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ABSTRACT

Nonlinearities in Sovereign Risk Pricing: The Role of CDS Index Contracts*

Is the pricing of sovereign risk linear during bearish episodes? Or can initial shocks on economic fundamentals be exacerbated by endogenous factors that create nonlinearities? We test for nonlinearities in the sovereign bond market of European peripheral countries during the debt crisis and explain them. Our estimates based on a panel smooth threshold regression model during January 2006 to September 2012 show four main findings: 1) Peripheral sovereign spreads are subject to significant nonlinear dynamics. 2) The deterioration of market conditions for financial names changes the way investors price risk of the sovereigns. 3) The spreads of European peripheral countries have been priced above their historical values, given fundamentals, because of amplification effects. 4) Two CDS indices on financial names unambiguously stand out as leading drivers of these amplification effects.

JEL Classification: C23, E44, F34, G12 and H63
Keywords: CDS indices, European sovereign crisis and panel smooth threshold regression models

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1 Introduction

From the beginning of the sovereign debt crisis in May 2010, the decade-long process of interest rate convergence in the euro-area reversed. Two distinct categories emerged, the peripheral and the core euro-area economies. Aside from Greece, the sharp rise of peripheral sovereign bond spreads and their volatility are hard to reconcile with the underlying economic fundamentals: spreads surged suddenly, while the economic conditions were deteriorating gradually[1]. We consider the hypothesis that amplification dynamics have driven sovereign risk during the crisis. Initial shocks on economic fundamentals may have been exacerbated by endogenous mechanisms. Is pricing of sovereign risk linear, or can we identify endogenous factors of amplification? The answers will help to monitor and price sovereign risk and to curb financial fragmentation.

There is now extensive theoretical research suggesting that the pricing of assets, including sovereign debt, may be nonlinear. Recent work stresses the importance of nonlinear effects and amplification dynamics through the price mechanism during financial crises (Brunnermeier and Oehmeke, 2009). On the one hand, the initial drop in asset prices will be exacerbated if it triggers fire-sale liquidations driven by the deterioration of the mark-to-market portfolio value. Theory suggests that relatively small shocks can imply large spillover effects (Brunnermeier and Pedersen, 2009). Moreover, Brock *et al.* (2009) show that proliferation of hedging instruments may produce nonlinear systems and destabilize markets. Here we test whether the credit derivatives market has amplifed the risk instead of mitigating it.

Previous empirical work has identified non-linearity in the spread determination model for euro-area peripheral sovereigns during the crisis (Aizen-

[1] In Spain, for example, the public debt amounted to less than 60% of GDP even by end 2009. The Italian primary budget surplus implied that if interest rates had stayed low, only modest fiscal adjustment would have been necessary to service the debt. Even invoking a broader set of economic fundamentals seems insufficient to explain the sudden eruption of the crisis. Unemployment and the trade deficit had been increasing gradually. And Ireland's trade balance had been improving at the time of the crisis.

man *et al.* (2011), Gerlach *et al.* (2010), Montfort and Renne (2012), Borgy *et al.* (2011), Favero and Missale (2011)). Two different regimes have been described, a crisis and a non-crisis regime, with additional fundamental factors important in the crisis regime. These papers usually attribute nonlinearities to the fiscal situation: they find that yield spreads have become much more sensitive to fiscal imbalances after 2008, with a deterioration of fiscal indicators generating a significant widening of the spreads after 2008. But the crisis may have other than fiscal roots. Attributing nonlinear dynamics only to fiscal imbalances, an exogenous driver, is questionable in the light of recent advances in *macro-finance*.

In this paper, we draw on recent research on financial crises to explore the endogenous drivers of nonlinearities in the sovereign bond markets of euro-zone peripheral countries. We explicitly test three hypotheses. First, we hypothesize that the nexus between sovereigns and banks observed during the crisis (Gennaioli *et al.*, 2010, Huizinga and Demirguc-Kunt, 2010, Acharya and Steffen, 2013) may have created a nonlinear relationship which goes both ways and features some amplification in the sovereign risk. Second, adverse liquidity effects on euro area banks have been documented during the crisis, including a significant fall of interbank loans after mid-2010 (Allen and Moessner, 2013). So we will examine the effects on sovereign risk of a negative externality due to fire-sale liquidation of assets by testing whether liquidity shocks have had self-reinforcing effects on sovereign bonds. Third, we explore the hypothesis that derivatives produce nonlinear systems (Brock *et al.* 2009, Simsek, 2013) by investigating the effects on the sovereign price of credit default swaps (CDS), the most active credit derivative market. Delatte *et al.* (2012) and Palladini and Portes (2011) have documented an adverse influence of the sovereign CDS on the underlying bond pricing when bearish investors use these instruments to express their views on the sovereign credit. But we know much less about the effects of corporate CDS on sovereign risk. In a down-cycl,e however, their effect on the cash market may feed back to the sovereign risk. To explore this hypothesis, we focus on *synthetic CDS indices* which cover default risk on various pools of corporate entities, because their standardization and liquidity make them the instrument chosen by investors to express views on market segments. We test whether a rise in corporate CDS spreads amplifies the risk of sovereigns. We compile a new set of financial variables that capture our three hypotheses, and we identify the best candidates explaining how initial shocks to fundamentals may be amplified.

Amplification can be modeled through increasing weights in the spread determination. In other words, the same change in a fundamental has a higher impact on the spread in the crisis period than it had previously. To capture this idea, we use the smooth transition regression model initially proposed by Terasvirta (1996) and developed in panel by Gonzalez *et al.* (2005). Contrary to the alternative family of nonlinear models, the Markov-switching (MS) models, STR model offers a parametric solution to account for nonlinearity by allowing the parameters to change smoothly as a function of an observable variable (MS specifications assume that the transition variable is unobserved). In this paper, we consider potential threshold variables to account for the time variability of the estimated coefficients, and we follow Gonzàlez *et al.* (2005) to identify the optimal threshold variable. In sum, our panel threshold regression framework establishes a ranking among three hypotheses that might give rise to amplification effects (Fouquau *et al.* 2008).

We estimate equations for the sovereign spreads of five European peripheral countries: Spain, Ireland, Italy, Portugal and Greece over the period January 2006 to September 2012. We deliberately end our sample at the beginning of the Outright Monetary Transactions (OMT) programme that has succesfully narrowed the spreads and blurred market signals[2].

Our estimates uncover four main findings. First, sovereign spreads are subject to significant nonlinear dynamics, a result that invalidates linear estimations of the sovereign spreads during this period. Second, the tests

[2] As Paris and Wyplosz (2013) have argued, "Spreads no longer show us what investors think about debt sustainability. They reflect a mix of debt-sustainability expectations and forecasts of ECB reactions".

reveal that *uncertainty* and *stress* on financial entities are major drivers of nonlinearities. The deterioration of market conditions for financial names changes the way investors price risk of the sovereigns. Third, we detect amplification effects in the spreads of the five peripheral countries, with heterogeneous dynamics, that our PSTR approach enables us to capture. Last, two CDS sub-indices on financial entities are leading drivers of nonlinearities. Their linearity rejection statistics are unambiguously higher than the twenty-two alternative variables. This result may stem from the high leverage created by these instruments. It seems that when active investors leverage their views on credit risk in the financial sector, this drives down the price of sovereigns. The financial risk feeds back to peripheral countries through CDS indices. This result suggests that sovereign bond investors and policymakers should carefully monitor the financial credit derivative markets.

The remainder of this paper is organized as follows. Section 2 reviews the determinants of the sovereign bond spread in a linear context and the theoretical arguments for the presence of nonlinearities. Section 3 reviews the existing empirical evidence of nonlinearities in the pricing of sovereign bonds during the European debt crisis. Section 4 introduces the PSTR specification methodology and the test procedure. Section 5 summarises our dataset, and Section 6 discusses the estimation results. Section 7 concludes.

2 The Determinants of Sovereign Bond Spreads

Linear context

The government bond yield spread represents the risk premium paid by governments relative to the benchmark government bond. The empirical literature has explored a large set of macroeconomic variables to explain sovereign spreads[3]. From a theoretical perspective, although sovereign debt is notably different from corporate debt, these instruments can be priced

[3]Early and influential empirical papers include Edwards (1986), Eichengreen and Portes (1989), Cantor and Packer (1995).

by decomposing the risk premium into credit risk and liquidity risk, a well-established distinction in the corporate context (Longstaff and Schwartz (1995), Duffie and Singleton (1999)).

Credit risk is influenced by variables that affect the sustainability of the debt and the likelihood of repayment. For a sovereign entity, these are macroeconomic variables determining internal and external balances, more precisely variables important in determining the budget deficit and the current account. The empirical evidence in the euro area context suggests that significant determinants include fiscal variables, activity-related and competitiveness-related variables (see Attinasi *et al.* 2009, Haugh *et al.* 2009, De Grauwe and Ji, 2012).

Liquidity risk is related to the size of the issuer, with an expected negative relationship due to larger transaction costs in small markets. In contrast with findings on credit risk, empirical evidence is mixed about the pricing of a liquidity premium in the sovereign bond spread[4].

Beyond these two theoretical risk premia, the growing influence of global factors on domestic financial conditions shown in recent work (Rey, 2013) points to the potential influence of international risk aversion. Borgy *et al.* (2012) find a significant impact of international risk aversion on the sovereign bond yield spread in the euro-area context.

