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Measuring bilateral spillover and testing contagion on sovereign bond markets in Europe

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ABSTRACT

The global financial crisis rapidly spread across borders and financial markets, and also distressed EU bond markets. The crisis did not hit all markets in the same way. We measure the strength and direction of linkages between 16 EU sovereign bond markets using a factor-augmented version of the VAR model in Diebold and Yilmaz (2009). We then provide a novel test for contagion by applying the multivariate structural break test of Qu and Perron (2007) on this FAVAR detecting significant sudden changes in shock transmission. Results indicate substantial spillover, especially between EMU countries, with Belgium, Italy and Spain being key markets during the financial crisis. Contagion has been a rather rare phenomenon limited to a few well defined moments of uncertainty on financial assistance packages for Greece, Ireland and Portugal. Most of the frequent surges in market co-movement are driven by larger shocks rather than by contagion.

1. Introduction

Losses on subprime loans in US banks have had global consequences, as uncovered debt positions created a snowball debt effect that brought down major financial institutions in both the US and Europe. The ensuing financial crisis called for policy intervention, not just by central banks, but also out of the deep pockets of the tax payer. Massive public aid in support of the financial sector, together with falling tax revenues and spending on recovery plans to withstand the economic fall-out of the financial collapse, unleashed a sovereign debt crisis.

Turbulence on European bond markets is just the latest chapter in this string of events. Rising sovereign spreads set off a sequence of fiscal bailouts, further trouble in the banking system, the downgrading of all EMU countries but Germany, and *de facto* IMF interventions in several EU countries. These events demonstrate the strong intertwining of European financial markets. Empirical studies confirm that sovereign bond yield spreads in EMU countries are driven by international financial market conditions, and dominate idiosyncratic risk factors such as default, liquidity and exchange rate premia.

Attributing spread movements to international factors underlines the importance of financial integration, but gives an unsatisfactory answer as to what causes those market developments in the first place. Even international developments must eventually be driven by events in some domestic market that then transmits to all other markets, and feeds back to the source market. The flaw of most studies is to proxy the external risk factors with an aggregate measure that is supposedly exogenous to domestic events, and affects all markets in a similar way. However, linkages are not equally strong between all markets simultaneously (Kaminsky and Reinhart, 2000). Those studies therefore have little

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to say on the strength, direction and time-variation of bilateral linkages between bond markets.[1]

Of particular concern is that those linkages could be driven by contagion. A surge in global risk aversion and risk of contagion were deemed to be significant factors behind the increase in European sovereign spreads (Codogno et al., 2003; Bernoth et al., 2012). Rescue packages for the banking sector and the economic fallout of the crisis cast in doubt the sustainability of public debt positions in several EU countries (Attinasi et al., 2010). Investors seem to have been particularly sensitive to such bad fundamentals and dropped bonds quickly at times of increased turbulence on financial markets (Favero and Missale, 2012). As a result, default and liquidity risk started to rise, also in bond markets starting from much sounder fiscal positions.

In this paper, we measure the strength and direction of all bilateral linkages for 16 EU sovereign bond markets, and propose a novel test for the presence of contagion. First, the spillover measure is based on the forecast error variance decomposition of a VAR model, as in Diebold and Yilmaz (2009, 2012). We augment the VAR with a common factor to control for EU-wide developments, and this introduces an additional feedback channel from common events to bond markets. The decomposition then shows the relative importance of domestic and foreign sources of sovereign bond spread dynamics, and provides a summary measure of the bilateral spillover between all markets. Second, the contagion test is based on the Qu–Perron (2007) (QP) break test for detecting sudden changes in the coefficients of the VAR model, conditional on possible changes in the size and volatility of the shocks hitting the markets. We call contagion a significant change in the transmission mechanism that amplifies the effects of a market subject to a shock to another one. Applying the QP test to the VAR gives an indication of the presence of contagion across EU bond markets. More interestingly, we can track the source of this contagion by considering a QP test on a subset of VAR coefficients. In this way, we may claim if contagion spreads from a particular source.

Results indicate that whereas the overall spillover between EU bond markets has always been substantial, it increased to a permanently higher level since the start of the financial crisis. Linkages between EMU countries are substantially stronger, reflecting their structural integration. Some EMU markets such as Italy, Spain or Belgium are major transmitters of shocks to EMU as a whole. Spillover is also important for countries outside the eurozone but considerably less so: Central European countries (Czech Republic, Hungary and Poland) seem to affect one another, Denmark, Sweden and the UK are rather insulated from other bond markets. On the other hand, contagion has been a rather rare phenomenon limited to periods of uncertainty about financial assistance to Greece, Ireland and Portugal. These events involved important economic and policy events at EU level, but their origins are to be tracked to idiosyncratic domestic shocks, like fiscal trouble. Therefore, the frequent surges in co-movement across bond markets is mainly driven by larger shocks rather than by increased in intensity of shock transmission.

The paper is structured as follows. In Section 2, we review our empirical approach to measuring sovereign bond spillover based on the VAR method of Diebold and Yilmaz (2009, 2012) and how to identify contagion in this model with the QP test. The main empirical results on spillover and contagion are discussed in Section 3. The final section summarises the main results and discusses some policy implications.

2. Methodology

2.1. Measuring spillover with a FAVAR

Diebold and Yilmaz (2009, 2012) propose a measure of spillover based on the forecast variance decomposition of a VAR model including prices of n different assets (x_t). A covariance stationary variable VAR(p):

$$x_t = \sum_{n=1}^{p} \Phi_n x_{t-n} + \varepsilon_t \tag{1}$$

with $\varepsilon_t \sim (0, \Sigma)$ a vector of independently and identically distributed disturbances can be rewritten in its moving average representation as:

$$x_t = \sum_{n=0}^{\infty} A_n \varepsilon_{t-n} \tag{2}$$

where some regularity conditions on the A_i matrices apply. These moving average coefficients are the key to understanding the dynamics of the VAR. The decomposition of the variance of the forecast error of one of the asset prices i at h steps ahead records how much of the variance is due to shocks in another variable included in the VAR, h periods after the shock. Call θ_{ij}^h this percentage contribution of a shock to one asset price to the time series variation in another asset price h steps ahead. Then $\lambda_{ij}^h = \theta_{ij}^h / \sum_{j=1}^{n} \theta_{ij}^h$ is the percentage contribution of θ_{ij}^h in the effect of error variances in forecasting x_i due to shocks to x_j, over all n asset prices included in the VAR.

These λ_{ij}^h are an index number between 0 and 100 that reflects the contribution of a shock originating in one market and flowing to another, and are the key measure of spillover between any two asset markets. Let us define *own variance shares* λ_{ii}^h as the fractions of the h-steps-ahead error variances in forecasting x_i due to shocks to x_i, for i = 1, 2,. . .,n, and *cross variance shares* λ_{ij}^h as the fractions of the h-steps-ahead error variances in forecasting x_i due to shocks to x_j, for i, j = 1, 2,. . ., n, such that $i \neq j$.

Diebold and Yilmaz (2009) suggest using the cross variance shares to measure the spillover from one series x_i to another x_j, i.e., the percentage contribution of a change in daily quoted asset prices on the variation in asset prices of each particular market. The matrix Λ contains all bilateral linkages λ_{ij} to and from two different markets:[2]

$$\Lambda = \begin{pmatrix} \lambda_{AA} & \lambda_{AB} & \cdots & \lambda_{AZ} \\ \lambda_{BA} & \lambda_{BB} & \cdots & \lambda_{BZ} \\ \vdots & & \ddots & \vdots \\ \lambda_{ZA} & \cdots & \cdots & \lambda_{ZZ} \end{pmatrix} \tag{3}$$

The first column of Λ, say for market A, contains all λ_{jA} and the column-elements are the contribution from a shock in market A to asset prices on other markets. The entry λ_{AA} is the percentage contribution of a shock in explaining the movement of the market's own asset price. The rows of Λ, for some market B, contain λ_{Bi} and can be read as the spillover market B receives from a shock to other markets.

The index is not a simple measure of co-movement of markets, but measures the importance of an idiosyncratic shock in a market

[1] Only a few recent studies on sovereign bond spreads have started to separate the role of global risk aversion and country-specific risk and measure the degree of spillover in the sovereign bond market using weight matrices. Caceres et al. (2010) calculate a country-specific spillover coefficient based on joint probabilities of distress extracted from CDS credit default swap spreads. Claeys et al. (2012) proxy the linkages between bond markets by economic distance measures to derive a spatial measure of financial integration and show that the spillover curbs around half of changes in domestic bond rates. Favero (2012) measures the effect of a global risk factor, whose weight depends on the distance of fiscal fundamentals vis à vis a benchmark country.

[2] This Λ is like the weight matrix measuring distance in spatial econometrics.

onto other markets, taking into account also its feedback. Prices move contemporaneously on different financial markets, but this spillover is stronger between markets that are more closely connected. As the matrix Λ measures the bilateral interdependence between markets, this method of Diebold and Yilmaz (2009, 2012) improves over partial equilibrium (regression) approaches since it does not suppose a market is affected only by some exogenous conditions. Each market's movement feeds back into the overall co-movement of markets. The decomposition of the VAR provides a general equilibrium effect that measures the strength and direction of bilateral transmission between markets.

