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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 differently across EMU (Economic and Monetary Union) member states (Claeys and Vašìcek, 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 analyse 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. If the answer to the latter is yes, we also ask 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] This study adopts the approach of Bekaert *et al.* (2005), where contagion is defined as 'correlation between markets in excess of that implied by economic fundamentals'. We complement this with a recent measure of local correlation, introduced 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 financial contagion during the Eurozone

[1] An additional classification of Eurozone 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.

[2] See King and Wadhwani (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 more recently by Forbes (2012).

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

The dataset includes 5-year sovereign bond yields and CDS spreads covering the period from January 2006 to April 2018. In the first stage, our analysis employs a factor model, where the bond and CDS data are conditioned on state variables (see Bekaert *et al*., 2005; De Bruyckere *et al*., 2013; Fontana and Scheicher, 2016). These variables reflect the underlying fundamentals that drive sovereign bonds and CDS contracts. By controlling for them, we avoid any issues associated with the bias correction for correlations and further exclude any alternative explanations of our findings, such as those stemming from changes in the underlying fundamentals (see Forbes and Rigobon, 2002).

In the second stage, we employ the method of Tjøstheim and Hufthammer (2013) to study the local Gaussian correlation dynamics of the filtered bond and CDS series. The approach of Galeano and Weid (2014) is employed 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 (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. The analysis is then replicated for our initial bond and CDS data without employing the first-stage factor model. We compare the estimates from the analysis of the fundamentals-filtered series with those from the analysis of the unfiltered series to distinguish between cases of "pure contagion" and cases of contagion due to changes in economic fundamentals ("wake-up call contagion") or short-lived contagion episodes ("limited contagion").

Typically, during crises, the larger the shock, the stronger the correlations among financial asset prices (Veldkamp, 2006; Aloui *et al*., 2011). As Table 1 suggests, the Eurozone crisis poses no exception: a number of negative shocks in different countries of the EMU South caused a sharp rise in the bond yields and CDS spreads of the source country (in bold). This rise is greater than the equivalent rise in the bond yields and CDS spreads of the remaining countries. Most importantly, this was accompanied by a rise in the cross-market correlations between changes in the source country's bond yields and CDS spreads after the event. The degree of comovement in this period ranges between 0.50 and 0.75, which indicates that sovereign bond and CDS markets do not respond uniformly to the same shocks. This is in line with conventional knowledge that correlation between financial assets becomes stronger as the market is going down and approaches one when the market crashes.

[Insert Table 1 about here.]

However, this one-number description is simplistic. Linear measures of dependence work well for approximately bivariate Gaussian variables, but tend to have low power when the dependence structure is nonlinear. Traditional tests that are based on rank correlation statistics, namely Spearman's rho and Kendall's tau, are more robust against non-Gaussian distributions, but do not generally capture non-monotone dependence. All these measures are measures of global dependence. During economic downturns there may be subsets of values exhibiting stronger dependence (positive or negative), while the dependence in other subsets might be weaker. A proposed solution refers to the concept of conditional correlation (see Longin and Solnik, 1998; Forbes and Rigobon, 2002; Hong *et al.*, 2007), 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.

The approach we adopt is in conceptual proximity to the traditional correlation analysis, but differs in several respects. 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. An additional challenge stems from the fact that linkages are not equally strong between all markets (Kaminsky and Reinhart, 2000). The effect of a shock on sovereign bond and CDS markets may be heterogeneous, varying with market conditions and the nature of the shock (e.g., a fiscal or financial shock, a sovereign downgrade, etc.). The local Gaussian correlation can detect these nonlinear and asymmetric changes in the dependence structure (for a discussion, see Støve and

[3]This is known as the bias problem for the 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.

Tjøstheim, 2014). It provides a direct measure of both average and upper-lower tail dependence that enhances our understanding of the interconnectedness between bonds and CDSs in the different segments of the distribution. The latter includes extreme market conditions (booming or crashing). Comparing the local correlation in the pre- and post-shock periods, allows us to test for contagion with a bootstrap procedure.

Empirical tests of contagion require some prior assumptions regarding the potential triggers. Thus far, these triggers were chosen exogenously and referred to country shocks that caused a significant increase in cross-market correlations (see, e.g., King and Wadhwani, 1990; Baig and Goldfajn, 1998; Calvo and Mendoza, 2000; Forbes and Rigobon, 2002).[4] Yet, these obvious shocks might not be suitable for identifying contagion: market volatility usually increases in unstable periods, and therefore these correlations will be larger than in stable periods and conclusions of contagion may be considered incorrect if synchronisation in volatility is not considered as a contagion effect (see, e.g., Forbes and Rigobon, 2002; Støve *et al.*, 2014). Furthermore, since these shocks refer to crisis periods with extreme asset price movements, this synchronization might not take place across the entire distribution; it can further materialize over consecutive days and not only on the day of the given shock. The local correlation measure captures changes of the correlation structure in the tails, the parts reflecting these extreme price movements.

The identification scheme is based on the Galeano and Weid (2014) algorithm for correlation change-point inference and can detect the number and position of multiple change points in the correlation structure of our data series. By assuming that expectations and variances are constant and that there are sudden shifts in the correlations during the period of the shock, this procedure identifies endogenously the number of the change points. These points are then employed to split the sample into a pre- and a post- correlation break period (endogenously) and on some a priori defined date. The consequent examination of contagion is then bound to focus on points more probable to reflect these structural changes.

We expect the interaction between the two markets to be significant, since by construction, sovereign CDS contracts and the underlying government bonds offer investors a similar exposure to the risk and return of sovereign debt. Their relative pricing is linked by a no-

[4]Such shocks include, inter alia, the 1987 US market crash, the 1994 Mexican devaluation, the 1997 Asian crisis, the onset of the Eurozone crisis in October 2009.

arbitrage condition: the CDS premium should be equal to the yield spread over a risk-free benchmark on a par floating-rate bond (Duffie, 1999). Contingent on this theoretical equivalence, traders try to exploit any short-term price differences between cash and synthetic markets to make risk–free profits. In the long–run, the no–arbitrage relation guarantees that the two markets are in equilibrium (for the euro area, see Palladini and Portes, 2011).

In the short–run, the two markets price credit risk differently to various degrees. These short-term inefficiencies are mainly attributed to three general mechanisms: (i) maturity transformation (see Diamond and Dybvig, 1983; Drechsler *et al.*, 2018), (ii) herding behavior (see Scharfstein and Stein, 1990), and (iii) leverage cycles (see Geanakoplos, 2009). For example, any entity (sovereign in our case) that faces a maturity mismatch between its expected revenues and debt obligations, must roll over its debt periodically. If investors are sufficiently pessimistic about its ability to refinance its debt, the sovereign may face a run on its bonds and/or a buyout of the underlying CDS. CDS contracts serve this purpose, by allowing pessimists to leverage without holding the underlying bond; this increases the cost of borrowing and leads to a further increase in the sovereign's CDS spreads and insolvency risk (see Geanakoplos, 2009). The aforementioned mechanisms are expected to exacerbate in periods of distress, such as those during the Eurozone crisis (owing to the herding behavior of investors in the CDS market).

Our results indicate that contagion is neither a single-source nor a single-market phenomenon in the context of the European sovereign debt crisis. In specific, "pure contagion" phenomena in the sovereign bond and CDS markets have undergone two major phases. The first, 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 indicating contagion are concentrated around important economic events. These include the Greek deficit's upward revision in November 2009, and the consequent fears about a possible Greek default, which ultimately led to the Greek government's official request of its first rescue plan in April 2010. Other events include the agreements on the Irish and Portuguese rescue plans (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 bond yields (a finding partly documented in Missio and Watzka, 2011; Metiu, 2012; De Santis, 2012; Beirne and Fratzscher, 2013; Claeys and Vašìcek, 2014; Caporin *et al.*, 2018). This opposite responsiveness between the bond and CDS markets with regards to the periphery bond-stemming contagion reveals an additional difference between the two: the capacity of the bond market to reduce shocks and/or losses within its context. This contrasts with the CDS market, where such shocks – reflected by the higher yields in the periphery bonds – are preserved and possibly amplified. The analysis of the unfiltered series reveals that European government bonds 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 bond pricing (Bekaert *et al.*, 2014; Claeys and Vašìcek, 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 bonds 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 as key countries for the evolution of euro area sovereign credit risk. This 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 an inclusive and detailed timeline of the amount and direction of contagion within the European sovereign debt and credit markets, where 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 was 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 confirm these arguments.

Our contribution further concerns the identification of the increased sensitivity of the CDS market to adverse economic shocks, especially in times of crisis. Contagion phenomena ignite and consequently propagate to countries with similar macroeconomic fundamentals. We thus highlight the prominence of sovereign CDS spreads in terms of price discovery and contagion capacity relative to their bond counterparts.

Importantly, we identify the potential mechanisms responsible for the emergence of contagion phenomena. As such, we focus on liquidity risk, which consisted 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, but only when stemming from the bond market. We thus content that credit risk is a stronger determinant of price discovery and contagion transmission between the two markets relative to liquidity risk.

On the same line, we further 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 nevertheless unable 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 (evident in Gyntelberg *et al.*, 2018), further restrict their contagion capacity.

Last, our study is the first to link the important policy changes on the regulatory and the monetary policy front during the Eurozone crisis with sovereign contagion dynamics. To this end, we provide evidence that the voting of the naked CDS ban exerted a containment effect on the transmission of CDS-bond contagion within the EMU South. We further consider the role of 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 the periphery countries' borrowing costs and restoring monetary policy transmission, rather than targeting Eurozone contagion.

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 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 Contagion identification scheme and the factor model

The concept of contagion is not unequivocal and alternative empirical tests for it exist. An intuitive starting point to measure a potential increase in interdependence could be looking at simple correlations between two default risk indicators. Seminal papers have focused on the cross-market correlations in stock markets before and after a financial shock (see, for instance, King and Wadhwani, 1990; Baig and Goldfajn, 1998; Calvo and Mendoza, 2000). Yet, simple correlations during crisis periods could be misleading, as one would expect higher correlations during periods of higher volatility (Forbes and Rigobon, 2002).

Conceptually, testing for contagion in our study imply the following five assumptions: (i) contagion is triggered by an adverse shock, (ii) the transmission of the shock is different from regular adjustments observed in tranquil times, (iii) the transmission of the shock differs across the segments of the distribution, (iv) contagion is sequential: cross-market linkages increase following a shock originated in the domestic sovereign bond or CDS market and (v) contagion occurs over and above what one would expect from economic fundamentals.

The last criterion, which is mainly attributed to Bekaert *et al.* (2005), allows us to differentiate between three contingencies. Assuming that bond and CDS spreads are adequate credit risk proxies and that changes follow a linear or a nonlinear factor structure, increased correlation between the two markets can be driven by three potential sources :

• an increase in the exposure of bond yields and/or CDS spreads to common factors, which is labelled as "wake-up call contagion". Practically, this occurs when there is evidence of increased local correlation on the unfiltered series (i.e., the fundamentals have not been accounted for).[5] The hypothesis of contagion between a given country-pair is verified when examining the unfiltered series but rejected when employing the fundamentals-filtered 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 unfiltered and the filtered series (i.e., it exists over and above economic fundamentals). The hypothesis of contagion is confirmed under both the unfiltered and the fundamentals-filtered series and this evidence should be interpreted as evidence of "pure contagion". 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 filtered series but not on the unfiltered ones. The hypothesis of contagion for a given country or country-pair is confirmed by the estimation of the fundamentals-filtered series but rejected by the estimation of the unfiltered 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.

