Do European Monetary Union Sovereign CDS and Bond Markets Share the Same Information to Price Credit Risk?

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Abstract

We analyze the extent to which the prices in the sovereign credit default swap (CDS) and bond markets reflect the same information on credit risk in the context of the European Monetary Union. The empirical analysis is based on the theoretical equivalence relation that should hold between the CDS and bond spreads in a frictionless environment. We first test and find evidence in favour of the existence of persistent deviations between both spreads during the crisis but not before. Such deviations are found to be related to some market frictions, like counterparty risk, market-illiquidity, and funding costs. We also find evidence suggesting that the price discovery process is state-dependent. Specifically, the levels of counterparty and global risk, funding costs, market liquidity, the volume of debt purchases by the European Central Bank in the secondary market, and the agreements of banks to accept losses on their Greek bonds are all found to be significant factors in determining which market leads price discovery.

Keywords: Sovereign Credit Default Swaps, Sovereign Bonds, Credit Spreads, Price Discovery.

JEL Codes: G10, G14, G15.

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1. Introduction

In the last years many studies have analyzed the relationship between CDS and bond spreads for corporate as well as for emerging sovereign reference entities. [1] However, the relation between sovereign CDS and bonds markets in developed countries has not attracted much interest until very recently, mainly for two reasons. First, sovereign CDS and bonds spreads in developed countries have been typically very low and stable given the perceived high credit quality of most issuers (see Table 1). Second, trading activity in this segment of the CDS market was typically low.

However, the global financial crisis that followed the collapse of Lehman Brothers in September 2008 triggered an unprecedented deterioration in public finances of the world's major advanced economies in a peacetime period. Since 2010, some countries in the euro area, including Greece, Ireland and Portugal and, to a lesser extent, Spain and Italy have faced some episodes of heightened turbulences in their sovereign debt markets. Against this context, the levels of perceived credit risk and the volume of trading activity in the sovereign CDS markets in many advanced economies have increased.

The previous literature have paid some attention to investigate the relationship between the corporate bond market and the corporate CDS markets but only a few papers have studied whether the empirical regularities identified in the corporate markets, including those related to price discovery, are also found in the case of sovereign reference entities. The aim of this paper is shed light on these latter issues within the context of the recent episodes of sovereign-debt crises in several countries in the European Monetary Union (EMU).

Specifically, we analyse the theoretical equivalence relation between the sovereign bond yield spread (with respect to a risk-free benchmark) and the corresponding CDS spread.[2] Abstracting from market frictions and other contractual clauses both spreads

[1] We discuss the related literature in Section 2.

[2] The results are obtained using the German bond as a proxy of the risk free asset, as in e.g. Geyer, Kossmeier, and Pichler (2004), Bernoth, von Hagen, and Schuknecht (2006), Delis and Mylonidis (2010), Favero, Pagano, and Von Thadden (2009), Foley-Fisher (2010), or Palladini and Portes (2011), among others

should reflect the same information on the credit risk of a given reference entity and therefore should be equal. In other words, the *basis*, defined as the difference between the CDS spread and the corresponding bond spread should be zero. Or in case it differs from zero, the differences should be purely random but not related to any specific factor. Moreover, in such a frictionless scenario; both spreads (or credit-risk prices) should incorporate the credit risk information in a similar way, i.e. both markets should equally efficient in terms of the process of credit-risk price discovery. The current European sovereign debt crisis poses a particularly interesting scenario to test for the previous hypotheses. In particular, we analyze the bond-CDS equivalence relation from three different perspectives.

First, we test the "no-arbitrage" theoretical frictionless relation that equates the bond and the CDS spreads to find that there are persistent deviations from that relation. Interestingly, we find that the deviations begin with the outset of the subprime crisis with no evidence of such deviations before then.

Second, based on the previous finding, we study the possible causes of the deviations between the bonds and the CDS spreads. To that aim, we estimate de effects of several factors on the basis. In particular, we find that the counterparty risk indicator has a negative and significant effect on the basis, especially after September 2008, when some of the most active sellers of CDS started to face financial difficulties. Funding costs have a negative effect on the basis due to their stronger effect on the demand for bonds relative to the demand for CDS, as the latter require less funding to take on the same risk position. A higher degree of liquidity in the bonds market relative to the CDS market has a positive effect on the basis given that *ceteris paribus* a more liquid bond implies a lower bond yield and spread. The volume of debt purchases by the European Central Bank (ECB) in the secondary market that have taken place after May 2010 affects positive and significantly to the basis. These purchases exert a negative effect on bonds spreads. The fact that the such effect is not present (or it is weaker) in the case of the CDS spreads may indicate that these interventions affect other components of the bonds prices other than credit risk (e.g. through a fall in the bond liquidity premium) or, simply, induce some over-price effect in the bonds market, for a given a level of credit-risk. Although the effect of the level of global risk, proxied by the VIX Index, is not significant, the country-specific risk premium, measured through the stock market index, affects the basis positively and significantly. This suggests that while global volatility is priced similarly in both markets, the idiosyncratic volatility is not, with the CDS market reacting more to changes in this latter case. Further, in periods in which shocks originated in the bond market are transmitted into the CDS market, which we label as *spillovers*, we find an overreaction-effect in this last market; that is, the information coming from the bond market has a more direct and stronger effect on the CDS market than the information flowing from the CDS market to the bond market. Finally, the effect of the lagged basis suggests a high degree of persistence and, hence, a relatively low speed of adjustment of the basis.

Third, motivated also by the previous evidence on persistent bases, we address the question of which market leads the credit-risk price discovery process. To this aim, we follow a dynamic price discovery approach, based on Gonzalo and Granger (1995). Our analysis reveals that the price discovery process is state-dependent. Specifically, the levels of counterparty and global risk and the successive agreements of private banks to accept losses on their holdings of Greek bonds worsen the ability of the CDS market to provide efficient information in the price discovery process. The effect of counterparty risk is due to the perception of a lower quality of the protection sold in the CDS market when this risk is high. The effect of global risk could be due to the fact that the information contained in bond spreads is more efficient during periods of high global risk. The agreements of private banks to accept losses on the Greek bonds could have caused a lack of confidence of investors in the CDS market after such agreements. On the other hand, the level of funding costs and the volume of sovereign debt purchased by the ECB worsens the efficiency of the bond market in the price discovery process. Regarding funding costs, they affect the bond buyers to a higher extent than to the CDS buyers, as the CDS market allows for more leveraged positions. The operations of the ECB could have limited the quality of the information contained in the bonds prices on the true credit risk of these assets.

The remainder of the paper is organized as follows: Section 2 discusses the related literature. Section 3 describes the data. Section 4 presents the methodology and the results based on the analysis of persistent deviations between CDS and bond spreads. Section 5 analyses of the determinants of the basis. Section 6 presents the results of the dynamic price discovery test. Section 7 contains some final remarks.

2. Related literature

In this section we focus on the branch of the literature on CDS and bonds spreads more closely related to the three questions approached in this paper: persistent deviations between bond and CDS spreads, determinants of such deviations, and the price discovery process in the bonds and CDS markets.

We investigate the existence and persistency of deviations between CDS and bond spreads on the basis of the notion of arbitrage introduced by Hogan, Jarrow, Teo, and Warachka (2004). Until now, this approach has only been applied to corporates CDS and bonds, in Mayordomo, Peña and Romo (2011a). In particular, they analyse the existence of persistent deviations between CDS and asset swap spreads of European corporations using the pre-crisis period (before 2008) and the crisis (after then). Their results show that there are persistent deviations both in the pre-crisis and the crisis periods.

There is an extensive literature addressing the determinants of corporate bond and CDS spreads.[3] Although this type of analysis is less frequent for the case of sovereign references, this topic is attracting an increasing attention since the inception of the EMU.[4] Our aim, however, is not to study the determinants of the CDS or the bond spreads but, rather, the determinants of the basis to test whether both markets reflect different information. Although the analysis of the determinants of the basis is less frequent than the analysis of the individual credit spreads, there are some earlier contributions in the literature on the sovereign credit markets. For instance, Fontana and Scheicher (2010) employ weekly data from January 2006 to June 2010 to analyse the determinants of the basis to find that the sovereign bases are significantly linked to the cost of short-selling bonds and to country specific and global risk factors. In his analysis of the CDS-bond parity, Levy (2009) finds that the frictionless parity relation does not hold for emerging markets sovereign debt but he argues that an important part of the deviations can be attributed to liquidity effects. Küçük (2010) relates the CDS-bond

[3] See, for instance, Elton, Gruber, and Agrawal (2001), Collin-Dufresne, Goldstein and Martin (2001), Chen, Lesmond and Wei (2007), among others study the determinants of the corporate bond spread. The studies analyzing the determinants of the corporate CDS spreads include Longstaff, Mithal and Neis (2005), and Ericsson, Jacobs and Oviedo-Helfenberger (2009).

