"Pure" or "wake-up-call" contagion? Another look at the EMU sovereign debt crisis

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Abstract

We test whether the sharp increase in sovereign spreads of Euro area countries with respect to Germany observed after the burst of the Greek crisis is due to a deterioration of the macroeconomic and fiscal fundamentals, or to some form of financial contagion. We include in the analysis indicators of domestic and external imbalances which were mostly disregarded by previous studies, and distinguish between an increased attention devoted by investors to the variables which ultimately determine the creditworthiness of a sovereign borrower ("wake-up call" contagion) and behaviour disconnected from fundamentals ("pure" contagion). We find evidence of "wake-up call" contagion but not of "pure" contagion.

JEL codes: E62, H62, H63

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1 Introduction

At the beginning of 2009, when the 10th anniversary of the launch of the Euro was being celebrated across Europe, many commentators viewed the single currency as a major success (see, e.g. European Commission, 2009). As the introduction of the common currency approached, the convergence between interest rates towards the low levels of the most creditworthy Member States had been fast: in the period 1992-1998, the average spread in long-term government bond yields with respect to Germany declined from about 200 basis points to 24 points. From 1999 onwards spreads continued to narrow and at the end of 2007 yield differentials were negligible (16 basis points on average). Due to the financial turmoil triggered by the Lehman Brothers' bankruptcy in September 2008, some tensions started to surface, but at the end of that year the average yield spread in the Euro area was still at about 100 basis points. Strains on the government securities markets became worrisome only towards the end of 2009 (Fig. 1). Concerns mainly focused on Greece. After a series of upward revisions, the last of which equal to nearly 3 percentage points of GDP in October 2009, the Greek government estimated that the net borrowing amounted to 12.7 per cent of GDP in 2009, up from 7.7 per cent in 2008. The tensions originated in Greece spread to the government securities of other Euro area countries, notably Ireland , Portugal and, to a lesser extent, Spain and Italy. After three years since these events, some States have basically lost access to the bond market[1], and more generally sovereign debt strains in the Euro area are still worrysome and widespread, despite important progresses in correcting public finances made in the meanwhile by national governments.

The debate concerning the causes of the European sovereign debt crisis inflames both politics and the academia. While some argue that deteriorated public finances and fundamental macroeconomic weeknesses are at the root of the crisis, others claim that spreads are quite above the levels justified by fundamentals, and invoke forms of "market irrationality" and/or "contagion". The aim of the present paper is to assess the relative merits of the competing opinions through a formal econometric analysis.

Needless to say, the answer to this question has significant policy implications. Evidence of sizable and systematic mispricing of sovereign credit risk would imply that it is ill-advised to rely on markets as a means to induce fiscal and macroeconomic discipine in Member Countries. Furthermore, it would strenghten the case for interventions of european institutions (such as the European Stability Mechanism or the ECB) in the sovereign bond markets. Indeed, the Eurogroup Summit (June 29, 2012) has decided to use the EFSF/ESM instruments in order to "stabilise markets" for Member States respecting all their European commitments and timelines. Soon afterwards, the ECB decided to undertake Outright Monetary Transactions (OMT) in the secondary markets for sovereign bonds in the euro area "to address severe distortions which originate from, in particular, unfounded fears of the reversibility of the euro" (September 6, 2012).

While several other papers have studied the relationship between spread and fiscal fundamentals

[1] Greece applied for financial support in May 2010, followed by Ireland (November 2010) and Portugal (April 2011).

in EMU, ours contributes to the discussion in three ways.

First, a broader set of fundamentals is considered. Indeed, one of the lessons of the EMU crisis is that even countries with low levels of public debt and deficit can incur in a sudden deterioration of their fiscal position, for example as an effect of financial sector bailouts (which may transform private liabilities into public debt). This risk was considered obvious for emerging markets at least since the Asian crisis of the late nineties, but it was not taken into account by the EMU rules and – as we show in our paper – by investors. Our second contribution to the literature is to distinguish between different forms of contagion and measure their relative importance in explaining the post-crisis behaviour of European sovereign spreads.[2] Our third contribution is methodological in nature, as we apply to sovereign debt spreads for the first time panel methodologies designed to detect and to tackle non-stationarity and cointegration.

To give a preview of our results, we find that the burst of the Greek crisis had a systematic impact on the other Euro area countries' sovereign spreads. However, such impact has been different across borrowers. In particular, investors penalized more those governments with weaker fiscal and macroeconomic fundamentals.

The rest of the paper proceeds as follows: in Section 2, we provide a review of the literature and clarify our definition of contagion; in Section 3 we present our dataset; in Section 4, we discuss our empirical strategy, show our results, discuss several extensions and perform some robustness checks; in section 5, we provide numerical estimates of the long-run values of the spreads, derived from our empirical analysis; we finally, in Section 6 we draw some tentative conclusions and policy implications from our results.

2 Literature review

Several papers assess the determinants of sovereign spreads in EMU, starting from Codogno et al. (2003), among which Favero et al. (2010), Beber et al. (2010), Schuknecht et al. (2009, 2011); Attinasi et al. (2009); Sgherri and Zoli (2009); Hallerberg and Wolff (2008); Akitoby Statmann (2008). Typically, they explore the role of (a) country-specific factors influencing the risk of default, namely fiscal fundamentals and market liquidity and (b) common factors, such as the market appetite for risk. In particular, they bring to the data an empirical model such as:

$$s_{it} = \alpha_0 + \alpha_1 s_{it-1} + \beta_0 Z_{it} + \beta_1 F_t + \varepsilon_{it} \quad , \ |\alpha_1| < 1 \tag{1}$$

where Z_{it} is a vector of country-specific variables and F_t is a vector of variables which are common across countries. The above-mentioned papers differ with respect to the frequency at which the spread is measured (from daily to yearly), the included regressors, and the estimation method (in

[2] Of course, the two contributions are related: to understand whether spreads are excessive given fundamentals, is necessary to take a stance concerning the relevant fundamentals.

particular, some adopt a pooled cross-section/time-series approach, others provided country-specific estimates). Of course, high-frequency studies (such as Favero et al., 2010 and Beber et al. 2010) do not consider the role of fiscal and macro fundamentals, which are mostly constant over the period that they analyze.

