Nadia Linciano, Luca Giordano, Paola Soccorso

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Sovereign Risk Premia in the Euro Area and the Role of Contagion

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Nadia Linciano
CONSOB, Research Department

Luca Giordano
CONSOB, Research Department

Paola Soccorso
CONSOB, Research Department

Abstract

This work estimates a reduced model of the determinants of the 10-year yield spreads relative to Germany for 10 Eurozone countries. Results show that since the inception of the 2007 crisis, spreads have exhibited a rising time-dependent component. Country specific estimated responses to financial turmoil highlight three major results. Core countries have not been affected by financial contagion during the subprime crisis, and from 2011 onwards, they have benefited from government yield spreads that are lower than what is explainable by the underlying fundamentals. Peripheral member countries (except Italy) – which from the outset of the EMU benefited from underpricing of their economic and fiscal fragility due to the implicit bailout insurance – have suffered from a revision of market expectations since 2010. Italy, penalised by its historically high debt-to-GDP ratio, has been hit by a rising contagion effect since 2010, which is estimated to account for 180 b.p. of the spread observable in the 1st semester of 2012.

Keywords: Government Yield Spreads; Sovereign Risk Premia; Financial Crisis; Sovereign Debt Crisis; Contagion.
JEL Codes: G12; E43; H63.

1 Introduction

Since the beginning of 2010, when irregularities in Greece's budget were disclosed, a relentless rise in the spreads against the German Bund occurred for Greece, Ireland, and Portugal. Since July 2011, other non-core countries, such as Spain and Italy, have recorded a strong increase in bond yields, while core countries, such as Germany, have benefited from a flight-to-quality effect. Overall as the crisis developed, the observed pattern of spreads appeared to be more sensitive to changes in global conditions rather than to actual changes in the country-specific fiscal position.

A strand of the empirical literature on contagion shows that Euro Area countries were hit by a contagion phenomenon during the last international crisis (Gentile and Giordano, 2012; Metiu, 2012), while, according to other studies, contagion affected only a subset of countries during the sovereign crisis (Caporin *et al.*, 2013).

Corresponding author: Nadia Linciano, n.linciano@consob.it, CONSOB, Research Department, Via G.B. Martini 3, 00198 Roma, +39068477360,
Luca Giordano, l.giordano@consob.it,
Paola Soccorso, p.soccorso@consob.it.

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Contagion is defined as the transmission of shocks from one country to others or the cross-country correlation, not explained by a change in fundamentals or common shocks[1].

Contagion occurs when cross-country correlations increase during «crisis times» relative to correlations during «tranquil times». Indeed a constant high degree of co-movement in a crisis period would only point out that markets are interdependent (Forbes and Rigobon, 2001)[2].

This paper analyses the determinants of government yields in the Euro Area from January 2002 to May 2012. The aim is to disentangle the role of country-specific fundamentals, driven by fiscal and macroeconomic factors, from what is referred to as «contagion», bringing together the empirical literature on sovereign risk premia and contagion.

According to our results, the contagion component explains almost one third of the spreads' dynamic in 2009-2010 and almost 10% since 2011. However, results at the country level are quite different between core and peripherals. For core countries (excluding Germany, which is our benchmark) the analysis shows that model-predicted spreads are basically in line with fundamentals, though since the onset of the debt crisis some countries exhibit spreads lower than what is predicted by fundamentals. For example, in the first quarter of 2012, France showed spreads lower than what was implied by fundamentals by an amount ranging from roughly 50 to 90 basis points, depending on the model specification, while for the Netherlands such a «discount» was estimated to be as high as roughly 60 basis points. On the other hand, since 2009, spreads of peripheral countries are on average significantly higher than what is predicted by fundamentals due to a contagion effect; for most of these countries, contagion has a role comparable to fundamentals in explaining the level of the spreads in 2012. For example, contagion accounts for an amount ranging from roughly 170 to 240 basis points for Spain, while for Italy it explains roughly between 150 and 180 basis points of the spread, depending on the model specification.

To our knowledge, existing studies on the determinants of government bond yields assume that whatever is unexplained by the model is considered as contagion (*e.g.* irrational behaviour, herding effects, panic, etc.). Instead, the present study treats contagion as an additional explanatory variable which is estimated separately by using country individual effects and time dummy variables. In other words, residuals are regarded purely as the unexplained component of the fitted model rather than as contagion per se.

The work is organised as follows. The next section recalls some stylised facts of the sovereign debt crisis. Section 3 reviews the recent empirical literature on the determinants of yield spread in the Euro Area. Section 4 presents the sample, the model and the estimation results. Conclusions are drawn in the last section.

[1] For example, Masson (2004) defines contagion as meaning only «those transmissions of crises that cannot be identified with observed changes in macroeconomic fundamentals». Using a different terminology, Eichengreen *et al.* (1996), argue that there is contagion if the probability of a crisis in a given country increases conditionally on the occurrence of a crisis elsewhere, after controlling for the standard set of macroeconomic fundamentals. This definition is sometimes referred to as excess co-movement – a correlation that remains even after controlling for fundamentals and common shocks. Herding behaviour is usually said to be responsible for co-movement beyond that explained by fundamental linkages.

[2] Forbes and Rigobon (2002) argue that «contagion is a significant increase in cross-market co-movements after a shock». This definition is sometimes referred to as «shift-contagion».

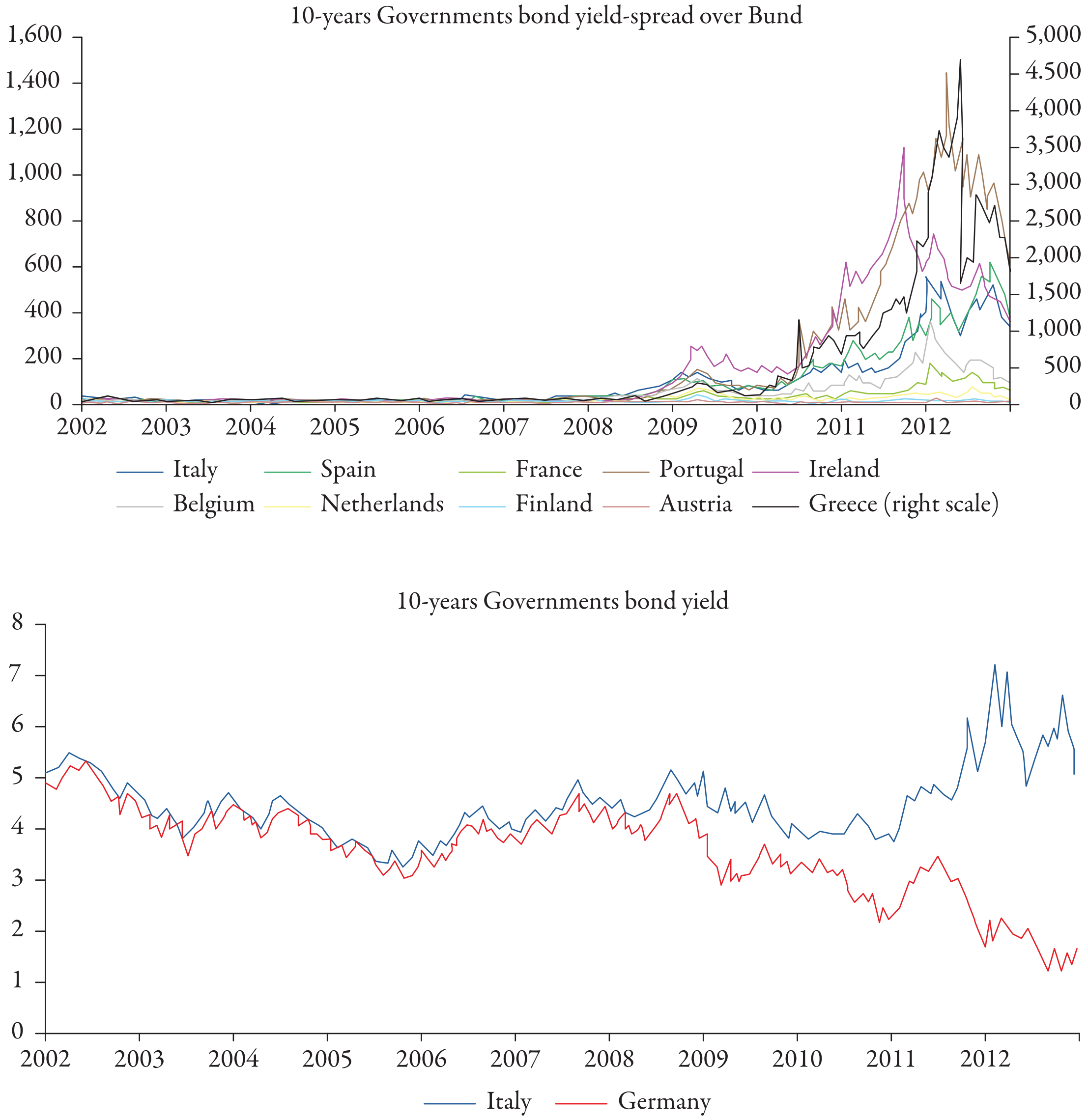

Figure 1: Ten year government bonds: yields and spreads relative to the German Bund for some Euro area countries (January 1st, 2002-September 14th, 2012).

Source: authors' calculations based on Thomson Reuters data.

2 The pattern of sovereign risk premia since the introduction of the euro

Sovereign risk premia for Eurozone countries have shown a strong convergence since the foundation of the Monetary union until January 2010[3]. From this point on, after the disclosure of the irregularities in Greek government budget accounting, Greek yields rose relentlessly followed by those of Ireland and Portugal. Since July 2011 other countries (namely Spain, Italy, and for a more limited time Belgium) have experienced a marked increase in their spreads relative to Germany (Figure 1).

[3] As documented by Pagano and Von Thadden (2004), the mean yield spread of the initial EMU participants over the German yield drop from «218 basis points in 1995 to 111 in 1996, 29 in 1997, 19 in 1998, and 20 in 1999». The downward trend resumed after 2002, following a slight rebound.

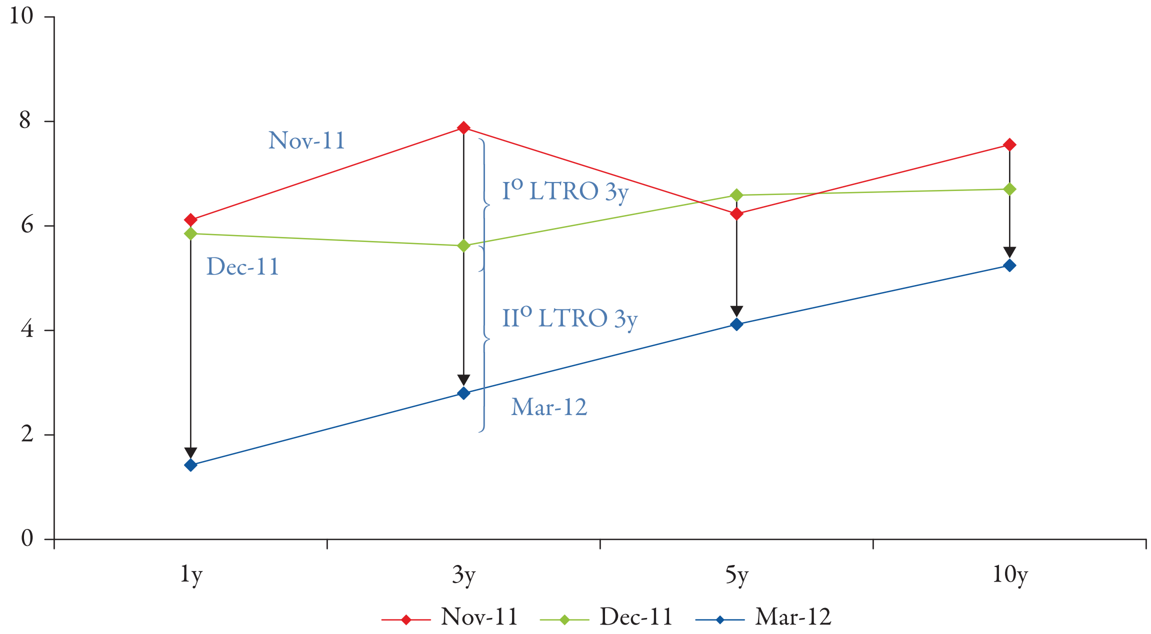

Figure 2: Yield curve of the Italian government bonds.