[4] See e.g. Codogno, Favero and Misale (2003), Geyer, Kossmeier and Pichler (2004), Bernoth, von Hagen and Schuknecht (2006), Favero, Pagano and Von Thadden (2009), Beber, Brandt and Kavajecz (2009), or Mayordomo, Peña, and Schwartz (2011).

basis for 21 emerging market countries between 2004 and 2008 to some factors capturing bond liquidity, speculation in CDS market, CDS liquidity, equity market performance and some global macroeconomic variables. Foley-Fisher (2010) studies the relation between bond and CDS spreads for ten EMU countries on the basis of a theoretical model of heterogeneous investors' expectations. He shows that the basis is consistent with a relatively small dispersion in the beliefs of investors on the probability that certain European countries would default.[5]

Finally, the most frequent analysis of the CDS-bond relation in corporate and sovereign credit markets is based on the concept of price discovery. Most of the recent papers study price discovery on the basis of either Hasbrouck's (1995) or Gonzalo and Granger's (1995) methodologies. Both approaches are supported by an empirical test based on a VAR with an Error Correction Term model. For the period before the subprime crisis a recurrent empirical finding is that the CDS market reflects the information more accurately and quickly than the bond market in the corporate sector (see Blanco, Brennan, and Marsh, 2005; or Zhu, 2006; among others). Most of the analyses of price discovery in sovereign markets have been applied to emerging markets. For instance, Ammer and Cai (2007) find that bond spreads lead CDS premia more often than had been found for investment-grade corporate credits. Chan-Lau and Kim (2004) find that it is difficult to conclude that one particular market dominates the price discovery process. Using data from eight emerging market countries for the period 2003-2006, Bowe, Klimaviciene, and Taylor (2009) find that the CDS market does not, in general, dominate price discovery, which appears to be country-dependent. The recent crisis has increased the interest on this question in the context of the European sovereign debt markets. For instance, Fontana and Scheicher (2010) find that since the outset of the crisis, the bond market has a predominant role in price discovery in Germany, France, the Netherlands, Austria, and Belgium while the CDS market is playing a major role in Italy, Ireland, Spain, Greece and Portugal. Palladini and Portes (2011) use data on six Euro-area countries over the period 2004-2011 and find that the CDS market moves ahead of the bond market in terms of price discovery. They also show that short-run deviations from the frictionless equilibrium relation between bond and CDS spreads persist longer than it would take the participants in one market to

[5] Analyses of the basis in the corporate credit market include Trapp (2009), Nashikkar, Subrahmanyam, and Mahanti (2008), and Bai and Collin-Dufresne (2009), among others.

observe the price in the other. Delatte, Gex, and Lopez-Villavicencio (2010) find that the bond market leads the price discovery process in the core European countries in periods of relative calm while in periods of turbulence the CDS market leads the price-formation process. In the high-yield European countries, the CDS spread reflect credit risk more adequately than the bond spreads in periods of both calm and tension but the leadership of the CDS spread is exacerbated by financial turmoil. All these analyses have been carried out based on static measures of price discovery, such that a single measure is obtained for the entire period under analysis. However, as argued by Longstaff (2010), the nature of the price-discovery process in financial markets can be state dependent. Thus, Delis and Mylonidis (2010) study the dynamic interrelation between bond and CDS spreads on the basis of a Granger causality test. They find feedback causality during periods of financial distress. As we will show later, Gonzalo and Granger (1995) test is more useful to determine who the leader is and who the follower is.

3. Data

The data consists of daily 5-year sovereign bond yields and CDS spreads for eleven EMU countries (Austria, Belgium, Finland, France, German, Greece, Ireland, Italy, The Netherlands, Portugal, and Spain) from January 2004 to October 2011. Bond yields are obtained from Reuters and CDS spreads from Credit Market Analysis (CMA), which reports data (bid, ask and mid) sourced from 30 buy-side firms, including major global investment banks, hedge funds, and asset managers.[6]

Table 1 reports the main properties of the data. As becomes evident from this table, average CDS rates vary substantially across countries and periods. For the period 2004-2008, the lowest average CDS spread was 5 basis points (bp) for Germany and the highest one was 23 bp points for Greece. For the same period, the lowest average bond spread is 4 bp for both France and The Netherlands, and the highest average is 25 bp for Greece. We note that CDS spreads are on average higher than bond spreads in most of the countries, i.e. the basis, defined as the difference between the CDS and bond spreads, is positive. We observe an increase in both the average and the volatility of

[6] Mayordomo, Peña and Schwartz (2011) compare the quality of the data on CDS from different providers and find that CMA produces, on average, the most reliable data.

CDS and bond spreads over the subsequent years (from 2009 on) in most of the countries and especially in the peripheral ones (Greece, Ireland, Portugal, Spain and Italy). The exceptions are Austria, Finland, and The Netherlands in which the average credit spreads in 2010 were lower than in 2009 but they increased again in 2011. For the period 2009-2011, the lowest annual average CDS spread was 31 bp for Finland in 2010 and the highest annual average was 2,075 bp for Greece in 2011 (being the maximum daily CDS spread at 6,752 bp the 26th of September, 2011). The lowest annual average bond spread was -6 bp for Finland in 2010 and the highest was 1,644 bp for Greece in 2011.[7] In the last three years, the CDS spreads are, again, on average, higher than the bond spreads in most of the countries. As for the period 2004-2008, the CDS spread is on average higher than the bond spread in 2009, 2010, and 2011 in most of the countries (the exceptions are: Ireland and Portugal in 2011 and Greece in 2009 and 2010).

< Insert Table 1 here>

As for the rest of the data used in the subsequent estimations, the country-stock and global risk indexes, which are proxied by means of the implied stock market volatility (we use the VIX for the global indicator)), are obtained from Reuters. To capture funding costs we use the difference between the 90-day US AA-rated commercial paper interest rates for financial companies and the 90-day US T-bill, both from Datastream. We employ liquidity measures for the sovereign CDS and bonds which are obtained from the bond and CDS bid-ask spreads. Bond bid-ask prices are obtained from Reuters while CDS bid-ask spreads come from CMA. To proxy for the counterparty risk on the side of CDS dealers, we employ the CDS spreads of the 14 banks most active as dealers in the CDS market. These CDS spreads are obtained from CMA. The information regarding the European Central Bank (ECB) bond purchases which took place after May 2010 was obtained from the ECB webpage.

4. Are there persistent deviations between CDS and bond spreads?

Suppose that an investor buys a bond at its par value with a maturity equal to T years and a yield-to-maturity equal to *ytm*. Also, assume that at the same time the investor

[7] The negative sign for the bond spread in Finland in 2010 is due to the fact that the average yield of the Finnish bond was lower than the one of the German bond.

buys protection on such reference entity for T years in the CDS market and the premium of such contract is s. The investor has eliminated the default risk associated with the underlying bond and the investor's net annual return is equal to $ytm - s$. Absent any friction, arbitrage forces would imply that the net return should be equal to the T-year risk-free rate, which we denote by r. Alternatively, if $ytm - s < r$, then by means of a short position in the bond, writing protection in the CDS market and buying the risk-free bond the investor could have obtained a positive profit without any risk. If, on the contrary, $ytm - s > r$, the investor could obtain a certain profit by buying the risky bond, buying protection in the CDS market and taking a short position in the risk-free bond. Hence, in equilibrium, $ytm - r = s$.

In order to investigate the existence and persistency of deviations between CDS and bond spreads, that would violate the previous equilibrium relation, we apply the statistical arbitrage test employed by Mayordomo, Peña and Romo (2011a). This test is based on the notion of arbitrage introduced by Hogan, Jarrow, Teo and Warachka (2004) according to which, absent market frictions, an arbitrage opportunity (in an statistical sense) represents a zero-cost, self-financing trading opportunity that has positive expected cumulative trading profits with a declining time-averaged variance and a probability of loss that converges to zero as time passes. Bearing in mind that, within the logic of this methodology, the existence of arbitrage opportunities is conditioned to the absence of market frictions, in our application of this test, we interpret the results in a rather agnostic way. In particular, we do not identify persistent deviations between both spreads with unexploited arbitrage opportunities. Indeed, when such deviations are found we relate them, in a statistical sense, to several potential market frictions (see Section 5).

