# Sovereign Credit Risk, Liquidity, and ECB Intervention: *Deus ex Machina*?[☆]

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**Abstract**

We examine the dynamic relation between credit risk and liquidity in the Italian sovereign bond market during the Euro-zone crisis and the subsequent European Central Bank (ECB) interventions. Credit risk drives the liquidity of the market: a 10% change in the credit default swap (CDS) spread leads to a 13% change in the bid-ask spread, the relation being stronger when the CDS spread exceeds 500 bp. The Long-Term Refinancing Operations (LTRO) of the ECB weakened the sensitivity of market makers' liquidity provision to credit risk, highlighting the importance of funding liquidity measures as determinants of market liquidity.

*Keywords:* Liquidity, Credit Risk, Euro-zone Sovereign Bonds, Financial Crisis, MTS Bond Market
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## 1. Introduction

The challenges facing the governments of the GIIPS countries (Greece, Ireland, Italy, Portugal and Spain) in refinancing their debt marked the genesis of the Euro-zone sovereign debt crisis. Following a series of credit rating downgrades of three countries on the Euro-zone periphery, Greece, Ireland and Portugal, in the spring of 2010, the crisis spread throughout the Euro-zone. The instability in the Euro-zone sovereign bond market reached its apogee during the summer of 2011, when the credit ratings of two of the larger countries in the Euro-zone periphery, Italy and Spain, were also downgraded. This culminated in serious hurdles being faced by several Euro-zone countries, causing their bond yields to spike to unsustainable levels. The crisis has abated to some extent, due in part to fiscal measures undertaken by the European Union (EU) and the International Monetary Fund (IMF), but mostly thanks to the intervention by the European Central Bank (ECB) through a series of policy actions, including the Long-Term Refinancing Operations (LTRO) program, starting in December 2011.

The discussion in the academic and policy-making literatures on the Euro-zone crisis has mainly focused on market aggregates such as bond yields, relative spreads, and credit default swap (CDS) spreads and the reaction of the market to intervention by the Troika of the ECB, the EU and the IMF. Although the analysis of yields and spreads is useful, it is equally relevant for policy makers and market participants to understand the dynamics of market liquidity in the European sovereign debt markets, i.e., the drivers of market liquidity, particularly given the impact market liquidity has on bond yields, as documented in the previous literature on asset prices.

In this paper, we address the latter issue and analyze the inter-relation between market liquidity and credit risk, the effect of the funding liquidity of the market makers, and how this inter-relation changed thanks to the ECB interventions. We drive our analysis by developing a simple model that formalizes several channels through which credit risk affects market liquidity. Our empirical analysis shows that credit risk affects market liquidity, and that this relation shifts conditional on the level of the CDS spread: it is stronger when the CDS spread exceeds 500 bp, a threshold used as an indicator by clearing houses in setting margins. Moreover, we show that the LTRO intervention by the ECB, which funneled funding liquidity into the banking system, weakened the sensitivity of market liquidity to credit risk.

The linkage between credit risk and market liquidity is an important topic because a liquid market is of paramount importance for both the success of the implementation of central bank interventions, whether in the form of interest rate setting, liquidity provision funding, or quantitative easing, and their unwinding. Moreover, as we show in this paper, monetary policy has an impact on the interplay between credit risk and market liquidity itself.

The main focus of our research in this paper is to determine the dynamic relation between market liquidity and credit risk, as well as other risk factors such as global systemic risk, market volatility, and the funding liquidity risk of market makers. We study the effects of the ECB measures in the context of this dynamic relation. We employ the time-series of a range of liquidity metrics, as well as CDS spreads, a measure of credit quality, to analyze the liquidity of Italian sovereign bonds during the period from July 1, 2010 to December 31, 2012. We allow the data to help us uncover how the relation between credit risk and liquidity depends on the endogenous level of the CDS spread. In addition, we examine how these relationships were influenced by the interventions of the ECB.

We motivate our empirical analysis with a simple model of a risk averse market maker, holding an inventory of a risky asset and setting her optimal marginal quotes (and, therefore, the optimal bid-ask spread), in the presence of margin constraints and borrowing costs. The margins, set by a clearing house, depend on the risk of the asset, as measured by the CDS spread, and the actions of the central bank. The CDS market is fundamental to the market maker's and the clearing house's decisions, since it is from the CDS market that they deduce the future volatility of the asset return. In addition, the market maker can pledge her assets at the central bank to finance her positions at rates influenced by the central bank's actions. The model provides several empirical predictions that we test in the empirical section of the paper.

First, we test the empirical prediction that the relation between the credit risk of a sovereign bond and its liquidity is statistically significant and, specifically, that the credit risk, as measured by the CDS spread, leads the liquidity, and not the other way around. We find that a 10% change in credit risk is followed by a 13% change in market liquidity. Further, we find that the coefficients of both contemporaneous and lagged changes in the CDS spread are statistically and economically significant in explaining the market liquidity of sovereign bonds, even after controlling for the lagged liquidity variable and the contemporaneous changes in other factors. In particular, we test whether global risk and funding liquidity factors also affect market liquidity.

Second, we examine whether the relation between credit risk and market liquidity is conditional on the level of the CDS spread, i.e., whether it is significantly altered when the CDS spread crosses a certain threshold. We let the data identify the presence of such a CDS threshold effect, and find that the relation between market liquidity and credit risk is different, depending on whether the Italian CDS spread is below or above 500 bp. We find not only that a change in the CDS spread has a larger impact on market liquidity when the CDS spread is above 500 bp, but that this relation is instantaneous, while the lead-lag relation is stronger for lower levels of the CDS spread. We interpret this finding, together with a change in the margins for bonds, in light of the predictions made by Brunnermeier and Pedersen (2009).

Third, we analyze the impact of ECB intervention on the relation between credit risk and liquidity. The threshold effect in CDS levels is present only until December 21, 2011. In fact, our test for an endogenous structural break indicates that, on December 21, 2011 (when the ECB allotted the funds of the LTRO program), the relation between the two variables changes significantly. Thereafter, during 2012, after the large amount of funding liquidity from the LTRO program has become available to market makers and market participants, changes in market liquidity still respond to changes in credit risk, but with a lagged effect, and with a significantly lower intensity, while the only contemporaneous variable that affects market liquidity significantly is the global funding liquidity variable proxied by the Euro-US Dollar cross-currency basis swap spread (CCBSS).[1]

The Euro-zone sovereign crisis provides us with an unusual laboratory in which to study how the interaction between credit risk and illiquidity played out, in a more comprehensive framework than has been used in previous studies of corporate or other sovereign bond markets. In contrast to research on corporate bonds, which are generally traded over-the-counter (OTC), we have the advantage of investigating an exchange-traded market, using a unique, tick-by-tick data set obtained from the Mercato dei Titoli di Stato (MTS), the world's largest electronic trading platform for sovereign bonds. With respect to the US Treasury and other sovereign bond markets, the presence of a common currency for sovereign issuers means that the ECB is completely independent of the Italian government. Hence, the central bank's monetary policy has a qualitatively different impact on its sovereign credit risk, as well as on the market liquidity of its sovereign bonds, compared to countries whose central banks are somewhat within the control of the sovereign.

To our knowledge, ours is the first paper to empirically investigate the dynamic relation between market liquidity and credit risk in the sovereign bond market, particularly during a period of crisis. The existing literature has highlighted the theoretical relation between bond yields and market liquidity, as well as that between funding liquidity and market liquidity (as modeled by Brunnermeier and Pedersen, 2009). We contribute to this literature

[1] This spread represents the additional premium paid per period for a cross-currency swap between Euribor and US Dollar Libor. Market participants view it as a measure of the macro-liquidity imbalances in currency flows between the Euro and the US Dollar, the global reserve currency.

by exploring the role of central bank interventions, and show both theoretically and empirically that they affect the relation between sovereign credit risk and market liquidity. The laboratory for our analysis is the Italian sovereign bond market, particularly around the Euro-zone crisis, starting from July 2010. Italy has the largest sovereign bond market in the Euro-zone (and the third largest in the world after the US and Japan) in terms of amount outstanding, and is also a market that experienced substantial stress during the recent crisis. It is important to emphasize that such an analysis cannot be performed in other large sovereign bond markets, such as those of Germany or France, since they were not as much affected by the sovereign credit risk concerns.

In Section 2 of the paper, we survey the literature on sovereign bonds, particularly the papers relating to liquidity issues. In Section 3, we present a model of market maker behavior in the setting of the bid-ask spread and derive its empirical implications. In Section 4, we provide a description of the MTS market architecture and the features of our database. In Section 5, we present our descriptive statistics. Our analysis and results are presented in Section 6, and Section 7 presents several robustness checks. Section 8 concludes.

## 2. Literature Survey

The dynamic relation between credit risk and the market liquidity of sovereign bond markets has received limited attention in the literature, thus far. The extant literature on bond market liquidity seldom focuses on sovereign bond markets, with the exception of the US Treasury bond market; yet, even in this case, most papers cover periods before the current financial crisis and address limited issues related to the pricing of liquidity in the bond yields.[2] It is, therefore, fair to say that the relation between sovereign credit risk and market liquidity has not yet been investigated in the US Treasury market, possibly because US sovereign risk was not an issue until the recent credit downgrade by Standard & Poor's. The liquidity in the US Treasury bond market has been investigated by Chakravarty and Sarkar (1999), using data from the National Association of Insurance Commissioners, and Fleming (2003), using GovPX data. Fleming and Remolona (1999), Pasquariello and Vega (2007), and Goyenko, Subrahmanyam, and Ukhov (2011) study the responses of the US Treasury markets to unanticipated macro-economic news announcements. In a related paper, Pasquariello, Roush, and Vega (2011) study the impact of outright (i.e.,

[2]Specifically, the existing literature documents the *direct* impact of liquidity (e.g., Dick-Nielsen, Feldhütter, and Lando, 2012, among others) on bond yields and prices, but not the impact of credit risk on liquidity, or how credit risk affects the bond yields through bond liquidity. In this spirit, we need to establish the relation between credit risk and liquidity in order to then, in turn, quantify its effect on bond yields. An effort in this direction is made by Jankowitsch, Nagler, and Subrahmanyam (2014).

permanent) open-market operations carried out by the Federal Reserve Bank of New York on the microstructure of the secondary US Treasury market. Furthermore, there are a few papers in the literature analyzing data from the electronic trading platform in the US known as BrokerTec, such as Fleming and Mizrach (2009) and Engle, Fleming, Ghysels, and Nguyen (2011).

There are a handful of papers on the European sovereign bond markets, and again, these papers generally examine a limited time period, mostly prior to the global financial crisis, and largely focus on the impact of market liquidity on bond yields; see for example Coluzzi, Ginebri, and Turco (2008), Dufour and Nguyen (2012), Beber, Brandt, and Kavajecz (2009), Favero, Pagano, and von Thadden (2010) and Bai, Julliard, and Yuan (2012). More recent work has highlighted the effects of ECB interventions on bond yields, market liquidity, and arbitrage relationships between fixed income securities. Ghysels, Idier, Manganelli, and Vergote (2014) study the effect of the Security Markets Programme (SMP) intervention on bond returns, while Corradin and Rodriguez-Moreno (2014) document the existence of unexploited arbitrage opportunities between European sovereign bonds denominated in Euros and Dollars, as a consequence of the SMP. Eser and Schwaab (2013) and Mesters, Schwaab, and Koopman (2014) show long- and short-term effects of the ECB interventions on European bond yields. Finally, Corradin and Maddaloni (2015) and Boissel, Derrien, Örs, and Thesmar (2014) investigate the relation between sovereign risk and repo market rates during the European sovereign crisis.