While the spread determination model is assumed to be constant and linear in most studies, there is now substantial theoretical research suggesting that the pricing process of assets, including sovereign debt, may be nonlinear. In the following we review the different arguments to guide our empirical exploration in the subsequent sections.

[4]For example, Geyer *et al.* (2004) finds that liquidity plays a minor role for the pricing of EMU government yield spreads. Favero *et al.* (2009) find that investors value liquidity, but they value it less when risk increases.

Theoretical arguments for nonlinear pricing of sovereign bonds

In the European sovereign crisis, we may observe a feedback loop between sovereigns and the banking sector, which may imply nonlinearities in the pricing of sovereign bonds. On the one hand, Gennaioli *et al.* (2010) argue that the sovereign risk affects the banks through their exposure to sovereign bonds. Huizinga and Demirguc-Kunt (2010) provide evidence in a large cross-country sample that bank CDS spreads responded negatively to the deterioration of government finances in 2007-08. Acharya and Steffen (2013) find that the Eurozone banks actively engaged in a 'carry trade' in the crisis period, increasing their exposure to risky sovereign debt. On the other hand, bank risk affects the sovereigns, which are expected to bail out systemically important institutions (Acharya *et al.* , 2011). That represents a significant risk given the size of banks compared to the size of the public backstop.

We hypothesize that the nexus between banks and sovereign creates a nonlinear relationship which goes both way and features some amplifications. Bernanke et al. (1999) show that an adverse shock to the economy is amplified through the credit channel. The resulting weakening of the economy may affect the sovereign risk in a nonlinear manner. This is more likely to happen when a shock on financial intermediaries forces them to restrict the quantity of credit, and that weakens the economy[5]. This reduces fiscal revenues and raises sovereign risk, which leads to a deterioration of bank balance sheets. Recently, Coimbra (2014) has explicitly modelled the resulting feedback loop. After a rise in sovereign risk, the banks' VaR constraint binds, which reduces their demand for sovereign bonds, thereby raising the sovereign risk premium. This in turn leads to adverse sovereign debt dynamics, which raises sovereign risk. The initial shock is exacerbated and feeds back to credit conditions. Borrowing costs deteriorate further, causing more credit restrictions. Highly-leveraged investors are more vulnerable to initial shocks and forced into credit restrictions to a greater extent.

[5]From 2010, US prime money market funds have strongly reduced their exposure to euro area banks and stopped lending to banks from peripheral countries (IMF, 2012b)

Although this model admits of quantification that supports the theory, we do not yet have a model of simultaneous determination of sovereign and bank risk that is amenable to econometric specification. In this work, we will explicitly test whether the rise in the risk of the banking sector amplifies the sovereign risk.

In addition, there are specific pricing mechanisms that may drive nonlinearities in the sovereign bond market. The most documented such mechanism is due to liquidity problems implying self-amplifying dynamics in asset prices, because of the negative externality due to fire-sale liquidation of assets[6]. An initial shock on sovereign bonds may trigger a liquidity spiral because it degrades the quality of collateral[7]. Banks facing liquidity problems will be forced to sell off assets to regain liquidity or restore their capital ratio. The emergence of asymmetric information frictions strengthens the dynamics (Brunnermeier *et al.*, 2009). The pricing of debt becomes more "information sensitive", and safe assets become less safe, so investors are more selective about the quality of assets they accept as collateral. Their demand for the sovereign bonds that are perceived to be more risky declines, thereby raising the sovereign risk premium. So there is a liquidity spiral: a falling sovereign bond market leads financial intermediaries to fly to liquidity, and this amplifies the effects of the initial price reduction. Relatively small shocks can cause liquidity suddenly to dry up, leading to a major correction of asset prices (Brunnermeier and Pedersen, 2009). In this work, we will test a second hypothesis, that liquidity shocks have self-reinforcing effects on sovereign pricing.

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[6]Stiglitz (1982) and Geanakoplos and Polemarchakis (1986) initially pointed out this externality.

[7]As an example of a fire-sale driven by more restrictive collateral requirement, the *Financial Times* reports in November 2010 "[...] a sell-off in Irish bonds was driven by a fire sale of positions by market participants who were unable to meet collateral requirements enforced by LCH.Clearnet- one of Europe's biggest clearing houses [...]. Ireland's banks were faced with an estimated \$1bn cash-call from LCH.Clearnet as a result of its decision to require a deposit of 15 per cent against all Irish bond positions as an indemnity against default."

Finally, we explore the hypothesis that derivatives produce nonlinear systems, and a proliferation of hedging instruments may destabilise markets (Brock *et al.*, 2009, Geanakoplos, 2010, Simsek 2013). Using credit derivatives, investors can hedge the credit risk of a reference entity and also express negative credit views more easily than with cash assets. Short sales of credit instruments can be executed with reasonable liquidity and a lower risk of suffering from a short squeeze (Tavakoli, 2008). In the corporate sector, where the liquidity in the cash market is sometimes limited, Blanco *et al.* (2005) and Baba and Inada (2009) have shown a lead for CDS prices over credit spreads in the price discovery process, implying that bearish investors who express their negative views on certain entities trigger self-reinforcing dynamics. Up-front principal is small or zero in most derivatives, implying that these instruments create high leverage. We know from Adrian and Shin (2010) that high leverage enhances the amplification of the initial correction. Thus by offering a new technology allowing bearish investors to express their negative views, credit derivatives amplify the effect of an initial correction on a market segment (Geanakoplos, 2010). Although we have evidence that the sovereign CDS market influences the underlying sovereign bond pricing during bearish episodes (Delatte *et al.* 2012, Palladini and Portes, 2011, and references cited there), we know much less about the effects of corporate CDS on sovereign risk. In a down-cycle, their effect on specific corporate markets may affect the sovereign because of the resulting weakening of the sector activity, thereby reducing fiscal revenues and potentially giving rise to counter-balancing expansionary fiscal measures that further deteriorates the fiscal balance. This is more likely to happen in corporate sectors with a large contribution to the domestic output and/or a large employment share, or in the financial sector because of the nexus between banks and sovereigns described before.

Given this background we explore the link from corporate CDS spreads to the sovereign spreads and test a third hypothesis: that the rise in corporate CDS prices has amplified the risk of sovereigns.

There are various theoretical reasons why the pricing of sovereign bonds may be nonlinear. In the following we briefly explore the empirical evidence of nonlinearities in the pricing of sovereign bonds during the European debt crisis.

3 Empirical evidence of nonlinear sovereign bond spreads during the European debt crisis

Non-linearity in the spread determination model of peripheral members of the euro-area during the crisis has been seen in previous work (Aizenman *et al.* (2011), de Grauwe and Ji (2013), Gerlach *et al.* (2010), Montfort and Renne (2011), Borgy *et al.* (2011), Favero and Missale (2012)). Two different regimes have been described, a crisis and a non-crisis regime, with additional fundamental factors important in the crisis regime.
Montfort and Renne (2011) model the joint dynamics of euro-area sovereign yields with a Markov-switching specification and find a regime-switching feature at the origins of the large fluctuations during the crisis. In particular, they identify a crisis regime that captures the rise in volatility experienced by the sovereign bond market since 2009.

Borgy *et al.* (2011) examine the macroeconomic determinants of risk premia in the sovereign yield spreads of six euro-area members and give special emphasis to fiscal sustainability measures. They show a structural break in the relationship between sovereign spreads and fiscal determinants in 2008: spreads have become much more sensitive to fiscal imbalance measures after 2008, so a given deterioration of the debt service ratio generates a significantly larger widening of the spreads after 2008. Overall the risk perception has changed during the sovereign crisis, with the deterioration of fiscal balances gaining a major role after 2008. This result underlines the increasingly important constraint on fiscal policy imposed by the financial markets, an observation of Haugh *et al.* (2009), who similarly find that incremental deteriorations in fiscal performance have led to larger increases in the spreads of euro area countries after 2008.

These empirical papers have detected nonlinearity in the sovereign bond spread determination model of euro-area members. The models employed in these papers reveal nonlinear dynamics on the weights of fiscal factors but do not explain the sources of nonlinearity. The regime shifts are due to an unobservable variable. Our objective is to relax linearity and allow the spread determination model to change according to an *observable signal* that sets off amplifying spirals. The next Section presents our empirical strategy.

4 Empirical strategy: specification and estimation

We estimate sovereign bond spread determination using a panel smooth threshold regression (PSTR) model developed by Gonzales *et al.* (2005). The choice of panel data is motivated by the low time dimension of macroeconomic data. The PSTR model allows us to characterize nonlinearity as a function of an observable variable. More precisely, the sovereign spread can be estimated as follows:

$$S_{it} = \mu_i + \beta_1' X_{it} + \beta_2' X_{it} g(q_{it}; \gamma, c) + u_{it} \tag{1}$$

for $i = 1, \ldots, N$ and $t = 1, \ldots, T$ where μ_i represents individual fixed effects, X_{it} is a set of variables that capture credit risk, liquidity risk and international risk aversion and u_{it} are i.i.d. errors. $g(.)$ is a continuous transition function bounded between 0 and 1. We use a logistic function of order 1 that has an S shape:

$$g(q_{it}; \gamma, c) = \frac{1}{1 + \exp\left[-\gamma(q_{it} - c)\right]} \quad , \gamma > 0. \tag{2}$$

where q_{it} is the observable threshold variable. The γ parameter determines the smoothness, *i.e.,* the speed of the transition from one regime to the other, and c the location parameter, which shows the inflexion point of the transition. The higher the value of the γ parameter, the faster (*i.e.*, sharper) the transition. This specification allows a smooth transition between two extreme regimes defined by the vectors β_1' and $\beta_1' + \beta_2'$. For example, if we take a threshold variable that proxies *flight to liquidity*, the higher this proxy, the closer the coefficient gets to $\beta_1' + \beta_2'$. The PSTR model is a way to account for individual heterogeneity (Fouquau *et al.*, 2008).