Source: authors' calculations based on Ministry of Finance and Bank of Italy data.

For the countries hit by the sovereign debt crisis the yield differentials relative to the German Bund declined in the first quarter of 2012, thanks to the successful private sector involvement in the Greek debt restructuring plan (which eased fears of a disorderly default by Greece), the fiscal adjustment and structural reforms undertaken by some Eurozone countries, and the actions carried out by the EU leaders to improve fiscal discipline and to contain the crisis. Moreover, the two long-term refinancing operations by the European Central Bank (ECB) – the first on December 26th 2011 for € 486 billion and the second on February 29th 2012 for € 530 billion – contributed to the decline in spreads. As for Italy, the government bond yield curve experienced a significant downward shift: in fact, following the ECB operations, net purchases of Italian bonds by domestic banks are estimated to have reached about € 80 billion (Figure 2)[4].

However, renewed tensions started to hit high debt countries at the beginning of April 2012, spurred by the developments of the Greek crisis, the difficulties experienced by the Spanish banking sectors, and the expectation of a negative growth rate for the Euro Area countries. These factors exacerbated the perception of the sovereign risk for peripheral countries. Spain, as well as Italy, recorded new pressures in government bond markets, while long term rates drop considerably for Germany, the Netherlands and France[5]. Such pressures eased again in September 2012, following the approval by the

[4] Source: Bank of Italy.

[5] For peripheral countries the increase in the spread relative to the German Bund was driven also by the decline of the yield of the German Bund itself (steadily lower than 2% since March 2012). Such a decline reflected both a flight to quality effect and investors' preferences for high rated government bonds, which led to lower refinancing costs with central counterparties and with the ECB.

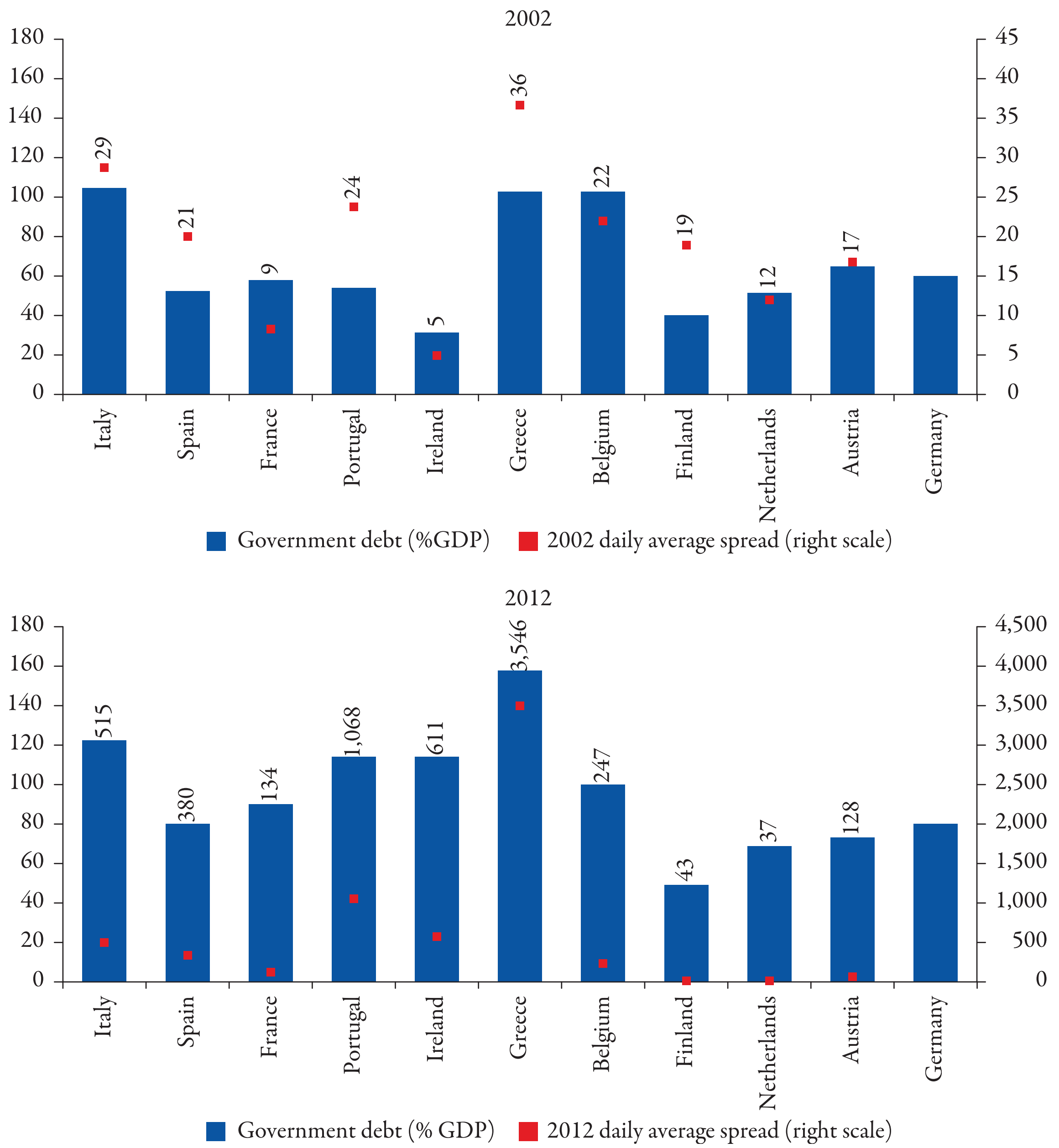

Figure 3: Ten year government bond yield spreads and public debt-to-GDP ratios for some Euro area countries (ten year government yield spreads are computed as averages of daily data; public debt-to-GDP ratios are end-of-period data; for 2012 the Spring economic forecast of the European Commission is considered).

Source: calculation on Thomson Reuters and European Commission data.

ECB of the «outright open market operation plan», contemplating unlimited buying of member States' bonds to drive down their borrowing costs.

Given these stylised facts, many researchers and practitioners have recently wondered to what extent the dramatic movements in government bond spreads that occurred in the Euro Area over the last few years are due to fundamental factors (as proxied by the countries fiscal position and other macroeconomic indicators) or rather to negative market sentiment (see the next section for a literature review). To this end, it is useful

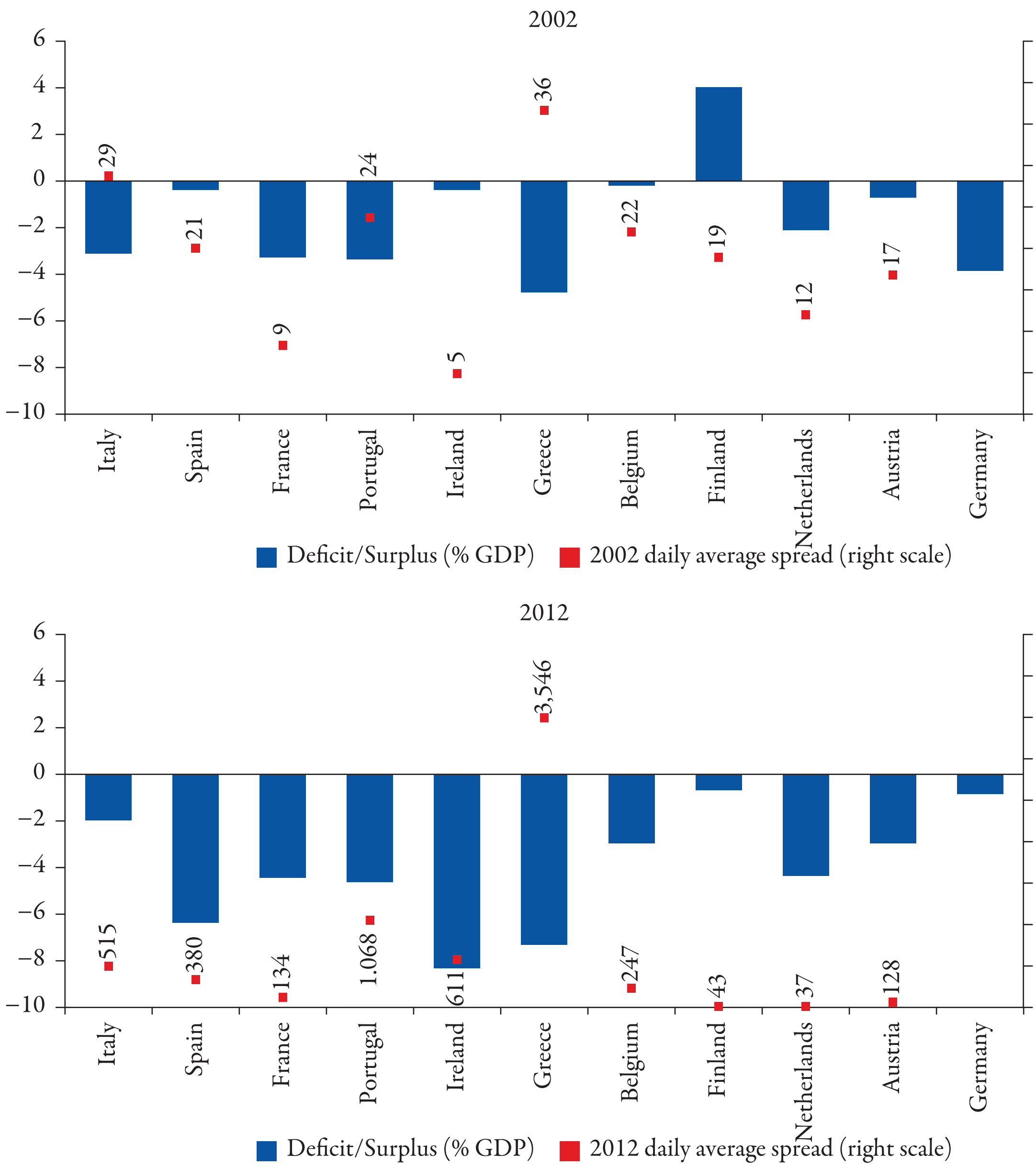

Figure 4: Ten year government bond yield spreads and deficit to GDP ratios for some Euro area countries (ten year government yield spreads are computed as averages of daily data; deficit to GDP ratios are end-of-period data; for 2012 the Spring economic forecast of the European Commission is considered).

Source: calculation on Thomson Reuters and European Commission data.

to look at the relationship between spreads and the countries specific default risk as proxied by debt-to-GDP ratios, deficit-to-GDP ratios and the fiscal space (*i.e.* debt-to-tax revenues ratio).

