9)). The STCCs are estimated assuming that $R_t = (1 - G_t)\cdot R_1 + G_t \cdot R_2$, $G_{,t}(\gamma, c, s_t) = (1 + \exp\{-\gamma(s_t - c)\})^{-1}$ and $\gamma > 0$ (see equations (7) and (8)). The chosen transition variable is the lagged value of the KfW-Bund spread ($s_t = KfW_{t-1}$). The two-step estimation refers to the DCC-GARCH approach where the conditional variances are estimated in a first step and the conditional correlations are estimated in a second step. The iterative estimation refers to the conditional maximum likelihood estimation proposed by Silvennoinen and Teräsvirta (2015). The joint estimation is the estimation of all the parameters in a single step after an extensive grid search. *Correlation Regime 1* and *Correlation Regime 2* report the off-diagonal elements of the 2x2 correlation matrices R_1 and R_2 (see equation (7)). The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. HAC-robust standard errors are reported in parenthesis. ***, ** and * indicate significant coefficients at the 1%, 5% and 10% levels, respectively.

	France	France	France	Germany	Germany	Germany	Italy	Italy	Italy
	Two-step	Iterative	Joint	Two-step	Iterative	Joint	Two-step	Iterative	Joint
Constant (Sov. Yield)	0.1255*	0.3848	0.5056*	0.1423***	0.2782**	0.216**	0.9813***	1.1454***	1.2153***
	(0.0943)	(0.328)	(0.3622)	(0.0599)	(0.1524)	(0.095)	(0.3536)	(0.3571)	(0.3746)
ARCH Sov. Yield	0.0515*	0.0595**	0.0653**	0.0488***	0.0574***	0.0542***	0.0587***	0.0673***	0.0691***
	(0.0341)	(0.036)	(0.0353)	(0.0111)	(0.0128)	(0.0111)	(0.0189)	(0.0199)	(0.0205)
GARCH Sov. Yield	0.9445***	0.9194***	0.9095***	0.9437***	0.9267***	0.9335***	0.8645***	0.8316***	0.8274***
	(0.0326)	(0.049)	(0.0486)	(0.0132)	(0.02)	(0.0155)	(0.0329)	(0.0403)	(0.0415)
Variance credit risk	0.0006	0.007	0.007	0.0226	0.0287	0.0292	0.044***	0.0683***	0.0702***
	(0.0057)	(0.007)	(0.0077)	(0.0242)	(0.0274)	(0.0241)	(0.0122)	(0.0229)	(0.0236)
Constant (OIS)	0.0326**	0.2173**	0.2905***	0.1092***	0.2268**	0.1759***	0.1092***	0.2549***	0.2764***
	(0.0157)	(0.0971)	(0.1148)	(0.0383)	(0.1116)	(0.0616)	(0.0346)	(0.0766)	(0.0796)
ARCH OIS	0.0399***	0.0474***	0.0514***	0.052***	0.062***	0.0576***	0.052***	0.0545***	0.0554***
	(0.014)	(0.0126)	(0.0127)	(0.0083)	(0.01)	(0.0094)	(0.0098)	(0.0093)	(0.0094)
GARCH OIS	0.9589***	0.9379***	0.9311***	0.9426***	0.9243***	0.9328***	0.9426***	0.9292***	0.9277***
	(0.0124)	(0.0174)	(0.0179)	(0.0082)	(0.0153)	(0.0113)	(0.0098)	(0.0125)	(0.0125)
Gamma	0.7009***	0.5278***	0.9168***	0.5252	0.3922***	21.3065***	1.025***	1.5178***	1.5616***
	(0.2311)	(0.0141)	(0.1626)	(0.5036)	(0.0464)	(8.4759)	(0.1318)	(0.1564)	(0.2307)
Threshold c	74.8585***	86.3615***	41.6898**	60.8099***	70.0903	7.6958***	50.7475***	23.8822***	24.6355***
	(15.519)	(17.3828)	(18.4845)	(11.0074)	(138.1398)	(0.8476)	(3.1072)	(4.3078)	(6.4456)
Correlation Regime 1	0.9473***	1***	1***	0.9592***	1***	0.9354***	0.9066***	1***	1***
	(0.1097)	(0.0333)	(0.0519)	(0.1792)	(0.0452)	(0.01)	(0.0822)	(0.0428)	(0.0823)
Correlation Regime 2	0.1745**	0.0176	0.4531***	0.5808***	0.5052	0.798***	-0.282***	0.1227**	0.1311**
	(0.1024)	(0.1726)	(0.1229)	(0.1779)	(0.67)	(0.0116)	(0.0922)	(0.0683)	(0.0759)
Log Likelihood	-13041	-13014	-13006	-13078	-13068	-13047	-14607	-14576	-14575

Table 5: Dynamic Conditional Correlation GARCH (DCC-GARCH)

This table shows the DCC-GARCH estimates for sovereing yield i and the OIS rate at time t at five-year maturity using $\triangle r_{i,t} = \beta_{i0}+\beta_{i1}\triangle r_{i,t-1}+\beta_{i2}\hat{\xi}_{i,t}+\varepsilon_{i,t}$ for the conditional mean (see equation (4)). $\hat{\xi}_{i,t}$ denotes credit risk obtained as the residuals of the following regression $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1}\triangle BA_{i,t} + \phi_{i2}\triangle KfW_t + \phi_{i3}\triangle VSTOXX_t + \xi_{i,t}$ (see equation (3)). To obtain the conditional variance-covariance matrix $H_t = D_t R_t D_t$ with $D_t = (h^{1/2}_{i,t}, h^{1/2}_{OIS,t})$, the conditional variance is estimated in the first step by employing a GARCH(p,q) with $p = q = 1$ of the standard form $h_{i,t} = a_{i0} + \sum_{j=1}^{q} a_{ij}\varepsilon^2_{i,t-j} + \sum_{l=1}^{p} b_{il}h_{i,t-l}$. The country coverage includes Austria (AUT), Belgium (BEL), Finland (FIN), France (FRA), Germany (GER), Ireland (IRE), Italy (ITA), the Netherlands (NED), Portugal (POR) ad Spain (ESP). The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. Note that the parameter estimates for the variance of the OIS rate differ only for the countries that have different sample sizes, and otherwise are identical. HAC-robust standard errors are reported in parenthesis. ***, ** and * indicate significant coefficients at the 1%, 5% and 10% levels, respectively.