Describing all bilateral linkages in Λ is arduous, especially as adding new markets inflates the dimensions of the spillover matrix, so we condense Λ into a few summary statistics. The *total spillover* index TS^h measures the contribution of the spillover of shocks between all the variables included in the VAR to the total forecast error variance. TS^h is basically the sum of the cross variance shares across all variables (at a certain forecast horizon h), expressed as a ratio to the total forecast error variation, i.e.:

$$TS^h = 100 \cdot \sum_{i \neq j}^{n} \lambda_{ij}^h \Big/ \sum_{i,j=1}^{n} \lambda_{ij}^h \tag{4}$$

The direction of the spillover a market i transmits to all other $n-1$ markets is measured by the sum of each column of the matrix Λ, not including the own contribution of each market:

$$DS_{\leftarrow i}^h = 100 \cdot \sum_{j \neq i}^{n} \lambda_{ji}^h \Big/ \sum_{i,j=1}^{n} \lambda_{ji}^h \tag{5}$$

Instead, a market i receives a spillover from all other $n-1$ markets, and this *directional spillover* DS^h can be expressed as follows:

$$DS_{\rightarrow i}^h = 100 \cdot \sum_{j \neq i}^{n} \lambda_{ij}^h \Big/ \sum_{i,j=1}^{n} \lambda_{ij}^h \tag{6}$$

Measure (6) is the sum of the row-elements of the matrix Λ. The directional spillover tells how much of the total spillover comes from, or goes to, a particular source. The *net spillover* from a market i to all other markets j is then the difference between the gross shock received from and sent to all other markets, i.e. $NS^h = DS_{\rightarrow i}^h - DS_{\leftarrow i}^h$. It is also possible to calculate then the *net pairwise spillover*, which shows how much each market i contributes to another market j in net terms. For this, we get:

$$NS_{i \leftrightarrow j}^h = 100 \cdot \left[\lambda_{ij}^h \Big/ \sum_{k=1}^{n} \lambda_{ik}^h - \lambda_{ji}^h \Big/ \sum_{k=1}^{n} \lambda_{jk}^h \right] \tag{7}$$

Co-movement of asset prices additionally reflect similar responses to common shocks. The origin of these shocks may not be tracked to any specific market but reflect endogenous developments common to all; or its origin can be exogenous to the system of all markets. We therefore additionally control for the existence of common factors in the VAR. Following Bernanke et al. (2005), we use a two-step strategy for estimating this factor-augmented VAR (FAVAR). In the first step, we use factor analysis to extract the common factors driving a significant part of the n observable random variables in x_t. The factor model assumes that x_t can be written in function of k unobservable common factors $F_1, \ldots, F_k$:

$$x_t - \mu_t = w_1 F_1 + \cdots + w_k F_k + \varepsilon_t \tag{8}$$

where μ_t is a variable mean, w_1 to w_k are k factor loadings and ε_t is an independently distributed error term with zero mean and finite variance. One can express n observable variables in terms of k unobservable common factors, where each factor determines x_t with a certain weight w. The factor model (8) requires additional moment and covariance restrictions in order to be estimated. In a first step, we impose the common assumption that the k factors are orthogonal, and use the minimum average partial (MAP) method to determine the number of factors, after which we apply the principal factor method to estimate the factor loadings. In a second step, we estimate the FAVAR, which, besides the original n variables x_t contains an additional k factors F_i. From the decomposition of the FAVAR, linkages occur between any two markets, and in addition the common factor(s). As a consequence, there is both a direct effect of market A on market B, and an indirect effect via the common factor(s). If some market A, for example, contributes strongly to common developments across all markets, and the common factors drives other markets in turn, then market A is indirectly linked to other markets.[3] Exogenous factors can simply be included in the VAR with control variables.

Another methodological caveat of analysing linkages between financial markets is the contemporaneous (intraday) correlation of asset prices. The variance decomposition depends on the ordering of the variables in the VAR, and the cholesky identification of the VAR imposes diagonal block restrictions on the contemporaneous feedback effect between markets. Exogeneity assumptions not allowing for simultaneous feedback are not realistic when testing spillover. In line with Diebold and Yilmaz (2012), we adopt the generalised impulse-response framework of Koop et al. (1996) and Pesaran and Shin (1998) that accounts for the correlation of shocks to all markets by using the historically observed distribution of the shocks. With his generalised approach, all measures derived from the FEVD are invariant to ordering. Our identification of the FAVAR gives a causal direction of the spillover from one market to another, and furthermore excludes simultaneity thanks to the use of common factors.

2.2. Measuring contagion with break tests in the FAVAR

The strength and direction of interdependence between markets evolves over time. Co-movements sometimes stay subdued for long periods, or evolve only in a smooth fashion to suddenly switch to abrupt jumps. We examine the evolution of spillover by estimating the FAVAR model over recursive sample windows. The resulting spillover series does not allow disentangling the underlying sources of time-variation. The period-by-period changes can be related to differently sized shocks (the VAR residuals) or to modifications of the transmission mechanism (VAR coefficients). Some of these changes in transmission are the result of fundamentals, which pass either through real channels like trade or financial links, or to sudden spells of contagion across financial markets (Kaminsky and Reinhart, 2000).

Typically, contagion is defined as a sudden shock in a crisis market that spreads to other markets, and whose transmission cannot be explained by a contemporaneous change in economic fundamentals (Pericoli and Sbracia, 2003). This definition covers various theoretical models of contagion. The spread of a crisis can be the consequence of a switch from one equilibrium to another, as sunspots trigger coordination problems between market participants, or changes in economic behaviour due to informational cascades or herding by market participants. Conceptually, contagion can be distinguished from normal interdependence as a co-movement between markets that cannot be tracked back to fundamental linkages (Kaminsky and Reinhart, 2000) or as excessive co-movements across markets in periods of high stress that reflect breaks in the usual transmission mechanism (Forbes and Rigobon, 2002).

Empirical tests of contagion identify sudden changes in co-movement of some statistics of linkages between markets.

[3] Our approach merges also the studies by Metiu (2012) who tests contagion on Value-at-Risk on CDS of EMU bonds, and Wing Fong and Wong (2012) who examine bilateral linkages with a CoVaR model.

Therefore, modelling contagion requires three steps: firstly, an identification of the common driving factors below market prices so as to isolate the shocks at some stress or crisis periods; secondly, the direction of transmission from the crisis market to another market; and finally, a measure of a significant change in this transmission across markets.

The seminal test for contagion by Forbes and Rigobon (2002) is based on a significant increase in the correlation – conditional on changes in market volatility – between two markets on some date defined *a priori* from market knowledge. Other tests first select the shocks from the residual outliers in a VAR defined on the asset prices to isolate the crisis-date, and consequently test the significance of a dummy series associated with these outliers in a structural model explaining asset prices to find evidence of contagion on these dates (Favero and Giavazzi, 2002; Pesaran and Pick, 2007; Bae et al., 2003).

The identification of contagion on EU sovereign bond markets with these tests, especially during the recent turmoil, is particularly fraught with those complications.

First, the procedures require some prior assumption on the existence of a crisis. Most studies take some prior stance on the occurrence of a crisis in a market. Favero and Giavazzi (2002) isolate crisis episodes from an ad-hoc assumption on the threshold size of the residuals.[4] Financial turmoil has however been ongoing for a couple of years so that normal and crisis times cannot easily be discerned.

Second, some episodes may appear as crises and spread across markets because of their importance, yet these events are not contagious *per se*. Crises may reflect changes in the size of shocks rather than in the transmission mechanism (Forbes, 2012). Contagion test do not account for the changes in behaviour of economic shocks.

Third, most studies suppose some market is at the origin of the contagion. This is a problematic assumption as several bond markets have been in a crisis contemporaneously. Favero and Giavazzi (2002) tackle this problem by setting up a VAR model that allows linkages between several markets. Defining *ex ante* what is the non-crisis country potentially being affected by this crisis is not obvious because of the many interlinkages between these economies. Additionally, the direction of transmission is not only bilateral. Common political and economic events at EU level have had a contemporaneous impact across EU countries.

As a result, many of the previously developed tests for contagion suffer from a simultaneity bias between correlated asset prices, and from changes in co-movement unrelated to transmission but due to modification of the size and volatility of the shocks (Caporin et al., 2013).[5] This invalidates their application to the eurozone crisis.

To address these challenges, we apply a multivariate test for multiple structural changes at unknown dates by Qu and Perron (2007) to the generalised FAVAR model. The QP test searches endogenously for possible changes in the VAR coefficients. It does so by first testing the null of coefficient stability in Eq. (1) against an alternative of $l \leqslant m$ breaks with m being the maximum number of breaks allowed. A Wald-type statistic –WD max(m) – is used to determine the number of breaks by testing the null of zero structural breaks against the alternative of at least one break. When this null is rejected, the sequential F-test SEQ($l+1|l$) for $l = 1,2,\ldots, m$ tests the null hypothesis of l breaks against the alternative of $l+1$ breaks. The additional break is sequentially tested on sample segments (of some minimal length) defined by the breakdates from the estimation with l breaks. This test eventually rejects the existence of a further break and so determines endogenously the number of breaks and their location.

The test indicates if a significant change in some market occurs, which is associated with a crisis moment for this market.

The QP test explicitly allows testing separately for structural breaks in the parameters of the VAR coefficients Φ and in the covariance matrix of VAR residuals ε_t.[6] We identify the abrupt jumps in the VAR coefficients and control for possible changes in the size and volatility of the shocks. In addition, the QP test controls for changes in heteroskedasticity, which eliminates a possible other flaw in detecting contagion.

Note that the QP test treats breaks symmetrically. The break could be associated with an intensification or a dampening of the transmission across markets. Contagion should be associated with the former only. We distinguish the QP breakdates with a significant amplification from those that see a significant reduction in the spread across markets. Another – more explicit – way to identify the source market as being in crisis or stress is to examine the level of financial stress in markets subject to contagion vis-à-vis those that were the source of contagion around the breakdates. We use a variant of the financial stress index by Slingenberg and de Haan (2011) to evaluate the degree of financial stress in pairs of countries with an identified change in transmission. As a rule of thumb, the degree of financial stress shall be relatively higher in the source market than in the recipient of the shock.

We so identify purely the changes in transmission. If this change results in a positive and significant spread across markets, we call significant breakdates evidence of contagion from a crisis market to other markets.

We can identify contagion at two different levels. A QP breaktest on the overall VAR model gives an indication of the presence of contagion between all markets. This is not irrelevant for EU bond markets given that common political and economic events may have caused multilateral contagion. But if the source of contagion is a specific market, the drawback of the overall QP test is that it does not trace its origin. Contagion is more easily understood as a bilateral phenomenon¸ spreading from a crisis market to other markets. Therefore, we need to trace back the overall break down to bilateral linkages that are responsible for the identified break in the VAR. The QP test allows testing a subset of the VAR coefficients for structural change. We can so determine (i) which equation is this dummy significant for (i.e. which bond markets were subject to contagion) and (ii) which market this dummy correspond to (i.e. which market emits contagion).

2.3. Specification

We estimate a FAVAR to compute the percentage contribution of a change in daily quoted EU government bond spreads on the variation in the spreads of 16 EU bond markets. The spreads are 10-year sovereign bond yield spreads over the corresponding German bond yield over the period May 2000 to February 2012 (closing price).[7] Bond spreads are to be preferred as a measure over sovereign CDS premia in our study as we prefer a longer sample period to examine interdependence in normal times. Prior to the financial crisis, sovereign CDS markets were often not liquid and for some sovereign issuers in Europe practically non-existent.[8] Moreover,

[4] Favero and Giavazzi (2002) exclude insignificant dummies and lags from the initial VAR model to achieve an over-identified model. Their structural model features very little interdependence in normal times, yet does display a nonlinear transmission of shocks between markets (i.e. contagion).

[5] Dungey et al. (2005) give an overview and discuss the sampling properties of these tests.

[6] The QP test extends the multivariate break test of Bai et al. (1998).