The estimation of the PSTR model consists of several stages. In the first step, a null hypothesis of linearity is tested against the alternative hypothesis of a threshold specification. Then, if the linear specification is rejected, the estimation of the parameters of the PSTR model requires eliminating the individual effects, μ_i, by removing individual-specific means and then applying nonlinear least squares to the transformed model (see Gonzàlez *et al.*, 2005).

In the Gonzàlez *et al.* (2005) procedure, testing linearity in a PSTR model (equation 4) can be done by testing $H_0 : \gamma = 0$ or $H_0 : \beta_0 = \beta_1$. In both cases, the test is non-standard since the PSTR model contains unidentified nuisance parameters under H_0 (Davies, 1987). The solution is to replace the transition function, $g(q_{it}; \gamma, c)$, with its first-order Taylor expansion around $\gamma = 0$ and to test an equivalent hypothesis in an auxiliary regression. We then obtain:

$$S_{it} = \mu_i + \theta_0 \ X_{it} + \theta_1 \ X_{it} q_{it} + \epsilon^*_{it}. \tag{3}$$

In these auxiliary regressions, parameter θ_1 is proportional to the slope parameter γ of the transition function. Thus, testing linearity against the PSTR simply consists of testing $H_0 : \theta_1 = 0$ in (3) for a logistic function with the usual LM test. The corresponding LM statistic has an asymptotic $\chi^2(p)$ distribution under H_0.

Before proceeding to the estimation, we present our data.

5 Data description

In this Section we present our dataset and sources used to estimate the linear model of sovereign bond spreads and to construct the threshold variables that capture the forces described in Section 3.

The estimation of the model of Eq.(4) is subject to two major data constraints. On the one hand, macroeconomic fundamentals have a low frequency (annual, quarterly or monthly), while our financial data are daily. Therefore we transform all series to monthly data. We calculate the monthly average of the daily series and we transform quarterly to monthly using a local quadratic transformation with the average matched to the source data[8]. On the other hand, the sovereign crisis started in late 2009, and the Outright Monetary Transactions (OMT) programme implemented in September 2012 succesfully narrowed the spreads and blurred market signals. So we have only three years during which the hypothesized transition might have occured. Therefore, to obtain a sufficient number of observations, our estimation is based on a balanced panel of the five peripheral European countries in which the sovereign yield has been most under pressure (Greece, Ireland, Italy, Spain and Portugal) between January 2006 and September 2012.

Determinants of the sovereign bond spread

Our dependent variable is the sovereign bond spread, which prices the default risk of a country. It is defined as the difference between the sovereign bond yield and the risk-free rate of the same maturity. For each country in the sample, we use the long-term German yield, which is the benchmark risk-free rate for the Euro area (Dunne *et al.*, 2007), and the government yield of this country at the same maturity. We rely on daily observations of 10-year bond yields provided by Bloomberg, from which we compute a monthly average[9]. All data described in this Section are plotted in Figure 1.

A key choice is the set of explanatory variables included in X_t in Eq (4). As mentioned in Section 1, we need variables to capture credit risk, liquidity risk and international risk aversion. We test the following variables: debt-to-GDP ratio, deficit, unemployment, unit labor cost, risk, liquidity.

First, the country's credit risk is traditionally related to fiscal sustainability. We include the debt-to-GDP ratio and fiscal deficit from Eurostat. We add the squared value of the debt-to-GDP ratio to capture non-linear dynamics that might be due to threshold effects of sovereign debt on real

[8]We used Eviews software for this transformation.

[9]For Ireland only 8-year bond yields are available, so we computed the spread using the 8-year German yield.

growth. The fiscal data are revised data, necessary because of the presence of Greece in the sample, although these are not the data initially observed by market participants. Other relevant variables are economic activity and the country's competitiveness. We proxy economic activity using the unemployment rate rather than GDP to avoid collinearity with the debt-to-GDP ratio. The unit labor cost and trade balance are included to proxy the country's competitiveness[10].

Second, we include a variable for liquidity risk, proxied as the bid-ask spread of the dependent variable and alternatively measured by market size, as the country's share of total outstanding Euro-denominated long-term government securities issued in the Euro zone. Data are available on a monthly basis from the European Central Bank (ECB).

Third, we include the CBOE Volatility Index (VIX) as a measure of international risk aversion, because it is often considered by many to be the world's premier barometer of investor sentiment and market volatility (*e.g.*, Rey, 2013).

Last we control for the effect on peripheral spreads of non-standard monetary measures adopted by the ECB during the crisis. In May 2010, the ECB decided to start the Securities Markets Programme (SMP) with large securities purchases in order to address tensions in certain market segments[11]. We use the amount of securities held for monetary purposes, as shown in the ECB's weekly financial statements, and including Securities Market Program, 1st and 2nd Covered Bond Purchase Programs (available in ECB Statistical Data Warehouse)[12].

[10]All data are available at a quarterly frequency, except for unemployment (monthly) and fiscal deficit (annual).

[11]The SMP was terminated in September 2012 in favour of Outright Monetary Transactions (OMTs) in sovereign secondary bond markets.

[12]On the other hand, the ECB provided in December 2011 and March 2012 more than 1 trillion Euros of additional liquidity to the financial system with the very longer-term refinancing operations (LTRO). Unfortunately publicly available data are not broken down by country, which makes the inclusion of the data composed of two observations irrelevant

Endogenous drivers of nonlinearities, three hypotheses

We construct a set of financial data to capture our three hypotheses presented in Section 2. They represent the set of threshold variables that we will include alternatively in our nonlinear estimations in the next Section. All threshold variables are plotted in Fig. 2.

1. *Feedback loop from banks to sovereigns*

   In order to test if the rise in the risk of the banking sector amplifies the risk of sovereigns, we create indicators of uncertainty and stress in this sector.

   - *IVolbank* used in Hakkio and Keeton (2009) denotes the idiosyncratic volatility of bank stock prices. It serves as an equivalent of the VIX for the banking industry rather than for the corporate sector as a whole. It is computed as the standard deviation of residual returns from a CAPM regression using an aggregate European banking sector price index and the S&P Europe 350 taken from Datastream.
   - *CMAXFin* is an indicator of stress widely used by market practictioners to identify periods of extreme price declines (Patel and Sarkar (1998)). We take the five domestic banking stock indices from Datastream and calculate *CMAXFin* as the maximum cumulated index losses over a moving two-year window with $Cmax_t = 1 - \frac{P_t}{max[P_{t-24}...P_t]}$. The more bearish the market, the closer to 1 the indicator.
   - An additional useful indicator of stress in the banking system is the *Euribor-OIS spread*, calculated as the difference between the Euro Interbank Offered rate and the overnight indexed swap rate. This indicator must be taken with some caution because of the alleged manipulation of the Euribor rate.

   We control for an overall effect of uncertainty and stress not due to the banking sector by including indicators on non-financial

in our panel estimates.

sectors:

- *Vstoxx* is the European equivalent of the VIX, considered by many to be the leading measure of market volatility[13].
- *FTSE300* and *S&P350* denote the return of the European aggregate stock market indices
- *DomsticIndex* is the matrix of the domestic stock returns indices of the five countries in our panel (PSI, IBEX, ATHEX, FTSEMIB, ISEQ).
- *RvolGerm* captures bond market volatility using the 10-year German government bond index. It is the realized volatility computed as the monthly average of absolute daily rate changes.
- *Rvol Nonfi* is the realized volatility of domestic non-financial sector stock market indices taken from Datastream.
- *Rvoldoll*, *Rvolyen* and *Rvolpound* are the realized volatility of three bilateral euro exchange rates for the US dollar, the Japanese yen and the British pound respectively.

2. *Nonlinear effects of credit derivatives*:

   In this paper we focus on the most active credit derivatives, the credit default swap market (CDS). In particular, a significant development has been the creation of synthetic CDS indices which cover default risk on a pool of entities as opposed to single-name CDS. Buying an index is equivalent to selling protectin. Therefore buying and selling the indices can be compared to buying and selling portfolios of bonds. A buyer takes on the credit exposure to the bonds, *i.e.* is exposed to defaults, similar to buying a cash portfolio. Investors can express views on a specific market segment via CDS indices. The main advantages of these new classes of credit derivatives are standardization and liquidity, which explain their growth. CDS indices accounted for 43% of gross notional amount of the CDS market in December 2012,

[13]We use Vstoxx to proxy the European market volatility, while we use VIX to capture international risk aversion.

up from 20% in 2004 (Vause, 2011)[14].

We take the i-Traxx Europe, a broad tradable credit default swap family of indices traded by Markit. Most widely traded is i-Traxx Europe Main index (125 investment grade credits), further segmented into sub-indices defined by industry groups and trading levels. We include:

- *i-Traxx Europe* comprises the most liquid 125 CDS referencing European investment grade credits
- *X-over* comprises the most risky 40 constituents
- *HiVol* is a subset of the main Europe index consisting of what are seen as the most risky 30 constituents
- *SenFin* comprises constituents from the sector of senior-financials and *SubFin* from the sector of sub-financials.