	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP
Constant (Sov. Yield)	0.1739**	0.1779	0.2052	0.1193	0.0354	0.8074**	0.3756**	0.0529	0.2139	0.1055*
	(0.0921)	(0.169)	(0.2517)	(0.1047)	(0.0403)	(0.429)	(0.1734)	(0.0528)	(0.2191)	(0.0705)
ARCH Sov. Yield	0.0517***	0.0472**	0.0467	0.0523***	0.0352***	0.1222***	0.0739***	0.0352***	0.0746***	0.0515***
	(0.0127)	(0.0225)	(0.041)	(0.0188)	(0.0072)	(0.0359)	(0.0145)	(0.0106)	(0.0218)	(0.0113)
GARCH Sov. Yield	0.9424***	0.9469***	0.9417***	0.9442***	0.9637***	0.8778***	0.9181***	0.9628***	0.9254***	0.9481***
	(0.0143)	(0.0268)	(0.0512)	(0.0209)	(0.008)	(0.0342)	(0.0154)	(0.0119)	(0.026)	(0.0104)
Constant (OIS)	0.0328	0.0328	0.0386	0.0311	0.0328	0.0328	0.0328	0.0307	0.0328	0.0328
	(0.0359)	(0.0359)	(0.0341)	(0.0302)	(0.0359)	(0.0359)	(0.0359)	(0.0355)	(0.0359)	(0.0359)
ARCH OIS	0.0395***	0.0395***	0.0471***	0.0395***	0.0395***	0.0395***	0.0395***	0.0392***	0.0395***	0.0395***
	(0.0091)	(0.0091)	(0.0136)	(0.0088)	(0.0091)	(0.0091)	(0.0091)	(0.0092)	(0.0091)	(0.0091)
GARCH OIS	0.9592***	0.9592***	0.9511***	0.9593***	0.9592***	0.9592***	0.9592***	0.9595***	0.9592***	0.9592***
	(0.0102)	(0.0102)	(0.0139)	(0.0094)	(0.0102)	(0.0102)	(0.0102)	(0.0102)	(0.0102)	(0.0102)
DCC a	0.0877***	0.0304***	0.033	0.0155*	0.0147*	0.0664***	0.0324**	0.0852***	0.0328***	0.0234***
	(0.0304)	(0.0108)	(0.0833)	(0.0113)	(0.0094)	(0.0187)	(0.0145)	(0.024)	(0.008)	(0.0087)
DCC b	0.8389***	0.9657***	0.8313***	0.98***	0.9792***	0.9211***	0.9658***	0.8068***	0.9659***	0.9755***
	(0.0471)	(0.0122)	(0.0489)	(0.0163)	(0.0159)	(0.0146)	(0.0158)	(0.0711)	(0.0085)	(0.0094)
Log Likelihood	-13501	-13831	-9250	-13040	-13028	-15677	-14588	-12790	-15723	-14604

Table 6: Smooth Transition Conditional Correlation GARCHX (STCC-GARCHX)

This table shows the STCC-GARCHX estimates for sovereign yield i and the OIS rate at five-year maturity using $\triangle r_{i,t} = \beta_{i0} + \beta_{i1}\triangle r_{i,t-1} + \beta_{i2}\hat{\xi}_{i,t} + \varepsilon_{i,t}$ for the conditional mean (see equation (4)). $\hat{\xi}_{i,t}$ denotes credit risk obtained as the residuals of the following regression $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1}\triangle BA_{i,t} + \phi_{i2}\triangle KfW_t + \phi_{i3}\triangle VSTOXX_t + \xi_{i,t}$ (see equation (3)). To obtain the conditional variance-covariance matrix $H_t = D_t R_t D_t$ with $D_t = (h_{i,t}^{1/2}, h_{OIS,t}^{1/2})$, the conditional variance is estimated as a GARCHX(p,q,r) model with $p = q = r = 1$ of the form $h_{i,t} = a_{i0} + \sum_{j=1}^{q} a_{ij}\varepsilon_{i,t-j}^2 + \sum_{l=1}^{p} b_{il}h_{i,t-l} + \sum_{k=1}^{r} \delta_{ik}\hat{\xi}_{i,t-k}^2$, $\varepsilon_{it} = h_{i,t}^{1/2} z_{i,t}$ *and* $\varepsilon_{i,t} \mid \psi_{i,t-1} \sim N(0, h_{i,t})$ (see equation (9)). The smooth transition conditional correlations are estimated assuming $R_t = (1 - G_t) \cdot R_1 + G_t \cdot R_2$, $G_t(\gamma, c, s_t) = (1 + \exp\{-\gamma(s_t - c)\})^{-1}$ and $\gamma > 0$ (see equations (7) and (8)). The transition variable is the lagged value of the KfW-Bund spread ($s_t = KfW_{t-1}$). *Correlation Regime 1* and *Correlation Regime 2* report the off-diagonal elements of the $2x2$ correlation matrices R_1 and R_2 (see equation (7)). The country coverage includes Austria (AUT), Belgium (BEL), Finland (FIN), France (FRA), Germany (GER), Ireland (IRE), Italy (ITA), the Netherlands (NED), Portugal (POR) ad Spain (ESP). The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. HAC-robust standard errors are reported in parenthesis. ***, ** and * indicate significant coefficients at the 1%, 5% and 10% levels, respectively. ***, ** and * reported for the log likelihood refer to the likelihood ratio test that compares the STCC-GARCHX versus the respective STCC-GARCH specification.

	AUT	BEL	FIN	FRA	GER	IRE	ITA	NED	POR	ESP
Constant (Sov. Yield)	0.8785***	1.9967***	0.5613	0.5055*	0.216**	1.6865***	1.2153***	0.9661**	0.2458	0.8149***
	(0.3008)	(0.6882)	(0.5591)	(0.3622)	(0.095)	(0.6307)	(0.3746)	(0.4886)	(0.2289)	(0.3045)
ARCH Sov. Yield	0.0602***	0.1122***	0.0778*	0.0653**	0.0542***	0.1068***	0.0691***	0.082***	0.0519***	0.0446***
	(0.0122)	(0.0227)	(0.0548)	(0.0353)	(0.0111)	(0.0358)	(0.0205)	(0.0255)	(0.0193)	(0.0127)
GARCH Sov. Yield	0.8828***	0.7452***	0.878***	0.9095***	0.9335***	0.7862***	0.8274***	0.861***	0.9203***	0.8888***
	(0.0246)	(0.0644)	(0.0852)	(0.0486)	(0.0155)	(0.0369)	(0.0415)	(0.0489)	(0.0406)	(0.0321)
Variance credit risk	0.0422*	0.1273**	0.072	0.007	0.0292	0.1555***	0.0702***	0.0886**	0.0314	0.0414***
	(0.0313)	(0.065)	(0.0985)	(0.0077)	(0.0241)	(0.0579)	(0.0236)	(0.0533)	(0.0254)	(0.0166)
Constant (OIS)	0.3814**	0.4171***	0.1774**	0.2905***	0.1759***	0.1819***	0.2764***	0.6288**	0.1922***	0.2797***
	(0.1871)	(0.1176)	(0.0876)	(0.1148)	(0.0616)	(0.0735)	(0.0796)	(0.3691)	(0.0631)	(0.0773)
ARCH OIS	0.0535***	0.0664***	0.0592***	0.0514***	0.0576***	0.0507***	0.0554***	0.077***	0.05***	0.0528***
	(0.0097)	(0.0118)	(0.0151)	(0.0127)	(0.0094)	(0.0097)	(0.0094)	(0.0212)	(0.0093)	(0.009)
GARCH OIS	0.9233***	0.9084***	0.9311***	0.9311***	0.9328***	0.9387***	0.9277***	0.8861***	0.938***	0.9297***
	(0.0187)	(0.0168)	(0.0176)	(0.0179)	(0.0113)	(0.0127)	(0.0125)	(0.0402)	(0.0118)	(0.0121)
Gamma	1.909	1.1364***	99.9897	0.9168***	21.3062***	13.8325	1.5617***	0.6684	2.5923***	2.0366***
	(6.5758)	(0.1727)	(252.0885)	(0.1626)	(8.4758)	(15.1717)	(0.2307)	(1.138)	(0.5225)	(0.3206)
Threshold c	16.2099	35.3935***	12.3282***	41.6891**	7.6958***	28.335***	24.6349***	58.9182	13.243***	18.4222***
	(18.1104)	(9.034)	(1.7731)	(18.4846)	(0.8476)	(2.0851)	(6.4454)	(50.3924)	(5.2972)	(4.8815)
Correlation Regime 1	1***	1***	0.8464***	1***	0.9354***	0.6048***	1***	0.9457***	1***	1***
	(0.279)	(0.0486)	(0.0283)	(0.0519)	(0.01)	(0.0468)	(0.0823)	(0.2777)	(0.1261)	(0.0884)
Correlation Regime 2	0.6347**	0.3213***	0.6704***	0.4531***	0.798***	0.1535***	0.1311**	0.5201*	0.1849***	0.2339***
	(0.3224)	(0.0866)	(0.0444)	(0.1229)	(0.0116)	(0.0394)	(0.0759)	(0.3798)	(0.0452)	(0.0548)
Log Likelihood	-13471	-13744	-9251	-13006	-13047	-15695	-14575	-12839	-15764	-14643