There is a vast literature on liquidity effects in the US corporate bond market, examining data from the Trade Reporting and Compliance Engine (TRACE) database maintained by the Financial Industry Regulatory Authority and using liquidity measures for different time periods, including the global financial crisis. This literature is relevant to our research both because it analyzes a variety of liquidity measures and because it deals with a relatively illiquid market with a vast array of securities. For example, Friewald, Jankowitsch, and Subrahmanyam (2012a) show that liquidity effects are more pronounced in periods of financial crisis, especially for bonds with high credit risk. Similar results have been obtained by Dick-Nielsen, Feldhütter, and Lando (2012), who investigate the effect of credit risk (credit ratings) on the market liquidity of corporate bonds.[3]

In a theoretical contribution to the literature on the relation between corporate credit risk and liquidity, Ericsson

[3]Other recent papers quantifying liquidity in this market provide related evidence. See, for example, Edwards, Harris, and Piwowar (2007), Mahanti, Nashikkar, Subrahmanyam, Chacko, and Mallik (2008), Zhou and Ronen (2009), Jankowitsch, Nashikkar, and Subrahmanyam (2011), Bao, Pan, and Wang (2011), Nashikkar, Subrahmanyam, and Mahanti (2011), Lin, Wang, and Wu (2011), Feldhütter (2012), and Jankowitsch, Nagler, and Subrahmanyam (2014).

and Renault (2006) show both theoretically and empirically that bond illiquidity is positively correlated with the likelihood of default. He and Milbradt (2014) provide a theoretical framework for the analysis of corporate bonds traded in OTC markets and show that a thinner market liquidity, following a cash flow decline, feeds back into the shareholders' decision to default, making a company more likely to default. A final theoretical paper related to our analysis is by Brunnermeier and Pedersen (2009), who investigate the relation between funding liquidity and market liquidity.

To the best of our knowledge, there are no theoretical models that investigate the relation between *sovereign* credit risk and market liquidity. The models in Ericsson and Renault (2006) and He and Milbradt (2014) cannot be applied straightforwardly to the sovereign framework because of the nature of the credit event. There are, in fact, no bankruptcy or strategic default choices in the sovereign context (see Augustin, Subrahmanyam, Tang, and Wang, 2014, Section 7.1), although the outcome of debt renegotiation, e.g., the recovery rate, could arguably be affected by the liquidity of the secondary market. From a theoretical perspective, one channel that definitely applies to the relation between sovereign credit risk and market liquidity is that of the market maker's inventory concerns, as in the model proposed by Stoll (1978). In this paper we extend Stoll's (1978) model by including further determinants of market liquidity, i.e., margins and a policy effect, whereby both margins and borrowing rates are influenced by the policy maker's actions (i.e., by the central bank). Our model is designed to specifically capture the effects that credit risk has on the market liquidity of bonds. A comprehensive theoretical model where sovereign credit risk, via debt renegotiations, affects market liquidity could be formulated; yet, such model lies beyond the scope of this paper. Nonetheless, in our empirical investigation, we allow and test for both the effects of credit risk on liquidity and liquidity on credit risk.

There are several important differences between the prior literature and the evidence we present in this paper. First, we are among the first to focus on the relation between liquidity (rather than yield spreads) in the cash bond market and credit risk, especially in the context of sovereign credit risk. Second, while most of the previous literature spans past, and thus more normal, time periods in the US and Euro-zone markets, the sample period we consider includes the most relevant period of the Euro-zone sovereign crisis. Third, our focus is on the *interaction* between credit risk and liquidity, i.e., how credit risk affects illiquidity and vice versa. Fourth, we examine the impact of monetary policy interventions on the linkage between credit risk and liquidity, in the context of ECB policies over the past few years, to measure and document their differential effects. Finally, we contribute to the literature a model that links the bid-ask spread in the bond market to the CDS market.

## 3. The Model and its Testable Implications

In this section, we review and extend the standard model by Stoll (1978), in order to guide and motivate our empirical analysis. The extension allows us to define some simple concepts and gain an intuition about the forces driving the choice, by a market maker of a sovereign bond, of what bid-ask spread to quote on the market. The market maker stands ready to buy from, or sell to, an external trader, extracts information regarding the risk of the sovereign bond from the CDS market, and faces margin constraints arising from her inventory. The players in our model are i) the market maker, ii) other (external) traders buying or selling the bonds, iii) the clearing house, and iv) the central bank. The main purpose of our model is to characterize how a change in the CDS spread is reflected in the bid-ask spread of a bond issued by the underlying entity.[4] Figure 1 summarizes the players and the mechanisms of our model.

INSERT FIGURE 1 HERE

Central to the development of our model is identifying how the actions of each of the actors are affected by the credit risk of the bond that we are considering, and how, in turn, these actions affect the liquidity provided by the market maker. The model in Stoll (1978) shows that an increase in the risk of the security is directly reflected in the market liquidity provision choice of the market maker (*Inventory Risk* in Figure 1). In addition to this direct channel, our model includes an *indirect* channel, through which the credit risk of the bond affects the liquidity provision choice of the market maker. The indirect channel relates to the dealer's cost of financing a bond in the repo market, including the margin requirements, when she has a non-positive inventory and she needs to sell a bond to a trader (*Margins* in Figure 1). In the indirect channel, credit risk affects the liquidity provision by the market maker through the clearing house's margin setting decision, which depends on the credit risk of the bond (*Margin Setting* in Figure 1). This hypothesis is motivated by the "Sovereign Risk Framework" adopted by LCH.Clearnet, the major European clearing house, and by other clearing houses, including Cassa di Compensazione e Garanzia, during the sovereign crisis: the framework states that the clearing house adjusts the margins based on a list of

[4]We thank the referee for suggesting we formalize our empirical predictions in a simple model.

indicators, which includes the CDS spread and the bond yield spread over the German bund, to account for losses incurred in case of default by the issuer of the security (LCH.Clearnet, 2011).

The margin setting decision by the clearing house is also affected by the policies of the central bank, i.e., by i) the central bank's key interest rates, ii) the central bank's interventions, and iii) its explicit requests to the clearing house (*Funding Rate* and *Margin Framework* in Figure 1). First, the (collateralized) borrowing rate, set by the central bank, affects the volume traded on the repo market, by affecting its supply and demand, and, thus, the risk bearing capacity of the clearing house (see Mancini, Ranaldo, and Wrampelmeyer, 2014, for a detailed account of the effects of the ECB's interventions on the European Repo market).

Second, during the European debt crisis, the ECB enacted several extraordinary interventions: i) the Security Market Program (SMP), initiated in May 2010, ii) LTRO, announced and implemented in December 2011, iii) policy guidance, and iv) the outright monetary transactions (OMT), also announced in December 2011.[5] These interventions could affect the credit risk of the Eurozone, the liquidity of its bond market, or the funding liquidity of its banks: any of these effects should be taken into consideration by the clearing house, when setting margins. A similar implication can be drawn from the model by Brunnermeier and Pedersen (2009): the provision of funding liquidity relaxes the market makers' borrowing constraints and, consequently, the impact of margins on market liquidity.

Third, our hypothesis that central banks can affect even more directly the relation between margin settings and credit risk is supported by documents from the International Monetary Fund (2013) and the Bank of Italy (2012). Following a substantial margin increase by the clearing house LCH.Clearnet at a time of high credit risk, the Italian and French central banks worked with the clearing house to propose a shared methodology to ensure that margin requirements would depend smoothly on the CDS spread. This prevents the clearing house from implementing abrupt margin increases, disrupting the liquidity of the sovereign bond market when the sovereign credit risk is already high (Bank of Italy, 2012). The central banks requested the clearing house to avoid the possibility for margins to become procyclical to sovereign risk. Finally, in our model, the central bank affects the dealer's option

[5]The SMP is a Eurosystem programme to purchase bonds—especially sovereign bonds—on the secondary markets. The last purchase under the SMP was made in February 2012. At its peak, in August 2011, the programme's volume totalled around 210 billion. The LTRO interventions provided three-year funding of €489 billion on December 21, 2011 and €523 billion on February 29, 2012. The long-term maturity of this massive funding action was unprecedented in ECB policy history, and even globally. By policy guidance we largely refer to the Mario Draghi speech on July 26, 2012, at the Global Investment Conference in London, where he stated: "The ECB is ready to do whatever it takes to preserve the euro. And believe me, it will be enough." Outright monetary transactions is the programme to purchase sovereign bonds that substitute the SMP programme.

to seek financing, by pledging the securities she holds, through changing the rate at which she can obtain funds (*Borrowing Costs* in Figure 1). One could also argue that the central bank's policy interventions themselves depend on the level of credit risk of the system (the dotted line in Figure 1). While we do not pursue this line of modelling, our predictions would be robust to the inclusion of this additional channel. Finally, our model aims at specifically capturing the effect of credit risk on bond market liquidity. While a model emphasizing the effect of a shock to market liquidity on credit risk in the sovereign context, possibly via debt renegotiation, could be developed, such a model lies beyond the scope of this paper.

We only model explicitly the behavior of the market maker, and assume as exogenous the other players' actions. In our model, we assume that the dealer, or market maker, is continuously making the market for a security; in this continuum in time, we choose an arbitrary point at which we model her optimal quote-setting decision. The dealer has an initial wealth of $W_0$ and an inventory made up of the bond with a dollar value equal to $I$. Moreover, she also invests a fraction $k$ of $W_0$ in the market portfolio. She invests the remainder of her wealth, $(1-k)W_0 - I$ at the risk-free rate $r_f$, if $I < (1-k)W_0$, i.e., in case there is a surplus. However, if $I > (1-k)W_0 > 0$, she borrows the residual amount, by pledging securities in her portfolio at the central bank, at a rate $r_b = r_f + b$. Additionally, if the inventory position $I$ is negative, she borrows the bond on the repo market, where it is subject to a margin requirement $m$. We model the margin, $m$, as an upfront cost of borrowing the specific bond rather than, for example, any bond under a general collateral agreement. In general, having to post margin constitutes a (opportunity) cost for the market maker, who would have otherwise allocated the required capital differently.

In light of our assumptions, we indicate the margin set by the clearing house as $m(b, CDS)$, i.e., a generic function of the CDS and the central bank liquidity policy, parametrized by the (collateralized) borrowing rate at the central bank. Following from the previous arguments, the margin setting decision depends on the credit risk and the policy arguments as follow: $\frac{\partial m(b,CDS)}{\partial CDS} > 0$, and $\frac{\partial m(b,CDS)}{\partial b} > 0$. We interpret the request of the central bank to avoid procyclical margin setting policies as a *shift* in the sensitivity of the margins to the level of the CDS spread, for a given level of borrowing rate, i.e., a shift in $\left.\frac{\partial m(b,CDS)}{\partial CDS}\right|_b$.

If the dealer does not trade on the chosen date, the terminal wealth from her initial portfolio will be

$$W_I = W_0 k (1 + r_M) + I (1 + r) + \begin{cases} ((1-k)W_0 - I)\left(1 + r_f\right) & \text{if } (1-k)W_0 > I > 0 \\ ((1-k)W_0 - (1-m)\, I)\left(1 + r_f\right) & \text{if } (1-k)W_0 > 0 > I \\ ((1-k)W_0 - I)(1 + r_b) & \text{if } I > (1-k)W_0 > 0 \end{cases},$$

where the market portfolio (expected) return is $r_M$ ($\overline{r_M}$) and variance $\sigma_m^2$, and the bond (expected) net return is $r$ ($\overline{r}$).[6] The (forward looking) variance of the bond return, which the market maker extracts from the CDS market, is $\sigma^2(CDS)$.