Figure 3 plots the yearly average spread and end-of-period public debt-to-GDP for major Euro Area countries in 2002 (left panel) and in the first half of 2012 (right panel). Over the time span considered all Euro Area countries, apart from Belgium, have experienced a sharp increase in the level of government debt relative to their GDP. This resulted mainly from the 2008 financial crisis, and the consequent government financed banking system

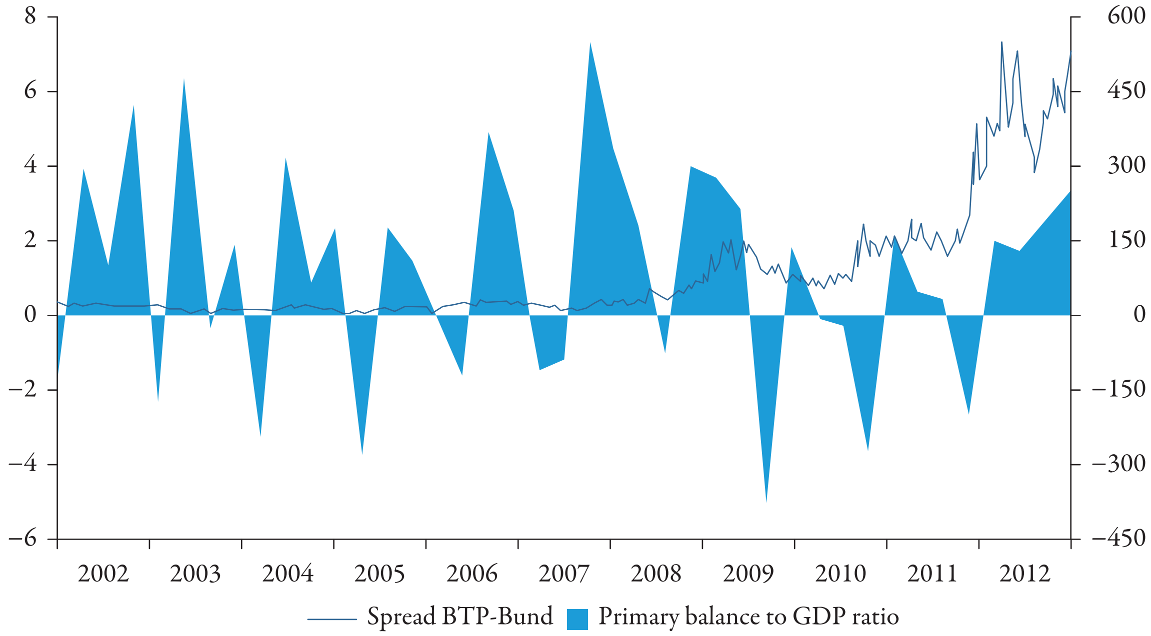

Figure 5: Italy: ten year government bond yield spread and primary balance to GDP ratio (January 1st, 2002-June 30th, 2012; data on daily spread in percentage points – right scale; quarterly primary deficit to GDP ratio – left scale; the Spring forecast of the European Commission is considered for 2012 annual value of primary deficit to GDP ratio).

Source: Thomson Reuters, ECB and European Commission.

rescue plans, and the recession that followed the financial crisis. Italy was less affected by the financial crisis and therefore recorded one of the lowest increases in debt-to-GDP ratio (roughly 18 percentage points, followed by Austria and Finland, whose ratio went up by 9 and 8 percentage points, respectively). Ireland, Greece and Portugal are at the opposite ends of the rankings (with debt-to-GDP ratio increases of around 84, 58 and 60 percentage points, respectively). As for the remaining countries, the least hit were the Netherlands (slightly more than 20) followed by Germany (almost 22), Spain (more than 28) and France (almost 32).

Another relevant indicator of fiscal fragility is the ratio of government deficit to GDP: Figure 4 shows this variable coupled with the yearly average spread at the end of 2002 (left panel) and in the first half of 2012 (right panel) for the 11 Euro Area countries considered. Apart from Italy and Germany, in 2012 all countries are expected to record a public deficit to GDP ratio that is higher than ten years before. In particular, Italy is expected to mark a deficit to GDP ratio equal to 2 percentage points (3.1% in 2012).

Figure 5 plots the evolution of the ratio of primary budget balance to GDP and of the sovereign bond spread for Italy.

Since 2010 the spread for Italian government bonds has shown a departure from the overall positive dynamics of the primary budget balance to GDP. In fact, Italy is penalised by its high stock of debt, which *ceteris paribus* requires larger primary surpluses to offset interest payments. Conversely, more virtuous Euro Area countries are able to run larger primary deficits or the same primary surpluses at a lesser cost. This clearly emanates

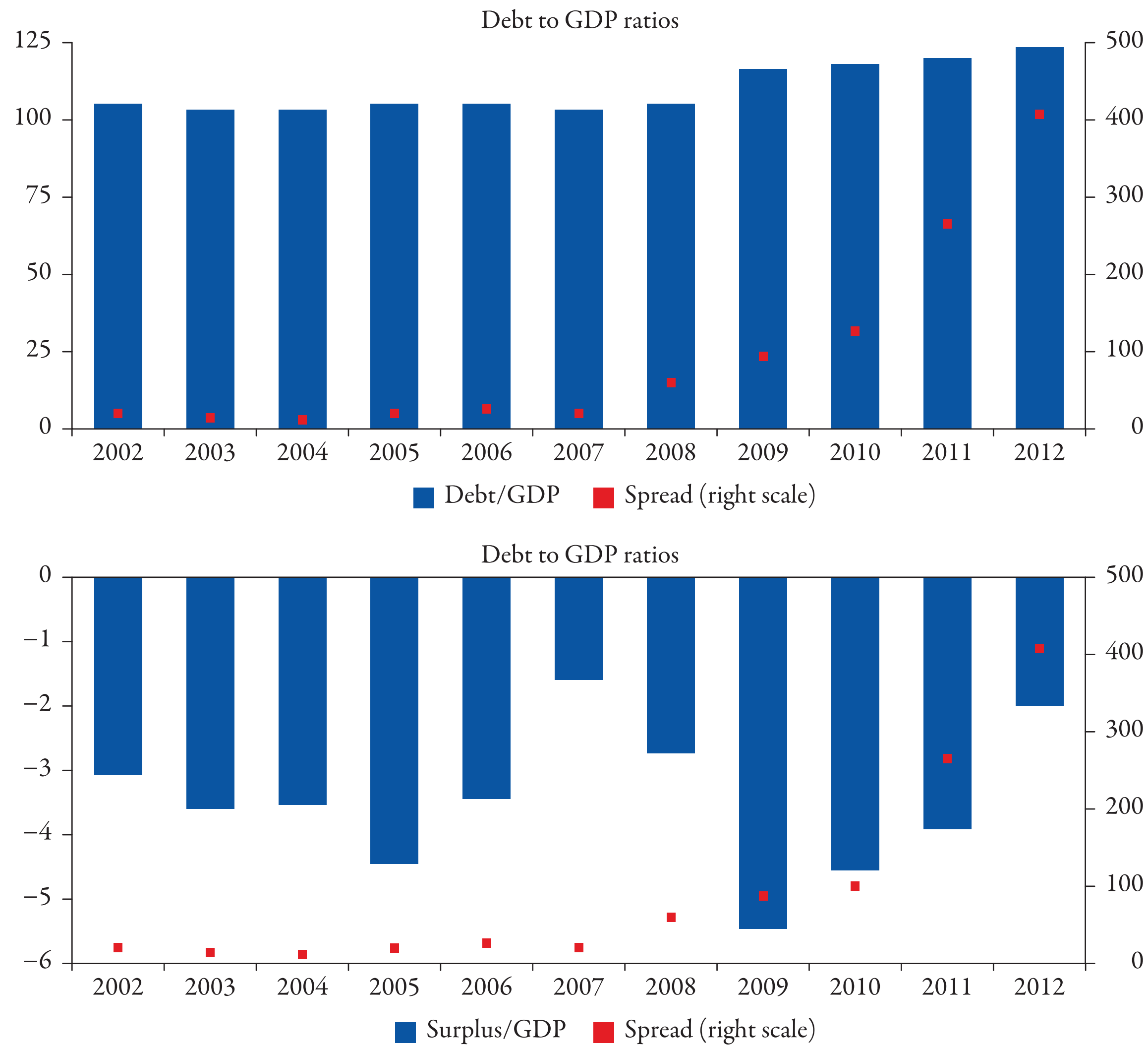

Figure 6: Italy: Ten year government bond yield spreads and fiscal fundamentals (ten year government yield spreads are computed as averages of daily data; public debt and deficit to GDP ratios are end-of-period data; for 2012 the Spring economic forecast of the European Commission is considered).

Source: Thomson Reuters and European Commission.

from the comparison between the fiscal position and spreads for Italy and France over the period January 2002 throughout June 2012 (Figures 6 and 7).

The inspection of Figure 6 allows us to draw two main considerations for Italy. Firstly, the relationship between debt-to-GDP ratio and the (average) sovereign spread, especially since 2008, shows a non-linear and convex pattern, implying that the impact on the spread of a one percentage point increase in the debt-to-GDP ratio is greater for higher levels of debt. This relationship is an empirical regularity, which generally holds for high debt countries (De Grauwe and Ji, 2012). Indeed, as public debt goes up the likelihood of a default grows too, thus leading investors in government bonds to demand a proportionally higher risk premium. Secondly, since 2010 the surge in the spread seems to be disconnected from the dynamics of the fiscal fundamentals; to a lesser extent this holds also for France and other non-core countries.

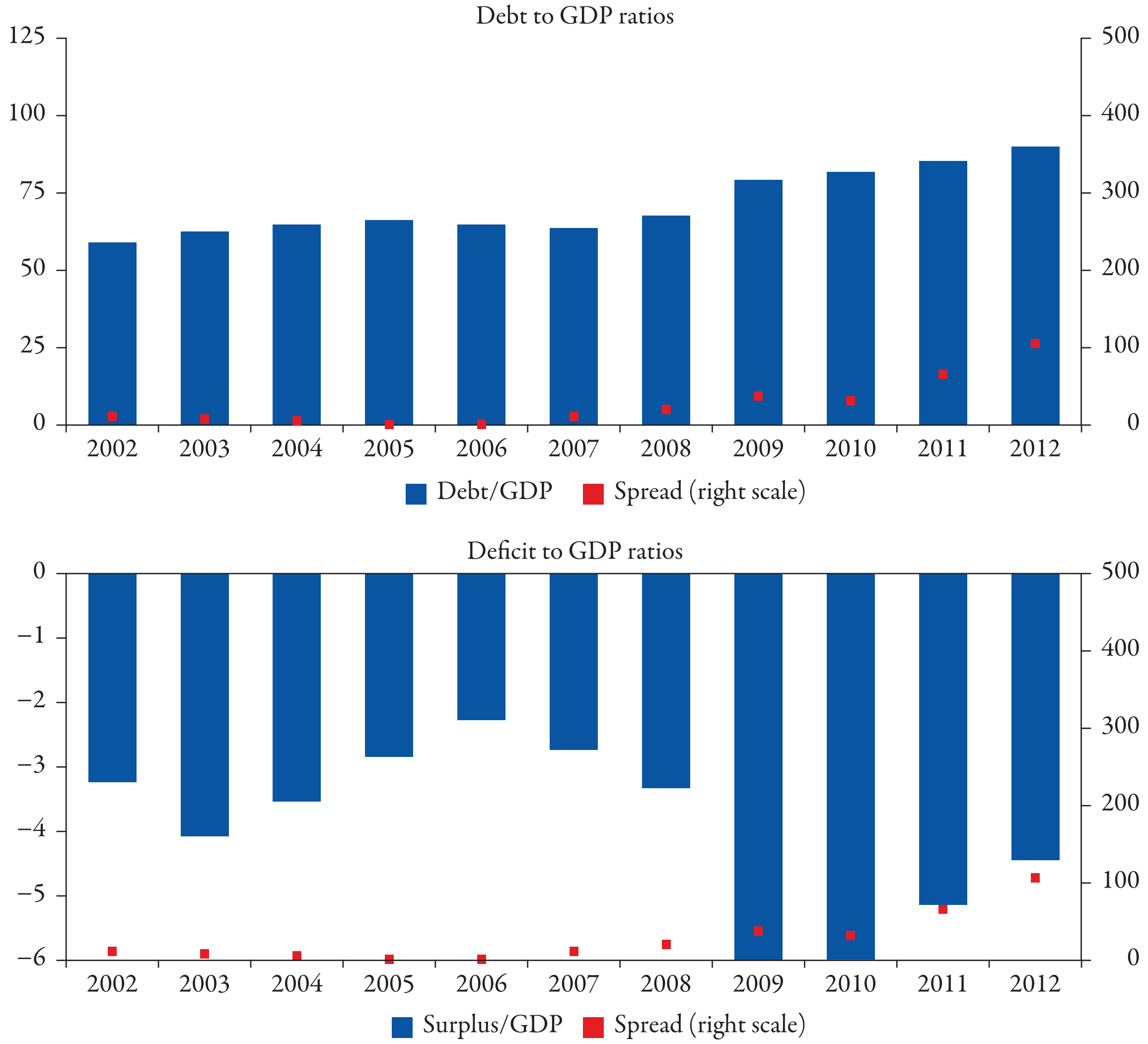

Figure 7: France: Ten year government bond yield spreads and fiscal fundamentals (ten year government yield spreads are computed as averages of daily data; public debt and deficit to GDP ratios are end-of-period data; for 2012 the Spring economic forecast of the European Commission is considered).