Table 7: Impact of the ECB Asset Purchase Programme on Conditional Correlations

This table shows the coefficients on the OMT and PSPP programmes proxied the former with announcement dummies and the latter wih Bloomberg news containing jointly the following keywords “Draghi, and QE or quantitative easing, and sovereign, and euro area”. The models also control for the FED news on asset purchases obtained collecting Bloomberg news containing jointly the following keywords “Bernanke or Yellen, and QE or quantitative easing, and US”. OLS is a pooled regression where the endogenous variable (the conditional correlations estimated either with DCC-GARCH or STCC-GARCHX) in first difference is regressed against the OMT dummy, the ECB and FED Bloomberg news, the lagged endogenus variable and the following regressors in first difference with their own respective lag: the US VIX, the three-month EURIBOR-OIS spread, the 10-year Greek sovereign-yield OIS spread, and the average quanto CDS spread of Italy and Spain relative to Germany. The VARX is the panel VAR described in equation (11) with in addition the OMT dummy and the ECB and FED Bloomberg news as exogenous variables. The panel VAR includes the following endoegenous variables with two lags: the conditional correlations (estimated either with DCC-GARCH or STCC-GARCHX), the US VIX, the three-month EURIBOR-OIS spread, the 10-year Greek sovereign-yield OIS spread, and the average quanto CDS spread of Italy and Spain relative to Germany. The STCC-GARCHX controls for credit risk in the mean and variance equations. The more vulnerable country’s group includes Ireland, Italy, Portugal and Spain. The less vulnerable country’s group includes Austria, Belgium, France, Germany and the Netherlands. The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. ***, ** and * indicate significant coefficients at the 1%, 5% and 10% levels, respectively.

	DCC	**STCC**	**STCCX**	**DCC**	**STCC**	**STCCX**
		not including US QE news			including US QE news	
			OLS			
PSPP	0.022***	0.020***	0.021***	0.022***	0.020***	0.021***
	(0.008)	(0.004)	(0.004)	(0.007)	(0.004)	(0.004)
OMT	0.047*	0.563**	0.499**	0.036*	0.553**	0.491**
	(0.508)	(0.279)	(0.238)	(0.508)	(0.279)	(0.238)
$adj.R^2$	0.001	0.063	0.059	0.001	0.063	0.059
			VARX			
PSPP	0.014*	0.026***	0.027***	0.014*	0.026***	0.027***
	(0.008)	(0.005)	(0.004)	(0.008)	(0.005)	(0.004)
OMT	0.208	0.667**	0.601**	0.159	0.618**	0.565**
	(0.507)	(0.279)	(0.238)	(0.312)	(0.280)	(0.239)
			VARX - more vulnerable country group			
PSPP	0.031**	0.044***	0.042***	0.031**	0.044***	0.040***
	(0.015)	(0.009)	(0.009)	(0.015)	(0.091)	(0.009)
OMT	-0.477	0.540	0.609	-0.511	0.486	0.552
	(0.908)	(0.544)	(0.513)	(0.909)	(0.544)	(0.513)
			VARX - less vulnerable country group			
PSPP	-0.006	0.013***	0.014***	-0.006	0.014***	0.014***
	(0.010)	(0.004)	(0.003)	(0.011)	(0.004)	(0.003)
OMT	0.326	0.981***	0.779***	0.247	0.950***	0.752***
	(0.639)	(0.210)	(0.189)	(0.640)	(0.210)	(0.289)

Figure 1: Sovereign Spreads and CDS Spreads

This figure shows the sovereign spreads at five-year maturity, calculated as the difference between the sovereign yield of a bond i and the OIS rate at time t at five-year maturity, the CDS spreads of a bond i at time t at five-year maturity and the respective orthogonalized CDS spreads estimated using $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1}\triangle BA_{i,t} + \phi_{i2}\triangle KfW_t + \phi_{i3}\triangle VSTOXX_t + \xi_{i,t}$, (see equation (3)), where $CDS_{i,t}$ denotes the CDS spread of a sovereign bond i at time t at five-year maturity, $BA_{i,t}$ the bid-ask spread of a sovereign bond i at time t at five-year maturity, $VSTOXX_t$ the implied volatility of option prices on EURO STOXX 50 at time t. The accompanying descriptive statistics are reported in Table 1. The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. The vertical bars denote 9 August 2007 (the interbank credit crisis), 15 September 2008 (Lehman), 16 October 2009 (the Greek fiscal crisis), 8 December 2011 (three-year LTROs), 26 July 2012 (Draghi's speech) and 11 June 2014 (negative deposit facility rate and additional non-standard measures).