After trading a dollar quantity $Q$, the dealer's post-trading wealth is

$$\begin{aligned} W_{I+Q} =& W_0 k (1 + r_M) + (I + Q)(1 + r) + C_Q (1 + r_f) + \\ & \begin{cases} ((1-k)W_0 - (I + Q))\left(1 + r_f\right) \\ \qquad \text{if } (1-k)W_0 > I + Q > 0 \\ ((1-k)W_0 - (1-h)(I + Q))\left(1 + r_f\right) \\ \qquad \text{if } (1-k)W_0 > I + Q > 0 > I \\ ((1-k)W_0 - (I + Q))(1 + r_b) \\ \qquad \text{if } I + Q > (1-k)W_0 > 0 \end{cases} \end{aligned}$$

where $C_Q$ is the dollar cost of entering into this transaction and depends on $Q$. These costs can be positive or negative, depending on whether the marginal trade in the bond raises or lowers the dealer's inventory-holding costs, and essentially captures the dealer's exposure cost of holding a non-optimal portfolio. The dealer has a constant absolute risk aversion utility function, $U(x) = -e^{-\gamma x}$, and she will trade and price the trade so that her expected

[6] Since we aim to gain an understanding of the day-to-day change in a liquidity measure, we model the return of the bond as normally distributed between one period (day) and the next. This is a plausible assumption as long as the bond is neither near the maturity date nor in default, which is reasonable for our sample of Italian sovereign bonds.

utility from maintaining the existing portfolio is equal to the expected utility from trading the dollar quantity $Q$:

$$E\left[U\left(W_I\right)\right] = E\left[U\left(W_{I+Q}\right)\right].$$

In Appendix A, we show that the absolute bid-ask spread, calculated as the relative bid-ask spread for purchasing a quantity $Q = p_0$ multiplied by the price of the bond $p_0$, is

$$\begin{aligned} BA = & \frac{\gamma p_0^2 \sigma^2(CDS)}{1+r_f} + b\frac{p_0 - W_0\,(1-k)}{1+r_f} \\ & + m(b, CDS)p_0. \end{aligned} \tag{1}$$

The market maker observes the CDS price ($CDS$) on the CDS derivative market and extracts the (forward looking) volatility of the bond $\sigma(CDS)$. We model the relation between the standard deviation of returns and the CDS price by approximating it with a linear function, as in Brenner and Subrahmanyan (1988), thus deriving $\sigma(CDS)$ as:

$$\sigma(CDS) = \left(1 + r_f\right)\frac{CDS}{p_0\, n\,(0)}, \tag{2}$$

where $n(0) \approx 0.4$ is the probability density function of the standard normal distribution evaluated at 0.[7]

Re-writing the absolute bid ask spread as a function of the CDS price, we obtain the relation between the dependent variable of interest, the absolute bid-ask spread, and its determinants, the CDS price, and the policy parameters set by clearing houses and the central bank:

$$BA(b, CDS) = \delta CDS^2 + m\,(b, CDS)\,p_0 + b\eta, \tag{3}$$

where $\frac{\gamma(1+r_f)}{n(0)^2} = \delta > 0$ and $\frac{p_0 - W_0(1-k)}{1+r_f} = \eta > 0$, and where we emphasize that the margin setting decision by the clearing house depends *both* on the borrowing cost set by the central bank and on the level of the CDS.[8]

Equation (3) features the two channels through which the first determinant of market liquidity, the CDS price,

[7] This is a partial equilibrium analysis; in a general equilibrium model, a change in volatility via *CDS* would also change $p_0$, as the underlying asset price would in a general version of the Black-Scholes model. In our model, therefore, we assume that the asset price is exogenous, and focus on changes in the return volatility. All detailed calculations deriving the model can be found in Appendix A.

[8] The second inequality follows from the requirement that the market maker borrows the residual amount, when buying a bond, by pledging the security at the central bank, as modeled in Appendix A.

affects the bid-ask spread. The first channel, represented by the first term in the equation ($\delta CDS^2$), is a direct one, arising from the market maker's update of the (forward looking) bond volatility, as extracted from the derivative market. The second channel, the second term in the equation ($m(b, CDS)\, p_0$), is an indirect effect of the CDS price through the margin setting decision by the clearing houses, since the clearing houses, like the market maker, extract information about the riskiness of the bond from the CDS market. Our model rationalizes how changes in margins, which depend on the level of the CDS spread (or price), affect the relation between credit risk and liquidity.

A second determinant of market liquidity is the central bank's monetary policy, which affects both the market maker's borrowing costs, through the third term in the equation ($b\eta$), and the second (indirect) channel through which the CDS price affects the liquidity: the margin settings. The monetary policy affects the margin setting decision by the clearing house, which influences the market maker's decision via the second term in the equation ($m(b, CDS)\, p_0$). In the next subsection, we derive the empirical predictions of the model that we test in the data.

### *3.1. Empirical Predictions*

**Empirical Prediction 1.** *The illiquidity of the bond market increases with credit risk.*

This follows from Equation (3), as $\frac{\partial BA}{\partial CDS} > 0$, since $\delta > 0$, $\eta > 0$, and $\frac{\partial m(b,CDS)}{\partial CDS} > 0$. We expect an increase in credit risk to raise the market illiquidity of the bond. As in the Stoll (1978) model, and in line with other inventory models of market microstructure, our model predicts that an increase in the risk of a security, e.g., credit risk, implies a riskier inventory, leading to a withdrawal of liquidity offered to the market by the market maker.

Since we expect the change in credit risk to be a relevant variable in characterizing the *dynamics* of liquidity in the market through the market makers' inventory concerns, we investigate the lead-lag relation between credit risk and illiquidity, and the directionality of this relation.[9]

Moreover, our first empirical prediction is in line with risk management practices based on value-at-risk (VaR) models used widely by market participants, particularly the market makers. A portfolio with an excessively large VaR, due to credit risk, erodes the dealers' buffer risk capacity, which results in the dealer setting higher bid-ask spreads.[10]

[9]We address the contemporaneous interaction between the two variables in detail in Section Int.1 of the internet appendix, via instrumental variables analysis.

[10]This link also has implications for the *dynamics* of the relation between credit risk and market liquidity: The VaR is calculated at the end of day $t-1$. In periods of market stress, however, the VaR is often monitored at an intraday frequency, implying that day-$t$ liquidity will depend on the contemporaneous, day-$t$, credit risk.

**Empirical Prediction 2.** *The dynamic relation between credit risk and market illiquidity shifts conditional on the level of the CDS spread.*

We derive from Equation (3) the sensitivity of the bid-ask spread to the CDS spread, $\frac{\partial BA}{\partial CDS} = 2\delta CDS + \frac{\partial m(b,CDS)}{\partial CDS}$. This sensitivity depends on the CDS spread through two channels: the direct risk channel, and the indirect margin setting channel; Empirical Prediction 2 focuses on the latter. As documented in LCH.Clearnet (2011), the "Sovereign Risk Framework" states that the margin-setting decisions depend on the level of CDS and, particularly, that the clearing house deems that the risk of a security has increased significantly if the 5-year CDS spread increases above 500bp. In our model, this dependence would translate into a shift in $\frac{\partial m(b,CDS)}{\partial CDS}$, when the CDS spread crosses the 500bp threshold.[11]

To test this empirical prediction, we employ the threshold test proposed by Hansen (2000) to investigate i) whether a structural break in the level of CDS is present in the relation between credit risk and liquidity, ii) if this threshold corresponds to 500 bp, and iii) how the relation between credit risk and market liquidity changes, below and above the threshold.[12]

**Empirical Prediction 3.** *The monetary policy interventions of the central bank affect the dynamic relation between credit risk and market liquidity.*

A central bank intervention that targets the access to funding liquidity by banks and market makers would, in our model, affect the sensitivity of the bid-ask spread to the CDS spread by changing the clearing houses margin setting decision, i.e., through $\frac{\partial m(b,CDS)}{\partial CDS}$. In the context of the relation between credit risk and liquidity, therefore, a successful intervention would be one that affects the sensitivity of the market makers to changes in credit risk by providing them with improved funding liquidity. Therefore, we especially expect the LTRO to have an impact, due to the nature of its large funding liquidity shock, qualifying it as a significant structural break, thus affecting the market liquidity in the sovereign bond market through the availability of funding liquidity to market makers. As in Brunnermeier and Pedersen (2009), we expect the margin channel to be have a larger impact on the market maker's liquidity provision when she is funding-liquidity constrained. The availability of massive amounts of medium-

[11] Other related conceptual arguments can be advanced for such a shift in the relation. First, during the Euro-zone crisis, the adverse change in credit quality was generally accompanied or followed by downgrades in the credit rating, altering the clientele of investors who were able to hold Italian sovereign bonds. Second, in the presence of a sharp decline in credit quality, internal (and external) models of risk-weighting and illiquidity used by banks, a major investor segment, would necessarily predict an increase in the capital required to support the higher level of risk.

[12] Appendix B presents the details of the econometrical procedure.

term funding from the ECB, at unusually low interest rates, should have shifted the incentives of dealers to hold sovereign bonds.

Our third empirical prediction investigates the presence of regime shifts in the estimated relation between credit risk and market liquidity around the dates of significant policy interventions by the ECB. Due to the large number of such interventions (SMP, LTRO, OMT, policy guidance, as described above) during the Euro-zone crisis, we choose to allow the data to endogenously inform us of the presence of structural breaks that indicates whether these interventions indeed affected the relation between credit risk and market liquidity. To investigate this issue, we perform a SupWald structural break test, a modified Chow test with an unknown break point (see Chow, 1960; Andrews, 1993; Hansen, 1997). Appendix B presents the procedure in detail.

As argued earlier, the ECB interventions and its moral suasion towards the clearing houses could affect the sensitivity of the market liquidity to the credit risk via the indirect margin channel and, thus, affect the findings established in the previous empirical predictions. Therefore, we replicate the analysis in Empirical Prediction 1 and 2, for the two periods identified by the statistical procedure. Thus, for the two periods separately, we i) quantify the sensitivity of the bid ask spread to the CDS spread, and ii) test whether the relation shifts, when the CDS spread is above a threshold.

## 4. MTS Market Structure and Description of Variables

Our data consist of all real-time quotes, orders, and transactions that took place on the MTS European sovereign bond market during our period of study, and are provided by the MTS Group. These high-frequency data cover trades and quotes for the fixed income securities issued by twelve national treasuries and their local equivalents: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Slovenia, and Spain. The MTS system is the largest interdealer market for Euro-denominated sovereign bonds and is made up of many markets, including the EuroMTS (the "European market"), EuroCredit MTS, and several domestic MTS markets. In this study, we will focus on the liquidity of Italian sovereign bonds, regardless of whether the trading or quoting activity took place on the domestic or the European market. The MTS trading system is an automated quote-driven electronic limit order interdealer market, in which market makers' quotes can be "hit" or "lifted" by other market participants via market orders. EuroMTS is the reference electronic market for European benchmark bonds.[13]

The sample period of our study is from July 1, 2010 to December 31, 2012.[14] The time period we analyze provides a good window in which to study the behavior of European sovereign bond markets during the most recent part of the Euro-zone sovereign debt crisis and the period leading up to it. Our data set consists of 189 Italian sovereign bonds. Table 1 presents the distribution of these bonds in terms of maturity and coupon rate, among original maturity groups as well as bond types. In terms of maturity groups, the bonds are grouped together based on the integer closest to their original maturity. As Table 1 shows, the large majority (in numbers) of the bonds analyzed have short maturities (from 0 to 5 years). All bonds considered in this analysis belong to one of the following types: Buoni Ordinari del Tesoro (BOT), which correspond to Treasury bills, Certificato del Tesoro Zero-coupon (CTZ), corresponding to zero-coupon bonds, Certificati di Credito del Tesoro (CCT), or floating notes, and Buoni del Tesoro Poliennali (BTP), which are coupon-bearing Treasury bonds. The vast majority of the bonds in our sample belong to the BOT and BTP types. We exclude inflation and index-linked securities from our analysis.