Source: Thomson Reuters and European Commission.

Overall, for the majority of high-debt countries, including Italy, fiscal fundamentals appear to have been underpriced in the period prior to the global financial crisis and overpriced during the crisis. Therefore, the departures from fiscal fundamentals appear to be time dependent. At the launch of the EMU, a positive market sentiment led to the convergence of government bond risk premia, which benefited high-debt countries; as the financial and sovereign crises erupted, a negative market sentiment on the resilience of the Euro Area favored the dispersion of spreads, impacting upon the high debt countries more and favouring countries that were perceived to be safer[6].

[6] The role of the perceived risk of a break-up of the Euro Area is also suggested by Di Cesare *et al.* (2012).

3 The determinants of government yield spreads: a review of recent empirical evidence

A large empirical literature has studied the determinants of government bond spreads in the Euro Area since the beginning of the EMU. Many of these studies estimate a reduced form model by regressing the sovereign spreads at certain maturities on a set of explanatory variables. These variables may be grouped into factors affecting the public debt sustainability, other macroeconomic factors, such as the external position of the economy, the liquidity of the sovereign bonds, international risk, and global risk aversion indicators.

Public debt sustainability, which proxies sovereign default risk, is affected by fiscal variables, economic growth, inflation rates, and interest rates. A rising budget deficit as well as a rising primary budget deficit are obvious indicators of increasing fiscal fragility. Also a high stock of debt weakens public finance sustainability, since it implies burdensome debt service payments and consequently a greater exposure to small changes in interest rates[7]. As the deficit and debt grow, sovereign default risk rises too, thus prompting a surge in the risk premium demanded by investors.

Empirical evidence for the Euro Area mostly confirms the role of fiscal fundamentals, although its significance varies across countries. As pointed out by earlier studies, at the onset of the EMU the ratio of debt-to-GDP was found to be relevant for some of Eurozone countries (namely, Spain and Italy), and to affect bond yields according to a non-linear relationship – only in the case of interaction with international risk indicators (Pagano and von Thadden, 2004). The relevance of fiscal fundamentals seems to change not only across countries but also over time. Most recent studies analysing the impact of the latest crises provide evidence in this sense (Von Hagen *et al.*, 2010; Favero and Missale, 2012). De Grauwe and Ji (2012) show that during 2010-2011 a significant portion of the rise in the spreads of Portugal, Ireland, Greece, and Spain was unrelated with the underlying fiscal fundamentals, being driven rather by the surge in negative market sentiment. Such sentiment did not act with respect to stand-alone countries, *i.e.* countries that issue debt in their own currencies, in spite of the fact that their debt-to-GDP ratios and fiscal space variables are as equally high and increasing. According to the authors this phenomenon is mainly due to the perceived fragility of the Euro Area, due to the fact that member countries issue debt in a currency that they cannot control[8].

According to more recent analyses conducted by the IMF, the observed sovereign spreads with respect to Germany of countries deemed more vulnerable to market tensions are well above what could be explained by fiscal and other long-term fundamentals (IMF, 2012). For Italy and Spain, in the first half of 2012 the estimated values of the spreads are around 200 basis points.

[7] In this regard it was pointed out that all the measures of fiscal fragility potentially suffer from an endogeneity problem, given that they are affected by changes in bond yields. However, as long as the average maturity of the debt is not too short, the contemporaneous impact of movements in interest rates on either the deficit to GDP ratio or the debt-to-GDP ratio is rather low.

[8] In other words, for Eurozone countries there is no guarantee that the central bank would step in to pay bondholders in the case of a liquidity crisis.

All the above mentioned studies assume that the coefficients of the relationship between fiscal fundamentals and spreads are time invariant until a discrete structural break occurs. Bernoth and Erdogan (2010) depart from this hypothesis and use a time varying coefficient model to capture the gradual shift of such a relationship affecting 10 EMU countries between 1999 and 2010. Attinasi *et al.* (2009) and Gerlach *et al.* (2010) identify the events that contributed to the re-pricing of the sovereign risk for some Euro Area countries since the eruption of the 2008 financial crisis. Alessandrini *et al.* (2012) show that a structural break occurred in 2010 leading to an upward re-assessment of the default risk of high debt countries. Giordano *et al.* (2013) use a set of multiplicative time dummies in order to detect three types of contagion effects («pure contagion», «wake-up-call effect» and «shift contagion»).

As recalled above, besides fiscal fundamentals, the overall state of the economy is of crucial importance in determining the country's ability to meet its payment obligation. In principle, a rising debt is not a problem as long as the economy grows at a faster pace than its public debt. In this sense the empirical evidence is mixed; however, most recent studies confirm the relevance of the negative impact of economic growth on spreads (Alessandrini *et al.*, 2012; De Grauwe and Ji, 2012).

The role of the external sector is investigated in several studies. Both the current account balance, *i.e.* exports minus imports, and the real effective exchange rate are found to be significant (Alessandrini *et al.*, 2012; De Grauwe and Ji, 2012; Maltriz, 2012). The current account balance is expected to negatively affect government bond yields, owing to its role as an indicator of competitiveness and of a country's ability to raise funds for debt servicing; therefore as it improves, the sovereign spreads should decline. Conversely, as pointed out by De Grauwe and Ji (2012), current account deficits signal an increase in net foreign debt which either directly (if spurred by public overspending) or indirectly (if due to private sector's overspending) undermines a government's ability to meet its payment obligations[9].

Sovereign yield spreads may also be influenced by liquidity risk, that is, the risk of having to sell or buy the asset in an illiquid market, at an unfair price, therefore bearing high transaction costs. The liquidity risk is usually measured through either bid-ask spreads or the size of the sovereign bond markets. The evidence presented by the empirical literature on this issue so far is controversial (Beber *et al.*, 2009; Haugh *et al.*, 2009; Favero and Missale, 2012; Bernoth and Erdogan, 2010).

Besides the mentioned country specific variables, there is strong evidence showing that spreads are driven by a common international factor (Codogno *et al.*, 2003). Such a relationship is usually captured though a proxy such as the spread between the yields of US corporate bonds and the yields of US Treasuries (Codogno *et al.*, 2003; Attinasi *et al.*, 2009; Bernoth and Erdogan, 2010; Gerlach *et al.*, 2010; Schuknecht *et al.*, 2010; Favero

[9] According to Maltriz (2012) the relationship between spreads and current account balance may also have a positive sign. A positive current account surplus, which for the balance of payment identity is coupled with net capital outflows, might in fact signal either the inability of a country to borrow from abroad or a capital flight. In both cases, sovereign spreads should rise. Such a relationship would reflect short-term liquidity issues, while the negative sign of the current account recalled above would be related to long-term solvency arguments.

and Missale, 2012; Maltriz, 2012)[10] or as a composite index of several measures of risk (Alessandrini *et al.*, 2012). As pointed out by Borgy *et al.* (2011), a principal component analysis regularly reveals that the first principal component (usually interpreted as time-varying risk aversion of international investors) accounts for more than 80% in the total variation of spreads.

4 Estimation and results

4.1 The model

This section introduces the empirical models used to estimate the determinants of sovereign bond yields in the Euro Area over the January 2002 to May 2012 period. The analysis refers to the monthly 10-year spreads relative to Germany for the following ten countries: Austria, Belgium, Finland, France, Greece, Ireland, Italy, the Netherlands, Portugal, and Spain.

We did not consider the CDS spreads, because this alternative measure for the default risk suffers from three main shortcomings: first of all, data of CDS premia are available only for a few countries and for the most recent years; moreover, CDS premia are driven not only by credit risk but also by counteparty risk; finally, during the crisis, CDS premia might have been affected by short-selling bans imposed in some countries (Aizenman *et al.*, 2011). To begin, we simply regress spreads on the country's fiscal position, economic growth, and external sector position as well as on a global risk aversion indicator according to the following specification:

$$\text{(1.a)}\quad \begin{aligned} Spread_{it} = \alpha &+ \beta_1 FS_{it} \\ &+ \beta_2 FS^2_{it} + \beta_3 Gr_{it-1} + \beta_4 IP_{it} + \beta_5 CA_{it-2} + \beta_6 REE_{it-1} + \beta_7 Liq_{it} \\ &+ \beta_8 GRA_t + u_{it} \end{aligned}$$

In (1.a) FS_{it} stands for *Fiscal space* (defined as the ratio between sovereign debt and tax revenues) of country i at time t; this variable enters both in level and quadratic terms (more details are given below on this point). Gr_{it} refers to the GDP growth rate while IP_{it} denotes the industrial production of country i at time t: both variables account for *economic activity*. *External competitiveness* variables are also included, that is CA_{it-2}, which stands for the current account balance relative to GDP, and REE_{it-1}, the real effective exchange rate. Liq_{it} refers to the share of country i public debt over the total debt outstanding in the Euro Area at time t. Finally, GRA_t (*Global risk aversion*) is an indicator of international risk. An alternative specification to (1.a) replaces the fiscal space with the debt-to-GDP ratio ($Debt_{it}$) as follows:

[10] Pagano and von Thadden (2004) recall that the appropriateness of such a measure as a proxy of the global risk factor is supported by empirical evidence showing significant spillovers between the volatilities of the return series of European and US bonds.

$$Spread_{it} = \alpha + \beta_1 Debt_{it} + \beta_2 Debt^2_{it} + \beta_3 Gr_{it-1} + \beta_4 IP_{it} + \beta_5 CA_{it-2} + \beta_6 REE_{it-1} + \beta_7 Liq_{it} + \beta_8 GRA_t + u_{it} \tag{1.b}$$

For the sake of brevity and clarity we will refer to (1.a) and (1.b) also as the *Basic models*.

We neglect other variables, such as the inflation rate and the short term interest rate, which according to some empirical contributions may be relevant (Alessandrini *et al.*, 2012), because they were never statistically significant; we also tested the relevance of the primary deficit/surplus over GDP and of the budget balance over GDP but they were never significant. Moreover, following De Grauwe and Ji (2012), we do not add sovereign ratings or other measures of systemic risk (such as the first component of the CDS of Euro Area countries or similar) because they might introduce an endogeneity bias, given that they tend to react to changes in government bonds yields.

Before estimating the model, we addressed the empirical issues raised by two features of the data set used: the first is the presence of seasonal cycles in the macro data; the second is the discrepancy between the frequency of the dependent variable and the frequency of the explanatory variables.

Cyclical fluctuations characterise many monthly or quarterly time series. If not removed, such fluctuations may hinder the understanding of the underlying trends; this problem is easily overcome by using adequate seasonal adjustment tools.

In the present work we applied a moving average (MA) filter to smooth both fiscal data (namely *tax revenues*, which is the numerator of the *fiscal space variable FS*), whose time series exhibit the typical step-shape due to the cyclicality in public finance data, and economic activity data (*e.g.* GDP growth, industrial production index and current account data), which are affected by external seasonality conditions, holydays, etc.