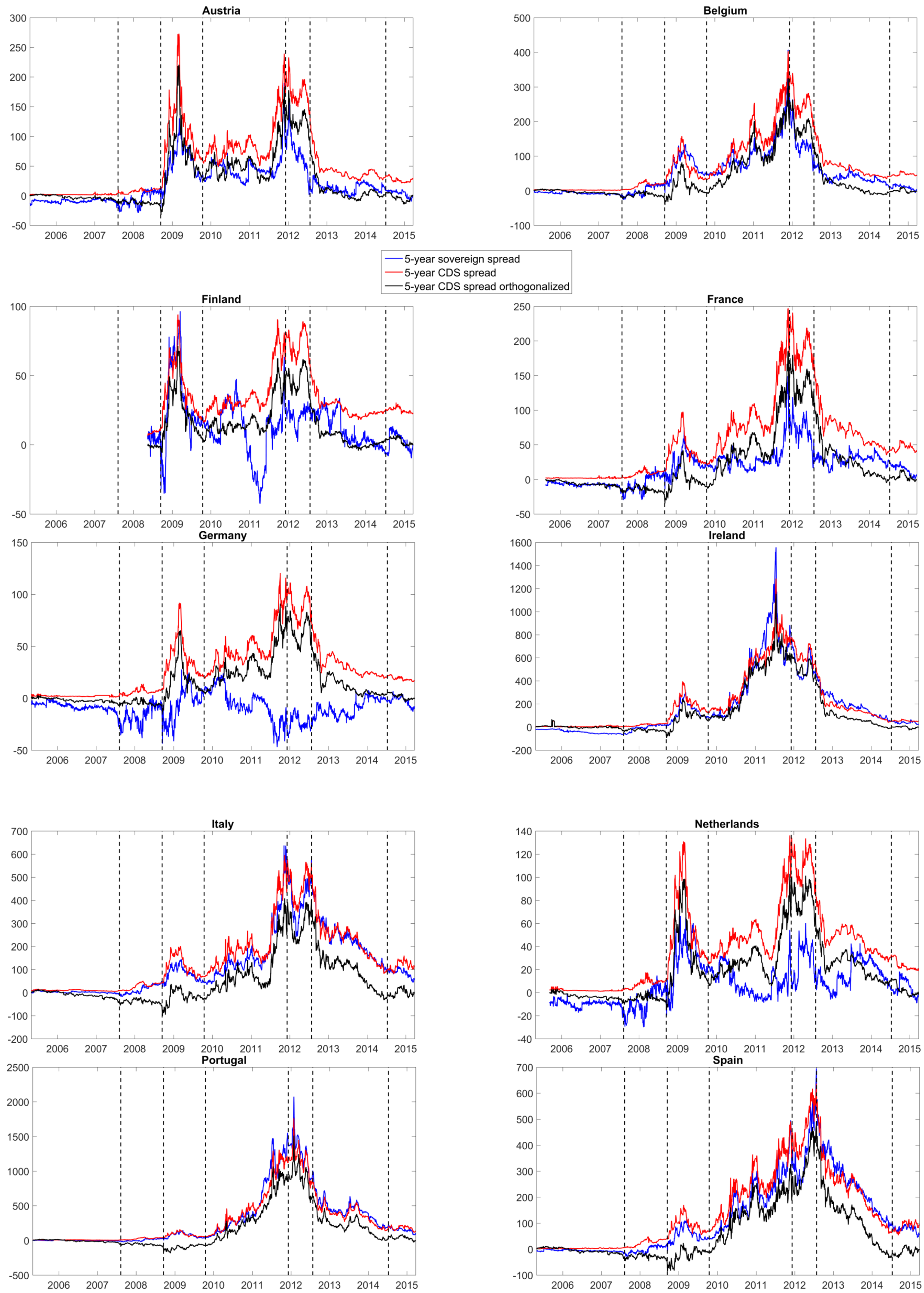


Figure 2: Economic Variables

This figure shows the OIS rate at time t at five-year maturity, the KfW-Bund spread at time t at five-year maturity, the implied volatility of option prices on the EURO STOXX 50 index (VSTOXX) at time t, and the implied volatility of option prices on the S&P 500 index (VIX) at time t. Zero-coupon yield curves for bonds issued by KfW and the German government are estimated using the Merrill Lynch exponential spline (MLES) model. The accompanying descriptive statistics are reported in Table 1. The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. The vertical bars denote 9 August 2007 (the interbank credit crisis), 15 September 2008 (Lehman), 16 October 2009 (the Greek fiscal crisis), 8 December 2011 (three-year LTROs), 26 July 2012 (Draghi's speech) and 11 June 2014 (negative deposit facility rate and additional non-standard measures).

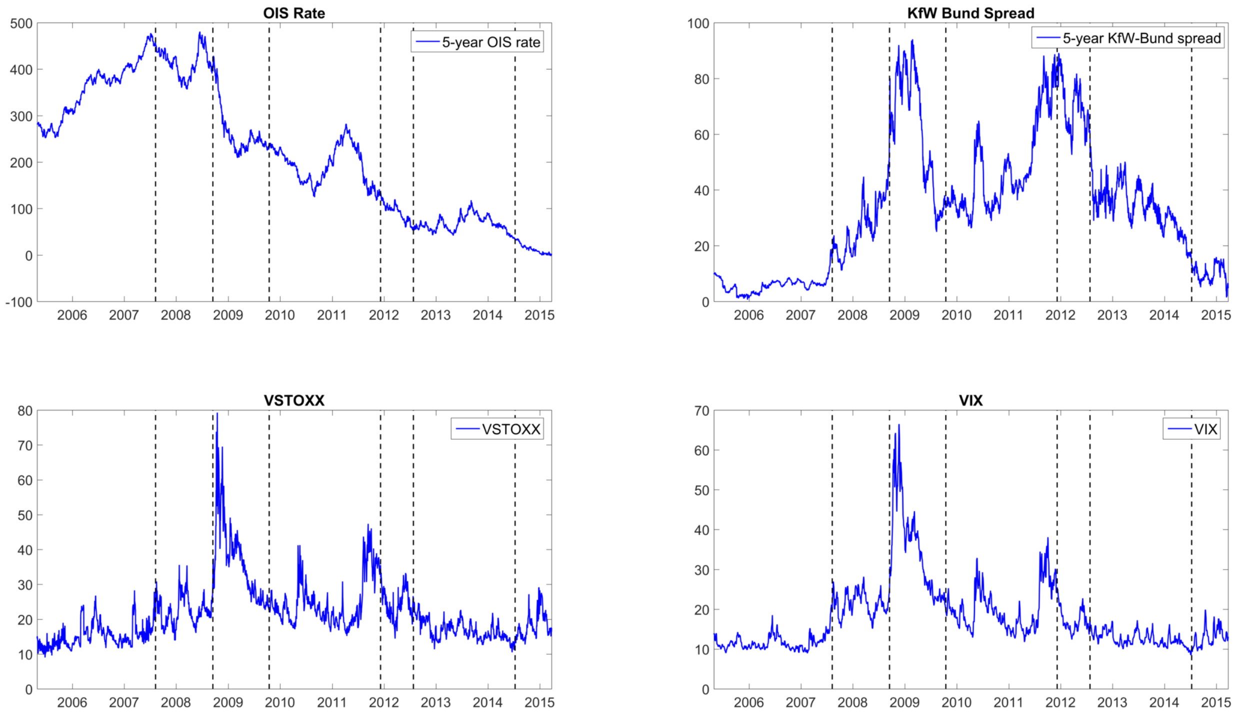


Figure 3: Conditional Correlation: DCC-GARCH versus STCC-GARCHX

This figure shows the correlations estimated with DCC-GARCH and STCC-GARCHX models between the change in sovereign yield of bond i and the change in the OIS rate at time t at five-year maturity. The KfW-Bund spread at five-year maturity at $t-1$ is used as transition variable in the STCC-GARCHX and $\hat{\xi}^2_{i,t-1}$ is used as an exogenous variable in the respective countries' variance equations (see Table 6 for parameter estimates). The estimation of the mean equation is based on $\triangle r_{i,t} = \beta_{i0} + \beta_{i1}\triangle r_{i,t-1} + \beta_{i2}\hat{\xi}_{i,t} + \varepsilon_{i,t}$, (see equation (4)). $\hat{\xi}_{i,t}$ denotes credit risk obtained as the residuals of the following regression $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1}\triangle BA_{i,t} + \phi_{i2}\triangle KfW_t + \phi_{i3}\triangle VSTOXX_t + \xi_{i,t}$ (see equation (3)). The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. The vertical bars denote 9 August 2007 (the interbank credit crisis), 15 September 2008 (Lehman), 16 October 2009 (the Greek fiscal crisis), 8 December 2011 (three-year LTROs), 26 July 2012 (Draghi's speech) and 11 June 2014 (negative deposit facility rate and additional non-standard measures).