To address these common risk factors, we follow De Bruyckere *et al.* (2013) and Fontana and Scheicher (2016) and condition bond yields and CDS spreads on seven state variables. More specifically, we use the three-month Euribor rate to proxy for the euro-wide risk-free rate; we expect a negative relationship between risk-free rate and credit spreads (Merton, 1974). By subtracting the three-month Euribor rate from the ten-year Euro Swap rate, we proxy for the slope of the term structure, since an increasing slope of the term structure should, ceteris paribus, lead to an increase in the expected future spot rate and to a decrease

[5]The results from this exercise are presented in the Appendix A.4. As in Forbes and Rigobon (2002), we have performed heteroscedasticity-filtering in the unfiltered series. If heteroscedasticity bias is ignored when testing for changes in correlation, then contagion is over-accepted.

in credit spreads through its effect on the drift of the asset value process. We expect the relationship between the term spread and each of the sovereign bond yields and CDS spreads to be negative.

To control for market-wide changes in business climate in each EMU country, we use the idiosyncratic equity returns, defined as the difference between the national equity-index returns and the benchmark Eurozone index return (STOXX Europe 600). We expect equity returns in a given country to be negatively related to that country's credit spreads. We calculate the idiosyncratic equity volatility as the annualized GARCH (1,1) volatility of idiosyncratic stock returns. We anticipate a positive relationship between idiosyncratic equity volatility and credit spreads. To control for market-wide credit risk, we consider the iTraxx Europe index, constructed as the equally weighted average of the 125 most liquid CDS series in the European market. A higher iTraxx indicates a higher overall default risk in the economy, thus pointing to a positive relationship between this index and each of the sovereign bond yields and CDS spreads (De Bruyckere *et al.*, 2013).

Global risk factors constitute an additional driver of sovereign credit spreads, as higher volatility is associated with higher economic uncertainty (Longstaff *et al.*, 2011). To isolate bond yields and CDS spreads from the impact of euro area volatility, we consider the implied volatility on the EuroStoxx 50, as reflected in the VSTOXX index. By deducting a GARCH-based estimate of volatility from the VSTOXX index, we obtain a proxy for the risk premium. The resulting variable measures the risk premium that investors in equity options require as compensation for bearing euro area equity market risk. This compensation (risk-premium) should be a positive function of credit spreads. We further control for exchange rate uncertainty, by using the 30-day implied EURO/USD exchange rate volatility index.[6] 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).

[6] The index is provided by the CBOE and follows the methodology of the VIX index.

We collect all state variables from Datastream and the regression specification of the factor model takes the following form:

$$\Delta Y_{it} = c + \Delta\beta_1 Rf_t + \Delta\beta_2 Term_t + \Delta b_3 RP_t + \Delta\beta_4 ERV_t + \Delta\beta_5 EqR_{it} \\ + \Delta\beta_6 EqV_{it} + \Delta\beta_7 iTraxx_{it} + e_{it} \qquad (1)$$

where ΔY_{it} is a vector representing the change in bond yield or in the CDS spread of a country i at time t, ΔRf_t is the change in the three-month Euribor rate, $\Delta Term_t$ is the change in the slope of the term structure, ΔRP_t is the change in risk premium, ΔERV_t is the change in the Euro/USD exchange rate volatility index, ΔEqR_{it} is the change in idiosyncratic equity returns, ΔEqV_{it} is the change in idiosyncratic equity volatility and the $\Delta iTraxx_{it}$ is the change in iTraxx Europe CDS index.[7]

However, most of the covariates in the different regressions of Equation (1) have a low explanatory power for our bond yield and CDS spread series. Fontana and Scheicher (2016), use a similar set of explanatory variables to proxy the determinants of CDS premia and bond spreads and obtain similar findings. Likewise, Groba *et al*. (2013) find that the explanatory ability of local and global factors is not as high for euro area CDS premia. This, may be attributed to the fact that the key drivers of sovereign credit risk have affected CDS premia and bond spreads in a different way during the crisis (see Fontana and Scheicher, 2016). In line with these studies, we employ a large set of control variables at the daily level, that theoretically should capture the time-varying effects of factors at the global, US, European, and domestic level. Other approaches however, highlight the possibility that common factors may change over time (see, e.g., Bekaert *et al*., 2014; Bekaert and Hoerova, 2016).[8] To confirm the stability of the parameters in our factor model, we perform a CUSUM test for each of our bond yield and CDS spread series.[9] Results from this test suggest parameter

[7] A battery of unit root tests for credit spreads and for the state variables expressed in levels, has not rejected the unit root hypothesis. Therefore we estimate this equation in changes only. Results are omitted for reasons of space but are available from the authors upon request.

[8] Bekaert *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.

[9] For brevity, the results of the CUSUM test are not presented here and are available upon request.

stability for nearly all bond yields and CDS spreads in our sample; the only exception is the Irish bond yields and CDS spreads in early 2011.[10]

The application of the bootstrap test for contagion necessitates the independence of each variable over time. To satisfy this, one could employ volatility filtered series that pick up any fat tails of the distribution. However, the use of standardized series may hide interesting dynamics that affect the results. To address this, we initially estimate the factor model of Equation (1) using OLS. Although this violates the assumption of independent and identical observations, we are nevertheless interested to see the effect on the results.[11] We subsequently estimate the factor model using a GARCH(1,1) model with normally distributed errors; although this will not pick up any fat tails in the return distribution as the Student's t distribution, it will remove heteroscedasticity (see Table 8 in the Appendix A.3).[12] By contrasting the two filtering methods, we observe that under OLS, the p-values (at the 5% level) are on average lower than under GARCH, due to the expected volatility effects. Most importantly, there is no difference in our conclusions about the presence of contagion between the two methods at the 5% level. Since both methods provide identical results, we estimate our factor model with OLS. We then employ the filtered series to examine changes in intra- and cross- market dependence and test for contagion effects.

3 Methodology

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 distribution that approximates that point (approximate the density locally rather than the correlation). The correlation of the approximating Gaussian distribution is taken as the local correlation in that neighbourhood.[13]

[10]For the case of Ireland, we estimate the factor model of Equation (1) by including a dummy variable for the respective period.

[11]In line with Støve *et al.* (2014), the filtering of the weight functions is wider. We experiment with different weight functions, but all provide qualitatively similar results.

[12]Both Ljung-Box test and the LM test for conditional heteroscedasticity imply that the fitted models are satisfactory.

[13]For the sake of brevity, a detailed description of the local Gaussian correlation procedure is provided in the

3.1 A bootstrap test for contagion

This section presents a test for contagion that uses the measure of local Gaussian correlation to examine whether cross-market linkages have increased (for more discussion and an application in financial markets see, Støve *et al.*, 2014). Contagion is confirmed if the local correlation for the crisis period has increased significantly compared to that before the crisis. The test was proposed by Støve *et al.* (2014) and is a bootstrap procedure.[14]

Denote $Z_t, t = 1, ..., T$ as the sovereign bond yields in country where the crisis started and $X_t, t = 1, ..., T$ the bond yields in another country. The data are filtered to remove time and volatility dependence (see also Forbes and Rigobon, 2002).[15] Let the standardised yields be written as $d_t = (X_t, Z_t)$. We then split the data in a pre-crisis period (NC) and a post-crisis period (C).[16] If the local correlation for the post-crisis period is significantly above the pre-crisis one, contagion is confirmed. Fixed gridpoints (x_i, z_i) for $i = 1, ..., n$ are used to estimate the local correlations. Thus, the null and the alternative hypothesis can be written as:

$$H_0 : \ \rho_{NC}(x_i, z_i) = \rho_C(x_i, z_i) \ \text{ for } i = 1, ..., n \ \text{ (no contagion)}$$
$$H_1 : \ \sum_{i=1}^{n} (\rho_C(x_i, z_i) - \rho_{NC}(x_i, z_i)) > 0 \ \ \text{(contagion)}$$

The bootstrap works by drawing observations $\{d_1, ..., d_T\}$ at random and replacing them in $\{d_1^*, ..., d_t^*\}$. Next, this resample is divided in NC and C and $\hat{\rho}^*_{NC}(x_i, z_i)$ and $\hat{\rho}^*_C(x_i, z_i)$ is computed on the grid (x_i, z_i) for $i = 1, ..., n$. The diagonal grid $(x_i = z_i)$ is employed in the subsequent analysis to minimize the computational time. The next step is to calculate:

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

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 other words, we avoid the estimation of local correlation in a gridpoint far away from any observations. Repeated resampling allows us to compute D_1^* for these resamples and to construct its distribution. Last, we calculate $\hat{\rho}_{NC}(x_i, x_i)$, $\hat{\rho}_C(x_i, x_i)$ and D_1 from the real

Appendix A.2.

[14] Similar ones are often used in a nonparametric setting, e.g., to for differences between quantities in non-parametric regressions (see, for example, Hall and Hart, 1990; Vilar-Fernandez *et al.*, 2007).

[15] In our empirical analysis an ARMA-GARCH(1,1) filter is used (see Section 3 for further details).

[16] Dungey *et al.* (2005), mention that tests of contagion can be affected by the predefined split of 'crisis' and 'non-crisis' periods. For this purpose, we use a number of alternative splits when we perform the bootstrap tests for contagion; all these give quantitatively similar results and are available from the authors upon request.

filtered observations $\{d_1, ..., d_T\}$. The p-value in terms of the D_1^* distribution is found and implies a rejection of H_0 if it is below a chosen significant level α.[17]

4 Data and descriptive statistics

4.1 Bond yields and CDS spreads

The dataset consists of bond yields and CDS spreads for sovereign bonds and sovereign CDS contracts with 5 years to maturity from January 2, 2006, to April 5, 2018.[18] The sample includes all EMU member states at the time of the introduction of the euro, 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 therefore, 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. Daily frequency is employed given that the 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. Moreover, our examination period includes some very tense periods and operating at a lower frequency would imply losing information. In sensitivity analysis, we further use the bid-ask spreads (calculated as: $ask\ price$ - $bid\ price$) on the sovereign bonds and sovereign CDSs with 5-year maturity and the CDS-bond basis for the same maturity from January 2 2006, to December 31, 2014.

Figure 1 (Panels A and C), 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 follow the opposite

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

[18]For Greece the sample ranges from January 2, 2006 to March 8, 2012 since the next day, on 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 is not included in the sample. The two-letter country codes that are used in abbreviation in our analysis are taken from the International Standard for country codes (ISO 3166).

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 of the periphery countries. As shown in Figure 1 (Panels B and D), CDS spreads in the periphery match closely the upward trend of the periphery bond yields, pointing to a close association between the two asset-markets. However, CDS spreads also increased in the core, indicating an overall rise in sovereign default probabilities across the entire Eurozone.

[Insert Figure 1 about here.]

Table 2, presents descriptive statistics for daily changes in sovereign 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.

