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Sovereign bond and CDS market contagion: A story from the Eurozone crisis*

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Abstract

We examine the asymmetric and nonlinear nature of the cross- and intra-market linkages of eleven EMU sovereign bond and CDS markets during 2006-2018. By adopting the excess correlation concept of [Bekaert *et al.* \(2005\)](#) and the local Gaussian correlation approach of [Tjøstheim and Hufthammer \(2013\)](#), we find that contagion phenomena occurred during two major phases. The first, extends from late 2009 to mid 2011 and concerns the outright contagion transmission from EMU South bond markets towards all European CDS markets. The second, is during the revived fears of a Greek exit in November 2011 and is characterized by contagion from (i) CDS spreads in the EMU South towards bond yields in the same bloc and Belgium, and (ii) from Italian and Spanish CDS spreads towards all European CDS spreads. Consistent with their “too big to bail out” status, Italy and Spain emerge as pivotal for the evolution of sovereign credit risk across the Eurozone. Our examination of the relevant mechanisms, highlights the importance of credit risk over liquidity risk, and the containment effect of the naked CDS ban.

Keywords: sovereign bond market, sovereign CDS market, nonlinear dependence, contagion, local Gaussian correlation

JEL Classification: G01; G14; G15; C1; C58

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“Contagion occurs when cross-country correlations increase during crisis times relative to correlations during tranquil times”. — World Bank

1 Introduction

Contagion emerges in times of crisis and was a prominent feature of the European sovereign debt crisis. Recent evidence suggests that negative shocks were diffused in a different way across EMU member states (Claeys and Vařicek, 2014; Broto and Pérez-Quirós, 2015; Caporin *et al.*, 2018); countries with higher debt/deficits (Greece, Italy, Ireland, Portugal and Spain, henceforth GIIPS) were immediately affected, whereas the direct impact on the rest of the European economies has appeared less severe.¹ This issue has revived the discussion on the transmission of shocks and contagion in the euro area, which has led to ambiguous conclusions (see, e.g., Metiu, 2012; De Santis, 2012; Beirne and Fratzscher, 2013; Blatt *et al.*, 2015; Caporin *et al.*, 2018). This paper considers all relevant adverse financial and economic shocks stemming from the GIIPS during the Eurozone crisis and analyzes their impact on the intra- and cross-market linkages of the European sovereign bonds and CDS markets. Having established that, we examine whether these linkages can generate contagion and whether this contagion is synchronous across different EMU countries.

Testing for contagion is not a straightforward exercise, as there is no broad consensus on what exactly constitutes contagion.² In this study, we adopt the approach of Bekaert *et al.* (2014) and use a factor model to examine unexpected factor exposures with respect to changes in fundamentals. Following Bekaert *et al.* (2005) we employ the residuals of the factor model to test for “residual contagion”, where contagion is defined as ‘correlation between markets in excess of that implied by economic fundamentals’. We complement this with a measure of local Gaussian correlation by Tjøstheim and Hufthammer (2013); this enables us to examine asymmetric and nonlinear changes in the dependence structure and test for contagion effects transmitted from the GIIPS to the entire Eurozone. We identify the dates where these effects occurred and provide a timeline of all the events that triggered

¹An additional classification of EMU (Economic and Monetary Union) countries is that between the EMU South (Greece, Ireland, Italy, Portugal, Spain) and the EMU North (Austria, Belgium, Finland, France, Germany, the Netherlands). Instead of the terms “EMU South” and “EMU North” the terms “EMU periphery” and “EMU core” respectively are also common. Throughout this paper these terms are used interchangeably.

²See King and Wadhvani (1990); Bekaert and Hodrick (1992); Forbes and Rigobon (2002); Bekaert *et al.* (2005); reviews are provided by Pericoli and Sbracia (2003), Dungey *et al.* (2005) and Forbes (2012).

financial contagion during the Eurozone crisis. We also reveal the direction of this contagion and the counterparties affected.

Our dataset includes 5-year sovereign bond yields and CDS spreads covering the period from January 2006 to April 2018. In the first step, our analysis employs a factor model, where the bond and CDS data are conditioned on certain state variables (see [Bekaert *et al.*, 2005](#); [De Bruyckere *et al.*, 2013](#); [Bekaert *et al.*, 2014](#); [Fontana and Scheicher, 2016](#)). We employ these variables within a Kalman filter model and calculate the residuals of each of our bond and CDS series. By estimating time-varying coefficients from the entire sample, we ensure that any findings of contagion are not erroneously attributed to higher correlations due to volatility bias.

In the second step, we employ the method of [Tjøstheim and Hufthammer \(2013\)](#) to study the local Gaussian correlation dynamics of the residual bond and CDS series. We adopt the approach of [Demetrescu and Wied \(2019\)](#) to detect endogenous break-points in the correlation structure of our series and identify the dates of changes in their intra- and cross-asset (inter)dependence. The next step is to estimate the local Gaussian correlation for each country-pair around the break-date and test for contagion effects via a bootstrap test (for an application, see [Støve *et al.*, 2014](#)). If the local correlation has increased significantly after the break-date, this is evidence of “pure contagion”, i.e., contagion over and above what one would expect from economic fundamentals.

We subsequently replicate the analysis for our initial bond and CDS data without employing the first-stage factor model. By contrasting the estimates from the analysis of the residuals with those from the analysis of the original series, we can further distinguish between cases of “pure contagion” and cases where markets react to changes in economic fundamentals (“wake-up call”) or short-lived contagion episodes (“limited contagion”).

Our results indicate that contagion is neither a single-source nor a single-market phenomenon in the context of the European sovereign debt crisis. “Pure contagion” phenomena in the sovereign bond and CDS markets have undergone two major phases. The first phase extends from late 2009 to mid 2011 and concerns the outright transmission of contagion from the sovereign bond markets of the EMU South towards the CDS markets of both EMU blocs. Most correlation break-points that indicate contagion are concentrated around important eco-

conomic events. These include the Greek deficit's upward revision in November 2009 and the consequent fears about a possible Greek default, which led to the Greek government's official request of its first rescue plan in April 2010. Other events include the agreements on the rescue plans of Ireland and Portugal (November 2010 and April 2011 respectively) and the negative economic developments in Italy and Spain (July-August 2011).

On the other hand, the European sovereign bond market appears immune to extreme negative developments in the periphery countries' bond yields (a finding partly documented in [Missio and Watzka, 2011](#); [Metiu, 2012](#); [De Santis, 2012](#); [Beirne and Fratzscher, 2013](#); [Claeys and Vařicek, 2014](#); [Caporin *et al.*, 2018](#)). There is opposite responsiveness between the bond and CDS markets with regards to the periphery bond-stemming contagion, revealing an additional difference between the two markets during that period: the capacity of the bond market to reduce shocks and/or losses. This is in contrast to the CDS market where such shocks – reflected by the higher yields in the periphery bonds – are preserved and even amplified. The analysis of the original series reveals that European sovereign bond yields were subject to a “wake-up call” during the late 2008–early 2009 period. This prompted investors to reassess the vulnerability of Eurozone countries, thereby leading to a repricing of the factors that affect sovereign credit risk ([Bekaert *et al.*, 2014](#); [Claeys and Vařicek, 2014](#)).

The second phase is around November 2011 and Greece's plan to hold a referendum on Eurozone membership. This prolonged political uncertainty has revived the fears of the country's Eurozone exit. During this period, CDS-bond contagion is mainly a “periphery” phenomenon: contagion stems from the CDS markets of the EMU South and is directed to the bond markets of the same bloc and Belgium. When intra-CDS contagion is concerned, we find evidence of outright contagion from the Italian and Spanish CDS markets towards the CDS markets of both blocs. The outbreak of CDS-stemming contagion during the respective period marks the development of the Greek debt crisis into a European debt crisis. In this regard, Italy and Spain appear to be the key countries for the evolution of euro area sovereign credit risk. This finding is consistent with their “too big to bail out” status and the fears that a Greek default would cause a domino effect across the Eurozone. The exposure of euro area CDS spreads to Italian and Spanish CDS spreads stands in contrast to previous findings within the intra-CDS context, where contagion was found to be either non-existent

(see Caporin *et al.*, 2018) or only a Eurozone periphery phenomenon (see Broto and Pérez-Quirós, 2015) and also to arguments about the limited capacity of the periphery CDS markets to generate contagion (see Kalbaska and Gatkowski, 2012).

We complement the above studies by offering a timeline of the magnitude and direction of contagion within the European sovereign debt and credit markets. Contagion is decomposed into its “pure” and “wake-up call” components. We show that the two Eurozone blocs were following a divergent path; the fiscal shocks across the EMU South were the driving cause of such divergence. On the contrary, the notion that contagion phenomena in the Eurozone were a consequence of the US financial meltdown and the resulting global financial crisis is rather weak. The occurrence of the vast majority of “pure contagion” episodes in the late 2009–late 2011 period confirms these arguments.

Importantly, we identify the potential mechanisms for the emergence of contagion. As such, we focus on liquidity risk, which was an important determinant of European sovereign bond yields during the Eurozone crisis (see Beber *et al.*, 2009; Brunnermeier and Pedersen, 2009; Monfort and Renne, 2014). By distinguishing between contagion transmission due to credit risk considerations and due to liquidity risk considerations, we show that liquidity risk is an important source of risk in the European CDS market, particularly when stemming from the bond market. We thus content that liquidity risk is an essential determinant of price discovery and contagion transmission between the two markets, complementing credit risk.

On the same line, we examine the role of arbitrage opportunities. Our findings reveal that although arbitrage forces may be present (see Fontana and Scheicher, 2016; Gyntelberg *et al.*, 2018), they were not able to fully close the pricing gaps between the two markets and affect contagion dynamics. We thus demonstrate that the different regimes and adjustment speeds which characterize the correcting mechanisms between the sovereign bond and CDS markets (also evident in Gyntelberg *et al.*, 2018), further restrict their contagion capacity.

Moreover, our study is the first to link the important policy changes on the regulatory and monetary policy fronts during the Eurozone crisis with sovereign contagion dynamics. We provide evidence that the agreement on the naked CDS ban exerted an easing effect on the transmission of CDS-bond contagion within the EMU South. We further consider the ECB’s large-scale monetary policy interventions (in the form of government bond purchases and

one-off liquidity injections) and find no sign of bond contagion either before or after their conduct; if anything, these interventions aimed at easing periphery countries' borrowing costs and restoring monetary policy transmission, rather than targeting contagion.

Finally, our work relates to the literature on tail market events and financial contagion. For this purpose, several estimation and testing strategies have been applied, including the analysis of the coincidence of tail returns (Bae *et al.*, 2003), the multivariate distribution modeling of tail returns (Longin and Solnik, 2001), the estimation of the expected market crashes after a crash in one market (Hartmann *et al.*, 2004) or the exceedance correlation concept (Ang and Bekaert, 2002; Ang and Chen, 2002; Hong *et al.*, 2007). The latter refers to the concept of conditional correlation, where the ordinary product-moment correlation is computed for certain regions of the distribution. However, the conditional correlation in a local region is not equal to the global correlation for a pair of jointly Gaussian variables.³ Moreover, this approach produces a measure of linear dependence locally, which is questionable in a nonlinear and non-Gaussian framework.

In this paper we use the local Gaussian correlation concept to assess the nonlinear dependence behavior of our series. The local Gaussian correlation does not suffer from the bias problem of the conditional correlation (e.g., as in the exceedance correlation concept), while the latter can be considered as a special case of the former. More closely related to our approach are the recent proposals for measuring and testing nonparametric dependence based on auto-distance correlation functions (Székely and Rizzo, 2009; Zhou, 2012), copula models (Oh and Patton, 2017) and the cross-quantilogram of Han *et al.* (2016). However, these test dependence under the assumption of independent and identically distributed series or estimate time series serial dependence. Little has been done in the perhaps most important situation of estimating the strength of cross-dependence, and testing for contagion between two general stationary time series. In this respect, the local Gaussian correlation is easier to interpret than auto-distance and copulas, providing a direct measure of both average and upper-lower tail dependence and a complete characterization of dependence structure.

The remaining of the paper is organised as follows: Sections 2 and 3 present the empirical methodologies and Section 4 describes the dataset. Section 5 examines contagion

³This is known as the bias problem for conditional correlation (see Forbes and Rigobon, 2002) and refers to the fact that in a Gaussian distribution, dependence is completely characterized by the correlation coefficient.

transmission from the sovereign bond and CDS markets of the EMU South, while Section 6 identifies the mechanisms and assesses the sensitivity of our findings. Section 7 concludes.

2 Methodology

The concept of contagion is not unequivocal, given the number of alternative methodological approaches. In this paper, we take an asset pricing perspective on measuring economic fundamentals and identify contagion through the correlation of an asset pricing model’s residuals (as in [Bekaert *et al.*, 2014](#)). These residuals are obtained from a regression, in which the dependent variable is the daily change in the bond yield (CDS spread) of a given country and the explanatory variables are six factors related to the bond and CDS markets. In this respect, our six-factor model encompasses different EMU-, global financial-, and US-specific factors. Given this, we set a benchmark for what sovereign bond and CDS market comovements should be, based on these fundamentals.

We use Kalman filtering to estimate time-varying coefficients for our factor model. We calculate the standardised signal errors from the Kalman filter (the residuals from the factor model) and test for increasing correlations between the residual bond/CDS series by using the local correlation approach. We endogenously detect the break-dates in correlation after considering changes in the variance of the residual series following the method of [Demetrescu and Wied \(2019\)](#). Subsequently, we test for contagion via a bootstrap test for each intra-market pair (bond-bond or CDS-CDS pair) and cross-market pair (bond-CDS or CDS-bond pair) that we observe an increase in local Gaussian correlation after the correlation break-date. Effectively, our approach is in conceptual proximity with the “excess correlation” concept of [Bekaert *et al.* \(2005\)](#), who define contagion as “correlation over and above what one would expect from economic fundamentals”. Finally, we perform the correlation break test and the bootstrap test for each possible pair from our original series (i.e., without accounting for fundamentals through our factor model). This leads to three different cases:

- an increase in the exposure of bond yields and/or CDS spreads to common factors, which is labelled as “wake-up call”. Practically, this occurs when there is evidence of increased local correlation on the original series. The hypothesis of “wake-up call” between a given country-pair is verified for the original series but rejected for the residual series.

- an increase in the exposure of bond yields and/or CDS spreads to common factors, which is further accompanied by an increase in the correlation across and/or between unexplained bond yields and CDS spreads. This is labelled as “pure contagion” and occurs when there is evidence of increased local correlation on both the original and the residual series. Accordingly, the hypothesis of contagion is confirmed for both the original and the residual series. Generally, such cases refer to potent and persistent shocks, whose impact is not limited to economic fundamentals, but goes beyond them.

- an increase in the correlation across and/or between unexplained bond yields and CDS spreads without a corresponding increase in the exposure of bond yields and/or CDS spreads to common factors. This is labelled as “limited contagion” and occurs when there is evidence of increased local correlation on the residual series but not on the original ones. The hypothesis of contagion for a given pair is confirmed by the estimation of the residual series, but rejected by that of the original series. This form of contagion generally refers to short-lived shocks that quickly recede: due to their magnitude they are reflected on the unexplained series, however their temporary nature is not reflected on economic fundamentals.

2.1 The factor model

We employ a six-factor model that controls for EMU-level developments, global risk aversion and U.S. conditions. Initially, we control for liquidity risk, as prior studies document the existence of a liquidity premium in sovereign bond and CDS prices (Favero *et al.*, 2010; Antón *et al.*, 2018). Our proxy for liquidity risk in the bond (CDS) market is the quoted bid-ask spread (*BidAsk*) on the 5-year sovereign bond (CDS);⁴ we include this as a standard factor in the regression for the country’s bond yields (CDS spreads). Our EMU-specific factor concerns the idiosyncratic equity returns in each country (*EqR*), to control for market-wide changes in business climate. This is defined as the difference between the country’s equity-index return and the benchmark Eurozone index return (STOXX Europe 600). We expect equity returns in each country to be negatively related to that country’s credit spreads.

Global risk factors constitute an additional driver of sovereign credit spreads, as higher volatility is associated with higher economic uncertainty (Longstaff *et al.*, 2011). We control

⁴The quoted bid-ask spread of the bond (CDS) is equal to $(Ask - Bid)/(0.5(Ask + Bid))$, where *Ask* and *Bid* are the ask price and bid price respectively of the bond (CDS); see Goyenko and Ukhov (2009).

for systemic risk with the 3-month Euribor-OIS spread ($3mEUR$). We employ this variable (in logarithmic form), as it is considered a reliable measure of systemic risk (see [Rodriguez et al., 2013](#)). Since systemic credit risk has become a considerably large fraction of total credit risk (for evidence from the subprime crisis, see [Bhansali et al., 2008](#)), we expect that an increase in systemic risk is reflected in rising bond yields and CDS spreads. We further control for global uncertainty with the VIX index (VIX). The choice of this variable (in logarithmic form), owes to its strong relationship with sovereign credit risk ([Pan and Singleton, 2008](#)). The fact that the VIX is highly correlated with spreads on sovereign entities, makes it a proxy for global “event risk” in credit markets. Given this, we expect a positive relationship between the VIX and sovereign CDS spreads.

Turning to US-specific factors, we control for exchange rate uncertainty, with the 30-day implied EURO/USD exchange rate volatility index (EVZ).⁵ In the event of higher uncertainty about the future path of the bilateral exchange rate, USD-quoted protection should be more expensive than the equivalent euro-quoted protection. This is due to the currency hedge provided by the USD-quoted protection against a potential sovereign default and a consequent euro depreciation; this indicates a positive correlation between bilateral exchange rate volatility and CDS spreads ([Fontana and Scheicher, 2016](#)). Finally, we control for overall economic conditions in the U.S. by employing the S&P500 index ($SP500$).⁶

The regression specification of the factor model takes the following form:

$$\Delta Y_{it} = \beta_{0,t} + \beta_{1,t} \Delta BidAsk_{it} + \beta_{2,t} \Delta EqR_{it} + \beta_{3,t} \Delta 3mEur_t + \beta_{4,t} \Delta VIX_t + \beta_{5,t} \Delta EVZ_t + \beta_{6,t} \Delta SP500_t + e_{it} \quad (1)$$

where ΔY_{it} is a vector representing the change in bond yield or in the CDS spread of a country i at time t , $\Delta BidAsk_{it}$ denotes the change in the quoted bid-ask spread on the 5-year sovereign bond and CDS market, ΔEqR_{it} is the change in idiosyncratic equity returns, $\Delta 3mEur_t$ is the change in the 3-month Euribor-OIS spread, ΔVIX_t is the change in the VIX index, ΔEVZ_t is the change in the Euro/USD exchange rate volatility index and

⁵The index is provided by the CBOE and follows the methodology of the VIX index.

⁶In alternative specifications we control for country credit risk with the average CDS spread of the country’s banking sector. This owes to the significant comovement between bank and sovereign CDS spreads during the Eurozone crisis, particularly after the announcement of financial sector bailouts ([Acharya et al., 2014](#); [Antón et al., 2018](#)). Since CDS data are only available for a limited number of banks in each country and for a shorter period than our examination period, we include this in sensitivity analysis.

$\Delta SP500_t$ is the change in the S&P500 stock market index.⁷

2.2 Evolution and stability of model factors

An identification challenge relates to the possibility that our selected factors have a varying explanatory power for our bond and CDS series; this may be partly attributed to the fact that the key drivers of sovereign credit risk have affected CDS premia and bond spreads differently during the global financial crisis (Groba *et al.*, 2013; Fontana and Scheicher, 2016). In fact, as volatilities tend to dramatically increase during crises, increased correlations are not necessarily indicative of contagion, a concept introduced by Forbes and Rigobon (2002). Hence, not controlling for the possibility that our bond and CDS series exhibit a time-varying exposure to fundamentals (as in Bekaert *et al.*, 2014; Bekaert and Hoerova, 2016), might lead to erroneous findings of contagion.⁸

In this respect, successful identification rests on ensuring that changes in factor loadings and changes in the correlation of the residual series do not compensate each other, leading to stability in the correlation across the original series. If factor loadings remain stable, any change in the correlation of the residual series would correctly point to contagion. To confirm the stability of the parameters in our factor model, we adopt two strategies.

First, we perform a CUSUM test for each of our bond and CDS series. The CUSUM test is based on the cumulative sum of the recursive residuals. This option plots the cumulative sum along with the 5% critical lines. The test finds parameter instability if the cumulative sum goes outside the area between the two critical lines. We present estimates from this test in Figure 5 (for the bond series) and Figure 6 (for the CDS series) of the Appendix A.3. Both figures indicate parameter stability for all bond yields and CDS spreads in our sample.

Second, we employ the Kalman filter and estimate time-varying coefficients for our set of bond and CDS factors. This approach enables us to determine the optimal degree of

⁷A battery of unit root tests for credit spreads and for the state variables expressed in levels do not reject the unit root hypothesis. Therefore, we estimate this equation in changes only (results are omitted for brevity and are available upon request).

⁸Bekaert *et al.* (2014) analyze the transmission of the 2007-09 crisis to 415 country-industry equity portfolios in 55 countries. They employ a three-factor model, to distinguish between a US-specific factor, a global financial factor and a domestic factor respectively. Bekaert and Hoerova (2016) employ techniques of time-varying risk aversion and uncertainty for Germany and the US from January 1992 to March 2008. They find that the variance premium contains a substantial amount of information regarding risk aversion in both countries, while the credit spread primarily contains information about economic uncertainty.

coefficient persistence and smoothing, thereby capturing any variation in the exposure of our bond and CDS series to our factors. The Kalman filter model with time varying-coefficients is specified as:

$$y_t = \alpha_t + Z_t s_t + \epsilon_t \quad (2)$$

$$s_{t+1} = l_t + T_t s_t + u_t \quad (3)$$

where Equation (2) is the measurement equation and Equation (3) is the transition or state equation. α_t , Z_t , l_t and T_t are conformable vectors and matrices, and ϵ_t , u_t are vectors of mean zero, Gaussian disturbances, $u_t \sim N(0, Q_t)$, and $\epsilon_t \sim N(0, H_t)$.