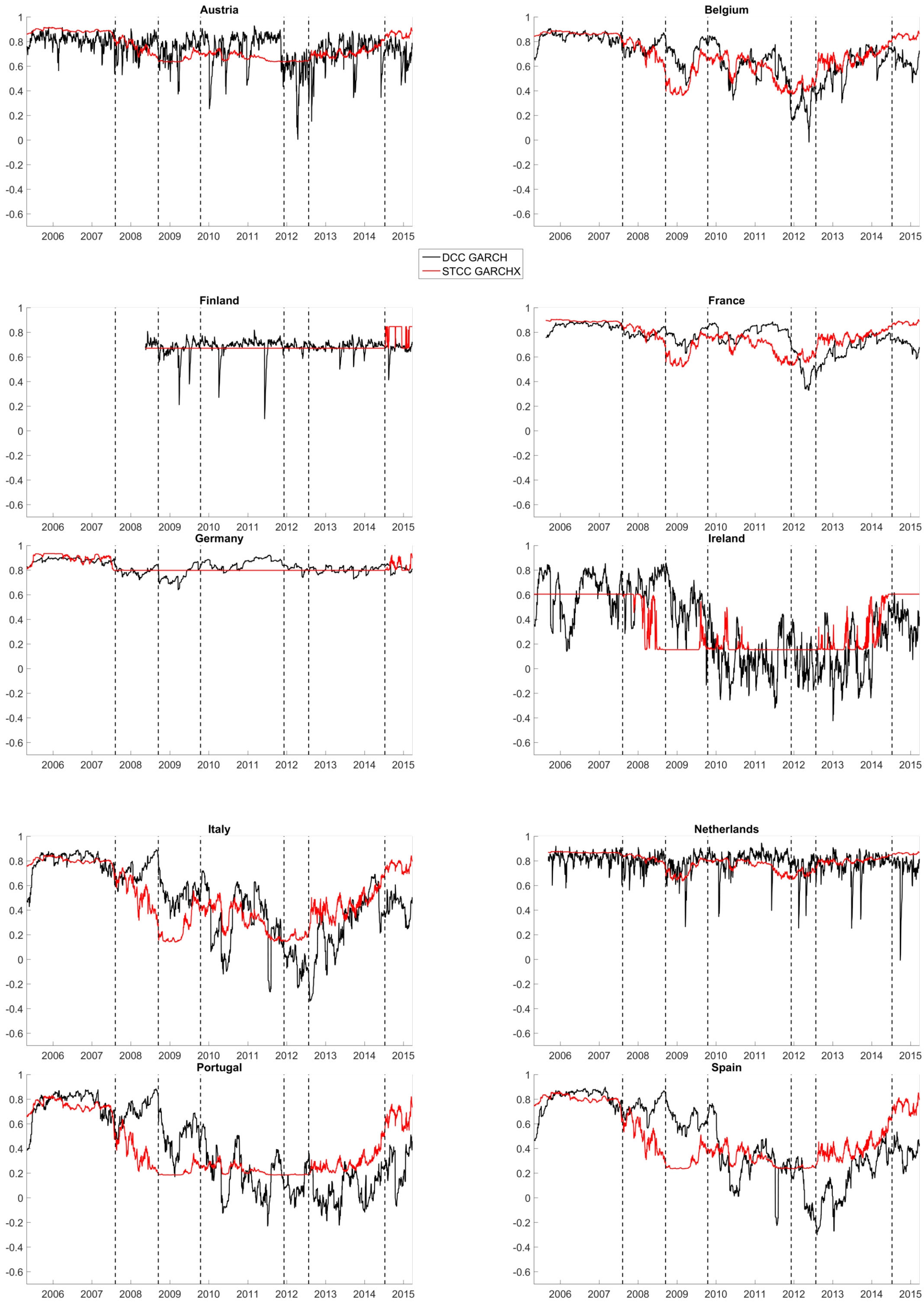


Figure 4: Conditional Volatility: DCC-GARCH versus STCC-GARCHX

This figure shows the conditional variance in basis points estimates obtained using DCC-GARCH and STCC-GARCHX for the change in the sovereign yield of a bond i at time t at five-year maturity. The KfW-Bund spread at five-year maturity at $t-1$ is used as a transition variable in the STCC-GARCHX and $\hat{\xi}^2_{i,t-1}$ is used as an exogenous variable in the respective countries' variance equations (see Table 6 for parameter estimates). The estimation of the mean equations are based on $\triangle r_{i,t} = \beta_{i0} + \beta_{i1}\triangle r_{i,t-1} + \beta_{i2}\hat{\xi}_{i,t} + \varepsilon_{i,t}$, (see equation (4)). $\hat{\xi}_{i,t}$ denotes credit risk obtained as the residuals of the following regression $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1}\triangle BA_{i,t} + \phi_{i2}\triangle KfW_t + \phi_{i3}\triangle VSTOXX_t + \xi_{i,t}$ (see equation (3)). The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. The vertical bars denote 9 August 2007 (the interbank credit crisis), 15 September 2008 (Lehman), 16 October 2009 (the Greek fiscal crisis), 8 December 2011 (three-year LTROs), 26 July 2012 (Draghi's speech) and 11 June 2014 (negative deposit facility rate and additional non-standard measures).