Few papers consider the issue of contagion among sovereign securities within EMU. Some of them simply augment equation (1) with a further Z_{it} variable which captures developments in all the other EMU countries different from i. In particular, Caceres et al. (2010) employ a measure of "distress dependence", which is built extracting from the vector of CDS premia the unconditional marginal probability of default for each country. They then infer from those marginal distributions the joint default probability, and build and add-up the default probability of country i conditional on the default of the other countries. Similarly, Hondroyiannis et al. (2012) add a "contagion variable" defined as a weighted combination of other countries' spreads. Neither Caceres et al. (2010) nor Hondroyiannis et al. (2010) consider the more recent years.

Our contribution borrows instead from a different stream of the literature, which discusses contagion focusing on developing countries. In this literature, more precise and circumscribed definitions of contagion are used.[3] We follow in particular, Eichengreen et al. (1996), Masson (1998) and Goldstein et al. (2000), who distinguish between three kinds of circumstances:

- "Wake-up-call" contagion, i.e. a situation in which the crisis initially restricted to one country provides new information that prompts investors to reassess the default risk of other countries (this concept is used, for example, by Goldstein, 1998, Masson, 1999, Goldstein et al., 2000). In this case, domestic fundamentals justified a flight from sovereign debt even before the crisis event, but investors did not priced/perceived correctly the risk. The wake-up call hypothesis has been first put forward by Goldstein (1998) to explain contagion from Thailand (a relatively small and closed economy) to the other Asian countries in the Asian crisis of the late nineties. He argues that the other countries were affected by the same structural and institutional weaknesses as Thailand (crony capitalism, weak banking system, etc.), but investors ignored those weaknesses until the Thai "wake-up call". This kind of behaviour is also consistent with forms of "rational inattention" (Tutino, 2011 and Wielderholt, 2010). According to rational inattention theory, given the existence of costs in aquiring and elaborationg information, rational agents could in some circumstances optimally choose to ingnore some bits of information, for example concering fundamentals.

- Shift-contagion, which occurs when the normal cross-market channel intensifies after a crisis in one country. It can be seen as analogous to wake-up call contagion except for the fact that it is due to an increased sensitivity to common factors such as global risk aversion (the F_t term in 1) instead of country-specific factors. We borrow the term and the concept from the work of Forbes and Rigobon (e.g. Forbes and Rigobon, 2002).

[3] This literature is surveyed in Pericoli and Sbracia (2003), and Dungey et al. (2005).

- "Pure" contagion. This residual category covers any instance of contagion that is completely unrelated not only to the change in fundamentals (as it is the case for wake-up call and shift contagion) but also to the level of fundamentals, be they country-specific (as in the case of the wake-up call contagion) or global (as in the case of shift-contagion). It may be due to self-fulfilling (and therefore individually rational) loss of confidence (Calvo, 1988), to irrational herding behaviour (Chari and Kehoe, 2003), or to margin calls and other wealth effects for investors, triggered by capital losses in the country which originated the crisis (Kodres and Pritsker, 2002, Calvo and Mendoza, 2000, Schinasi and Smith, 2000).

In distinguishing between the three types of contagion, our contribution is similar to the recent paper by Bekaert et al. (2011). The authors use an international asset pricing framework with global and local factors to predict equity returns, defining unexplained increases in factor loadings as indicative of contagion, and find evidence of systematic contagion with its severity inversely related to the quality of countries' economic fundamentals and policies. They conclude that the "wake-up call" holds for equity markets, with markets and investors focussing substantially more on country-specific characteristics during the crisis.

We also see the approach pioneered by Gande and Parsley (2005) as very relevant and complementary to ours. They consider a sample of emerging countries and allow rating news from any one of them to influence the sovereign spreads in the others. In section 4.1 below, we also consider a unidirectional version of their methodology, substituting our crisis dummy with a variable summarizing Greek rating developments.[4]

Finally, let us remark that in this paper we are only concerned with contagion across sovereign bond markets, thus we leave aside the issue of contagion from the sovereign debt market to other financial markets or to the banking sector, which is addressed among others by Acharya et al, 2011, Alter and Schuler, 2011, Angeloni and Wolff, 2012.[5]

3 Data and descriptive statistics

We cover nine Euro area countries (Austria, Belgium, Finland, France, Ireland, Italy, Portugal, Spain and The Netherlands) using monthly data from January 2000 to December 2011. As it is customary in the literature, we exclude Greece (the "ground zero" country) from the analysis.[6]

Our dependent variable is the 10-year government bond yield spread with respect to the corresponding German Bund.[7]

[4] A couple of recent papers on the EMU use multi-equation econometric techniques, which can be seen as multi-equation extensions of Gande and Parsley (2005). Arezki et al. (2011) estimate a VAR allowing for the mutual interdependence of sovereign debt markets and the stock market. De Santis (2012) allows for a long-run co-integrating relationship between spreads and other variables.

[5] However, as we will see below we take into account the possibility that the state of banks may have an impact on sovereign spreads.

[6] We have verified that our main results do not change if Greece is included in the regressions.