3. *Negative externality due to fire-sale liquidiation*

We use standard indicators of *flight to liquidity* complemented by indicators of *flight to quality* and *asymmetry of information* because they occur simultaneously during a liquidity run and strengthen self-amplifying dynamics (as put in Section 2).

- *Aaa/10-year Treasury spread* denotes the spread between European corporate bonds rated Aaa and the 10-year German Treasury bond. It is a standard measure of liquidity premium, because even the highest-rated corporate bonds tend to be less liquid than Treasury securities. All corporate bond indices are Markit i-boxx European corporate bonds, taken from Datastream.

[14]CDS trading has continued to grow after 2007 (IOSCO, 2012). At the end of 2012, the gross notional value of outstanding CDS contracts amounted to approximately 25 trillion US dollars, and the corresponding net notional value to approximately 2.5 trillion US dollars. The fact that the gross notional value of the CDS contracts has more than halved since the peak of 2007 (with 60 trillion US dollars) is mostly attributed to the development of compression mechanims that eliminate legally redundant contracts (Vause 2011).

- *High-yield bond/Baa spread* denotes the spread between "junk bonds", *i.e.* bonds with too low a rating to be considered investment-grade, and Baa-rated corporate bonds, the lowest-rated bonds considered as investment-grade. High-yield bonds are issued in smaller quantities and traded by a limited set of investors (institutional investors are banned from the market) in comparison with Baa-rated bonds, implying a liquidity premium to compensate investors for holding the less liquid asset.
- *10-year swap spread*. The fixed-rate payment leg of a swap is expressed as the Treasury yield plus a spread that compensates investors for the fact that claims on fixed-rate payments are considerably less liquid than Treasury securities.

These variables also capture *flight-to-quality* because they all include a default risk premium. In addition, the indicator *High-yield bond/Baa spread* captures also *asymmetry of information*, because a rising value partially stems from adverse selection problems when investors have difficulty in determining which bonds are riskier than others during financial crisis episodes (Hakkio and Keeton, 2010).

- *Flight to quality* is also proxied with the indicator *StockbondsCorr* which measures the correlation between domestic stock total return indices and the total return German Treasury index. It is well-documented that the correlation between stock and government bond returns is usually significantly negative during financial crises, because investors consider government bonds safer (Andersson *et al.* 2008). We compute the correlation over rolling three-month periods using the domestic stock index of each country of our panel and the 10-year German government bond index taken from Datastream. We use the negative values of the correlations, so that an increase in the measure corresponds to higher *flight-to-quality*.
- The *asymmetry of information* is also measured using *cross-section*

dispersion bank proposed in Hakkio and Keeton (2010) and computed as the cross-section dispersion of bank stock returns to capture uncertainty about the relative quality of banks. The intuition is that the larger the cross-section dispersion, the larger proportion of returns is unexpected, so the larger the information asymmetry. It is calculated using daily data on the S&P Europe 350 and the stock prices of the 82 largest commercial banks in terms of market value[15].

It is worth making a general comment looking at the set of threshold variables plotted in Fig. 2 : most variables experienced a first peak during the subprime crisis, followed by a second peak due to the sovereign debt crisis in Europe. Thus our financial series capture two episodes of crisis, contrary to our dependent variable, which is mostly affected by the second episode. This pattern represents a methodological challenge to detect the drivers of nonlinearity during the European crisis. In the following, we present our results.

6 Estimation results: Nonlinear dynamics in the European sovereign market.

We recall that the PSTR specification of the spread is as follows:

$$S_{it} = \mu_i + \beta_1' X_{it} + \beta_2' X_{it} g(q_{it}; \gamma, c) + u_{it} \tag{4}$$

for $i = 1, ..., n$ and $t = 1, ..., T$, X represents the vector of determinants, μ_i the country fixed effects, $g(.)$ the threshold function, q_{it} the threshold variable, γ the smooth parameter, c the location parameter.

[15] More precisely we estimate a CAPM regression of the daily return on each bank's stock index against the daily return on the S&P Europe 350 index, using data for the previous 12 months. The estimated coefficients are then used to calculate the forecast errors of the current month. Last we calculate the interquartile range for these residuals in order to keep the central 50%. The lower the interquartile value, the smaller the dispersion across banks.

Selection of the optimal linear model

First, we proceed to the linear estimation using a panel estimation with fixed effects. The first step is to select the optimal linear model. We use alternative series in the vector of explanatory variables and select the optimal combination based on standard selection criteria. Results displayed in Table 1 suggest that our specifications are robusts with similar estimated values in different specifications. The information criteria suggest that specifications 1 and 2 could both be considered as optimal (Schwarz = 0.207, AIC = 0.197), and we keep specification 2 because it includes only factors found significant.

With a negative and a positive coefficient respectively, the evolution of the sum of Debt and squared Debt is ambiguous, while trade balance is not significant. In turn, as expected, unemployment and international risk aversion have an upward impact on the spread: a rise in unemployment and in the VIX increase the sovereign spread. Liquidity effects are properly captured by our measures based on the bid-ask spread (an increase in the bid-ask spread increases the sovereign spread) and volume (a reduction of outstanding issues increases the spread). We keep both in the vector of determinants because information criteria are systematically better when both measures are included.

In addition, as in other studies (De Grauwe and Ji, 2012, Wyplosz, 2013), we find that competitiveness is not relevant: the unit labor cost has an unexpected sign (higher labour cost reduces the spread) while the trade deficit is never significant.

Last, in all specifications the secondary market program adopted by the ECB has no effect, suggesting that the liquidity injections have no direct effect on the spread. In sum, market forces still dominate.

In the following we adopt a parsimonious approach and proceed to the tests and nonlinear estimation of specification 2.

Linearity tests: the feedback loop played a significant role

In the second step, we test this linear specification of the spread (spec 2) against a specification with threshold effects. We select the best threshold variables, with the objective of identifying the drivers of nonlinear effects. As suggested by Gonzàlez et al. (2005), the "optimal" threshold variable corresponds to the variable that leads to the strongest rejection of the linearity hypothesis.

The linearity test results reported in Table 2 clearly reject the null hypothesis of a linear relationship, regardless of which threshold variable is included in the specification. The remarkably high level of rejection in most models makes the presence of nonlinear dynamics unambiguous. This is consistent with previous empirical work mentioned above and confirms that it is inappropriate to use linear models to estimate sovereign spreads during this period.

The ranking of the test statistics reveals that the feedback loop hypothesis unambiguously stands out (*CmaxiFi*, Euribor-OIS and *IVolBank* reject linearity with 194.2, 119.4 and 116.2 resp.). Second, two CDS indices composed exclusively of financial constituents *CDSSnrFin* and *CDSSubFin* rank among the highest rejection statistics (148.3 and 130.9 resp.).

In sum, the tests reveal that investors are sensitive to the risk in the banking sector, and this triggers nonlinear dynamics. The pricing model is a nonlinear function of fundamentals, where the weight of these fundamentals varies with the risk of banks (we examine the evolution of the estimated coefficient below). The deterioration of market conditions for banks changes the way investors price risk of the sovereigns. It is interesting to observe that indicators of uncertainty about the *non*-financial sector, *rvol NonFin* and *rvol Germ*, rank among the last with low statistics (39.5 and 30.4 resp.). These results illustrate how deeply the sovereign debt crisis is intertwined with the banking crisis (Lane, 2012). This pattern is confirmed with the financial CDS indices, which price the risk on financial entities. It is interesting to observe that CDS are not only a good measure of risk but short positioning vehicles used by investors to express their views on credit. Their strong rejection statistics may not only confirm the feedback loop hypothesis but also indicate an adverse influence of these instruments on the sovereign risk. We discuss this below in the reported estimates.

The hypothesis about adverse liquidity effects does not get much empirical support. Most indicators of such effects included in our model get low rejection statistics in comparison with first-ranked indicators analyzed previously (see column 1-3 in Table 2). This result suggests a major difference between the US subprime and the European debt crises: while *flight to liquidity* and *quality* and *asymmetry of information* are unambigously relevant factors of amplification in the subprime crisis (Gorton and Metrick, 2012)[16], our estimates suggest that such concerns have played a minor role during the European debt crisis, in comparison with the banking sector's balance-sheet effects and the subsequent feedback-loop dynamics to the sovereigns.

Last the tests reveal that the volatility of different market segments play a minor role in nonlinear dynamics. While the volatility of FTSE and S&P get a fairly high rejection statistics (111.8 and 111.6), other volatility measures such as Vstoxx do not confirm the effect of overall volatility (LM= 64.2). This suggests that aggregate equity indices correlate with bank stocks indices and thus convey a similar information. Last, volatility of the foreign exchange market is not a relevant factor of nonlinearity (*rvol Pound*, *rvol Doll* and *rvol Yen* get 49.5, 40.0 and 43.9 resp) because intra-Euro zone, not extra-Euro zone capital transfers have been relevant since 2010 (IMF, 2012a). Peripheral countries have suffered massive capital flight back to the core countries, resulting in monetary fragmentation of the euro-zone. But the aggregate external position of the eurozone has not deteriorated significantly.