To test for the existence of persistent deviations from the zero-basis benchmark, we first compute the increase in the discounted trading profits that an investor would obtain under the assumption of no trading and funding costs. Specifically, the profits from a given investment strategy, in the sense just stated, are defined as the basis times the contract notional value. We compute such profits quarterly and the payment on a given date t is added to the trading profits accumulated from the first investing date to the last date, t-1. The accumulated profits constructed in this way are assumed to have been invested or borrowed at the risk-free rate in the interim, from t-1 to t. The cumulative trading profits are then discounted up to the initial date. The increase in the discounted cumulative trading profits at a given date *t* is denoted by Δv_t and is assumed to evolve according to the following process:

$$\Delta v_t = \mu t^{\theta} + \sigma t^{\lambda} z_t \qquad (4)$$

for $t = 0, 1, 2, \ldots, n,$ with n denoting the last investment date and where z_t are innovations. We assume the following initial conditions: $z_0 = 0$ and $v_0 = 0$ (i.e. the strategy is self-financed). Parameters θ and λ determine whether the expected trading profits and the volatility, respectively, are decreasing or increasing over time and their intensity. Under the assumption that z_t is an *i.i.d. N(0,1)* variable, the expectation and variance of the discounted incremental trading profits in equation (4) are $E[\Delta v_t] = \mu t^{\theta}$ *and* $Var[\Delta v_t] = \sigma^2 t^{2\lambda}$, respectively. Then, the discounted cumulative trading profits generated by a given strategy satisfy:

$$v_n = \sum_{t=0}^{n} \Delta v_t \sim N\left(\mu \sum_{t=0}^{n} t^{\theta}, \sigma^2 \sum_{t=0}^{n} t^{2\lambda}\right) \qquad (5)$$

We then define the log-likelihood function for the increments in equation (5) and estimate the parameters of interest $(\mu, \theta, \sigma, \lambda)$ by maximizing that function using a non-linear optimization method based on a Quasi-Newton-type algorithm. Then, we implement formally the notion of statistical arbitrage test outlined before through the specification and testing of the following three simultaneous hypotheses:

$$H1: \lim_{t\to\infty} E^{P}[v(t)] > 0 \quad \Rightarrow \quad \mu > 0, \quad and$$
$$H2: \lim_{t\to\infty} P(v(t) < 0) = 0 \quad \Rightarrow \quad \lambda < 0 \quad or \quad \theta > \lambda, \quad and$$
$$H3: \lim_{t\to\infty} Var[\Delta v(t) \mid \Delta v(t) < 0] = 0 \quad \Rightarrow \theta > \max\left\{\lambda - \frac{1}{2}, -1\right\}.$$

Statistical arbitrage requires that the expected cumulative discounted profits, *v(t),* are positive (*H1*), the probability of loss converges to zero (*H2*), and the variance of the incremental trading profits *v(t)* also converges to zero (*H3*).[8]

Hence, these three conditions must be simultaneously satisfied to have support for the existence of persistent non-zero basis. In practice, this implies an intersection of several sub-hypotheses. To maximize the power of the test, instead of testing whether the

[8] Implicit in hypothesis *H3* is the idea that investors are only concerned about the variance of a potential decrease in wealth. Whenever the incremental trading profits are non-negative, their variability is not penalized.

previous hypotheses are simultaneously satisfied, we redefine the null hypothesis as the absence of persistent non-zero basis and so, our test is based on the following union of sub-hypotheses which are given by the complementary of the previous hypotheses (see Jarrow, Teo, Tse, and Warachka, 2007):

$$
\begin{aligned}
&H1^C : \mu \le 0, \quad or \\
&H2^C : \lambda \ge 0 \quad and \quad \theta - \lambda \le 0, \quad or \\
&H3_1^C : \theta - \lambda + \frac{1}{2} \le 0, \quad or \\
&H3_2^C : \theta + 1 \le 0,
\end{aligned}
$$

where $H1^C$ and $H2^C$ are the complementary of hypotheses *H1* and *H2*, respectively, while $H3_1^C$ and $H3_2^C$ come from the complementary of hypothesis *H3*. If, at least, one of the last four hypotheses above is satisfied, we conclude that no persistent deviations exist.

To test these hypotheses we need to estimate the *p-values* for the previous restrictions. To this aim, we follow the methodology developed by Politis, Romano, and Wolf (1997 and 1999). This technique provides an asymptotically valid test under weak assumptions. Specifically, our analysis leads to two one-tail tests:

a) H_0: no persistent deviations and H_A: negative deviations (the bond spread is significantly higher than the CDS spread);
b) H_0: no persistent deviations and H_A: positive deviations (the CDS spread is significantly higher than the bond spread).

The results of these tests are summarized in Table 2. Panels A and B report the results for the period ranging from January 2004 to September 2008 for negative and positive bases, respectively. Panels C and D report the corresponding results for the period ranging from September 2008 to October 2011. As shown in Panels A and B, we cannot reject the null hypothesis (no persistent deviations) at any standard significance level. This result holds irrespectively of whether we consider either positive or negative bases. However, after September 2008, the CDS spread is persistently higher than the bond spread in six cases (see Panel D) while none of the countries analysed presents a persistent negative basis, as shown in Panel C.

< Insert Table 2 here >

As a conclusion, the above results reveal that the zero-basis hypothesis cannot be rejected when we consider the pre-crisis period although temporary non-zero basis are not rare during the crisis. This last result must be interpreted with caution since, as argued before; a non-zero basis cannot be understood mechanically as an opportunity for arbitrage. For instance, Schonbucher (2003) and Mengle (2007) emphasize that shorting a bond with a required maturity is not always a feasible option. Moreover, the fact that recurrent non-zero bases seem to be common during the crisis period may be symptomatic of the presence of other restrictions and frictions that prevent a perfect timeless alignment between the CDS and the bond spreads and whose relevance may have been exacerbated by the crisis itself. This could be the case, for instance, of funding costs, differences in liquidity across markets and counter-party risk in the CDS market. In the following section we test for the significance of these (and other) factors, as potential explanatory variables for the cases of non-zero basis detected during the crisis.

5. The determinants of the basis

In this section we test whether the differences between the CDS and bond spreads are purely random or, alternatively, whether they are related to any market-specific or global factors. Due to the observations corresponding to the last months of 2010 and the ones of 2011 included in the sample, the basis is a non-stationary variable. For this reason, instead of analysing the determinants of the basis we study the determinants of the relative basis, that is defined as the difference between the CDS and bond spreads relative to the average credit spread which is obtained as the simple mean of the CDS and the bond spreads. We consider the following potential explanatory factors:

a. Counterparty Risk. In principle, the higher the counterparty risk of the seller of a CDS is, the lower should be the CDS price charged as a result of the lower quality of the protection. We test for this effect by using the first principal component obtained from the CDS spreads of the main 14 banks which act as dealers in that market.[9] The first principal component series should reflect the common default probability and, hence, it is akin to an aggregate measure of counterparty risk.[10] Actually, the first principal component for the series of CDS spreads of this set of dealers explains 87.5% of the total variance of the observed variables. We use the counterparty risk variable lagged one period to avoid any problem of endogeneity derived from the potential contemporaneous effects of the banks' activity on the sovereign credit spreads.

b. Liquidity. In theory, one would expect that higher liquidity in the bond market relative to the CDS market would go hand in hand with a higher basis, since a more liquid bond implies a lower spread in that market. To test for this relative liquidity effects, we construct a ratio of relative liquidity between the CDS and the bond. Specifically, the degree of liquidity in the CDS market is proxied by the bid-ask spread of the CDS premium. The higher this spread is, the lower is the degree of liquidity in the CDS market. A similar measure of liquidity is computed for the bond market. The ratio between both measures is taken as indicative of the relative liquidity in the CDS market vis-à-vis the bond market. As this ratio rises, liquidity in the CDS market relative to the bond market falls and so, the basis would, in principle, increases.

c. Financing Costs: One would expect that higher financing costs would lower the demand for bonds and could lead to a decrease in prices, and hence, to higher bond spreads. The effect of funding costs on CDS spreads should be lower given that in this case the required amount of funding to get the same (gross) risk position is lower (i.e. risk-leverage is higher in the case of the CDS investment). For this reason, in principle, an increase in financing costs would have a negative effect on the basis. Due to the difficulty in obtaining data on institution-level funding constraints, we use the spread between financial commercial paper and T-bill rates as a common proxy for the funding constraints faced by financial intermediaries, as in Acharya, Schaefer and Zhang (2007).