[7] An often-used alternative would have been the Credit Default Swap (CDS) spread. However, for our purposes it

Concerning our independent variables, in our baseline specification we consider as common factor (the F_t variable in equation 1) the VIX, which is the most used indicator of the propensity of investors to bear the credit risk.[8] Concerning country-specific fundamentals (Z_{it} in equation 1), we include the general government debt, the private debt (defined as household plus non financial corporation debt), GDP growth and the current account surplus.[9] We also control for liquidity, measured by the bid-ask spread.[10]Given we are dealing only with advanced economies, we are not concerned with differences in debt characteristics such as inflation-indexation and currency denomination. Indeed, throughout the sample debt in all countries is almost completely nominal and issued in euros.[11]

The inclusion of the private debt and the current account balance, while non-standard in the literature on advanced economies (an exception is Gourinchas and Obstfeld, 2012), is frequent in studies concerning emerging economies, and has strong economic rationale. They are both indicators of an economy's leverage (respectively domestic and external). While a current account deficit does not mean per se a higher sovereign vulnerability, it is often associated with competitiveness imbalances and therefore worrisome macroeconomic developments. Furthermore, external capital inflows (which are the mirror image of the current account deficit) may trigger a boom in the non-tradable sector (particularly, the housing market), increasing the risk of a subsequent burst.[12] A similar reasoning can be applied to the private sector debt: if households and firms turn out to be unable to repay their debt, this might jeopardize public finances either because they are directly bailed-out by the government or – as it often happens – because the government bails-out the domestic banks that lent to households and firms in the first place. In any case, countries

suffers from several shortcomings. First, a well-developed CDS market exists only for few countries in our sample, and even for those countries data are available only for the more recent years. Second, CDS prices are not only driven by credit risk considerations but also by counteparty risk. Third, during the crisis in some countries CDS markets have been subject to policy interventions, such as short-selling restrictions, which are likely to have had an impact on CDS prices.

[8]The VIX, a shortname for the Chicago Board Options Exchange Market Volatility Index, is a measure of the implied volatility of the S&P 500 stock index; it is considered "a good indicator of the level of fear or greed in U.S. and global capital markets. When investors are fearful, the VIX level is significantly higher than normal. Market participants require additional compensation in the form of above-average excess returns for riskier assets." (see Antognelli et al., 2000). Data on government bond yields and VIX are taken from Thomson Financial Reuters. These data are released daily, and we compute monthly averages of them.

[9]Like we do with our left-hand variable, all independent variables are taken as a ratio to GDP and in difference with respect to the corresponding German variables.

[10]This measure for liquidity is common in the literature (see, among others, Codogno, Favero and Missale, 2003, and Favero, Pagano and von Thadden, 2010). Our variable is computed as the difference between the minimum bid yield and the maximum ask yield observed at daily frequencies for benchmark bonds; this computational method implies limited variability over time of this difference. Favero, Pagano and von Thadden (2010) use instead the best 5 bid and ask prices.

[11]As it is well known (see e.g. the contributions in Eichengreen and Hausmann, 2004) this is not true for emerging economies. Concerning debt duration, we observe in our sample moderate cross-countrydifferences, but they are basically time-invariant and therefore mostly captured by the country fixed effects.

[12]This would in turn induce sizable output gaps and revenue shortfalls, increasing public debts and jeopardizing its sustainability. This is the interpretation of the EMU crisis given by Spaventa and Giavazzi (2011) for some EMU countries.

with a very low initial public debt might increase it significantly and overnight. Interestingly, these variables are to be monitored at the European level within the context of the new Macroeconomic Imbalances Procedure (European Commission, 2012).[13]

Fiscal and macroeconomic variables are taken from the Eurostat quarterly database. These data are generally released with a delay of one quarter. Our monthly series are obtained keeping the value of the variable constant in each month of the quarter. In our specification we thus assume that spreads react contemporaneously to liquidity and volatility factors but they react with a 3-month lag to fiscal and macroeconomic variables. This also limits endogeneity problems, and thus concerns about possible reverse causation between the current spread and the left-hand variables.

In Table 1 we report some descriptive statistics of the variables used in our benchmark specification, distinguishing between two sub-periods (before and during the crisis). In the upper part of each panel we summarize the evolution of our dependent variable, i.e. the average yield spread, and the financial factors that in our specification are assumed to influence it. In the bottom part we summarize the development of our fiscal and macroeconomic fundamentals. All statistics refer to the sub-sample of countries excluding Germany and Greece.

The spread between the government bond yields in these nine Euro area countries and the German one increased on average from 19 basis points in the period before the crisis to 175 points from 2009-M10 onwards. The increase has been significantly larger in the sub-group of peripheral countries (Portugal, Ireland, Italy and Spain), from 25 to 330 points. Liquidity, measured by the bid-ask spread, worsened on average in the last part of our sample (on average the spread increased from 1.0 to 5.5). The evolution of the VIX indicates that global risk aversion has increased on average during the Euro area sovereign crisis; however, as considerable financial markets tensions already emerged following the Lehman Brothers' bankruptcy, the difference across sub-periods is not sizeable.

Turning to fundamentals, both fiscal and macroeconomic conditions deteriorated significantly during the sovereign crisis. Among domestic imbalances, the general government debt in our nine euro area countries increased on average by 17 percentage points of GDP (almost 30 in the peripheral countries); the increase in the private debt has been even larger (42 points in the overall sample and 57 in the peripheral countries). GDP slowed on average from 1.8 to 1.1 per cent, reflecting a negligible acceleration in the "virtuous" countries and a marked slow down in the others (from 2 to almost 0). Also the countries' external position worsened: on average the current account deficit increased from 0.5 to 0.7 per cent of GDP; with respect to Germany the deterioration has been greater (about 2.5 percentage points of GDP), reflecting strongly diverging competitiveness paths between Germany, on one side, and the other countries, on the other.