In the last step of our empirical investigation, we examine more precisely

[16]Gorton and Metrick (2012) have documented spillovers and endogenous responses of other market participants due to concerns about market liquidity in the Fall of 2008.

the impact of these variables on the determination of sovereign spreads. We consider which determinants have their weight changed most when the identified drivers of nonlinearity deteriorate.

Estimation results: a rise in the risk of CDS financial subindices amplifies the sovereign risk

Table 3 reports the linearity test statistics, the smooth parameter, γ, the location parameter c and the residual sum of squares in the three specifications that best reject linearity.

The optimal specification is the one in which *CmaxFi* is the threshold variable because it rejects linearity with the highest statistics (Gonzàlez *et al.*, 2005). In addition, this threshold variable minimizes the information criterion (Schwarz : -0.485 *versus* -0.267 and -0.140 for *CDSSnrFin* and *CDSSubFin* respectively). In this specification the smooth parameter is high ($\gamma = 111.4$), implying a sharp transition between two extreme regimes. This variable *CmaxFi* captures the heterogeneity in our sample. In fact, Italy, Spain and Portugal remain exclusively in the first regime (in these countries *CmaxiFi* is always lower than the estimated location parameter $c = 0.819$ as shown in Fig 2, graph *Cmax Financials*), while Ireland and Greece went from the first to the second regime (47 and 12 observations respectively as shown in Fig 2). Heterogeneity is confirmed in the other specification including an individual threshold variable, *Ivol Bank*, with similar patterns: the transition is sharp ($\gamma = 141.0$), and only Ireland and Greece went from the first to the second regime (27 and 12 observations respectively as shown in Fig 2, graph *Ivol Bank*).

Therefore, while the five peripheral countries are usually gathered in the same bundle, our estimates suggest that their spreads have a different dynamics. This finding leads us to split our sample into two sub-samples, one including Italy, Spain and Portugal, the other Greece and Ireland. The smaller sub-sample still has 162 observations, which is sufficient for reasonably precise and stable estimates.

We re-estimate the model using the three previous threshold variables in each sub-sample (Table 4). Linearity is strongly rejected again, but the sub-sample estimates indicate a different ranking from the full sample. In fact, in both samples, *CDSSnrFin* and *CDS Sub-Fin* best reject linearity (LM =88.2/82.8 and 67.3/61.9 resp), while *CmaxFi* ranks lower. This result confirms that the individual variable *CmaxFi* was mostly capturing heterogeneity in our previous estimates (as probably was *IvolBank*). In turn, *CDSSnrFin* and *CDS Sub-Fin*, which are two homogeneous variables, account for the time-instability in the spread determination model. In other words, the prominent driver of nonlinearity in the bond determination model is the price of a corporate CDS index that covers financial names. Now we examine the evolution of the coefficients to identify whether amplification effects have affected the spreads. To do so, we adopt a general-to-specific modelling approach where we eliminate non-specific variables based on their statistical significance and the Schwartz information criterion.

Italy, Portugal and Spain

Results in Table 5 report the estimated coefficients in regime 1 and regime 2 ($\hat{\beta}_1$ and $\hat{\beta}_1 + \hat{\beta}_2$) of the bond spread determination models of Italy, Portugal and Spain, in which *CDSSnrFin* and *CDS Sub-Fin* are threshold variables. We examine the transition of the estimated coefficients along the CDS indices variation. Table 5 indicates that the transition from the first to the second regime is sharp ($\gamma = 53.7$) and the threshold value, c, representing the switching date of the transition process, is located in autumn 2010. Our model thus correctly captures the increase in market tensions about the European sovereigns in 2010 after the Greek crisis broke. The spread determination model for these countries appears to have changed radically in autumn 2010[17]. Recall that amplification can be modeled through increasing weights in the spread determination.

[17]In the alternative model $\gamma = 2.18$ which corresponds to a sharp transition too, see Table 5.

In fact, our estimates suggest amplification effects that operate in regime 2 through a stronger influence on the spread of all macroeconomic determinants: debt, fiscal balance and unemployment as well as the international risk aversion ($| \hat{\beta}_1' + \hat{\beta}_2' | > | \hat{\beta}_1' |$). In other words, when the price of the sub-index iTraxx *CDSSnrFin* deteriorates and exceeds 135.7 bp, the weight of these fundamentals increases in the determination model, so the shocks to fundamentals have more effect on the bond spread. In turn, the influence of liquidity is ambiguous because the coefficients of both variables capturing liquidity show two contrary movements in regime 2 : we find a stronger negative influence of the relative stock of outstanding debt (implying that a deterioration of liquidity affects the spread more in regime 2 than in regime 1), while the influence of the bid-ask spread is lower in the second regime ($| \hat{\beta}_1' + \hat{\beta}_2' | > | \hat{\beta}_1' |$, implying that a rise in the bid-ask spread affects the spread less in regime 2). In addition, we observe that the sign on unit labor cost is contrary to the expected sign, as in the linear estimates (see Table 1)[18]. Last, we observe that the influence of the SMP program has not changed during the crisis and remains null.

To check the robustness of our estimates, we proceed to alternative estimates. First, overall amplification effects are confirmed when *Cmax Fin* is used as a threshold variable in an alternative specification (see Table 6)[19]. Second, financial CDS and sovereign bonds may price the same information, which would raise an endogeneity bias due to simultaneity. To address this, we re-estimate our optimal model by lagging the threshold variable. Linearity is rejected with a similar statistic ($LM = 63.2$ versus 62 in the core estimate), and amplification effects are confirmed. Last, we check that our nonlinearity finding does not result from omitting the CDS index as an explanatory variable. Our results are not affected by the introduction of

[18]As in the linear estimates, models excluding this variable have a lower RSS, so we decide to keep it in the vector of explanatory variables.

[19]We observe that the combined influence of debt and squared debt increases in regime 2 as well as the weight of fiscal balance and unemployment. Only the influence of VIX, which is found to be stable, differs from the core estimate.

the CDS index in the specification, a result that confirms that this variable nonlinearly affects the sovereign bond pricing[20].

Greece and Ireland

Results of the second sub-sample including Greece and Ireland are reported in Table 7. The slope parameter is low ($\gamma = 0.43$), and this transition occurs in autumn 2010, consistently with the previous sub-sample[21].
Amplification effects also operate through a stronger influence of unemployment. The effects of debt and squared debt are contradictory and compensate for each other. The effects of the VIX and of the bid-ask spread are positive, as expected, but they remain stable in the second regime, contrary to the previous sub-sample. As in the previous sub-sample and in the linear estimate, the unit labor cost has the same unexpected sign. Last, contrary to the previous sample, we observe that the SMP has a negative effect on the spread in the second regime ($\hat{\beta}_1' + \hat{\beta}_2' < 0$). In other words, our estimates suggest that the bond purchases carried out by the ECB have counterbalanced amplification effects on the bond spreads of Greece and Ireland. Similarly to the previous panel, we have proceeded to alternative estimates displayed in Table 8. Model 1 confirms the stronger influence of debt and unemployment and indicates a stronger influence of liquidity, a result not uncovered in the core estimates. The downward influence of the SMP is confirmed too.

Thus the spread determination model changed during the crisis, and amplification effects are detected in both sub-panels, where initial shocks on fundamentals are exacerbated when the price of financial CDS sub-indices increases. The higher the CDS price, the more risk-averse are investors towards the peripheral countries. Figure 2, which plots the evolution of both

[20] Results available on request.

[21] The lower slope parameter may indicate a slower transition than in the previous sub-sample but it is also important to note that the transition speed depends on γ and the distance between the threshold variable and the threshold parameter c. The fact that CDS indices increase strongly during the crisis implies that the transition from one regime to the other is fairly fast, as in the other sub-sample.

financial CDS sub-indices, shows that their prices experience a first peak due to the subprime crisis and then rise progressively to reach a second peak, significantly higher in 2012, when peripheral sovereign risk holdings of European banks put the entire euro system at risk.

We mentioned in Section 2 that high leverage enhances the amplification of the initial correction. Recall that the up-front principal in buying CDS indices is small or zero, implying a high leverage created by these instruments. Market anecdotal evidence reports that from May 2010, some traders have taken positions on the i-Traxx Financials to leverage their views on credit risk in the financial sector due to rising sovereign risk[22]. Our results suggest that this has reinforced the risk of sovereigns because the financial risk feeds back to the peripheral countries through CDS indices. We conjecture that the large amplification effects detected by our model result from high leverage created by CDS indices and their late introduction into the market. In the context of the subprime crisis, Geanakoplos (2010) stresses that the late introduction of standardized CDS contracts into the mortgage market in 2005 precipitated its downturn because the derivatives allowed the pessimists to leverage their credit views[23]. Similarly, it is interesting to observe that standardized CDS contracts on European corporate names were introduced in 2008 when Markitt launched the Europe i-Traxx index. The implication was that bearish investors had an opportunity to leverage after the market reached a peak, which magnified the depression of financial name prices in the context of the feedback loop between banks and sovereigns[24]. This result suggests that sovereign bond investors should carefully monitor the credit derivative market.

[22] As an example, ETF.com, a publication focused on financial indices, reports in June 2011: "The two indices have been closely correlated -sovereigns have bailed out banks and banks are holding government debt." (the i-Traxx SovX Western Europe includes the 15 most liquid sovereign CDS constracts)

[23] Geanakoplos (2010) points that the "midstream" introduction of CDS magnifies the fall in prices while their introduction from the beginning of the market moves the markets closer to completeness.