The MA smoothing allowed us also to extract observations with a higher time variability from the aggregated observations of the low moving variables (for example, monthly values from the quarterly data of the GDP growth rate). This helped to address the biases that may have resulted from the combination of the daily data of the government bond spreads with the quarterly data of the fiscal and macroeconomic variables. In empirical work this combination is usually accomplished by lowering the frequency of the variables with higher moving periodicity through aggregation, and by keeping the low frequency variables constant until a new observation occurs. However, on statistical grounds this is equivalent to introducing a measurement problem, which may bias the estimated coefficients of the explanatory variables towards zero (Gerlach *et al.*, 2010). Hence, we preferred to extract monthly observations from the quarterly observed information through the application of MA smoothing. This, in turn, allowed us to limit the extent of the aggregation of the daily data on the spreads (to the monthly rather than to the quarterly frequency) and to have all the variables in the model at a monthly frequency.

Finally, the MA smoothing also helped us to collapse together in every single observation the values at time *t*, one or more lagged values and one or more leading values recorded at some future dates, depending on the width of the time-window which was appropriately chosen on a case-by-case basis[11]. In this way we averaged across the different values which may have been relevant for the investors at time t[12]. In other words, the spread at time t may have reacted to the GDP growth recorded in t, to its past values to the extent past realizations of the GDP growth affect the country credit risk with a delay (more detail is provided on this point below) and to the expected value of the GDP growth, proxied by the values observed after t. For example, we apply a MA (2,1,2) filter to the *industrial production* variable (IP_{it}) and a MA (4,1,4) filter to the GDP. As we chose a symmetric time-window, past and future values are weighed equally.

The models specified above regress the spread at time t on a set of variables observed at either t, $t-1$ or $t-2$. In fact, it may take some time before the change in a macro variable impacts the sovereign default risk, depending on the features of the transmission mechanism in place. For example, a current fall of the GDP growth rate will lower tax revenues in the future, which in turn will result in a future deterioration of country solvency. The same line of reasoning holds for the degree of competitiveness, as captured by the current account balance and the real effective exchange rate, affecting both the GDP growth (and hence tax revenues and country solvency) and the ability of a country to raise external funds to meet its payment obligations.

The estimation results are robust with respect to the choice of different lags, as confirmed by the fact that they remain qualitatively the same using lags different from those applied in (1.a) and (1.b) (more details in Section 4.2).

Let us now turn to a deeper analysis of the variables included in (1.a) and (1.b) and of their expected sign (see also Table A.1 in the Appendix for details on the definition, the source and the frequency of the variables).

Fiscal position. As outlined in the previous section, the role of the variables accounting for country fiscal position has long been investigated in the literature. In particular, we followed Aizenman *et al.* (2012) and De Grauwe and Ji (2012), who advocate that fiscal space, defined as the ratio of debt-to-total tax revenues, is a better measure of debt sustainability because it takes into account the government's ability to raise taxes: in fact a low-debt country can face as many difficulties as a high-debt country if it takes a lot of time to generate the revenues necessary to meet its payment obligation[13]. Therefore, in this study fiscal space (*FS*) and debt-to-GDP (*Debt*) were used as alternative measures of country fiscal fragility. Moreover, following the literature and given the evidence substantiated by the descriptive analysis reported in section 2, these variables are included both in levels and quadratic terms to account for a non-linear relationship. As commented on by Grauwe and Ji (2012), theoretical studies model the default decision as a

[11] As a robustness check we run equations (1.a) and (1.b) using MA filters with different time width and results hold.

[12] High frequency financial data (such as the bond yields) reflect the investors' reaction to an information set which may differ from the one available to researchers, commonly using revised macro data. However macro data are subject to revisions, which are made available to the public with a lag. Therefore, the market can still react in t to the release of the information referring to past periods if it differs significantly from the forecasted value. Revisions may be substantial especially during turbulent periods.

[13] As pointed out by Borgy *et al.* (2011), the choice of the most appropriate measure of the fiscal fundamentals is a matter of debate. For instance, Bernoth *et al.* (2006) argue that debt service (*i.e.* the ratio of gross interest payments to current government revenue) is preferable since governments have less incentives to manipulate it than other indicators that are used officially to monitor the individual country's fiscal position.

discontinuous one, becoming more and more likely as the debt-to-GDP ratio rises. This in turn implies that the higher the debt-to-GDP ratio, the more sensitive investors are to a given increase in the ratio itself.

Economic activity. In line with the literature, we included variables capturing the overall state of the economy such as GDP growth rate (*Gr*), lagged by one period, and the industrial production index (*IP*). Both these variables are expected to contribute negatively to the spread, given that the higher they are the better the country's fiscal position. We use the industrial production index because it is a leading indicator and as such plays an important role in the formation of investors' expectation. Indeed, it is released at a higher frequency than the GDP growth rate and, as a contributor to an economy's growth, it is regarded as an early indicator of the state of the economy[14].

Due to the economic linkages between *Gr* and *IP* on one hand and between *CA* and *REE* on the other, we tested for the existence of collinearity which might affect the significance of the estimated parameters. A pairwise collinearity test (*Gr-IP* and *CA-REE*) rejected the null hypothesis of collinearity.

External competitiveness. Following the literature, we included both the current account balance relative to GDP (*CA*) and the real effective exchange rate (*REE*). We included the lagged values of such variables under the hypothesis that, as mentioned above, their impact on the spread may exhibit a certain sluggishness.

Liquidity. As a measure of the market liquidity of government bonds (*Liq*) we use countries' debt relative to the overall debt of all EMU countries in order to take into account the countries' market size with respect to the whole Euro Area. For lack of data, we did not use the bid-ask spread; however our measure is often used in the empirical literature, which also shows that it is highly related to other liquidity proxies (see Maltriz, 2012, for a deeper discussion of this issue). The expected sign of the impact of liquidity on spreads is negative: the deeper the secondary markets of government bonds, the lower the liquidity premium priced into sovereign spreads.

Global risk aversion. As already stated in the previous section, sovereign bond spreads are driven not only by country specific factors but also by a time-varying international risk factor (*GRA*), which in turn affects international risk appetite. Following the literature, in our analysis we capture such a factor with the spread between the yield on AAA and BBB US corporate bonds. A widening of this spread signals shifts in investors' preferences from the riskier to the safer private sector assets. We also run the model with alternative international risk indicators, such as the VIX, obtaining results similar to those reported in Table 1 (see section 4.2).

We also expanded the *Basic model* in order to account for time dependency and for country fixed effects. As showed by the descriptive analysis in section 2 and as documented by the most recent empirical contributions outlined in section 3, both the convergence of sovereign spreads recorded since the start of the EMU and the dispersion which arose after the eruption of the Greek crisis signal a mispricing of fundamental fiscal factors. Up until 2010 the market was not excessively worried about the vulnerabilities of high debt

[14] However we also tested the industrial production significance in $t-1$ and the results were basically unchanged with respect to those reported in Table 1 (see section 4.2).

countries. Since the beginning of 2010, however, the market has over-reacted to fiscal position factors by penalising the non-core member countries in particular. To account for a possible mispricing of fundamental fiscal factors, we included yearly time dummies in (1.a) and in (1.b). Moreover, in order to capture non-linearities in the contribution of the debt-to-GDP ratio driven by the evolution of global conditions, we combined the debt-to-GDP ratio with the global risk aversion by using an interacted variable from mid-2011 onwards. In this way we tested whether changes in the perception of the countries' default risk, and hence of their fiscal fundamentals, can also be traced back to the evolution of international risk factors, thus introducing another source of non-linearity in the relationship between fiscal variables and spreads. Finally, we also added country dummies, in order to capture country fixed effects due to institutional and structural features which are time invariant and may impact the spread:

$$
\begin{aligned}
(2.a) \qquad Spread_{it} = {} & \alpha + \beta_1 FS_{it} \\
& + \beta_2 FS_{it}^2 + \beta_3 Gr_{it-1} + \beta_4 IP_{it} + \beta_5 CA_{it-2} + \beta_6 REE_{it-1} + \beta_7 Liq_{it} + \beta_8 GRA_t \\
& + \beta_8 Debt_{it} * GRA_t * D_{post\ July\ 2011} + \sum_{t=1}^{10} \mu_t D_t + \sum_{i=1}^{10} \delta_i Z_i \\
& + u_{it}
\end{aligned}
$$

where D_t stands for a vector of unit quarter time dummies, covering the interval from 2003 to the first semester of 2012, and Z_i stands for the dummy for country *i*; the term $Debt_{it} * GRA_t * D_{post\ July\ 2011}$ is the interacted variable between the debt-to-GDP ratio and the global risk aversion indicator from the second semester of 2011 onwards. We will refer to (2) as the *Time dependent model*. The *Basic model* was also run by using two alternative measures of the country's fiscal position, that is the fiscal space variable (model 2.a) and the debt-to-GDP ratio (model 2.b; this does not include the debt-risk aversion interacted variable to prevent collinearity problems).

Finally, we took into consideration a well known salient feature of most economic time series, that is the inertia (or sluggishness) which may make consecutive observations interdependent. Time series data on government yield spreads exhibit trend. Therefore, we performed a variety of test for unit roots (or stationarity) in panel datasets which confirmed that the government yield spreads variable has a unit root (see Appendix, Table A.2). Moreover, in order to prevent the instance of spurious regression due to the same order of integration of the dependent variable (spread) and some other explanatory variables, we performed a Fischer-type panel unit root test for *IP* (industrial production) and *Gr* (GDP growth; Choi, 2001). For *IP* we rejected the null hypothesis (under which all panels contain unit roots) with 1% error (inverse chi-squared statistic equals to 40.15). On the contrary, *Gr* turns out to be non-stationary of order I(1) as we cannot reject the Fischer test (inverse chi-squared statistic equals to 19.39 and p-value equals to 46%). As a second step, we test for the existence of a cointegration relationship between the spread and the above explanatory variables (*IP* and *Gr*) and we find that for almost all countries there are no cointegration vectors, thus excluding the case for spurious regression between the spread and its determinants (see Table A.3 in the Appendix). In order to avoid the misspecification problems due to the omission of the lagged value of the dependent variable in the model we use the feasible generalised least square estimator (FGLS) accounting for the presence of AR(1) autocorrelation within panels. Notwithstanding the fact that we adopt a FGLS with AR(1) residuals and use country fixed effects and time dummies in order to alleviate the high persistency of the dependent variable (which is non stationary or order 1), residuals still remain non stationary (see Appendix, Table A.4). To account for the non-stationarity problem we also estimate a model in first difference (by introducing the lagged value of spread as an additional explanatory variable; see Appendix, Table A.5). In the same empirical framework De Grauwe and Ji (2012) and Favero and Missale (2012) also find that spreads are non-stationary. While De Grauwe and Ji perform a panel unit root test (strongly rejecting the null-hypothesis of stationarity; Breitung, 2000), Favero and Missale run a Dickey-Fuller unit root test for time series model reaching the same conclusion, namely that the spread follows a I(1) stochastic process. In spite of this evidence, authors do not use the first difference of the spread (*i.e.* they do not add the lagged dependent variable to the explanatory variables) since they argue that an interest rate – such as the spread – may be non-stationary only in the short run, while in the long run it cannot move upward or downward infinitely.

4.2 The estimation results

This section presents the estimation results (Table 1). The variables accounting for countries' fiscal position, that is the debt-to-GDP ratio (*Debt*) and the fiscal space (*FS*), are statistically significant in all specifications. Moreover, the non-linear relationship between these factors and the spread is confirmed.

Consistently with previous studies, the variables proxing countries' economic activity, namely GDP growth (*Gr*) and industrial production (*IP*), have a significant and negative effect in all instances.

In addition, the variables accounting for the external position of a country, namely the current account balance (*CA*) and the real effective exchange rate (*REE*), are significant. However these variables lose significance when time dependency is accounted for (*i.e.* in the *Time dependent model* 2.a and 2.b); this result is not unexpected since the current account (*CA*) and the real effective exchange rate (*REE*) are the most seasonal explanatory variables and therefore their significance could have been lessened in the estimated equations that include time dummies.