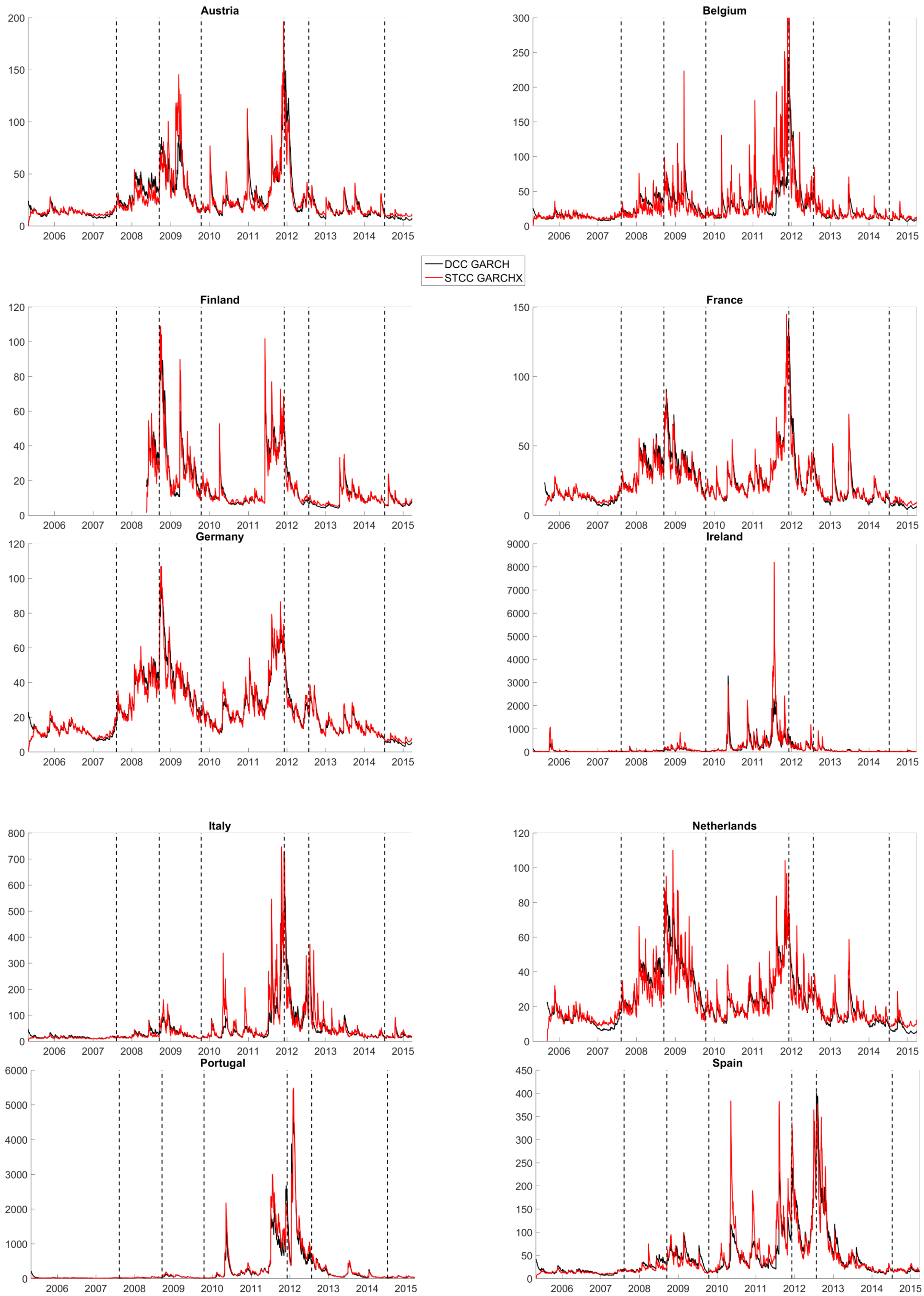


Figure 5: Macro Factors and Correlation Breakdown: STCC-GARCHX

This figure shows (i) the three-month EURIBOR-OIS spread (top-left), (ii) the 10-year Greek sovereign yield-OIS spread (middle-left) and (iii) the average quanto CDS spread of Italy and Spain relative to Germany (bottom-left) and their respective IRFs on the conditional correlations obtained using a panel VAR (see equation (11)), which includes the US VIX, the estimated conditional correlation with the STCC-GARCHX model, the three-month EURIBOR-OIS spread, the 10-year Greek sovereign yield-OIS spread, and the average quanto CDS spread of Italy and Spain relative to Germany. The identification restriction is recursive and the variables are ordered as described above. The mean equation of the STCC-GARCHX model is $\triangle r_{i,t} = \beta_{i0} + \beta_{i1}\triangle r_{i,t-1} + \beta_{i2}\hat{\xi}_{i,t} + \varepsilon_{i,t}$, (see equation (4)). $\hat{\xi}_{i,t}$ denotes credit risk obtained as the residuals of the following regression $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1}\triangle BA_{i,t} + \phi_{i2}\triangle KfW_t + \phi_{i3}\triangle VSTOXX_t + \xi_{i,t}$ (see equation (3)). The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability. The vertical bars denote 9 August 2007 (the interbank credit crisis), 15 September 2008 (Lehman), 16 October 2009 (the Greek fiscal crisis), 8 December 2011 (three-year LTROs), 26 July 2012 (Draghi's speech) and 11 June 2014 (negative deposit facility rate and additional non-standard measures).