[7] The main source for the data is Thomson Reuters Datastream. For reasons of data availability we did not include Luxembourg or smaller CEE countries, which have quoted bond yields only in recent years.

[8] Since the prohibition of short run speculative swap positions (the naked swaps), the CDS market has practically dried up, which makes a comparison to recent developments increasingly hard.

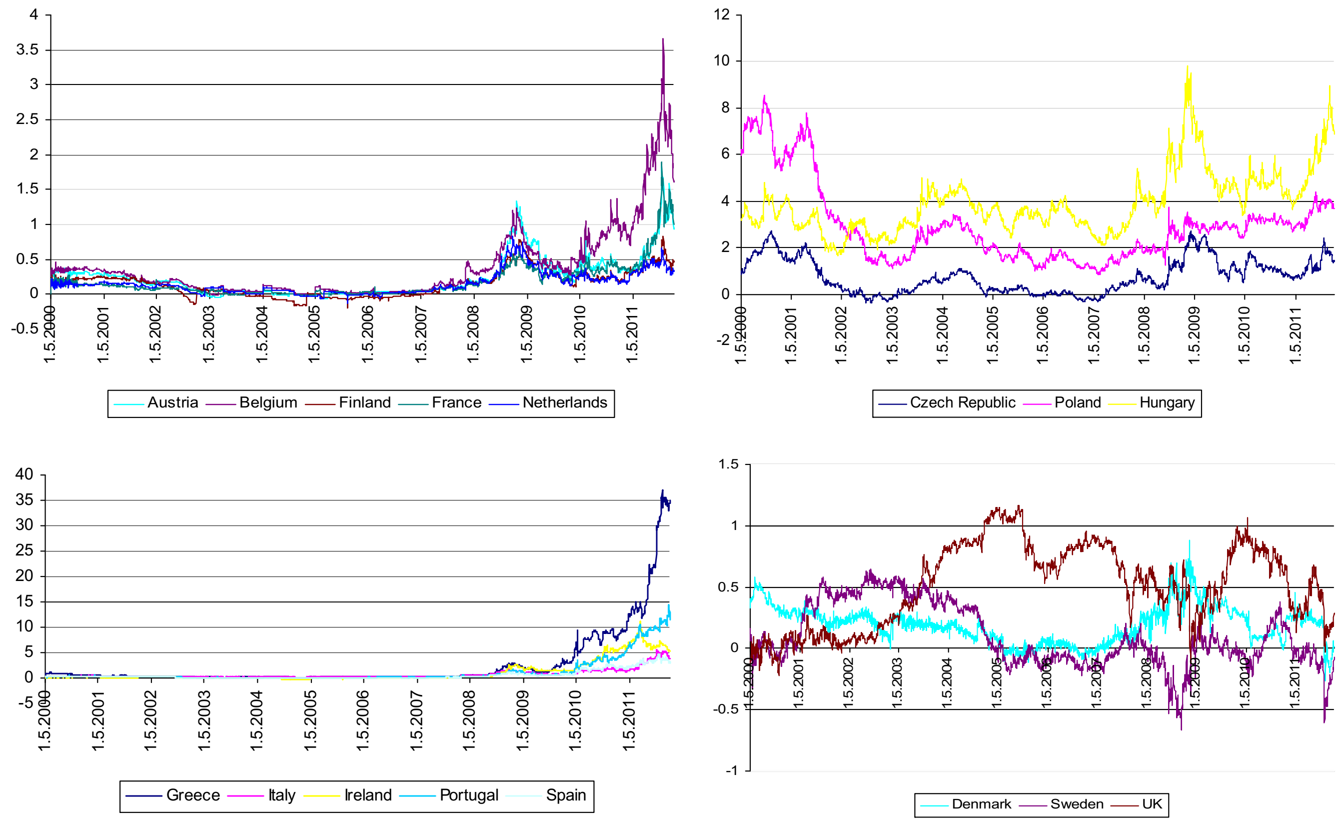

Fig. 1. Bond spreads on the German 10-year bond yield.

although CDS arguably provide a closer measure of sovereign credit risk, we look at interdependence between sovereign bond markets regardless of the source of the underlying risk. Next to credit risk, developments in bond market liquidity, inflation and exchange rates can be an additional factor of divergence between EU bond markets. However, investors effectively bear these risks and interdependence is determined by these factors too.[9]

Fig. 1 shows the spreads for four different groups of countries. First, we have a group of core EMU countries (Austria, Belgium, France, Finland and the Netherlands), where the spreads are moderate but have nonetheless risen a lot since the onset of the financial crisis and again as the sovereign debt crisis started, except in the Netherlands or Finland. A second group of the EMU periphery (or GIIPS) countries (Greece, Ireland, Italy, Portugal and Spain) have seen spreads booming to rates that made refinancing impossible. Third, we have a group of Central European (CE) countries (Czech Republic, Hungary and Poland) where spreads are moderately above EMU levels. Finally, the eurozone 'opt-outs' (Denmark, the UK and Sweden) whose spreads have been much lower and fluctuating.

As noted in previous research (Codogno et al., 2003; Bernoth et al., 2013) significant variability in EU sovereign yield spreads is related to external developments. Therefore, as a first step in estimating the FAVAR model, we apply factor analysis to extract

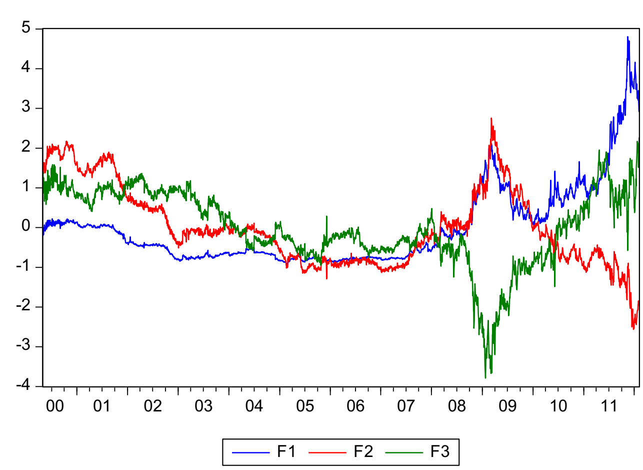

Fig. 2. Time evolution of factors.

common factors from the 16 EU sovereign bond yield spreads. The MAP-method selects three factors as common drivers.[10] Their evolution is very smooth until the onset of the financial crisis in

[9] A potential drawback of using spread above a risk free rate (German bond yield) is that we cannot assess the spillover to and from the reference country. This can be particularly relevant if the reference country enjoys safe haven status when investors fly to less risky and more liquid markets. German bonds have enjoyed this status on several occasions during the financial crisis. Nonetheless, they are a natural proxy of a risk-free rate for the EU sovereigns as perceived by investors. As a robustness check, we performed the FAVAR analysis also using sovereign yield spreads – also including the German one – over US 10 year bonds. The results confirm that the common EU factor is dominated by developments in the German bond market, which makes it a good proxy.

[10] We used alternative methods to determine the number of factors and estimate their loadings and these checks all provide similar results. A particular restriction is that the factor analysis assumes fixed loadings over time. Given the significant changes in European sovereign debt markets, we performed the factor analysis on two subsamples with a breakdate of January 1st 2010. Although the results pointed to some differences between the two periods, the first factor consistently explains at least 65% of the variance and the factor loadings did not vary notably, i.e. the loadings for EMU countries were close to one, the loadings for CEE countries smaller and the loadings for Denmark, Sweden and the UK small or even negative. Evidence that the relative importance of different factors varies over time, albeit not greatly, is also reported in Broto and Perez-Quiros (2011) and ECB (2012).

Table 1
Factor loadings.

	Factor 1	Factor 2	Factor 3	Communality	Uniqueness
Czech Republic	0.64	0.58	−0.08	0.75	0.25
Poland	0.33	0.66	0.37	0.69	0.31
Hungary	0.75	0.06	−0.45	0.78	0.22
Austria	0.94	0.18	−0.15	0.94	0.06
Finland	0.87	0.37	−0.12	0.91	0.09
Netherlands	0.84	0.29	−0.32	0.90	0.10
France	0.96	−0.11	0.00	0.94	0.06
Belgium	0.97	−0.14	0.08	0.97	0.03
Spain	0.92	−0.28	0.17	0.96	0.04
Italy	0.95	−0.28	0.07	0.98	0.02
Greece	0.86	−0.39	0.24	0.96	0.04
Portugal	0.88	−0.37	0.25	0.97	0.03
Ireland	0.85	−0.29	0.15	0.84	0.16
Denmark	0.29	0.79	−0.11	0.72	0.28
Sweden	−0.37	0.21	0.44	0.38	0.62
UK	−0.26	−0.64	−0.47	0.70	0.30
	Variance	Cumulative	Difference	Proportion	Cumulative
Factor 1	9.60	9.60	6.92	0.72	0.72
Factor 2	2.68	12.28	1.58	0.20	0.92
Factor 3	1.10	13.38	-	0.08	1
Total	13.38	35.25		1	

Table 2
Bilateral linkages, Λ matrix, full sample (May 2000–February 2012).