INSERT TABLE 1 HERE

### 4.1. *Description of Variables*

We measure bond liquidity for the MTS market by the daily *Bid-Ask Spread*, defined as the difference between the best ask and the best bid, per €100 of face value, proxying for the cost of immediacy that a trader will face when dealing with a small trade. We measure the bid-ask spread per bond at a five-minute frequency from the market open to the market close, namely from 8 AM to 5.30 PM, then average it per bond throughout the day, and finally average the daily bond measures across bonds to obtain a market-wide daily liquidity measure.

The Italian-sovereign-specific credit risk is measured by the spread of a senior five-year dollar-denominated CDS contract obtained from Bloomberg. The choice of this proxy for sovereign credit risk is debatable. An alternative potential proxy for Italian sovereign risk could be the BTP-Bund yield spread. We prefer to avoid using the BTP-Bund yield spread because this variable is likely to be intimately connected to the bond quote and

[13]Benchmark bonds are bonds with an outstanding value higher than €5 billion. Section Int.2 of the internet appendix provides details of the market architecture, trading protocol, and data released for the MTS market; see also Dufour and Skinner (2004).

[14]Our data set from July 2010 to May 2011 includes only intraday updates of the three best bid and ask quotes. From June 1, 2011, we have detailed tick-by-tick, second-by-second, data. The end date is dictated by a major change in market structure that was implemented in December 2012, and that changed the role of market makers acting in the European section of the MTS market. Fortuitously, the period we consider covers a large part of the Euro-zone crisis. A more detailed description of the differences between the datasets can be found in the internet appendix, Section Int.2.

transaction prices that are also used to calculate our liquidity measures. CDS spreads are obviously related to the BTP-Bund yield spread (as Figure 2 shows), through arbitrage in the basis between them, but at least are determined in a different market.[15]

INSERT FIGURE 2 HERE

Finally, in order to control for and characterize the effect of global credit risk and funding liquidity, we employ several macro-economic indicators, most of which are common in the academic literature. The Euribor-DeTBill yield spread captures the (global) counterparty and credit risk and, thus, an increase in the cost of funding, and is measured as the difference between the three-month Euro-area Inter-Bank Offered Rate (Euribor) for the Euro, covering dealings from 57 prime banks, and the three-month yield of the three-month German Treasury bill. As banks are more uncertain, they charge each other higher rates on unsecured loans; similarly, looking for high-quality collateral, they purchase safe Treasury bills, lowering their yields. This measure is the European counterpart of the TED spread used by, among others, Brunnermeier (2009). The USVIX, measuring global systemic risk, is the implied volatility index of S&P 500 index options calculated by the Chicago Board Options Exchange (CBOE) and used widely as a market sentiment indicator. The CCBSS represents the additional premium paid per period for a cross-currency swap between Euribor and US Dollar Libor, and serves as a proxy for funding liquidity.[16] All these variables were obtained from Bloomberg.

## 5. Descriptive Statistics

Table 2 presents the summary statistics for the market activity measures for Italian sovereign bonds traded on the MTS market and system variables, between July 2010 and December 2012, spanning the period of the Euro-zone sovereign crisis. The table reports statistics for the daily time-series of the market-wide variables: *Trades*,

[15]We show in Section Int.3 of the internet appendix that there is no statistically significant lead-lag relation between the two daily series, because the adjustment between them takes place on the same day. Also, in Section Int.4 of the internet appendix, we investigate whether the intraday volatility of the bond yield, as measured using the MTS transaction data, and the liquidity of the CDS market affect the liquidity, while controlling for the credit risk. These modifications do not significantly change the results, supporting our choice of the CDS spread as a measure of credit risk.

[16]The CCBSS can be thought of as the spread of the longer-term, multi-period equivalent of deviations from uncovered interest rate parity. When liquidity is available to arbitrageurs in all currencies, deviations from the (un)covered interest rate parity will be closed and profited on, while lasting deviations can be interpreted as a sign of lack of funding liquidity. Baba, Packer, and Nagano (2008) and Baba (2009) show that cross-currency basis swaps are used by banks to finance themselves in foreign currencies when the interbank market in the home currency is illiquid, Brunnermeier, Nagel, and Pedersen (2008) show that deviations from uncovered interest rate parity are partially explained by shocks to funding liquidity. Acharya and Steffen (2015) and Ivashina, Scharfstein, and Stein (2012) investigate the funding liquidity needs of European banks and relate them to the (un)covered interest rate parity.

*Volume*, and *Bid-Ask Spread* were calculated on a daily bond basis and then averaged across bonds to obtain the time-series. *Quoted Bonds* is the time-series of the number of bonds quoted each day.

INSERT TABLE 2 HERE

The mean (median) number of bonds quoted each day on the MTS is 89 (88), and the daily volume of trading in the market is slightly below €2.9 billion (€2.6 billion), which translates into a daily traded volume for each quoted bond of about €32.6 million (€28.7 million). Based on these numbers, the daily trading volume in the Italian sovereign bond market (as represented by the MTS) is much smaller than in the US Treasury market, by a couple of orders of magnitude, with the average traded quantity in the latter being around $500 billion per day (Bessembinder and Maxwell, 2008). The average daily trading volume in the MTS Italian bond market is even smaller than in the US municipal market (around $15 billion), the US corporate bond market (around $15 billion), and the spot US securitized fixed income market (around $2.7 billion in asset-backed securities, around $9.1 billion in collateralized mortgage obligations, and around $13.4 billion in mortgage-backed securities).[17]

Our volume statistics are in line with the stylized facts documented in the previous literature, taken together with the consistent shrinkage of overall market volumes since the Euro-zone crisis began. Darbha and Dufour (2013) report that the volume of the Italian segment of the MTS market as a whole, over their 1,641-day sample, was €4,474 billion. This translates into an average daily volume of about €3.8 billion. Darbha and Dufour report that the daily volume per bond shrank from €12 million in 2004 to €7 million in 2007. Their sample includes only coupon-bearing bonds; thus, their figures for overall market volume are not directly comparable to ours.

The daily number of trades on the MTS Italian sovereign bond market is 352 in total (or about 4 per bond), which is similar to the 3.47 trades a day per corporate bond on TRACE, as reported in Friewald, Jankowitsch, and Subrahmanyam (2012a). Dufour and Nguyen (2012) report an average of 10 trades per day per Italian bond in an earlier period, between 2003 and 2007. As with the trading volume, the number of trades declined during the crisis period compared to earlier years. Our sample period covers the most stressed months of the Euro-zone crisis, when the creditworthiness of several European countries was seriously questioned by market participants. As we will show later, the liquidity in the MTS market was intimately related to the evolution of spreads in the sovereign CDS market, and varied just as drastically, as the time-series plots of the CDS spread and the *Bid-Ask Spread* in Figure

[17]Details for the corporate bond, municipal bond, and securitized fixed income markets are provided in Friewald, Jankowitsch, and Subrahmanyam (2012a), Vickery and Wright (2010), and Friewald, Jankowitsch, and Subrahmanyam (2012b), respectively.

2 show. Up to the end of 2011, at the peak of the crisis, the two series share a common trend, which is not repeated in the second half of our sample.

The commonality in the two series in Figure 2 becomes particularly evident, for example, when one considers the highest spike for the *Bid-Ask Spread* (€4.48 per €100 of face value), which happened on November 9, 2011. On the previous day, after the markets had closed, the Italian Prime Minister, Silvio Berlusconi, lost his majority in the parliament, which led to his resignation. The spike in the *Bid-Ask Spread* corresponds to a similar spike in the *CDS Spread*. The event clearly had medium-term effects, as both the *Bid-Ask Spread* and the *CDS Spread* persisted at high levels for about two months, before returning to more moderate quantities in January 2012. In mid-2012, however, the *CDS Spread* reached levels close to 500 bp, while the *Bid-Ask Spread* oscillated around the time-series median value of €0.30.

The reasons for choosing to present our results based on the bid-ask spread as a measure of market liquidity bear mention. First, the quoted bid-ask spread is the most familiar and widespread measure of market liquidity. Thus, it allows for a direct comparison with the previous and contemporaneous literature on liquidity. Second, the large number of quotes that are aggregated into a single daily bid-ask spread time-series suggests that market makers are very active, and ensures that the computed spread is a precise estimate of their willingness to trade, since the quotes are firm. Finally, high-frequency quote updates indicate that accurate quoting in the MTS market is important for primary dealers under the supervision of the Bank of Italy. These quotes are, moreover, also used by officials at the Italian Treasury to evaluate (and eventually even disqualify) sovereign bond market makers.[18]

The results of the Dickey-Fuller unit root test for the variables used in our empirical investigation are presented in Table 2 under the "Unit Root Test" columns for the levels of and differences in the variables. All our tests for the control variables and the CDS spread support the existence of a unit root, while the bid-ask spread and the USVIX show a mean-reverting property. However, (i) the first-order auto-correlation for the liquidity measure is 81%, and (ii) the unit root test did not reject the unit root null hypothesis when it was performed on the first part of the sample, for the period when the Euro-zone crisis first unfolded. In light of this fact, and in order to have a consistent, unique model for the whole data sample and to ensure well-behaved residuals, we perform our analysis

[18]From July 1, 2010 until May 31, 2011, we use the MTS database that provides only the three best bid and ask prices. However, we have an overlapping sample of seven months of both the databases, and perform a comparison of the bid-ask liquidity measure we calculate, using the two databases. The results show that there is almost no difference between the two, for the purpose of computing the bid-ask spread; see Section Int.2 of the internet appendix.

in first differences.

As shown in Figure 2, the Italian CDS spread for our sample period ranges from 127 bp to 592 bp, with a mean of 321 bp and a standard deviation of 138 bp, indicating the large changes in this variable during the period under study. Figure 3 shows the evolution of the macro variables. The Euribor-DeTBill spread (Panel (a)) also presents a significant level of volatility, with a daily standard deviation of 0.36%, while the USVIX (Panel (b)) ranges from 13.45% to 48%. The CCBSS variable (Panel (c)), which captures the general level of funding liquidity in the system, and which should be close to zero in the absence of funding constraints, ranges from 12bp to 107bp, indicating a large variability in the global liquidity conditions in the Euro-zone in the period considered. All the funding and credit variables suggest that the conditions in the Euro-zone financial system were at their worst around the third quarter of 2011, but improved somewhat during the first quarter of 2012, then worsened, although to a lesser extent, around June 2012, and continued to decline towards the end of that year.

INSERT FIGURE 3 HERE

The correlations between the credit, funding liquidity and market liquidity variables are shown in Table 2 Panel C. The correlations between the variables in levels are presented above the diagonal, while those for the variables in differences are below the diagonal. In differences, bond market liquidity is most highly correlated with the Italian *CDS Spread* and the CCBSS.

## 6. Results

In Section 3 we derived three empirical predictions and, in this section, we investigate them, focusing on the dynamic relationships between credit risk and market liquidity and the effect of the ECB's *deus ex machina*. In order to test the first empirical prediction, regarding the dynamics of the relation between the credit risk of Italian sovereign bonds, as measured by the *CDS Spread*, and the liquidity of the Italian sovereign bonds, as measured by their *Bid-Ask Spread*, we first investigate, in Section 6.1, whether there is a lead-lag relation between the two variables, using a Granger-causality test in a Vector Auto Regression (VAR) setting.[19]

[19]We conduct our analysis in changes, after winsorizing the data at the 1% level to diminish the importance of outliers, such as the large changes in bid-ask spread in the second half of 2011, in particular that of November 9. For robustness, we repeat the analysis after winsorizing the data at the 5% level. The results are mostly unchanged and reported in the internet appendix Section Int.5.

In Section 6.2, we focus on Empirical Prediction 2, and test for the presence of a threshold in the level of the CDS spread that shifts the relation between credit risk and market liquidity. We perform this analysis using the threshold test proposed by Hansen (2000), and characterize how the relation between credit risk and market liquidity changes below and above this threshold. Finally, in Section 6.3, we investigate Empirical Prediction 3 and test whether and how the dynamics of the relation are affected by the ECB interventions. We use an endogenous structural break test described in detail in Appendix B, and study whether the injection of funding liquidity by the central bank lowered the sensitivity of market liquidity to the worsening credit conditions of the Italian sovereign.