[13] Concerning external imbalances, the European scoreboard also includes the net investment position (the flow counterpart of the current account balance), the change in export market shares, the change in unit labour costs, and the change in the real effective exchange rate. Concerning internal imbalances, the scoreboard includes the private sector credit flow (the flow counterpart of domestic debt), the change in the house price index, and the unemployment rate.

4 Empirical analysis

In our analysis we use two alternative empirical models. The first one (section 4.1) is akin to equation (1), as it assumes that the spread is a stationary variable, even if it has an auto-regressive component. Stationarity is assumed by all the previous literature, and we provide estimates of this model for comparability's sake. However, as we will discuss below, there are good empirical reasons to question the stationarity hypothesis, and also to conjecture the existence of a long-run cointegrating relationship between the spread and the other covariates (section 4.2). Therefore, we will subsequently focus on the estimation of such a long-run relationship (section 4.3).

4.1 Stationary case

The empirical model. - We enrich the specification in (1) in order to take into account the three different kinds of contagion effects outlined in Section 2. We estimate the following model:

$$s_{it} = \alpha_{i0} + \alpha_1 s_{it-1} + \beta_0 Z_{it} + \beta_1 F_t + \gamma_0 D_t + \gamma_1 D_t s_{it-1} + \gamma_2 D_t Z_{it} + \gamma_3 D_t F_t + \varepsilon_{it},\ |\alpha_1|, |\alpha_1 + \gamma_1| < 1 \quad (2)$$

where the error term is assumed zero-mean, stationary and independent across countries (but we allow for heteroskedasticity and auto-correlation), and D_t is a dummy variable taking value one after the outbreak of the Greek crisis, that is, the revision of the official public finance figures by the new government in October 2009.

Therefore, γ_0 captures "pure contagion", coefficients γ_2 capture the "wake-up call" effect (a more pronounced post-crisis sensitivity to country-specific fundamentals), γ_3 captures shift-contagion (an increased sensitivity to common factors).

Notice also that in our specification we allow for country-specific fixed effects, to control for time-invariant unobserved characteristics. Indeed, the previous literature has pointed to some very slow moving features, which influence a sovereign's credit worthiness, such as the political system (Akitoby and Strassman, 2008) or debt intolerance (Quian et al. 2011). We also allow for a change in the auto-correlation coefficient in the post-crisis period (γ_1).

Baseline results. - The Least Square Dummy Variables (LSDV) estimates of equation (2) are shown in the first column of Table 2. In the pre-crisis period the only statistically significant coefficients are those of GDP growth and of the VIX index. Instead, in the post-crisis period the relationship becomes significant for all the fundamental variables, except for the private debt and the bid-ask spread. This suggests that a wake-up call effect exists for EMU countries. In particular, current account imbalances and public indebtness were not relevant in the pre-crisis period, whereas in the post-crisis period they are positively related to the sovereign spreads. On the contrary, the crisis dummy (γ_0) is not significant, suggesting that no "pure contagion" effect is present, and the same is true concerning "shift-contagion". Finally, the estimated auto-correlation parameter is relatively high (with no change in the coefficient after the Greek crisis), which points to possible

non stationarity.

Taking different proxies into account. - As a first robustness exercise, we consider two alternative measures of liquidity: the first is the share of long and medium term sovereign bond issuance in the Euro area (used by many authors, among which Attinasi et al. 2009); the second is the traded volumes of total government securities maturing at 9- to 11-years relative to Germany, which are available from Thomson Financial Reuters at a monthly frequency (also used by many, e.g. Codogno et al., 2003). In both cases, we found liquidity statistically insignificant (both alone and interacted with the crisis dummy).

Second, we use the funding ratio of banks, computed as the amount of loans to Euro area residents minus the amount of deposits (standardized by dividing the difference by the latter variable) as a proxy of banks' balance sheet fragility instead of private debt. However, also this measure turns out to be insignificant.

The definition of the contagious event. - A possible pitfall of our analysis is that it relies on a sharp hypothesis concerning the start of the EMU sovereign crisis. Moreover, our 0/1 dummy might also be seen as an inappropriate way to model the post-crisis development of the Greek fiscal situation. We address both problems by using, instead of our crisis dummy, a variable summarizing the Greek credit rating (we borrow this approach from Gande and Parsley, 2005, and De Santis, 2012). In particular, we transform the sovereign credit rating information (expressed in letters) of the three major credit rating agencies (Fitch, Standard & Poor's and Moody's) into a numerical variable using a linear scale. The variable takes 22 values from 1 (triple-A) to 22 (default). We also take into consideration the changes in the credit watch: a negative credit watch increases the value of the variable by 0.5 while a positive credit watch corresponds to a decrease of 0.5.[14] The results are analogous to our baseline regression (Table 2, column 2). In particular, the only fundamental variable which is statistically significant when taken in isolation is GDP growth. When interacted with the "Greek rating" variable, instead, also government debt and the current account surplus become significant, as it is the case in the baseline regression. In particular, the analysis makes clear that a worsening of the situation in Greece magnifies the positive effect of a current account surplus and the negative effect of public debt on the spreads of the other EMU countries. Finally, as in our baseline model, the Greek fiscal situation index, taken per se, has no effect on other countries' spreads.

Considering only the Periphery countries. - Another possible critique is that things could be different if one only considered peripheral euro area countries. First, for these countries it is more likely that investors' attention was already high before the crisis, given that their fiscal reputation was already undeniably worse. This reduces the probability of observing "wake-up call" contagion. Second, to the extent that investors consider those countries as more similar to Greece, this increases the probability of observing "pure" contagion. However, even restricting the sample to those countries (in particular, we focus on Portugal, Spain, Ireland and Italy) this is not the case.