[Insert Table 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 Galeano and Weid (2014).[19] This 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 the relationship between those markets. The identification of a structural change in the cross- and intra-asset correlations further allows us to split the sample into a pre- and a post-event period, or more properly into a pre- and post-contagion period.[20] We then quantify the im-

[19]The Galeano and Weid (2014) methodology for the identification of structural changes in correlation is given in Appendix A.1.

[20]The test results for detecting structural changes in correlation are presented in Appendix A.4.

pact of the structural change by estimating the transmission of a shock to the bond and CDS markets during the respective periods.

Consistent with our priors, we conduct the analysis by employing the bond and CDS data after accounting for fundamentals via the factor model of Equation (1). However, we conduct a similar analysis for the original bond and CDS data without the removal of fundamentals (results from this exercise are presented in the Appendix A.5).[21] By contrasting the results from the two methods, we can differentiate between cases of "pure contagion" from cases of "wake-up call contagion" and "limited contagion". In this respect, if the hypothesis of contagion between a given country-pair is verified by the estimation of unfiltered series but rejected by the estimation of fundamentals-filtered series, this would serve as evidence of "wake-up call contagion". In contrast, findings of contagion under both the unfiltered and fundamentals-filtered series should be interpreted as evidence of "pure contagion". Last, if contagion is verified by only the filtered 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

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; bootstrap test for contagion is not able to reject the null hypothesis of no-contagion in practically all pairs (columns 1 to 5 in Panel A). This absence of shift contagion stemming from the distressed countries of the European periphery (also partly documented in Missio and Watzka, 2011; Metiu, 2012; De Santis, 2012; Beirne and Fratzscher, 2013; Claeys and Vašìcek, 2014; Caporin *et al.*, 2018) is not necessarily a sign of widespread disassociation between the sovereign debt markets.[22] It can serve as (i) supporting evidence of the investors' flight(s) away from the risky periphery

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

[22] The 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.

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 government bond yields can be explained in terms of the – enduring– interdependence between them.

[Insert Table 3 about here.]

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. Most importantly, these results are derived from the analysis of the fundamentals-filtered series, therefore pointing to contagion over and above that suggested by fundamentals ("pure contagion"). Interestingly, this transmission occurs over a series of different phases, all of them 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.[23] 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.[24]

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

[23]The 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 during his first Eurogroup meeting. In the budget draft for fiscal year 2010 that was submitted to the Hellenic Parliament for consideration in the November 5, 2009, the 2009 deficit was revised to 12.7%, while in the final draft of November 16, 2009 (that was voted by the Parliament), the budget was revised to 13.6%.

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

and tax increases in order to reduce its budget deficit from 31% of GDP to 3% by 2014. These 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.[25]

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.[26] 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, (July 15, 2011) the European Banking Authority (EBA) announced that five Spanish banks failed its "stress tests", while seven other Spanish banks barely passed.[27] 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 EU/IMF bailout package totaling €109 billion. However, this agreement was not able to contain speculations on a potential 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 the continuing speculations about a looming Greek exit were diffused across the Euro-

[25] On November 21, 2010, the Irish Prime Minister Brian Cowen announced that Ireland has applied for aid from the EU and IMF. On November 28, 2010, the Irish government accepted a €67.5 billion joint EU/IMF bailout package.

[26] On April 6, 2011, the Portuguese Prime Minister José Sócrates extended a request to the European Union for a financial assistance programme. On May 6, 2011, the Portugal reached an agreement with the EU and the IMF for a financial assistance programme of €78 billion.

[27] The 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%.

zone CDS markets, via the rising Italian and Spanish sovereign bond yields. This period of contagion transmission is further consistent with the onset of the upward trend in the Italian and Spanish government bond yields, evidenced in Panel B of Figure 1.

We consequently contrast the results in Table 3 with the results from the unfiltered series in Table 11 of the Appendix. Contrary to Panel A in Table 3, estimates in Panel B of Table 11 provide evidence of widespread contagion from each of the GIIPS bond markets to practically all European CDS markets. Since these findings are derived from the analysis of the unfiltered series, but not confirmed from the analysis of the filtered ones, they point to a form of "wake-up call contagion". Importantly, most correlation change-dates fall within the late 2008-early 2009 period. Thus, we can infer that movements in the European sovereign bond markets were part of the general repricing of sovereign credit risk after the global financial crisis and during the evolution of the Eurozone crisis. Turning to Panel B in Table 11, we observe that estimates exhibit only marginal deviations from those of Panel B in Table 3.

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

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 bounds of the EMU South and Belgium. The majority of correlation change-points indicating contagion are observed during the third quarter of 2011 and in particular November 2011. Initially, contagion appears to be stemming 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 the respective period, which had implications for the viability of the Eurozone itself. 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.[28] Shortly after the referendum call, efforts for the formation of government of national unity in Greece temporarily collapsed (November 4, 2011),

[28]The 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.

only to resume successfully seven days later.[29]

[Insert Table 4 about here.]

Our estimates reveal that this heightened period generated contagion transmission from the Greek CDS spreads towards the periphery bond yields. Shortly after, contagion phenomena further emerged from the rest of the periphery economies. Results in columns 2-5 of Panel A show that the CDS markets of Ireland, Italy, Portugal, and Spain transmit 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. Most importantly, this transmission is not verified by the analysis of the unfiltered 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 12 in the Appendix). This in turn reveals that this was a short-lived episode of "limited contagion" owing to the abrupt political uncertainty in Greece that shortly receded.

The analysis at the intra-CDS market level (see 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 the accompanied political developments in Greece – were found responsible for the transmission of CDS-stemming contagion to periphery bond markets. According to 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

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

leading to the Greek debt restructuring of March 2012.[30] 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, the Netherlands) and Belgium. Notwithstanding their size, the increasing influence of Italian and Spanish CDSs can also be explained when considering the nature of the CDS contracts. By construction, CDSs mainly reflect sovereign credit risk. Hence, a possible default by either Italy and/or Spain could trigger domino effects that could eventually lead 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 fundamentals-filtered series in Table 4 with those from the unfiltered series in Table 12 of the Appendix. Panel A of Table 12, points to certain cases of CDS-bond contagion from the periphery (Italy, Portugal, Spain) towards the core (Germany, Finland, France, the 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 bond yields across the core and attributed to the investors' upward repricing of those countries' fundamentals, prompted by the rising CDS spreads across the periphery. Last, estimates from Panel B of Table 12, do not reveal additional cases of intra-CDS contagion that are not consequently confirmed by the main estimations in Panel B of Table 4.

[30]On 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 involvement of the private sector, whereby 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.

5.3 Discussion

Our analysis identifies two key phases for the transmission of contagion within the European sovereign financial market 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 during these 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 therefore appears that the European sovereign bond and CDS markets differ with regards to their capacity to preserve or amplify potential shocks occurring in the periphery.

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 blocs' CDS spreads. The respective 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.

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. Importantly, 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.[31] Due to their economic size (third and fourth EMU econ-

[31] In Italy, the low rates of productivity and output growth were not keeping up with an increasing debt load

omy respectively), both countries are considered too expensive for the EMU members 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 had enough funds to accomplish such a task, thereby constituting either countries as too big to bail out. Hence, the respective 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 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 unfiltered series (Tables 11 and 12 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 crises provided new information, prompting investors to reassess the vulnerability of other countries, leading to a repricing of sovereign bond pricing factors (Bekaert *et al*., 2014; Claeys and Vašìcek, 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.

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 Eurozone. 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.

6 Further analysis

Our results in the previous section show that the countries of the EMU South acted as transmitters of contagion not only within their own country-bloc but most importantly, towards the countries of the EMU North. Moreover, 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 enlighten our findings, and the heterogeneity around certain fiscal and regulatory events. The role of liquidity and basis deviations for contagion transmission is also examined.

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 remaining ten 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 being the break-date. We observe that following 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 the respective period. For what matters, 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 consequently test for contagion from November 16, 2009 until April 5, 2018 (the end of our sample), with break-date April 23, 2010. Interestingly, 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. Overall we verify an increased sensitivity of the European CDS market to the negative developments in Greece. The estimates between the Greek sovereign 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 pre-crisis period has moved up. It should further be noted that in most cases there is a similar increase in local correlation over large segments of the distribution. This in turn, points to the existence of strong linkages between the Greek bond market and the European CDS markets during the post-crisis 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 pre-April 2010 period, that declines post-April 2010. Interestingly, 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 in turn points to the existence of strong linkages (albeit not contagious) between the Greek bond and these countries' 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 sovereign bond driven by 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 joint EU/IMF financial assistance 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.[32] 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 our analysis in 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 might be the case that the intense political events in Greece curtain the easing effect of the regulatory ban brought about by the passage of the law, especially on the government bonds of Ireland, Italy, Portugal, and Spain.

To ensure that we adequately isolate the effect of this regulatory change, we further examine its impact on the transmission of contagion. We expect that the main operating channel of transmission is from the CDS market to the bond market. Moreover, although the regulation's effective date was known in advance, the voting outcome was fairly unanticipated. To this end, Table 6 examines differences in the transmission of contagion from the GIIPS CDSs to all European bonds in the periods before and after the regulation's passage in November 15, 2011. Results from this exercise provide no evidence of contagion from the periphery CDS markets to any of the two blocs following the EU Parliament's voting: all p-values generated by our test are non-statistically significant after November 15, 2011.

[Insert Table 6 about here]

To fully detect whether this reflects a change from the pre-November 2011 period, we further examine the evolution of the dependence structure between our series. Figure 4,

[32] The 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 contacts 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.

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 November 15, 2011 (green line) and the period after (red line). Evidently, there is a strong positive correlation between the CDSs and bonds of GIIPS pre-November 2011, which recedes 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 passage of 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 usually referred to as "flight-to-quality" or "flight-to-liquidity" (see Beber *et al.*, 2009; Brunnermeier, 2009; Monfort and Renne, 2014; Fontana and Scheicher, 2016; 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 (an illiquidity contagion) across the same bloc, 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).[33] 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 (filtered data) 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.

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 the 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; importantly, 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, 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 only 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 if anything, liquidity risk is a material source of risk in the context of the European sovereign CDS market only when stemming from the sovereign bond market. Moreover, credit risk emerges as a stronger determinant of price discovery between the two markets relative to liquidity risk.

[33]The 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).

6.4 Additional tests

We conduct additional sensitivity tests, the results of which are available on request. First, we consider the role of arbitrage opportunities as explained by the deviations of the CDS-bond basis (the 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. To accomplish this, we estimate Equation (1) with the 5-year sovereign CDS-bond basis as dependent variable, and calculate the residuals. We subsequently test for contagion between our filtered bond yields and the filtered 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 the two blocs. 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).

Second, 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 the monetary policy transmission in the euro area (see Eser and Schwab, 2016). As such, we expect that these purchases contributed to the minimization of the 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 consequently 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 the respective 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 the 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, the results in this section suggest that sovereign bond and CDS contagion dynamics have exhibited heterogeneities during the Eurozone crisis. These heterogeneities are mainly associated with the occurrence of certain fiscal and regulatory events during the main phase of the crisis. Furthermore, the 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 European 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 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 either 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 bond pricing (Bekaert *et al.*, 2014; Claeys and Vašìcek, 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, but only when stemming from the bond market. We content that credit risk is a stronger determinant of price discovery and contagion transmission between the two markets relative to liquidity risk. 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. Last, by studying the regulatory response during the crisis, we provide evidence that the voting of 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 that to future research.