### *6.1. The Dynamics of Credit Risk and Liquidity*

**Empirical Prediction 1.** *The illiquidity of the bond market increases with credit risk.*

In this section, we investigate Empirical Prediction 1, testing whether the increase in credit risk drives the reduction of market liquidity or vice versa. While our theoretical model has been explicitly designed to characterize the effects that a change in the credit risk has on the market liquidity, we cannot rule out that market liquidity has, in turn, an effect on credit risk. Therefore, to allow for this feedback loop, we implement this analysis by estimating a VAR system that allows us to perform a Granger-causality test. Since global risk factors could affect market liquidity, on top of security-specific credit risk concerns, we include USVIX, the Euribor-DeTBill spread, and the CCBSS in our VAR specification as "exogenous variables". These variables are exogenous in that we are not interested in studying the effect of the endogenous variables on their dynamics, only the opposite effect. We thus describe the system using a VAR with eXogenous variables (VARX) model.

The mathematical formulation of this Granger-causality test is based on linear regressions of the change in the *Bid-Ask Spread*, $\Delta BA_t$, and the change in the *CDS Spread*, $\Delta CDS_t$, on their $p$ lags. Specifically, let $\Delta BA_t$ and $\Delta CDS_t$ be two stationary daily time-series, and $X_t$ a time-series $m-$vector of stationary exogenous variables. We can represent their linear inter-relationships using the following VARX model:

$$\begin{pmatrix} \Delta BA_t \\ \Delta CDS_t \end{pmatrix} = \begin{pmatrix} K_{BA} \\ K_{CDS} \end{pmatrix} + \sum_{i=1}^{p} \begin{pmatrix} a_{11_i} & a_{12_i} \\ a_{21_i} & a_{22_i} \end{pmatrix} \begin{pmatrix} \Delta BA_{t-i} \\ \Delta CDS_{t-i} \end{pmatrix}$$
$$+ \sum_{j=0}^{q} B_j \begin{pmatrix} \Delta X1_{t-q} \\ \Delta X2_{t-q} \\ \vdots \\ \Delta Xm_{t-q} \end{pmatrix} + \begin{pmatrix} \epsilon_{BAt} \\ \epsilon_{CDS\,t} \end{pmatrix}, \tag{4}$$

where $\epsilon_{\mathbf{t}} \sim N(\mathbf{0}, \mathbf{\Omega})$, the $B_j$s are 2-by-m matrices, and the $a_{ij_p}$s are the $p$-lag coefficients of the model. This formulation allows for the presence of $m$ contemporaneous, and lagged (up to $q$), exogenous variables to control for factors that might affect the dynamics of the endogenous variables. We can conclude that $\Delta CDS$ Granger-causes $\Delta BA$ when the $a_{12_p}$s are contemporaneously different from zero. Similarly, we can surmise that $\Delta BA$ Granger-causes $\Delta CDS$ when the $a_{21_p}$s are contemporaneously different from zero. When both these statements are true, there is a feedback relation between the two time-series.

The lag length was chosen based on the corrected Akaike criterion, which suggests a lag length of 3 for the endogenous variables and no lagged exogenous variables. The results of the Granger-causality test, with $p = 3$ and $q = 0$, for the relation between the changes in the *CDS Spread* and the *Bid-Ask Spread*, are reported in Table 3, where we report the $F$-test test statistics for the contemporaneous significance of the cross-variable terms for each equation (the $a_{12}$s for the bid-ask spread equation under $\Delta BA_t$, and the $a_{21}$s for the CDS spread equation under $\Delta CDS_t$).[20]

INSERT TABLE 3 HERE

As the table shows, in line with Empirical Prediction 1 in Section 3, the *CDS Spread* Granger-causes liquidity in the bond market at a 1% level (the heteroskedasticity-robust $F$-test is 6.01 and the 1% confidence value is 3.81, and the bootstrapped results provide identical significance levels), while the opposite directionality is not significant at any of the usual confidence levels (the $p$-value is 0.70). This result confirms Empirical Prediction 1 and supports

[20]Throughout the paper, statistical significance is always determined on the basis of $t$-tests that are calculated using heteroskedasticity-robust standard errors.

the inventory risk channel as a driver of the relation between credit risk and market liquidity.

The macro variables are significant in explaining the two variables. Specifically, the bond market illiquidity depends positively on the availability of funding liquidity for European banks and on the sentiment of the market, as measured by the $CCBSS$ and $USVIX$, respectively. In untabulated results, however, the contemporaneous dependence of the macro variables does not lower the significance of the effect of (lagged) credit risk on market liquidity, although it contributes towards lowering the residual cross-correlation.

In order to interpret the dynamics of the system, we calculate the impulse response functions (IRF) for the relationships between the variables. We do this for the rescaled variables, so that they have a mean of 0 and a standard deviation of 1, for ease of interpretation. Figure 4 presents the results, for which the 5% confidence bands were bootstrapped based on 5,000 repetitions. As shown in Panel (a) of the figure, a one-standard-deviation shock to the *CDS Spread* at time 0, corresponding to a 4.1% change, is followed by a change of 0.26 standard deviations in the *Bid-Ask Spread*, corresponding to a 5.2% increase in the same direction, and is absorbed by both variables in two days. Alternatively, the parameters imply that a 10% change in the *CDS Spread* (corresponding to a change of $10\%/4.1\% = 2.43$ standard deviations) is followed by a $2.43 \cdot 5.2 = 12.7\%$ change in the *Bid-Ask Spread*. The results are, hence, both statistically and economically significant, and confirm the results of the Granger-causality tests presented above. The IRF in Panel (b) shows that a shock at time 0 to market liquidity lasts until time 1, but only affects market liquidity itself, indicating that the reaction of the *CDS Spread* to a shock in market liquidity is never different from zero, in line with the findings of the Granger-causality tests.

INSERT FIGURE 4 HERE

Since the focus of this study is the dynamics of the credit risk and bond market liquidity in relation to each other, and past values of bid-ask spread do not affect credit risk, as per Table 3, we focus solely on the bid-ask spread regression in the VARX system, augmenting it with the contemporaneous change in credit risk. This corresponds to a shift from a reduced-form to a structural approach for the VAR, where the contemporaneous causation runs from credit to liquidity. As the ordering of the variables in this causation chain cannot be tested in the VAR setting (see, e.g., Lütkepohl, 1993), we turn to instrumental variable (IV) methods to establish whether feedback between the contemporaneous *CDS Spread* and *Bid-Ask Spread* changes—or, alternatively, other forms of endogeneity—is supported by the data. We do so to ensure that our specification does not disqualify the structural approach we take, or otherwise suggest the opposite relation. In Section Int.1 of the internet appendix, we show using several cohorts of valid and strong instruments that the *CDS Spread* is indeed not endogenous to the system, and hence its inclusion as a regressor is justified: the regression parameter attached to it in the bid-ask spread regression is unbiased and consistently estimated.

As both the lead-lag and the contemporaneous relation indicate the direction of the Granger-causality, we only need focus in the rest of the paper on the causal effects on the liquidity measure (i.e., the $\Delta BA_t$ equation), in order to determine the dynamics of the system. This will be sufficient to capture the dynamics of the credit-liquidity relation (including the effect of ECB interventions), given the lack of statistical support for causality in the opposite direction. Therefore, we regress changes in the liquidity measure, *Bid-Ask Spread*, on the contemporaneous changes in the *CDS Spread*, and their respective lags, and on the contemporaneous macro variables. Equation (5) presents our baseline regression specification for the remainder of the paper:

$$\begin{aligned}\Delta BA_t = \alpha_0 + \sum_{i=1}^{3} \alpha_i \Delta BA_{t-i} + \sum_{j=0}^{1} \beta_j \Delta CDS_{t-j} \\ + \beta_2 CCBSS + \beta_3 USVIX_t + \epsilon_t,\end{aligned} \tag{5}$$

where $\Delta BA_t$ is the change in the bond-market-wide bid-ask spread from day $t-1$ to day $t$, and $\Delta CDS_t$ is the change in the CDS spread, as before. The statistically insignificant lags of the CDS measure and $\Delta EuriborDeTBill_t$ have been dropped due to their lack of statistical significance. The results for Equation (5) are reported in Table 4 Panel A.

Comparing the parameters in Table 4 Panel A to those in Table 3 shows that adding the contemporaneous change in the *CDS Spread* does not modify our findings, with the exception of a lower level of statistical significance for the other contemporaneous variables. This was to be expected, since these other variables potentially proxy for changes in the credit risk. Moreover, the dynamics of the bid-ask spread are well accounted for, since the residuals show no autocorrelation according to the Durbin $h$-test and the Breusch-Godfrey serial correlation test (never significant at the 10% level or lower for lags up to 10, with one exception).

INSERT TABLE 4 HERE

As for the dynamics of the system, the change in the *CDS Spread* has a lagged effect on market liquidity, i.e., the reaction of market liquidity, measured by the *Bid-Ask Spread*, to changes in the *CDS Spread*, occurs on the next day. The *Bid-Ask Spread* also shows evidence of an autoregressive component, being strongly related to the change in the *Bid-Ask Spread* that took place the day before, with a negative sign: this suggests an overreaction adjustment dynamic in the *Bid-Ask Spread*, as shown already in the IRF of Figure 4 Panel (b). This effect can be ascribed to the actions of the market makers, who adjust their quotes as a reaction, not only to the changes in the traded price, but also to the changes in the quotes of the other primary dealers. A 10% increase in the CDS spread on day $t$ results in an increase in the bid-ask spread of 5.41% on day $t$ and a further increase of $-0.352 \cdot 5.41\% + 7.94\% = 6.04\%$ on day $t + 1$, for a cumulative increase of 11.45%.

Regarding the significance of the lagged $\Delta CDS$ term, a partial explanation can be found in the timing of VaR-based models in practice. Since the calculation of the dealer's VaR generally takes place at the end of the day, the exposure to the credit risk is taken into account by the dealer when deciding how much liquidity to offer only on the day following the credit shock, which implies the significance of the lagged change in credit risk.[21]

### *6.2. The relation between Credit Risk and Liquidity Conditional on the Level of Credit Risk*

**Empirical Prediction 2.** *The dynamic relation between credit risk and market illiquidity shifts conditional on the level of the CDS spread.*

Turning to our Empirical Prediction 2, Equation (5) above implicitly assumes that the estimated relation holds independent of the *level* of credit risk, in particular when the *CDS Spread* is above a particular threshold level. For the reasons highlighted in the theoretical model presented in Section 3, on account of margin setting, and downgrade concerns, it is possible that the market makers' liquidity provision would be more sensitive to changes in credit risk when the *CDS Spread* breaches a particular threshold. We investigate this empirical prediction by allowing the data to uncover the presence of a threshold in the level of the *CDS Spread*, above which a *different* relation between changes in CDS and changes in market liquidity is observed. We use the test proposed by Hansen (2000), described in detail in Appendix B, to examine this hypothesis, estimating Equation (6) for different $\gamma$, where $I[CDS \leq \gamma]$ equals 1 if the condition is satisfied and 0 otherwise:

[21] One variable that could also affect the inventory levels of market makers (e.g., through the risk management practices of dealer desks), and therefore market liquidity, is the volatility of the bond yield. In Section Int.4 of the internet appendix, we repeat the analysis including this variable and our results are robust to this inclusion. Moreover, we also test whether the *CDS Spread* drives both changes in market liquidity and bond return volatility or whether the effects are the other way around, and show that it is the former relation that prevails, confirming that the analysis we perform in this section is correct and robust to the insertion of volatility into the pool of endogenous variables.

$$
\begin{aligned}
\Delta BA_t = \alpha_0 &+ \sum_{i=1}^{3} \alpha_i \Delta BA_{t-i} \\
&+ I\left[CDS \leq \hat{\gamma}\right]\left(\sum_{j=0}^{1} \beta_j \Delta CDS_{t-j}\right) \\
&+ I\left[CDS > \hat{\gamma}\right]\left(\sum_{j=0}^{1} \tilde{\beta}_j \Delta CDS_{t-j}\right) \\
&+ \beta_2 USVIX_t + \beta_3 CCBSS + \epsilon_t. \qquad (6)
\end{aligned}
$$

Figure 5 shows, on the $y$-axis, the sum of squared residuals for the regression in Equation (6) as $\gamma$, shown on the $x$-axis, changes (the sum of squared residuals for Equation (5) is plotted at $\gamma = 0$). The sum of squared residuals is minimized when $\gamma = 496.55$. We test for the identity between parameters above and below the threshold, or, equivalently, for the presence of the threshold, $H_0 : \beta_0 = \tilde{\beta}_0, \beta_1 = \tilde{\beta}_1$ and, since the test statistic asymptotic distribution is non-pivotal, we bootstrap it, as described in Hansen (1996). The test statistic for the presence of the threshold we observe (25.05) is significant at better than the 1% level, thus confirming the presence of a threshold.[22]

INSERT FIGURE 5 HERE

While the previous paragraphs confirm the presence and location of the threshold, $\hat{\gamma} = 496.55$ bp, Figure 6 shows the test statistic needed to determine the confidence bounds around the point estimate we find. The threshold has a point estimate of 496.55, with a 5% confidence interval between 485 and 510, and is almost identical for various alternative specifications of the relation (including whether or not lagged or macro variables are included).