	Czech Republic	Poland	Hungary	Austria	Finland	Netherlands	France	Belgium	Spain	Italy	Greece	Portugal	Ireland	Denmark	Sweden	UK	Factor	From others
Czech Republic	52.52	7.51	6.65	2.51	0.52	0.74	1.65	2.74	3.48	4.01	0.80	0.83	1.94	4.04	0.91	0.03	9.14	47.48
Poland	6.94	61.17	6.38	1.10	0.21	0.22	0.77	1.78	2.44	2.97	1.09	1.12	1.95	5.32	0.79	0.02	5.74	38.83
Hungary	6.86	8.79	54.43	2.35	0.46	0.42	0.63	3.00	2.99	3.60	1.66	1.30	3.10	3.60	0.09	0.06	6.68	45.57
Austria	1.69	1.54	2.56	21.79	3.83	6.49	9.60	11.01	7.44	9.18	2.00	1.50	3.72	0.39	0.09	0.08	17.09	78.21
Finland	1.53	0.96	0.79	8.52	26.30	10.77	8.83	7.96	4.45	5.05	1.38	1.38	3.59	0.87	0.41	0.60	16.62	73.70
Netherlands	1.60	0.84	1.61	7.77	8.39	25.56	8.39	7.68	5.44	5.29	1.59	2.30	4.36	1.35	0.47	0.97	16.39	74.44
France	1.54	1.33	1.54	9.58	3.84	6.54	18.97	11.77	8.16	11.49	2.33	1.36	3.27	0.98	0.28	0.25	16.79	81.03
Belgium	1.67	1.41		7.12	2.56	4.51	8.10	20.94	13.34	13.60	1.89	2.28	5.65	0.22	0.14	0.07	14.74	79.06
Spain	1.36	1.04	1.15	5.24	1.43	3.45	6.40	10.64	27.19	14.85	2.93	3.61	7.79	0.13	0.13	0.27	12.39	72.81
Italy	1.75	1.39	1.39	3.93	1.27	2.62	4.25	12.33	17.65	26.29	3.02	3.68	6.67	0.18	0.06	0.07	13.46	73.71
Greece	1.12	0.79	0.76	2.59	1.56	1.89	4.81	9.29	9.69	7.78	35.52	6.04	9.02	0.01	0.01	0.11	9.02	64.48
Portugal	0.79	0.67	0.98	2.19	0.27	0.82	1.30	8.52	10.00	6.53	5.93	37.73	16.43	0.01	0.15	0.03	7.63	62.27
Ireland	1.07	0.79	1.00	3.23	1.78	2.44	3.69	7.79	9.77	4.99	5.33	10.31	38.32	0.01	0.05	0.03	9.41	61.68
Denmark	3.99	4.13	4.75	1.25	2.20	2.26	2.25	0.56	0.24	0.32	0.30	0.23	0.33	64.17	5.24	0.18	7.60	35.83
Sweden	1.25	1.01	0.56	0.15	0.58	0.84	0.38	0.23	0.46	0.31	0.04	0.13	0.09	4.70	87.21	0.63	1.44	12.79
UK	0.20	0.15	0.14	0.33	0.83	1.89	0.53	0.25	1.97	0.84	0.30	0.92	0.80	0.13	1.14	87.63	1.94	12.37
Factor	3.15	2.27	2.62	8.78	4.42	6.31	8.05	11.53	10.03	11.51	2.86	3.56	6.70	1.18	0.28	0.28	16.46	83.54
To others	36.51	34.60	34.64	66.65	34.14	52.20	69.64	107.09	107.54	102.33	33.43	40.57	75.39	23.11	10.23	3.67	166.07	997.82
To others (+ own)	89.03	95.76	89.07	88.44	60.44	77.76	88.61	128.03	134.73	128.61	68.96	78.30	113.72	87.28	97.43	91.30	182.53	59%
From others	47.48	38.83	45.57	78.21	73.70	74.44	81.03	79.06	72.81	73.71	64.48	62.27	61.68	35.83	12.79	12.37	83.54	
Net spillover	10.97	4.24	10.93	11.56	39.56	22.24	11.39	−28.03	−34.73	−28.61	31.04	21.70	−13.72	12.72	2.57	8.70	−82.53	

2008, but then spikes to diverge later on (Fig. 2). The first factor starts to increase in 2008 as the global financial crisis hit and there was a significant increase of yield spreads, notably in the eurozone. The second spike appears during the latest acute phase of the debt crisis in the autumn of 2011. The principal factor method shows that the first of these factors is able to explain over 70% of the variance of the spreads (Table 1). The factor loadings are close to unity for the eurozone countries, which suggests that this factor mostly identifies developments common to the EMU. Non-eurozone countries have substantially lower loadings on this factor. The financial crisis hit all EU countries, and the second factor also reaches a peak in late 2008 and early 2009, but has steadily declined since. The third factor reaches a minimum in 2008/09 and has been rising since. Both explain much less of the overall variance and although their loadings are somewhat higher for the non-eurozone countries, they do not have an intuitive interpretation.

The basic FAVAR model contains two lags of the domestic bond spreads of 16 EU countries and the prime common factor obtained in the first step.[11] In line with Diebold and Yilmaz (2009) we compute the forecast error variance decomposition at a horizon of 10 days (one and a half weeks), which should be sufficient to capture the horizon at which spillover across markets occurs. We additionally control the VAR for exogenous factors, and include a short-term interest rate (EONIA) to control for the possible effects of monetary

[11] As a robustness check we have included also the other two factors but the results remained practically unchanged as the additional factors track a relatively small share of the overall bond market variance.

policy on the short end of the term structure. Another control variable by which we also capture the role of global bond markets is the Chicago Board Options Exchange Index (VIX). This index is often used to measure risk aversion on global markets. Volatility on markets outside Europe, especially the US, is argued to be a main driver of bond spreads.[12]

3. Empirical results

3.1. Bilateral linkages across bond markets

We first estimate the FAVAR on the full sample, and analyse the bilateral linkages between all 16 EU sovereign bond markets. Fig. 1 already suggested the presence of important interlinkages between sovereign bond markets, but also that these linkages might not be equally strong between all of them. Table 2 reports the own and cross variance shares of a shock to bond spreads onto bond markets and the common factor (each main entry displays the coefficient λ_{ij}). Table 2 further contains a set of summary statistics. We sum the effect of shocks to market *A* on all others (either including the own effect or not) in the two rows following the country effects. The right-hand column sums the effect country *B* receives from all other markets.

Table 2 captures the direction and intensity of linkages between different sovereign bond markets, as well as the spillover between individual bond markets and the common factor. The total spillover on 16 EU bond markets amounts to 59%, meaning that more than half of the variation in sovereign bond spreads can be explained by shocks to bond spreads in other countries. The remaining 41% of all movements are caused by purely domestic factors, i.e. the idiosyncratic dynamics of the domestic spread in the past. This finding is in line with what other studies find: a major part of the bond spreads is determined not by domestic factors but by international bond markets (Codogno et al., 2003; Longstaff et al., 2011; Favero and Missale, 2012). In contrast to these studies, our result is not derived from a partial equilibrium assumption, in which global conditions cause domestic changes, but it fully accounts for the feedback of domestic markets to international markets.

There are quite some differences between the strength of the idiosyncratic effect for different groups of countries. The diagonal entries in Table 2 show that the own variance share (λ_{ii}) is not equally strong for each country. For the eurozone opt-outs the country-specific effect accounts for over two-thirds of the changes in the bond spread. This idiosyncratic effect also dominates the spillover effect for the CE countries as it ranges between one-half and two-thirds. By contrast, the idiosyncratic change amounts to just one-quarter for the eurozone countries (with a slightly higher share for Greece, Portugal and Ireland). The EMU bond markets are more strongly integrated and shocks to spreads mostly affect other markets rather than being idiosyncratic.[13] The total spillover disguises the large variety of pairwise spillover effects. We can see this from the cross variance shares in Table 2. The bilateral linkages between countries are quite distinct between countries inside and outside the eurozone. For the three opt-out countries, the bilateral linkages both among them and with the other EU countries are weak. Less than 15% of the shocks to bond spreads to these three countries spills over to other markets. The most extreme cases are Sweden and the UK, whose sovereign borrowing costs hardly have any effect on other EU countries at all. Denmark is relatively less insulated from bond markets in the eurozone probably because of its stronger integration and its participation in the ERM II. Similarly, CE countries have limited effects on other bond markets, although their bilateral linkages are stronger than for the opt-outs. About one-third of all the spillover to other markets only occurs between the Czech Republic, Hungary and Poland themselves, which makes them a specific group. The same applies to the spillover these countries receive.

The common factor affects – and is affected by – all bond markets. As we might expect from the high factor loadings of EMU countries in the estimation of the factor model (Table 1), most of its impact flows to eurozone countries. Common EMU bond market developments mainly have their source in Belgian, Italian and Spanish bond markets. Although the common factor transmits strongly to all EMU markets – in particular on Austrian, Finnish, French and Dutch bonds – in no market does the spillover to the common factor dominate the direct effect onto other markets. This implies that the indirect effect is weak and does not add much to the direct spillover effect of domestic shocks.

Among the eurozone countries, we can identify three groups of countries by the strength of their bilateral spillover: (i) the core eurozone (Austria, Finland, France and the Netherlands), where domestic factors are of minor importance and countries affect each other, (ii) Portugal, Ireland and Greece, where domestic dynamics are slightly more important than bilateral (mutual) effects, and (iii) Belgium, Italy and Spain who have strong mutual bilateral effects, but both receive and pass on the effects of shocks to all eurozone markets.

The share of each of these three markets in the overall spillover is more than 50%, both on the sending and receiving end. The results for Italy and Spain are probably not surprising given the size of those bond markets, and the turbulence through which both sovereigns have gone in recent years. Other studies also find that Italy and Spain have been crucial transmitters of shocks on bond markets to other countries. For example, using CDS series, Broto and Perez-Quiros (2011) find that both Italy and Spain are more affected by events on other EMU markets than by domestic events. The Belgian bond market has not received the same attention by analysts as eurozone countries in the periphery. It is therefore not typically considered to create systemic links on sovereign bond markets. Still, a few recent studies conclude that Belgium, together with Spain and Italy, is a systemic bond market in Europe (e.g. Ang and Longstaff, 2011). De Santis (2012) or Metiu (2012) find that events in the EMU periphery mainly spread to other periphery countries, but also to Belgium and France. In the latter two studies, Belgium is the market that is particularly hit by events in other EU countries. This result endorses the use of the FAVAR approach since it accounts for both the transmission and feedback effects, and therefore ranks markets on their relative linkage to other markets.

Numerous explanations can be given for the direction and strength of all bilateral linkages. We do not attempt an exhaustive analysis of those channels, as other studies have discussed this at length, like Bekaert et al. (2011) or Forbes (2012). A few channels are important to explain the strength of spillover we find. First, several studies have argued that the financial crisis has made investors rediscover macroeconomic fundamentals in bond pricing (Bekaert et al., 2011). Investors have received a wake-up call

[12] In this way, we implicitly benchmark the spillover between EU bond markets also on the evolution of global bond markets.