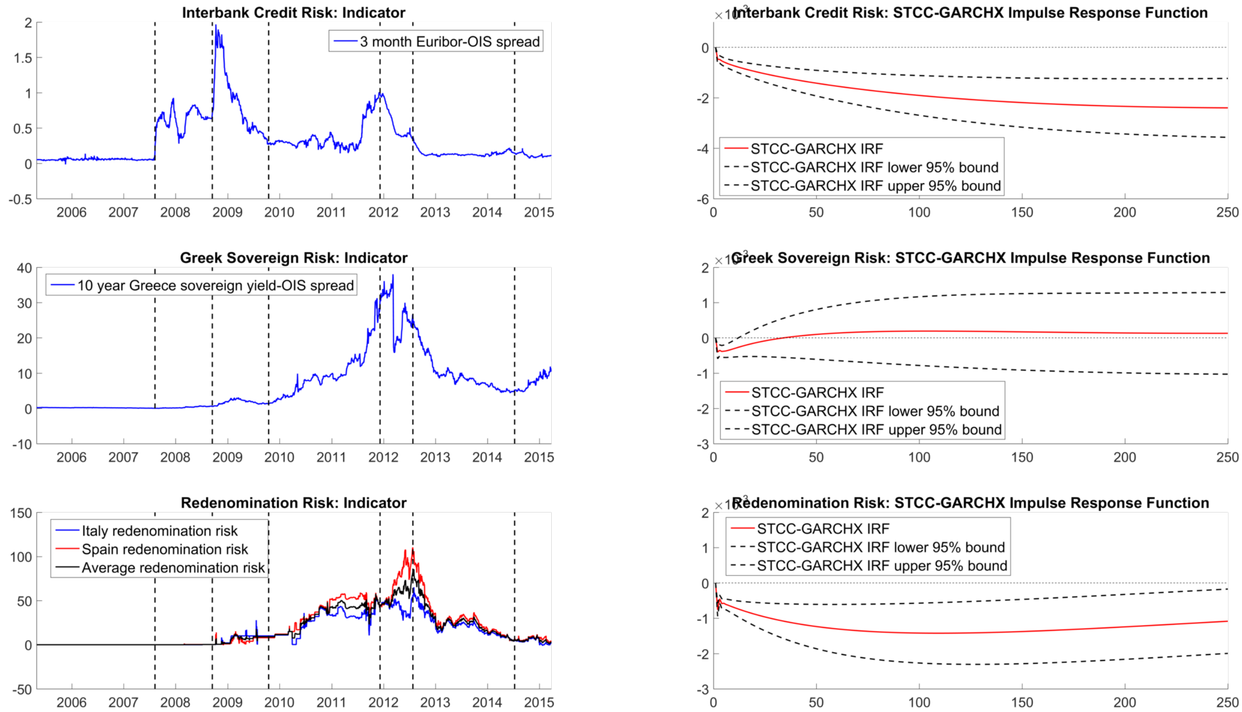


Figure 6: Macro Factors and Correlation Breakdown: Before and After October 2011

This figure shows the IRFs of (i) the three-month EURIBOR-OIS spread (top), (ii) the 10-year Greek sovereign yield-OIS spread (middle) and (iii) the average quanto CDS spread of Italy and Spain relative to Germany (bottom) on the conditional correlations obtained using a panel VAR (see equation (11)), which includes the US VIX, the estimated conditional correlation with the STCC-GARCHX model, the three-month EURIBOR-OIS spread, the 10-year Greek sovereign yield-OIS spread, and the average quanto CDS spread of Italy and Spain relative to Germany. The identification restriction is recursive and the variables are ordered as described above. The panel VAR is estimated over the period 1 August 2007 - 30 September 2011 (left) and over the period 1 October 2011 - 25 March 2015 (right). The mean equation of the STCC-GARCHX model is $\triangle r_{i,t} = \beta_{i0} + \beta_{i1}\triangle r_{i,t-1} + \beta_{i2}\hat{\xi}_{i,t} + \varepsilon_{i,t}$, (see equation (4)). $\hat{\xi}_{i,t}$ denotes credit risk obtained as the residuals of the following regression $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1}\triangle BA_{i,t} + \phi_{i2}\triangle KfW_t + \phi_{i3}\triangle VSTOXX_t + \xi_{i,t}$ (see equation (3)). The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability.

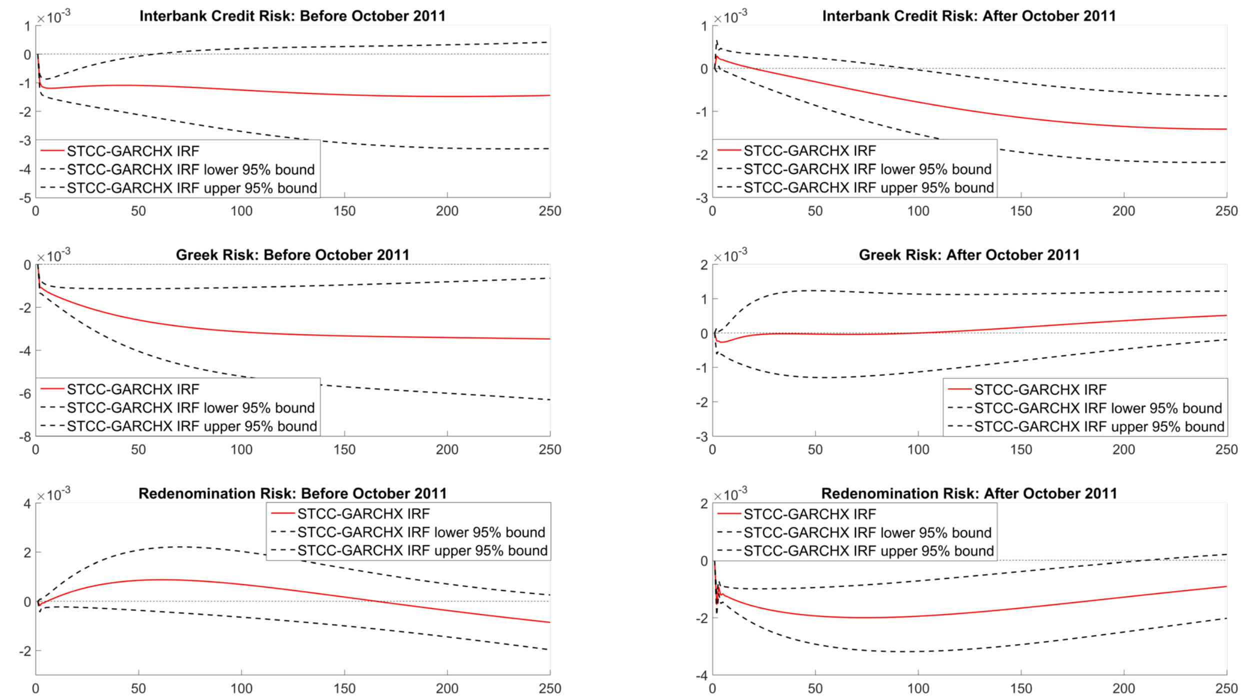


Figure 7: Macro Factors and Correlation Breakdown: More Vulnerable Country Group

This figure shows the IRFs of (i) the three-month EURIBOR-OIS spread (top), (ii) the 10-year Greek sovereign yield-OIS spread (middle) and (iii) the average quanto CDS spread of Italy and Spain relative to Germany (bottom) on the conditional correlations obtained using a panel VAR for Ireland, Italy, Portugal and Spain (see equation (11)), which includes the US VIX, the estimated conditional correlation with the STCC-GARCHX model, the three-month EURIBOR-OIS spread, the 10-year Greek sovereign yield-OIS spread, and the average quanto CDS spread of Italy and Spain relative to Germany. The identification restriction is recursive and the variables are ordered as described above. The panel VAR is estimated over the period 1 August 2007 - 30 September 2011 (left) and over the period 1 October 2011 - 25 March 2015 (right). The mean equation of the STCC-GARCHX model is $\triangle r_{i,t} = \beta_{i0} + \beta_{i1} \triangle r_{i,t-1} + \beta_{i2} \hat{\xi}_{i,t} + \varepsilon_{i,t}$, (see equation (4)). $\hat{\xi}_{i,t}$ denotes credit risk obtained as the residuals of the following regression $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1} \triangle BA_{i,t} + \phi_{i2} \triangle KfW_t + \phi_{i3} \triangle VSTOXX_t + \xi_{i,t}$ (see equation (3)). The more vulnerable countries include Ireland, Italy, Portugal and Spain. The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability.