Government bond liquidity (*Liq*), as proxied by each country debt market share over the debt of all the EMU members, is almost always estimated to be important. It gains significance in the *Time dependent model* (2.b), thus confirming existing empirical evidence claiming that investors value liquidity more during turbulent periods.

Finally, as expected, the time dummies are strongly significant in the aftermath of the 2008 financial crisis, that is when the countries, which were previously perceived as safe, became involved in the sovereign debt crisis (Spain, Italy and Belgium). Moreover, the

Table 1: Estimation results

Variables	Basic model (1)		Time dependent model (2)	
	1.a (using fiscal space)	1.b (using debt ratio)	2.a (using fiscal space)	2.b (using debt ratio)
fiscal space	0.0453		−0.4561 (***)	
fiscal space squared	0.0138 (*)		0.0563 (***)	
debt / GDP		−0.0212 (*)		−0.0300 (**)
debt / GDP squared		0.0345 (***)		0.0551 (***)
GDP growth (in $t-1$)	−0.1952 (**)	−0.2543 (***)	−0.2423 (***)	−0.3377 (***)
industrial production	−0.0068 (**)	−0.0071 (***)	−0.1059 (***)	−0.0082 (***)
current account (in $t-2$)	−0.0391 (**)	−0.0381 (**)	0.0077	0.00846
real effective exchange rate(in $t-1$)	0.0206 (**)	0.0224 (**)	0.0121	0.0139
liquidity (debt share)	−0.0457 (**)	−0.0679 (***)	−0.0539	−0.1967 (**)
GRA	0.1373 (***)	0.1516 (***)	0.0929 (*)	0.1573 (***)
debt * GRA post July 2011			0.9457 (***)	
time component				
2003			−0.0000	0.0004
2004			0.0002	0.0012
2005			0.0010	0.0027 (**)
2006			0.0016	0.0034 (**)
2007			0.0021	0.0040 (***)
2008			0.0033 (**)	0.0056 (***)
2009			0.0060 (***)	0.0076 (***)
2010			0.0092 (***)	0.0100 (***)
2011			0.0108 (***)	0.0106 (***)
2012			0.0115 (***)	0.0119 (***)
constant	−0.0098	−0.0059	0.0060	0.0139
country fixed effect			controlled	controlled
Wald chi2 (degrees of freedom in parentheses)	(8) 90.00	(8) 142.16	(28) 875.01	(27) 539.27

Note: fiscal space, GDP growth, industrial production and current account balance were seasonally adjusted through a moving average (MA) filter; the length of the moving window was appropriately chosen depending on the time series. Such smoothing allowed us to obtain monthly estimated values for the variables, which were used in the estimation. (***) significant at 1% ($p < 0.01$), (**) significant at 5% ($p < 0.05$), (*) significant at 10% ($p < 0.1$).

inclusion of the time dummies causes the fiscal variables to gain statistical and economic significance. This supports the hypothesis that investors' valuation of a country's fiscal position is time varying, and that it is dependent on the level of the international risk (*GRA*), which is significant in all specifications. Along the same line of argument we can interpret the significance in the specification (2.a) of the debt-risk aversion interacted variable ($Debt_{it} * GRA_t * D_{post\ July\ 2011}$), which turns out to be relevant.

Figure 8 plots the observed spreads and the fitted spreads resulting from the *Basic Model* 1.a and the *Time dependent model* 2.a for Italy, Spain, France and the Netherlands (the fitted values look similar when using other specifications, that is 1.b and 2.b; the results for the other countries are available upon request from the authors).

For Italy and Spain, the *Basic model* predicts that their sovereign risk should have been priced higher up until 2010 and much lower from then on. This provides evidence supporting the hypothesis that investors demanded a premium which, relative to the economic and financial fundamentals, was too low up until the financial crisis and too high thereafter 2010. Hence, a relevant fraction of the relentless increase in both the Italian and Spanish spreads is explained by the contagion phenomenon: the *Time dependent model*, accounting for the impact of negative market sentiment, quite closely tracks the pattern of observed spreads.

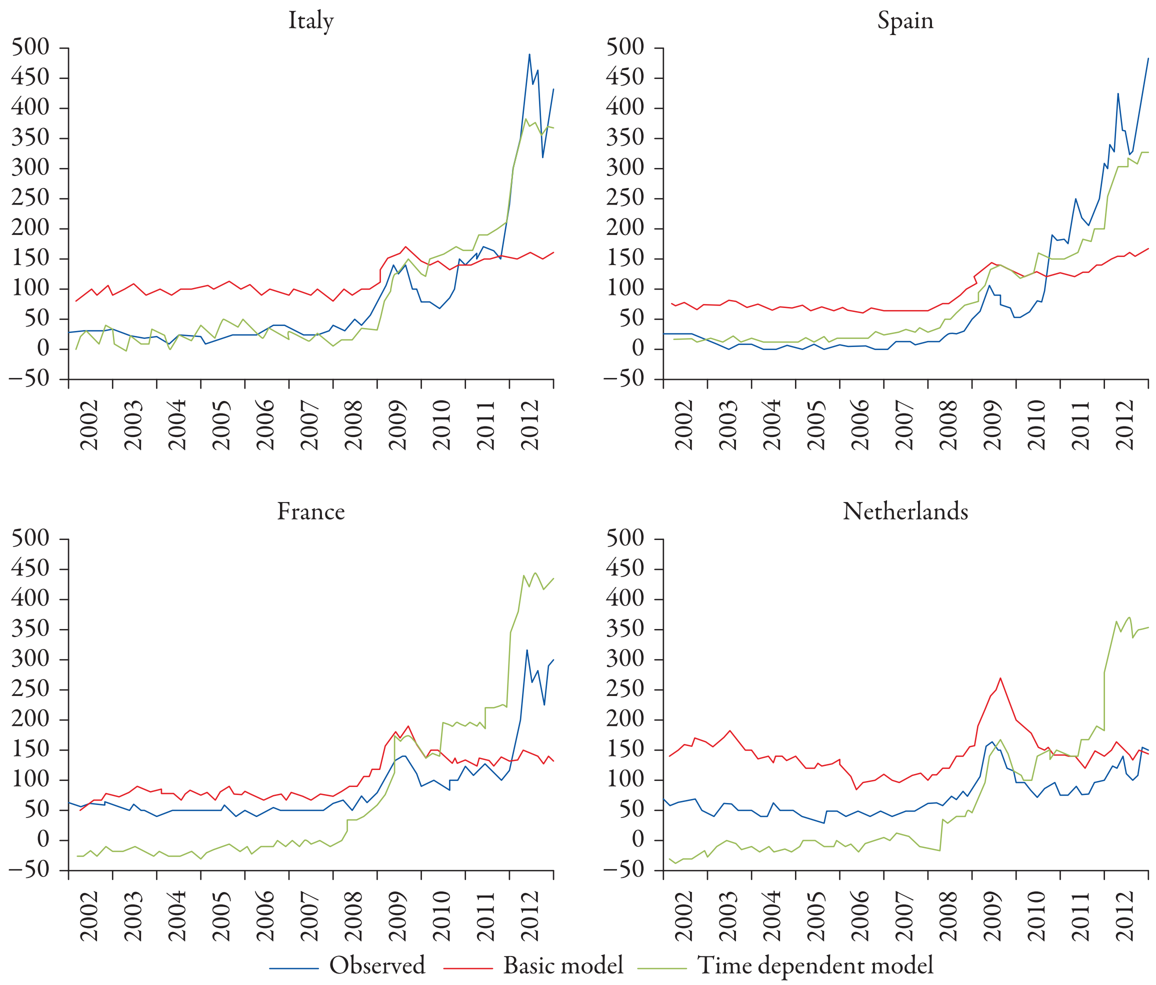

Figure 8: Actual and fitted values of sovereign spreads for Italy, Spain, France and the Netherlands (values in basis point).

In order to disentangle the role of country-specific contagion effects from fundamentals factors, we estimate the share of predicted spreads due to each component (macroeconomic and fiscal variables *versus* contagion) without assuming that contagion is equal to the difference between observed and fitted spreads (*i.e.* residuals), but rather implementing specific econometric tools (margins and marginal effects) that investigate how much of the total predicted spreads can be accounted for by each component included in the model.

The calculation of margins of responses and derivatives of responses (marginal effects) allowed us to obtain the percentage share of average annual variation of spreads due to contagion for all Euro Area countries (so called *systemic contagion*) and the amount of spread that for each single country is solely ascribed to contagion (so called *idiosyncratic contagion*).

Margins are statistics calculated from predictions of a previously fitted model at fixed values of some covariates and averaging or otherwise integrating over the remaining covariates (Searle *et al.*, 1980). In our model the covariates are the time dummies which incorporate the effects of contagion. For instance, after a regression fit on time t and

Table 2: Percentage share of annual spreads' variation due to systemic contagion

	Time dependent model (2.a) Margins of Responses (ΔS)	Time dependent model (2.b) Margins of Responses (ΔS)
2007	–	9.63%
2008	19.13%	19.96%
2009	35.57%	21.11%
2010	31.57%	22.03%
2011	11.56%	3.62%
2012	4.05%	9.09%

$t+1$, the marginal mean for time t is the predicted mean of dependent variable ($Spread_{it}$) where every observation is treated as if it were observed at time t[15].

In other words, margins of responses give us the magnitude of the contagion effect within the sample, that is the percentage share of the annual variation of the spreads due to time-varying market sentiment (*systemic contagion*), keeping constant all other economic fundamentals.

Table 2 shows for the selected two models previously estimated (*Time dependent models* 2.a and 2.b) the percentage share of total annual variation of observed spreads which can be ascribed to systemic contagion, that is the annual movement of spreads solely due to the impulse transmitted by time dummies. As already mentioned, these contagion effects were computed following Searle *et al.* (1980).

Both models confirm that systemic contagion reached its peak during 2009-2010, in the aftermath of the subprime crisis, when it explains almost one third and almost one fourth of the increase in the spreads. According to specification (2.a), almost 36% of the increase in spreads during 2009 was due to contagion, which occurred as a consequence of the financial turmoil, rather than resulting from the deterioration of the credit risk or the solvency risk of single countries.

Coefficients for time determinants increased rapidly during the financial crises and seem to flatten in the last two years of the estimation period. However, according to model (2.b) the impact of systemic contagion rebounds in the first semester of 2012 (accounting for a 9.09% increase against the 3.6% in 2011)[16].

In order to obtain a country specific measure of contagion (*idiosyncratic contagion*), we calculate the derivatives of the responses (marginal effects), which are an informative way of summarising the fitted results[17].

[15] Standard errors are obtained by the delta method which assumes that the values at which the covariates are evaluated to obtain the marginal responses are fixed.

[16] Note also that we have only 5 monthly observations for 2012 (January-May 2012).

[17] The change in a response for a change in the covariate is not equal to the parameters estimated; one should take into account interactions between country and time specific covariates (country dummies * time dummies). In order to overcome this complication we need to run the fitted model, compute the partial derivatives and make inference on these (Buis, 2010; Baum, 2010). Consider a very simple model, such as:

$$y = \beta_0 + \beta_2 x + \beta_2 d_{time} + \beta_2 d_{country} + \beta_4 (d_{time} * d_{country}) + \varepsilon$$

The partial derivative in d_{time} is:

$$dy / d_{time} = \beta_2 + \beta_4 d_{country}$$

that is the sum of two components, a time effect which is common to all of samples (β_2) and a time effect that changes

To compute these marginal effects (*idiosyncratic contagion*) we include nine multiplicative time-country dummies in our models (Italy, Spain, France, Portugal, Ireland, Greece, Finland, the Netherlands, Austria)[18] and obtain nine specific country coefficients for each year, representing the specific country's response to the time effects (Tables 3 and 4)[19].