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[9] The 14 main dealers are: Bank of America, Barclays, BNP Paribas, Citigroup, Credit Suisse, Deutsche Bank, Goldman Sachs, HSBC, JP Morgan, Morgan Stanley, Royal Bank of Scotland, Societé Generale, UBS, and Wachovia/Wells Fargo. These dealers are the most active global derivatives dealers and are known as the G14 (see for instance ISDA Research Notes (2010) on the Concentration of OTC Derivatives among Major Dealers).

[10] The use of the dealers' CDS spreads as a proxy of counterparty risk is based on the Arora, Ghandi, and Longstaff (2009) study which analyses the existence of counterparty risk in the corporate CDS market.

Specifically, we use the spread between the 90-day US AA-rated commercial paper interest rates for the financial companies and the 90-day US T-bill.

d. Domestic and global risk premiums: As additional potential explanatory variables for the basis, we consider a measure of the country and global risk premium. If both the CDS and bond spreads are prices for the same credit risk, the effect of the country-specific and global risk premia on the basis should be no significantly different from zero. However, a significant non-zero effect would suggest that one of the markets reacts more to changes in the risk premiums than the other. To control for the fact that this idiosyncratic and global volatility could be priced differently in the two markets, we use the previous risk factors as additional explanatory variables. The country-specific risk premium is proxied by means of the stock market volatility, defined as the absolute value of the stock index returns. The global risk premium or global risk is proxied by means of the VIX Index. The correlation between the VIX Index and the counterparty risk variable is around 0.8. Thus, in order to avoid any multicollinearity problem, we modify the VIX Index variable and define it as the residual of the regression of this variable onto the first principal component CDS spreads corresponding to the main CDS dealers such that counterparty risk and the VIX are now orthogonal variables.

e. Bond-CDS Spillovers: We here use the notion of spillovers between the CDS and the bond markets as the variation in the CDS (bond) spread that is not attributable to its past values but to contemporary shocks to the bond (CDS) spread. To measure such spillovers or contagion-effects, we use a procedure based on Diebold and Yilmaz's (2010) methodology (see Appendix A.1 for details). The Bond-CDS spillovers variable is obtained after computing the ratio between the spillovers flowing from the changes in bond spread to the CDS spread relative to spillovers from the CDS to the bond spread changes. This variable reflects the increase in either spread due to a direct effect coming from the other market.[11] A positive sign implies that when the ratio increases, then the basis widens, or in other words, that the CDS spread increases with respect to the bond spread.

[11] We work in relative terms because the dependent variable is the difference between both credit spreads.

f. Volume of debt purchased by the European Central Bank: We use as an additional explanatory variable the amount of sovereign debt purchased by the ECB in the secondary market from May 2010 onwards. This information is available only on a weekly basis and so the effect of the ECB debt purchases is proxied by means of the amount of debt that was purchased the week previous to the current date. These purchases are supposed to decrease both CDS and bond spreads but, in principle, it is natural to think that they would have a more direct effect on the bond price. Thus, following these interventions, the bond spread would decrease to a higher extent than the decrease in the CDS spread, thus affecting positively to the basis.

g. Lagged basis: The lag of the basis should absorb any past information transmitted into the current observation. Due to the existence of persistent deviations between the CDS and bond spreads documented in Section 4, we expect a positive sign.

We estimate the coefficients for the above variables of interest for the period which spans from January 2004 to October 2011 using a fixed-effects estimation procedure that is robust to heteroskedasticity. We use the bootstrap methodology to correct for any potential bias in the standard errors due to the use of generated regressors. The results are reported in Table 3. Column 2 of that table reports the standardized coefficients (i.e., the regression coefficient as in Column 1 multiplied by one standard deviation of the corresponding explanatory variable). All the variables, including the dependent variable, are expressed on a per unit basis with the exception of the volume of bonds purchased by the ECB in a given week which is defined in billions of euros.

The counterparty risk proxy has a negative and significant effect, as expected, which confirms its relevance on the levels of the CDS spreads. Funding costs have a negative effect due to their stronger effect on the demand for bonds relative to the demand for CDS that require a lower amount of funding. A low degree of liquidity in the CDS market relative to the bond market has a positive effect given that a more liquid CDS implies a lower CDS spread. The global risk variable is not significant which may suggest that both markets reflect global risk to a similar extent. Nevertheless, the country-specific risk measure seems to be priced differently in both markets. The positive and significant effect could be explained by the fact that the CDS spread is more affected by increases in the country specific risk premium. We also find that the shock-spillovers ratio has a positive and significant effect. That is, when the shock transmission from the bond to the CDS market dominates the shock transmission in the opposite direction, then the basis widens. One economic implication of this empirical finding is that in periods in which the ratio of spillovers from the bond market to the CDS market increases we observe *ceteris paribus* a relative overreaction in the a CDS market and thus, the CDS spread will signal a stronger default probability for a given reference name than the bond spread. In line with the results of the previous section, we find a high level of persistency in the relative basis. That is, there is a relatively low speed of adjustment towards the long-run bond-CDS equivalence relation. Finally, the constant term reflects whether the relative basis differs, on average, from zero and the magnitude of such deviation. We observe that the relative basis is on average significantly positive, suggesting that the bond-CDS equivalence relation does not hold even when we take into account the market frictions described above and the costs that are needed to trade the basis. Nevertheless, its magnitude is low relative to the average relative basis during the period 2004 – 2011 (2% relative to 32%). Thus, when we take into account the determinants of the basis, the magnitude of the average relative deviation is close to zero and is reduced in 93.75%. This result confirms the strong influence on these determinants to guarantee the bond-CDS equivalence relation.

The relatively high R-square of this regression is mainly due to the effect of the lagged basis. However, it should be noted that the explanatory variables retain a relatively high explanatory power even when we ignore the lagged basis, in which case the R-square is 0.29. Actually, this is of a similar magnitude of the one reported by Trapp (2009) on a daily basis for corporates using firm fixed-effects but ignoring the effect of the lagged basis.

In Column 2 we report the standardized coefficients. As expected, the strongest effect is the one corresponding to the lagged relative basis. Among the rest of the variables, the strongest effects are the ones coming from the counterparty risk and the financing costs. An increase equal to one standard deviation in any of the previous variables would diminish the relative basis in -0.7%.

< Insert Table 3 here >

6. Price-discovery analysis

An efficient price discovery process is characterized by a quick adjustment of market prices from the old to the new equilibrium as new information arrives (see e.g. Yan and Zivot, 2007). The previous literature that has tried to measure this form of market-efficiency has focused on static price-discovery analyses. In contrast, we here show that the price discovery process in the markets for sovereign credit risk in the EU does not show a time-invariant pattern (Section 6.1). Given this finding, we then try to identify the effect of several potential explanatory variables of the price-discovery metrics obtained in the previous step (Section 6.2).

6.1. A dynamic price-discovery metric

To estimate a time-variant price discovery metric we extend the Gonzalo and Granger (1995) price-discovery analysis using rolling windows. The Gonzalo and Granger's model of price-discovery is based on the following vector error correction model (VECM) specification:

$$\begin{pmatrix} \Delta BSpr_t \\ \Delta CDSSpr_t \end{pmatrix} = \alpha(BSpr_{t-1} - \beta_2 - \beta_3 CDSSpr_{t-1}) + \begin{pmatrix} \sum_{i=1}^{p} \lambda_{1,i} \Delta BSpr_{t-i} \\ \sum_{i=1}^{p} \lambda_{2,i} \Delta BSSpr_{t-i} \end{pmatrix} + \begin{pmatrix} \sum_{i=1}^{p} \delta_{1,i} \Delta CDSpr_{t-i} \\ \sum_{i=1}^{p} \delta_{2,i} \Delta CDSSpr_{t-i} \end{pmatrix} + \begin{pmatrix} u_{1,t} \\ u_{2,t} \end{pmatrix} \quad (6)$$

The above empirical model is a vector autoregressive (VAR) system formed by two equations defined from the vector which includes the bond and CDS spreads of the same underlying country, denoted by $BSpr_t$ and $CDSSpr_t$, respectively, and an error correction term (ECT) which is defined by the expression $\alpha(BSpr_{t-1} - \beta_2 - \beta_3 CDSSpr_{t-1})$, where β_2 and β_3 are estimated in an auxiliary cointegration regression and the parameter vector $\alpha' = (\alpha_1, \alpha_2)$ contains the error-correction coefficients measuring each price's expected speed of adjustment. The estimation of the VECM specification is restricted to the existence of a cointegration relation between the bond and CDS spreads. This cointegration relation appears in the ECT as $(BSpr_{t-1} - \beta_2 - \beta_3 CDSSpr_{t-1})$. The parameters $\lambda_{1,i}$, $\lambda_{2,i}$, $\delta_{1,i}$, and $\delta_{2,i}$ for $i = 1,..p$, with p indicating the total number of lags, contain the coefficients of the VAR system that measure the effect of the lagged first difference in the CDS and bond spreads on the first difference of such spreads at time *t*.[12] Finally, u_t denotes a white noise vector.