[13] The potency of spillover across European bond markets should not come as a surprise. The gradual process of economic and financial integration, stimulated by several rounds of capital account liberalisation, financial deregulation, and the introduction of the euro has not been limited to capital and stock markets (Lane and Milesi-Ferretti, 2008). Whereas in the past, governments relied on domestic savings held in the national banking system, they can now tap into international capital markets. In the eurozone, integration has made bond portfolios increasingly internationally diversified. Issuance in a common currency has motivated debt managers to compete for investors from other countries willing to diversify their portfolio by increasing the volume of new debt issues. Improved transparency and the elimination of some technical obstacles (such as trading systems and tax differences) has further reduced home bias and promoted integration of bond markets (Hartmann et al., 2009). As a consequence, half of public debt is held by a pool of mostly European creditors (BIS, 2011).

during the crisis. The systemic importance of Belgium in EU bond markets could be due to the high debt to GDP ratio, even though its volume is small relative to the debt issues of Italy or Spain. This similarity would not explain the strong spillover *per se*, unless investors react in an identical way to fundamentals. Second, some studies argue that markets are mispricing sovereign bonds due to perceptions of investors that spillover across EMU countries is strong, both through trade linkages and the financial sector, but the policy reactions have been weak. Several papers have looked at the joint dynamics of sovereign and bank risk (Aizenman et al., 2013; Acharya et al., 2011). In the case of Belgium, it economically belongs to the core EMU countries, and has (had) an internationally developed banking system with strong exposure to the EMU periphery. For example, the Belgian banks Fortis and Dexia were among those with the highest exposure to US subprime loans and Greek public debt respectively (BIS, 2011). Spanish banks are exposed to problems in the domestic financial sector and could transmit this to their foreign branches. Belgian, Italian and Spanish banks moreover mutually hold large portions of public debt (BIS, 2011). Second, some papers have argued that sovereign defaults may occur in sequence as investors recompose their portfolios to reduce risk and cut losses once one sovereign might go in default. Recent papers in the literature of endogenous sovereign risk include Hatchondo and Martinez (2009) and Lizarazo (2013).

Commonalities in the EU sovereign bond market are mostly common developments within the eurozone. By contrast, non-eurozone countries and especially the UK, stand rather apart. The evidence suggests that this separation might be driven by exchange rate differences relative to the euro area. This suggests that, non-eurozone countries have been able to contain mispricing on their sovereign risk (de Grauwe and Ji, 2013; Poghosyan, 2012).

3.2. Time-variation in spillover in EU sovereign bond markets

The full sample estimates summarise the linkages between markets. Arguably, the strength of these linkages evolves over time. Spreads in all EU countries moved closely together since early 2002 until the financial crisis made spreads rise and the the European debt crisis made spreads EMU countries diverge. As in Diebold and Yilmaz (2009), we look at the time-evolution of these linkages and therefore run the VAR model over a 200-day rolling window.

We summarise the time evolution of the total spillover index between all bond markets in Fig. 3. We observe that the linkages between markets have not been limited to the recent periods of financial turmoil. The spillover has been substantial most of the time, as the index never falls below 50%.[14] The most relevant change occurs at the start of the financial crisis. Prior to the crisis, the total sovereign bond spillover oscillates between 55% and 70%, and we observe a few specific spikes, for example after 9/11, the application of the Excessive Deficit Procedures (EDP) to some EU countries (March 2003) and the revision of the Stability and Growth Pact in March 2005. The high overall level of spillover confirms the evidence of other studies on the crucial role of external factors in driving bond spreads (Codogno et al., 2003; Bernoth et al., 2012).

The decline in overall spillover in 2006 indicates a period in which investors on bond markets started to perceive sovereign issuers as distinct. The start of the financial crisis in mid-2007 reversed this development by raising the co-movement of sovereign bond spreads. As of late 2008, the total spillover index shifts permanently to higher level until the end of the sample. We

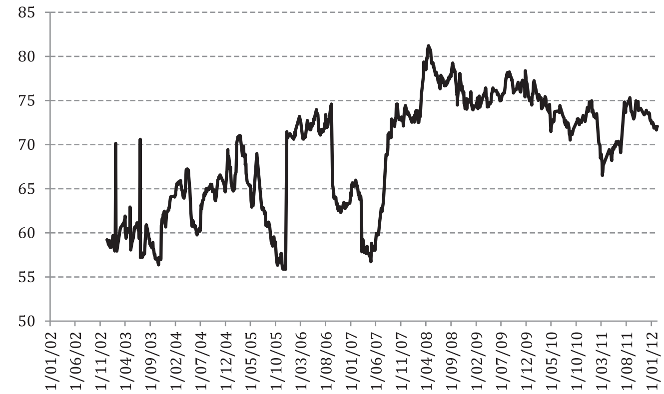

Fig. 3. Total spillover plot, 200-day window, 10-steps-ahead forecast, full sample (May 2000–February 2012).

observe how the spillover peaks at the height of the financial crisis in 2008, when the crisis starts on European bond markets in late 2009 and as the eurozone sovereign debt crisis unfolds during the spring of 2010. It thus seems that interdependence between markets has been at a constantly high level for EU markets. This should not come as a surprise, since financial integration in Europe has accelerated after the introduction of the euro (ECB, 2012). Despite integration, yield spreads on long term bonds of EMU countries are still not completely aligned (Hartmann et al., 2009). Some differences in EU government bond markets remain due to differences in liquidity and the availability of derivatives markets for these assets, and a different response of national markets to global factors (Favero et al., 2010).

Even over a period of strong interlinkages, there have been some major sudden fluctuations. In order to provide a closer view on the evolution of spillover during the financial crisis, Fig. 4 shows a close-up image of Fig. 3 starting in January 2008. Many of these peaks and troughs in the spillover index are associated with major (global or European) financial or political events. We indicate those events on Fig. 4:

A. The collapse and subsequent sale of Bear Stearns to JP Morgan Chase (March–May 2008).
B. The collapse of Lehman Brothers (September 2008).
C. Greek government revealed corrected data on fiscal stance (November 2009).
D. The first Eurogroup meeting on Greece (May 2010).
E. The agreement by EU leaders on the establishment of the European Systemic Risk Board (December 2010) and set up of the European Stability Mechanism (March 2011).
F. The spread of the debt crisis to Spain and Italy (June 2011), and the measures adopted in August and September 2011 by the ECB.

The time-variation in total spillover index sums numerous changes in bilateral linkages across markets, some of which may even go in opposite directions. Space constrains do not allow reporting time-varying indices of bilateral spillover as there are 17^2 pairs for each direction,[15] It seems interesting to look at least at one particular case: Greece was the first EMU country to run into fiscal trouble in November 2009 as its fiscal balances were announced by the then recently installed Papandreou government to be much larger than previously published. This revision was followed by a series of events in Greece and the eurozone, such as fiscal bailouts and trouble in the balance sheets of banks.

[14] We can compare our estimate, which varies between 50% and 80%, with Diebold and Yilmaz (2009), who estimate the spillover between stock markets (1995–2007) at between 40% and 55%. While our total sovereign bond spillover from the whole sample analysis is 59%, their stock market spillover index is 35%.

[15] The time-varying plot for any other pair of countries can be obtained from the authors upon request.

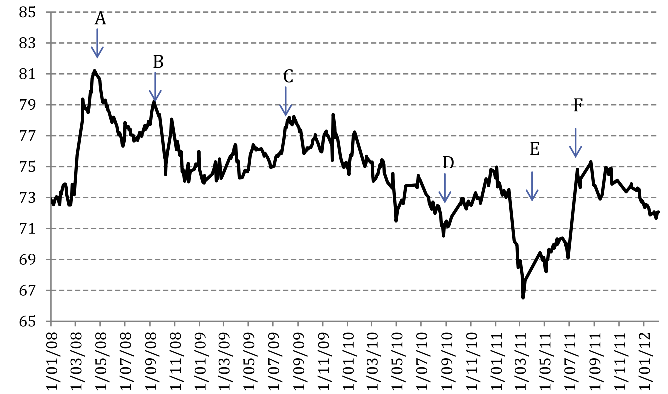

Fig. 4. Total spillover plot, 200-day window, 10-steps-ahead forecast, January 2008–February 2012.

We noted in Table 2 that the overall spillover from Greece onto other bond markets is rather small. This evidence based on the whole sample might seem rather counterintuitive given the political and economic events since 2009 (Mink and De Haan, 2013). In Fig. 5, we present the time-varying estimates of the spillover from Greece to other markets (the columns of Table 2). In order not to clutter the graph with a decomposition of all 16 markets, we have assembled the spillover into 4 groups of countries following the panels in Fig. 1 (but Greece itself is excluded from the EMU periphery group).

Two observations on Fig. 5 stand out. The contribution of changes in sovereign spreads in Greece on other markets fluctuates significantly over time (the variation is much more pronounced than in the overall spillover index) and the transmission is quite different across groups of EU countries. The spillover remains rather stable until the onset of the global financial crisis. While the non-EMU countries are just marginally affected, most of the Greek spillover goes to other EMU periphery and core countries in equal proportions. Whereas some pre-crisis spikes do not have any clear interpretation and can be related to market microstructure effects, some others can be related to ongoing events such as the first EDP against Greece (May 2004) or to doubts about the implementation of the Stability and Growth Pact (March 2005). Although these events cause minor jumps in sovereign spreads, their impact across markets is much larger as markets were at that time much less volatile. The financial crisis immediately magnifies the spillover to other markets early in 2008 as the global financial crisis hits Europe. All other markets suffer this increased spillover in a similar way. Even though doubts about the budgetary situation of Greece had started to rise since late 2009, the spillover fluctuated at higher levels but decreased to pre-crisis levels in early 2010. A comparison to Fig. 4 shows that there was no particular event at EMU level to create this disconnect between Greece and other bond markets. Mink and De Haan (2013) argue that after acknowledging the bad state of Greek public finances in early 2010, investors started to put a higher weight on the domestic fiscal position and discerned the problems of Greece from other EU sovereigns. This arguably reduced the spillover from Greece to the rest of the EU. This disconnect has not continued to hold as a strong reversal in the degree of spillover has taken place in the spring of 2011. This increase in the effect of Greek fiscal problems mainly affects the EMU countries. During the summer of 2011, the impact on other periphery countries, in particular Italy and Spain, rises strongly. These reversals reflect the ongoing discussions at the EU level regarding the bailout of Greece. The rescue package of July 2011 does not seem to have separated the fiscal trouble in Greece from other bond markets as the spillover remains at a high level since.[16] The latter part of the sample displays market behaviour that might be consistent with de Grauwe and Ji (2013). They argue that the surge in spreads is disconnected from the rise in public debt ratios and is a sign of mispricing of sovereign risk, which might be due to uncertainty on a solution to the debt crisis (Aizenman et al., 2013).

Our results show that sovereign bond spillover may flare up as strong underlying interdependence raises vulnerabilities to shocks. We can see this ripple effect even better by isolating the spillover from Greece to Spain and Italy (Fig. 6), which are the largest countries being to date affected by the sovereign debt crisis. Fig. 6 again confirms that the time variation in spillover is very significant. In particular, spillover rose strongly up until the first deal on bailing out Greece. Investors did not continue to discern Greece from the other countries in the EMU periphery throughout 2011, and spillover to Spain and Italy flared up again strongly in June 2011, until the first intervention of the ECB in August 2011. Interestingly but not entirely surprisingly, the spillover towards both countries has an almost identical temporal pattern. Moreover, Fig. 6 also demonstrates that the mood of the market can reverse very quickly in either direction.