In the empirical application of linear regression, regression coefficients are presented by the state vector. The state vector of the Kalman filter is expressed as:

$$s_t = \begin{bmatrix} \beta_{0,t} \\ \beta_{1,t} \\ \beta_{2,t} \\ \beta_{3,t} \\ \beta_{4,t} \\ \beta_{5,t} \\ \beta_{6,t} \end{bmatrix}$$

and the states (the regression coefficients) evolve according to random walk:

$$\beta_{i,t+1} = \beta_{i,t} + u_{i,t}, \quad u_{i,t} \approx N(0, \sigma_{ui}^2), \quad i \in [0, 6].$$

The model can be rewritten as follows:

$$\Delta Y_t = [1 \ \Delta BidAsk_{it} \ \Delta EqR_{it} \ \Delta 3mEur_t \ \Delta VIX_t \ \Delta EVZ_t \ \Delta SP500_t] \begin{bmatrix} \beta_{0,t} \\ \beta_{1,t} \\ \beta_{2,t} \\ \beta_{3,t} \\ \beta_{4,t} \\ \beta_{5,t} \\ \beta_{6,t} \end{bmatrix} + \epsilon_t$$

$$\begin{bmatrix} \beta_{0,t+1} \\ \beta_{1,t+1} \\ \beta_{2,t+1} \\ \beta_{3,t+1} \\ \beta_{4,t+1} \\ \beta_{5,t+1} \\ \beta_{6,t+1} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \beta_{0,t} \\ \beta_{1,t} \\ \beta_{2,t} \\ \beta_{3,t} \\ \beta_{4,t} \\ \beta_{5,t} \\ \beta_{6,t} \end{bmatrix} + \begin{bmatrix} u_{0,t+1} \\ u_{1,t+1} \\ u_{2,t+1} \\ u_{3,t+1} \\ u_{4,t+1} \\ u_{5,t+1} \\ u_{6,t+1} \end{bmatrix}$$

where $Z_t = [1 \ \Delta BidAsk_{it} \ \Delta EqR_{it} \ \Delta 3mEur_t \ \Delta VIX_t \ \Delta EVZ_t \ \Delta SP500_t]$, $\alpha_t = l_t = 0$, $H_t = \sigma_\epsilon^2$, and $Q_t = diag\{\sigma_{u0}^2, \sigma_{u1}^2, \sigma_{u2}^2, \sigma_{u3}^2, \sigma_{u4}^2, \sigma_{u5}^2, \sigma_{u6}^2\}$. We let the state variances for

the time-varying coefficients, σ_{ui}^2 , and the measurement equation variance, σ_ϵ^2 , to vary over time. A more detailed analysis of the Kalman filter is provided in the Appendix A.4.

We use the Kalman filter for each country's bond yields vis-à-vis each of the different bond market factors and present results in Figure 7 of the Appendix. Similarly, in Appendix Figure 8 we present results from the Kalman filter for each country's CDS spreads and each of the relevant CDS market factors. Turning to Figure 7, we observe that the assumption of parameter constancy holds for almost all countries. According to our estimates, the sign and the magnitude of the correlations between bond yields and the six factors in each country remain stable over our sample period. This is further evident for the CDS series in Figure 8.

The analysis from the Kalman filter further points to a substantial asymmetry in the importance of each of the bond and CDS fundamentals in our factor model. Moreover, we observe that the pricing of these factors is different between the bond and CDS markets, which further motivates our examination of cross-asset contagion in Section 5; for example, an increase in the level of systematic risk is accompanied by an increase in the Greek CDS spreads, but not in the Greek bond yields.⁹ Overall, findings from this exercise indicate that there are only small time variations in the exposures of the bond and CDS series to our six model factors (the β' s in our factor model).

Having tested the stability of our factor model's coefficients, we subsequently obtain the standardised signal errors through Kalman filtering (i.e., the residuals from the factor model). To identify increasing correlations between the residual bond/CDS series, we use a local correlation approach and apply a bootstrap test for contagion; we discuss these methods in Section 3 and provide more technical details in Sections A.1 and A.2 of the Appendix.

3 Contagion identification scheme

Tjøstheim and Hufthammer (2013) introduce a new measure of nonlinear dependence inherent to the concept of local correlation (cf. also Teräsvirta *et al.*, 2010). The central idea of the new approach is to approximate an arbitrary bivariate return distribution by a family of Gaussian bivariate distributions. At each point of the return distribution there is a Gaussian

⁹Additionally, we would expect that when risk aversion is high, spreads are relatively high. This is not confirmed by our time-varying estimations for none of the two markets. This is in line with the arguments about limited pricing of risk in the run-up to the global financial crisis (see D'Agostino and Ehrmann, 2014).

distribution that approximates that point (approximates the density locally rather than the correlation). The correlation of the approximating Gaussian distribution is taken as the local correlation in that neighbourhood.¹⁰

3.1 A bootstrap test for contagion

This section presents another application of the local Gaussian approximation proposed by [Støve et al. \(2014\)](#), who measure and test for financial contagion. Following the concept of [Forbes and Rigobon \(2002\)](#) contagion is defined as "a significant increase in cross-market linkages after a shock to one market (or a group of markets)".

According to [Støve et al. \(2014\)](#), the application of the bootstrap test for contagion necessitates the independence of each variable over time. To satisfy this, one could employ GARCH-family models that pick up any fat tails of the distribution. However, this type of standardisation may remove interesting dynamics that affect the results. To ensure the *i.i.d.* assumption of the local Gaussian correlation approach, we apply to the standardised residual series of each factor model the bootstrap test of independence of [Lacal and Tjøstheim \(2017\)](#), and the Ljung-Box test along with the ARCH-LM test of [Engle \(1982\)](#) for conditional heteroscedasticity (the results of the tests are presented in the Appendix Tables [A.6](#) and [A.7](#)).¹¹ All tests show that our residual series are *iid* so they can be employed in the standard bootstrap procedure.¹²

Denote that $d_t = (Z_t, X_t), t = 1, \dots, T$, as the pair of standardised residual series of two bond markets. Subsequently, the standardised residuals series are split into two sub-periods, the period before the correlation break-date (denoted as BB) and the period after the correlation break-date (denoted as AB). The chronology of the different event periods is identified through a correlation break test. The window size for the periods surrounding each event (pre- and post- event periods) is set to 300 observations for the pre-event period and to 150 observations for the post-event period. This follows [Støve et al. \(2014\)](#), who show that

¹⁰For brevity, a detailed description of the local Gaussian correlation procedure is provided in Appendix [A.2](#).

¹¹For the serial independence test we use the `ind.test()` function ("*lg-package*") employing the R-project program.

¹²For additional robustness, we further estimate the factor model using a GARCH(1,1) model with normally distributed errors. This should remove the GARCH heteroscedasticity from the series, although it will not pick up any fat tails in the yields distribution (as would the Student's *t* distribution). This method provides identical results with the factor model filtered residuals.

the bootstrap test of contagion using a timeline setting of at least 300 observations before an event (crisis) and 100 observations after the event is robust in terms of significance and power. For cases where two close events/crashes cause large increases in correlations and the pre-event period of the first event largely overlaps with the post-event period of the other event, we set different window sizes ranging from 90 to 300 observations (with increments of 30 observations) and test the robustness of the results for each event.

Contagion is confirmed if the local correlation for the post-break period has increased significantly compared to that before the break period. By comparing the two estimates (BB and AB), the null and the alternative hypothesis can be written as:

$$H_0 : \rho_{BB}(x_i, z_i) = \rho_{AB}(x_i, z_i) \text{ for } i = 1, \dots, n \text{ (i.e., no contagion)}$$

$$H_1 : \sum_{i=1}^n (\rho_{AB}(x_i, z_i) - \rho_{BB}(x_i, z_i)) > 0 \text{ (i.e., contagion)}$$

where (x_i, z_i) for $i = 1, \dots, n$ are the selected fixed gridpoints used to estimate the local correlations. For the choice of the gridpoints, we follow [Støve et al. \(2014\)](#) and employ the diagonal grid ($x_i = z_i$) to minimize the computational time. The next step is to compare the two estimates calculating the following test statistic:

$$D_1 = \frac{1}{n} \sum_{i=1}^n [\hat{\rho}_{AB}(x_i, x_i) - \hat{\rho}_{BB}(x_i, x_i)] w_i(x_i, x_i),$$

where w_i denotes a weight function that allows to concentrate on a certain region. The weight function is chosen to minimize the distance between the gridpoints and the observations. In this regard, we avoid the estimation of local correlation in a gridpoint far away from any observations. Since the proposed test statistic has a non-analytical asymptotic distribution, [Støve et al. \(2014\)](#) suggested the following bootstrap procedure in order to obtain the p -value of the test:

Step 1: Randomly drawing observations of the standardised residuals $\{d_1, \dots, d_T\}$ and replacing them in $\{d_1^*, \dots, d_t^*\}$.

Step 2: Split the resampled data into BB and AB and calculate $\hat{\rho}_{BB}^*(x_i, z_i)$ and $\hat{\rho}_{AB}^*(x_i, z_i)$ on the grid (x_i, z_i) for $i = 1, \dots, n$.

Step 3: Calculate the bootstrapped test statistic:

$$D_1^* = \frac{1}{n} \sum_{i=1}^n [\hat{\rho}_{AB}^*(x_i, x_i) - \hat{\rho}_{BB}^*(x_i, x_i)] w_i(x_i, x_i),$$

Repeating Step 1 to Step 3 5000 times allows us to compute D_1^* and construct its distri-

bution. The p -value of the test can be found by comparing D_1 in terms of the distribution of bootstrapped statistics D_1^* , and a decision of rejecting H_0 (no contagion hypothesis) can be made if it is below a given significance level α .¹³

4 Data

4.1 Bond yields and CDS spreads

The dataset consists of bond yields and CDS spreads for sovereign bonds and CDS contracts with 5-year maturity from January 2, 2006, to April 5, 2018.¹⁴ The sample includes all EMU member states at the time of the euro's introduction, i.e., Austria (AT), Belgium (BE), Finland (FI), France (FR), Greece (GR), Germany (DE), Ireland (IE), Italy (IT), the Netherlands (NL), Portugal (PT), and Spain (ES). The 5-year tenor constitutes the most liquid and frequently quoted part of the credit curve and thus, the most traded maturity for CDS contracts. Daily bond yields are from Thomson Reuters Datastream and daily CDS spreads are from Markit. We further convert the change in bond yields into basis points (bps), due to the CDS spread changes (calculated as: $\Delta Y_{it} = Y_{it} - Y_{it-1}$) being already in basis points. We employ the daily frequency given that comovements in the bond and CDS markets are not constant as investors shift their assets; in addition arbitrage opportunities, which represent a significant driver (and consequence) of these comovements are likely to be diminished at lower frequencies. In alternative specifications, we use as dependent variables the quoted bid-ask spread on the sovereign bonds and CDSs with 5-year maturity and the CDS-bond basis for the same maturity.

4.2 Descriptive statistics

Figure 1 (Panel B) graphs the EMU periphery bond yields, which have been soaring since late-2009 and after mid 2010, when Greece reached an agreement with the European Union (EU) and the International Monetary Fund (IMF) for a €110 billion financing package to recover from its debt crisis. Bond yields in the EMU core (Panel A) follow the opposite

¹³The authors would like to thank Dag Bjarne Tjøstheim and Bård Støve for providing the R codes for the contagion bootstrap test.

¹⁴For Greece the sample ranges from January 2, 2006 to March 8, 2012 since the next day (March 9, 2012), after the agreement on sovereign-debt restructuring, the Greek bonds stopped trading. Data for Luxembourg is available from February 2009 onwards, except for the period from January 2010 to February 2011 and therefore, the country is not included in the sample.

course; they have been falling during the post-2008 period, with only a slight rise in the first half of 2011, appearing to have benefited from the skyrocketing borrowing costs in the periphery. As shown in Figure 1, CDS spreads in the periphery (Panel D) match closely the upward trend of the periphery bond yields (Panel B), pointing to a close association between the two asset-markets. However, CDS spreads also increased in the core (Panel C), indicating an overall rise in sovereign default probabilities across the entire Eurozone.

[Insert Figure 1 about here.]

Table 1 presents descriptive statistics for daily changes in bond yields and CDS spreads. The average bond yield change (standard deviation), expressed in basis points, is -0.121 (4.54) for the core and 0.395 (17.87) for the periphery. Similarly, the mean CDS spread change (standard deviation), is 0.005 (8.15) and 3.086 (137.2) for the core and periphery respectively. Overall, GIIPS bond yields and CDS spreads are higher and more volatile than their core counterparts. Table 2 presents the response of bond yields and CDS spreads in each country to a number of negative events in the EMU South. We observe a sharp rise in the bond yields and CDS spreads of the source country (in bold), which exceeds the equivalent rise in the remaining countries. Notably, this is accompanied by a rise in the cross-asset correlations between the source country's bond yields and CDS spreads after each event: the degree of comovement in this period ranges between 0.50 and 0.75, indicating that sovereign bond and CDS markets do not respond uniformly to the same shocks.

[Insert Tables 1 and 2 about here.]

5 Empirical results

We proceed to the examination of the degree of pair-wise conditional correlations between and across the European sovereign bond and CDS markets considering as source of contagion each of the European periphery countries, namely Greece, Ireland, Italy, Portugal and Spain. We detect endogenous break-dates on which contagion transmission is initiated by using the algorithm for correlation change-point inference of Demetrescu and Wied (2019).¹⁵

¹⁵The Demetrescu and Wied (2019) methodology for the identification of structural changes in correlation is described in detail in Appendix A.1.

This is a test for constant correlations that allows for breaks at unknown times in the marginal moments (means and variances). This also enables us to locate and identify a change in the correlation between the bond and CDS markets and thus determine the exact date associated with a fundamental change in their relationship. The identification of a structural change in the cross- and intra-asset correlations further allows us to split the sample into a pre- and post-event period, or more properly into a pre- and post-contagion period. We then quantify the impact of the structural change by estimating the transmission of a shock to the bond and CDS markets during the respective periods.

In line with our discussion in Section 2, we conduct the analysis by employing the bond and CDS series after accounting for fundamentals via the factor model of Equation (1); we present results from this exercise in Tables 3 and 4. We then conduct the same analysis for the original bond and CDS series and present results in Appendix A.8 (Tables 10 and 11).¹⁶ By contrasting the results from the two methods, we can differentiate between cases of “pure contagion” and cases of “wake-up call” and “limited contagion”. In this respect, if the hypothesis of contagion between a given pair is verified by the estimation of the original series but rejected by the estimation of the residual series, this would serve as evidence of “wake-up call”. In contrast, findings of contagion under both the original and residual series should be interpreted as evidence of “pure contagion”. Last, if contagion is only verified by the residual series, this points to a short-lived episode of “limited contagion”.

5.1 Transmission of contagion from the GIIPS bond markets to the EMU bond and CDS markets

5.1.1 Intra-bond contagion

Table 3, presents the results on the transmission of contagion from the bonds of the EMU South towards the bonds (Panel A) and CDSs (Panel B) in each of the eleven EMU member states. When the analysis is concentrated solely within the bond market, estimates in Panel A reveal that bonds yields across the Eurozone have been immune to contagion phenomena during the pre- as well as post-financial crisis period: the bootstrap test for contagion is not able to reject the null hypothesis of no-contagion in practically all pairs (columns 1 to 5

¹⁶As in Forbes and Rigobon (2002), we have performed heteroscedasticity-filtering in the original series. If heteroscedasticity bias is ignored when testing for changes in correlation, then contagion is over-accepted.

in Panel A). This absence of contagion stemming from the distressed periphery countries (also partly documented in [Missio and Watzka, 2011](#); [Metiu, 2012](#); [De Santis, 2012](#); [Beirne and Fratzscher, 2013](#); [Claeys and Vařicek, 2014](#); [Caporin *et al.*, 2018](#)) is not necessarily a sign of widespread disassociation between the sovereign debt markets.¹⁷ It can serve as (i) supporting evidence of the investors' flight(s) away from the risky periphery bonds to the safer core bonds and therefore, of a negative correlation between bond yield movements in the two blocs (see [Beber *et al.*, 2009](#)) and (ii) an indication that the common shifts of periphery bond yields can be explained in terms of the –enduring– interdependence between them.

[Insert Table 3 about here.]

5.1.2 Bond-CDS contagion

This disassociation is not confirmed for the bond-CDS nexus: results in Panel B of Table 3, point to transmission of bond-originated contagion from all countries of the EMU South to nearly all European countries. These results are derived from the analysis of the residual series, thus indicating contagion over and above what can be explained by fundamentals (“pure contagion”). Notably, this transmission occurs over a series of different phases, all coinciding with major economic and political events during the late 2009–mid 2011 period.

The first phase is during November 2009, shortly after the upward revision of the Greek government's budget deficit: estimates in the first column of Panel B, indicate the transmission of contagion from the Greek bond to each of the European CDS markets.¹⁸ The Greek-bond stemming contagion does not appear to be a one-off phenomenon, as it further emerges in April 2010. The respective month is characterized by the growing fears about a possible Greek default, which ultimately led to the Greek government's official request for the activation of the joint EU/IMF aid package.¹⁹

¹⁷The aforementioned studies find no evidence of shift contagion across the European sovereign bond markets when the role of contagion originator is primarily assumed to be Greece.

¹⁸The Greek budget deficit was initially revised upward from 6.0% of GDP to 12.5% on October 19, 2009 by the new Greek Minister of Finance in his first Eurogroup meeting. In the budget draft for fiscal year 2010 submitted to the Hellenic Parliament for consideration in November 5, 2009, the 2009 deficit was revised to 12.7%, while in the final draft of November 16, 2009 (voted by the Parliament), it was revised to 13.6%.

¹⁹On April 23, 2010, Greek Prime Minister George Papandreou made a live broadcast announcement from the Greek island of Kastelorizo on the request of the €60 billion EU/IMF financial aid programme.

The next phase occurs in November 2010 and includes the transmission of contagion from the Irish bond yields to almost all European CDS spreads (second column of Panel B). This period marks the re-eruption of the Irish crisis (back from the 2008 banking crisis and the €64 billion bailout of Irish banks) and resulted in a joint EU/IMF financial assistance programme. The Irish government's request for the programme was made on November 21, 2010, while on November 24, 2010 the government outlined €15 billion in spending cuts and tax increases to reduce its budget deficit from 31% of GDP to 3% by 2014. These events were followed by massive rallies and protests in Dublin three days later, before the €67.5 billion bailout package being accepted on November 28, 2010.

The fourth and more intense phase of contagion occurred during the second and third quarter of 2011. Starting from April 2011, the Portuguese government requested financial assistance from the EU, which in early May was agreed to be provided jointly by the EU and the IMF.²⁰ According to estimates in the fourth column of Panel B, the period shortly before the Portuguese request (late March 2011) until the final bailout agreement (early May 2011) was characterized by the transmission of contagion from Portuguese bond yields to nearly all EMU CDS spreads. Contagion phenomena were even more prevalent during the third quarter of 2011. Results in the third and fifth column of Panel B, point to contagion from the bond markets of Italy and Spain respectively towards each of the EMU CDS markets. Most of the correlation break-dates that indicate contagion are observed between mid July and early August of 2011, a period of significant turmoil in the Italian and Spanish economies.

In specific, on July 14, 2011, Italy raised €3 billion from selling government bonds, albeit at a record interest rate of 5.9%. One day later, the European Banking Authority (EBA) announced that five Spanish banks failed its “stress tests”, while seven other Spanish banks barely passed.²¹ During the same period, talks abounded that Greece would become the first country to be forced to exit the Eurozone. Indeed, a few breaking points are observed on July 20, just one day before the agreement between EU and IMF for a second bailout package totaling €109 billion. However, this was not able to contain speculations on a potential

²⁰On April 6, 2011, Portuguese Prime Minister José Sócrates extended a request to the EU for a financial assistance programme. On May 6, 2011, Portugal reached an agreement for a €78 billion EU/IMF programme.

²¹The banks that failed the EBA's stress tests were Catalunya Caixa, Caja de Ahorros de Mediterraneo, Banco Pastor, Unnim, and Group Caja3. Seven banks, i.e., Banco Sabadell, Banco Popular, and Bankinter and the savings banks Novacaixagalicia, Caja Ontinyent, Banca Civica, and Bankia, just achieved the minimum requirement of core equity Tier 1 (CET1) capital ratio of 5.0%.

Greek default, which were echoed in the European Commission President’s warning that the sovereign debt crisis was spreading beyond the Eurozone periphery. According to our findings, developments in Italy and Spain as well as continuing speculations about a looming Greek exit were diffused across the Eurozone CDS markets, via the rising Italian and Spanish bond yields. This period of contagion transmission is further consistent with the onset of the upward trend in the Italian and Spanish bond yields, evidenced in Panel B of Figure 1.

We subsequently contrast the results in Table 3 with those from the original series in Table 10 of the Appendix. Contrary to Panel A in Table 3, estimates in Panel A of Table 10 provide evidence of widespread contagion from each of the GIIPS bond markets to practically all European bond markets. Since these findings are derived from the analysis of the original series, but not confirmed by that of the residual ones, they point to a form of “wake-up call”. Importantly, most correlation change-dates fall within the late 2008–early 2009 period. Thus, we can infer that movements in the European bond markets were part of the general repricing of sovereign credit risk after the global financial crisis and during the onset of the Eurozone crisis. Turning to Panel B in Table 10, we observe that estimates exhibit only marginal deviations from those of Panel B in Table 3, confirming the transmission of “pure contagion” from the GIIPS bond markets to all European CDS markets.

5.2 Transmission of contagion from the GIIPS CDS markets to the EMU bond and CDS markets

5.2.1 CDS-Bond contagion

Table 4, presents results when contagion is assumed to be stemming from the CDS market. Estimates from the examination of the CDS-bond transmission in Panel A, reveal that contagion phenomena are restricted only within the EMU South and Belgium. The majority of correlation change-points indicating contagion are observed during the fourth quarter of 2011 and in particular, November 2011. Initially, contagion appears to originate from Greece in early November (column 1 in Panel A). The Greek CDS-stemming contagion could be linked to the political developments in Greece during that period, which had implications for the Eurozone’s viability. On October 31, 2011, the Greek Prime Minister called for a referendum on the EU/IMF rescue plan for Greece agreed only days earlier, which on

November 2, 2011 was modified to be a referendum on Greece’s Eurozone membership.²² Shortly after the referendum call, efforts for the formation of a national unity government in Greece temporarily collapsed (November 4, 2011), only to resume successfully seven days later.²³

[Insert Table 4 about here.]