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Table 1: This table presents the changes in basis points on 5-year sovereign bond yields and CDS spreads after a specific adverse event on 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 triple B plus 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 Δ(Bond) vs Δ(CDS)	
		AT	BE	DE	ES	FI	FR	GR	IE	IT	NL	PT	pre-event	post-event
	Panel A: Δ(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: Δ(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 2: 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 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 Galeano and Weid (2014) based on the standardized yields–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 1000 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
Panel A : Bond Markets										
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	22/11/2009				
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	22/11/2009					0.635	20/5/20111	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	20/6/2011
							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
Panel B : CDS Markets										
AT	**0.039**	**2/11/2009**	**0.035**	**15/11/2010**	**0.008**	**18/7/2011**	**0.000**	**28/3/2011**	**0.000**	**1/8/2011**
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 Galeano and Weid (2014) based on the standardized yields–spreads. 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 1000 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
Panel A : Bond Markets										
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**	**1/8/2011**	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**
Panel B : CDS Markets										
AT			0.992	4/1/2013	**0.032**	**21/11/2011**	0.822	15/9/2008	0.988	15/9/2008
							0.996	1/2/2012	**0.032**	**3/1/2012**
BE			0.994	28/1/2012	**0.000**	**9/1/2012**	0.963	13/1/2012	**0.045**	**9/1/2012**
DE			0.788	13/1/2012	**0.000**	**16/1/2012**	0.975	15/9/2008	**0.021**	**18/1/2012**
							0.992	13/1/2012		
ES	**0.041**	**18/11/2011**			**0.000**	**21/11/2011**	0.987	31/10/2011		
FI	**0.000**	**1/10/2008**	0.899	19/1/2012	**0.000**	**1/12/2011**	0.641	2/1/2012	**0.010**	**14/11/2011**
FR					**0.019**	**9/1/2012**	0.418	12/11/2009	**0.046**	**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**	**14/9/2010**			**0.009**	**18/11/2011**
IE	**0.000**	**11/11/2011**			**0.021**	**15/11/2011**	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	**0.000**	**21/11/2011**
			0.784	31/8/2015			0.994	3/9/2012	0.885	
NL	**0.000**	**12/9/2008**	0.760	7/12/2012	**0.005**	**20/1/2012**	0.716	5/9/2008	0.772	15/9/2008
	0.911	13/9/2010					0.988	13/1/2012	**0.000**	**18/1/2012**
PT			0.967	5/1/2012	**0.000**	**14/11/2012**			**0.009**	**31/10/2011**
					0.991	3/9/2012				

Table 5: This Table presents the estimated p-values from the bootstrap test for contagion from the Greek bond yields (Column 2) and CDS spreads (Column 4) to the rest of the European sovereign CDS markets. Greece (GR) is considered as the country of origin for the European sovereign debt crisis. 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 1000 replications.

	Deficit Revision November 2009		Financial Request April 2010	
Origin:Greece	$Bonds$(GR)	Contagion?	CDS(GR)	Contagion?
$European\ CDS\ Markets$				
Austria	0.001***	Yes	0.992	No
Belgium	0.004***	Yes	0.885	No
Finland	0.001***	Yes	0.912	No
France	0.000***	Yes	0.709	No
Germany	0.005***	Yes	0.763	No
Ireland	0.050**	Yes	0.446	No
Italy	0.021**	Yes	0.000	No
Netherlands	0.009***	Yes	0.770	No
Portugal	0.013**	Yes	0.414	No
Spain	0.000***	Yes	0.558	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. The bootstrap test for contagion is based on 1000 replications.

	GR (1)	IE (2)	IT (3)	PT (4)	ES (5)
$European\ Bond\ Markets$					
AT	0.512	0.899	0.377	0.788	0.521
BE	0.487	0.525	0.614	0.411	0.947
DE	0.215	0.395	0.881	0.632	0.628
ES	0.855	0.884	0.934	0.941	–
FI	0.860	0.647	0.865	0.548	0.488
FR	0.412	0.912	0.954	0.684	0.923
GR	–	0.662	0.266	0.985	0.912
IE	0.350	–	0.744	0.325	0.624
IT	0.975	0.789	–	0.478	0.784
NL	0.455	0.246	0.989	0.658	0.998
PT	0.990	0.411	0.221	–	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 Galeano and Weid (2014) 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 1000 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
Panel A : Bond Markets										
AT	0.912	7/11/2011					0.890	22/9/2008		
BE	0.988	6/10/2009	0.690	4/1/2010	0.885	29/6/2011	0.540	6/10/2010	0.964	22/06/2011
DE	0.717	2/10/2009	0.754	14/12/2009	0.996	7/7/2011			0.556	5/7/2011
ES	0.991	17/09/2009	0.662	18/3/2010	0.785	13/7/2011	0.412	12/1/2011		
FI	0.998	3/11/2011	0.993	10/2/2010						
FR	0.944	21/09/2009	0.991	5/3/2010			0.911	11/11/2009		
GR			0.090	14/10/2009	0.680	29/6/2011	0.630	22/09/2009	0.993	17/09/2009
IE	0.431	14/10/2009			0.545	29/3/2010	0.997	28/1/2009	0.525	18/3/2010
IT	0.989	29/6/2011	0.756	2/2/2010	0.689	16/11/2009	0.754	25/4/2011	0.669	13/7/2011
NL	0.955	24/11/2011	0.460	10/12/2009	0.830	10/5/2010	0.674	6/10/2009		
PT	0.990	22/09/2009	0.889	28/1/2009	0.332	25/4/2011			0.285	12/1/2011
Panel B : CDS Markets										
AT	0.995	12/10/2009	0.344	25/11/2010	0.511	21/7/2011	0.878	21/3/2011	0.691	25/7/2011
BE	0.110	20/11/2009					0.996	30/3/2011	0.785	15/8/2011
DE	0.443	30/11/2009	0.675	8/11/2010	0.997	11/7/2011	0.994	5/4/2011	0.887	15/8/2011
ES	**0.000**	**16/11/2009**	**0.002**	**24/11/2010**	**0.006**	**2/8/2011**	**0.004**	**14/4/2011**		
FI	0.998	24/11/2009	0.887	2/11/2010			0.422	28/4/2011	0.998	12/8/2011
FR	0.898	23/11/2009			0.832	4/7/2011			0.991	2/8/2011
GR			**0.000**	**28/10/2009**	**0.000**	**2/11/2009**	**0.012**	**12/11/2009**	**0.036**	**16/11/2009**
IE	**0.004**	**28/10/2009**			**0.032**	**20/6/2011**	**0.044**	**9/12/2010**	**0.047**	**24/11/2010**
IT	**0.000**	**2/11/2009**	**0.008**	**9/12/2010**			**0.007**	**18/7/2011**	**0.022**	**2/8/2011**
NL	0.687	19/11/2009	0.541	1/12/2010	0.745	7/7/2011	0.132	9/5/2011	0.998	21/7/2011
PT	**0.040**	**12/11/2009**	**0.000**	**22/11/2010**	**0.021**	**18/7/2011**			**0.044**	**14/4/2011**

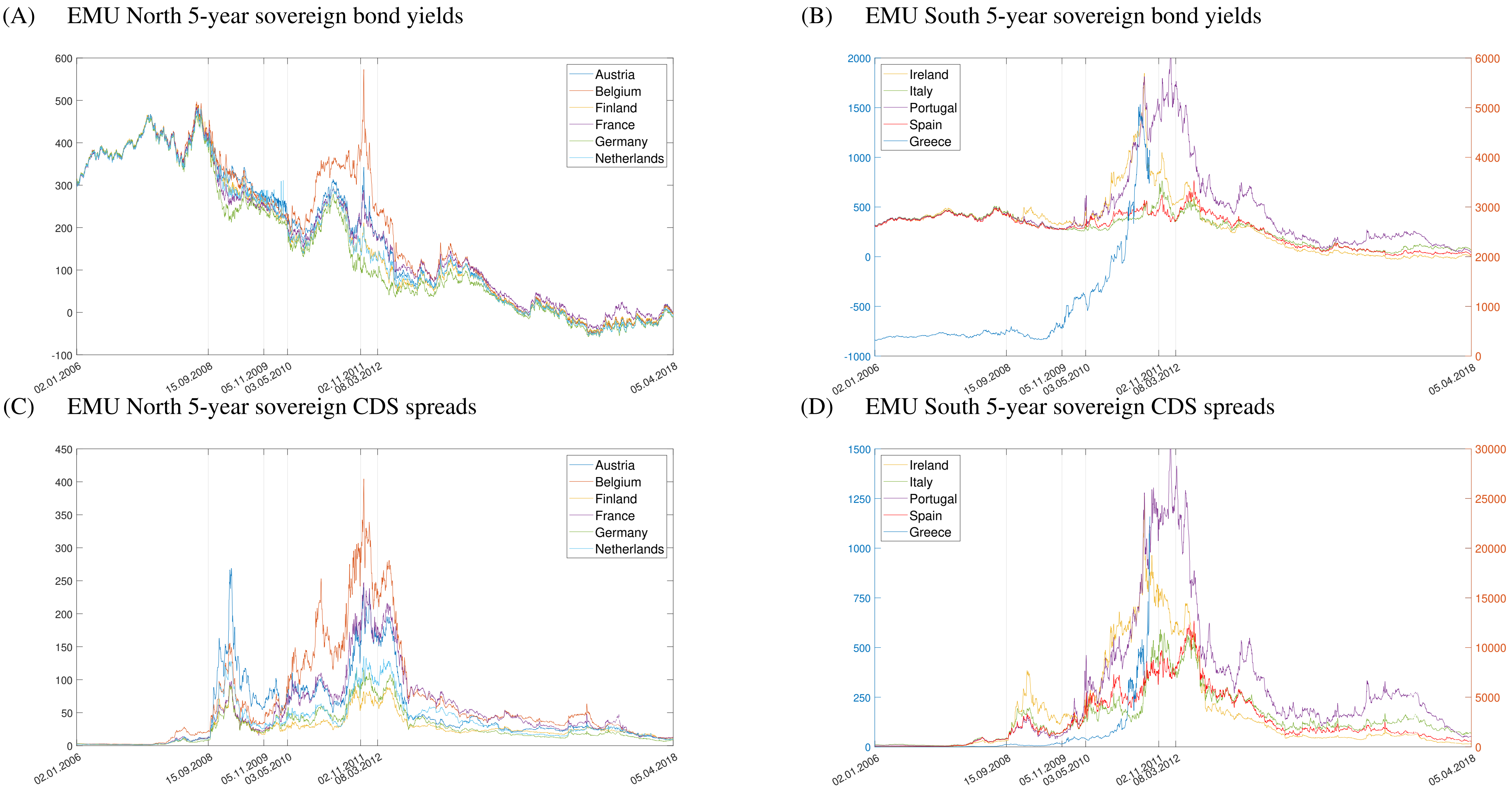

Figure 1: Evolution of sovereign bond yields and CDS spreads (in basis points). Panels A and C show the 5-year sovereign bond yields and CDS spreads for the countries of the EMU North (Austria, Belgium, Finland, France, Germany, Netherlands) and Panels B and D show the 5-year sovereign bond yields and 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.