[14] The value is an average of the results for the three main rating agencies.

On the contrary (Table 2, column 3) results are quite similar to the baseline estimation.

Bias-corrected estimates. - Since Nickell (1981), it is well known that the LSDV estimator is biased when used in dynamic panels. While the fact that bias is decreasing in the length of the panel should be reassuring given our very long sample period, we also experimented with the Kiviet (1995) estimation technique, which appears to be particularly appropriate for macroeconomic (i.e. big T/small N) panels (Judson and Owen, 1999). It turns out that the bias-corrected estimates are basically identical to our baseline.

4.2 Testing for unit roots and cointegration

A legitimate concern with the econometric analysis presented in section 4.1, given the observed relatively high persistence of the spreads, is that they could actually be non-stationary. Indeed, performing common panel unit root tests such as those by Levin Lin and Chu and by Pesaran, Im and Shin (see Banerjee, 1999 and Baltagi, 2008, ch. 12, Choi, 2006), we could not reject the null of integration (see Table 3). This result is robust even if we compute the relevant test statistics using different lag structures and different time spans. In particular, unit roots appear to be present not only if we look at the full sample, or at the post-crisis period, but also if we restict the analysis to the pre-crisis sub-sample.[15]

We also tested for the existence of a cointegrating relationship between the spread and its determinants. In particular, we adopted the residual-based approach by Kao and Pedroni (Banerjee, 1999 and Baltagi, 2008, ch. 12). While the results are in line with the existence of a cointegrating vector, they are not clear cut (Table 3).

4.3 Non stationary case

In this section, we model the long-run relationship between spreads and fundamentals as:

$$s_{it} = \alpha_{i0} + \beta_0 Z_{it} + \beta_1 F_t + \gamma_0 D_t + \gamma_1 D_t Z_{it} + \gamma_2 D_t F_t + \varepsilon_{it}, \tag{3}$$

therefore allowing for a structural change in the relationship in the post-Greek crisis part of the sample, and for the different kinds of contagion effects highlighted in the previous sections. As before, the error term is assumed independent across countries but possibly heteroskedastic and auto-correlated.

To estimate equation 3, we resort to different methods, in order to check the robustness of the results to different statistical assumptions.

First, we run a simple LSDV regression. Indeed if the spread is I(1) and there is no cointegrating relationship between the spreads and the fundamentals, i.e., ε_{it} in equation 3 is I(1), this estimator can be shown to deliver consistent estimates of the long-run average relationship between them,

[15] This suggests some caution in interpreting the results of previous papers, which did not consider the issue.

contrary to the pure time-series case (Phillips and Moon, 1999; see also Phillips and Moon, 2000, and Baltagi, 2008, Ch. 12).

Results are shown in the fourth column of Table 2 and are qualitatively similar to those obtained with the stationary model, but much more pronounced and clear cut. Before the crisis, all the fundamentals are significant with economically meaningful signs, except GDP growth (which is not significant) and the current account surplus (which has the "wrong" sign). After the crisis, the effect on the spread is magnified and with the expected for all the fundamentals. In particular, the effect of growth and of the current account surplus becomes significant and negative, as it should be if markets assess correctly sovereign creditwothiness.

If the spread is I(1) but there exists a cointegrating relationship between spreads and fundamentals (ε_{it} in equation 3 is I(0)), it can be shown that OLS estimates are inconsistent. We therefore estimate 3 using the panel dynamic least square (DOLS) estimator proposed by Kao and Chiang (2000) (which extend to panel data the approach of Saikkonen, 1991 and Stock and Watson, 1993).

That is, estimates of the coefficients of interest are found by running the following OLS regression:

$$s_{it} = \alpha_{i0} + \beta_0 Z_{it} + \beta F_t + \gamma_0 D_t + \gamma_2 D_t Z_{it} + \gamma_3 D_t F_t + \sum_{j=-2}^{2} \delta_{0j} \Delta Z_{it+j} + \sum_{j=-2}^{2} \delta_{1j} \Delta F_{t+j} + \varepsilon_{it}, \quad (4)$$

where the inclusion of ΔZ_{it+j} and ΔF_{t+j} among the regressors helps to get a consistent estimate of the βs and the γs.

The results, reported in Table 2 (col. 5) are remarkably similar to those of the previous exercise.

As a final exercise, we consider a model with random, instead of fixed individual effects. As shown by Baltagi et al. (2008, 2011), to this aim the best available option is to estimate equation (3) with feasible generalized least squares (this is true both if ε_{it} is I(0) and if ε_{it} is I(1)). Results, shown in Table 2 (col. 6) are qualitatively similar to those obtained with the fixed-effects specification.

5 Computing the long-run level of sovereign spreads.

Equation 3 can be rewritten applying the Oaxaca-Blinder decomposition to the crisis-induced change in spreads, as in Eichengreen and Mody (2000). That is, the difference between the pre-crisis and the post-crisis spread can be decomposed in two parts: one due to a change in the regressors, and the other due to a change in the coefficients. The change in the constant term is what we identify as the "pure" contagion effect. Indeed, conditional on the occurrence of the crisis, one gets:

$$\begin{aligned} E(s_{it}^{LR}|D_{it} &= 0) = \alpha_{0i} + \beta_0 E(Z_{it}|D_{it}=0) + \beta_1 E(F_t|D_{it}=0) \ , \\ E(s_{it}^{LR}|D_{it} &= 1) = \alpha_{0i} + \gamma_0 + (\beta_0+\gamma_1) E(Z_{it}|D_{it}=1) + (\beta_1+\gamma_2) E(F_t|D_{it}=1), \end{aligned}$$

where the LR superscripts serve as a reminder that we are considering here the long run equilibrium values of the spread. Therefore, the post-crisis long-run value of the spread is equal to:

$$\begin{aligned} E(s_{it}^{LR}|D_{it} \quad &= \quad 1) = E(s_{it}^{LR}|D_{it}=0) + \\ &\beta_0 \left[E(Z_{it}|D_{it}=1) - E(Z_{it}|D_{it}=0)\right] + \beta_1 \left[E(F_t|D_{it}=1) - E(F_t|D_{it}=0)\right] + \\ &+ \underbrace{\gamma_0}_{pure} + \underbrace{\gamma_2 E(Z_{it}|D_{it}=1)}_{wake-up\ call} + \underbrace{\gamma_3 E(F_t|D_{it}=1)}_{shift} . \end{aligned} \tag{5}$$