3.3. Robustness checks

We have controlled for common bond market developments by including a common factor in a FAVAR. The importance of this factor can be seen from calculating matrix Λ for a VAR including only the 16 bond spread series without the common factor (see Table A.1 in the Appendix). The total spillover is about 5% lower, since the feedback from the common factor to each market is now incorporated into the evolution of the domestic spread. This feedback is particularly stronger for the eurozone countries. The own variable shares (i.e. the diagonal elements of Λ) are therefore larger, as is the spillover from the domestic to other markets. Therefore, omitting this common factor might cause upward bias in the own variance share, as the feedbacks of common EMU events are not taken into account. This again attests to the importance of the commonalities between EMU bond markets (see Fig. 7).

An alternative way to take the common factor into account is to de-factorise the spread series for each country and retain only the idiosyncratic evolution of the spread in the FAVAR. The spillover should just reflect the transmission across bond markets of idiosyncratic shocks, at least if the transmission of the common factor is identical across markets. For markets like Belgium, Italy or Spain that share common developments (Table 2), the own variance share indeed increases, and the spillover to other markets is limited (see Table A.2 in the Appendix). In contrast, the model mostly owes to country-specific shocks the deviation from a common factor in markets that do not have much in common with the others. The consequence is that the spillover from these markets to the others is much stronger now (at 66%). Two contrasting cases are the periphery and core EMU countries. Since the former have been driving rising spreads in the eurozone, the common factor absorbs most of the spillover. Any other country-specific deviation has affected the spread only domestically (for more than 50% in Greece, Ireland or Portugal). By contrast, the spread in the core EMU has not followed the rise of the EMU periphery to the same extent, but it is still correlated with the spreads in other EU countries (Fig. 1). The spillover between the core EMU bond market and the other bond markets is therefore much higher (as is the total spillover in this model). For the same reason, the importance of

[16] Favero and Missale (2012) show similar evidence on the interaction of generalised risk aversion and a worsening domestic fiscal situation.

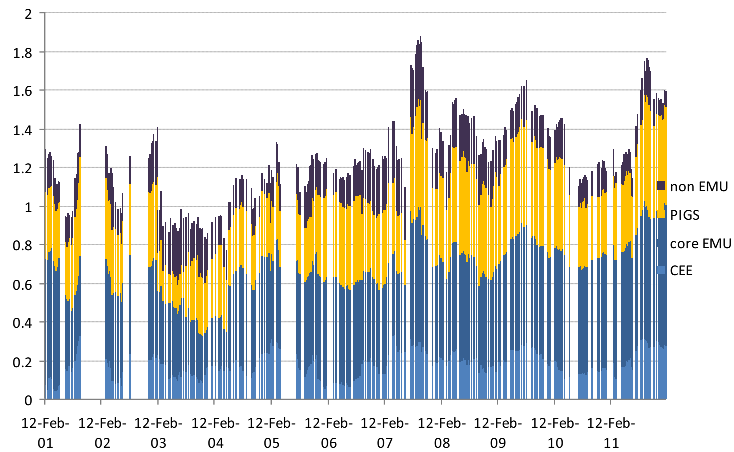

Fig. 5. Decomposition of the effect of Greek bond spreads on other markets.

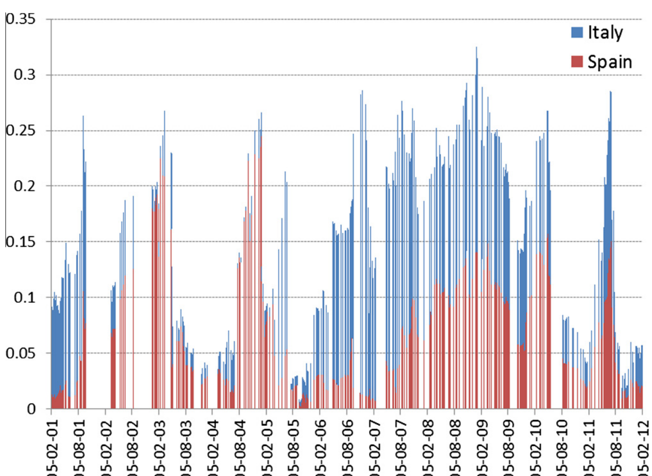

Fig. 6. Spillover from Greece to Italy and Spain.

the opt-out countries (Denmark, Sweden and the UK) on the spillover transmission increases substantially.

In addition to alternative specifications of the FAVAR model, we check the baseline model on some parameters. These confirm that our previous results are robust to (i) changes in the number of lags included in the VAR, (ii) the number of steps ahead when making the forecast, and (iii) the sample window. VAR models with 4 lags (instead of 2), a 20-days-ahead (instead of 10-days-ahead) forecast and a 400-day (instead of 200-day) rolling window all depict a similar evolution of the spillover over time (Fig. 8).

3.4. Testing for contagion on the EU sovereign bond markets

We did not discuss the nature of interdependence between markets so far. The sudden spikes in the spillover index suggest an excess co-movement that is disconnected from fundamentals, and spreads from a bond market in crisis to other markets. This could be contagion as changes in economic fundaments are unlikely to occur at such abrupt fashion. Some of the identified spikes also correspond to periods when market contagion has been often mentioned in the media.

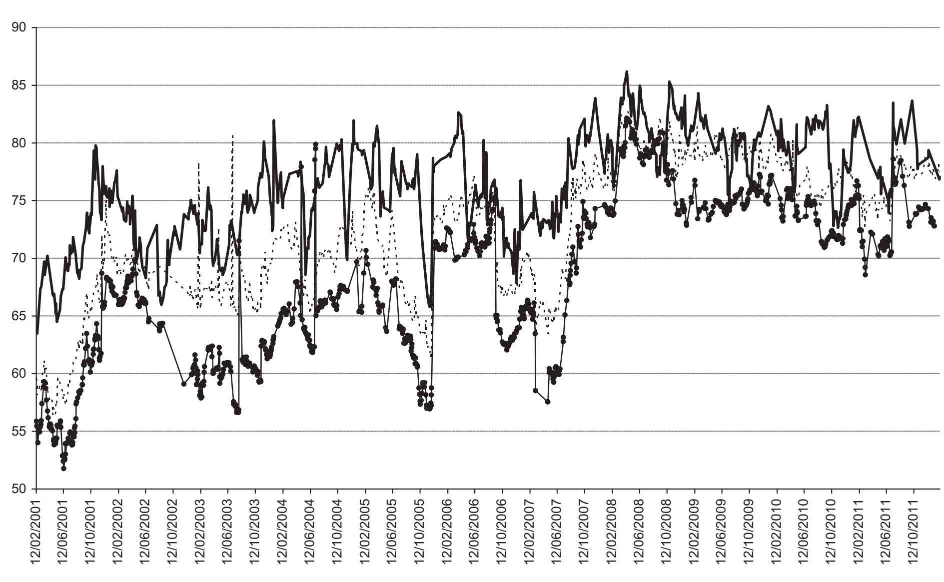

Fig. 7. Robustness checks on the VAR model.

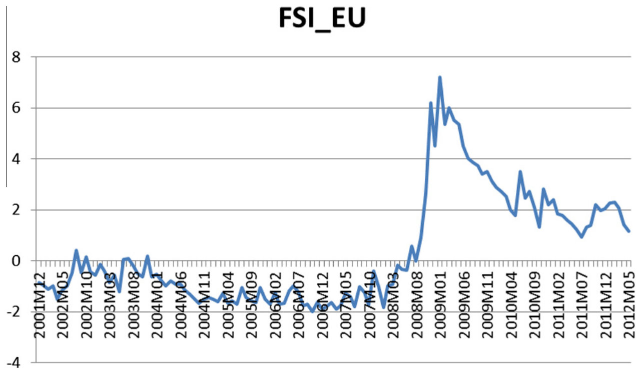

Fig. 8. Financial stress index for EU (sum of individual indices for 16 EU countries).

The fluctuations in the spillover series cannot be used as such for inferring contagion. The spillover series is just a particular transformation of the VAR variance decomposition but does not reveal the underlying sources of these changes. We adopt a formal statistical procedure to identify structural breaks driven by a sudden change in the transmission mechanism as evidence of contagion.

In the remainder of the analysis, we focus on a subset of EU countries over the period of the European debt crisis. The first reason is that we found the linkages between the eurozone bond markets to be much stronger than with those markets that still bear exchange rate risk (the CE countries and those EU countries opting out of the euro). The second reason is that for the eurozone countries, the period of the financial crisis is the more interesting one. Several studies have debated evidence of contagion on bond markets. From the discussion in Section 3.2, we also learned that the total spillover increased substantially after 2007. Formally, we can test whether a structural break occurs in the generalised FAVAR model with the Bai et al. (1998) test. This BLS test checks recursive estimates of the VAR and tests for a break with the sequential Sup-Quandt Andrews likelihood ratio test. We apply a correction for a possible change in volatility in the residuals before and after the breakdate. The optimal search should concentrate on the central 70% of the sample (Bai, 1997), but with on a sample of 3020 daily observations, we can reserve just 5% at the start and end of sample to ensure stable outcomes of the test. A significant break occurs within a rather wide confidence interval from July 9th, 2008 to September 30th, 2008 (the point break is on September 26th, 2008, with p-value 0.00). This break is clearly associated with the financial turmoil occurring in the summer of 2008 and culminating in the collapse of Lehman Brothers on September, 15th, 2008. We will look at subsamples in this post-Lehman period.[17]

In a first step, we apply the multivariate Qu–Perron test for a change in all parameters of the VAR to a dataset of weekly observations of the 10 EMU bond yield spreads, starting in the first week of January 2009 till the last week of January 2012. We set the trimming percentage for the minimum length of a sequence to 20% of the sample, and allow for a maximum of five breaks. The test statistics, together with the breakdates and their confidence intervals are reported in Table 3. The WD max test rejects the null of zero structural breaks and the sequential *F*-test further indicates three breaks are significant (out of a maximum of five) in the VAR coefficients. The sum of the change in coefficients before and after the breakdate are reported for each bond market, and measure the change in transmission each market spreads to others (the column elements in (3)).

A first significant break occurs in the first two weeks of August 2009. This break does not correspond to any particular event we detected before in Fig. 4). The sum of the coefficients of other markets shows a particularly marked – but negative – change for Belgium. This suggests that the effects of other EMU countries on the Belgian market declined considerably. This could be related to the resolution of the banking crises at Fortis and Dexia bank earlier in 2009. We would not label this event as contagion though, since the change from regime 1 to regime 2 implies a dampening of the effect.