[24] This result is consistent with previous findings that CDS have destabilizing effects on the underlying market (Delatte *et al.* (2012), Palladini and Portes (2011))

7 Concluding remarks

We estimated the sovereign spread of five peripheral members of the euro-area using panel non-linear estimation methods. Two important objectives were to test empirically for the presence of nonlinear dynamics and to identify what may have driven the non-linear effects during this crisis. Our PSTR estimation confirms that the model of determination of sovereign bond spreads is not linear during the European crisis. Investors have priced the European sovereigns differently since Fall 2010. The contagion from Greece to the rest of the peripheral countries has probably operated through simultaneous dynamics in asset prices. On the other hand, our hypothesis about the role of liquidity shocks is rejected: they do not seem to have had self-reinforcing effects in peripheral European sovereign pricing.

Our methodology allows us to emphasize individual dynamics inside the panel. The sovereign bond of Italy, Spain and Portugal have not been driven by exactly the same dynamics as the bond of Grecce and Portigal. But we do we find that, in all countries, initial shocks on fundamentals are amplified when volatility and stress on financial entities increase. In addition our results reveal that when active investors use credit default swaps indices to leverage their views on credit risk in the financial sector, this amplifies changes in the spreads of European peripheral countries. The returns of peripheral government bonds have been driven up in excess of what the fundamentals would normally justify by large directional positions in CDS indices. CDS indices have been broadly used in the securitization process such as index-CDO. This suggests further exploration of this family of instruments.

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Table 1: Selection of the optimal linear model

	spec 1	spec 2	spec 3	spec 4	spec 5	spec 6	spec 7	spec 8
$Debt - to - GDP$	-0.127^{***}	-0.128^{***}	-0.165^{***}	-0.213^{***}	-0.128^{***}	-0.165^{***}	-0.089^{***}	-0.115^{***}
	(-7.83)	(-8.73)	(-11.19)	(-8.56)	(-8.11)	(-11.19)	(-5.62)	(-9.10)
$Debt - to - GDP^2$	0.001^{***}	0.001^{***}	0.001^{***}	0.001^{***}	0.001^{***}	0.001^{***}	0.001^{***}	0.001^{***}
	(13.35)	(13.62)	(15.96)	(12.07)	(14.1)	(15.96)	(11.16)	(14.53)
Fiscal balance	0.025	0.025	0.007	0.113^{***}	0.051^{***}	0.007	0.07^{***}	0.056^{***}
	(1.35)	(1.35)	(0.38)	(3.88)	(2.73)	(0.38)	(3.91)	(3.06)
Unemployment	0.459^{***}	0.457^{***}	0.381^{***}	0.586^{***}	0.362^{***}	0.381^{***}	0.438^{***}	0.377^{***}
	(14.1)	(14.85)	(12.26)	(10.66)	(10.61)	(12.26)	(12.88)	(11.62)
Unit Labor Cost	-0.105^{***}	-0.104^{***}	-0.101^{***}	-	-	-0.101^{***}	-	-
	(-6.42)	(-6.56)	(-5.95)			(-5.95)		
Trade balance	-0.004	-	-	0.325^{***}	0.055	-	0.05	-
	(-0.1)			(5.38)	(1.35)		(1.31)	
Vix	0.033^{***}	0.033^{***}	0.034^{***}	0.021^{**}	0.016^{***}	0.034^{***}	0.014^{**}	0.016^{**}
	(5.22)	(5.24)	(5.14)	(2.2)	(2.56)	(5.14)	(2.35)	(2.55)
Bid-Ask	3.643^{***}	3.639^{***}	3.683^{***}	-	3.803^{***}	3.683^{***}	3.77^{***}	3.86^{***}
	(26.11)	(26.83)	(25.44)		(24.77)	(25.44)	(26.00)	(26.11)
Outstanding stock	-59.76^{***}	-59.70^{***}	$-$	-65.25^{***}	-	-	-58.16^{***}	-
	(-7.53)	(-7.53)		(-4.78)			(-6.98)	
Unconventional Monetary Policy	0.0032^{**}	0.0032^{**}	0.0068^{***}	0.0040^{*}	0.0055^{***}	0.0068^{***}	0.0018	0.0051^{***}
	(2.35)	(2.38)	(5.14)	(1.73)	(3.98)	(5.14)	(1.31)	(3.80)
AIC	0.197	0.197	0.329	1.284	0.409	0.329	0.294	0.414
Schwarz	0.207	0.207	0.339	1.294	0.419	0.339	0.304	0.424

Note: (*): significant at the 10% level; (**): significant at the 5% level and (***): significant at the 1% level.

Table 2: Linearity Tests wth a PSTR model

	H1: Fire-sale liquidation			H2: Feedback loop	H3: CDS indices	Control
	Flight to liquidity	Flight to quality	Asymetry information			
AAA/ 10-year Treasury spread	111.4***					
10-year Swap spread	78.2***	78.2***				
A/ 10-year Treasury spread	84.8***	84.8***				
High-Yield bond/ Baa spread	70.8***	70.8***	70.8***			
StockbondsCorr		91.4***				
Cross-Section dispersion banks			51.2***			
IVOL bank				116.2***		
CmaxFin				194.2***		
Euribor-ois				119.4***		
I-traxx Europe					108.4***	
X-over					91.9***	
Hivol					79.0***	
CDS Snr-Fin					148.3***	
CDS Sub-Fin					130.9***	
Vstoxx						64.2***
RVOL Germ						29,5*
RVOL Nonfin						41,4***
RVOL Pound						47,9***
RVOL Doll						36,6***
RVOL Yen						39.6*
FTSE 300						111.8***
S&P 350						111.6***
Domestic indices returns						37.7***

Notes: The corresponding LM statistic has an asymptotic $\chi^2(p)$ distribution under H_0. (*): significant at the 10% level; (**): significant at the 5% level and (***): significant at the 1% level.

Table 3: Estimation of the sovereign bond model with a PSTR model (Full Sample)

	Model 1	Model 2	Model 3
	Cmax Fin	CDS Snr-Fin	CDS Sub-Fi
Linearity Stat	194.2***	148.3***	130.9***
Smooth Parameter	111.4	0.266	0.090
Loc Parameter	0.819	239.7	391.6
RSS	175.7	218.5	247.9
Schwarz Crit.	-0.485	-0.267	-0.140

Notes: (*): significant at the 10% level; (**): significant at the 5% level and (***): significant at the 1% level.

Table 4: Estimation of the sovereign bond model with a PSTR model (two sub-samples)

	Model 1	Model 2	Model 3
	Cmax Fin	CDS Snr-Fin	CDS Sub-Fi
Sub-panel Italy, Spain and Portugal			
Linearity Stat	54.2***	88.2***	82.8***
Smooth Parameter	40.22	56.9	1.9
Loc Parameter	0.544	135.8	228.0
RSS	26.2	27.1	26.3
Schwarz Crit.	-1.68	-1.65	-1.68
Sub-panel Grece and Ireland			
Linearity Stat	42.0***	67.3***	61.9***
Smooth Parameter	140.8	0.62	7.55
Loc Parameter	0.863	155.2	261.7
RSS	75.1	115.7	110.5
Schwarz Crit.	-0.001	0.430	0.384

Notes: (*): significant at the 10% level; (**): significant at the 5% level and (***): significant at the 1% level.

Table 5: Estimates of the sovereign bond model with a PSTR model for Italy, Spain & Portugal

	Model 1		Model 2		Model 3	
	CDS Snr Fin		CDS Sub Fin		CMax Fi	
	β_1	β_2	β_1	β_2	β_1	β_2
Debt	0.0212^{***} (3.54)	0.035^{***} (4.28)	0.011^{*} (1.94)	0.042^{***} (5.23)	0.129^{***} (5.91)	-0.031^{***} (-3.48)
Squared Debt	-	-	-	-	-0.0004^{***} (-3.45)	0.0002^{***} (3.48)
Fiscal Balance	0.004 (0.30)	0.154^{**} (2.40)	0.053^{***} (3.43)	0.137^{**} (2.12)	-0.053^{*} (-1.80)	0.170^{***} (3.90)
Unemployment	0.015 (0.54)	0.112^{***} (2.70)	0.100^{***} (3.50)	0.116^{***} (2.65)	-0.155^{***} (-3.04)	0.179^{***} (4.62)
Unit Labor Cost	0.004 (0.48)	-0.024^{**} (-2.54)	-0.006 (-0.66)	-0.028^{***} (-2.89)	-	-
VIX	0.015^{***} (5.67)	0.035^{***} (4.71)	0.014^{***} (4.9)	0.038^{***} (5.27)	0.021^{***} (5.99)	0.002 (0.25)
Bid-Ask	15.18^{***} (12.9)	-10.44^{***} (-8.14)	43.05^{***} (6.16)	-38.33^{***} (-5.49)	7.71^{***} (10.66)	-2.515^{***} (-3.86)
Outstanding Stock of gov	-4.61 (-0.78)	-9.32^{***} (-5.70)	-11.672^{**} (-1.99)	-9.607^{*} $(-6.57$	-	-
Unconv. Monet. Policy	0.0074^{***} (4.20)	0.0013 (0.57)	0.0072^{***} (5.79)	0.0003 (-0.15)	0.0071^{***} (7.34)	0.0019 (1.38)
Smooth Parameter γ	60.3		2.18		39.62	
Loc Parameter c	135.7		227.9		0.545	
Linearity Stat.	96.9^{***}		90.0^{***}		62.0^{***}	
RSS	27.6		26.81		26.4	
Schwarz Crit.	-1.685		-1.716		-1.786	