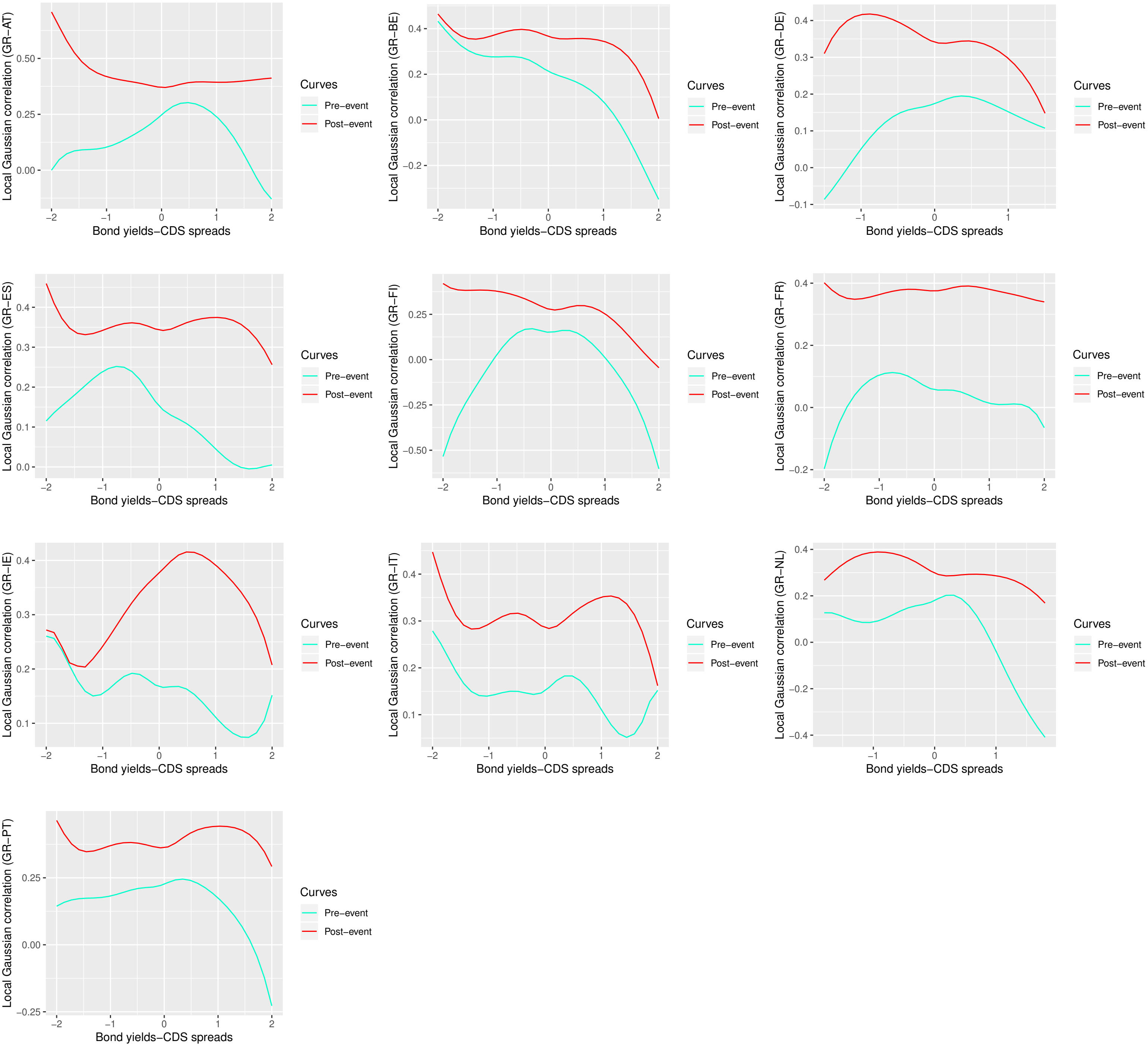

Figure 2: Local Gaussian correlation estimates between the Greek sovereign bond yields and each of the European sovereign CDS yields following Greece's upward deficit revision on November 16, 2009.

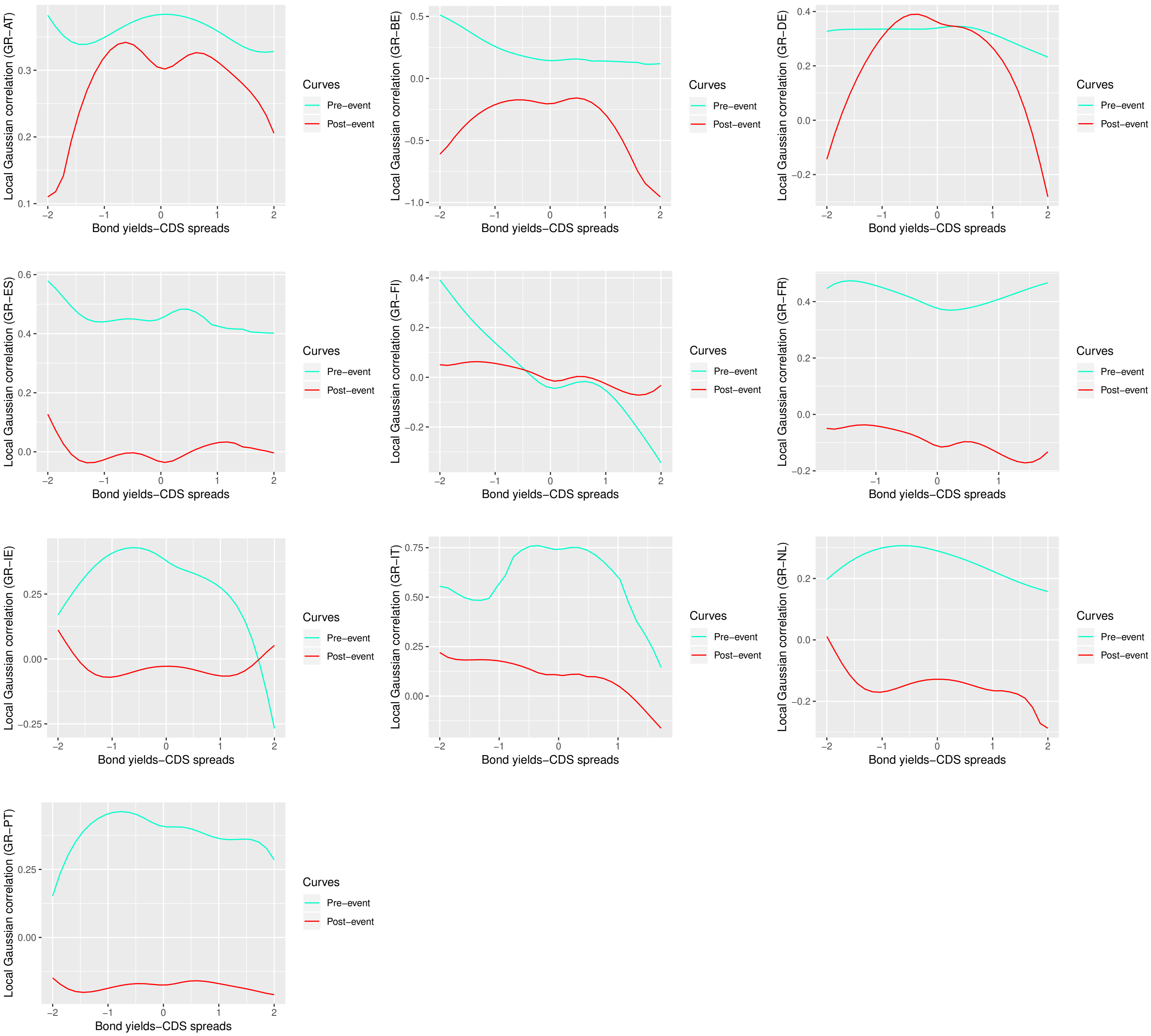

Figure 3: Local Gaussian correlation estimates between the Greek sovereign bond yields and each of the European sovereign CDS yields following Greece's request for financial assistance in April 23, 2010.

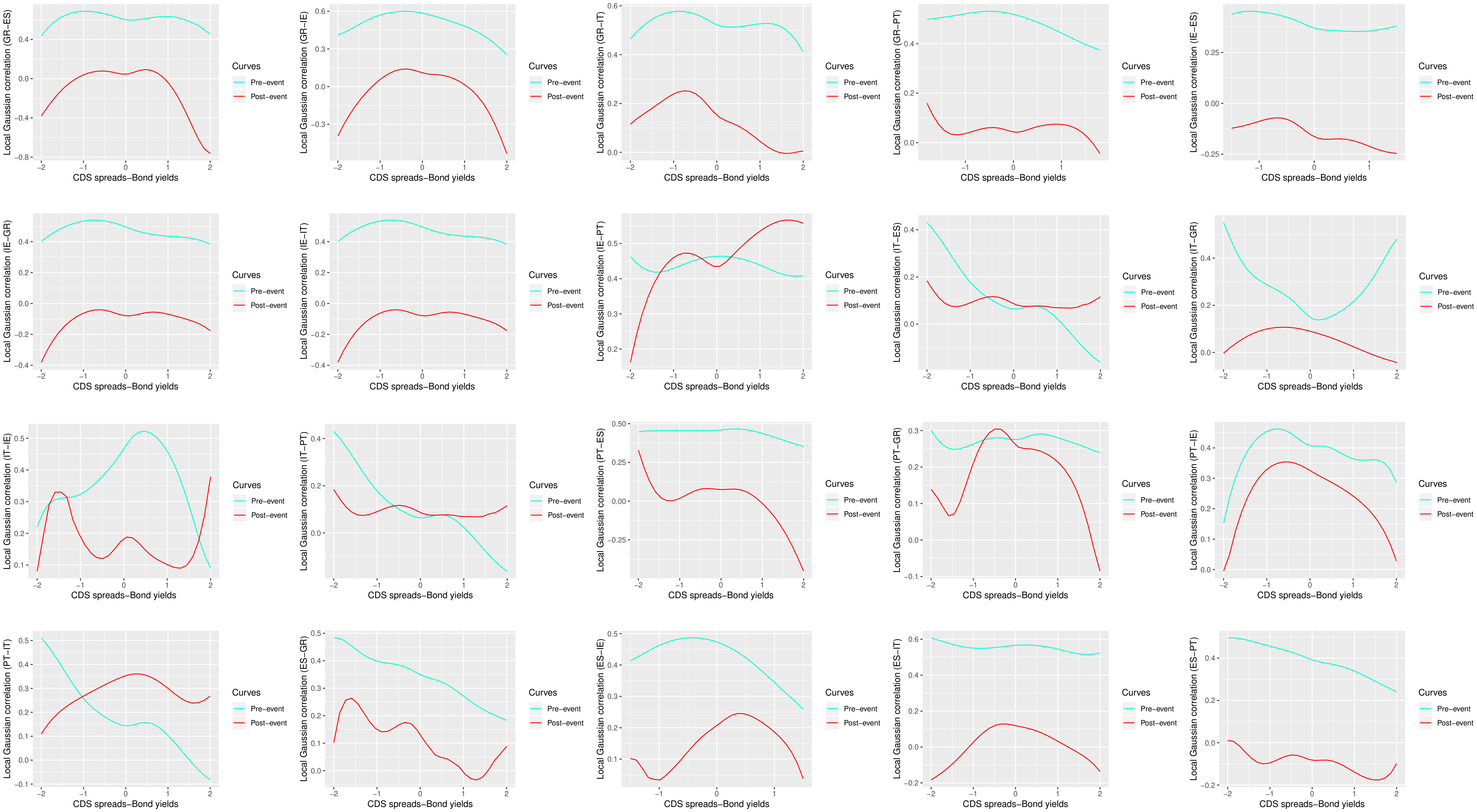

Figure 4: Local Gaussian correlation estimates between the GIIPS CDS spreads and each of the European sovereign bond yields following the voting of the naked CDS ban on November 15, 2011.