A second break falls in between March 12th and 26th, 2010 which spans a period of uncertainty on the possible assistance to Greece. On March 28th, eurozone leaders agree to offer financial support, with the aid of the IMF. This seems to imply a particularly strong change for the core EMU countries, who are subject to large and positive changes in transmission. By contrast, the GIIPS receive now much less transmission after the event. A third significant breakpoint is identified late October 2010. At that moment, it was becoming increasingly likely that Ireland would request financial support after the bailout of Anglo Irish bank suddenly added 32% to the debt to GDP ratio. The break now indicates a substantial drop in transmission to France and the Netherlands, most likely related to the position of the banking sector in both countries. Since the break implies a drop in transmission, we would not classify this as contagion.

Of the three possible dates with evidence of contagion, we retain a single one related to the situation of Greek public finance, because of the strengthening of the transmission effect on other EMU countries (i.e. coefficients increase for Greece in the VAR equations of the other countries). This result shows that contagion is not prevalent during the eurozone crisis. The result seems at odds with the eyeball evidence from the spillover index in Fig. 3 where several very apparent breaks are present. It demonstrates that most of this spillover is driven by larger sized shocks being transmitted in pretty much the same way as prior to the crisis. Caporin et al. (2013) show similar evidence with Bayesian quantile regressions of a rather constant propagation of shocks on European CDS markets.

The breakdates suggested by the QP test are an indication of an overall change in the FAVAR model. The QP test can equally be applied on a subset of parameters in the model. Trying all possible combination would be infeasible for space constraints; therefore we concentrate our attention on the common suspects (GIIPS) both in terms of markets that emit contagion and those subject to contagion. We take the same example of Greece to check the spread of contagion from the trouble on the Greek bond market to other eurozone markets. As the previous results suggested a single event with contagion coming from Greece, we look for a single break on the VAR coefficients showing the transmission from Greece to other markets. The results in Table 4 show a single significant break, located towards the end of May 2010. This finding is in line with the previous result that the uncertainty surrounding the aid packages to Greece created on eurozone bond markets. The VAR coefficients for the subsamples delimited by the breakdate show no evidence of an upwards jump in coefficients in the direction of a particular market, but rather that contagion is evenly spread across all markets.

We now consider a more limited subsample of daily data on bond spreads in 10 EMU markets running from August 2nd, 2010 till August 31st, 2011. This period was characterised by strong turbulence on eurozone bond markets, with requests for assistance by Ireland and Portugal, IMF programmes, problems in the banking sector and the political responses in setting up the Systemic Risk Board, the European Stability Mechanism, the Pact for the euro,

[17] A practical reason is that the QP test is a sequentially recursive test for several breaks. Even for a small set of parameters to test for, the number of possible combinations rises exponentially. The algorithm demands big computing power and must be limited to a small number of series over rather short samples.

Table 3
QP test for contagion on all coefficients of the FAVAR model, weekly data, January 2nd, 2009–February 6th, 2012.

		Test statistic	Critical value
WD max test		4.03	1.21
Sequential test	For 2 breaks	105.00	10.31
	For 3 breaks	32.03	5.68
Break dates	Week August 7th, 2009	Week March 19th, 2010	Week October 29th, 2010
With confidence interval (*p*-value)	July 31st 2009–August 14th, 2009	March 12th–March 26th 2010	October 22th–November 5th, 2010
Coefficients	From regime 1 to 2	From regime 2 to 3	From regime 3 to 4
Constant	−0.4039	−0.2502	0.252
France	0.5554	0.5494	−1.0543
Netherlands	0.2225	0.9751	−0.7834
Spain	0.1433	−0.7886	1.3118
Italy	0.0632	0.3619	−0.1781
Belgium	−2.4792	0.5724	0.966
Greece	1.059	−0.68	−0.3447
Portugal	0.2709	−0.0674	−0.593
Ireland	−0.505	0.4615	0.143
Finland	−0.1175	−0.1761	0.2402
Austria	−0.0969	0.4445	−0.0495

Note: the table shows the WD max test for the existence of a break in the coefficients, and the sequential test statistics for comparing the significance of the *n*th + 1 break against the *n*th break. Estimated break dates are shown together with the 90% confidence interval. Coefficients shown are the sum of the changes in the VAR coefficients associated with the transmission from each bond market, before and after the breakdate.

Table 4
QP test for contagion from Greece, FAVAR model, weekly data, January 2nd, 2009–February 6th, 2012.

	Test statistic	Critical value
WD max test	199.58	1.21
Break dates	Week May 21st, 2010	
With confidence interval	May 14th 2010–May 28th, 2010	
Coefficients	From regime 1 to 2	
Constant	−0.0931	
France	0.1068	
Netherlands	0.5840	
Spain	0.0665	
Italy	−0.0043	
Belgium	−0.0401	
Greece	−0.1146	
Portugal	−0.0452	
Ireland	0.2134	
Finland	0.0292	
Austria	0.0858	

Note: the table shows the WD max test for the existence of a break in the coefficients, and the sequential test statistics for comparing the significance of the *n*th + 1 break against the *n*th break. Estimated break dates are shown together with the 90% confidence interval. Coefficients shown are the sum of the changes in the VAR coefficients associated with the transmission from each bond market, before and after the breakdate.

Table 5
QP test for contagion from Greece, FAVAR model, daily data, August 2nd, 2010–August 31st, 2011.

		Test statistic	Critical value
WD max test		5.69	1.19
Sequential test	For 2 breaks	4.13	1.68
Break dates	December 9th, 2010	April 22nd, 2011	
With confidence interval	December 6th–10th 2010	April 21st–23rd 2011	
Coefficients	From regime 1 to 2	From regime 2 to 3	
Constant	−0.0659	0.5130	
France	−0.0154	0.1503	
Netherlands	0.8364	−0.8213	
Spain	0.2961	1.0661	
Italy	−0.6156	1.6767	
Belgium	−1.1583	−0.0225	
Greece	−1.0512	1.2354	
Portugal	−5.6826	3.9546	
Ireland	7.7466	−1.3134	
Finland	−1.5716	0.2576	
Austria	0.4832	−0.9805	

Note: the table shows the WD max test for the existence of a break in the coefficients, and the sequential test statistics for comparing the significance of the *n*th + 1 break against the *n*th break. Estimated break dates are shown together with the 90% confidence interval. Coefficients shown are the sum of the changes in the VAR coefficients associated with the transmission from each bond market, before and after the breakdate.

and the intervention by the ECB. This period has also been the subject of analysis in several studies debating contagion on bond markets.

The multivariate QP test finds evidence of two significant breakdates (Table 5). The first in early December 2010, and a second at the end of April 2011. The first break is related to the Irish request for EU financial assistance. On December 7th, EU finance ministers agree on a joint EU-IMF assistance package. The coefficients in the second column of Table 5 show how forcefully this event changes the transmission from Ireland to other countries. The sum of all coefficients is positive and large. The only exception is Portugal, which now sees a contemporaneous drop in its transmission. Fiscal trouble in the Portuguese budget became already anticipated by financial markets, and unsurprisingly, Portugal is going to request aid in April 2011. The second breakdate is directly associated with this event. Over the month of April 2011, EU leaders negotiate the terms of the Portuguese aid package. The coefficients reported in column 3 show how strongly this modifies the transmission of Portugal to other markets.

The very concept of contagion refers to the situation when the transmission strengthens from a market under high stress to a market under relatively lower stress. The lengthy nature of the eurozone crisis makes a distinction between normal and crisis times rather complicated. In fact, looking at a financial stress index, developed by Slingenberg and de Haan (2011), for the whole EU (i.e. sum for 16 countries included in our analysis) in Fig. 8 would classify the whole period since late 2008 as high stress. The index is continuously several standard deviations above its mean, which is normalised to zero. The structural breaks in the VAR coefficients detected by the QP test in this period suggest limited contagion. Descending from the overall VAR to bilateral

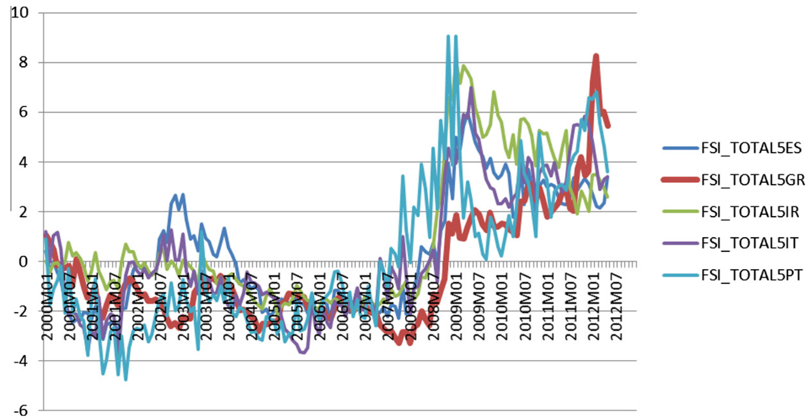

Fig. 9. Financial stress index for GIIPS.

linkages, we can measure the degree of financial stress from a source and receptor market around the time of the breaks in transmission. Fig. 9 shed some light on the cross-country variation of financial stress with horizontal lines drawn at the dates when the QP test reveals contagion in March 19th 2010 (Greece), April 22nd 2010 (Portugal) and December 9th 2010 (Ireland). In the first case, the FSI in the source market (Greece) was relatively lower than in the potential receptor markets (Ireland, Italy, Portugal and Spain); in the second case the stress in the source market (Portugal) was at a similar level than stress in one receptor market (Greece) and again substantially lower than in other receptor markets (Ireland, Italy and Spain), while in the third case the stress recoded in the source market (Ireland) exceeds the one of all the receptor markets (Greece, Italy, Portugal and Spain). From this point of view only the third case could be classified as contagion. However, the contemporaneous comparison of the FSI values recorded for different countries is not entirely appropriate given that values are normalised for each country.[18] Therefore, an alternative view could be to look at trends of FSI around those dates. Doing this we get slightly different view on the first two episodes. Financial stress in Greece in March 2010 and Portugal in April 2010 was relatively low vis-à-vis other markets, but closer inspection of Fig. 9 reveals that while stress in the receptor markets (Ireland, Italy and Spain) was declining, stress in Greece and Portugal jumped up in early 2010. This jump was reflected almost contemporaneously in the FSI of the other three countries (Ireland, Italy and Spain). This evolution is most visible in the FSI of Portugal. Stress increases sharply from early 2010 and is followed with a few months' lag by a rising FSI in Ireland, which in mid-2010 sharply reverses its previous downward-sloping trend. Even these two periods arguably show a sign of contagion.