The price discovery metric for the bond and CDS markets, denoted by GG_{bond} and GG_{CDS}, respectively, can then be constructed from the elements of the vector α', which contains the coefficients that determine each market's contribution to price discovery:

$$GG_{Bond} = \frac{\alpha_2}{-\alpha_1 + \alpha_2}; \qquad GG_{CDS} = \frac{-\alpha_1}{-\alpha_1 + \alpha_2}$$

Given that $GG_{Bond} + GG_{CDS} = 1$, we would conclude that the bond (CDS) market leads the price discovery process whenever GG_{Bond} is higher (lower) than 0.5. The intuition for this is the following. The larger the speed in eliminating the price difference from the long-term equilibrium attributable to a given market, the higher the corresponding α according to (6), and the higher is the price discovery metric.

In order to apply the methodology outline above to provide a dynamic metric of price-discovery leadership in the two markets at stake, we estimate the system in equation (6) using rolling windows with different lengths: 500, 750, and 1,000 days. To do so, we first need to check for the order of integration of the CDS and bond spreads and then for the existence of a cointegration relation. A first interesting feature is that, on average, the estimated metrics do not seem to be very sensitive to the window length. Based on this, in the remaining of the paper we focus on 1,000-day windows.[13] Using rolling windows with a length of 1,000 observations, we find that both CDS and the bond spreads are found to be non-stationary in all the countries and dates.[14] We next apply the cointegration test to a total of 990 1,000-day windows for each of the ten countries to find cointegration between both spreads in 6,295 cases (64% of the total). In particular, the country with the lowest (highest) percentage of cointegration relations is Finland

[12]The optimal number of lags is determined by means of the Schwarz information criteria.

[13] On the one hand, the 500-day windows enable us to update the influence of the new observations quicker than the 1,000-day windows and to estimate the price discovery metrics for the years 2006 and 2007. On the other hand, the 1,000-day windows enable us to consider a higher number of price-discovery metrics per day and country due to the existence of a higher percentage of cointegration relations for these windows. Actually, the years 2006 and 2007 present the lowest number of days with cointegration relations per year.

[14] The number of lags employed in the unit root test is chosen according to the Schwarz information criterion.

(Portugal) with cointegration in 41% (83%) of the windows. As we increase the window length, we find a higher number of cointegration relations and unit roots.

Figure 1 shows the estimated price-discovery metric for the 1,000-day windows for two groups of countries in the sample, peripheral and central.[15] In particular, we report a 30-day moving average of the mean price-discovery metrics which is obtained as an equally weighted mean across the ten euro area countries. An important message steaming from Figure 1 is that the price discovery metrics are not static but rather evolve over time, with the relative leadership of the CDS market in the process of price-discovery being more pronounced around some specific dates.

Specifically, before the Lehman Brothers collapse the CDS market leads sovereign risk price discovery. This finding is consistent with the results reported by, e.g., Blanco, Brennan and Marsh (2005), or Zhu (2006) in the context of the corporate debt markets. The first noticeable rise in the relative leadership of the bond market took place around February 2008, around the collapse of Bear Stearns. Afterwards, in September 2008, coinciding with the fall of Lehman Brothers and AIG the bond price discovery metric again jumps to reach its highest value at the end of 2008. This pattern suggests that during these two specific episodes, Bear Stearns and Lehman-AIG, the bond spread led, although by a small margin, the price discovery process. The price discovery metrics show the last significant rebound around the end of June 2011, some days before the proposal for agreement made by some private banks in July 2011 to accept a loss of 21% on their Greek bonds under the implicit presumption that such a voluntary deal would not trigger a credit event that would call for the activation of the CDS protection. Following this, in August 2011 the bond market became the leader in terms of price discovery confirming the role this market during the most recent phase of the crisis as the fairest source of information of credit risk.[16]

From a different perspective, we observe a decoupling between the price-discovery measure for the core and the peripheral countries from the end of 2008 until mid-2011. Starting from the end of 2008, it is worth noting that for most of the time the relative

[15] The peripheral group includes Ireland, Italy, Greece, Portugal and Spain. The core group includes Austria, Belgium, Finland, France and The Netherlands. The window length in all cases is 500 days.
[16] This result is in line with the role of the bond market as an appropriate credit risk measure in crisis periods found by Mayordomo, Peña, and Romo (2011b) for the corporate case.

efficiency of the CDS market in the peripheral economies is significantly higher than for the core countries, where the bonds market indeed leads the price discovery process for most part of 2009. The difference between the estimated price discovery metric for both groups of countries widens further during 2010. This pattern, which is mainly motivated by a sharp increase in the relative efficiency of the CDS market, may be reflecting the ECB policy of buying sovereign debt issued by peripheral countries. Another potential explanation is the use of the core countries sovereign debt as a reserve asset during periods of heightened aggregate uncertainty during which investors' perception of risks change dramatically in a way that may not be easily reconciled with risk attitudes observed in normal times.[17]

This core-peripheral countries decoupling persists until May 2011 when we observe a similar trend in both groups of countries improving generally the power of the bond market to lead price discovery. In particular, after July 2011 the bond market moves ahead the CDS market. This result could be due to the increase of overall risk perception but also to a lack of confidence of institutional investors in the CDS market after the aforementioned proposal for an agreement of banks to accept a voluntary haircut on their Greek bonds without activation of the CDS contracts. This explanation would fit well with the fact that the price discovery metric for Greece jumps from a value near zero in April 2011 to one in July 2011.

< Insert Figure 1 here >

Figure 2 shows the estimated price-discovery metrics for the 1,000-day windows for the same two groups of countries (peripheral and core) but showing the number of countries for which we can implement the Gonzalo and Granger's methodology (i.e. unit roots in both the CDS and bond spreads series are found and both spreads are cointegrated). The darker the line is, the higher is the number of countries for which the analysis can be implemented. We observe that for most of the time, the analysis can be implemented on four or more countries. Actually, it can be implemented on eight or more countries for some rather long time-windows from the beginning of 2008 to the end of 2009.

[17] Caballero and Krishnamurthy (2008) provide a theoretical framework that exploits the notion of Knightian uncertainty to rationalize, among other phenomena, the intense accumulation of high-quality assets during episodes of macroeconomic instability and Caballero and Kurlat (2009) illustrate that idea within the context of the Great Recession.

< Insert Figure 2 here >

The Granger-causality test has also been commonly employed in some previous price-discovery analyses. This methodology requires the credit spreads to be integrated of order one and for this reason it can be applied to longer time-windows than the Gonzalo and Granger (1995) test due to the absence of the cointegration requirements in the former method. Hence, we repeat the analysis using a rolling windows Granger causality test in which the optimal lag length of the VAR is obtained by means of the Schwarz information criteria. In general, we find that before 2008 none of the credit spreads causes the other. Afterwards we find a bidirectional causation relation in the peripheral countries. Thus, attending to the Granger causality test we cannot clearly differentiate whether the bond or the CDS market lead the process of price discovery in the peripheral countries. In the core countries we observe that the CDS causes the bond spread while there is no causation effect from the bond spread from the beginning of 2008 to the end of the sample.

The Granger causality test is the approach followed by Delis and Mylonidis (2010) to estimate the dynamic price discovery process. Our results are consistent with the ones that the previous authors found for the countries they study (Greece, Italy, Portugal, and Spain): a bidirectional causation of credit spreads in the peripheral countries after 2008. However, these results confirm that Granger causality test is not so informative to determine which market leads the price discovery process, as it is the case in the peripheral countries, while Gonzalo and Granger's methodology enables us to conclude which market leads at every date.