Marginal effects measure to what extent spreads are greater or lower than the fitted values predicted by the model on the basis of economic and fiscal factors only.

Results can be summarised as follows:

– Core countries (France, Finland, the Netherlands and Austria) were not affected by the upsurge in financial turmoil during the subprime crisis: in fact the share of predicted spreads attributable to contagion is estimated to be equal to zero. Since the eruption of the sovereign debt cris, such countries have experienced a lower spread than what would be justifiable by their economic fundamentals (in the first half of 2012 France and Netherland are predicted to have benefited from a discount of 53 and 57 b.p. respectively).

– Some peripheral countries (Spain, Portugal, Ireland) suffered an abrupt revision of their credit risk since the insurgence of the sovereign debt crisis, which triggered the market revision of their already known economic fragility. As a consequence, they experienced an overpricing phenomenon on their spreads due to contagion from the start of 2010 onwards (for Spain, the contagion effect reached its peak during the first few months of 2012, with values ranging from 167 to 242 basis points, depending on the specification adopted).

– Italy experienced a rising contagion effect that in the first semester of 2012 reached a value ranging, depending on the specification adopted, between 147 to 181 basis points. This penalisation may be explained by its historically high debt-to-GDP ratio, which makes Italy particularly exposed to the reversals of market sentiment[20].

by country (β_4) and represents the specific country response to the time fluctuations (in other words, how severe the impact of financial contagion to one country is compared with the responses of others).

[18] Belgium is the country omitted.

[19] The derivatives of responses showed in the following tables are obtained from the fitted models illustrated above and use statistical properties of covariates to make inference. Tables 3 and 4 highlight the marginal effects results (derivatives of responses) referred to specifications (2.a) and (2.b) for the period 2007-2012. Based on the theoretical predictions of a fair value of government yield spreads for each country in every year – that is the value of the spreads derived from fundamentals variables incorporating credit risk of debtors – the marginal effects methodologies give us the share (in basis points) of that fair value which is ascribable to idiosyncratic contagion. We don't take this value from residuals of estimations (as residuals are unexplained components and it's open to some degree of subjectivity to impute those to specific roots), but we estimate coefficients of time dummy, country dummy and interactions between them as they correctly represent the weight of non-core fundamental variables in transmission of financial contagion to the selected countries.

[20] To test the robustness of the results we re-estimated equations (1.a, 1.b, 2.a and 2.b) by applying different lags to the independent variables. For the sake of brevity we only state the main differences which emerged (detailed results are available upon request from the authors). If the industrial production variable is taken at time $t - 1$ (instead of t) different results are observable depicted in table 3 (we only report 2012): Spain 176 b.p., France –47 b.p. and the Netherlands –63 b.p., all other results being equal. If we take GDP growth at time t (instead of $t - 1$) we get: Italy 192 b.p., Spain 180 b.p. and Portugal 671 b.p. In the case of the fiscal position variables (fiscal space and fiscal space squared) taken at $t - 1$ (rather than t) results will change to a greater degree: Italy –240 b.p., France –90 b.p., Spain 150 b.p., the Netherlands –52 b.p. and Portugal 620 b.p.. We did the last robustness check by taking the external competitiveness variables (current account and real effective exchange rate) at time t instead of $t - 1$ and we obtained, other things being equal, that for Italy the amount of spread due to contagion is 195 b.p. (181 in the basic model of Table 3), for France is –46 b.p. and for the Netherlands –70 b.p.

Table 3: Idiosyncratic contagion effects – *Fiscal space model* (*2.a*)

	Italy	Spain	France	Portugal	Ireland	Greece	Finland	Netherlands	Austria
2007	65 (**)	0.00	0.00	0.00	0.00	–59 (**)	0.00	0.00	0.00
2008	86 (***)	0.00	0.00	0.00	0.00	–48 (*)	0.00	0.00	0.00
2009	80 (*)	0.00	0.00	0.00	124 (***)	–74 (**)	0.00	0.00	53 (*)
2010	113 (**)	110 (***)	0.00	168 (***)	166 (***)	164 (***)	0.00	0.00	48 (*)
2011	155 (***)	155 (***)	0.00	446 (***)	247 (***)	340 (***)	0.00	0.00	0.00
2012	181 (***)	167 (***)	–53 (*)	665 (***)	0.00	–	0.00	–57 (*)	0.00

Note: amount of spread (in basis points) due to country specific contagion. (***) α = 0.001, (**) α = 0.005, (*) α = 0.01, (--) not estimable

Table 4: Idiosyncratic contagion effects – *Debt model* (*2.b*)

	Italy	Spain	France	Portugal	Ireland	Greece	Finland	Netherlands	Austria
2007	67 (***)	0.00	0.00	0.00	0.00	–78 (***)	0.00	0.00	0.00
2008	81 (***)	0.00	–41 (*)	0.00	0.00	–80 (***)	0.00	0.00	52 (*)
2009	0.00	0.00	–70 (***)	0.00	91 (**)	–126(***)	0.00	0.00	56 (**)
2010	0.00	117 (***)	–98 (***)	79 (***)	137 (***)	63 (*)	0.00	0.00	53 (**)
2011	91 (***)	193 (***)	–100 (***)	307 (***)	242 (***)	128 (***)	47 (*)	0.00	62 (**)
2012	147 (***)	242 (***)	–86 (***)	507 (***)	86 (*)	–	0.00	0.00	79 (***)

Note: amount of spread (in basis points) due to country specific contagion. (***) α = 0.001, (**) α = 0.005, (*) α = 0.01, (--) not estimable.

Figure 9 shows the share of annual average predicted spread due to fundamentals and due to contagion for each country. The left panel refers to what we labeled as a *Time dependent model with fiscal space* (2.a) and points out that in 2012 Italy suffered from a contagion which accounts for almost 50% of the total predicted spread (*i.e.* predicted spread was equal to 369 b.p., of which 181 b.p. is due to contagion). According to the *Time dependent model with debt* (2.b – right panel), the share of annual predicted spread related to contagion is equal to 147 b.p. which accounts for around 43% of the total.

We conclude our analysis by presenting, only for Italy from 2007 onwards, the disaggregation of the predicted average spread, obtained through the *Time dependent model 2.a* (hereafter $\overline{\widehat{Spread}}_{ITA}$), into two components:

– the contribution of contagion ($\widehat{Spread}_{ITA,C}$), *i.e.* the *time marginal effect* for Italy computed as above,

– the component of the fitted spread driven by fundamentals (*i.e.* excluding the time dummies): $\widehat{Spread}_{ITA,F} = \widehat{Spread}_{ITA} - \widehat{Spread}_{ITA,C}$.

Alternatively $\widehat{Spread}_{ITA,F}$ can be computed as the sum of the relative contributions of all the statically significant variables included in (2):

$$\overline{\widehat{Spread}}_{ITA,F} = \hat{\beta}_1\overline{FP} + \hat{\beta}_2\overline{Gr} + \hat{\beta}_3\overline{IP} + \hat{\beta}_7\overline{GRA} + \hat{\beta}_8\overline{Debt_{it} * GRA * D_{post\ July\ 2011}} \tag{2}$$

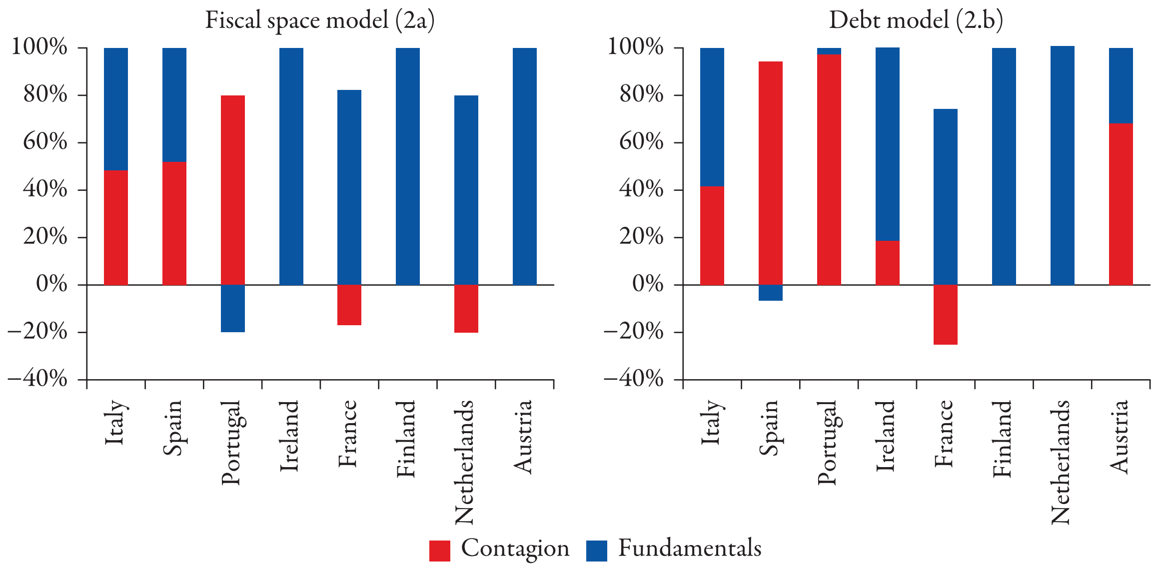

Figure 9: Percentage contribution of fundamentals and contagion to the model-predicted spreads of some European countries – 2012 estimates (labels indicate the contributions of fundamentals and contagion in basis points)

where for each regressor the yearly average is taken into account[21].

Figure 10 shows the estimated relative contributions of contagion and fundamental factors. For 2007 and 2008, fundamentals are estimated to have reduced the (fitted yearly average) spread; this is quite plausible given that at that time the overall state of economy remained still unaffected by the financial crisis. From 2009 onwards, as the general economic conditions deteriorated, fundamentals are estimated to have raised the spread.

Figure 11 disaggregates the contributions of all the fundamental regressors and of the global risk aversion to the (fitted yearly average) spread. The estimated impact of the fiscal position considered on its own (*i.e.* neglecting the post-2011 interaction with the international risk aversion) increased until 2010 (to 225 b.p. from about 68 b.p. in 2007) and then decreased (to about 120 b.p. in the first half of 2012). However, when accounting for the interaction with international aversion, the overall impact of the fiscal components (*i.e.* the sum of fiscal space and *debt* * *GRA*) has risen consistently (reaching almost 260 b.p.). Finally, the positive contribution of the industrial production shrinks as it slows down.

5 Conclusion

Since the occurence of the sovereign debt crisis at the beginning of 2010, peripheral countries of the Euro Area have experienced a relentless rise in the spread against the German Bund. On the other hand, the core countries have benefited from a flight-to-quality effect, leading to a considerable reduction in their government bond yields.

[21] As an example, the relative contribution of IP is equal to $\frac{\hat{\beta}_3 \overline{IP}}{\hat{\beta}_1\overline{FP}+\hat{\beta}_2\overline{Gr}+\hat{\beta}_3\overline{IP}+\hat{\beta}_{72}\overline{GRA}+\hat{\beta}_8\overline{Debt_{it}} * \overline{GRA} * \overline{D_{post\ 2011}}}$. This ratio is then multiplied by $\widehat{Spread}_{ITA,\,F}$ to get the contribution in basis points.