According to our estimates, this heightened period generated contagion from the Greek CDS spreads towards the periphery bond yields. Shortly after, contagion phenomena further emerged from the periphery economies. Results in columns 2-5 of Panel A show that the CDS markets of Ireland, Italy, Portugal and Spain transmitted contagion to the bond markets in each country of the EMU South and Belgium. The dates of this transmission are located in the days right after the formation of the Greek government unity. Evidently, the political uncertainty in Greece sparked a contagion wave within the EMU South. Importantly, this transmission is not verified by the analysis of the original series: most of the break-dates surrounding the Greek developments of November 2011 do not appear or enter with a non-significant sign (Panel A of Table 11 in the Appendix). This points to a short-lived episode of “limited contagion” owing to the abrupt political uncertainty in Greece that shortly receded.

5.2.2 Intra-CDS contagion

The analysis at the intra-CDS market level (Panel B in Table 4), provides evidence of contagion stemming from the EMU South’s biggest economies, i.e., Italy and Spain, to practically all European CDS markets. Estimates in columns 3 and 5 of Panel B, show that contagion transmission took place over two phases: the first and more intense was during November 2011, where Italian and Spanish CDS spreads directed contagion towards each Eurozone country. These results complement those of Panel A in Table 4, where the same period (and

²²The rescue plan included a 50% debt write-off for private sector investors and €130bn of new bailout loans to Greece. The initial referendum call on the proposed EU/IMF rescue plan was modified after pressures from the French President Nicolas Sarkozy and the German Chancellor Angela Merkel in the G20 Cannes summit of November 2, 2011. The call for referendum was abandoned on November 4, 2011.

²³On November 4, 2011, the leaders of the two largest political parties engaged in talks for the formation of a government of national unity, but talks collapsed within the same day. The following day, Greek Prime Minister George Papandreou resigned and succeeded by Lucas Papademos on November 11, 2011, who led a new government of national unity.

the accompanied political developments in Greece) were found responsible for the transmission of CDS-stemming contagion to periphery bond markets. Based on our estimates, this period further establishes Italy and Spain as the absolute transmitters of CDS contagion towards all European CDS markets. Indeed, during the respective period we observe a surge in the CDS spreads of both EMU blocs (Panels C and D in Figure 1).

The second phase of this transmission is observed in the first weeks of 2012, the period leading to the Greek debt restructuring of March 2012.²⁴ Although contagion phenomena are not as intense as in late 2011, they are still evident and mainly directed towards the largest economies of the core (France, Germany, Netherlands) and Belgium. Notwithstanding their size, the increasing influence of Italian and Spanish CDSs can also be explained when considering the nature of CDS contracts. By construction, CDSs mainly reflect sovereign credit risk. Hence, a possible default by either Italy and/or Spain could trigger domino effects, eventually leading to the collapse of the EMU. Panel B, shows that these concerns and the consequent contagion transmission were more prevalent in late 2011 (primarily) and early 2012 (secondarily). The exposure of European CDS markets to the Italian and Spanish CDS spreads stands in contrast to previous findings within the intra-CDS context, where contagion was found to be either non-existent (see Caporin *et al.*, 2018) or only a European periphery phenomenon (see Broto and Pérez-Quirós, 2015), and to arguments about the limited capacity of the GIIPS to generate contagion (see Kalbaska and Gatkowski, 2012).

We further compare the results from the residual series in Table 4 with those from the original series in Table 11 of the Appendix. Panel A of Table 11 points to certain cases of CDS-bond contagion from the periphery (Italy, Portugal, Spain) towards the core (Germany, Finland, France, Netherlands) that are not evident in Panel A of Table 4. As such, these cases, which are mainly concentrated in the second half of 2010, cannot be classified as “pure contagion”. They can be rather perceived as a “wake-up call” for government bonds across the core and attributed to the investors’ upward repricing of those countries’ fundamentals,

²⁴On February 12, 2012, the Greek parliament voted in favour of a second bailout package for Greece totalling €130 billion. The parliamentary approval of this package was a prerequisite for the debt restructuring of March that occurred between March and April 2012. Under the restructuring, the Greek government amended the conditions of bonds under Greek law with a total face value of €177 billion. The restructuring included the private sector’s involvement, where investors were required to accept a 53.5% haircut of the face value of Greek government bonds. As a result, the nominal value of Greek debt was reduced by €107 billion, approximately 50% of GDP. It constituted the world’s biggest debt restructuring, involving securities of €206 billion.

prompted by the rising CDS spreads across the periphery. Last, estimates from Panel B of Table 11, do not reveal additional cases of intra-CDS contagion that are not subsequently confirmed by the main estimations in Panel B of Table 4.

5.3 Discussion

Our analysis identifies two key phases for the transmission of contagion within the European sovereign financial markets framework. The first extends from late 2009 to mid 2011 and is characterized by the outright transmission of contagion from the bond markets of the EMU South towards the CDS markets of all member states. Contagion phenomena emerge in the periods corresponding to the Greek deficit's upward revision (November 2009) and the fears of a Greek default that culminated in the country's first bailout package (April 2010), the rescue programmes in Ireland and Portugal (November 2010 and April 2011 respectively) and the negative economic developments in Italy and Spain (July-August 2011).

However, during these intense periods there is absence of bond-stemming contagion towards any of the European government bonds whatsoever. The rising CDS spreads in both the periphery and core in the same periods (also evident in Panels C and D of Figure 1) indicate that the negative developments in the periphery were only transmitted across the CDS market. It thus appears that the European bond and CDS markets differ in their capacity to preserve or amplify potential shocks occurring in the periphery. This is further an indication that investors in the sovereign bond market behave according to the deteriorating economic fundamentals of the periphery countries. Implicit in this view is that during the Eurozone crisis, bond markets were more efficient in the pricing of credit risk relative to CDS markets. This complements the findings of *Arce et al. (2013)*, where sovereign bond spreads tend to reflect credit risk more efficiently than CDS spreads during periods of high global risk.

The next phase is around November 2011, and includes Greece's referendum announcement and the prolonged political uncertainty that revived the fears of the country's Eurozone exit. In this period, we observe a) the emergence of CDS-bond contagion in the periphery, and b) the transmission of contagion from the Italian and Spanish CDS spreads to either bloc's CDS spreads. This period marks the onset of the triggering capacity of the CDS market; for the first time during our examination, rising sovereign default probabilities are

transmitted to the periphery bond yields as well as to both blocs' CDS spreads. Evidently, this is the critical point where the Greek debt crisis developed into a European debt crisis.

These developments were associated with rising sovereign default risk, which in turn aggravated the risk aversion of market participants (as in [Mink and De Haan, 2013](#); [Fontana and Scheicher, 2016](#); [Caporin *et al.*, 2018](#)). Since CDS contracts are written on government bonds, they instantly reflected the potential sovereign defaults across the periphery, thereby rendering the CDS market the main field of contagion phenomena. Importantly, due to the high degree of financial integration in the euro area, the markets quickly reappraised the risks for the entire Eurozone. This led to a stronger connection between the European sovereign CDS markets, explaining the transmission of contagion to the core, shortly after.

Our estimates further elevate Italy and Spain to key determinants of sovereign credit risk across the Eurozone: Italian and Spanish CDS spreads are the only transmitters of contagion to all countries' CDS spreads. This transmission mainly occurs during the late 2011–early 2012 period. In addition to speculations about a Greek Eurozone exit, this period was also characterized by increasing market concerns over the economic outlook and debt sustainability of Italy and Spain.²⁵ Due to their economic size (third and fourth EMU economy respectively), both countries are considered too expensive to be realistically bailed out. Only for Italy, where public debt stands at approximately €2 trillion (the world's fourth largest), a 3-year rescue support program was estimated by the IMF at €600 billion. The European Financial Stability Fund evidently did not have enough funds to accomplish this, thereby rendering either country too big to bail out. Hence, this period reflected the fears that a Greek default would cause a domino effect, causing Italy and Spain to fall as well, with resulting implications for Eurozone's stability.

Interestingly, this heavy influence of the Italian and Spanish CDS markets appears to ease after March 2012. According to our estimates, the presence of contagion phenomena regardless of their source is rather limited thereafter. A potential explanation can be offered by an important development in the beginning of 2012, namely the Greek debt restructuring

²⁵In Italy, the low rates of productivity and output growth were not keeping up with an increasing debt load of almost 120% of GDP, the second highest in the Eurozone, behind Greece. In Spain, the property bubble eventually turned to bust, resulting in the country's banks accumulating a mounting pile of bad mortgage debts, and the highest unemployment rate in the EMU. Along these lines, Italy's credit rating was decreased by three notches in late 2010, while Spain's credit rating was cut three times (one notch each time) from 2010 to 2011.

of March 9, 2012. This historical deal prevented a Greek default and demonstrated the EU’s willingness to preserve the Eurozone. For what matters, this deal is associated with the minimization of contagion phenomena across the Eurozone entering the second quarter of 2012, particularly those directed from the periphery CDS markets.

The analysis of the original series (Tables 10 and 11 of the Appendix), provides evidence of a “wake-up call” in the European bond markets in late 2008–early 2009. This period was marked by the spillovers of the financial crisis and the early seeds of the Eurozone crisis. Either crisis provided new information, prompting investors to reassess the vulnerability of other countries, leading to a repricing of sovereign bonds (Bekaert *et al.*, 2014; Claeyes and Vařicek, 2014). The consequent rise in GIIPS long-term government bond yields was further fueled by recession and government announcements of bank rescue operations that exacerbated investor perceptions of sovereign credit risk. This was reflected in the opposite evolution of government bond yields between the periphery and core (Monfort and Renne, 2014; Fontana and Scheicher, 2016), a fact also evident in Panels A and B of Figure 1. The concentration of correlation break-dates in this period, indicates that these developments were priced by the bond markets, only to the extent that affected bond fundamentals.

6 Further analysis

The previous section revealed that the countries of the EMU South acted as transmitters of contagion not only within their own country-bloc but importantly, towards the EMU North. This transmission took place over two distinct phases of the European sovereign debt crisis. Having established that, in this section we perform a number of tests to further understand our findings and assess their heterogeneity around certain fiscal and regulatory events. Moreover, we examine the role of liquidity and CDS-bond basis deviations for contagion transmission.

6.1 The developments in Greece

Estimates in Section 5.1 pointed to a surge in contagion transmission from the Greek bond yields to almost all European CDS spreads in the days surrounding two important events: the November 2009 Greek upward deficit revision (primarily) and the April 2010 Greece’s

request for financial assistance (secondarily). Since these essentially marked the onset of the Eurozone crisis, we examine the transmission of contagion from the Greek bonds to the rest EMU CDSs in the different subperiods associated with these developments. In specific, we estimate our bootstrap test for the period between the deficit revision and the financial assistance request, as well as for the periods preceding and succeeding each event.

Table 5 presents estimates for the period extending from January 2, 2006 (the beginning of our sample) until April 23, 2010 (the Greek Prime Minister's announcement on the EU/IMF financial aid request), with November 16, 2009 as break-date. We observe that after the deficit revision and until the financial assistance request, the negative movements in the Greek bond yields are transmitted to each of the remaining countries' CDS spreads: all p -values are statistically significant at conventional levels, pointing to the contagious nature of the Greek bonds during this period. Importantly, the negative movements in the Greek bonds led to a rise in default probabilities across the Eurozone, as reflected in euro area CDS spreads. We then test for contagion from November 16, 2009 until March 8, 2012 (the last day of trading for Greek government bonds), with April 23, 2010 as break-date. We observe that contagion phenomena recede in the period following the Greek government's request for the EU/IMF economic adjustment programme (non-statistically significant p -values).

[Insert Table 5 about here.]

This pattern is further confirmed when examining the evolution of the dependence structure between our series. Figure 2 illustrates the local Gaussian correlation estimates for the Greek bond yields vis-à-vis each of the remaining countries' CDS spreads, for the period before November 2009 (green line) and for the period between November 2009 and April 2010 (red line). Overall, we observe an increased sensitivity of the CDS market to the negative developments in Greece. The estimates between the Greek bond and the European CDS markets provide strong evidence of increased dependence for all country pairs during the Eurozone crisis. For all the Greek bond–European CDSs pairs, the entire local correlation curve for the post-crisis period has moved up. We should further note that in most cases, there is a similar uniform increase in local correlation over the different segments of the distribution. This points to the existence of strong linkages between the Greek bond market and the European CDS markets during the post-November 2009 period.

[Insert Figure 2 about here.]

On the same line, we contrast local Gaussian correlation estimates for the period between November 2009 and April 2010 (green line) against estimates for the period after April 2010 (red line) and plot them in Figure 3. Again, we observe a positive correlation across all country-pairs in the pre-April 2010 period, that declines thereafter. This decline is more potent for the core relative to the periphery (including Belgium); for the latter, it lies above zero even after April 2010. This points to the existence of strong linkages (albeit not contagious) between the Greek bond and the periphery CDSs after April 2010.

Overall, results from this exercise highlight the dominant role of the November 2009 Greek developments and the contagious capacity of the Greek bond (due to a potential Greek default) for the rising sovereign credit risk across the Eurozone. Moreover, they demonstrate the easing effect exerted by the agreement on the EU/IMF aid programme in April 2010.

[Insert Figure 3 about here.]

6.2 The naked CDS ban

In an attempt to curb destabilizing speculation on distressed Eurozone countries' default, the EU banned the purchase of naked CDS contracts, effective November 1, 2012.²⁶ Because bond and CDS markets are complementary, the naked CDS buyers might inflate sovereign CDS spreads, thereby driving up sovereign bond yields (Silva *et al.*, 2016; Gyntelberg *et al.*, 2018). As Section 5.2 revealed, contagion from the CDS market to the bond market was mainly observed during the Greek Eurozone exit discussions in November 2011.

Although the Greek-stemming contagion subsided following the cancellation of the referendum and the formation of a national unity government in Greece, contagion phenomena were further evident for the remaining countries of the EMU South. During the same period, the regulation on naked CDS contracts was voted into law by the European Parliament (November 15, 2011). Since these events coincide, it may be that the intense political events

²⁶The Regulation on Short Selling and certain aspects of Credit Default Swaps was published in the Official Journal of the European Union on March 24, 2012. The regulation prohibits any person or legal entity in the European Union from entering into uncovered ("naked") CDSs on sovereign debt and restricting uncovered short sales on shares and sovereign debt after November 1, 2012. Effectively, the Regulation bans CDS contracts on sovereign debt that do not hedge exposure to the sovereign debt itself, or to assets or liabilities whose value is correlated to the value of the sovereign debt.

in Greece curtail the easing effect of the regulatory ban brought about by the passage of the law, especially on the bonds of Ireland, Italy, Portugal and Spain.

In this section, we identify the effect of this regulatory change on contagion transmission. We adopt three steps to ensure that any findings on contagion are free from other developments during the same period and thus, could be attributed to the naked CDS ban. First, we choose as our break-date the day when EU lawmakers reached an agreement on the naked CDS ban, namely October 11, 2011. Given that negotiations for the ban were deadlocked for months, the reaching of an agreement (and its timing) was fairly unanticipated (as opposed to the voting in November 15, 2011). Moreover, as the November 2011 political instability in Greece is characterised by contagion phenomena from the periphery CDS markets to that bloc's bond markets and from Italian and Spanish CDS to all CDS markets, we exclude these dates from our sample. Finally, we limit our post-event window to 90 days after October 11, 2011 in order to isolate the effect of the agreement from other events.

Given that the regulation targeted uncovered CDS positions, we expect that the main transmission channel is from the CDS market to bonds. To this end, Table 6 examines differences in contagion transmission from the GIIPS CDSs to all European bonds in the periods before and after October 11, 2011. We find no evidence of contagion from the periphery CDS markets to any of the two blocs following the agreement on the ban: all p -values generated by our test are non-statistically significant.

[Insert Table 6 about here]

To fully detect whether this reflects a change from the pre-October 2011 period, we further examine the evolution of the dependence structure between our series. Figure 4, illustrates the local Gaussian correlation estimates for the GIIPS CDS spreads vis-à-vis each of the remaining countries' bond yields, for the period before October 11, 2011 (green line) and the period after (red line). Evidently, there is a strong positive correlation between the CDSs and bonds of GIIPS pre-October 2011, which recedes close to zero in the period after. On the other hand, the correlation between GIIPS CDSs and core bonds is very close to zero in both subperiods. Taken together, results from this exercise confirm the containment effect of the regulatory ban on the transmission of CDS-bond contagion within the EMU South. The agreement on the regulation appears to have strengthened the easing effect exerted by

the reversal of the negative developments in Greece in early November 2011, namely the country’s political instability and intention about a Eurozone membership referendum.

[Insert Figure 4 about here]

6.3 The role of liquidity

In periods of market distress investors tend to rebalance their portfolios towards less risky and more liquid securities, a phenomenon referred to as “flight-to-quality” or “flight-to-liquidity” (Beber *et al.*, 2009; Brunnermeier, 2009; Monfort and Renne, 2014; Gyntelberg *et al.*, 2018). This is particularly important in the context of the European sovereign bond market, since the destination of large flows into (and out of) this market is determined almost exclusively by liquidity (Beber *et al.*, 2009; Monfort and Renne, 2014).

In practice, it is difficult to disentangle the two phenomena in the Eurozone crisis setting. If investors decrease their periphery bond holdings in favour of core countries’ bonds, it is not clear whether they do so because of concerns about credit risk or liquidity risk (also given the strong correlation between the two). However, if contagion transmission from the EMU South (due to rising bond yields and/or CDS spreads) is also accompanied by a general drop in liquidity across the same bloc (an illiquidity contagion), this would be a supporting argument that liquidity is an additional driver of investors’ actions during the crisis.

To examine this premise, we test for contagion between sovereign bond and CDS liquidity by replacing our bond yields and CDS spreads with a measure of liquidity, namely the quoted bid-ask spread (see Goyenko and Ukhov, 2009). To this end, we estimate Equation (1) with the 5-year sovereign bond (CDS) quoted bid-ask spread as dependent variable and use the residuals to test for contagion between our liquidity measures. In essence, we assume that if the negative developments in the periphery affect investors’ liquidity risk considerations (raising bid-ask spreads), this should also be reflected in the transmission of illiquidity across that bloc. Furthermore, if investors search for liquidity in the core, the latter should be relatively immune to liquidity developments in the periphery.

However, liquidity conditions in the CDS market (primarily) and the bond market (secondarily) are susceptible to two opposing forces: while a surge in default probabilities should raise bid-ask spreads, an increase in the demand for CDS contracts (as a means to insure

against a potential default) should lower bid-ask spreads. As such, any test of illiquidity contagion should control for the possibility that, in the presence of elastic demand, higher trading volume should lead to greater liquidity (reflected in tighter bid-ask spreads). Given this, our factor model for bond yields (CDS spreads) includes the volume of sovereign bonds (CDSs) as an additional factor.²⁷ This aims at controlling for any demand-side forces that shape bond (CDS) liquidity.

Results from this exercise are presented in Table 7, where we initially focus on the transmission of illiquidity from the GIIPS bonds to all European bonds (Panel A) and CDSs (Panel B). According to Panel A, an increase in quoted bid-ask spreads in any of the GIIPS bonds is not transmitted to the bid-ask spreads in either of the two country-blocs; almost all p -values are non-statistically significant at conventional levels.

[Insert Table 7 about here]

When turning to Panel B, we notice that a rise in bid-ask spreads in any of the GIIPS bonds is transmitted to that bloc's CDSs; notably, the transmission dates match closely those of the transmission of price contagion in Panel B of Table 3. Nevertheless, this rise in periphery bond illiquidity is not transmitted to the core bloc's CDSs. If this is combined with the results about the transmission of bond-CDS price contagion in Section 5.1.1, we can argue that a rise in credit risk in the periphery has contagious effects on both blocs' CDS markets, whereas a rise in liquidity risk is confined within the periphery CDS markets.

On the other hand, testing for illiquidity contagion stemming from the GIIPS CDS markets (either to all bond markets or to all CDS markets) did not yield significant results (not presented for brevity). We conclude that liquidity risk is a material source of risk in the context of the European sovereign CDS market. Importantly, this risk is more potent when stemming from the bond market. Altogether, results from this exercise reveal that liquidity risk is an essential determinant of price discovery between the two markets, complementing credit risk.

²⁷We employ the natural logarithm of the notional amount (in million) for each 5-year sovereign bond (CDS). Due to data availability, data on trading volume is employed at weekly frequency.

6.4 Additional tests

We conduct additional sensitivity tests, the results of which are available on request. First, we estimate our baseline factor model by including the average CDS spread of the country's banking sector as an additional factor. This controls for the strong sovereign-bank nexus, reflected in the comovement between sovereign and bank CDS spreads during the Eurozone crisis (Acharya *et al.*, 2014; Antón *et al.*, 2018). A limitation of this approach is that CDS data are only available for a limited number of banks in each country and for a shorter period than our examination period. Estimates from this exercise are qualitatively and quantitatively similar to our baseline.

Second, we consider the role of arbitrage opportunities as explained by the deviations of the CDS-bond basis from its zero equilibrium value. In fact, arbitrage opportunities during the Eurozone crisis were primarily present due to “funding frictions” and “short-selling frictions”. While the former type of frictions made it difficult for arbitrageurs to finance the purchase of the bond (via repo transaction) for implementing a “negative basis trade”, the latter prevented arbitrageurs to short-sell the bond (in a “positive basis trade”) in order to profit from the relative mispricing (see Fontana and Scheicher, 2016).

Since arbitrage opportunities affect the equilibrium relationship between bonds and CDS contracts, we test for contagion transmitted from the GIIPS bond yields towards the CDS-bond basis of all countries. We do so, by estimating Equation (1) with the 5-year sovereign CDS-bond basis as dependent variable, and calculate the residuals. We subsequently test for contagion between our residual bond yields and the residual basis series. Across our sample period, we find no dates that indicate contagion transmission from the GIIPS bonds to the basis of any country in any bloc. This is further evident, when the reverse direction is considered, i.e., from the GIIPS basis to all European bond yields.