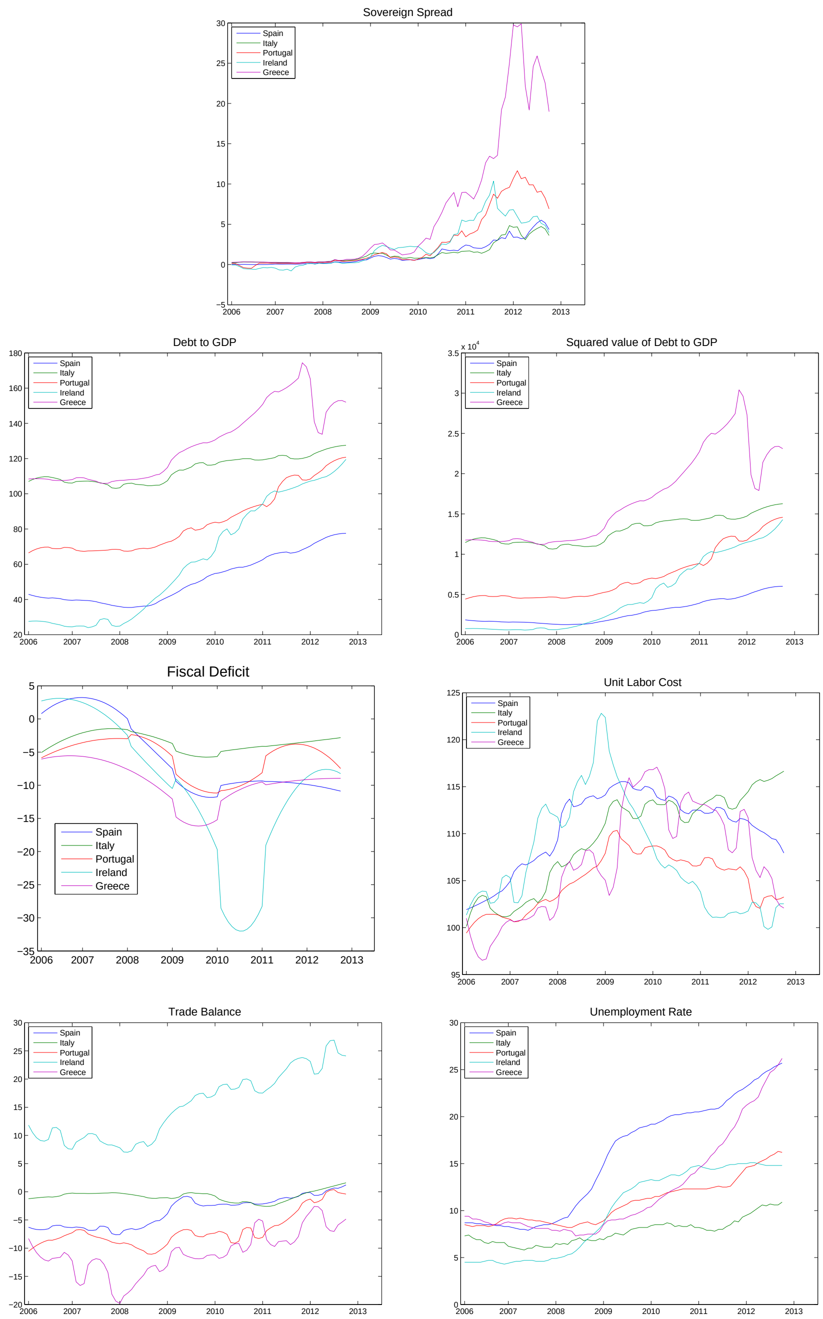
Notes: The T-stat in parentheses are corrected for heteroskedasticity. (*): significant at the 10% level; (**): significant at the 5% level and (***): significant at the 1% level.β_1 and β_2 correspond to the coefficient in Eq (11). β_1 is the coefficient in the first extreme regime . The coefficient in the second extreme regime is $\beta_1 + \beta_2$.

Table 6: Estimates of the sovereign bond model with a PSTR model for Greece & Ireland

	Model 1		Model 2		Model 3	
	CDS Snr Fin		CDS Sub Fin		Cmax Fi	
	β_1	β_2	β_1	β_2	β_1	β_2
Debt	-0.101^{***} (-4.20)	0.086^{***} (2.13)	-0.114^{***} (-4.99)	0.067 (1.53)	-0.296^{***} (-9.28)	0.242^{***} (4.05)
Squared Debt	0.0005^{***} (4.89)	-0.0004^{*} (-1.82)	0.001^{***} (6.5)	0.000 (-1.49)	0.001^{***} (10.89)	0.000 (-0.48)
Fiscal Balance	0.057^{***} (2.631)	0.031 (0.653)	0.056^{***} (2.91)	-0.047 (-0.81)	-0.092^{***} (-2.61)	0.109^{**} (2.28)
Unemployment	0.57^{***} (8.98)	0.69^{***} (4.30)	0.602^{***} (9.48)	0.639^{***} (3.99)	0.849^{***} (7.85)	0.02 (0.1)
Unit Labor Cost	0.03^{***} (1.73)	-0.008^{***} (-4.13)	0.022 (1.45)	-0.082^{***} (-4.02)	-0.036^{**} (-2.05)	-0.048^{**} (-2.10)
VIX	0.03^{***} (3.84)	-0.0135 (-0.44)	0.033^{***} (4.53)	-0.007 (-0.23)	0.025^{***} (3.99)	0.019 (0.83)
Bid-Ask	4.55^{***} (3.5)	-1.7 (-1.27)	4.735^{***} (4.19)	-1.904 (-1.61)	2.67^{***} (6.97)	1.175^{**} (2.05)
Outstanding Stock of gov	-	-	-	-	$161.9 * **$ (3.31)	-632.3^{***} (-5.96)
Uncon. Monet. Policy	0.026^{***} (6.26)	-0.038^{***} (-5.48)	0.024^{***} (6.49)	-0.033^{***} (-4.49)	0.0275^{***} (7.2)	-0.0459^{***} (-7.20)
Smooth Parameter γ	0.43		6.57		140.8	
Loc Parameter c	154.8		262.0		0.863	
Linearity Stat.	63.4***		60.2***		42.0***	
RSS	116.3		111.1		75.1	
Schwarz Crit.	0.358		0.312		-0.001	

Notes: The T-stat in parentheses are corrected for heteroskedasticity. (*): significant at the 10% level; (**): significant at the 5% level and (***): significant at the 1% level.β_1 and β_2 correspond to the coefficient in Eq (11). β_1 is the coefficient in the first extreme regime . The coefficient in the second extreme regime is $\beta_1 + \beta_2$.