6.2. An analysis of the determinants of market leadership in price-discovery

In this section we aim at shedding some light on the dynamic pattern of the price-discovery metrics estimated before, by regressing them on some potential explanatory factors.

Specifically, for each country, we construct a dummy variable that takes a value of 1 when the bond market reflects information more efficiently than the CDS market and 0 otherwise. This dummy is constructed on the basis of a rolling window estimation using

1,000 observations,[18] and then it is used as the dependent variable in a Logit regression that includes as regressors the same used in the regression contained in Table 2, with the exception of the lagged basis and the relative spillovers measure. This last variable is not included due to potential endogeneity problems. We consider these regressors because they have been found to have a significant effect on the deviations from a zero-basis. Hence, as this shows that the effect of such regressors is not reflected in the same way in the two markets, it seems natural to consider that one market could capture better than the other the effect of each determinant of the basis. Additionally, as argued before, the role of the two markets providing efficient information on credit risk could have been affected by the agreement of banks to accept losses on their Greek bonds. For this reason, we also use as an additional regressor a dummy variable that is equal to one after the agreement of banks the 21st of July 2011 to accept 21% losses on their Greek bonds and zero before that date.

The results are reported in Table 4. Columns 1 and 3 report the results obtained when we use daily and monthly price discovery metrics, respectively, for the period which spans from December 2007 – October 2011. Column 2 contains the marginal effects of the coefficients reported in Column 1. The marginal effects of the dummy variables are obtained as the discrete change (from 0 to 1) in the corresponding variable. All the explanatory variables, including the dependent variable, are expressed in percentages with the exception of the bonds purchased by the ECB in a given week, which is defined in billions of Euros, and the dummy haircut agreement by banks, which is equal to one after the 21st of July 2011 and zero before then.

One might expect that the higher the counterparty risk is, the lower should be the power of the CDS market to reflect adequately the credit risk due to the lower quality of the protection sold in this market and, perhaps, the uncertainty around such quality. The sign is the expected (positive) and significant. An increase equal to one percentage point

[18] When faced with missing values (due to a lack of cointegration relation between the CDS and the bond spreads), the value of the dummy is imputed according to the Granger-causality analysis performed earlier. For the remaining missing values, the value of the dummy is imputed whenever there is another observed value within the next month which coincides with the previous price discovery value observed and persists at least in the next ten observations. If after a given date there are not cointegration relations up to the end of the sample we do not impute any value. For sake of the robustness of this procedure, we use the 1,000-observation estimation. We do not employ the price discovery metric directly, which is a concrete value comprised between 0 and 1, but instead assign a value one or zero to such metrics, thus softening such strong assumption. We impute 16% of the total observations.

in the counterparty risk measure would increase the contribution of the bond market to the leadership of the price discovery process by a 3.5%.

As argued before, funding costs affect more negatively the bond market relative to the CDS, as this last market allows for higher-leveraged positions. This could explain that the funding costs affect negatively to the ability of the bonds market for anticipating the price of credit risk relative to the CDS market. Specifically, we find that an increase of 1% in financing cost would decrease the estimated price-discovery metric by 23.6%. Somewhat surprisingly, the degree of liquidity in the CDS market relative to the bond market does not affect significantly to the price discovery metric.

In line with the results obtained by Mayordomo, Peña and Romo (2011b) the bond spreads tend to reflect credit risk more efficiently than CDS spreads during periods of high global risk (high values of the VIX Index). On the contrary, the country specific risk premium proxied by means of the absolute value of the returns of the national stock index is not significant at any standard significance level.

If the ECB's demand for debt relatively inelastic with respect to its price then the information embedded in the prices formed in that market could become less revealing of the fundamental value of the corresponding bonds. This hypothesis is confirmed by the significant and negative sign of the variable representing the total amount of sovereign debt purchased by the ECB. Contrary to the analysis of the determinants of the basis in which we employ the amount of debt purchased the previous week, we use the cumulated amount of debt purchased until the date given that we are using historical information on bond and CDS spreads to construct the price discovery metrics. An increase of $1 bn in the total amount of debt purchased by the ECB lead to a decrease of -0.3% of the ability of the bond market in terms of price discovery. The total amount purchased by the ECB at the end of the sample period (October 2011) was $180.5 bn. Such an increase could have contributed to a 54% fall in the estimated price-discovery metric, thus leading to a large deterioration of the degree of efficiency in the bond market.

The proposal for an agreement among private banks in July 2011 to accept a voluntary loss on their Greek bonds seemed to play in favor of the relative efficiency of the bond markets, according to the positive and significant effect of the corresponding dummy variable. Specifically, after July 21st 2011, we find a fall of 17.8% in the contribution of the CDS market to the price discovery process.

Finally, we note that the use of daily price discovery metrics obtained with 1,000-day rolling windows implies that the new information added in a given day is small relative to the information of the other 999 days which persists from one estimation to the next one. Faced with this, we check whether the results hold independently on whether we use daily or monthly price discovery metrics. The corresponding results are reported in Column 3. The comparison of the results reported in Columns 1 and 3 confirms that the results are robust to the use of daily or monthly metrics.

< Insert Table 4 here >

7. Conclusions

This paper analyzes the extent to which the sovereign Credit Default Swap (CDS) and bond markets reflect the same information on their prices in the context of the European Monetary Union. The main results can be summarized as follows.

We first test the "no-arbitrage" theoretical relation that should exist between the bond and the CDS spreads in a frictionless environment since both spreads are supposed to be the prices for the same credit risk. Our results show that after the subprime crisis there are persistent deviations from that theoretical parity relation that were absent before. In particular, we find evidence in favour of a persistent positive basis for the crisis period in a number of countries.

Based on the previous finding, we analyse the role of some potential determinants of the basis, including several sources of risk (counterparty, country-idiosyncratic and global), and market frictions. In particular, we find that the counterparty risk indicator has a negative and significant effect on the basis. Funding costs have a negative effect on the basis while a high liquidity in the bond market relative to the CDS market has a positive effect. Although the effect of the global risk variable is not significant, the country specific risk premium measured by means of the stock market index affects positively to the basis. In periods in which shocks originated in the bond market are transmitted into the CDS market, which we interpret as spillovers or contagion-effects, we find an overreaction-effect in this last market. The European Central Bank's purchases of sovereign debt in the secondary market initiated in May 2010 are found to affect positive and significantly to the basis.

Finally, we conduct a dynamic analysis of market leadership in the price discovery process. An important result here is that the price discovery process is clearly state-dependent. Specifically, the levels of counterparty and global risk, funding costs, the volume of debt purchases by the European Central Bank following the rescue of the Greek economy and the subsequent proposal for an agreement among private banks to voluntarily accept losses of their Greek bonds are all found to be significant factors in determining which market leads price discovery.

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Appendix A.1

Estimation of the spillovers between the CDS and bonds markets

We use a notion of spillover effects according to which such effects are defined from a variance decomposition associated with an N-variable vector auto regression following the methodology employed by Diebold and Yilmaz (2010) to measuring directional spillovers in a generalized VAR framework that eliminates the possible dependencies of results on ordering. The spillovers between the CDS and bond spreads here estimated can be interpreted as the degree of variation in the changes of the CDS (bond) spreads that is not attributable to their historical information but to contemporary shocks (innovations) in the changes of the bond (CDS) spreads. This indicator of contagion takes higher values as the intensity of the contagion effect which is caused by the specific shocks of the bond (CDS) market increases. In the extreme case in which there is no contagion from the bond to the CDS market the indicator series is equal to zero.

In particular, we first consider a covariance stationary *N-variable VAR (p)*:

$$X_t = \sum_{i=1}^{p} \Phi_i X_{t-i} + \varepsilon_t \qquad (1)$$

where X_t denotes a vector of stationary changes in the CDS and bond spreads of a given country and $\varepsilon \sim (0, \Sigma)$ is a vector of independently and identically distributed disturbances such that the moving average representation is $X_t = \sum_{i=0}^{\infty} A_i \varepsilon_{t-1}$, where the *NxN* coefficient matrices A_i obey the recursion $A_i = \Phi_1 A_{i-1} + \Phi_2 A_{i-2} + ... + \Phi_p A_{i-p}$, with A_0 being an *NxN* identity matrix and $A_i=0$ for $i<0$. Thus, the error from the forecast of X_t at the *H-step-ahead* horizon, conditional on information available at *t-1*, can be expressed as $\xi_{t,H} = \sum_{h=0}^{H} A_h \varepsilon_{t+H-h}$, and the variance covariance matrix of the total forecasting error is computed as $Cov(\xi_{t,H}) = \sum_{h=0}^{H} A_h \Sigma A_h'$, where Σ is the variance-covariance matrix of the error term in equation (1), ε_t.