We conclude that although arbitrage forces might have been present in the context of the Eurozone crisis, they have nevertheless been unable to affect contagion dynamics. In fact, frictions and imperfections, such as illiquidity and high trading costs, often prevent arbitrage forces from fully closing the pricing gaps between the two markets; if markets are subject to such frictions, it is possible that the correcting mechanisms may have different regimes with different adjustment speeds (Gyntelberg *et al.*, 2018).

Third, we investigate the impact of the ECB's government bond purchases under the Securities Markets Programme (SMP). These purchases aimed at lowering yields and liquidity premia in the distressed countries' sovereign bonds and restoring monetary policy transmission (Eser and Schwab, 2016). As such, we expect that these purchases contributed to the minimization of contagion transmission from the GIIPS bond markets. To examine this, we test for differences in intra-bond contagion before and after the SMP's implementation in May 10, 2010. We find no evidence of contagion following the May 2010 period. We then test for changes in the correlation of our series with May 10, 2010 as break-date; again the correlation is around zero in both subperiods. Since the programme was reactivated in August 7, 2011 to enable the purchases of Italian and Spanish bonds, we further test for contagion before and after this date; again, results confirm the absence of intra-bond contagion between almost all country-pairs.

Finally, we consider the impact of the two LTRO auctions (December 8, 2011 and February 29, 2012) on contagion transmission from the GIIPS bonds to all European bonds. If the massive liquidity injections to the euro area banking sector eased sovereign default concerns via the sovereign-bank nexus, we expect the minimization of contagion phenomena following these auctions. We fail to find any contagion phenomena within the European sovereign bond market in the periods before and after each of the auction dates.

6.5 Summing up

Overall, results in this section reveal certain heterogeneities in the sovereign bond and CDS contagion dynamics during the Eurozone crisis. These are mainly associated with the occurrence of certain fiscal and regulatory events during the main phase of the crisis. Moreover, contagion phenomena primarily arise due to credit risk concerns, although liquidity risk also plays a non-trivial role, particularly when stemming from the GIIPS bond markets.

7 Conclusion

By adopting the correlation concept of contagion by Bekaert *et al.* (2005) and employing a new measure of local Gaussian correlation by Tjøstheim and Hufthammer (2013), we examine asymmetric and nonlinear changes in dependence structure and test for contagion in

the European sovereign bond and CDS markets during the 2006–2018 period. Our empirical findings suggest that contagion phenomena in the bond and CDS markets have undergone two major phases.

The first phase extends from late 2009 to mid 2011 and concerns the outright transmission of contagion from the bond markets of the EMU South towards the CDS markets of both EMU blocs. Most correlation break-points that indicate contagion are concentrated around important economic events. These include the Greek deficit's upward revision (November 2009) and the consequent first rescue plan for Greece (April 2010), the financial assistance programmes in Ireland and Portugal (November 2010 and April 2011 respectively) and the negative economic developments in Italy and Spain (July–August 2011). In contrast, during these intense periods, bond markets in both blocs appear immune to contagion stemming from the periphery bond markets; they were rather subject to a “wake-up call” during the late 2008–early 2009 period. This “wake-up call” prompted investors to reassess the vulnerability of Eurozone countries, leading to a repricing of the factors that affect sovereign bonds (Bekaert *et al.*, 2014; Claeys and Vařicek, 2014)

The second phase is during November 2011 and the Greece's referendum announcement that prolonged political uncertainty and revived the fears of the country's Eurozone exit. This period is characterized by contagion stemming a) from periphery CDS spreads towards bond yields in the periphery and Belgium and b) from Italian and Spanish CDS spreads towards CDS spreads in both EMU blocs. The outbreak of CDS-stemming contagion during the respective period marks the development of the Greek debt crisis into a European debt crisis. During this escalation, Italy and Spain emerge as key countries for the evolution of sovereign credit risk across the Eurozone. This is consistent with the “too big to bail out” status of either country. Arguably, the fears that a Greek default would generate a domino effect, causing Italy and Spain to fall as well, were well-founded.

Our examination of the mechanisms of contagion transmission shows that liquidity risk is an important source of risk in the European CDS market, particularly when stemming from the bond market. We further show that although arbitrage forces may be present, they were unable to fully close the pricing gaps between the two markets and affect contagion dynamics. Finally, by studying the regulatory response during the crisis, we provide evidence

that the agreement on the naked CDS ban exerted a containment effect on the transmission of CDS-bond contagion within the EMU South.

A clear implication from our analysis concerns the management of the financial institutions' exposures, which should correspond to the nature and channels of contagion during crises. In addition to first-order/direct exposures, regulation should encourage financial institutions to also manage second-order risks, such as those related to intra- and cross-asset correlations, particularly when being largely exposed to sovereign debt issues by countries under fiscal strain. To this end, the identification of the extent to which contagion phenomena depend on bilateral and multilateral exposures between countries would be of interest. We leave this to future research.

References

- Acharya, V.V., Drechsler, I. and Schnabl, P. (2014). A pyrrhic victory? Bank bailouts and Sovereign credit risk. *Journal of Finance*, 69, 2689–2739.
- Antón, M., Mayordomo, S. and Rodriguez-Moreno, M. (2018). Dealing with dealers: Sovereign CDS comovements in Europe. *Journal of Banking & Finance*, 90, 96–112.
- Ang, A. and Bekaert, G. (2002). International Asset Allocation With Regime Shifts. *The Review of Financial Studies*, 15, 1137–1187.
- Ang, A. and Chen, J. (2002). Asymmetric Correlations of Equity Portfolios. *Journal of Financial Econometrics*, 63, 443–494.
- Arce, O., Mayordomo, S. and Peña, J. I. (2013). Credit-risk valuation in the sovereign CDS and bonds markets: Evidence from the euro area crisis. *Journal of International Money and Finance*, 35, 124–145.
- Bae, K. H., Karolyi, G. A. and Stulz, R. M. (2003). A New Approach to Measuring Financial Contagion. *The Review of Financial Studies*, 16, 717–763.
- Beber, A., Brandt, M. W. and Kenneth, A. K. (2009). Flight-to-quality or flight-to-liquidity? Evidence from the euro-area bond market. *Review of Financial Studies*, 22, 925–957.
- Beirne, J. and Fratzscher, M. (2013). The pricing of sovereign risk and contagion during the European sovereign debt crisis. *Journal of International Money and Finance*, 34, 60–82.
- Bekaert, G. and Hodrick, J. R. (1992). Characterizing Predictable Components in Excess Returns on Equity and Foreign Exchange Markets. *Journal of Finance*, 467–509.
- Bekaert, G., Harvey, C. and Ng, A. (2005). Market integration and contagion. *Journal of Business*, 78, 39–69.
- Bekaert, G., Ehrmann, M., Fratzscher, M. and Mehl, A. (2014). The Global Crisis and Equity Market Contagion. *The Journal of Finance*, 69, 2597–2649.
- Bekaert, G. and Hoerova, M. (2016). What do asset prices have to say about risk appetite and uncertainty?. *Journal of Banking & Finance*, 67, 103–118.
- Berentsen, G. D. and Tjøstheim, D. (2014). Recognizing and visualizing departures from independence in bivariate data using local Gaussian correlation. *Statistics and Computing*, 24, 785–801.
- Berentsen, G. D., Kleppe, T. S. and Tjøstheim, D. (2014). Introducing local-gauss, an R package for estimating and visualizing local Gaussian correlation. *Journal of Statistical Software*, 56(1), 1–18.
- Bhansali, V., Gingrich, R. and Longstaff, F. A. (2008). Systemic credit risk: what is the market telling us. *Financial Analysts Journal*, 64, 16–24.
- Blatt, D., Candelon, B. and Hans Manner, H. (2015). Detecting contagion in a multivariate time series system: An application to sovereign bond markets in Europe. *Journal of Banking & Finance*, 59, 1–13.
- Broto, C. and Pérez-Quirós, G. (2015). Disentangling contagion among sovereign CDS spreads during the European debt crisis. *Journal of Empirical Finance*, 32, 165–179.
- Brunnermeier, M. K. (2009). Deciphering the liquidity and credit crunch 2007–2008. *Journal of Economic Perspectives*, 23, 77–100.
- Brunnermeier, M. K. and Pedersen, L. H. (2009). Market liquidity and funding liquidity. *Review of Financial Studies*, 22, 2201–2238.
- Caporin, M., Pelizzon, L., Ravazzolo, F. and Rigobon, R. (2018). Measuring sovereign contagion in Europe. *Journal of Financial Stability*, 34, 150–181.
- Claeys, P. and Vašíček, B. (2014). Measuring bilateral spillover and testing contagion on sovereign bond markets in Europe. *Journal of Banking & Finance*, 46, 151–165.
- D’Agostino, A. and Ehrmann, M. (2014). The pricing of G7 sovereign bond spreads—The times, they are a-changin. *Journal of Banking & Finance*, 47, 155–176.
- De Bruyckere, V., Gerhardt, M., Schepens, G. and Vander Vennet, R. (2013). Bank/sovereign risk spillovers in the European debt crisis. *Journal of Banking & Finance*, 37, 4793–4809.
- De Santis, R.A. (2012). The Euro area sovereign debt crisis: Safe haven, credit rating agen-

- cies and the spread of the fever from Greece, Ireland and Portugal. Tech. rep., European Central Bank.
- Demetrescu, M. and Wied, D. (2019) Testing for constant correlation of filtered series under structural change. *The Econometrics Journal*, 22, 10–33.
- Dungey, M., Fry, R., González-Hermosillo, B. and Martin, V. L. (2005). Empirical modelling of contagion: a review of methodologies. *Quantitative Finance*, 5, 9–24.
- Engle, R.F., 1982. Autoregressive conditional heteroskedasticity with estimates of the variance of United Kingdom inflation. *Econometrica*, 50, 987–1007.
- Eser, F. and Schwab, B. (2016). Evaluating the impact of unconventional monetary policy measures: Empirical evidence from the ECB’s Securities Markets Programme. *Journal of Financial Economics*, 119, 147–167.
- Favero, C., Pagano, M. and Von Thadden, E. L. (2010). How does liquidity affect government bond yields?. *Journal of Financial and Quantitative Analysis*, 45(1), 107–134.
- Fontana, A. and Scheicher, M. (2016). An analysis of euro area sovereign CDS and their relation with government bonds. *Journal of Banking & Finance*, 62, 126–140.
- Forbes, K. and Rigobon, R. (2002). No contagion, only interdependence: Measuring stock markets comovements. *The Journal of Finance*, 57, 2223–2261.
- Forbes, K. (2012). The ‘Big C’: Identifying contagion. NBER Working Paper No. 18465.
- Galeano, P. and Wied, D. (2014). Multiple break detection in the correlation structure of random variables. *Computational Statistics and Data Analysis*, 76, 262–282.
- Goyenko, R. Y. and Ukhov, A. D. (2009). Stock and bond market liquidity: A long-run empirical analysis. *Journal of Financial and Quantitative Analysis*, 44.1, 189–212.
- Groba, J., Lafuente, J. A. and Serrano, P. (2013). The impact of distressed economies on the EU sovereign market. *Journal of Banking & Finance*, 37, 2520–2532.
- Gyntelberg, J., Hördahl, P., Ters, K. and Urban, J. (2018). Price discovery in euro area sovereign credit markets and the ban on naked CDS. *Journal of Banking & Finance*, 96, 106–125.
- Hall, P. and Hart, J. D. (1990). Bootstrap test for difference between means in nonparametric regression. *Journal of the American Statistical Association*, 85, 1039–1049.
- Hall, P., Racine, J. S. and Li, Q. (2004). Cross-validation and the estimation of conditional probability densities. *Journal of the American Statistical Association*, 99, 1015–1026.
- Hamilton, James D. (1994a). Time Series Analysis, Princeton University Press.
- Hamilton, James D. (1994b). “State Space Models,” Chapter 50 in Robert F. Engle and Daniel L. McFadden (eds.), *Handbook of Econometrics*, Volume 4, Amsterdam: Elsevier Science B.V.
- Han, H., Linton, O., Oka, T. and Whang, W.-J. (2016). The CrossQuantilogram: Measuring Quantile Dependence and Testing Directional Predictability between Time Series. *Journal of Econometrics*, 193, 251–270.
- Hartmann, P., Straetmans, S. and de Vries, C. G. (2004). Asset Market Linkages in Crisis Periods. *The Review of Economics and Statistics*, 86, 313–326.
- Harvey, Andrew C. (1989). Forecasting, Structural Time Series Models and the Kalman Filter, Cambridge: Cambridge University Press.
- Hjort, N. L. and Jones, M. C. (1996). Locally parametric nonparametric density estimation. *Annals of Statistics*, 24, 1619–1647.
- Hong, Y., Tu, J. and Zhou, G. (2007). Asymmetries in stock returns: statistical tests and economic evaluation. *Review of Financial Studies*, 20, 1547–1581.
- Kalbaska, A. and Gatkowski, M. (2012). Eurozone sovereign contagion: Evidence from the CDS market (2005–2010). *Journal of Economic Behavior & Organization*, 83, 657–673.
- King, M. and Wadhwani, S. (1990). Transmission of volatility between stock markets. *Review of Financial Studies*, 3, 5–33.
- Koopman, Siem Jan, Neil Shephard, and Jurgen A. Doornik (1999). “Statistical Algorithms for Models in State Space using SsfPack 2.2,” *Econometrics Journal*, 2(1), 107–160.
- Lacal, V. and Tjøstheim, D. (2017). Local Gaussian autocorrelation and tests of serial de-

- pendence. *Journal of Time Series Analysis*, 38(1):51–71.
- Longin, F. and Solnik, B. (2001). Extreme correlations in international equity markets. *The Journal of Finance*, 56, 649–676.
- Longstaff, F.A., Pan, J., Pedersen, L. and Singleton, K. (2011). How sovereign is sovereign risk? *American Economic Journal: Macroeconomics*, 3, 75–103.
- Metiu, N. (2012). Sovereign risk contagion in the Eurozone. *Economics Letters*, 117, 35–38.
- Mink, M. and De Haan, J. (2013). Contagion during the Greek sovereign debt crisis. *Journal of International Money and Finance*, 34, 102–113.
- Missio, S. and Watzka, S. (2011). Financial contagion and the European debt crisis. CESifo, Munich, Germany. *Working Paper* 3554.
- Monfort, A., and Renne, J.-P. (2014). Decomposing euro-area sovereign spreads: credit and liquidity risks. *Review of Finance*, 18, 2103–2151.
- Pan, J. and Singleton, K. J. (2008). Default and recovery implicit in the term structure of sovereign CDS spreads. *The Journal of Finance*, 63, 2345–2384.
- Pericoli, M. and Sbracia, M. (2003). A primer on financial contagion. *Journal of Economic Surveys*, 17, 571–608.
- Oh, D. H. and Patton A. J. (2017). Modeling Dependence in High Dimensions With Factor Copulas. *Journal of Business & Economic Statistics*, 35, 139–154.
- Rodriguez-Moreno, M. and Peña, J. I. (2013). Systemic risk measures: the simpler the better? *Journal of Banking and Finance*, 37, 1817–1831.
- Silva, P.P., Vieira, C. and Vieira, I. V. (2016). The EU ban on uncovered sovereign credit default swaps: Assessing impacts on liquidity, volatility, and price discovery. *The Journal of Derivatives*, 23, 74–98.
- Støve, B., Tjøstheim, D. and Hufthammer, K. (2014). Using local Gaussian correlation in a nonlinear re-examination of financial contagion. *Journal of Empirical Finance*, 25, 62–82.
- Székely, G. J. and Rizzo, M. L. (2009). Brownian Distance Correlation. *Annals of Applied Statistics*, 3, 1236–1265.
- Teräsvirta, T., Tjøstheim, D. and Granger, C. W. (2010). *Modelling Nonlinear Economic Time Series*. Oxford University Press.
- Tjøstheim, D. and Hufthammer, K. (2013). Local Gaussian correlation: a new measure of dependence. *Journal of Econometrics*, 172, 33–48.
- Vilar-Fernandez, J. M., Vilar-Fernandez, J. A. and Gonzalez-Manteiga, W. (2007). Bootstrap tests for nonparametric comparison of regression curves with dependent errors. *Test*, 16, 123–144.
- Wied, D., Arnold, M., Bissantz, N. and Ziggel, D. 2012. A new fluctuation test for constant variances with applications to finance. *Metrika*, 75, 1111–27.
- Zhou, Z., 2012. Measuring Nonlinear Dependence in Time-Series, a Distance Correlation Approach. *Journal of Time Series Analysis*, 33, 438–457.

Table 1: This table reports means, standard deviations, maximums and minimums for bond yield and CDS spread changes (in basis points). Bond yields is the yield on 5-year on-the-run sovereign bonds. CDS spread is the change in 5-year sovereign CDSs. The sample includes Austria, Belgium, Finland, France, Ireland, Italy, Germany, Hellenic Republic (Greece), the Netherlands, Portugal, Spain. Bond yields are from Datastream, and CDS spreads are from Markit. The sample period for Greece is January 2, 2006 to March 8, 2012.

	Δ (Bond yields)				Δ (CDS spreads)				Obs.#
	Mean	Std. Dev	Max.	Min.	Mean	Std. Dev	Max.	Min.	
AT	-0.119	4.41	37.32	-34.43	0.003	4.23	44.28	-28.11	3090
BE	-0.122	4.79	43.77	-50.30	0.003	5.38	40.55	-57.38	3090
DE	-0.124	3.73	19.12	-22.51	0.002	2.41	20.81	-20.27	3090
ES	-0.113	7.08	60.51	-88.94	0.008	10.94	96.88	-102.66	3090
FI	-0.121	4.63	51.17	-41.18	0.003	2.64	25.49	-25.73	3090
FR	-0.119	3.93	34.54	-46.41	0.003	4.66	46.79	-44.93	3090
GR	2.420	45.86	516.57	-551.57	15.398	643.86	9025.77	-9888.18	1505
IE	-0.120	10.86	102.04	-148.00	0.004	11.48	113.79	-152.45	3090
IT	-0.102	7.25	69.20	-91.05	0.018	16.32	162.73	-161.76	3090
NL	-0.122	5.77	85.29	-89.16	0.014	29.56	274.84	-275.94	3090
PT	-0.108	18.30	409.19	-222.46	0.002	3.67	79.10	-74.23	3090
GIIPS	0.395	17.87	231.50	-220.40	3.086	137.26	1895.65	-2075.85	
Core	-0.121	4.54	45.20	-47.33	0.005	8.15	75.46	-75.39	

Table 2: This table presents the changes in basis points on 5-year sovereign bond yields and CDS spreads after a specific adverse event in GIIPS (Greece, Ireland, Italy, Portugal and Spain). (1) On November 5, 2009, the Greek government revealed a revised budget deficit of 12.7% of GDP for 2009. (2) On January 15, 2009, Ireland abandons plans to inject €1.5 bn into third largest bank Anglo Irish Bank and nationalises the commercial lender amid fears it could collapse. (3) On April 18, 2012, the Italian government cut its growth forecast for the economy in 2012, predicting a further shrink by 0.8%. (4) On April 6, 2011, Portugal requests a bailout from the EU. (5) On April 27, 2012, the rating agency Standard and Poor's has lowered the rating of Spain by two notches to BBB+ and maintained a negative outlook. The last two columns present the correlation coefficient between changes on bond yields and CDS spreads in basis points before and after a specific event.

													Correlation	
													$\Delta(\text{Bond})$ vs $\Delta(\text{CDS})$	
													pre-event	post-event
Panel A: $\Delta(\text{Bond})$														
GR	(1) Budget revision	0.04	0.08	-1.06	1.75	0.08	-2.16	40	-2.33	-2.51	-3.94	5.3	0.36	0.59
IE	(2) Rescue AIB	-3.29	-1.74	-5.54	-1.27	-0.95	1.22	-3.39	5.32	-5.94	4.29	-2.52	-0.13	0.55
IT	(3) Growth forecast revision	-1.64	-2.86	-2	-8.83	-2.89	-0.7	-4.77	-0.51	-0.25	-2.33	-18.45	0.57	0.72
PT	(4) Bailout request	0.79	1.27	1.42	0.38	1.36	1.24	-2.53	-21.43	-0.67	1.61	10.21	0.48	0.51
ES	(5) S&P downgrade	-1.85	-0.33	0.55	11.65	-2.16	0.26	-15.32	0.51	4.54	-2.12	-62.04	0.53	0.75
Panel B: $\Delta(\text{CDS})$														
GR	(1) Budget revision	2.99	8.66	6.80	23.74	2.01	6.94	127.69	24.15	20.67	3.36	75.53		
IE	(2) Rescue AIB	4.43	-0.15	2.10	3.27	-0.17	1.35	5.69	35.54	1.48	2.11	1.24		
IT	(3) Growth forecast revision	4.54	9.72	4.50	6.56	2.52	1.10	//	5.25	18.99	6.79	-0.68		
PT	(4) Bailout request	-0.74	-2.54	-0.74	-3.00	-0.91	-1.47	2.31	-15.46	-0.28	-0.58	10.94		
ES	(5) S&P downgrade	-3.42	-2.67	-1.60	4.31	-0.82	-1.86	//	-4.84	0.07	-4.66	-37.97		

Table 3: Bootstrap test for contagion considering GIIPS as the countries of origin. This table shows p -values from the bootstrap test for contagion according to the time dates of correlation change-points estimated with the algorithm of [Demetrescu and Wied \(2019\)](#) based on the residuals from the Kalman filter model. The null hypothesis indicates no contagion between the GIIPS (GR, IE, IT, PT, ES) bonds and the European bond (Panel A) and CDS (Panel B) markets. The bootstrap test for contagion is based on 5000 replications. The sample period starts on January 2, 2006 and ends on March 8, 2012 for Greece and on April 5, 2018 for the rest EMU South countries.