A Appendix

A.1 Galeano and Weid (2014) algorithm for the identification of structural changes in correlation

Given a sample of $1...T$ observations of the returns vector $(y_{1,t}, y_{2,t})^{'}$, let ρ_t denote the true but unknown unconditional correlation between $y_{1,t}$ and $y_{2,t}$ at time t.

The algorithm starts with testing the null hypothesis of constant correlations against the alternative hypothesis of a change-point t^c:

$H_0 : \rho_t = \rho$ for all $t \in \{1, ..., T\}$ vs $H_1 : \exists t^c \in \{1, ..., T-1\}$ such that $\rho_{t^c} \neq \rho_{t^c+1}$

This is accomplished using the model-free fluctuation-type test originally proposed by Wied *et al.* (2012). The test statistic is defined as:

$Q_T := \widehat{D} \underset{2\leq t\leq T}{max} \frac{t}{\sqrt{T}} \left|\widehat{\rho}_t - \widehat{\rho}_T\right|$ (A.1)

where $\widehat{\rho}_t$ is the sample correlation over the period 1 to t. The scalar coefficient $\widehat{D}$ is needed to adjust for correlation breaks that appear at the beginning of the sample where $\widehat{\rho}_t$ is more volatile, and is constructed as follows:

Let $\left\{(y_{1,t}, y_{2,t})'\right\}$ be the bivariate time-series with $E\left((y_{1,t}, y_{2,t})'\right) = 0$. For $i = 1, 2$ denote $\overline{y_i} = T^{-1}\sum_{i=1}^{T} y_{i,t}$, $\overline{y_i^2} = T^{-1}\sum_{i=1}^{T} y_{i,t}^2$ and $\widehat{\sigma}_{y_i} = \sqrt{\overline{y_i^2} - \overline{y_i}^2}$. Furher, denote $\overline{y_1 y_2} = T^{-1}\sum_{t=1}^{T} y_{1,t}y_{2,t}$ and $\widehat{\sigma}_{y_1 y_2} = \overline{y_1 y_2} - \overline{y_1}\cdot\overline{y_2}$. Let $k(\cdot)$ be the Bartlett kernel fuction. The scalar $\widehat{D}$ is then given by:

$\widehat{D} = \sqrt{\widehat{D}_3 \widehat{D}_2 \widehat{D}_1 \widehat{D'}_3 \widehat{D'}_3}$, where $\widehat{D}_1 = \sum\limits_{t=1}^{T}\sum\limits_{u=1}^{T} k\left(\frac{t-u}{\lfloor \log T \rfloor}\right) V_t V_u'$ with

$V_t = T^{-1/2}\left(y_{1,t}^2 - \overline{y_1^2}, y_{2,t}^2 - \overline{y_2^2}, y_{1,t} - \overline{y_1}, y_{2,t} - \overline{y_2}, y_{1,t}y_{2,t} - \overline{y_1 y_2}\right)^{'}$,

$\widehat{D}_2 = \begin{pmatrix} 1 & 0 & -2\overline{y_1} & 0 & 0 \\ 0 & 1 & 0 & -2\overline{y_1} & 0 \\ 0 & 0 & -\overline{y_2} & -\overline{y_1} & 1 \end{pmatrix}$

and

$\widehat{D}_3 = \left(-\frac{1}{2}\frac{\widehat{\sigma}_{y_1 y_2}}{\widehat{\sigma}_{y_2}}\widehat{\sigma}_{y_1}^{-3} - \frac{1}{2}\frac{\widehat{\sigma}_{y_1 y_2}}{\widehat{\sigma}_{y_1}}\widehat{\sigma}_{y_2}^{-3} \cdot \frac{1}{\widehat{\sigma}_{y_1 y_2}}\right)$

The purpose of the scalar $\widehat{D}$ is to appropriately rescale the cumulated sum of empirical correlation coefficients in such a way that convergence of Q_T to the asymptotic null distribution is achieved.

Under the null hypothesis and several reasonable moment and dependency restrictions, the test statistic Q_T is asymptotically Kolmogorov distributed (Wied *et al.*, 2012, Theorem 1). If Q_T stays below the upper critical value (see Eq.A.1) the null hypothesis of constant correlation cannot be rejected and the algorithm stops. Otherwise, H_0 is rejected and there is at least one change-point t^c within the sample period. The estimator for the single change-point is given by:

$t^c = \arg \underset{t}{max}\ \widehat{D}\ \frac{t}{\sqrt{T}} \left|\widehat{\rho}_t - \widehat{\rho}_T\right|$

To identify further change-points, the sample is split into the two subsamples $\left[1, ..., \widehat{t^c}\right]$ and $\left[\widehat{t^c} + 1, ..., T\right]$. These subsamples are then both tested individually. This procedure is repeated until no further change-points are detected. Galeano and Weid (2014) show that the presence of multiple change-points can affect the test's efficiency in identifying the true number of change points.

The last step of the algorithm therefore consists of a refining process in which the vector of the n detected change-points $\tau = \left[\widehat{t_1^c}, ..., \widehat{t_n^c}\right]$, sorted in ascending date order $\widehat{t_1^c} \leq ... \leq \widehat{t_n^c}$, is verified in subsamples containing only a single change point. Define the first observation of the sample as $\widehat{t}_0^c = 0$, the last observation as $\widehat{t}_{n+1}^c = T$, and form the subsamples $\left[\widehat{t}_{i-1}^c + 1, ..., \widehat{t}_{i+1}^c\right]$ for $i = 1, ..., n$. Each subsample starts at the first observation following the previous change-point $\widehat{t}_{i-1}^c$, includes change point $\widehat{t}_i^c$,but ends just before the next change-point $\widehat{t}_{i+1}^c$. These subsamples are tested individually. If the null hypothesis is not rejected the change-point contained in the subsample is removed from τ.

The following brief example should clarify the test procedure. To test for a break in the daily correlation between Greek and German bond yields we start with the full sample from January 2, 2006 (02/01/2006) to April 5, 2018 (05/04/2018). In the first iteration, the procedure detects a change in the correlation at time point 28/10/2008 (October 28, 2008). Following the proposed procedure, we split the series into two subperiods and look for changes in the subintervals [02/01/2006–28/10/2008] and [29/10/2008–05/04/2018], respectively. In the second subinterval, the procedure detects a change at time point 13/11/2009. The test statistic in the first subsample is insignificant. Then, we split the subinterval [29/10/2008–05/04/2018] into two subintervals and look for changes in the subintervals [29/10/2008, 13/11/2009] and [13/11/2009, 05/04/2018]. No more changes were found in these two subintervals. Then, we pass to step 3 and refine the search. For that we estimate the location of the change points in the intervals [02/01/2006–28/10/2008] and [13/11/2009, 05/04/2018], respectively. The test statistics for both subsamples remain significant and confirms the presence of both changepoints:τ = [28/10/2008,13/11/2009]. According to Galeano and Weid (2014), this procedure detects the correct number of correlation change-points.

A.2 Local Gaussian approximation and local correlation

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

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

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 $n_S \to \infty$, $\hat{\rho}_{\mathbf{c}}(S)$ would converge to $\hat{\rho}_{\mathbf{c}}(S) = corr(X, Z \mid (X, Z) \in S)$.

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,

$$\phi(u, \theta(y)) = \frac{1}{2\pi\sigma_1(y)\sigma_2(y)}\times$$

$$\exp\left\{-\frac{1}{2(1-\rho^2(y))}\left[\left(\frac{u_1-\mu_1(y)}{\sigma_1(y)}\right)^2 + \left(\frac{u_2-\mu_2(y)}{\sigma_2(y)}\right)^2 - 2\rho(y)\left(\frac{u_1-\mu_1(y)}{\sigma_1(y)}\right)\left(\frac{u_2-\mu_2(y)}{\sigma_2(y)}\right)\right]\right\} \tag{3}$$

where $u = (u_1, u_2)^\intercal$ 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,[34]

$$q = \int K_b(u-y)\left[\phi(u, \theta(y)) - \log\phi(u, \theta(y)) f(u)\right] du \tag{4}$$

where $K_b = (b_1 b_2)^{-1} K\left(b_1^{-1}(u_1 - y_1)\right) K\left(b_2^{-1}(u_2 - y_2)\right)$ 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 $\rho(y) \succ 0\ (\rho(y) \prec 0)$. The bias of conditional correlation is accommodated since the same Gaussian f fits every point.

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

$$L(Y_1, ..., Y_n, \theta_b(y)) = n^{-1}\sum_i K_b(Y_i - y)\log\phi(Y_i, \theta_b(y)) - \int K_b(u-y)\,\phi(u, \theta_b(y))\,du \tag{5}$$

[34]This 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:[35]

$$\frac{\partial q}{\partial \theta_j} = \int K_b(u-y) \frac{\partial}{\partial \theta_j} \{\log(\phi(u,\theta(y)))\} [\phi(u,\theta(y)) - f(u)] \, du, \; j = 1, ..., 5 \tag{6}$$

Using the notation,

$$\gamma_j(\cdot,\theta) = \frac{\partial}{\partial \theta_j} \{\log \phi(\cdot,\theta)\}, \tag{7}$$

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

$$\frac{\partial L}{\partial \theta_j} = n^{-1} \sum_i K_b(Y_i - y) \gamma_j(Y_i, \theta(y)) - \int K_b(u-y) \gamma_j(u,\theta(y)) \phi(u,\theta(y)) \, du$$

$$\to \int K_b(u-y) \gamma_j(u,\theta(y)) [f(u) - \phi(u,\theta(y))] \, du = -\frac{\partial q}{\partial \theta_j}, \; j = 1, ..., 5 \tag{8}$$

as $n \to \infty$, we see that (7) can be identified with (5). Also note that as $b \to \infty$ (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 $b_i \to 0, i = 1, 2$ 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 $\theta_{n,b}(y) \to \theta_b(y)$ for b fixed, and $\theta_{n,b}(y) \to \theta(y)$ almost surely for $b = b_n$ tending to zero.[36]

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).[37] Since both approaches provide qualitatively similar results, we present the bandwidth choice based on the normal-reference rule-of-thumb.[38]

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

[36] The 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).

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

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

A.3 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 GARCH(1.1) model with normal errors for different lags (p= 5, 10, 15 and 20). 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 (p= 5, 10, 15 and 20). These are based on the Lagrange multiplier test for conditional heteroscedasticity of Engle (1982).