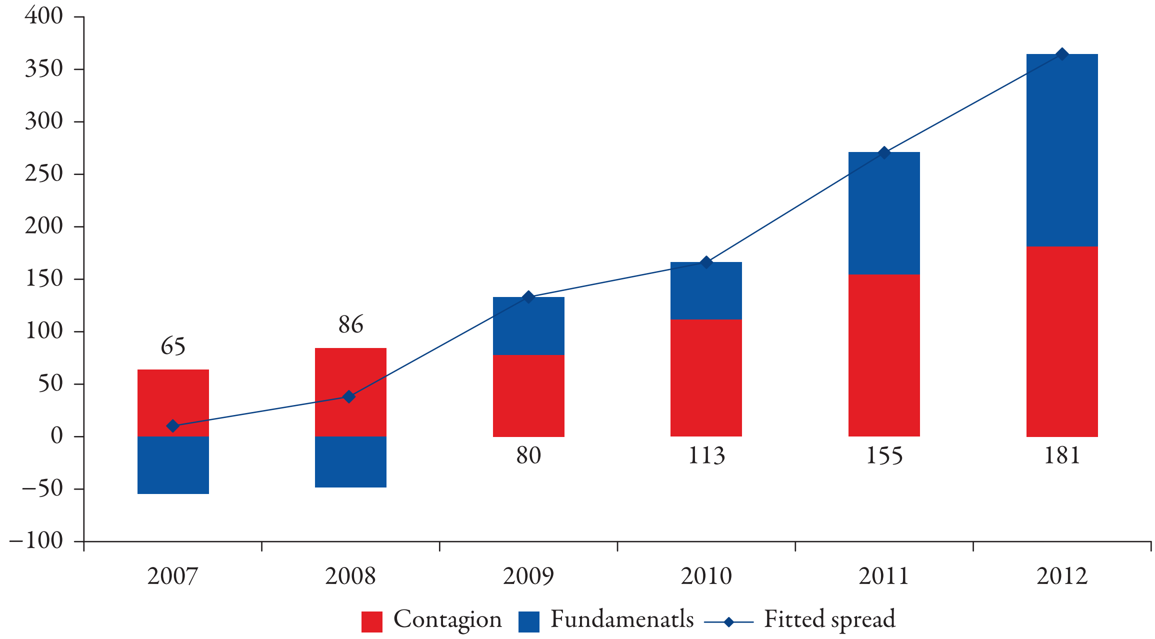

Figure 10: Estimates of the contribution of fundamentals and contagion to the Italian government bond spread (fitted values of the spread as estimated through the Time dependent model – 2.a).

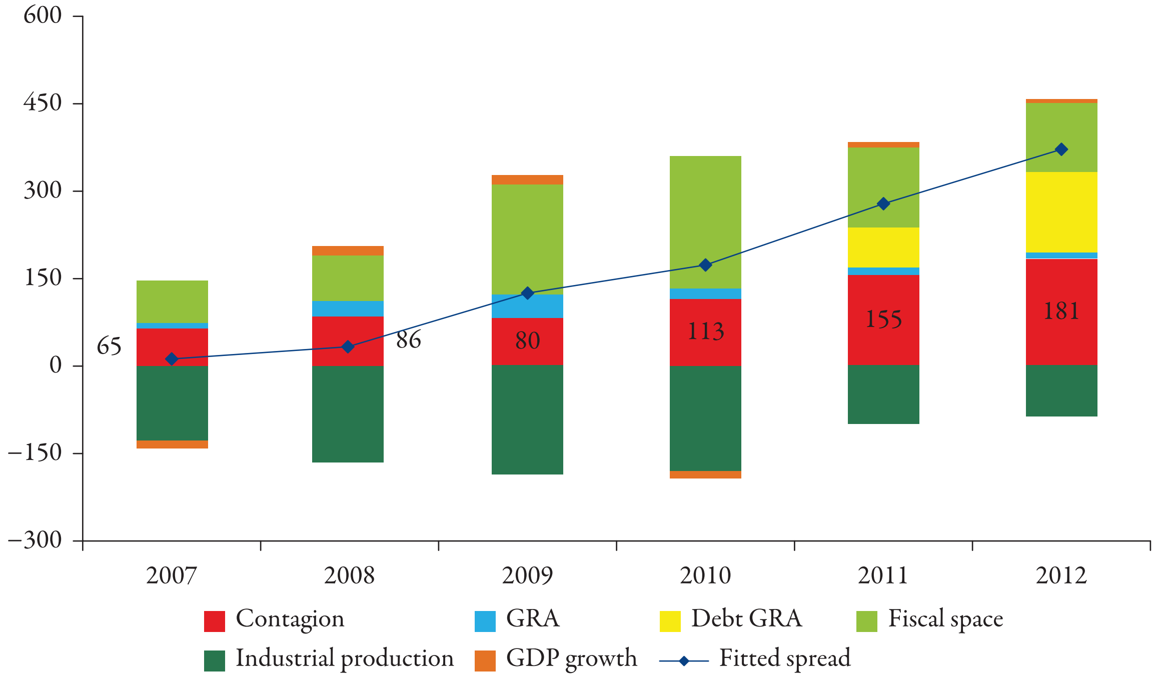

Figure 11: Estimates of the disaggregated contribution of fundamentals to the Italian government bond spread (fitted values of the spread as estimated through the Time dependent model – 2.a).

This paper analyses the determinants of sovereign spreads in the Euro Area from January 2002 to May 2012. The objective is to disentangle the role of country-specific fundamentals, driven by fiscal and macroeconomic factors, from what is referred to as «contagion».

Following the existing empirical literature, the work estimates a model of the determinants of the 10-year yield spreads relative to Germany for ten Eurozone countries. The results show that since the eruption of the financial crisis in 2007-2008, sovereign spreads have shown a time-dependent contagion component. On average, such a component explains almost one third of the spreads dynamic in 2009-2010 and almost 10% since 2011.

However, results at the country level are quite different between core and peripherals. As shown by the analysis, core countries (excluding Germany, which is our benchmark to measure spreads) were not affected by contagion until 2011; since the worsening of the sovereign debt crisis they seem to have benefited from a flight-to-quality effect. For example, in the first few months of 2012, France showed spreads lower than what was implied by fundamentals at the time by an amount ranging from roughly 50 to 90 basis points, depending on the model specification, while for the Netherlands such a «discount» was measured to be as high as around 60 basis point.

Peripheral countries, which at the inception of the European Monetary Union took advantage of a mispricing of their actual economic and fiscal fragility, have suffered from the abrupt revision of market expectations since 2009, showing spreads significantly higher on average than what is justified by macroeconomic and fiscal factors. In 2012, for most of these countries contagion had a role comparable to fundamentals in explaining the level of the spreads. For example, it accounts for an amount ranging from roughly 170 to 240 basis points for Spain, while for Italy – most likely penalised by its historically high debt-to-GDP ratio – contagion explains between roughly 150 and 180 basis points of the spread, depending on the model specification.

6 Appendix

Table A.1: The explanatory variables: description and sources

Variables	Definition	Frequency	Source
Spread	difference in yields to maturity of 10-year government bonds of ten euro member countries relative to Germany's	Monthly	Thomson Reuters
Fiscal position	gross government debt over GDP	Quarterly	Eurostat, EC
	primary balance over GDP	Quarterly	ECB, EC
	government budget deficit / surplus over GDP	Quarterly	ECB, EC
	fiscal space: Gross government debt over total tax revenues	Quarterly(1)	Eurostat
Economic activity	GDP growth; percentage change with respect to previous quarter	Quarterly(1)	Thomson Reuters
	industrial production	Monthly(1)	Thomson Reuters
External sector	current account balance over GDP	Monthy(1)	Thomson Reuters
	real effective exchange rate	Monthly	Thomson Reuters
Global risk aversion indicator	spread between the yield of US AAA corporate bonds and the yield of US BBB corporate bonds	Monthly	Fred database
Debt share	countries' debt relative to the overall debt of all EMU countries	Quarterly	Eurostat

Note: (1) Fiscal space, GDP growth, industrial production and current account balance were seasonally adjusted through a moving average (MA) filter; the length of the moving window was appropriately chosen depending on the time series. Such smoothing allowed us to obtain monthly estimated values for the variables, which were used in the estimation.

Table A.2: Unit root test (H_0 hypothesis: Panels contain unit roots)

Variable	LLC test		Harris-Tsavalis test		Breitung test	
	adjusted t*	P-value	rho	P-value	lambda	P-value
spread	7.57	1.00	1.009	1.00	3.98	1.00

We considered a simple panel-data model with a first-order autoregressive component:

$$y_{it} = \rho_i y_{it-1} + z_{it}^{'}\gamma_i + \varepsilon_{it}$$

where $i = 1,\ldots,N$ indexes panels; $t = 1,\ldots,T_i$ indexes time; y_{it} is the variable being tested (government yield spreads) and ε_{it} is a stationary error term. By default we set $z_{it} = 1$ so that the term $z_{it}^{'}\gamma_i$ represents panel-specific means (fixed effects). Panel unit-root tests are used to test the null hypothesis $H_0 : \rho_i = 1$ for all i versus the alternative $H_0 : \rho_i < 1$. We adopted three alternative specification tests proposed by Levin-Lin-Chu (2002), Harris-Tzavalis (1999) and Breitung (2000).

Table A.3: Johansen tests for cointegration between spreads and economic activity explanatory variables

Countries	5% critical values for industrial production	5% critical values for GDP growth
	15.41	15.41
	3.76	3.76
Italy	6.8267	4.92
Spain	8.9085	8.56
France	6.5483	6.65
Portugal	7.6669	14.01
Ireland	13.4056	4.84
Greece	15.0613	0.01(2)
Belgium	7.3353	5.58
Finland	14.3417	13.42
Netherlands	2.6453(1)	12.56
Austria	11.8163	12.12

Notes: (1) Trace statistic for the Netherlands points out the existence of a cointegration vector between *Spread* and *IP*. (2) Trace statistic for Greece points out the existence of a cointegration vector between *Spread* and *Gr*.

Table A.4: Unit root test for residuals (H0 hypothesis: All Panel contain unit roots)

	Time dependent model (2a)		Time dependent model (2b)	
		(first difference model)		(first difference model)
inverse chi-squared (20)	21.934	673.933	14.614	645.444
	(0.344)	(0.000)	(0.798)	(0.000)
inverse normal	1.417	−24.724	1.909	−24.070
	(0.922)	(0.000)	(0.972)	(0.000)
inverse logit t(54)	1.431	−59.311	2.248	−56.804
	(0.921)	(0.000)	(0.986)	(0.000)
modified inv. Chi-squared	0.306	103.396	−0.852	98.891
	(0.379)	(0.000)	(0.803)	(0.000)

We considered a Fisher-type unit-root test for residuals based on augmented Dickey-Fuller tests. P-values in parentheses.

Table A.5: First difference model – estimation results

Variables	time dependent model (2)	
	2.a (using fiscal space)	2.b (using debt ratio)
Spread (in $t-1$)	0.974 (***)	0.964 (***)
fiscal space	−0.061 (***)	
fiscal space squared	0.007 (***)	
debt / GDP		−0.005 (**)
debt / GDP squared		0.008 (***)
GDP growth (in $t-1$)	−0.142 (***)	−0.156 (***)
industrial production	−0.004 (***)	−0.003 (**)
current account (in $t-2$)	−0.003	−0.004
real effective exchange rate (in $t-1$)	−0.000	0.001
liquidity (debt share)	−0.007	−0.004
GRA	0.002	0.005
debt*GRA post July 2011	−0.057	
time component		
2003	0.000	0.000
2004	0.000	0.000
2005	0.000	0.000
2006	0.001	0.001
2007	0.001	0.001
2008	0.001	0.000
2009	−0.001 (**)	−0.001 (**)
2010	0.001 (***)	0.001 (**)
2011	0.002 (***)	0.001 (***)
2012	0.000	−0.001
constant	0.006	−0.001
country fixed effect	controlled	controlled
Wald chi2 (degrees of freedom in parentheses)	(29) 66015.31	(28) 63038.52

Note: fiscal space, GDP growth, industrial production and current account balance were seasonally adjusted through a moving average (MA) filter; the length of the moving window was appropriately chosen depending on the time series. Such smoothing allowed us to obtain monthly estimated values for the variables, which were used in the estimation. (***) significant at 1% ($p < 0.01$), (**) significant at 5% ($p < 0.05$), (*) significant at 10% ($p < 0.1$).

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