	GR (1)	Date	IE (2)	Date	IT (3)	Date	PT (4)	Date	ES (5)	Date
<i>Panel A : Bond Markets</i>										
AT	0.995	13/10/2008	0.099	29/1/2009	0.612	20/7/2011	0.999	9/10/2009	0.694	8/12/2009
	0.987	30/11/2009	0.871	25/3/2010			0.677	22/3/2010	0.230	11/6/2012
BE	0.993	6/8/2009	0.450	3/3/2010			0.451	12/4/2011		
DE	0.998	16/10/2008	0.958	25/2/2009	0.997	22/10/2008	0.999	26/9/2008	0.853	29/12/2008
	0.962	26/11/2009	0.348	2/12/2010	0.995	27/4/2010	0.954	26/11/2009	0.768	3/8/2011
					0.965	7/7/2011				
ES	0.412	23/12/2009	0.045	1/12/2010	0.000	20/7/2011	0.013	24/3/2011		
					0.887	30/6/2015	0.899	3/7/2015		
FI	0.975	26/10/2009	0.931	2/12/2010	0.999	2/7/2008	0.985	22/4/2011	0.974	3/12/2009
			0.812	5/3/2014	0.996	19/5/2010			0.063	20/7/2011
FR	0.870	16/11/2009	0.765	1/4/2010	0.986	18/7/2011	0.956	29/1/2010	0.350	14/7/2009
			0.631	9/12/2010			0.349	20/4/2011	0.512	18/7/2011
GR			0.112	9/12/2009	0.371	15/10/2008	0.350	17/2/2009	0.630	23/12/2009
					0.968	23/11/2009				
IE	0.034	9/12/2009					0.980	30/11/2010	0.154	1/12/2010
IT	0.941	15/10/2008					0.662	10/2/2010	0.012	20/7/2011
	0.855	23/11/2009					0.635	20/5/2011	0.312	30/6/2015
NL	0.991	14/10/2009	0.978	23/1/2009	0.830	24/6/2011	0.891	30/9/2009	0.794	21/6/2011
							0.208	1/4/2011		
PT	0.944	17/12/2009	0.750	30/11/2010	0.332	10/2/2010			0.614	24/3/2011
					0.428	20/5/2011			0.431	3/7/2015
<i>Panel B : CDS Markets</i>										
AT	0.039	2/11/2009	0.035	15/11/2010	0.008	18/7/2011	0.000	28/3/2011	0.000	2/8/2011
BE	0.000	6/11/2009	0.000	25/11/2010	0.019	22/7/2011	0.000	1/4/2011	0.000	3/8/2011
	0.000	21/4/2010								
DE	0.004	5/11/2009	0.041	23/11/2010	0.033	20/7/2011	0.000	18/4/2011	0.032	20/7/2011
	0.662	22/04/2010								
ES	0.000	9/11/2009	0.006	2/12/2010	0.000	20/7/2011	0.000	11/4/2011		
	0.000	22/4/2010	0.540	25/07/2011						
FI	0.000	13/11/2009			0.002	2/8/2011	0.037	4/5/2011	0.000	3/8/2011
FR	0.002	5/11/2009	0.001	12/11/2010	0.000	20/7/2011	0.014	18/4/2011	0.000	18/7/2011
	0.431	21/04/2010								
GR			0.269	2/1/2008	0.212	6/11/2009	0.961	17/11/2009	0.881	9/11/2009
			0.057	12/11/2009	0.571	20/4/2010	0.412	22/4/2010	0.417	22/4/2010
IE	0.502	2/1/2008			0.615	3/12/2010	0.312	30/11/2010	0.770	2/12/2010
	0.009	12/11/2009			0.000	15/7/2011	0.041	29/3/2011	0.000	25/7/2011
IT	0.000	6/11/2009	0.008	3/12/2010			0.000	12/4/2011	0.000	20/7/2011
	0.000	20/4/2010	0.415	15/7/2011						
NL	0.083	5/11/2009	0.000	10/12/2010	0.016	2/8/2011	0.012	4/5/2011	0.016	18/7/2011
PT	0.001	17/11/2009	0.000	30/11/2010	0.000	12/4/2011	0.000	5/5/2011	0.045	11/4/2011
	0.000	22/4/2010	0.565	29/3/2011						

Table 4: Bootstrap test for contagion considering GIIPS as the countries of origin. This table shows p -values from the bootstrap test for contagion according to the time dates of correlation change-points estimated with the algorithm of [Demetrescu and Wied \(2019\)](#) based on the residuals from the Kalman filter model. The null hypothesis indicates no contagion between the GIIPS (GR, IE, IT, PT, ES) CDS and the European bond (Panel A) and CDS (Panel B) markets. The bootstrap test for contagion is based on 5000 replications. The sample period starts on January 2, 2006 and ends on March 8, 2012 for Greece and on April 5, 2018 for the rest EMU South countries.

	GR (1)	Date	IE (2)	Date	IT (3)	Date	PT (4)	Date	ES (5)	Date
<i>Panel A : Bond Markets</i>										
AT	0.233	2/11/2009	0.148	13/10/2011	0.912	31/10/2011			0.064	25/10/2011
	0.411	4/11/2011							0.000	21/11/2011
BE	0.000	3/11/2011	0.000	14/11/2011			0.137	30/11/2010		
DE	0.872	22/4/2010	0.341	12/10/2011	0.832	26/9/2008	0.985	26/9/2008	0.946	26/9/2008
					0.307	25/10/2011	0.204	15/5/2013	0.742	26/2/2010
ES	0.018	8/11/2011	0.000	15/11/2011	0.000	9/11/2011	0.885	6/4/2010		
			0.996	7/8/2015	0.991	22/6/2015	0.000	16/11/2011		
FI	0.311	13/11/2009	0.962	9/1/2008	0.865	23/4/2010	0.994	10/9/2008	0.807	23/4/2010
	0.411	15/11/2011	0.981	23/4/2010	0.411	21/10/2013	0.962	15/5/2013	0.639	15/5/2013
FR	0.452	21/4/2010	0.948	17/7/2008	0.757	13/11/2009	0.912	26/9/2008	0.720	6/4/2010
			0.127	10/10/2011	0.412	18/11/2011	0.002	2/8/2011	0.518	8/11/2011
GR			0.011	9/11/2011	0.000	10/11/2011	0.011	9/11/2011	0.000	8/11/2011
IE	0.009	9/11/2011			0.012	12/10/2010	0.000	18/11/2011	0.050	15/11/2011
					0.997	17/11/2011			0.619	7/8/2015
IT	0.000	10/11/2011	0.912	12/10/2010			0.000	14/11/2011	0.000	9/11/2011
			0.003	17/11/2011					0.004	22/6/2015
NL	0.981	5/11/2009	0.965	1/7/2008	0.992	23/1/2009	0.975	10/9/2008	0.963	26/9/2008
	0.757	7/11/2011	0.438	3/8/2010	0.659	26/3/2013			0.589	26/3/2013
PT	0.046	9/11/2011	0.000	18/11/2011	0.018	14/11/2011			0.589	6/4/2010
									0.000	16/11/2011
<i>Panel B : CDS Markets</i>										
AT			0.992	4/1/2013	0.001	22/11/2011	0.822	15/9/2008	0.988	15/9/2008
							0.996	1/2/2012	0.021	3/1/2012
BE			0.994	28/1/2012	0.005	10/1/2012	0.963	13/1/2012	0.033	9/1/2012
DE			0.788	13/1/2012	0.000	17/1/2012	0.975	15/9/2008	0.019	19/1/2012
							0.992	13/1/2012		
ES	0.034	16/11/2011			0.010	22/11/2011	0.987	31/10/2011		
FI	0.411	25/9/2008	0.899	19/1/2012	0.001	30/11/2011	0.641	2/1/2012	0.010	14/11/2011
FR					0.005	10/1/2012	0.418	12/11/2009	0.011	5/1/2012
					0.966	1/2/2013	0.847	13/1/2012	0.968	1/2/2013
GR			0.411	11/11/2011					0.000	18/11/2011
IE	0.015	9/11/2011			0.003	16/11/2011	0.872	5/1/2012		
					0.958	31/8/2015				
IT	0.983	14/9/2010	0.672	15/11/2011			0.868	14/11/2012	0.041	22/11/2011
			0.784	31/8/2015			0.994	3/9/2012	0.885	
NL	0.311	22/9/2008	0.760	7/12/2012	0.033	19/1/2012	0.478	5/9/2008	0.912	15/9/2008
	0.471	22/9/2010					0.988	13/1/2012	0.020	16/1/2012
PT			0.967	5/1/2012	0.000	15/11/2012			0.017	7/11/2011

Table 5: This table presents the estimated p -values from the bootstrap test for contagion from the Greek bond yields to the rest of the European sovereign CDS markets. The first test refers to the period from January 2, 2006 until April 23, 2010, with November 16, 2009 as break-date. The second test refers to the period from November 16, 2009 until March 8, 2012 (the last day of trading for Greek government bonds), with April 23, 2010 as break-date. Significance levels at 10%, 5% and 1% are denoted by *, **, ***. Yes indicates that the null of no contagion is rejected at 5% level. The bootstrap test for contagion is based on 5000 replications.

Origin:Greece	Deficit Revision November 2009		Financial Request April 2010	
	<i>Bonds</i> (GR)	Contagion?	<i>Bonds</i> (GR)	Contagion?
<i>CDS Markets</i>				
Austria	0.002***	Yes	0.391	No
Belgium	0.007***	Yes	0.845	No
Finland	0.001***	Yes	0.601	No
France	0.000***	Yes	0.559	No
Germany	0.008***	Yes	0.713	No
Ireland	0.041**	Yes	0.681	No
Italy	0.022**	Yes	0.215	No
Netherlands	0.007***	Yes	0.770	No
Portugal	0.011**	Yes	0.423	No
Spain	0.000***	Yes	0.718	No

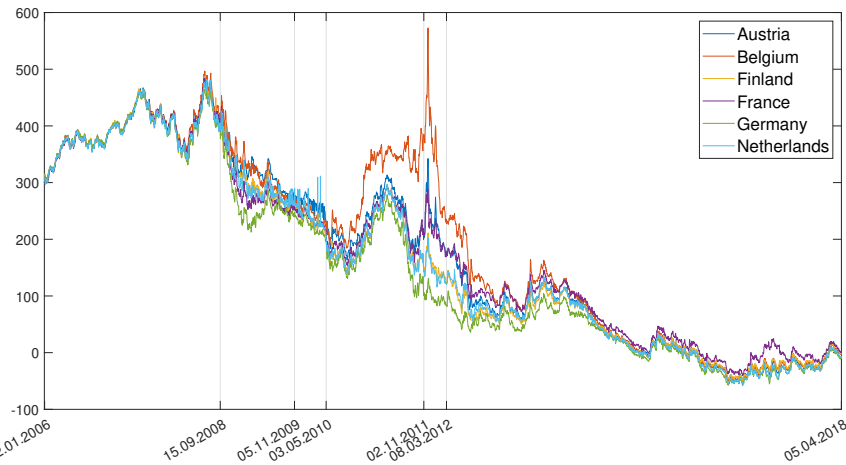
Table 6: This table presents the estimated p -values from the bootstrap test for contagion from the GIIPS CDS spreads to the rest of the European sovereign bond markets in the period before and after October 11, 2011 (the date when EU lawmakers reached an agreement on the naked CDS ban). The bootstrap test for contagion is based on 5000 replications.

	GR (1)	IE (2)	IT (3)	PT (4)	ES (5)
<i>Bond Markets</i>					
AT	0.607	0.741	0.451	0.711	0.899
BE	0.431	0.293	0.694	0.441	0.792
DE	0.213	0.575	0.923	0.242	0.891
ES	0.315	0.641	0.952	0.921	—
FI	0.471	0.488	0.435	0.578	0.398
FR	0.542	0.790	0.944	0.414	0.813
GR	—	0.412	0.214	0.545	0.931
IE	0.551	—	0.741	0.225	0.611
IT	0.625	0.899	—	0.418	0.704
NL	0.405	0.216	0.902	0.601	0.915
PT	0.490	0.290	0.331	—	0.444

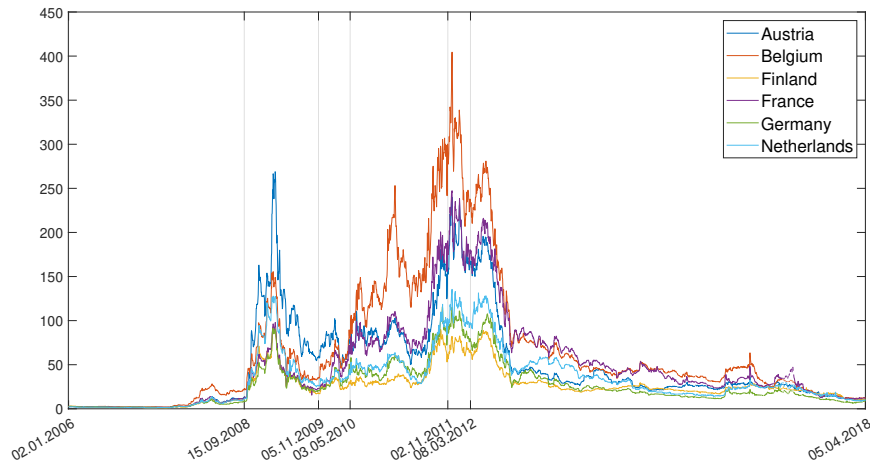
Table 7: Bootstrap test for contagion considering GIIPS as the countries of origin. This table shows p -values from the bootstrap test for contagion according to the time dates of correlation change-points estimated with the algorithm of Demetrescu and Wied (2019) based on the quoted bid-ask spreads. The null hypothesis indicates no contagion between the GIIPS (GR, IE, IT, PT, ES) bonds and the European bond (Panel A) and CDS (Panel B) markets. The bootstrap test for contagion is based on 5000 replications. The sample period starts on January 2, 2006 and ends on March 8, 2012 for Greece and on April 5, 2018 for the rest EMU South countries.

	GR (1)	Date	IE (2)	Date	IT (3)	Date	PT (4)	Date	ES (5)	Date
<i>Panel A : Bond Markets</i>										
AT	0.930	7/11/2011					0.775	22/9/2008		
BE	0.910	6/10/2009	0.665	4/1/2010	0.885	29/6/2011	0.491	6/10/2010	0.950	22/06/2011
DE	0.811	2/10/2009	0.678	14/12/2009	0.393	7/7/2011			0.189	5/7/2011
ES	0.740	17/09/2009	0.655	18/3/2010	0.545	13/7/2011	0.311	12/1/2011		
FI	0.910	3/11/2011	0.732	10/2/2010						
FR	0.644	21/09/2009	0.913	5/3/2010			0.778	11/11/2009		
GR			0.095	14/10/2009	0.540	29/6/2011	0.673	22/09/2009	0.499	17/09/2009
IE	0.633	14/10/2009			0.423	29/3/2010	0.998	28/1/2009	0.312	18/3/2010
IT	0.970	29/6/2011	0.728	2/2/2010	0.477	16/11/2009	0.555	25/4/2011	0.611	13/7/2011
NL	0.987	24/11/2011	0.480	10/12/2009	0.790	10/5/2010	0.864	6/10/2009		
PT	0.892	22/09/2009	0.900	28/1/2009	0.425	25/4/2011			0.385	12/1/2011
<i>Panel B : CDS Markets</i>										
AT	0.999	12/10/2009	0.435	25/11/2010	0.620	21/7/2011	0.895	21/3/2011	0.784	25/7/2011
BE	0.260	20/11/2009					0.970	30/3/2011	0.720	15/8/2011
DE	0.515	30/11/2009	0.628	8/11/2010	0.984	11/7/2011	0.950	5/4/2011	0.881	15/8/2011
ES	0.003	18/11/2009	0.025	25/11/2010	0.006	3/8/2011	0.000	14/4/2011		
FI	0.644	24/11/2009	0.396	2/11/2010			0.412	28/4/2011	0.998	12/8/2011
FR	0.661	23/11/2009			0.714	4/7/2011			0.411	2/8/2011
GR			0.003	29/10/2009	0.006	3/11/2009	0.021	11/11/2009	0.031	17/11/2009
IE	0.002	29/10/2009			0.033	22/6/2011	0.030	8/12/2010	0.025	25/11/2010
IT	0.006	3/11/2009	0.020	10/12/2010			0.021	20/7/2011	0.001	2/8/2011
NL	0.521	19/11/2009	0.475	1/12/2010	0.910	8/7/2011	0.226	9/5/2011	0.964	20/7/2011
PT	0.035	11/11/2009	0.007	24/11/2010	0.003	19/7/2011			0.012	12/4/2011

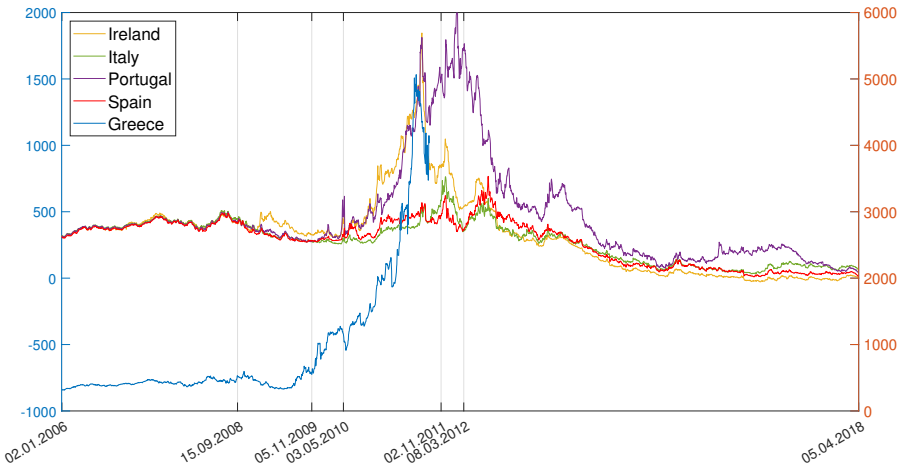
(A) EMU North 5-year sovereign bond yields



(C) EMU North 5-year sovereign CDS spreads



(B) EMU South 5-year sovereign bond yields



(D) EMU South 5-year sovereign CDS spreads

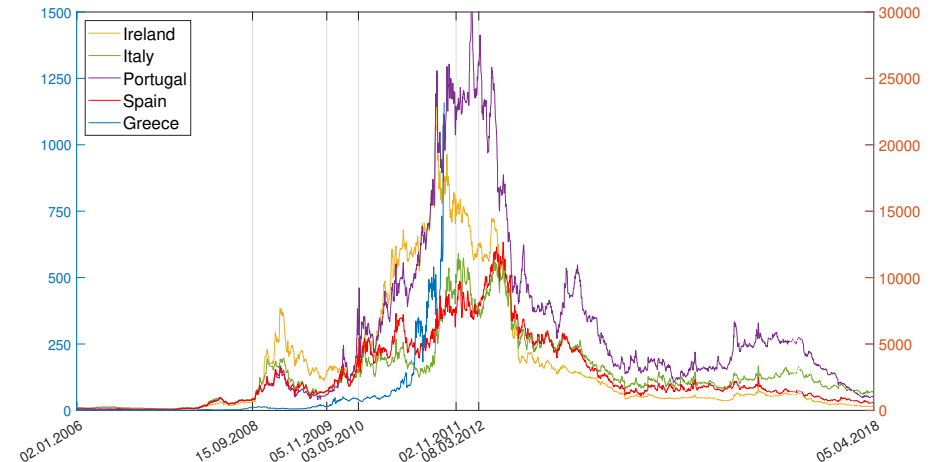


Figure 1: Evolution of sovereign bond yields and CDS spreads (in basis points). Panels A and C show the 5-year sovereign bond yields and CDS spreads for the countries of the EMU North (Austria, Belgium, Finland, France, Germany, Netherlands) and Panels B and D show the 5-year sovereign bond yields and CDS spreads for the countries of the EMU South (Greece, Ireland, Italy, Portugal, Spain). In Panels B and D, Greece is shown on the right axis while the rest EMU South countries are shown on the left axis.

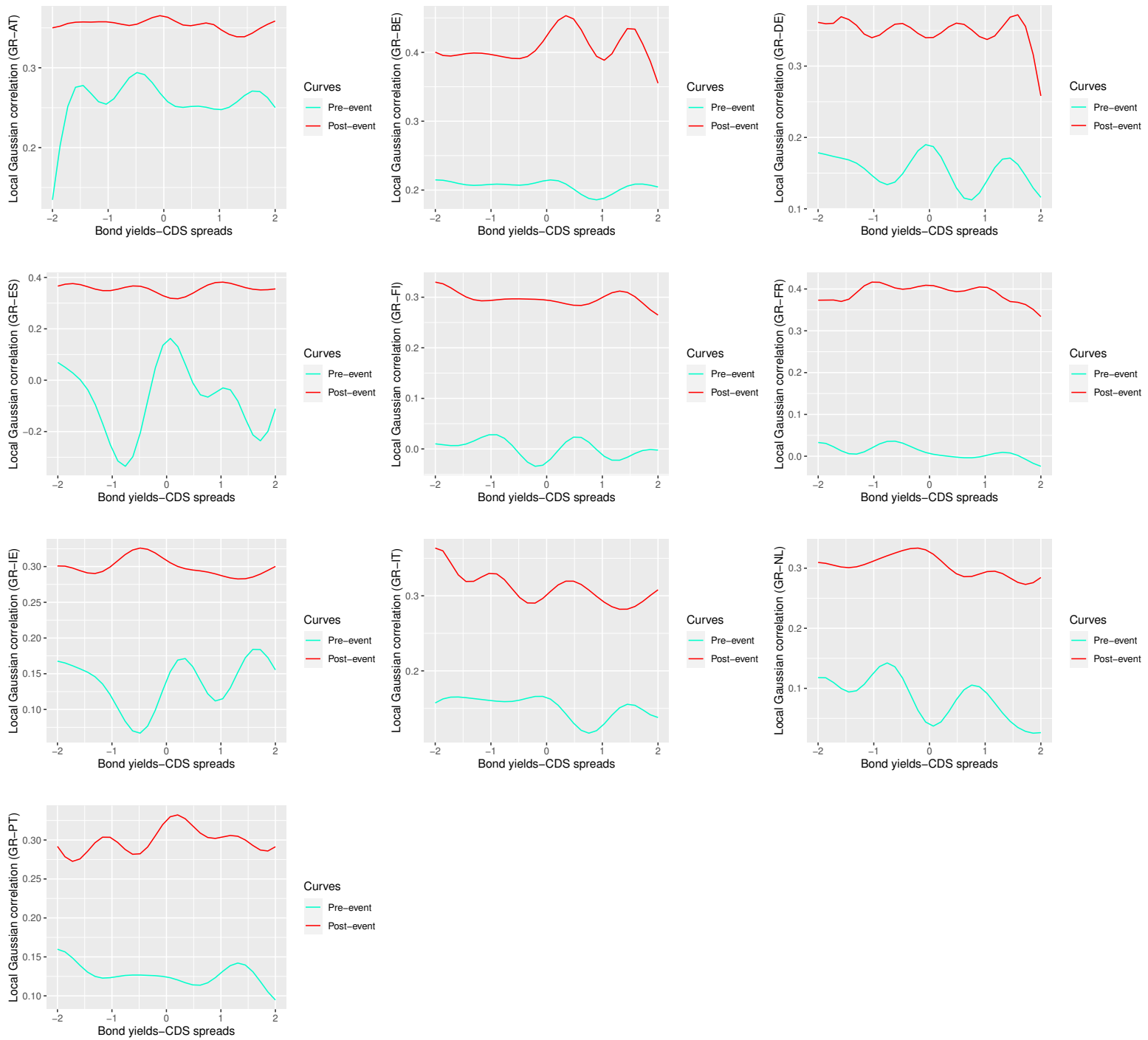


Figure 2: Local Gaussian correlation estimates between the Greek sovereign bond yields and each of the European sovereign CDS yields following Greece's upward deficit revision on November 16, 2009.