INSERT FIGURE 6 HERE

The confirmation of the presence of a structural shift in the data when the CDS spread crosses a certain threshold is, therefore, robust and strongly supported by the data, and indicates how important the level of the *CDS Spread* is for market liquidity.

[22]The histogram of the bootstrapped test distribution for this and similar tests referred to throughout the paper can be found in the internet appendix, Section Int.6.

This result confirms Empirical Prediction 2, and shows that the rules adopted by the clearing house to set margins as a function of the level of the CDS spread have an impact on the relation between credit risk and market liquidity. The application of the margin setting rule is shown in Figure 7, which depicts the time-series of bond-market bid-ask spread, CDS spread, and the average margin on Italian bonds with between 3 months and 30 years to maturity, charged by a major clearing house, Cassa Compensazione e Garanzia, which uses the same margins as those charged by LCH.Clearnet. The margin requirements changed only slightly between June 2010 and November 2011, from 3.26% to 4.53%, while the CDS spread rose threefold from about 150 bp to 450 bp. However, the same clearing house nearly doubled the margins to slightly below 9% on November 9, 2011, the second time the spread hits and stays consistently above 500 bp: in the sovereign risk framework, distributed by the LCH.Clearnet in October 2010 (see LCH.Clearnet, 2011), one of the indicators used to justify a hike in margin is indeed "a 500bp 5 year CDS spread".

It is important to stress that market participants were aware of the rule adopted by the clearing house, which had already enforced this margin setting rule for Irish sovereign bond on November 17, 2010, when the margins on repo transactions were raised from 16-18% to 31-33%. In that instance, LCH.Clearnet argued that this decision had been taken "in response to the sustained period during which the yield differential of 10 year Irish government debt against a AAA benchmark has traded consistently over 500 bp."[23]

INSERT FIGURE 7 HERE

The very day that the clearing houses changed the margins charged on sovereign bonds, their market liquidity suddenly worsened, corresponding to a shift in the level of the bid-ask spread, as predicted by Equation (3) in our model. Brunnermeier and Pedersen (2009) derive a similar prediction, that an increase in margins has an effect on the security's market liquidity, if the market makers' budget constraint is binding. As Figure 3 Panel (c) shows, the CCBSS, measuring the funding liquidity needs of the market makers, was at its highest during the second half of 2011, when the margin changes took place. We interpret our findings as a confirmation of Brunnermeier and Pedersen (2009): In the second half of 2011, when the funding liquidity of the market makers was at its lowest and their budget constraint was binding, a change in the margins charged on sovereign bonds led to a tightening of their market liquidity.

[23]Source: `http://www.lchclearnet.com/risk_management/ltd/margin_rate_circulars/repoclear/2010-11-17.asp` and `http://ftalphaville.ft.com//2010/11/17/407351/dear-repoclear-member/`

Having now identified the presence of a threshold and the effect that it has on the level of bid-ask spread, we need to determine how the *sensitivity* of market liquidity to credit risk is modified when the threshold is breached. Panel B of Table 4 reports the results for Equation (6), when $\gamma = \hat{\gamma}$, or the threshold is the point estimate found in the previous paragraphs, what we call for simplicity the 500 bp threshold. The column "Test" in Panel B reports the test statistic for whether each pair of parameters above and below the threshold is equal; e.g., the test statistic for H0: $\beta_0 = \tilde{\beta}_0$ is 11.33, significant at the 1% level.

As the panel shows, the relationships below and above 500 bp are rather different from each other: contemporaneous changes in the *CDS Spread* have a significantly larger economic impact on market liquidity above the threshold of 500 bp than below. In particular, the regression in Panel B indicates that the coefficient of the contemporaneous change below the threshold is 0.32, but not significant, while that above it is 2.85 and statistically significant. Looking at the lagged CDS variable, we find that, below the 500 bp threshold, market liquidity reacts with a lag to changes in the *CDS Spread*, with a significant impact of the autoregressive component and the lagged component of the change in the CDS. Above 500 bp, the relation is rather different: market liquidity reacts immediately to changes in the *CDS Spread*, with the impact being largely contemporaneous, since the change in the CDS spread has no impact on the change in the market liquidity the following day. The parameters suggest that an increase in the *CDS Spread* of 10% on day $t$, below (above) the threshold of 500 bp, induces a contemporaneous increase in the *Bid-Ask Spread* of 3.2% (28.5%) on day $t$, and an increase (decrease) of $-0.332 \cdot 3.2\% + 9.83\% = 8.77\%$ ($-0.332 \cdot 28.5\% - 8.54\% = -18\%$) on day $t + 1$, for a cumulative increase of 11.96% (10.46%). Although the cumulative $t + 1$ effects of a 10% increase in CDS spread are similar above and below the 500 bp threshold, the dynamics of the system are very different: Above 500 bp, the market overreacts by increasing the bid-ask spread instantaneously, while below 500 bp the market reacts moderately, and with a lag, to the increase in credit risk.

The results that we derive in this sub-section for market-wide measures are confirmed by the robustness analysis we perform in Section 7.1, where we group bonds with similar maturities, as determined by counterparty clearing houses with regard to margin requirements, and repeat the analysis regressing each group maturity on the corresponding maturity CDS spread.

Our conclusion, therefore, is that Empirical Prediction 2 is verified and that the dynamic relation between credit risk and market liquidity differs depending on the level of the CDS spread; specifically, in a stressed environment, credit shocks have an immediate impact on market liquidity.[24]

### 6.3. Policy Intervention and Structural Breaks

Our third empirical prediction is that the various interventions that occurred during the period could have generated a structural break in the relation between credit risk and market liquidity. Therefore, the third research aim of this paper is to examine whether such a structural break can be detected statistically and related to policy changes. Again, we let the data alert us to the presence of a structural break over time.

**Empirical Prediction 3.** *The monetary policy interventions of the central bank affect the dynamic relation between credit risk and market liquidity.*

As we have described above, the period that we investigate has been characterized by many events: the onset of the Euro-zone sovereign debt crisis, several sovereign credit downgrades, a political crisis that induced changes in Euro-zone governments, and several interventions by European central banks, and, in particular, by the ECB. Of course, by virtue of its status as the central bank of the Euro-zone, the ECB has a major influence on its sovereign bond markets. As described in Section 3, the ECB's monetary intervention takes many forms, ranging from formal guidance by its board members, in particular its president, to the injection of liquidity into the major banks in the Euro-zone, which themselves hold these bonds, to direct purchases of sovereign bonds in the cash markets.

The purpose of this section is not to quantify the direct effect of these interventions on the Euro-zone credit risk (see Eser and Schwaab, 2013), or its bond market liquidity (see Ghysels, Idier, Manganelli, and Vergote, 2014), but to examine whether the relation between credit risk and liquidity was significantly altered by one or more of these interventions, as exemplified in the theoretical model presented above, by testing for the presence of a structural break. The scant public availability of data concerning the quantity, issuer nationality, and timing of purchases of bonds in the SMP framework prevents us from quantifying the specific effect of those purchases. Similarly, in the absence of details of the extent of banks' access to LTRO funding and its usage, we are unable to investigate how the refinancing operation affected liquidity provision by the market makers (most of which belong to major international and national banks). However, since the several interventions and policy-relevant events took place over finite and non-overlapping periods of time, we can investigate econometrically whether a structural break in the

[24]Since we have determined the presence of parameter discontinuity, we should verify how that discontinuity affects the lead-lag relation investigated in Empirical Prediction 1 for the two samples. Our analysis shows that the same result applies whether the CDS level is below or above the threshold, as shown in Section Int.7 of the internet appendix.

relation between the two variables of interest occurred around the time of the announcement or implementation of the interventions. This analysis is relevant for our second empirical prediction for two main reasons: first, because if the data indeed exhibit structural breaks, our results will be biased if we ignore them, and second, because it will shed light on the relevant combination of conditions that affects the relation between credit risk and liquidity.

We investigate Empirical Prediction 3 by performing the "structural change breaks" test proposed by Andrews (1993) (the *supF* test in that paper), on Equation (6), the details of which are presented in Appendix B. Briefly, the test corresponds roughly to a Chow (1960) test but, while in the Chow test the structural change break is specified exogenously, this "structural change break" allows us to leave the structural break date unknown *a priori*. The test corresponds to performing a Chow test for the relation in question on each date in the sample. The date that is most likely to constitute a break in the data sample is found endogenously, identified as the date with the largest Chow test value, and the presence of a break itself is tested by comparing that date's (Chow) $F$-test statistic to a non-standard distribution. The test, therefore, verifies whether there is a structural break, at all, in the specified relation. If the null hypothesis of "no structural break" can be rejected, the date with the largest corresponding Chow test statistic will be selected as the structural break. Figure 8 shows the values of the Chow $F$-test statistic calculated on each date, with the horizontal line showing the confidence band for the highest $F$-value.

We find that, from a statistical perspective, the test indicates a break, on December 21, 2011, for the relation between the *Bid-Ask Spread*, and the *CDS Spread*, its lag, and the macro variables, and that this structural break is significant at the 10% level. Although December 21 is identified purely based on the statistical evidence as the date for which the (Chow) *supF* test is most significant for the relevant relationships between the *Bid-Ask Spread* and the *CDS Spread*, it coincides *exactly* with the date of the allotment and the day before the settlement of the LTRO program by the ECB.[25]

Our evidence suggests that the relation between credit risk and liquidity changed when the ECB provided LTRO funding to the banks. To the extent that the relation measures the sensitivity of the market makers' behavior to changes in the (credit) risk of their portfolios, our finding supports our empirical prediction, that the market makers were wary about providing liquidity to the sovereign bond market.

They were particularly concerned that, should an adverse credit event have occurred, their inventory would

[25]The policy implementation announcement of December 8, 2011 with all the important dates for this measure can be found online at `http://www.ecb.europa.eu/press/pr/date/2011/html/pr111208_1.en.html`

have suffered and they would have been left with no available funding liquidity. The large provision of funding from the ECB constituted a structural break in that relation and had a clear impact on the sensitivity of market makers to changes in the credit riskiness of their inventories, as we quantify in the following paragraphs.