4. Conclusion

The speed with which fiscal problems have spread across eurozone countries has come as a surprise. A crisis was not expected to initiate a string of fiscal bailouts, a collapse of the banking system and the involvement of the IMF in EU countries. Events since the start of the debt crisis in November 2009, coupled with a very rapid rise in bond spreads and the downgrading of all EMU countries but Germany, show how intertwined eurozone sovereign bond markets are. The strength of spillover is the result of economic and financial integration, which gradually proceeded since the start of the EMU. Additionally, stronger linkages, combined with economic imbalances and a weak economic policy, show the EU is also at risk of contagion on bond markets.

In this paper, we first measure spillover across the EU sovereign bond markets on a bilateral basis adapting the generalised VAR approach of Diebold and Yilmaz (2009) to account for common developments. We use daily data on bond spreads of 16 EU sovereigns (of both EMU ins and outs) over a long sample (2000–2012) that covers both the tranquil period as well as the financial crisis. Our results indicate that spillover has been a common feature of European sovereign bond markets but since the start of the financial crisis interdependence increased substantially. Nevertheless, there is a lot of heterogeneity in the bilateral linkages sent and received between specific markets.

Consequently, we propose a test of contagion applying the QP break test to the FAVAR model. The novelty of our approach is to test sudden changes in transmission by the endogenous search for breaks in transmission controlling for the size and volatility of shocks to markets, and to determine the direction of contagion. Strong interdependence does not imply increased susceptibility to contagion. There is some evidence of contagion during the eurozone fiscal crisis on the occasion of requests by Greece, Ireland and Portugal. These events created policy uncertainty that spread across EMU bond markets. Yet, most of the other increases in comovement across bond markets are driven by stronger shocks, for an otherwise identical propagation mechanism.

The political response to the eurozone crisis has been hesitant as it was often held up by doubts on the consequences of a sovereign default in Greece, Ireland or Portugal. Proponents of a bailout fear an immediate default of other eurozone sovereigns, as the spread of financial trouble to Spain and Italy might trigger an implosion of the European banking system and even the end of the eurozone itself, causing the "*mother of all financial crises*" (Eichengreen, 2010). Opponents to intervention hammer on the need to ensure fiscal discipline first, and avoid the moral hazard that might favour fiscal profligacy in the future. According to this view, a cascade of default is unlikely as debt positions have been unwound and can anyway be contained by intervention in banking markets. Our results show that both views are consistent. In normal times, an orderly solution to a debt crisis is possible. Strong underlying interdependence raises vulnerabilities to shocks so that contagion may suddenly flare up and lead to mispricing of sovereign risk in turbulent times (Aizenman et al., 2013; de Grauwe and Ji, 2013). A Greek default would have likely have a cascade effect across EMU if no other common policy actions are taken to stem contagion across markets.

[18] For example, the FSI increase that Greece experienced along late 2008 is very substantial given previous levels. For Portugal, the FSI was already very high in mid-2008 and the subsequent increase is relativaly smaller than in Greece.

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Appendix A.

See Tables A.1 and A.2.

Table A.1
Spillover table, no factor, full sample (May 2000–February 2012).

	Czech Republic	Poland	Hungary	Austria	Finland	Netherlands	France	Belgium	Spain	Italy	Greece	Portugal	Ireland	Denmark	Sweden	UK	From others
Czech Republic	57.80	8.26	7.32	2.76	0.57	0.81	1.81	3.02	3.83	4.42	0.88	0.91	2.13	4.45	1.00	0.03	42.20
Poland	7.36	64.89	6.77	1.17	0.22	0.24	0.82	1.89	2.59	3.15	1.16	1.19	2.06	5.64	0.83	0.02	35.11
Hungary	7.35	9.42	58.32	2.52	0.49	0.45	0.68	3.21	3.20	3.86	1.77	1.40	3.32	3.86	0.09	0.06	41.68
Austria	2.04	1.85	3.09	26.28	4.63	7.83	11.58	13.28	8.97	11.07	2.41	1.82	4.48	0.47	0.11	0.09	73.72
Finland	1.83	1.15	0.94	10.22	31.55	12.91	10.59	9.55	5.34	6.06	1.65	1.66	4.30	1.04	0.49	0.72	68.45
Netherlands	1.92	1.00	1.92	9.29	10.04	30.56	10.04	9.18	6.51	6.33	1.90	2.75	5.22	1.61	0.56	1.16	69.44
France	1.85	1.60	1.85	11.51	4.61	7.86	22.79	14.15	9.80	13.81	2.80	1.63	3.93	1.17	0.33	0.31	77.21
Belgium	1.96	1.65	2.07	8.35	3.00	5.29	9.50	24.56	15.64	15.95	2.22	2.67	6.63	0.26	0.16	0.09	75.44
Spain	1.55	1.18	1.31	5.98	1.63	3.94	7.30	12.15	31.04	16.95	3.34	4.12	8.90	0.15	0.15	0.31	68.96
Italy	2.02	1.60	1.61	4.54	1.46	3.03	4.91	14.25	20.39	30.38	3.49	4.26	7.70	0.21	0.07	0.08	69.62
Greece	1.24	0.87	0.84	2.84	1.71	2.08	5.29	10.21	10.65	8.55	39.04	6.64	9.91	0.01	0.02	0.12	60.96
Portugal	0.85	0.72	1.06	2.37	0.29	0.89	1.41	9.23	10.83	7.07	6.42	40.85	17.79	0.01	0.16	0.04	59.15
Ireland	1.18	0.87	1.10	3.57	1.97	2.69	4.08	8.60	10.78	5.50	5.88	11.38	42.30	0.01	0.05	0.03	57.70
Denmark	4.32	4.47	5.14	1.35	2.38	2.44	2.44	0.61	0.26	0.35	0.32	0.25	0.35	69.44	5.68	0.19	30.56
Sweden	1.27	1.02	0.57	0.15	0.59	0.85	0.39	0.23	0.47	0.31	0.04	0.13	0.09	4.77	88.48	0.64	11.52
UK	0.21	0.15	0.14	0.34	0.85	1.93	0.54	0.26	2.01	0.86	0.31	0.94	0.82	0.13	1.16	89.37	10.63
To others	36.95	35.83	35.74	66.98	34.44	53.23	71.37	109.81	111.27	104.24	34.59	41.74	77.63	23.79	10.86	3.88	852.35
To others (+own)	94.75	100.73	94.06	93.26	65.98	83.79	94.16	134.36	142.31	134.62	73.64	82.59	119.94	93.23	99.34	93.24	53.3%
From others	42.20	35.11	41.68	73.72	68.45	69.44	77.21	75.44	68.96	69.62	60.96	59.15	57.70	30.56	11.52	10.63	
Net spillover	5.25	−0.73	5.94	6.74	34.02	16.21	5.84	−34.36	−42.31	−34.62	26.36	17.41	−19.94	6.77	0.66	6.76	

Table A.2
Spillover table, de-factorised spread series, full sample (May 2000–February 2012).

	Czech Republic	Poland	Hungary	Austria	Finland	Netherlands	France	Belgium	Spain	Italy	Greece	Portugal	Ireland	Denmark	Sweden	UK	From others
Czech Republic	34.17	3.67	1.47	7.17	8.29	7.86	6.89	2.99	0.61	0.05	0.70	0.05	0.12	9.81	9.10	7.03	65.83
Poland	6.20	52.08	2.63	3.88	4.84	4.34	3.76	1.63	0.40	0.03	0.13	0.03	0.05	9.21	6.26	4.53	47.92
Hungary	4.96	8.34	82.36	0.23	0.24	0.10	0.01	0.01	0.04	0.15	0.04	0.01	0.07	2.91	0.36	0.17	17.64
Austria	1.85	0.33	0.01	16.39	13.88	14.00	13.68	8.03	1.45	0.14	0.60	0.02	0.09	8.05	10.63	10.87	83.61
Finland	2.20	0.38	0.03	12.65	15.77	14.37	13.25	7.66	1.46	0.10	0.54	0.03	0.10	8.81	11.09	11.57	84.23
Netherlands	2.23	0.36	0.01	12.53	14.42	15.42	13.17	7.63	1.62	0.11	0.50	0.07	0.11	9.01	11.10	11.72	84.58
France	1.89	0.33	0.03	12.75	13.72	13.82	15.64	8.30	1.75	0.37	0.46	0.01	0.06	8.69	10.93	11.23	84.36
Belgium	1.15	0.16	0.06	10.86	11.51	11.76	12.66	19.25	5.55	1.90	0.80	0.05	0.19	5.59	9.37	9.13	80.75
Spain	0.59	0.05	0.37	6.94	7.21	8.00	8.96	8.10	33.68	6.27	0.03	0.59	1.54	3.31	6.77	7.62	66.32
Italy	0.59	0.08	0.32	4.38	6.15	6.35	5.44	9.77	17.27	33.91	0.02	0.52	0.63	2.87	5.64	6.06	66.09
Greece	1.61	0.59	0.30	6.30	5.26	5.73	3.97	0.87	0.20	0.05	55.60	2.08	2.37	6.09	5.03	3.95	44.40
Portugal	0.09	0.11	0.05	0.04	0.04	0.05	0.10	1.99	4.37	0.47	4.39	71.12	16.79	0.22	0.10	0.08	28.88
Ireland	0.08	0.18	0.20	0.30	0.73	0.63	0.53	1.36	3.70	1.74	3.29	12.77	73.61	0.20	0.23	0.46	26.39
Denmark	2.90	0.86	0.09	11.44	13.42	13.18	12.23	6.35	0.84	0.03	0.55	0.02	0.03	15.40	11.91	10.77	84.60
Sweden	2.71	0.67	0.02	11.26	12.91	12.84	11.70	6.82	1.89	0.19	0.43	0.10	0.09	9.60	17.96	10.82	82.04
UK	2.32	0.46	0.01	11.22	12.98	13.21	11.79	6.98	2.36	0.31	0.39	0.17	0.16	8.27	10.89	18.49	81.51
To others	31.37	16.56	5.59	111.97	125.60	126.24	118.12	78.48	43.50	11.91	12.85	16.51	22.39	92.63	109.40	106.01	1029.14
To others (+ own)	65.55	68.64	87.94	128.35	141.37	141.67	133.76	97.73	77.19	45.82	68.45	87.63	96.00	108.03	127.36	124.50	64%
From others	65.83	47.92	17.64	83.61	84.23	84.58	84.36	80.75	66.32	66.09	44.40	28.88	26.39	84.60	82.04	81.51	
Net spillover	34.45	31.36	12.06	−28.35	−41.37	−41.67	−33.76	2.27	22.81	54.18	31.55	12.37	4.00	−8.03	−27.36	−24.50	

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