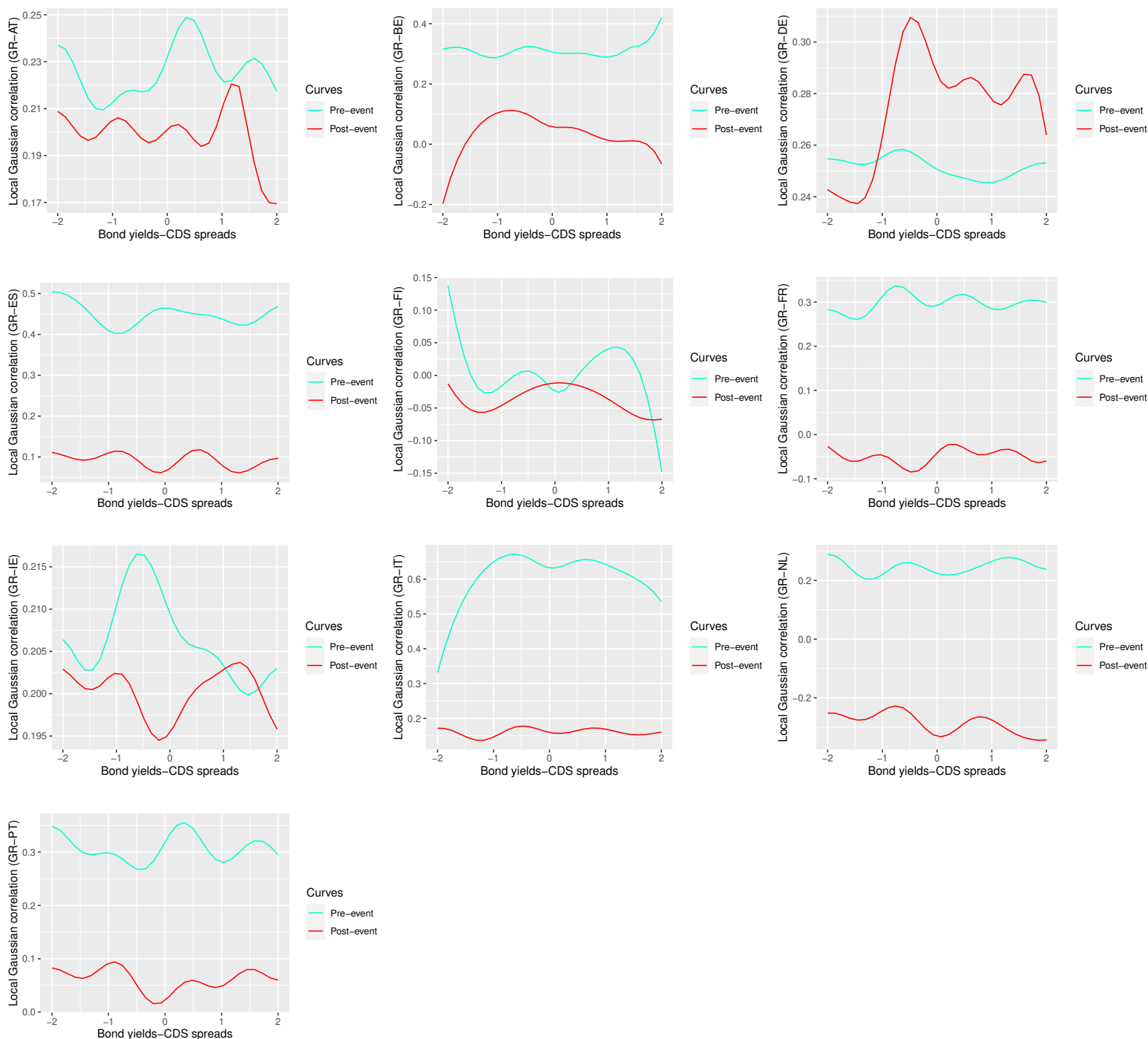


Figure 3: Local Gaussian correlation estimates between the Greek sovereign bond yields and each of the European sovereign CDS yields following Greece's request for financial assistance in April 23, 2010.

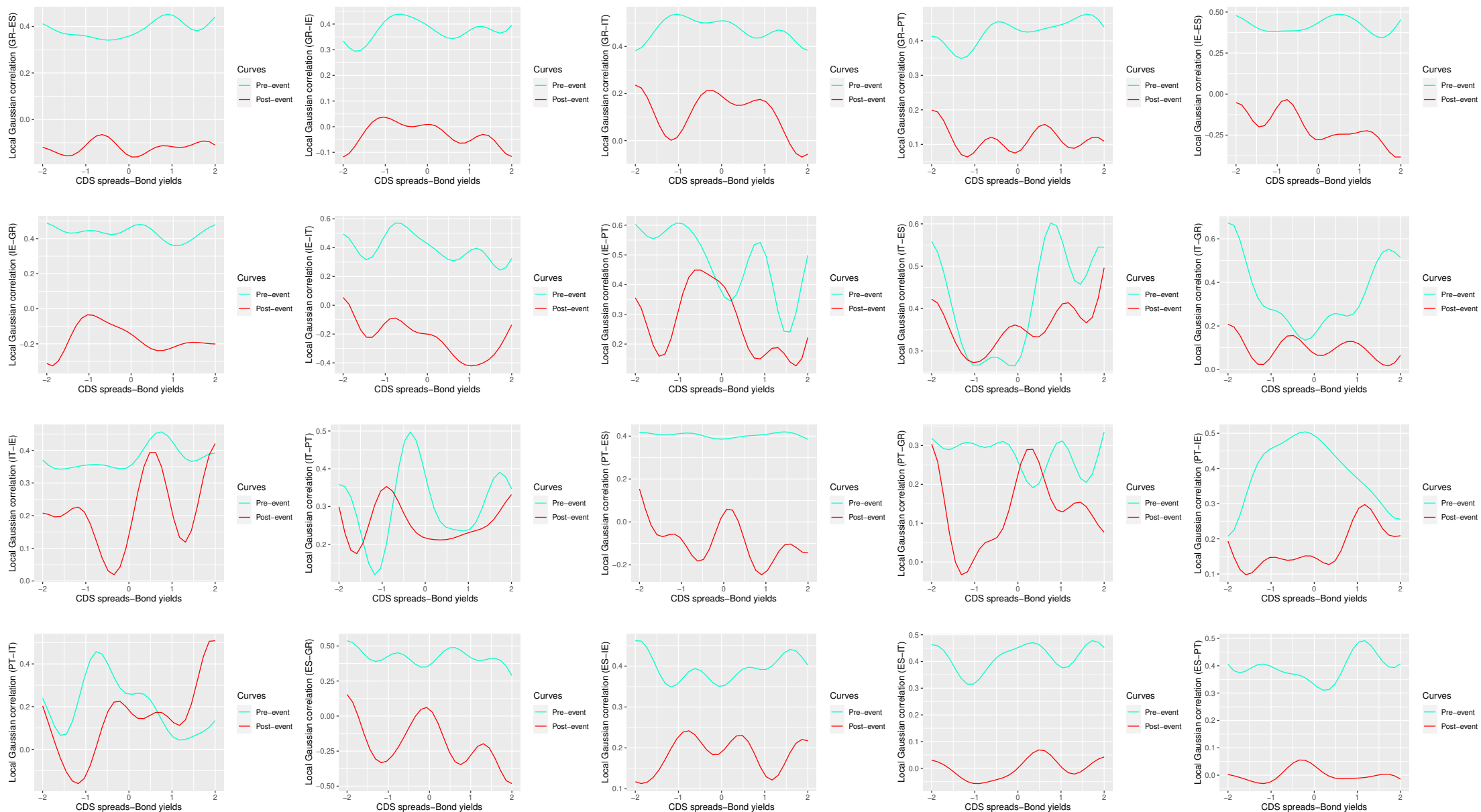


Figure 4: Local Gaussian correlation estimates between the GIIPS CDS spreads and each of the European sovereign bond yields following the EU’s agreement on the naked CDS ban in October 11, 2011.

A Appendix

A.1 Demetrescu and Wied (2019) algorithm for the identification of structural changes in correlation

Suppose Y_t is a sequence of random variables for $t = 1, 2, \dots$ with finite absolute moments. Our purpose is to test whether the variance of Y_t is constant over time.

For example we test, $H_0 : Var(Y_t) = \sigma^2 \forall t \in T$ against $H_1 : \exists t \in \{1, \dots, T-1\} : Var(Y_t) \neq Var(Y_{t+1})$ for a constant σ^2 . The test statistic is then formed as,

$$Q_T(Y) = \max_{1 \leq i \leq T} \left| \hat{D} \frac{i}{\sqrt{T}} ([VarY]_i - [VarY]_T) \right|$$

where the empirical variance is calculated from the first s observations as,

$$[VarY]_s = \frac{1}{s} \sum_{t=1}^s Y_t^2 - \left(\frac{1}{s} \sum_{t=1}^s Y_t \right)^2 = \overline{Y_s^2} - (\overline{Y_s})^2.$$

The scalar $\hat{D} = \left((1, -2\overline{Y_T}) \hat{D}_1 (1, -2\overline{Y_T})' \right)^{-1/2}$, which mainly captures the fluctuations resulting from estimating the expected values is necessary for the asymptotic null distribution. The test rejects the null hypothesis of constant variance in the case that the empirical variances fluctuate too much, given by:

$$\max_{1 \leq i \leq T} |[VarY]_i - [VarY]_T|.$$

The weighting factor $\frac{i}{\sqrt{T}}$ is scaling down deviations at the beginning of the sample when the $[VarY]_i$ are more volatile.

A transformation is needed in order to examine the large sample properties,

$$Q_T(Y) = \sup_{l \in [0,1]} \left| \hat{D} \frac{\tau(l)}{\sqrt{T}} ([VarY]_{\tau(l)} - [VarY]_T) \right| \text{ with } \tau(l) = [1 + l(T-1)]$$

Under the null hypothesis and assumptions A.1 – A.4 in Wied *et al.* (2012),

$$Q_T(Y) \rightarrow \sup_{l \in [0,1]} |B(l)|, \text{ where } B(l) \text{ is a one-dimensional Brownian Bridge.}$$

The quantiles and the limit distribution of $Q_T(Y)$ provide an asymptotic test.

In our paper we first apply the aforementioned test of Wied *et al.* (2012) in each series to take the estimated variance change point locations, in combination with a binary segmentation algorithm applied in a similar way as in Galeano and Weid (2014). Applied on the two time series, the test yields a variance change-point at January 2, 2008 for the Greek bond yields series and at November 5, 2007 for the German bond yield series. Next, the data are split into the interval before the change-point (including the point) and after so as to test in both segments again. If the smallest of the two p-values is smaller than $1-0.95^{1/2}$, then a new change point is found at the argmax of the corresponding series, which is split at this point again. The procedure is repeated until no further change points can be found. The variance breaks found are considered as fixed at the following the correlation break test.

The Demetrescu and Wied (2019) test for constant correlation employs a multivariate CUSUM statistic given by,

$$\hat{Q}_n = \max_{1 \leq j \leq n} \frac{j}{\sqrt{n}} \sqrt{(\hat{\mathbf{S}}_j - \hat{\mathbf{S}}_n)' \hat{\Omega}^{-1} (\hat{\mathbf{S}}_j - \hat{\mathbf{S}}_n)} \text{ with } \hat{\mathbf{S}}_j = \frac{1}{j} \sum_{t=1}^j \mathbf{g}(\hat{\mathbf{Z}}_t)$$

in combination with the robustified constant-correlation test based on \hat{Q}_n ,

$$g(\hat{Z}_{t1}, \hat{Z}_{t2}) = \hat{Z}_{t1} \hat{Z}_{t2} \text{ and } \hat{Z}_{ti} = \frac{X_{ti} - \hat{\mu}_{1,i}(1-D_{t,\lambda}) - \hat{\mu}_{2,i}D_{t,\lambda}}{\sqrt{\hat{\sigma}_{i,1}^2(1-D_{t,\lambda}) + \hat{\sigma}_{i,2}^2D_{t,\lambda}}}$$

Under some assumptions it holds that as $n \rightarrow \infty$,

$$\hat{Q}_n \xrightarrow{d} \sup_{s \in [0,1]} \sqrt{\left(\hat{\Gamma}(s) - s\hat{\Gamma}(1) \right)' \left(\hat{\Gamma}(s) - s\hat{\Gamma}(1) \right)} \text{ given that } \hat{\lambda} = \lambda_0 + O_p(n^{-1}), \text{ where}$$

$$\hat{\Gamma}(s) = \Gamma(s) + \Omega^{-\frac{1}{2}} \tau_{\lambda_0}(s) \left(\Pi'_{\lambda_0}(1) W \Pi_{\lambda_0}(1) \right)^{-1} \Pi'_{\lambda_0}(1) W \sum_{\lambda_0}^{1/2} \Theta(1)$$

The above test allows for a two regime model in the variances. Once we eliminate variance breaks stationarity is desirable under the null hypothesis of constant correlations. To perform the test we rely on a bootstrap approximation with 4,999 bootstrap replications.

A.2 Local Gaussian approximation and local correlation

Assume two return series with observed values $\{(X_t, Z_t) \mid t = 1, \dots, T\}$. The correlation between them conditionally on being in the region S can be written as:

$$\hat{\rho}_c(S) = \frac{\sum_{(X_t, Z_t) \in S} (X_t - \hat{\mu}_{X,c})(Z_t - \hat{\mu}_{Z,c})}{\left[\sum_{(X_t, Z_t) \in S} (X_t - \hat{\mu}_{X,c})^2 \right]^{\frac{1}{2}} \left[\sum_{(X_t, Z_t) \in S} (Z_t - \hat{\mu}_{Z,c})^2 \right]^{\frac{1}{2}}} \quad (4)$$

where $\hat{\mu}_{X,c} = \frac{1}{n_S} \sum_{(X_t, Z_t) \in S} X_t$ and $\hat{\mu}_{Z,c} = \frac{1}{n_S} \sum_{(X_t, Z_t) \in S} Z_t$ with n_S being the number of pairs with $(X_t, Z_t) \in S$. For ergodic series $\{X_t, Z_t\}$ as $n_S \rightarrow \infty$, $\hat{\rho}_c(S)$ would converge to $\hat{\rho}_c(S) = \text{corr}(X, Z \mid (X, Z) \in S)$.

A general bivariate density f for the variables (X, Z) would be fitted locally in a neighborhood of each point $y = (x, z)$, by a bivariate Gaussian density,

$$\phi(u, \theta(y)) = \frac{1}{2\pi\sigma_1(y)\sigma_2(y)} \times \exp \left\{ -\frac{1}{2(1-\rho^2(y))} \left[\left(\frac{u_1 - \mu_1(y)}{\sigma_1(y)} \right)^2 + \left(\frac{u_2 - \mu_2(y)}{\sigma_2(y)} \right)^2 - 2\rho(y) \left(\frac{u_1 - \mu_1(y)}{\sigma_1(y)} \right) \left(\frac{u_2 - \mu_2(y)}{\sigma_2(y)} \right) \right] \right\} \quad (5)$$

where $u = (u_1, u_2)^\top$ is the running variable in the Gaussian distribution and $\theta(y) = \phi(\mu_1(y), \mu_2(y), \sigma_1(y), \sigma_2(y), \rho(y))$, with $\mu_i(y)$, $i = 1, 2$ the local means, $\sigma_i(y)$, $i = 1, 2$ the local standard deviation and $\rho(y)$, the local correlation at the point $y = (x, z)$. The population values of the local parameters $\theta_b(y) = \theta(y)$ are obtained by minimizing the local penalty function,²⁸

$$q = \int K_b(u - y) [\phi(u, \theta(y)) - \log \phi(u, \theta(y)) f(u)] du \quad (6)$$

where $K_b = (b_1 b_2)^{-1} K(b_1^{-1}(u_1 - y_1)) K(b_2^{-1}(u_2 - y_2))$ is a product kernel with bandwidth $b = (b_1, b_2)$, and the local Gaussian correlation $\rho_b(y) = \rho(y)$ is defined as the last element of the vector $\theta(y)$ that minimizes q . Moving to another point $y' = (x', z')$ of f another Gaussian $\phi(u, \theta(y'))$ is required to approximate f in a neighbourhood S' of y' . In this way f may be represented by a family of Gaussian bivariate densities as y varies and in each specific neighborhood of y , the local dependence properties are described by $\rho(y)$. The (local) dependence may be defined to be positive (negative) if $\rho(y) > 0$ ($\rho(y) < 0$). The bias of conditional correlation is accommodated since the same Gaussian f fits every point.

Given the observations $Y_i = (X_i, Z_i)$, $i = 1, \dots, n$ from f the corresponding estimates $\hat{\theta}(y)$ are obtained by maximizing the local log-likelihood function (see Hjort and Jones, 1996),

$$L(Y_1, \dots, Y_n, \theta_b(y)) = n^{-1} \sum_i K_b(Y_i - y) \log \phi(Y_i, \theta_b(y)) - \int K_b(u - y) \phi(u, \theta_b(y)) du \quad (7)$$

²⁸This type of penalty function q was used in Hjort and Jones (1996) for density estimation purposes and later by Tjøstheim and Hufthammer (2013) in the development of local Gaussian correlation. The former argue that q can be interpreted as a locally weighted Kullback-Leibler criterion for measuring the distance between $f(\cdot)$ and the chosen parametric distribution (in our case $\phi(\cdot, \theta(y))$).

For the local likelihood function (4) to be consistent with the penalty function q , the $\theta(y)$ is chosen to minimize q , such that it satisfies the following 5-dimensional set of equations:²⁹

$$\frac{\partial q}{\partial \theta_j} = \int K_b(u - y) \frac{\partial}{\partial \theta_j} \{\log(\phi(u, \theta(y)))\} [\phi(u, \theta(y)) - f(u)] du, \quad j = 1, \dots, 5 \quad (8)$$

Using the notation,

$$\gamma_j(\cdot, \theta) = \frac{\partial}{\partial \theta_j} \{\log \phi(\cdot, \theta)\}, \quad (9)$$

and assuming that $E\{K_b(Y_i - y) u_j(Y_i, \theta(y))\} < \infty$, the law of large numbers gives,

$$\begin{aligned} \frac{\partial L}{\partial \theta_j} &= n^{-1} \sum_i K_b(Y_i - y) \gamma_j(Y_i, \theta(y)) - \int K_b(u - y) \gamma_j(u, \theta(y)) \phi(u, \theta(y)) du \\ &\rightarrow \int K_b(u - y) \gamma_j(u, \theta(y)) [f(u) - \phi(u, \theta(y))] du = -\frac{\partial q}{\partial \theta_j}, \quad j = 1, \dots, 5 \end{aligned} \quad (10)$$

as $n \rightarrow \infty$, we see that (7) can be identified with (5). Also note that as $b \rightarrow \infty$ (4) reduces to the ordinary log-likelihood for a Gaussian distribution ϕ plus a constant, and hence $\rho(y)$ reduces to the ordinary global Gaussian correlation. For more details about the local $b_i \rightarrow 0, i = 1, 2$ and estimation of standard errors, we refer to [Tjøstheim and Hufthammer \(2013\)](#). The numerical maximization of the local likelihood (4) leads to local likelihood estimates $\theta_{n,b}(y)$, including estimates $\rho_{n,b}(y)$ of the local correlation. It is shown in [Tjøstheim and Hufthammer \(2013\)](#) that under relatively weak regularity conditions $\theta_{n,b}(y) \rightarrow \theta_b(y)$ for b fixed, and $\theta_{n,b}(y) \rightarrow \theta(y)$ almost surely for $b = b_n$ tending to zero.³⁰

[Berentsen and Tjøstheim \(2014, Section 3.4\)](#) argue that the bandwidth choice depends on the nature of the question. To investigate the local dependence structure in the dataset can be quite informative to compute several bandwidths to obtain information about the dependence structure on different scales of locality. In some cases it would be beneficial to have a data-driven bandwidth choice similar to a bandwidth choice for density kernel estimation. In our empirical analysis, we employ two methods for bandwidth selection, the normal-reference rule-of-thumb as in [Støve et al. \(2014\)](#) and the methodology of likelihood cross-validation proposed by [Hall et al. \(2004\)](#).³¹ Since both approaches provide qualitatively similar results, we present the bandwidth choice based on the normal-reference rule-of-thumb.³²

²⁹More details concerning the local Gaussian theory can be found in [Berentsen and Tjøstheim \(2014\)](#) and [Tjøstheim and Hufthammer \(2013\)](#).

³⁰The R-package 'localgauss' has been used for estimating $\rho_{n,b}(y)$. An introduction to the R package 'localgauss' for estimation and visualization of local dependence is available in [Berentsen et al. \(2014\)](#).

³¹The R-package 'MASS' and 'np' have been used for rule-of-thumb and the data-driven bandwidth selection methods respectively.

³²For further discussion regarding bandwidth selection, see [Tjøstheim and Hufthammer \(2013\)](#).