In order to account for this structural break in our estimations, we split the sample into two periods, and again perform the threshold test as per Equation (6) in both subsamples. That is, we test whether the relation between the changes in the bid-ask spread, and the changes in the CDS spread and its lag, varies above and below an endogenously found threshold. The bootstrap procedure for the threshold test confirms the presence of different relationships below and above the threshold level of 500 bp for the CDS spread, in the first subsample (July 1, 2010 to December 21, 2011), but fails to identify a threshold for the second subsample. Figure 9 reports the test to identify confidence bands around the threshold point estimates, for the first and second subsample, in Panel (a) and (b), respectively: the threshold can be identified around 500 bp for the first subsample, while no threshold can be found in the second subsample.

This result suggests that, thanks to the assurance of a massive amount of liquidity from the ECB and the ECB's request to the clearing house to avoid the possibility for margins to become procyclical to sovereign risk, the relation between changes in the CDS spread and market liquidity was not altered when the Italian CDS Spread breached the level of 500 bp after the LTRO intervention, in contrast to the period before the intervention.

INSERT FIGURES 8 AND 9 HERE

Panel A of Table 5 presents the results of the estimation for the first subsample, before December 21, 2011, and confirms the results we presented above. The main difference is that, for the split sample, the relation between the change in the *CDS Spread* and market liquidity, when the *CDS Spread* is above 500 bp, is even stronger in the pre-LTRO regime, with a 10% increase in the *CDS Spread* translating into a 39% contemporaneous increase in the *Bid-Ask Spread*.

INSERT TABLE 5 HERE

Table 5 Panel B presents the results of the estimation for the second subsample, after December 21, 2011, and shows that the presence of the autoregressive component in market liquidity is still apparent. However, the contemporaneous relation between changes in the *CDS Spread* and changes in market liquidity is no longer significant, while there is a lagged adjustment of market liquidity related to changes in the *CDS Spread* on the previous day, with an economic intensity that is smaller that in the full sample reported in Table 4, Panel A (0.566 vs. 0.794), and about a half of the corresponding parameter for the 2011 subsample, when the CDS is below 500 bp, reported in Table 5 Panel A (0.566 vs. 1.028). Moreover, our analysis shows that the global risk variable, USVIX, affects market liquidity only for the 2011 subsample, while, after the ECB intervention, the only significant variable is the funding liquidity measure, CCBSS.

The previous literature (e.g., Eser and Schwaab, 2013; Ghysels, Idier, Manganelli, and Vergote, 2014) shows that the SMP had an effect on the yields of the bonds chosen for the program, following the large buying pressure exerted by the central bank purchases. However, to the extent that the risk levels of the market makers were maintained, the relation between credit risk and liquidity would have remained unaltered. Hence, the SMP, which was implemented in 2010, did not, in fact, constitute a structural break for that dependence. The LTRO, on the contrary, constituted a massive intervention targeting the availability of funding liquidity and, as such, was ideal for affecting how the banks disposed of their available capital, making them less sensitive to changes in credit risk, when providing liquidity to the market. We tested whether other structural breaks would emerge from the data after December 21, 2011, and no date emerged as statistically significant.

It is worth stressing that, although margins were increased again in June, July, and August 2012 (in August to the same level as in November 2011), Figure 7 shows that the market illiquidity did not increase then as it did in November 2011, as a result of the hike in margins, but rather stayed constant. The large infusion of funding liquidity resulting from the LTRO, confirmed by the low levels of CCBSS after January 2012 shown in Figure 3 Panel (c), loosened the market makers' funding constraints, so that, consistent with Brunnermeier and Pedersen's (2009) prediction, we show empirically that the change in margins in 2012 did not affect the market makers' provision of market liquidity, since their budget constraints were not binding.

The results of the analysis of the structural break in the time series confirm what we posited in Empirical Prediction 3 and allow us to argue that LTRO intervention was very effective in severing the strong connection between credit risk and market liquidity. It is interesting to observe that both the SMP and LTRO interventions generated injections of liquidity into the system by the ECB. However, the magnitudes were completely different (€103 billion in August 2011 versus €489 billion in December 2011) and so were the mechanisms: in the first case, the ECB bought the sovereign bonds directly, while, in the second case, it provided money to reduce the funding liquidity constraints of the banks, which perhaps used some of the released liquidity to purchase sovereign bonds.

## 7. Robustness Checks

### *7.1. Results for Bonds with Different Maturities*

In the body of the paper, we report the results based on the daily bid-ask spread, obtained from MTS data by averaging the quoted bid-ask spread on a bond-day basis, and then averaging them across bonds. The reader may wonder about the robustness of our results with regard to the data composition. One direction for investigating the robustness of the results is that of exploiting the cross-section of bonds. In fact, the liquidity of bonds with different maturities could relate to the CDS spread of corresponding maturity in different ways: Prices of short-term bonds are less sensitive to changes in credit risk and, similarly, their relevance for inventory concerns and VaR considerations should be mitigated by their short time-to-maturity. To characterize the heterogeneity of the effect of credit risk on market liquidity with respect to bond maturity, we split the bonds into different maturity groups and investigate whether i) the effects of credit risk on liquidity are smaller for shorter maturity bonds and ii) our main results hold similarly for all maturity groups.[26]

We consider 11 maturity buckets, based on the classification used by Cassa Compensazione e Garanzia when setting margins. Bonds are grouped together daily if they have the following time-to-maturity: from 0 to 1 month, from 1 to 3 months, from 3 to 9 months, from 9 months to 1.25 years, from 1.25 to 2 years, from 2 to 3.25 years, from 3.25 to 4.75 years, from 4.75 to 7 years, from 7 to 10 years, from 10 to 15 years, and finally from 15 to 30 years. We calculate a liquidity measure per group-day by averaging the bid-ask spreads of the bonds in each group. For each day, we interpolate the CDS spread curve provided by Markit for the Italian sovereign entity and extract, per each maturity bucket, the CDS spread for a contract that has maturity equal to the average between the lower and higher maturity boundaries, e.g., we interpolate the CDS curve, obtain the spread for the 4-year maturity contract and attribute it to the bucket including bonds with 3.25 to 4.75 years to maturity. Due to the lack of a CDS spread estimate for maturities below 3 months, we drop the observations for the first two groups. Table 6 reports the average bid-ask spread and CDS spread, together with the correlations between changes in the two variables, for each maturity group. The illiquidity measure is decreasing in time-to-maturity, with the exception of the 10-year benchmark bonds in group 9.

[26]We thank the anonymous referee for suggesting we pursue this direction in our analysis.

INSERT TABLE 6 HERE

Panel (a) of Figure 10 reports the evolution of the (log-) bid-ask spreads for the nine remaining maturity buckets from July 1, 2010 to December 2012, while Panel (b) reports the term structure of (log-) CDS spread for the nine corresponding maturities. Figure 11 shows the margin evolution for each maturity bucket. Panel (a) of Figure 10 shows that the liquidity series for different maturities comoved to a very large extent, and so did the CDS spreads in Panel (b). Moreover, when the 5-year CDS contract reached 500 bp (6.215 on the y-axis), the term structure became flat, so that all CDS contracts exhibited a spread above 500 bp, regardless of their maturity. That is exactly the time when the clearing houses raised their margins for all maturities, as shown in Figure 11.

INSERT FIGURES 10 AND 11 HERE

We first perform a pooled OLS panel regression corresponding to Equation (5), with the changes in the bid-ask spreads for maturity group $g$ on day $t$, $\Delta BA_{g,t}$, as the dependent variable and changes in CDS contracts for maturity $g$ on day $t$, $\Delta CDS_{g,t}$, as regressors, allowing the coefficients to differ across maturities:

$$\begin{aligned}\Delta BA_{g,t} = \alpha + \sum_{i=1}^{3} \alpha_i \Delta BA_{g,t-i} + \beta_{0g} \Delta CDS_{g,t} \\ + \beta_{1g} \Delta CDS_{g,t-1} + \beta_2 \Delta CCBSS_t \\ + \beta_3 USAVIX_t + \epsilon_t. \end{aligned} \tag{7}$$

The results for Equation (7) are reported in Table 7 Panel A. The table shows that the changes in the bid-ask spread for all the maturities are positively related to changes in the CDS contracts with one lag, so that the results for the average of the bid-ask spread reported above are confirmed. Moreover, the coefficients are increasing with maturities up to the 8th bucket, so that the effects of credit risk on illiquidity are smaller for shorter maturities. However, the parameters for longer maturities are decreasing. At least for bucket 9, we can attribute this effect to the fact that the 10-year bond is the most liquid bond, and has a lower bid-ask spread than the other maturities, while instead the term structure of CDS has a positive slope most of the time.

INSERT TABLE 7 HERE

We investigate whether the result regarding the threshold level of 500 bp for the 5-year CDS contract $CDS_t$ is confirmed, when we allow maturity groups to have different sensitivities to their corresponding CDS spread. We thus estimate Equation (8):

$$
\begin{aligned}
\Delta BA_{g,t} &= \alpha + \sum_{i=1}^{3} \alpha_i \Delta BA_{g,t-i} \\
&+ I\left[CDS_t < \gamma\right]\left(\beta_{0g}\Delta CDS_{g,t} + \beta_{1g}\Delta CDS_{g,t-1}\right) \\
&+ I\left[CDS_t > \gamma\right]\left(\tilde{\beta}_{0g}\Delta CDS_{g,t} + \tilde{\beta}_{1g}\Delta CDS_{g,t-1}\right) \\
&+ \beta_2 \Delta CCBSS_t + \beta_3 USAVIX_t + \epsilon_t. \qquad (8)
\end{aligned}
$$

The test statistic for the presence of the threshold has an estimate of 78.9 and is significant at the 1% level. Figure 12 shows that the results regarding the shift in the relation between the bid-ask and CDS spread, when the CDS spread crosses 500 bp, is confirmed. Therefore, the threshold effect we find for the market-wide bid-ask spread measure is the same for all maturities, as expected given that the term structure of the CDS spread is flat above 500 bp and all margins change significantly when the CDS cross the 500 bp level.[27]

The results of the panel regression for the subsamples in which the 5-year CDS spread is above and below 500 bp are reported in Table 7 Panel B. The results confirm those obtained in the previous sections for the market-wide bid-ask spread measure: below 500 bp, the relation with the lagged changes in the CDS is positive and significant, while when the CDS spread is above the threshold it is the contemporaneous change in credit risk that is significant. In summary, our main results hold when we group bonds in maturity buckets, providing robustness to the main results of the paper.

### *7.2. Results for Different Liquidity Measures*

In the main body of the paper, we conducted the analyses focusing on a single measure for the (il)liquidity of the bond market, the *Bid-Ask Spread*, since it is both the most familiar and most indicative of market conditions. As a final robustness effort, and since there is no consensus in the academic or policy-making literatures regarding the best metrics for assessing the liquidity of an asset, using a shorter data set, we repeat our regressions for the other liquidity measures that have been used extensively in the literature. We establish, in Section Int.8 of the internet

[27] In Section Int.6 of the internet appendix we estimate Equation (8) separately for each maturity group, i.e., we estimate Equation (6) for each maturity bucket, and we show that the same threshold is present in *all* maturity buckets.

appendix, that our results are robust to the choice of liquidity measure.[28]

## 8. Conclusion

The sovereign debt crisis in the Euro-zone has been the most important development in the global economy in the past five years. The crisis stemmed from both liquidity and credit risk concerns in the market and led to a sharp spike in CDS and sovereign bond yield spreads in late 2011, particularly in the Euro-zone periphery. It was only after the launch of the LTRO program and after Mario Draghi's "whatever it takes" comment in July 2012 that the market's alarm diminished: CDS spreads and sovereign bond yields had dropped to sustainable levels in most Euro-zone countries by late 2012. Hence, there is no doubt, *prima facie*, that the ECB programs were a crucial factor in, at least partially, abating the crisis.