	Bond markets				CDS markets			
	$Q^2(5)$ [$p-value$]	$Q^2(10)$ [$p-value$]	$Q^2(15)$ [$p-value$]	$Q^2(20)$ [$p-value$]	$Q^2(5)$ [$p-value$]	$Q^2(10)$ [$p-value$]	$Q^2(15)$ [$p-value$]	$Q^2(20)$ [$p-value$]
AT	13.855 [0.017]	16.477 [0.087]	18.557 [0.234]	31.062 [0.060]	4.515 [0.478]	5.973 [0.817]	8.062 [0.921]	16.636 [0.676]
BE	7.403 [0.192]	10.625 [0.388]	15.566 [0.411]	19.023 [0.520]	8.182 [0.146]	20.403 [0.126]	27.387 [0.325]	29.923 [0.071]
DE	0.989 [0.320]	1.252 [0.535]	1.294 [0.730]	1.587 [0.811]	5.283 [0.382]	8.085 [0.621]	12.531 [0.638]	13.644 [0.848]
ES	0.998 [0.963]	3.864 [0.953]	7.634 [0.938]	18.070 [0.583]	0.653 [0.985]	1.172 [1.000]	3.038 [1.000]	3.599 [1.000]
FI	30.752 [0.078]	30.861 [0.099]	30.931 [0.124]	30.959 [0.155]	3.799 [0.579]	6.866 [0.738]	7.892 [0.928]	10.228 [0.964]
FR	7.953 [0.159]	11.395 [0.328]	14.986 [0.452]	18.526 [0.553]	1.189 [0.946]	3.002 [0.981]	3.691 [0.999]	5.589 [0.999]
GR	10.437 [0.064]	17.358 [0.067]	20.369 [0.158]	21.279 [0.322]	3.984 [0.552]	6.444 [0.777]	12.522 [0.639]	13.763 [0.842]
IE	7.183 [0.207]	8.638 [0.567]	21.041 [0.136]	22.510 [0.314]	11.082 [0.05]	14.196 [0.164]	15.023 [0.450]	16.998 [0.653]
IT	1.889 [0.864]	2.803 [0.986]	6.515 [0.970]	30.275 [0.07]	4.606 [0.466]	7.765 [0.652]	9.440 [0.853]	10.950 [0.948]
NL	1.654 [0.895]	2.327 [0.993]	7.356 [0.947]	7.660 [0.994]	12.917 [0.024]	15.148 [0.127]	26.006 [0.053]	28.020 [0.109]
PT	1.893 [0.864]	2.682 [0.988]	3.594 [0.999]	5.401 [0.999]	3.594 [0.609]	6.310 [0.789]	7.878 [0.929]	10.824 [0.951]
	$ARCH(5)$	$ARCH(10)$	$ARCH(15)$	$ARCH(20)$	$ARCH(5)$	$ARCH(10)$	$ARCH(15)$	$ARCH(20)$
AT	0.0126	0.0787	0.2115	0.0617	0.5079	0.8607	0.9494	0.7364
BE	0.1838	0.3755	0.4085	0.5324	0.1467	0.1612	0.1399	0.1093
DE	0.3204	0.5411	0.7302	0.8133	0.3891	0.6625	0.6594	0.8762
ES	0.9628	0.9529	0.9246	0.6076	0.9851	0.9996	0.9995	1.0000
FI	0.0845	0.1067	0.1337	0.1565	0.5937	0.7445	0.9327	0.9709
FR	0.1463	0.3143	0.4399	0.5621	0.9464	0.9806	0.9984	0.9992
GR	0.0648	0.0893	0.2375	0.4178	0.5484	0.7889	0.5959	0.8303
IE	0.1803	0.5132	0.1208	0.3087	0.0534	0.2120	0.5096	0.7038
IT	0.8632	0.9861	0.9677	0.0679	0.4724	0.7074	0.8755	0.9644
NL	0.8981	0.9931	0.9504	0.9950	0.0214	0.1148	0.0334	0.1081
PT	0.8684	0.9887	0.9988	0.9991	0.6051	0.7667	0.9183	0.9466

A.4 Galeano and Weid (2014) test for structural changes in correlation.

Table 9: This table shows test statistic Q_T as defined in Eq.(1) for the fundamentals-filtered yields between the GIIPS (GR, IE, IT, PT, ES) bonds and the European bonds (Panel A) or CDS (Panel B) markets. The null hypothesis of constant correlations is rejected if the test statistic exceeds the critical value at the 5% significance level. * denotes statistically significant change points at 5% level. Critical values are: 10%: 1.22, 5%: 1.36, and 1%: 1.63. 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
Panel A : Bond Markets										
AT	1.60*	13/10/2008	1.77*	29/1/2009	2.93*	20/7/2011	3.17*	9/10/2009	2.61*	8/12/2009
	1.84*	30/11/2009	2.16*	25/3/2010			1.50*	22/3/2010	1.76*	11/6/2012
BE	1.99*	6/8/2009	1.84*	3/3/2010			2.05*	12/4/2011		
DE	2.65*	16/10/2008	3.01*	25/2/2009	2.60*	22/10/2008	2.00*	26/9/2008	1.66*	29/12/2008
	1.74*	26/11/2009	3.32*	2/12/2010	2.72*	27/4/2010	1.43*	26/11/2009	2.35*	3/8/2011
					1.62*	7/7/2011				
ES	2.16*	23/12/2009	1.74*	1/12/2010	4.00*	20/7/2011	2.59*	24/3/2011		
					2.15*	30/6/2015	2.37*	3/7/2015		
FI	3.10*	26/10/2009	2.34*	2/12/2010	2.54*	2/7/2008	1.81*	22/4/2011	2.15*	3/12/2009
			3.30*	5/3/2014	2.60*	19/5/2010			2.66*	20/7/2011
FR	1.54*	16/11/2009	3.13*	1/4/2010	1.91*	18/7/2011	3.43*	29/1/2010	2.59*	14/7/2009
			1.75*	9/12/2010			2.24*	20/4/2011	2.37*	18/7/2011
GR			1.83*	9/12/2009	1.82*	15/10/2008	1.72*	17/2/2009	2.16*	23/12/2009
					1.41*	22/11/2009				
IE	1.83*	9/12/2009					3.19*	30/11/2010	1.74*	1/12/2010
IT	1.82*	15/10/2008					2.27*	10/2/2010	4.00*	20/7/2011
	1.41*	22/11/2009					3.11*	20/5/20111	2.15*	30/6/2015
NL	1.92*	14/10/2009	3.44*	23/1/2009	1.83*	24/6/2011	3.61*	30/9/2009	1.79*	20/6/2011
							2.14*	1/4/2011		
PT	1.72*	17/12/2009	3.19*	30/11/2010	2.27*	10/2/2010			2.59*	24/3/2011
					3.11*	20/5/2011			2.37*	3/7/2015
Panel B : CDS Markets										
AT	1.75*	2/11/2009	1.79*	15/11/2010	2.86*	18/7/2011	1.66*	28/3/2011	3.50*	1/8/2011
BE	1.62*	6/11/2009	2.42*	25/11/2010	2.38*	22/7/2011	1.64*	1/4/2011	2.27*	3/8/2011
	1.91*	21/4/2010								
DE	2.60*	5/11/2009	2.32*	23/11/2010	3.07*	20/7/2011	2.77*	18/4/2011	3.87*	20/7/2011
	1.55*	22/04/2010								
ES	2.60*	9/11/2009	1.74*	2/12/2010	2.63*	20/7/2011	1.69*	11/4/2011		
	1.70*	22/4/2010	2.14*	25/07/2011						
FI	2.43*	13/11/2009			3.52*	2/8/2011	2.60*	4/5/2011	3.37*	3/8/2011
FR	1.74*	5/11/2009	1.75*	12/11/2010	3.01*	20/7/2011	1.91*	18/4/2011	2.43*	18/7/2011
	2.14*	21/04/2010								
GR			2.30*	2/1/2008	3.10*	6/11/2009	1.94*	17/11/2009	2.60*	9/11/2009
			2.05*	12/11/2009	2.45*	20/4/2010	2.77*	22/4/2010	1.70*	22/4/2010
IE	2.30*	2/1/2008			2.14*	3/12/2010	2.70*	30/11/2010	1.74*	2/12/2010
	2.05*	12/11/2009			2.76*	15/7/2011	1.67*	29/3/2011	2.14*	25/7/2011
IT	3.10*	6/11/2009	2.14*	3/12/2010			1.75*	12/4/2011	2.63*	20/7/2011
	2.45*	20/4/2010	2.76*	15/7/2011						
NL	3.19*	5/11/2009	1.57*	10/12/2010	2.92*	2/8/2011	2.99*	4/5/2011	2.10*	18/7/2011
PT	1.94*	17/11/2009	2.70*	30/11/2010	1.75*	12/4/2011	2.83*	5/5/2011	1.69*	11/4/2011
	2.77*	22/4/2010	1.67*	29/3/2011						

Table 10: This table shows test statistic Q_T as defined in Eq.(1) for the fundamentals-filtered yields between the GIIPS (GR, IE, IT, PT, ES) CDS and the European bond (Panel A) or CDS (Panel B) markets. The null hypothesis of constant correlations is rejected if the test statistic exceeds the critical value at the 5% significance level. * denotes statistically significant change points at 5% level. Critical values are: 10%: 1.22, 5%: 1.36, and 1%: 1.63. 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
Panel A : Bond Markets										
AT	1.35*	2/11/2009	1.65*	13/10/2011	1.77*	31/10/2011			2.33*	25/10/2011
	2.29*	4/11/2011							2.77*	21/11/2011
BE	2.92*	3/11/2011	2.23*	14/11/2011			1.91*	30/11/2010		
DE	1.46*	22/4/2010	2.07*	12/10/2011	1.86*	26/9/2008	1.42*	26/9/2008	2.01*	26/9/2008
					1.40*	25/10/2011	1.90*	15/5/2013	2.83*	26/2/2010
ES	2.01*	8/11/2011	2.56*	15/11/2011	2.58*	9/11/2011	2.82*	6/4/2010		
			2.13*	7/8/2015	1.85*	22/6/2015	1.92*	16/11/2011		
FI	2.62*	13/11/2009	2.56*	9/1/2008	1.98*	23/4/2010	1.68*	10/9/2008	1.81*	23/4/2010
	2.73*	15/11/2011	2.12*	23/4/2010	2.76*	21/10/2013	2.68*	15/5/2013	1.66*	15/5/2013
FR	2.33*	21/4/2010	2.73*	17/7/2008	2.73*	13/11/2009	2.40*	26/9/2008	2.74*	6/4/2010
			2.32*	10/10/2011	2.32*	18/11/2011	2.86*	1/8/2011	2.22*	8/11/2011
GR			1.59*	9/11/2011	3.53*	10/11/2011	1.71*	9/11/2011	2.01*	8/11/2011
IE	1.59*	9/11/2011			1.90*	12/10/2010	1.57*	18/11/2011	2.56*	15/11/2011
					1.56*	17/11/2011			2.13*	7/8/2015
IT	3.53*	10/11/2011	1.90*	12/10/2010			2.60*	14/11/2011	2.58*	9/11/2011
			1.56*	17/11/2011					1.85*	22/6/2015
NL	2.22*	5/11/2009	2.11*	1/7/2008	2.95*	23/1/2009	1.46*	10/9/2008	1.66*	26/9/2008
	1.40*	7/11/2011	3.28*	3/8/2010	2.20*	26/3/2013			1.64*	26/3/2013
PT	1.71*	9/11/2011	1.57*	18/11/2011	2.60*	14/11/2011			2.82*	6/4/2010
									1.92*	16/11/2011
Panel B : CDS Markets										
AT			2.03*	4/1/2013	2.05*	22/11/2011	1.56*	15/9/2008	2.30*	15/9/2008
							1.45*	1/2/2012	2.36*	3/1/2012
BE			1.39*	28/1/2012	2.58*	9/1/2012	1.61*	13/1/2012	2.03*	9/1/2012
DE			2.16*	13/1/2012	3.84*	16/1/2012	2.14*	15/9/2008	2.00*	18/1/2012
							3.09*	13/1/2012		
ES	1.44*	18/11/2011			2.11*	21/11/2011	1.65*	31/10/2011		
FI	1.78*	1/10/2008	1.80*	19/1/2012	1.50*	1/12/2011	2.28*	2/1/2012	1.94*	14/11/2011
FR					4.42*	9/1/2012	3.82*	12/11/2009	2.10*	5/1/2012
					4.31*	1/2/2013	2.06*	13/1/2012	2.67*	1/2/2013
GR			1.45*	11/11/2011	1.90*	14/9/2010			1.44*	18/11/2011
IE	1.45*	11/11/2011			2.98*	15/11/2011	1.92*	5/1/2012		
					1.82*	31/8/2015				
IT	1.90*	14/9/2010	2.98*	15/11/2011			1.68*	14/11/2012	2.11*	21/11/2011
			1.82*	31/8/2015			2.08*	3/9/2012		
NL	2.98*	12/9/2008	1.58*	7/12/2012	2.18*	20/1/2012	2.19*	5/9/2008	1.70*	15/9/2008
	1.82*	13/9/2010					2.67*	13/1/2012	2.64*	18/1/2012
PT			1.92*	5/1/2012	1.68*	14/11/2012			1.65*	31/10/2011
					2.08*	3/9/2012				