A.3 The CUSUM test for stability of the parameters

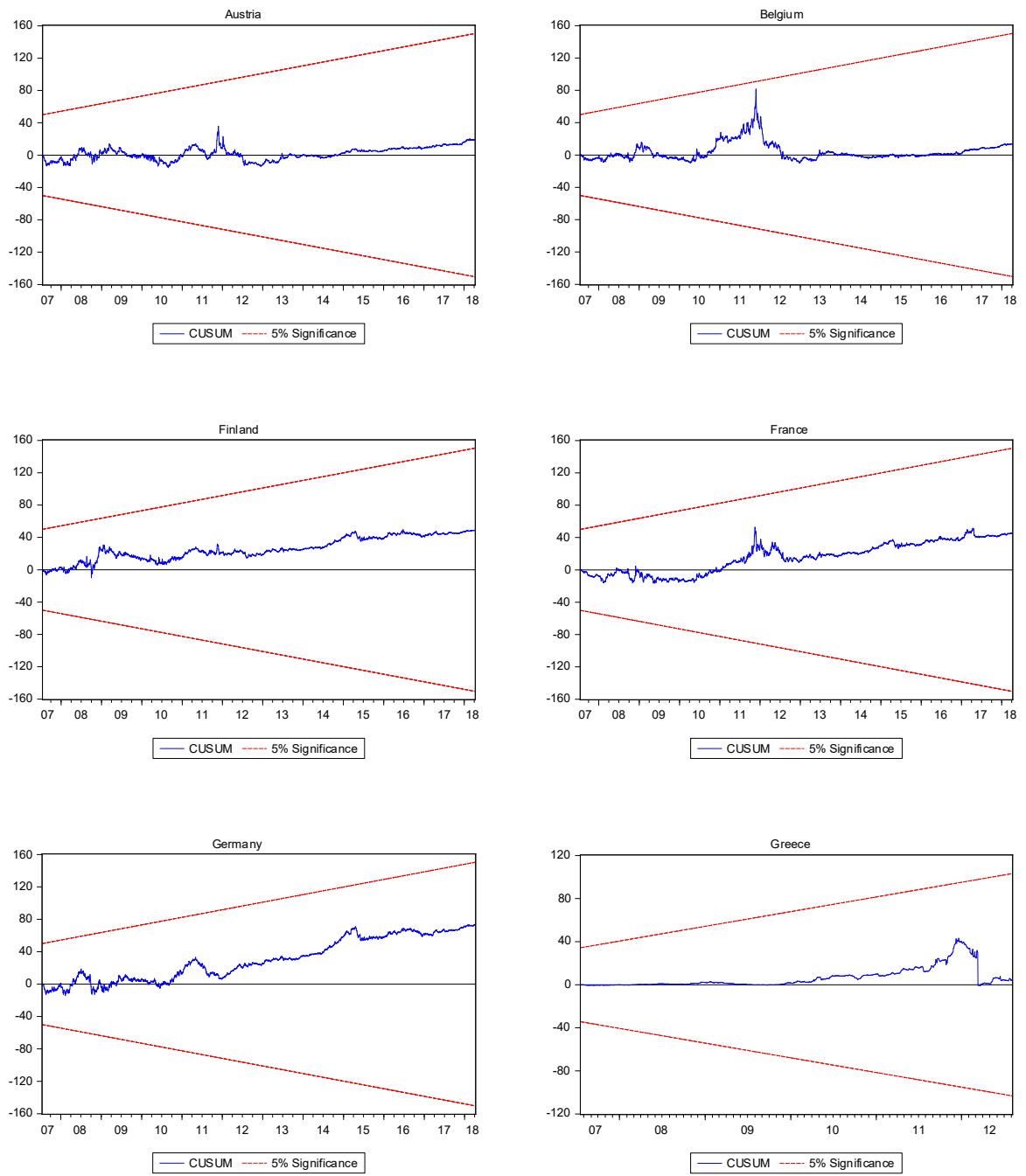


Figure 5: Cumulative Sum test for the bond market factor model together with the 5% critical lines. The test finds parameter instability if the cumulative sum goes outside the area between the two critical lines.

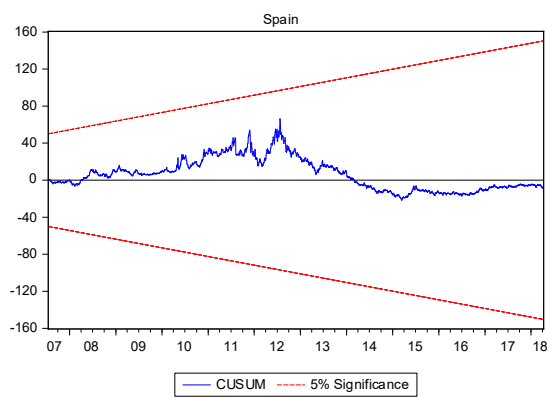
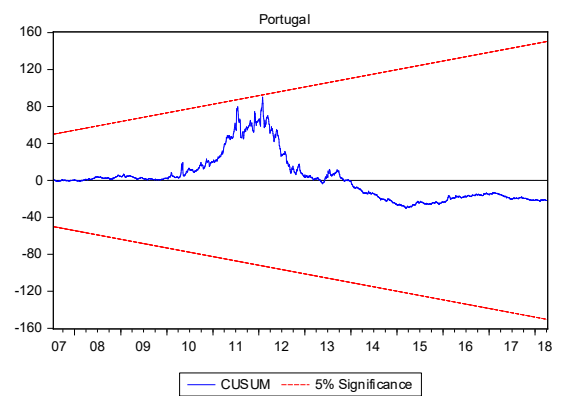
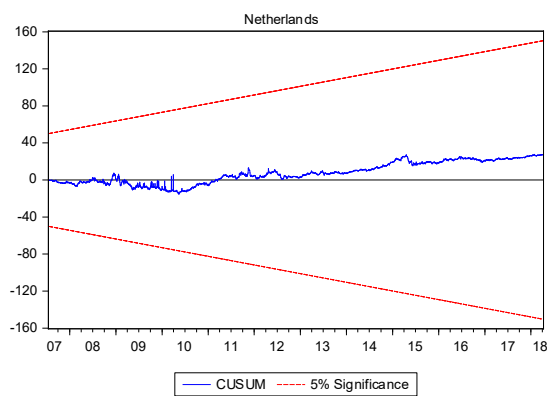
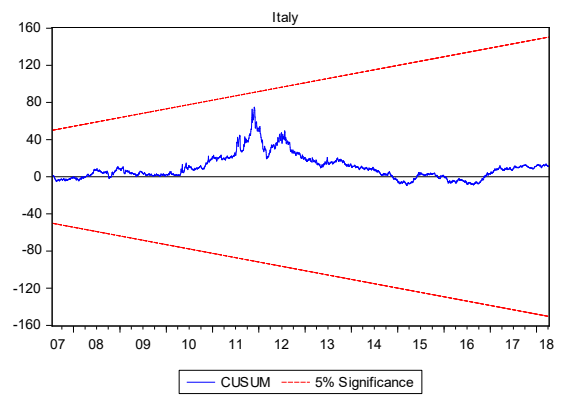
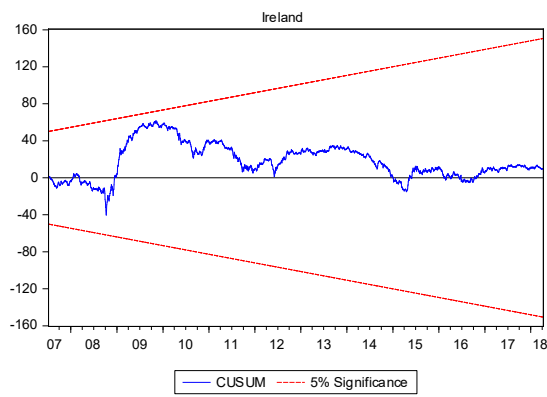


Figure 5 (cont.): Cumulative Sum test for the bond market factor model together with the 5% critical lines. The test finds parameter instability if the cumulative sum goes outside the area between the two critical lines.

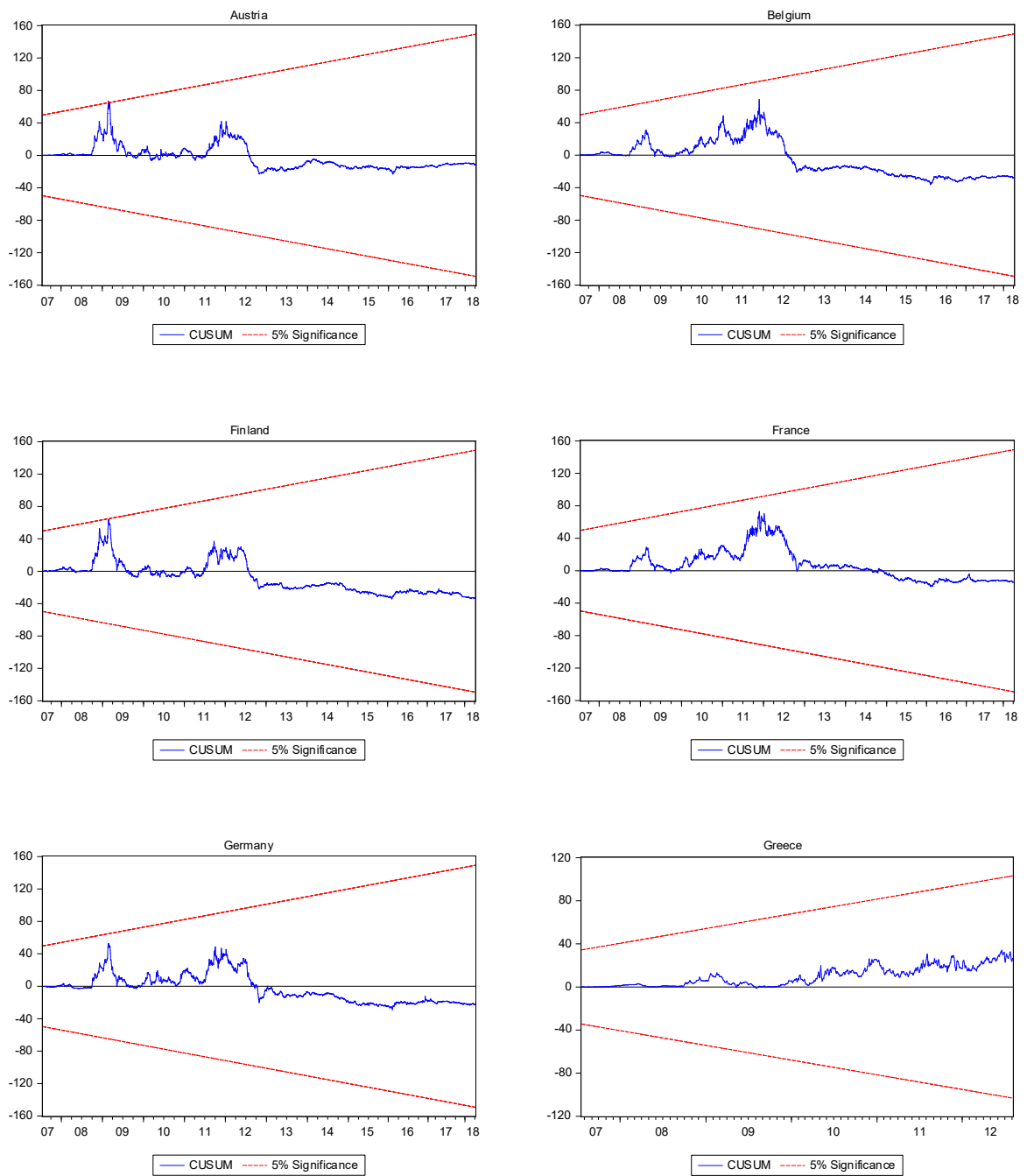


Figure 6: Cumulative Sum test for the CDS market factor model together with the 5% critical lines. The test finds parameter instability if the cumulative sum goes outside the area between the two critical lines.

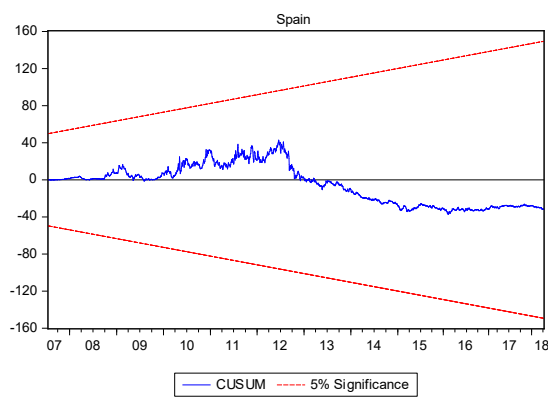
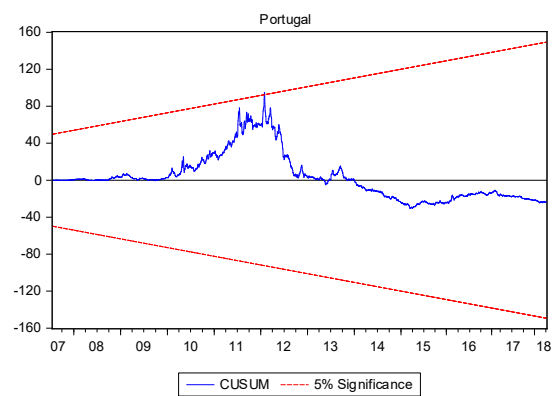
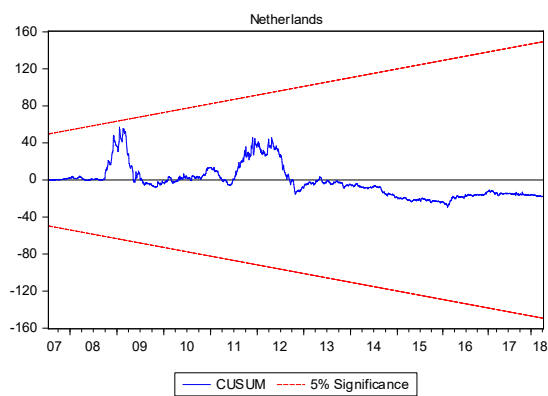
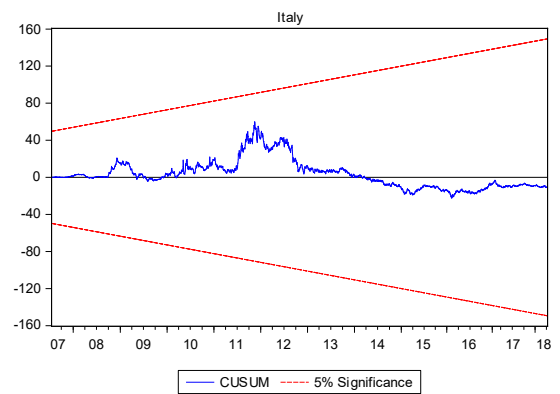
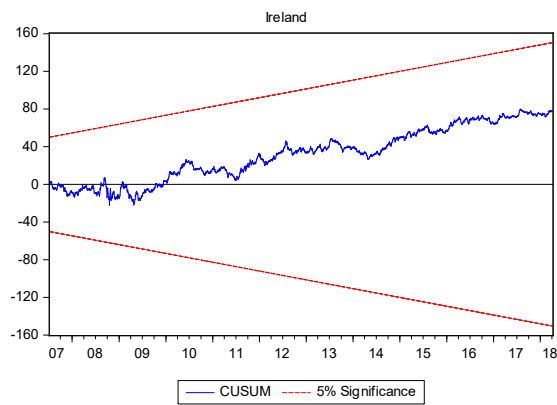


Figure 6 (cont.): Cumulative Sum test for the CDS market factor model together with the 5% critical lines. The test finds parameter instability if the cumulative sum goes outside the area between the two critical lines.

A.4 State space models and the Kalman filter

In this section, we present a brief discussion of the specification and estimation of a linear state space model and kalman filtering method. Our estimation method is in conceptual proximity to [Koopman *et al.* \(1999\)](#).³³

A linear state space representation of the dynamics of the $n \times 1$ vector y_t is given by the system of equations:

$$y_t = a_t + Z_t s_t + \epsilon_t \quad (11)$$

$$s_{t+1} = l_t + T_t s_t + u_t \quad (12)$$

where s_t is an $m \times 1$ vector of possibly unobserved state variables, a_t , Z_t , l_t and T_t are conformable vectors and matrices, and ϵ_t , u_t are vectors of mean zero, Gaussian disturbances. We will refer to the first set of equations as the “signal” or “observation” equations and the second set as the “state” or “transition” equations. The disturbance vectors ϵ_t and u_t are assumed to be serially independent, with contemporaneous variance structure:

$$\Omega_t = \text{var} \begin{bmatrix} \epsilon_t \\ u_t \end{bmatrix} = \begin{bmatrix} H_t & G_t \\ G_t' & Q_t \end{bmatrix} \quad (13)$$

where H_t is an $n \times n$ symmetric variance matrix, Q_t is an $m \times m$ symmetric variance matrix, and G_t is an $n \times m$ matrix of covariances. Generalizing the specification given in Eq. (11)—(13), the system matrices and vectors $\Xi_t \equiv \{a_t, l_t, Z_t, T_t, H_t, Q_t, G_t\}$ are allowed to depend upon observable explanatory variables X_t and unobservable parameters ϕ .

Considering the conditional distribution of the state vector s_t given information available at time k , the mean and variance matrix of the conditional distribution can be defined as:

$$s_{t|k} \equiv E_k(s_t) \quad (14)$$

$$P_{t|k} \equiv E_k \left[(s_t - s_{t|k})(s_t - s_{t|k})' \right] \quad (15)$$

where the subscript below the expectation operator indicates that expectations are taken using the conditional distribution for that period.

By setting $k = t - 1$, one important conditional distribution is obtained, so that we obtain the one-step ahead mean $s_{t|t-1}$ and one-step ahead variance, $P_{t|t-1}$, of the states s_t . Under the Gaussian error assumption, $s_{t|t-1}$ is also the minimum mean square error estimator of s_t and $P_{t|t-1}$ is the mean square error (MSE) of s_{t-1} . If the normality assumption is violated, $s_{t|t-1}$ is still the minimum mean square linear estimator of s_t .

Given the one-step ahead state conditional mean, we can also form the (linear) minimum MSE one-step ahead estimate of y_t :

$$y_t = y_{t|t-1} \equiv E_{t-1}(y_t) = E_{t-1}(y_t | s_{t|t-1}) = a_t + Z_t s_{t|t-1} \quad (16)$$

The one-step ahead prediction error is given by,

$$\tilde{\epsilon}_t = \epsilon_{t|t-1} \equiv y_t - y_{t|t-1} \quad (17)$$

and the prediction error variance is defined as:

$$V_t = V_{t|t-1} \equiv \text{var}(\epsilon_{t|t-1}) = Z_t P_{t|t-1} Z_t' + H_t \quad (18)$$

³³A comprehensive analysis can also be found in [Hamilton \(1994a\)](#), [Hamilton \(1994b\)](#) and [Harvey \(1989\)](#).

The Kalman filter is a recursive algorithm for sequentially updating the one-step ahead estimate of the state mean and variance given new information. Details on the recursion are provided in Koopman *et al.* (1999), Hamilton (1994a) and Harvey (1989). To implement the Kalman filter, we replace any unknown elements of the system matrices by their estimates. Under the assumption that the ϵ_t and u_t are Gaussian, the sample log likelihood:

$$\log L(\phi) = -\frac{nT}{2} \log 2\pi - \frac{1}{2} \sum_t \log |V_t(\phi)| - \frac{1}{2} \sum_t \tilde{\epsilon}_t'(\phi) V_t(\phi)^{-1} \tilde{\epsilon}_t(\phi) \quad (19)$$

may be evaluated using the Kalman filter. Using numeric derivatives, standard iterative techniques may be employed to maximize the likelihood with respect to the unknown parameters ϕ .

Moreover, it is sufficient to note that given initial values for the state mean and covariance, values for the system matrices Ξ_t , and observations on y_t , the Kalman filter may be used to compute one-step ahead estimates of the state and the associated mean square error matrix, $\{s_{t|t-1}, P_{t|t-1}\}$, the contemporaneous or filtered state mean and variance, $\{s_t, P_t\}$, and the one-step ahead prediction, prediction error, and prediction error variance, $\{y_{t|t-1}, \epsilon_{t|t-1}, V_{t|t-1}\}$. Consider also that we may obtain the standardised prediction residual, $\epsilon_{t|t-1}$, by dividing $\epsilon_{t|t-1}$ by the square-root of the corresponding diagonal element of $V_{t|t-1}$.