Terms in the second row capture the post-crisis change in fundamentals, while terms in the third row are those capturing the different kinds of contagion: γ_0 is what we call the "pure" contagion, and is unrelated to country characteristics; $\gamma_2 E(Z_{it}|D_{it}=1)$ captures "wake-up call" contagion, is country-specific and a function of fundamentals; $\gamma_3 E(F_t|D_{it}=1)$ is the "shift-contagion" component.

We use the estimates presented in section 4.3 to compute the various pieces of equation (5). In Table 4 we first consider, for each country, the estimated value of $E[s_{it}^{LR}|D_t=0]$ (column 1). We then add to this value the terms in the second line of equation 5 (column 2). To compute those values one needs to assess the pre- and post-crisis values of the fundamentals and of the VIX index. In the table, we put them equal to their respective sample counterparts. Finally, we add also the contagion terms, and therefore we get to $E[s_{it}^{LR}|D_t=1]$ (column 3).

According to our calculations, for most countries the spreads observed at the end of the sample (end-2011) are very close to their estimated long run levels. However, for two countries, namely Spain and Italy, they are quite above their equilibrium values (figure 2).

6 Conclusions and policy implications

The analyses presented in the paper suggest that, while macroeconomic indicators were largely ignored in pricing sovereign bonds before October 2009, after that date investors discriminated among sovereigns based on the quality of their fundamentals. In particular, countries with worse fiscal conditions and worse external positions recorded higher spread levels. In the terminology adopted in this paper, the sharp increase in spreads observed for some countries after the Greek crisis is the result of a "wake-up call", rather than of a "pure" form of contagion, as if the Greek crisis increased the convenience/willingness of investors to pay attention to the fundamentals of the other Euro Area countries.

Concerning the policy implications of our results, the fact that for some countries the current spread levels are above their long-run values supports the adoption of policy instruments meant to speed up the convergence of the spreads towards their long-run levels, such as those recently adopted by the European Stability Mechanism and the ECB's OMTs.[16]

[16] It is worthwhile to repeat that the estimated log-run levels of the spread incorporate the effects of the deterioration of fundamentals and the "wake-up" of the market.

However, it must be stressed that the absence of pure contagion, per se, does not settle the normative issue concerning the investors' ability to price sovereign bonds correctly. We cannot say, for example, whether the increased post-crisis sensitivity to fundamentals is "appropriate" or "excessive" (for example, it could reflect an increase in the perceived risk of the break-up of the euro area[17]). Answering this question would be important in implementing the OMTs. More broadly, it would help to settle the debate about the relative merits of market-based vs rule-based fiscal and macroeconomic discipline, which is as old as the very idea of the European Monetary Union. Indeed, already in 1989 the Delors' report worried that market forces "might be either too slow and weak or too sudden and disruptive". Further research on this issue, both theoretical and empirical, is warranted.

Another interesting and related question is whether in the future we will enter again a regime in which investors do not pay attention to fundamentals. To avoid disruptive cycles of excessive complacency and sudden wake-up calls, it seems advisable to push for market-friendly policies which highlight the fundamental imbalances of EMU countries even in good times, such as the timely dissemination of the Commission "scoreboards" and of the results of the Macroeconomic Imbalance Procedure.

[17] However, Di Cesare et al. (2012) point out that this risk begins to be perceived by investors in 2012, therefore after the end of our sample. Notice also that our coefficients capture an *average* long-run relationship (Phillips and Moon, 1999, 2000), which therefore should be not very affected by the perceived exchange rate risk: the appreciation of the new northern currency (or currencies) is likely of the same order of magnitude of the depreciation of the southern currency (or currencies).