A.5 Bootstrap test for contagion using GARCH filtered series

Table 11: 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 Galeano and Weid (2014) based on the standardized yields–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 1000 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
Panel A : Bond Markets										
AT	0.999	14/10/2008	**0.002**	**10/11/2009**	0.997	11/5/2009	0.999	22/10/2008	0.661	11/5/2009
	0.999	30/11/2009	**0.000**	**27/4/2010**	**0.000**	**2/4/2013**	0.987	30/3/2011		
BE	0.999	1/9/2009	0.912	11/11/2009	0.998	14/8/2009	0.999	5/4/2010	0.987	1/2/2010
			0.000	**15/11/2011**	0.983	20/8/2012	0.971	19/4/2011		
DE	0.999	22/10/2008	0.996	16/11/2009	0.979	26/12/2008	0.992	26/12/2008	0.990	29/12/2008
	0.999	13/11/2009	**0.000**	**26/4/2010**	0.980	7/4/2010	0.999	6/4/2010	0.997	11/7/2011
			0.000	**28/4/2015**	0.999	13/6/2011	0.999	31/5/2011		
ES	0.996	16/12/2009	**0.000**	**16/11/2010**	0.109	22/10/2008	0.987	26/12/2008		
			0.000	**22/6/2011**	**0.000**	**16/6/2011**	0.891	6/6/2011		
FI	0.999	13/11/2009	0.999	25/11/2009	0.998	19/1/2010	0.926	21/4/2008	0.982	19/1/2010
			0.000	**19/4/2010**	0.989	19/5/2011	0.999	1/2/2010		
FR	0.999	13/11/2009	**0.000**	**7/4/2010**	**0.000**	**18/4/2011**	0.999	7/1/2010	0.324	19/1/2010
			0.981	2/4/2013	0.952	15/5/2013	0.615	27/5/2011	**0.000**	**15/6/2011**
GR			**0.000**	**30/11/2009**	0.710	23/10/2008	0.352	23/12/2009	**0.000**	**16/12/2009**
					0.000	**29/10/2009**				
IE	0.912	30/11/2009			0.813	26/4/2010	0.999	10/6/2011	**0.000**	**16/11/2010**
					0.000	**20/6/2011**			**0.000**	**22/6/2011**
IT	0.990	23/10/2008	**0.000**	**26/4/2010**			0.999	26/12/2008	0.109	22/10/2008
	0.999	29/10/2009	**0.000**	**20/6/2011**			0.997	30/5/2011	**0.000**	**16/6/2011**
NL	0.991	5/10/2009	0.995	19/9/2008	0.981	17/6/2009	0.999	30/9/2009	0.767	5/10/2009
			0.000	**31/3/2010**	0.083	19/5/2011	0.189	31/5/2011	**0.000**	**1/4/2013**
PT	0.998	23/12/2009	**0.000**	**10/6/2011**	0.997	26/12/2008			0.987	26/12/2008
					0.000	**30/5/2011**			**0.000**	**6/6/2011**
Panel B : CDS Markets										
AT	**0.000**	**17/11/2009**	**0.014**	**30/9/2008**	**0.000**	**6/1/2009**	**0.001**	**8/12/2008**	**0.000**	**20/1/2009**
					0.990	17/10/2013			0.921	14/10/2013
BE	**0.015**	**6/11/2009**	0.274	18/12/2008	**0.000**	**12/1/2009**	**0.000**	**12/1/2009**	**0.000**	**2/2/2009**
	0.000	**22/4/2010**	0.999	24/3/2015	**0.000**	**6/5/2011**			0.917	31/8/2015
DE	**0.000**	**17/11/2009**	**0.000**	**22/12/2008**	**0.000**	**12/1/2009**	**0.000**	**29/1/2009**	**0.000**	**26/1/2009**
					0.999	21/1/2015			0.878	10/11/2014
ES	**0.000**	**6/11/2009**	**0.001**	**20/10/2008**	**0.000**	**13/1/2009**	**0.000**	**19/1/2009**		
	0.000	**21/4/2010**			**0.000**	**13/6/2011**	**0.000**	**6/6/2011**		
FI	**0.000**	**16/11/2009**	**0.026**	**18/12/2008**	**0.000**	**5/11/2010**	**0.000**	**11/1/2009**	**0.000**	**5/1/2009**
					0.998	30/3/2015			0.834	22/11/2010
FR	**0.036**	**5/11/2009**	**0.042**	**12/11/2008**	**0.000**	**19/1/2009**	**0.000**	**12/1/2009**	**0.000**	**20/1/2009**
					0.999	28/4/2010			0.998	7/4/2010
GR			**0.012**	**12/11/2009**	**0.003**	**6/11/2009**	**0.045**	**17/11/2009**	**0.032**	**6/11/2009**
									0.049	**21/4/2010**
IE	**0.047**	**12/11/2009**			**0.001**	**19/1/2009**	**0.028**	**12/10/2008**	**0.000**	**20/10/2008**
					0.858	7/5/2012	0.998	5/11/2011		
IT	**0.000**	**6/11/2009**	**0.003**	**19/1/2009**			**0.000**	**18/3/2009**	**0.000**	**13/1/2009**
			0.416	7/5/2012			0.899	31/8/2012	**0.000**	**13/6/2011**
NL	**0.019**	**16/11/2009**	**0.055**	**12/11/2009**	**0.000**	**12/1/2009**	**0.000**	**17/12/2008**	**0.000**	**29/12/2008**
			0.905	16/9/2014	0.556	7/8/2014			0.999	7/4/2010
PT	**0.000**	**17/11/2009**	**0.000**	**12/10/2008**	**0.000**	**19/1/2009**			**0.000**	**19/1/2009**
			0.000	**5/11/2010**	**0.002**	**18/6/2012**			**0.000**	**16/6/2011**

Table 12: 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 Galeano and Weid (2014) based on the standardized yields–spreads. 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 1000 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
Panel A : Bond Markets										
AT	0.952	8/11/2011			0.955	23/4/2010			0.344	4/10/2011
BE	**0.000**	**7/1/2011**			**0.000**	**4/1/2010**	**0.000**	**12/11/2010**	0.365	15/10/2012
					0.938	15/10/2012	0.593	20/4/2015		
DE					**0.062**	**7/6/2010**	**0.007**	**15/5/2010**	**0.000**	**30/9/2010**
					0.612	16/4/2013	0.640	25/5/2015	0.199	16/4/2013
ES	0.212	9/11/2011	**0.001**	**20/4/2010**	**0.000**	**12/3/2009**	0.512	10/11/2011		
			0.943	14/11/2011	**0.095**	**10/11/2011**	**0.000**	**18/6/2012**		
FI			0.907	23/4/2010	**0.040**	**7/6/2010**	**0.004**	**25/10/2010**	**0.010**	**15/9/2010**
			0.390	19/4/2013	**0.000**	**15/5/2013**	0.459	25/5/2015	**0.000**	**19/4/2013**
FR			0.763	6/4/2010	**0.011**	**6/7/2010**	**0.006**	**19/11/2010**	**0.004**	**6/7/2010**
			0.349	8/11/2011	0.772	8/11/2011	0.452	21/4/2015	0.459	8/11/2011
GR			0.394	9/11/2011			0.610	9/11/2011	0.112	9/11/2011
IE	0.334	9/11/2011			0.261	22/1/2010	**0.000**	**10/11/2011**	0.211	20/4/2010
									0.254	14/11/2011
IT	**0.072**	**7/11/2011**	**0.033**	**22/1/2010**			**0.000**	**18/3/2009**	**0.000**	**12/3/2009**
									0.047	**10/11/2011**
NL			0.758	3/5/2010	**0.000**	**22/7/2010**	**0.003**	**17/6/2010**	**0.007**	**3/5/2010**
			0.140	27/3/2013			0.206	19/4/2013	0.655	1/4/2013
PT	0.174	9/11/2011	0.141	10/11/2011	**0.000**	**18/3/2009**			0.112	10/11/2011
									0.001	**18/6/2012**
Panel B : CDS Markets										
AT			0.933	4/1/2013	**0.000**	**11/9/2008**	0.998	28/8/2008	**0.000**	**11/9/2008**
									0.999	4/1/2012
BE					0.652	27/11/2009	0.995	13/1/2012	**0.003**	**6/1/2009**
					0.999	16/11/2012			0.999	16/11/2012
DE			0.796	29/12/2011	0.163	6/1/2009	0.994	30/1/2012	0.999	7/12/2011
					0.990	13/12/2012				
ES			0.595	29/10/2010	0.195	12/11/2009				
					0.000	**14/11/2011**				
FI			0.998	11/2/2010	0.520	12/11/2009	0.997	15/8/2011	**0.048**	**12/11/2009**
					0.999	22/10/2012			0.998	14/11/2011
FR			0.991	21/10/2009	0.304	12/11/2009	0.316	12/11/2009	0.137	12/11/2009
							0.981	13/1/2012		
GR					0.880	14/9/2010				
IE							0.544	29/8/2010	**0.030**	**29/10/2010**
IT	0.928	14/9/2010							0.997	12/11/2009
									0.255	14/11/2011
NL	**0.000**	**12/9/2008**	0.808	4/1/2010	0.183	10/9/2008	0.803	18/9/2009	**0.026**	**10/9/2008**
	0.962	13/9/2010			0.996	7/12/2012			0.999	11/12/2012
PT			0.872	29/8/2010						