A.5 Kalman filter model estimations

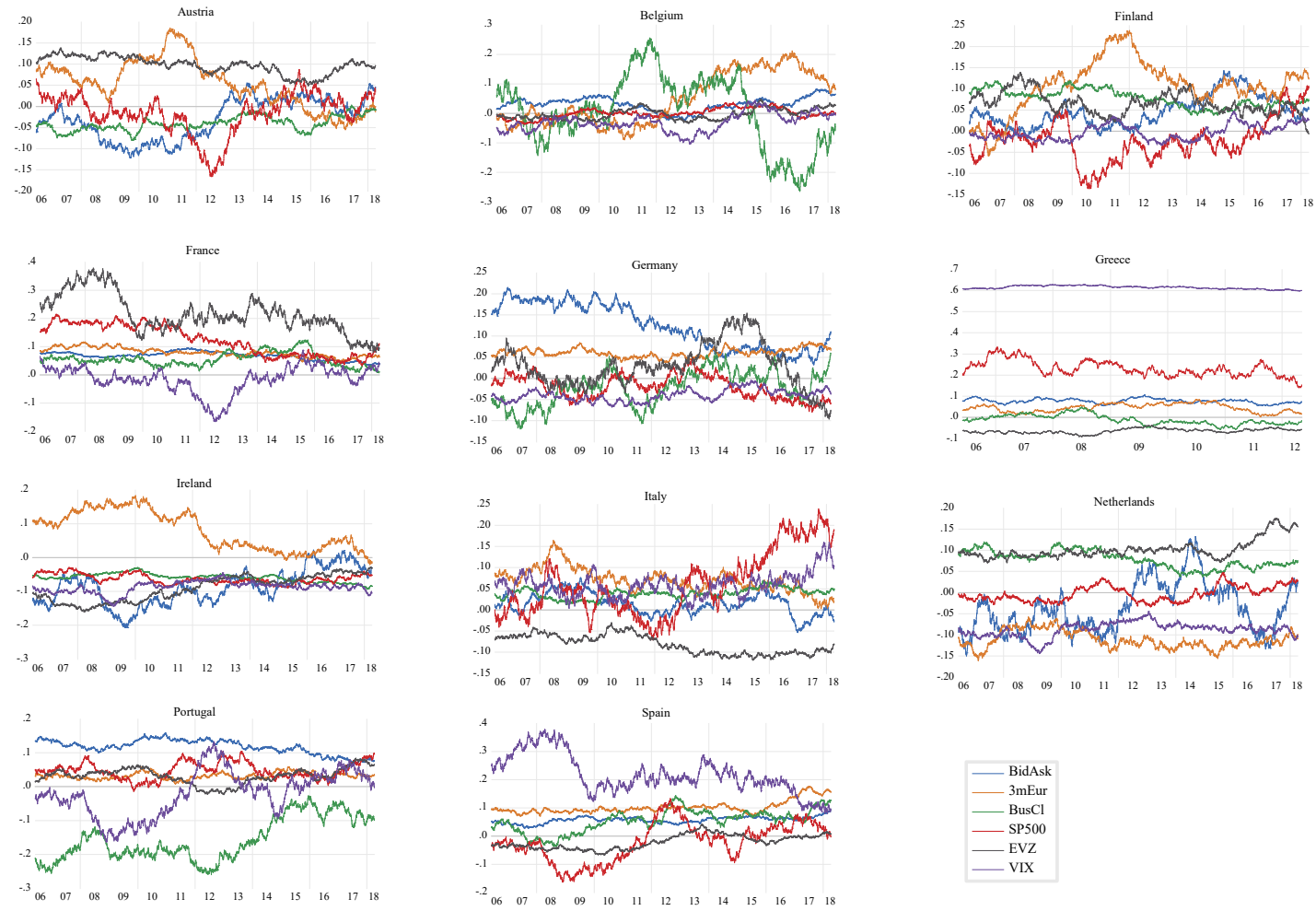


Figure 7: Kalman filter time-varying estimates of factor loadings $(\beta_{1,t}, \beta_{2,t}, \beta_{3,t}, \beta_{4,t}, \beta_{5,t}, \beta_{6,t})$ from the bond factor model (Eq. 1).

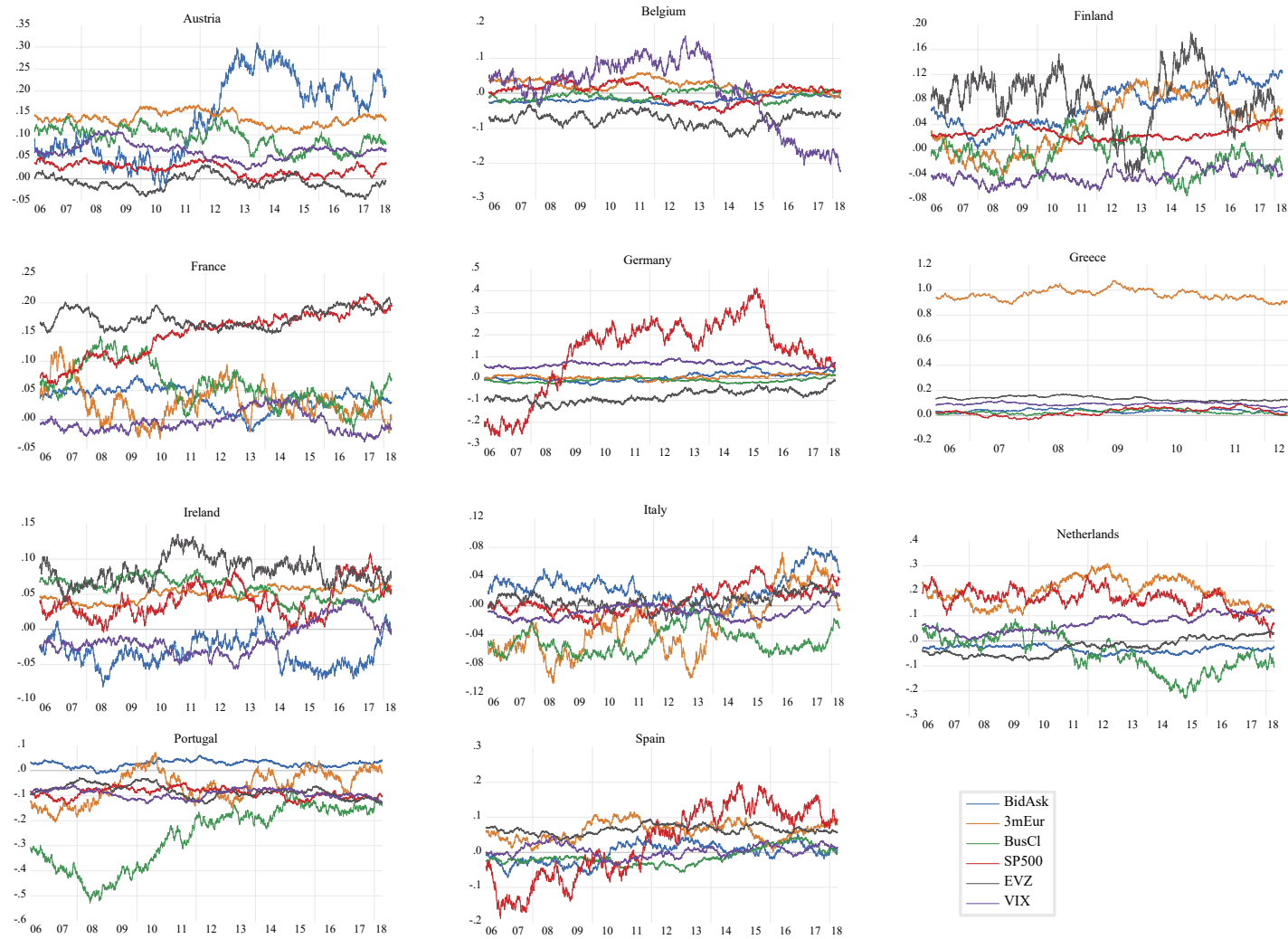


Figure 8: Kalman filter time-varying estimates of the factor loadings ($\beta_{1,t}, \beta_{2,t}, \beta_{3,t}, \beta_{4,t}, \beta_{5,t}, \beta_{6,t}$) from the CDS factor model (Eq. 1).

A.6 Diagnostic checks in standardised residuals.

Table 8: $Q^2(p)$ denotes the Ljung–Box statistics for tests of lack of correlation of squared standardised residuals derived from each Kalman filter model for different lags ($k= 1, 5, 10, 15$ and 20 lags). ARCH LM test tests the null hypothesis of no ARCH effects in standardised residuals. We report the ARCH test p -value for different lag orders ($k= 1, 5, 10, 15$ and 20 lags). These are based on the Lagrange multiplier test for conditional heteroscedasticity of Engle (1982).

	Bond markets					CDS markets				
	$Q^2(1)$	$Q^2(5)$	$Q^2(10)$	$Q^2(15)$	$Q^2(20)$	$Q^2(1)$	$Q^2(5)$	$Q^2(10)$	$Q^2(15)$	$Q^2(20)$
	[p -value]	[p -value]	[p -value]	[p -value]	[p -value]	[p -value]	[p -value]	[p -value]	[p -value]	[p -value]
AT	8.519 [0.133]	13.855 [0.311]	16.477 [0.087]	18.557 [0.234]	31.062 [0.060]	3.798 [0.851]	4.515 [0.478]	5.973 [0.817]	8.062 [0.921]	16.636 [0.676]
BE	3.987 [0.931]	7.403 [0.192]	10.625 [0.388]	15.566 [0.411]	19.023 [0.520]	5.529 [0.211]	8.182 [0.146]	20.403 [0.126]	27.387 [0.325]	29.923 [0.071]
DE	6.088 [0.465]	0.989 [0.320]	1.252 [0.535]	1.294 [0.730]	1.587 [0.811]	4.744 [0.340]	5.283 [0.382]	8.085 [0.621]	12.531 [0.638]	13.644 [0.848]
ES	6.123 [0.694]	0.998 [0.963]	3.864 [0.953]	7.634 [0.938]	18.070 [0.583]	0.122 [0.655]	0.712 [0.985]	1.172 [1.000]	3.038 [1.000]	3.599 [1.000]
FI	12.394 [0.150]	30.752 [0.078]	30.861 [0.099]	30.931 [0.124]	30.959 [0.155]	2.778 [0.661]	3.799 [0.579]	6.866 [0.738]	7.892 [0.928]	10.228 [0.964]
FR	5.744 [0.190]	7.953 [0.159]	11.395 [0.328]	14.986 [0.452]	18.526 [0.553]	2.190 [0.753]	1.189 [0.946]	3.002 [0.981]	3.691 [0.999]	5.589 [0.999]
GR	9.842 [0.103]	10.437 [0.064]	17.358 [0.070]	20.369 [0.158]	21.279 [0.322]	1.314 [0.369]	3.984 [0.552]	6.444 [0.777]	12.522 [0.639]	13.763 [0.842]
IE	6.906 [0.693]	7.183 [0.207]	8.638 [0.567]	21.041 [0.136]	22.510 [0.314]	8.436 [0.071]	11.082 [0.05]	14.196 [0.164]	15.023 [0.450]	16.998 [0.653]
IT	2.147 [0.117]	1.889 [0.864]	2.803 [0.986]	6.515 [0.970]	30.275 [0.07]	2.382 [0.775]	4.606 [0.466]	7.765 [0.652]	9.440 [0.853]	10.950 [0.948]
NL	3.810 [0.512]	1.654 [0.895]	2.327 [0.993]	7.356 [0.947]	7.660 [0.994]	9.891 [0.147]	12.917 [0.624]	15.148 [0.127]	26.006 [0.053]	28.020 [0.109]
PT	1.320 [0.281]	1.893 [0.864]	2.682 [0.988]	3.594 [0.999]	5.401 [0.999]	3.120 [0.596]	3.594 [0.609]	6.310 [0.789]	7.878 [0.929]	10.824 [0.951]
	ARCH (1)	ARCH (5)	ARCH (10)	ARCH (15)	ARCH (20)	ARCH (1)	ARCH (5)	ARCH (10)	ARCH (15)	ARCH (20)
AT	0.5040	0.1035	0.0787	0.2115	0.0617	0.7019	0.5079	0.8607	0.9494	0.5314
BE	0.8665	0.1838	0.3755	0.4085	0.5324	0.8708	0.1467	0.1612	0.1399	0.6810
DE	0.6217	0.3204	0.5411	0.7302	0.8133	0.9082	0.3891	0.6625	0.6594	0.8762
ES	0.9628	0.9628	0.9529	0.9246	0.6076	0.9251	0.9851	0.9996	0.9995	0.9991
FI	0.8186	0.0845	0.1067	0.1337	0.1565	0.9901	0.5937	0.7445	0.9327	0.9467
FR	0.8438	0.1463	0.3143	0.4399	0.5621	0.5547	0.9464	0.9806	0.9984	0.1658
GR	0.7910	0.0648	0.0893	0.2375	0.4178	0.6547	0.5484	0.7889	0.5959	0.9467
IE	0.2148	0.1803	0.5132	0.1208	0.3087	0.8741	0.0534	0.2120	0.5096	0.7787
IT	0.2776	0.8632	0.9861	0.9677	0.0679	0.8246	0.4724	0.7074	0.8755	0.2922
NL	0.4217	0.8981	0.9931	0.9504	0.9950	0.9940	0.0214	0.1148	0.0334	0.7751
PT	0.3641	0.8684	0.9887	0.9988	0.9991	0.5111	0.6051	0.7667	0.9183	0.8457

A.7 Bootstrap test for serial independence within the Kalman filter model residual series

Table 9: This table shows p -values from the bootstrap test for serial independence of [Lacal and Tjøstheim \(2017\)](#). The small p -value indicates that we reject the null hypothesis of independence between X_t and X_{t-k} ($k= 1, 5, 10, 15$ and 20 lags). The bootstrap test for serial independence is based on 1000 replications.

<i>Bond Markets</i>					
	$k = 1$	$k = 5$	$k = 10$	$k = 15$	$k = 20$
AT	0.505	0.991	0.932	0.687	0.403
BE	0.670	0.754	0.970	0.945	0.288
DE	0.469	0.433	0.375	0.384	0.257
ES	0.230	0.954	0.147	0.958	0.632
FI	0.465	0.774	0.931	0.667	0.336
FR	0.341	0.159	0.889	0.961	0.692
GR	0.739	0.211	0.733	0.959	0.545
IE	0.064	0.884	0.815	0.794	0.287
IT	0.527	0.789	0.705	0.260	0.909
NL	0.520	0.246	0.895	0.927	0.370
PT	0.902	0.830	0.987	0.605	0.820
<i>CDS Markets</i>					
AT	0.779	0.429	0.697	0.384	0.472
BE	0.811	0.824	0.806	0.777	0.297
DE	0.545	0.914	0.907	0.395	0.775
ES	0.629	0.863	0.894	0.411	0.541
FI	0.713	0.734	0.930	0.688	0.133
FR	0.242	0.847	0.741	0.671	0.761
GR	0.648	0.474	0.694	0.512	0.230
IE	0.552	0.899	0.855	0.857	0.476
IT	0.475	0.465	0.447	0.332	0.443
NL	0.964	0.312	0.849	0.287	0.328
PT	0.671	0.698	0.923	0.864	0.865

A.8 Bootstrap test for contagion using the original series

Table 10: This table shows p -values from the bootstrap test for contagion according to the time dates of correlation change-points estimated with the algorithm of [Demetrescu and Wied \(2019\)](#) based on the original series. The null hypothesis indicates no contagion between the GIIPS (GR, IE, IT, PT, ES) bonds and the European bond (Panel A) and CDS (Panel B) markets. The bootstrap test for contagion is based on 5000 replications. The sample period starts on January 2, 2006 and ends on March 8, 2012 for Greece and on April 5, 2018 for the rest EMU South countries.

	GR (1)	Date	IE (2)	Date	IT (3)	Date	PT (4)	Date	ES (5)	Date
<i>Panel A : Bond Markets</i>										
AT	0.000	13/10/2008	0.020	30/1/2009	0.029	20/7/2011	0.031	9/10/2009	0.000	8/12/2009
	0.037	30/11/2009	0.008	25/3/2010			0.012	22/3/2010	0.230	11/6/2012
BE	0.003	6/8/2009	0.007	3/3/2010			0.451	12/4/2011		
DE	0.010	16/10/2008	0.011	25/2/2009	0.000	21/10/2008	0.013	26/9/2008	0.042	30/12/2008
	0.009	25/11/2009	0.009	2/12/2010	0.046	27/4/2010	0.019	26/11/2009	0.011	3/8/2011
					0.965	7/7/2011				
ES	0.002	23/12/2009	0.045	1/12/2010	0.000	21/7/2011	0.013	24/3/2011		
					0.587	1/7/2015	0.591	2/7/2015		
FI	0.000	26/10/2009	0.931	2/12/2010	0.016	2/7/2008	0.011	22/4/2011	0.000	3/12/2009
			0.721	7/3/2014	0.019	19/5/2010			0.063	20/7/2011
FR	0.004	16/11/2009	0.010	1/4/2010	0.519	20/7/2011	0.029	29/1/2010	0.032	14/7/2009
			0.548	1/12/2010			0.009	21/4/2011	0.512	18/7/2011
GR			0.021	9/12/2009	0.048	15/10/2008	0.030	17/2/2009	0.044	23/12/2009
					0.000	23/11/2009				
IE	0.034	9/12/2009					0.980	30/11/2010	0.005	1/12/2010
IT	0.000	15/10/2008					0.033	10/2/2010	0.012	20/7/2011
	0.012	23/11/2009					0.045	20/5/2011	0.425	29/6/2015
NL	0.007	14/10/2009	0.007	23/1/2009	0.008	24/6/2011	0.040	30/9/2009	0.009	21/6/2011
PT	0.048	17/12/2009	0.012	30/11/2010	0.003	10/2/2010			0.027	24/3/2011
					0.042	20/5/2011			0.582	2/7/2015
<i>Panel B : CDS Markets</i>										
AT	0.045	2/11/2009	0.045	15/11/2010	0.011	18/7/2011	0.000	28/3/2011	0.000	2/8/2011
BE	0.010	5/11/2009	0.005	26/11/2010	0.042	21/7/2011	0.000	1/4/2011	0.007	3/8/2011
	0.011	21/4/2010								
DE	0.021	6/11/2009	0.045	23/11/2010	0.027	21/7/2011	0.004	18/4/2011	0.035	20/7/2011
	0.452	20/04/2010								
ES	0.003	9/11/2009	0.030	3/12/2010	0.010	20/7/2011	0.027	11/4/2011		
	0.008	22/4/2010	0.398	22/7/2011						
FI	0.005	12/11/2009			0.009	2/8/2011	0.041	5/5/2011	0.012	3/8/2011
FR	0.002	5/11/2009	0.001	12/11/2010	0.019	20/7/2011	0.017	18/4/2011	0.000	18/7/2011
	0.712	22/04/2010								
GR			0.422	3/1/2008	0.490	4/11/2009	0.873	18/11/2009	0.730	19/11/2009
			0.091	14/11/2009	0.731	21/4/2010	0.554	22/4/2010	0.497	21/4/2010
IE	0.249	3/1/2008			0.457	3/12/2010	0.354	29/11/2010	0.831	2/12/2010
	0.013	12/11/2009			0.000	15/7/2011	0.045	28/3/2011	0.010	25/7/2011
IT	0.000	6/11/2009	0.012	3/12/2010			0.014	12/4/2011	0.002	20/7/2011
	0.000	20/4/2010	0.539	15/7/2011						
NL	0.072	5/11/2009	0.000	10/12/2010	0.019	3/8/2011	0.037	5/5/2011	0.009	19/7/2011
PT	0.000	17/11/2009	0.023	30/11/2010	0.018	12/4/2011	0.029	5/5/2011	0.034	11/4/2011
	0.005	21/4/2010	0.348	30/3/2011						

Table 11: This table shows p -values from the bootstrap test for contagion according to the time dates of correlation change-points estimated with the algorithm of [Demetrescu and Wied \(2019\)](#) based on the original series. The null hypothesis indicates no contagion between the GIIPS (GR, IE, IT, PT, ES) CDS and the European bond (Panel A) and CDS (Panel B) markets. The bootstrap test for contagion is based on 5000 replications. The sample period starts on January 2, 2006 and ends on March 8, 2012 for Greece and on April 5, 2018 for the rest EMU South countries.

	GR (1)	Date	IE (2)	Date	IT (3)	Date	PT (4)	Date	ES (5)	Date
<i>Panel A : Bond Markets</i>										
AT	0.233	2/11/2009	0.148	13/10/2011	0.912	31/10/2011			0.089	25/10/2011
	0.411	4/11/2011							0.152	21/11/2011
BE	0.122	3/11/2011	0.187	20/10/2011			0.137	30/11/2010		
DE	0.872	22/4/2010	0.341	12/10/2011	0.012	15/7/2010	0.005	22/7/2010	0.000	29/7/2010
					0.307	25/10/2011	0.204	15/5/2013	0.742	26/2/2010
ES	0.471	10/10/2011	0.561	5/10/2011	0.745	9/11/2011	0.885	6/4/2010		
			0.996	7/8/2015	0.991	22/6/2015				
FI	0.311	13/11/2009	0.962	9/1/2008	0.015	1/9/2010	0.022	2/9/2010	0.004	8/9/2010
	0.411	15/11/2011	0.981	23/4/2010	0.411	21/10/2013	0.962	15/5/2013	0.639	15/5/2013
FR	0.452	21/4/2010	0.948	17/7/2008	0.007	22/7/2010	0.032	23/7/2010	0.030	5/8/2010
			0.127	10/10/2011	0.412	18/11/2011	0.041	4/8/2010	0.518	8/11/2011
GR			0.712	9/11/2011	0.301	10/11/2011	0.541	9/11/2011	0.738	8/11/2011
IE	0.497	9/11/2011					0.430	4/11/2011	0.351	15/11/2011
					0.997	17/11/2011			0.619	7/8/2015
IT	0.230	10/11/2011	0.912	12/10/2010			0.631	14/11/2011	0.350	9/11/2011
			0.259	10/11/2011					0.341	15/6/2015
NL	0.981	5/11/2009	0.965	1/7/2008	0.025	15/7/2010	0.001	28/7/2010	0.047	1/7/2010
	0.757	7/11/2011	0.438	3/8/2010	0.659	26/3/2013			0.589	26/3/2013
PT	0.329	9/11/2011	0.341	18/11/2011	0.338	14/11/2011			0.589	6/4/2010
<i>Panel B : CDS Markets</i>										
AT			0.812	26/1/2012	0.001	23/11/2011	0.722	15/9/2008	0.458	15/9/2008
							0.716	2/2/2012	0.021	3/1/2012
BE			0.291	25/1/2012	0.005	10/1/2012	0.370	12/1/2012	0.033	9/1/2012
DE			0.489	10/1/2012	0.000	18/1/2012	0.341	18/9/2008	0.019	19/1/2012
							0.992	13/1/2012		
ES	0.003	16/11/2011			0.000	23/11/2011	0.579	1/11/2011		
FI	0.411	25/9/2008	0.899	19/1/2012	0.013	29/11/2011	0.511	3/1/2012	0.010	14/11/2011
FR					0.005	10/1/2012	0.420	18/11/2009	0.015	4/1/2012
					0.966	1/2/2013	0.847	13/1/2012	0.968	1/2/2013
GR			0.741	10/11/2011					0.020	18/11/2011
IE	0.015	9/11/2011			0.003	16/11/2011	0.482	5/1/2012		
					0.958	31/8/2015				
IT	0.983	14/9/2010	0.712	16/11/2011			0.581	15/11/2012	0.000	22/11/2011
			0.349	31/8/2015			0.918	5/9/2012		
NL	0.401	19/9/2008	0.311	12/12/2012	0.009	18/1/2012	0.558	10/9/2008	0.421	16/9/2008
	0.391	23/9/2010					0.820	12/1/2012	0.020	16/1/2012
PT			0.571	12/1/2012	0.004	14/11/2012			0.008	8/11/2011