The moving average coefficients are the key to understanding the dynamics of the system. We rely on variance decompositions, which allow us to parse the forecast error variances of each variable into parts attributable to the various system shocks. By means of this variance decomposition we can obtain the proportion of the *H-step-ahead* error variance in forecasting X_i that is due to shocks to X_j, $\forall j \neq i$, for each *i.*

We first compute the variance shares which are defined as the fractions of the *H-step-ahead* error variances in forecasting X_i due to shocks to X_i, for $i= 1, 2,\ldots, N$. we then derive the cross variance shares, or spillovers, defined as the fractions of the *H-step-ahead* error variances in forecasting X_i due to shocks to X_j, for $i, j = 1, 2, \ldots, N$ such that $i \neq j$. The *H-step-ahead* forecast error variance decompositions are denoted by $\theta_{ij}^{g}(H)$, for $H = 1, 2, \ldots$, i.e.:

$$\theta_{ij}^{g}(H) = \frac{\sigma_{ii}^{-1} \sum_{h=0}^{H-1} \left(e_i^{'} A_h \Sigma e_j \right)^2}{\sum_{h=0}^{H-1} \left(e_i^{'} A_h \Sigma A_h^{'} e_i \right)} \qquad (2)$$

where Σ is the variance matrix for the error vector ε, σ_{ii} is the standard deviation of the error term for the *i-th* equation, and e_i is the selection vector with one as the *ith* element and zeros otherwise. The sum of the elements of each row of the variance decomposition table is not equal to 1, i.e. $\sum_{j=1}^{N} \theta_{ij}^{g}(H) \neq 1$. Each entry of the variance decomposition matrix can be normalized such that the elements of each row sum 1 as:

$$\tilde{\theta}_{ij}^{g}(H) = \frac{\theta_{ij}^{g}(H)}{\sum_{j=1}^{N} \theta_{ij}^{g}(H)} \qquad (3)$$

We compute the spillovers from shocks in the first differences in CDS and bond spreads. That is, we estimate spillovers between the bond and CDS spreads for a given country. We use first differences instead of percentage changes to minimize the effect of the outliers which are obtained when we use in the denominator the bond spread which is very close to zero, and even negative, in some countries at some points in time in which the bond yield does not differ materially from the German bond yield.[19]

[19] Spillovers can be computed from either shocks in the mean or the volatility. Diebold and Yilmaz (2010) estimate spillovers in volatility using as a measure for such volatility daily high and low prices. We have end-of-day CDS and bond spreads and for this reason, the only possibility to compute volatility is as the square of the measure employed to compute the mean spillovers (changes in credit spreads). We find that the variable refereed to shock spillovers in the mean is highly correlated with the equivalent measure for the variance.

Table 1: CDS and Bond Spreads Descriptive Statistics

Table 1 reports the CDS and bond spreads main descriptive statistics (mean, and standard deviation) for different time periods (2004 - 2008, 2009, 2010, and 2011). The bond spreads are obtained as the difference between the country A's yield and the German yield.

			Bond	CDS
Austria	2004-2008	Mean	7	8
		Std.Dev.	12	21
	2009	Mean	52	104
		Std.Dev.	25	49
	2010	Mean	42	79
		Std.Dev.	12	13
	2011	Mean	59	90
		Std.Dev.	14	33
Belgium	2004-2008	Mean	9	8
		Std.Dev.	17	14
	2009	Mean	49	63
		Std.Dev.	29	33
	2010	Mean	57	110
		Std.Dev.	27	43
	2011	Mean	136	187
		Std.Dev.	47	51
Finland	2004-2008	Mean	5	6
		Std.Dev.	12	9
	2009	Mean	34	37
		Std.Dev.	21	19
	2010	Mean	-6	31
		Std.Dev.	13	3
	2011	Mean	18	44
		Std.Dev.	26	18
France	2004-2008	Mean	4	6
		Std.Dev.	9	9
	2009	Mean	23	40
		Std.Dev.	12	20
	2010	Mean	24	70
		Std.Dev.	8	18
	2011	Mean	41	107
		Std.Dev.	18	39
Germany	2004-2008	Mean		5
		Std.Dev.		7
	2009	Mean		36
		Std.Dev.		18
	2010	Mean		40
		Std.Dev.		8
	2011	Mean		58
		Std.Dev.		20

Table 1 (Cont.): CDS and Bond Spreads Descriptive Statistics

Greece	2004-2008	Mean	25	23
		Std.Dev.	36	36
	2009	Mean	166	165
		Std.Dev.	78	54
	2010	Mean	779	682
		Std.Dev.	290	242
	2011	Mean	1644	2075
		Std.Dev.	506	1452
The Netherlands	2004-2008	Mean	4	6
		Std.Dev.	8	12
	2009	Mean	30	53
		Std.Dev.	18	30
	2010	Mean	19	45
		Std.Dev.	7	8
	2011	Mean	24	55
		Std.Dev.	16	22
Ireland	2004-2008	Mean	9	13
		Std.Dev.	20	31
	2009	Mean	151	190
		Std.Dev.	56	62
	2010	Mean	262	302
		Std.Dev.	164	154
	2011	Mean	821	730
		Std.Dev.	235	145
Italy	2004-2008	Mean	19	19
		Std.Dev.	25	27
	2009	Mean	74	103
		Std.Dev.	32	39
	2010	Mean	109	165
		Std.Dev.	40	42
	2011	Mean	218	246
		Std.Dev.	101	115
Portugal	2004-2008	Mean	13	14
		Std.Dev.	20	19
	2009	Mean	71	76
		Std.Dev.	37	27
	2010	Mean	251	293
		Std.Dev.	111	116
	2011	Mean	918	767
		Std.Dev.	366	273
Spain	2004-2008	Mean	8	12
		Std.Dev.	15	20
	2009	Mean	54	89
		Std.Dev.	30	26
	2010	Mean	152	205
		Std.Dev.	74	67
	2011	Mean	257	295
		Std.Dev.	66	64

Table 2: Statistical Arbitrage Test for the Existence of Persistent Mispricings

This table reports the p-value obtained from the statistical arbitrage methodology of Mayordomo, Peña, and Romo (2011a). A p-value lower than 0.05 indicates that a significance level of 5% there are persistent mispricings between the 5-year CDS and bond spreads. The bond spread is obtained as the difference between the country A's bond yield and the risk-free rate which is equal to the German bond yield. Panels A and B report the results for the period ranging from January 2004 to September 2008 for CDS-bond negative and positive bases, respectively. Panels C and D report the results for the period which spans from the collapse of Lehman Brothers (September 2008) to October 2011 for CDS-bond negative and positive bases, respectively. *** (** and *) indicates the existence of persistent mispricings at a significance level of 1% (5% and 10%, respectively).

Panel A: Persistent Negative Basis Before Lehman Brothers Collapse

	P-value	Persistent Mispricing
Austria	1.000	No
Belgium	0.961	No
Finland	1.000	No
France	1.000	No
Greece	0.999	No
The Netherlands	0.988	No
Ireland	1.000	No
Italy	0.678	No
Portugal	1.000	No
Spain	0.988	No

Panel B: Persistent Positive Basis Before Lehman Brothers Collapse

	P-value	Persistent Mispricing
Austria	1.000	No
Belgium	0.957	No
Finland	1.000	No
France	1.000	No
Greece	1.000	No
The Netherlands	0.987	No
Ireland	0.706	No
Italy	0.378	No
Portugal	1.000	No
Spain	0.988	No

Table 2 (Cont.): Statistical Arbitrage Test for the Existence of Persistent Mispricings

Panel C: Persistent Negative Basis After Lehman Brothers Collapse

	P-value	Persistent Mispricing
Austria	1.000	No
Belgium	1.000	No
Finland	1.000	No
France	1.000	No
Greece	0.641	No
The Netherlands	1.000	No
Ireland	1.000	No
Italy	0.650	No
Portugal	0.734	No
Spain	0.753	No

Panel D: Persistent Positive Basis After Lehman Brothers Collapse

	P-value	Persistent Mispricing
Austria	0.016	Yes**
Belgium	1.000	No
Finland	0.963	No
France	0.043	Yes**
Greece	0.968	No
The Netherlands	0.003	Yes***
Ireland	0.047	Yes**
Italy	0.034	Yes**
Portugal	1.000	No
Spain	0.041	Yes**

Table 3: Determinants of the basis

This table reports the effect of the potential determinants of the basis based on a fixed effects regression robust to heteroskedasticity. Column (1) reports the effect of such determinants for the period which spans from January 2004 – October 2011. This column contains the explanatory variables' coefficients and the standard errors between brackets. *** (** and *) indicates whether the coefficients are significant at a significance level of 1% (5% and 10%). The bootstrap methodology is employed to correct any potential bias in the standard errors due to the use of generated regressors. Column 2 reports the standardized coefficient (i.e., the regression coefficient as in Column 1 multiplied by the standard deviation of the corresponding explanatory variable).