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7 Figures and tables

Fig. 1 – Yield spreads between ten-year government bonds and the German Bund

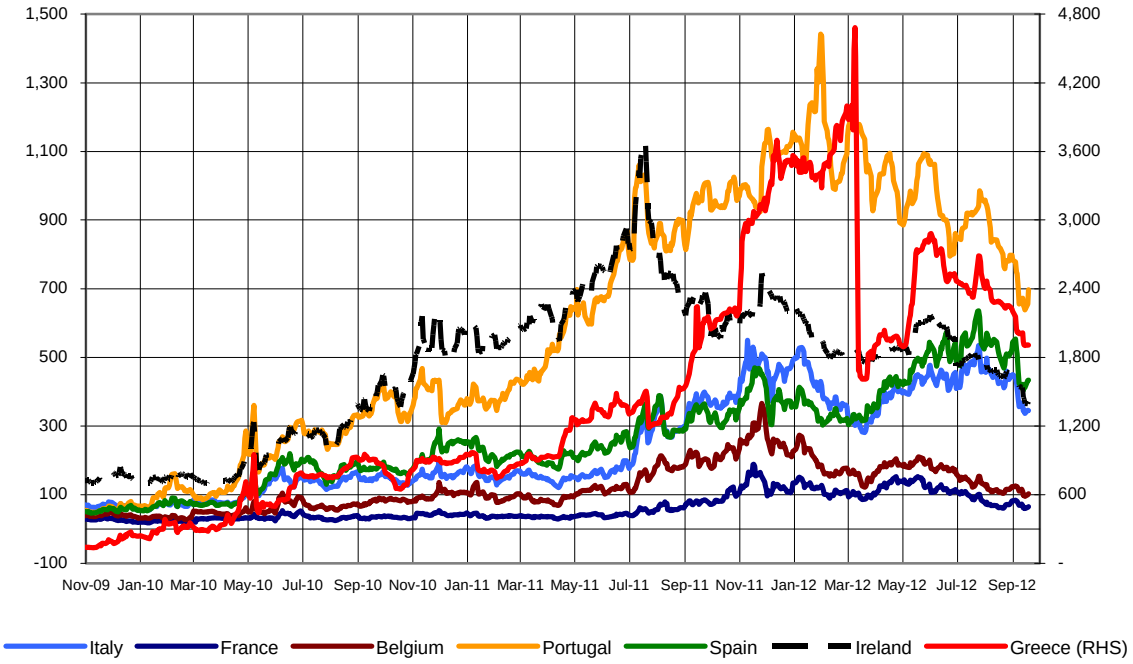

Table 1 – Descriptive statistics

	Mean	St.dev	Min	Max	Mean	St.dev	Min	Max
	January 2000 – October 2009				November 2009 – December 2011			
				Overall sample				
Sovereign spread (bp)	19.3	27.9	22.1	242.4	174.9	220.0	12.3	1109.3
Bid-Ask spread (bp)	1.0	0.8	0.2	6.0	5.5	14.8	0.3	85.4
Risk aversion (VIX)	25.9	10.9	12.6	63.3	28.0	7.1	20.1	45.6
Public debt/GDP×100	64.0	24.5	24.5	117.0	81.3	22.0	43.5	121.0
Private debt/GDP×100	162.0	42.8	75.2	303.1	204.4	49.3	125.3	303.4
GDP growth %	1.8	3.0	9.8	12.4	1.1	2.0	6.5	5.8
Current account surplus/GDP×100	0.5	5.5	13.3	11.9	0.7	5.0	13.3	11.7
				Ireland, Italy, Spain and Portugal				
Sovereign spread (bp)	25.7	35.6	22.1	242.4	328.1	253.5	52.7	1109.3
Bid-Ask spread (bp)	1.3	0.9	0.3	3.7	11.0	20.9	0.6	85.4
Risk aversion (VIX)	25.9	10.9	12.6	63.3	28.0	7.1	20.1	45.6
Public debt/GDP×100	63.2	28.8	24.5	117.0	92.1	22.2	53.9	121.0
Private debt/GDP×100	164.8	52.6	75.2	303.1	222.2	61.1	125.3	303.4
GDP growth %	2.0	3.3	8.3	12.4	0.1	1.5	5.5	2.2
Current account surplus/GDP×100	4.8	4.1	13.3	1.9	3.9	4.0	13.3	4.2
				Austria, Belgium, Finland, France and the Netherlands				
Sovereign spread (bp)	14.2	18.3	15.8	108.2	52.4	45.8	12.3	292.0
Bid-Ask spread (bp)	0.9	0.6	0.2	6.0	1.2	0.9	0.3	4.1
Risk aversion (VIX)	25.9	10.9	12.6	63.3	28	7.1	20.1	45.6
Public debt/GDP×100	64.7	20.4	29.9	115.6	72.7	17.7	43.5	100.0
Private debt/GDP×100	159.8	34.2	16.2	98.7	190.1	30.8	156.8	242.3
GDP growth %	1.7	2.6	9.8	6.4	1.9	1.9	6.5	5.8
Current account surplus/GDP×100	2.9	3.6	8.6	11.9	1.9	4.2	6.0	11.7

Table 2 – Regression results

	(1)		(2)		(3)		(4)		(5)		(6)	
Spread L1	0.927	***	0.947	***	0.930	***						
	(0.035)		(0.040)		(0.037)							
General government debt	-0.018		0.179		-0.088		1.211	***	1.120	***	0.337	***
	(0.116)		(0.112)		(0.147)		(0.295)		(0.258)		(0.0671)	
Private debt	0.050		0.087	**	0.043		0.926	***	0.939	***	0.167	***
	(0.040)		(0.043)		(0.031)		(0.077)		(0.080)		(0.039)	
GDP growth	-0.542	**	-1.172	**	-1.062	***	-0.077		-1.276		-2.341	***
	(0.27)		(0.516)		(0.408)		(0.639)		(0.783)		(0.825)	
Curr. acc.surplus	0.147		0.068		0.416		2.610	***	2.619	***	-0.351	
	(0.135)		(0.166)		(0.308)		(0.369)		(0.392)		(0.246)	
Bid-Ask	0.422		1.413		1.480	*	7.751	***	7.659	***	10.998	***
	(0.561)		(1.144)		(0.835)		(1.342)		(1.454)		(1.824)	
VIX	0.152	***	0.101		0.191	***	0.676	***	0.603	***	0.960	***
	(0.027)		(0.068)		(0.046)		(0.077)		(0.107)		(0.131)	
Dummy crisis	-15.128				-43.819		-84.738	***	-85.365	***	-95.619	***
	(10.377)				(35.894)		(25.716)		(23.346)		(15.467)	
*Spread L1 * crisis*	0.083				0.061							
	(0.052)				(0.073)							
*Public debt*crisis*	0.151	*			0.543	*	1.381	***	1.300	***	1.388	***
	(0.091)				(0.294)		(0.275)		(0.247)		(0.168)	
*Private debt*crisis*	0.044				0.139		0.337	**	0.293	**	0.649	***
	(0.047)				(0.115)		(0.138)		(0.121)		(0.080)	
*GDP growth*crisis*	-3.193				-7.274		-26.123	***	-21.603	***	-29.393	***
	(2.090)				(5.019)		(3.614)		(3.231)		(1.965)	
*Curr. acc.*crisis*	-0.871	*			-1.909		-4.597	***	-4.249	***	-5.282	***
	(0.524)				(1.333)		(1.219)		(1.124)		(0.673)	
*Bid-Ask*crisis*	-0.594				-1.657		0.065		0.064		-2.470	
	(0.769)				(1.018)		(1.507)		(1.463)		(1.840)	
*VIX*crisis*	0.198				0.192		2.174	**	2.204	***	2.007	***
	(0.345)				(0.893)		(0.882)		(0.825)		(0.462)	
Greek rating			-0.238									
			(1.626)									
*Public debt*Greek rating*			0.028	*								
			(0.016)									
*Private debt*Greek rating*			0.009									
			(0.011)									
*GDP growth*Greek rating*			-0.743									
			(0.484)									
*Curr.acc. surplus*Greek rating*			-0.126	*								
			(0.076)									
*Liquidy * Greek rating*			-0.132									
			(0.097)									
*VIX*Greek rating*			-0.010									
			(0.044)									
R2	0.98		0.98		0.98		0.87		0.89		0.85	
Observations	1269		1269		564		1269		1242		1269	