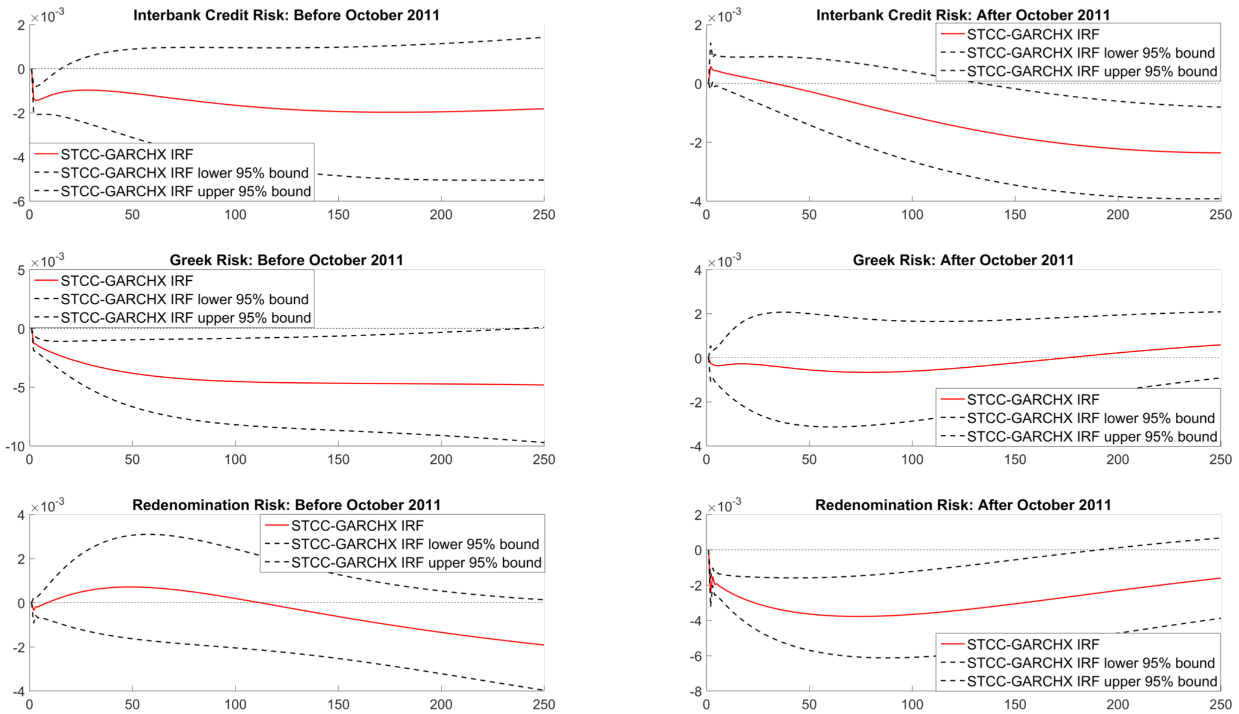
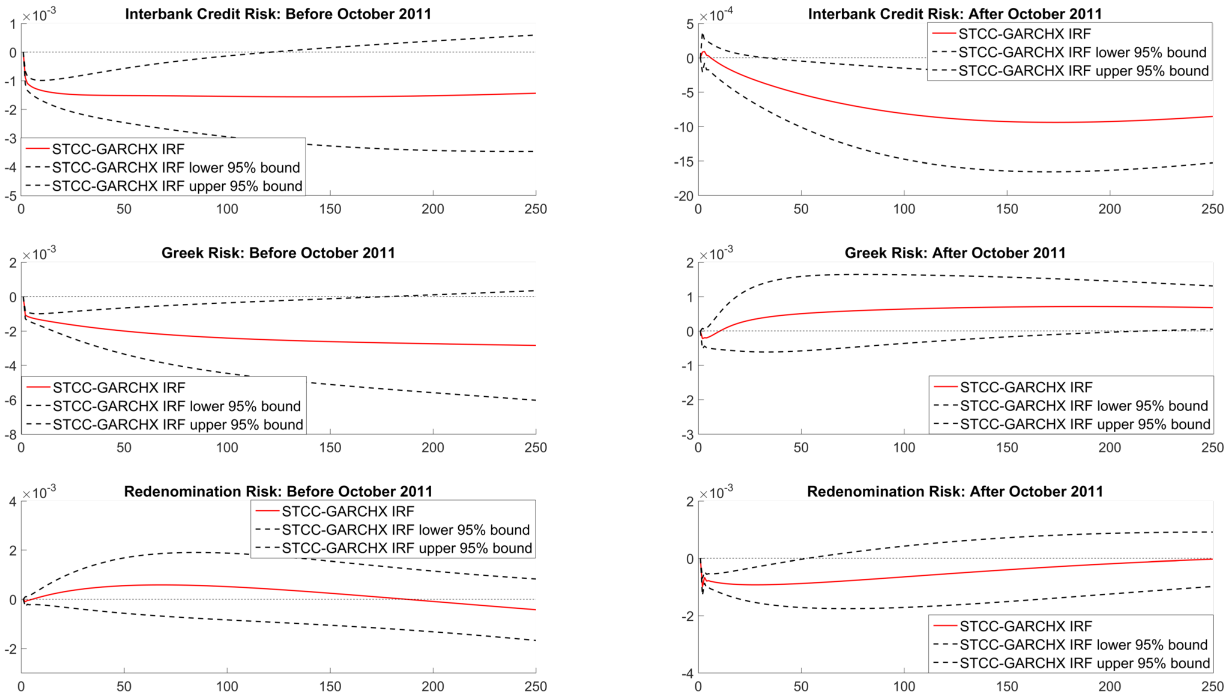


Figure 8: Macro Factors and Correlation Breakdown: Less Vulnerable Country Group

This figure shows the IRFs of (i) the three-month EURIBOR-OIS spread (top), (ii) the 10-year Greek sovereign yield-OIS spread (middle) and (iii) the average quanto CDS spread of Italy and Spain relative to Germany (bottom) on the conditional correlations obtained using a panel VAR for all countries except Ireland, Italy, Portugal and Spain (see equation (11)), which includes the US VIX, the estimated conditional correlation with the STCC-GARCHX model, the three-month EURIBOR-OIS spread, the 10-year Greek sovereign yield-OIS spread, and the average quanto CDS spread of Italy and Spain relative to Germany. The identification restriction is recursive and the variables are ordered as described above. The panel VAR is estimated over the period 1 August 2007 - 30 September 2011 (left) and over the period 1 October 2011 - 25 March 2015 (right). The mean equation of the STCC-GARCHX model is $\triangle r_{i,t} = \beta_{i0} + \beta_{i1}\triangle r_{i,t-1} + \beta_{i2}\hat{\xi}_{i,t} + \varepsilon_{i,t}$, (see equation (4)). $\hat{\xi}_{i,t}$ denotes credit risk obtained as the residuals of the following regression $\triangle CDS_{i,t} = \phi_{i0} + \phi_{i1}\triangle BA_{i,t} + \phi_{i2}\triangle KfW_t + \phi_{i3}\triangle VSTOXX_t + \xi_{i,t}$ (see equation (3)). The less vulnerable countries include Austria, Belgium, Finland, France, Germany and the Netherlands. The daily frequency sample period is 28 April 2005 - 25 March 2015. The initial sample period differs for Finland (15 May 2008), France (17 August 2005) and the Netherlands (8 September 2005) due to data availability.



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Roberto De Santis
European Central Bank, Directorate General Economics, Monetary Policy, Capital Markets and Financial Structures Division, Frankfurt am Main, Germany; email: roberto.de_santis@ecb.europa.eu

Michael Stein (Corresponding author)
University of Duisburg-Essen, Faculty for Economics and Business Administration,
Department of Financial Market Econometrics, Essen, Germany; email: michael.stein@steinpage.com

Postal address 60640 Frankfurt am Main, Germany
Telephone +49 69 1344 0
Website www.ecb.europa.eu

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