Figure 1: Dependent and Explanatory Variables



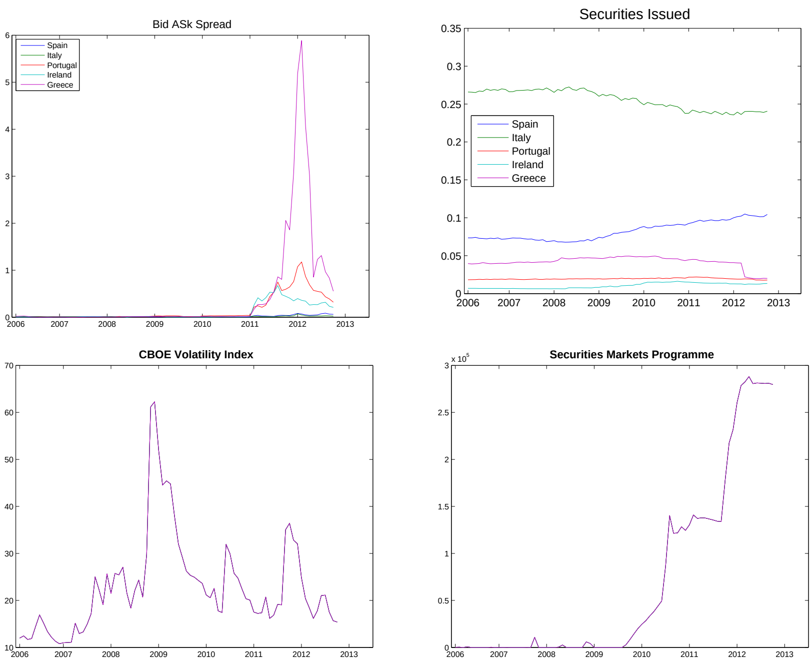
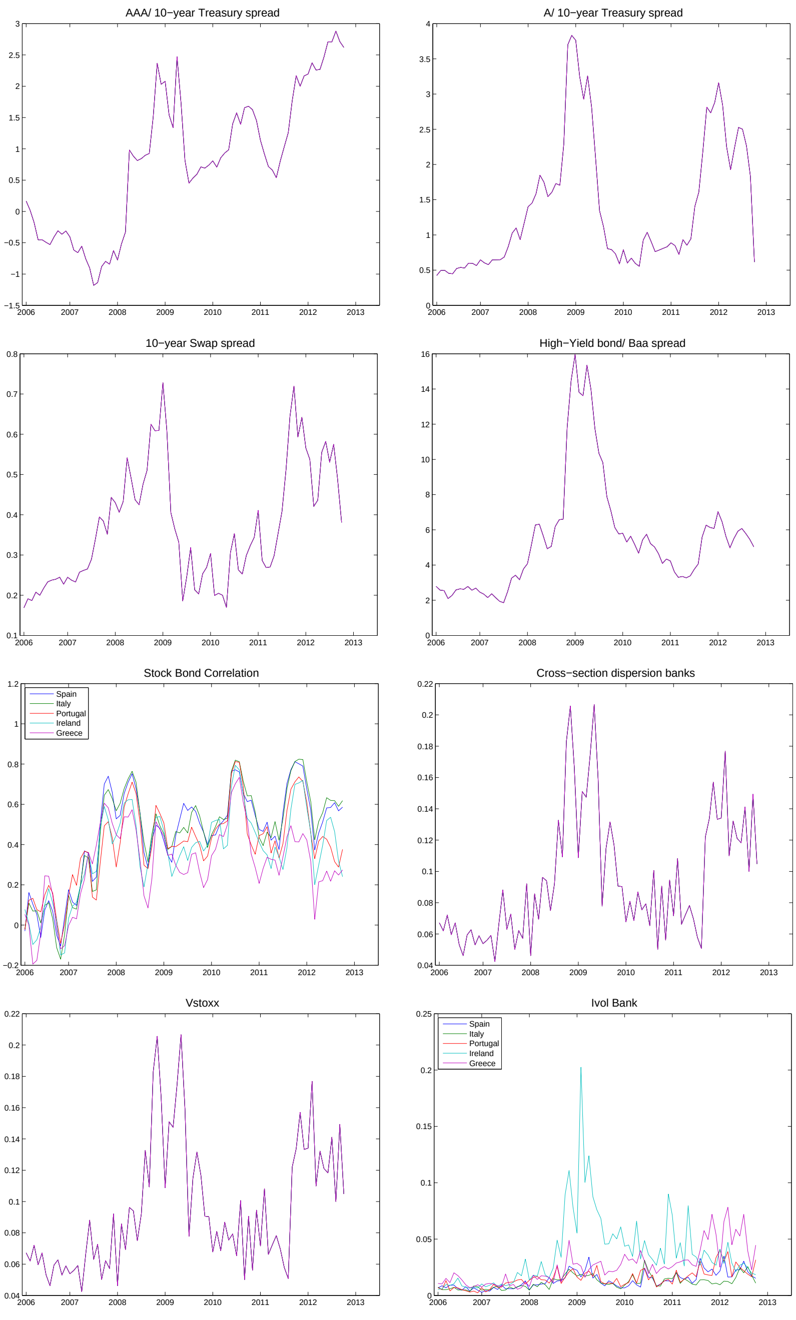
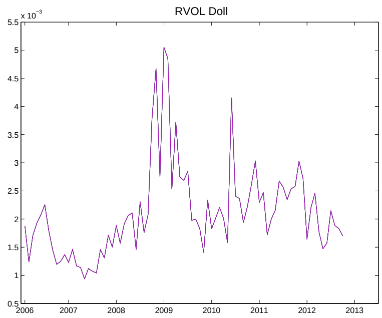

Bid ASk Spread
Spain
Italy
Portugal
Ireland
Greece
Securities Issued
Spain
Italy
Portugal
Ireland
Greece
CBOE Volatility Index
Securities Markets Programme
x 10^5

Figure 2: Threshold Variables

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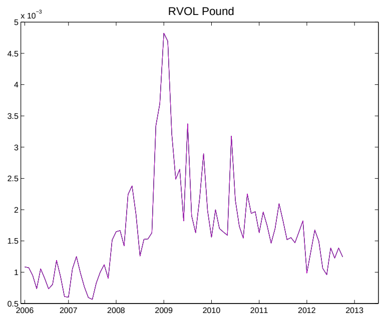
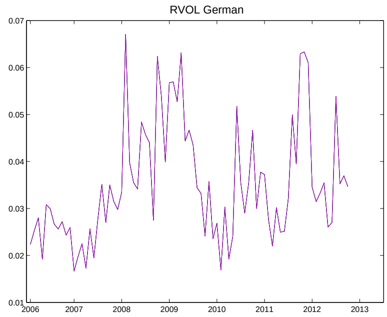
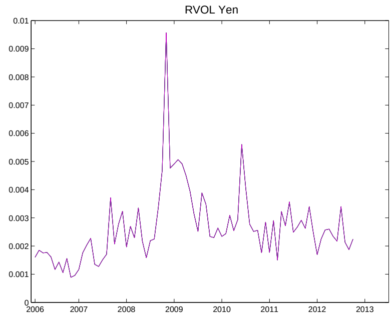
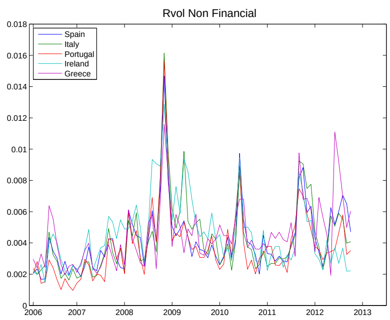
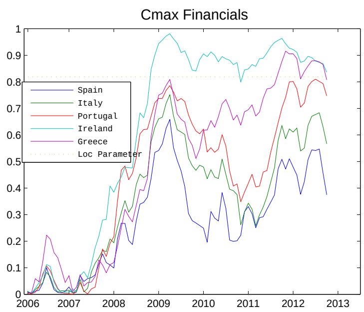
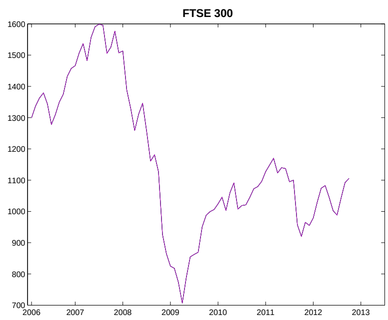

RVOL Pound

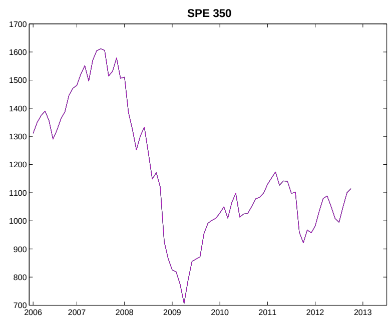

RVOL German

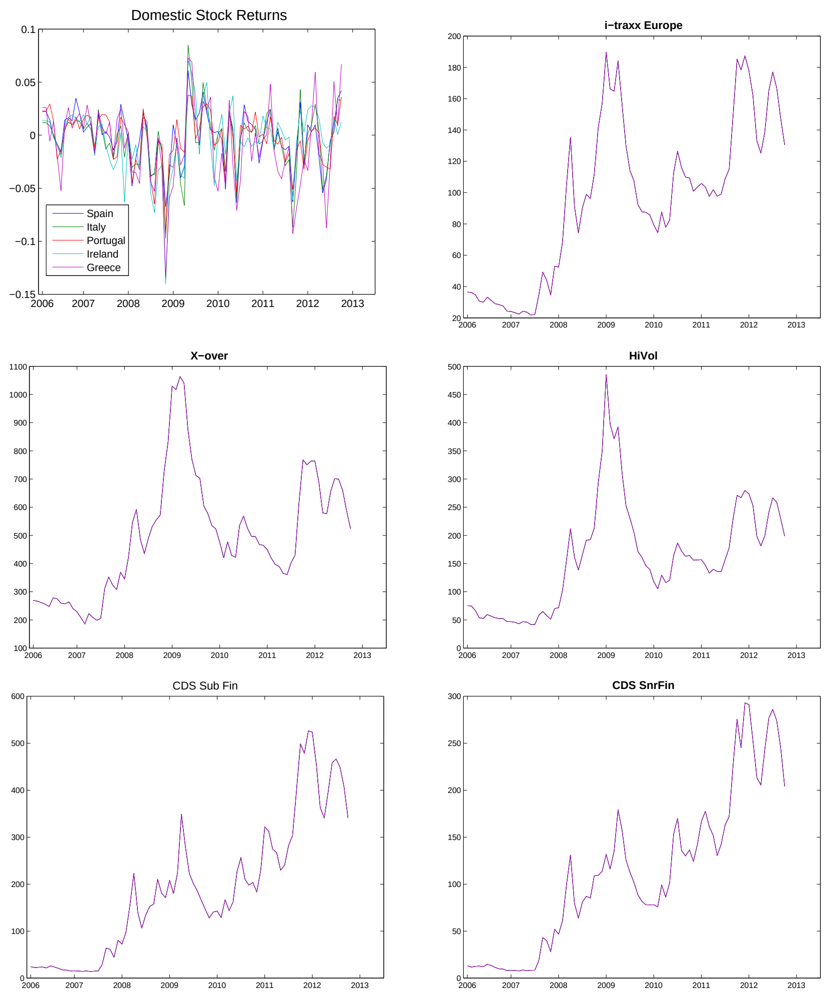

RVOL Yen


Rvol Non Financial
Spain
Italy
Portugal
Ireland
Greece


Cmax Financials
Spain
Italy
Portugal
Ireland
Greece
Loc Parameter


FTSE 300


SPE 350


Domestic Stock Returns
Spain
Italy
Portugal
Ireland
Greece
i−traxx Europe
X−over
HiVol
CDS Sub Fin
CDS SnrFin
2006
2007
2008
2009
2010
2011
2012
2013