Columns 1,2,3,4: LSDV; Col. 5: DOLS (1 lead and 1 lag added for each variable; country dummies incl.); Col. 6: FGLS. All estimations except Col. 6: Huber-white robust standard errors in parentheses. All estimations except Col. 3: full sample (Col. 3: includes only the periphery countries: PT, IT, IR, ES). *: significant at the 10% level; ** at the 5%; *** at 1%.

Table 3 – Unit root and Cointegration Tests

Panel unit root tests	
*Levin, Lin and Chou t**	15.940
H0: unit roots for all i's(H1: no unit root)	(1,000)
Im, Pesaran and Shin W-stat	11.970
H0: unit roots for all i's (H1: some unit roots)	(1,000)
Panel cointegration tests	
ADF statistic (Pedroni 1)	-1.642
H0: no cointegration (H1 assumes common autocorr. coefficient)	(0,0503)
ADF statistic (Pedroni 2)	-1.170
H0: no cointegration (H1 allows country-specific autocorr. coefficients)	(0,121)

Notes: P-values in parentheses; number of lags =1.

Figure 2 – Cointegrated model: predicted values

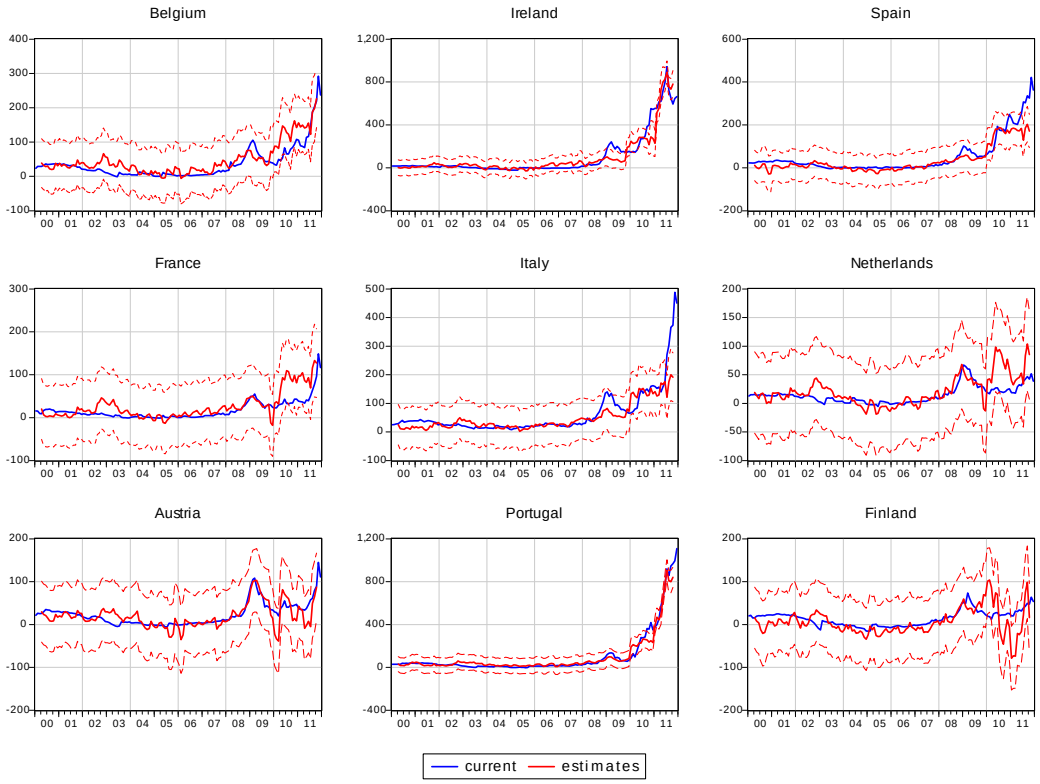

Table 4 – Long-run values of the spread

	Fitted values		
Coefficients	*pre-crisis*	*pre-crisis*	*post-crisis*
Fundamentals	*pre-crisis*	*post-crisis*	*post-crisis*
Italy	24	47	247
Austria	23	43	131
Belgium	21	45	210
Finland	0	16	81
France	7	32	175
Ireland	35	335	558
Portugal	46	257	507
Spain	28	97	269
Netherlands	10	35	134

Notes: Coefficients taken from Table 2, col. 5 (DOLS).