	(1)	(2)
Counterparty risk	-0.264***	-0.007
	(0.06)	
Ratio CDS/bond liquidity	0.003**	0.002
	(0.00)	
Financing costs	-1.325***	-0.007
	(0.34)	
Global risk (VIX) net of counterparty risk	-0.021	-0.001
	(0.02)	
Volatilty of country stock index returns	0.372***	0.004
	(0.12)	
Bonds purchased by ECB	0.001***	0.002
	(0.00)	
Shock spillovers from bond to CDS spreads relative to spillovers from CDS to bond spreads	0.005***	0.003
	(0.00)	
Lagged relative basis	0.938***	0.569
	(0.01)	
Constant	0.022***	
	(0.00)	
Number of observations	15,152	
Wald Chi2 statistic	46,315	
Prob>Wald Chi2	0	
Adjusted R-squared	0.913	

Table 4: Determinants of the Price Discovery Metrics

This table reports the effect of the potential determinants of the price discovery metrics using a panel fixed effects Logistic regression robust to heteroskedasticity. The standard errors are clustered by country.The price discovery metrics are obtained from the Gonzalo and Granger's (1995) methodology using rolling windows of 1,000 observations. The dependent variable takes value 1 when the bond spread reflects the information more efficiently than the CDS spread while a value equal to 0 indicates that the CDS spread leads the price discovery process. The bond spread is defined as the difference between the country A's yield and the German yield. Column (1) reports the results obtained when we use daily price discovery metrics for the period which spans from December 2007 – October 2011. Column (2) includes the marginal effects of the coefficients in Column (1). Column (3) reports the results obtained when we use monthly price discovery metrics for the same period. The sample length is due to the use of the first 1,000 observations to estimate the price discovery metric. The table contains the explanatory variables' coefficients and the standard errors between brackets. *** (** and *) indicates whether the coefficients are significant at a significance level of 1% (5% and 10%).

	(1)	(2)	(3)
Counterparty risk	0.287***	0.035	0.263***
	(0.07)		(0.10)
Ratio CDS/bond liquidity	0.000	0.000	-0.003
	(0.00)		(0.00)
Financing costs	-1.932***	-0.236	-1.897***
	(0.53)		(0.48)
Global risk (VIX) net of counterparty risk	0.050**	0.006	0.092**
	(0.02)		(0.04)
Volatility of country stock index returns	-0.001	0.000	-0.251
	(0.05)		(0.44)
Bonds purchased by ECB	-0.024***	-0.003	-0.012**
	(0.01)		(0.01)
Dummy haircut agreement by banks (July 11)	1.453**	0.178	2.452***
	(0.73)		(0.84)
Constant	-1.987***		-1.223
	(0.35)		(0.82)
Number of observations	8295		379
LR Chi2 statistic	3750		101
Prob>Wald Chi2	0		0
Log pseudolikelihood	-2888		-142
Pseudo R2	0.394		0.398

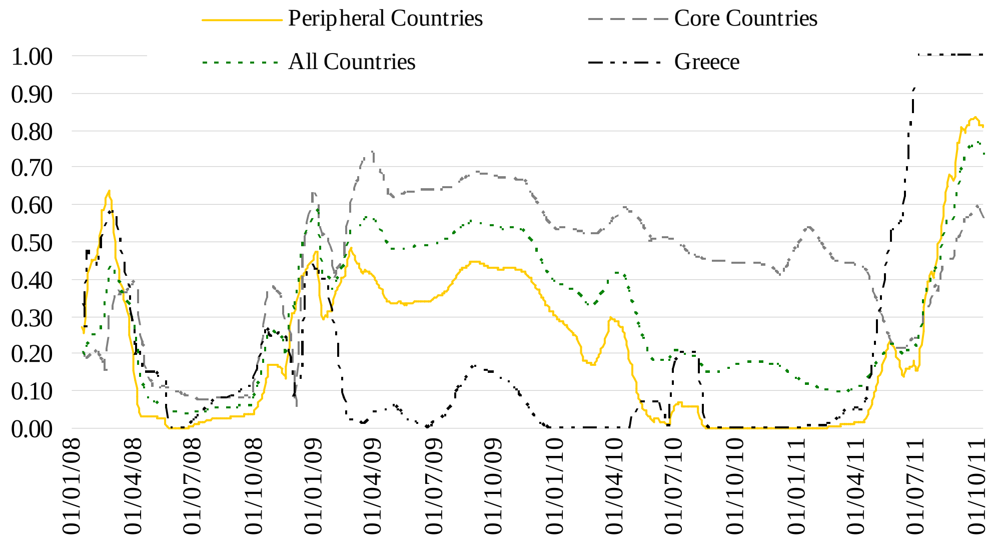

Figure 1: Price Discovery Metrics for Groups of EMU Countries with 1,000-day rolling windows. This figure shows the 30-days moving average of the price discovery metrics for the peripheral, core, and all the EMU countries. The price discovery metrics are estimated using 1,000-day rolling windows. These metrics for the groups of countries are obtained as the equally weighted average of the country specific price discovery metrics. The price discovery metrics for Greece correspond to the 30-days moving average of the metrics for this country.

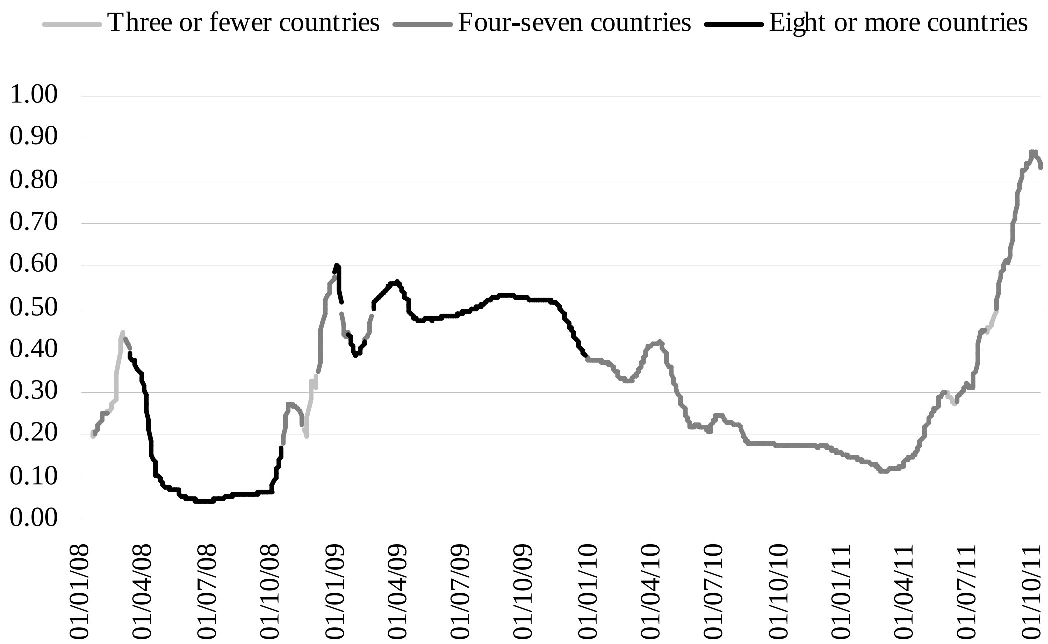

Figure 2: EMU Price Discovery Metrics and Number of Countries Employed in their Calculation: This figure shows the 30-days moving average of the EMU countries price discovery metrics that are obtained using 1,000-day rolling windows. These metrics are obtained as the equally weighted average of the country specific price discovery metrics. The line shows the number of countries that are employed to calculate the average metric such that the darker the line, the higher is the number of countries employed in its calculation.