These events provide us with an unusual laboratory in which to study how the interaction between credit risk and illiquidity played out, in a more comprehensive framework than has been used in previous studies of corporate or other sovereign bond markets, for the reasons we highlighted in the introduction. We investigate several hypotheses about the main drivers of the *dynamic relation* between credit risk and market liquidity, controlling for global systemic factors and funding liquidity. We conclude that credit risk was one of the main driving forces in determining the liquidity of the bond market, based on a Granger-causality analysis aimed at investigating whether liquidity risk drives credit risk or vice versa. We verify the robustness of our results by testing the same hypothesis in a panel-data setting, and by repeating the analysis using other liquidity measures. In addition to the specific Italian sovereign risk, other global factors such as the USVIX and the funding liquidity measure CCBSS are relevant to the dynamics of market liquidity.

A second important finding is that, prior to ECB intervention, the relation between credit risk and market liquidity was strong, and depended not simply on the changes in credit risk, but also on the *level* of credit risk. Using an econometric methodology that allows us to identify the threshold above which the relation is altered, we estimate that this level corresponds to a CDS spread of 500 bp. This break point of 500 bp is employed in the setting of margin requirements, which fundamentally alters the relation between changes in credit risk and market liquidity. We link our findings to the growing literature on funding liquidity, providing a fitting example of the

[28]The bid-ask spread is correlated by more than 60% with other liquidity variables, making it an appropriate representation of market liquidity. Pelizzon, Subrahmanyam, Tomio, and Uno (2013) study several liquidity proxies in the context of the cross-section of the Italian sovereign bonds.

Brunnermeier and Pedersen (2009) theoretical prediction on the effect of funding liquidity on market liquidity.

We also examine the improvement in market liquidity following the intervention by the ECB. Our analysis indicates that there is a clear structural break following the allotment and settlement of the LTRO on December 21, 2012. Remarkably, the data show that, following the ECB intervention, the improvement in funding liquidity available to the banks strongly attenuated the dynamic relation between credit risk and market liquidity. Although the CDS spread breached the 500 bp mark and margins were raised once again, market liquidity and the relation between credit risk and market liquidity did not change significantly between the regimes below and above this level. Actually, the only variable that still has an impact on market liquidity after the ECB intervention is the global funding liquidity variable, CCBSS. Thus, the ECB intervention not only vastly improved the funding liquidity of the market, but also substantially loosened the link between credit risk and market liquidity.

Our results will be of interest to the Euro-zone national treasuries, helping them to understand the dynamic nature of the relation between credit risk, funding liquidity, and market liquidity, which has strong consequences for the pricing of their issues in the auctions as well as in secondary markets. The ECB may also derive some insights from our analysis that could help them to better understand the impact of the unconventional instruments of new monetary policy. Apart from targeting both funding and market liquidity, the central bank ought also to focus on the market's perceptions of sovereign credit risk.

**Appendix A: The Model**

In this Appendix, we present our theoretical model in detail and make explicit the steps leading to the results reported in Section 3. In our model, the market maker (dealer) has an investment account, holding other securities, and a trading account, holding the bond in which she is making a market. At time $t$, the initial wealth $W_0$ is split between the investment account and the trading account, while the remainder, when positive, is invested in the risk free rate.[29] If the dealer does not trade during the period, at the end of the time interval $t$ to $t+1$, the terminal wealth of her initial portfolio will be

$$\begin{aligned} W_I =& W_0 k \left(1 + r_M\right) + I\left(1 + r\right) \\ & + \left(W_0\left(1 - k\right) - \left(1 - M_I\right) I\right)\left(1 + B_I + r_f\right), \end{aligned}$$

where $k$ is the fraction of her wealth invested in her preferred portfolio with (expected) return $r_M$ ($\overline{r_M}$), $I$ is the true dollar value of current inventory of the stock with (expected) net return $r$ ($\overline{r}$) and variance $\sigma^2$, $r_f$ is the (net) risk free rate over the interval. In this appendix, we make use of indicator functions to simplify the exposition, so that $B_I = bi(W_0\left(1-k\right) - \left(1-M_I\right) I < 0)$, where $i$ is the indicator function, equals $b > 0$ (0), when the cash position $W_0(1-k) - \left(1 - M_I\right) I$ is negative (positive), due to borrowing costs, and $M_I = mi(I < 0)$, due to margins. All returns are assumed to be normally distributed. The borrowing rate is higher than the lending rate and equal to $r_b = r_f + b$.

To better understand the wealth equation, let us consider the following examples (the chosen parameters being $\overline{r_M} = 10\%$, $r_f = 5\%$, $\gamma = 1$, $\sigma_M^2 = 1$, $W_0 = 1000$, $k = \frac{\overline{r_M} - r_f}{\gamma W_0 \sigma_M^2} = \frac{5\%}{1000} = 0.005\%$):

**Case 1:** $I = 500$ is invested in the inventory. The market maker is long in the bond, and so no margins have to be

[29] We do not use the time subscript, $t$, in the following, to avoid clutter in the notation.

taken into consideration; moreover, her total cash position is positive, and hence no borrowing is needed.

$$\begin{aligned} W_I =&1000 \cdot k \cdot (1 + r_M) + \underbrace{500 \cdot (1 + r)}_{\text{Inventory position}} \\ &+ \left( (1 - k)\, 1000 - \underbrace{500}_{\substack{\text{Cash paid} \\ \text{to the customer}}} \right) \cdot (1 + 5\%) \end{aligned}$$

**Case 2:** $I = 1500$ is invested in the inventory. She is long in the bond, so no margins have to be taken into consideration, however her total cash position is negative, so she needs to borrow at the central bank's lending facilities, where she pledges the bond as collateral. There, she can borrow the full amount, but, however, she will have to pay an interest rate $r_b = r_f + b > r_f$

$$\begin{aligned} W_I =&1000 \cdot k \cdot (1 + r_M) + \underbrace{1500 \cdot (1 + r)}_{\text{Inventory position}} \\ &+ \left( (1 - k)\, 1000 - \underbrace{1500}_{\substack{\text{Cash paid} \\ \text{to the customer}}} \right) \cdot \left( \underbrace{1 + b + 5\%}_{\text{Cost of borrowing}} \right) \end{aligned}$$

**Case 3:** $I = -500$, and thus, she is short 500 worth of the bond. She adds to her cash position $(1 - m)\, 500$, because of her short position in the bond (inventory). She borrows the bond at a cost that is a fraction $m$ of the face value.

$$\begin{aligned} W_I =&1000 \cdot k \cdot (1 + r_M) - \underbrace{500 \cdot (1 + r)}_{\text{Inventory position}} \\ &+ \left( \begin{array}{l} (1 - k)\, 1000 + \overbrace{500}^{\substack{\text{Cash received} \\ \text{from the customer}}} \\ - \underbrace{m500}_{\substack{\text{Cost of borrowing the specific} \\ \text{bond, paid upfront}}} \end{array} \right) \cdot (1 + 5\%)\,. \end{aligned}$$

The dealer trades so that her expected utility from maintaining the time-0 portfolio, or trading the dollar quantity

$Q$, are equal, with the post trading wealth being

$$\begin{aligned} W_{I+Q} =& W_0 k\,(1+r_M) + (I+Q)\,(1+r) \\ &+ \left(W_0\,(1-k) - (1-M_{I+Q})\,(I+Q)\right) \\ &\cdot \left(1 + B_{I+Q} + r_f\right) + C\left(1+r_f\right), \end{aligned}$$

where $C$ is the dollar cost of entering into this transaction, and the last term includes the cost of carrying the inventory (profit from borrowing out) in case of a buy (sell) trade. These costs can be positive or negative, depending on whether the trade in the stock raises or lowers the dealer's inventory holding costs, and essentially capture the dealer's exposure cost of holding a non-optimal portfolio. The indicator functions $B_{I+Q} = bi(W_0 - kW_0 - (1-M_{I+Q})\,(I+Q) < 0)$ and $M_{I+Q} = mi(I+Q<0)$ serve the same purpose of $B_I$ and $M_I$. She will trade if

$$E\left[U\left(W_I\right)\right] = E\left[U\left(W_{I+Q}\right)\right]. \tag{9}$$

The market maker is assumed to have a constant absolute risk-aversion utility function $U(x) = -e^{-\gamma x}$. The marginal condition in Equation (9) implies that the relative cost of trading for a quantity $Q$ is

$$\begin{aligned} \frac{C_Q}{Q} =& \frac{\gamma\left(\frac{Q}{2}+I\right)\sigma^2 - (1+\overline{r}) + \gamma k W_0 \sigma_{iM}}{1+r_f} \\ &+ \frac{I}{Q}\left[\begin{array}{c} (1-M_{I+Q})\,\frac{1+B_{I+Q}+r_f}{1+r_f} \\ -(1-M_I)\,\frac{1+B_I+r_f}{1+r_f} \end{array}\right] \\ &+ \frac{W_0}{Q}\,(1-k)\,\frac{B_I - B_{I+Q}}{1+r_f} \\ &+ (1-M_{I+Q})\,\frac{1+B_{I+Q}+r_f}{1+r_f}, \end{aligned}$$

where $\sigma_{iM} = COV\left[r_M, r\right]$ is the covariance between the bond and the market returns. We add the subscript in $C_Q$ to highlight the dependence of $C$ on $Q$. The dealer chooses $k$ optimally so that, when $I = 0$, $\frac{\partial W_I}{\partial k} = 0$, or $k = \frac{\overline{r_M} - r_f}{\gamma W_0 \sigma_M^2}$.[30] The choice of $k$, together with the mean-variance capital asset pricing model (CAPM) portfolio

[30]We follow Stoll (1978) and assume that the market maker chooses the fraction of her wealth invested in the market portfolio before building an inventory.

equilibrium condition $\overline{r} - r_f = \left(\overline{r_M} - r_f\right) \frac{\sigma_{iM}}{\sigma_M^2}$, allows us to rewrite $\frac{C_Q}{Q}$ as

$$
\begin{aligned}
\frac{C_Q}{Q} = & \frac{\gamma\left(\frac{Q}{2} + I\right)\sigma^2 - \left(1 + r_f\right)}{1 + r_f} \\
& + \frac{I}{Q}\left[\begin{array}{c} (1 - M_{I+Q})\,\frac{1+B_{I+Q}+r_f}{1+r_f} \\ -(1 - M_I)\,\frac{1+B_I+r_f}{1+r_f} \end{array}\right] \\
& + \frac{W_0}{Q}(1 - k)\,\frac{B_I - B_{I+Q}}{1 + r_f} \\
& + (1 - M_{I+Q})\,\frac{1 + B_{I+Q} + r_f}{1 + r_f}.
\end{aligned}
$$

The relative bid ask spread for a dollar quantity $|Q| > 0$ is the summation (since they are signed quantities) between buying a stock from the market maker at the ask price $-|Q| + C_{-|Q|}$ (i.e., the price at which the market maker sells $|Q|$), and selling it at the bid price $|Q| + C_{+|Q|}$, so that the relative bid-ask spread becomes

$$
\begin{aligned}
& \frac{-|Q| + C_{-|Q|} + |Q| + C_{+|Q|}}{|Q|} \\
& = \frac{C_{+|Q|} + C_{-|Q|}}{|Q|} = \frac{C_{+|Q|}}{|Q|} - \frac{C_{-|Q|}}{-|Q|}.
\end{aligned}
$$

We restrict our attention to the case when the market maker incurs costs both when she accumulates a long and a short position, i.e., $B_{I+|Q|} = b$, $B_{I-|Q|} = B_I = 0$ and $M_{I-|Q|} = m$, $M_{I+|Q|} = M_I = 0$. Moreover, we assume that $I = 0$, and the two components of the relative bid-ask spread for a quantity $|Q|$ become

$$
\begin{aligned}
\frac{C_{+|Q|}}{|Q|} &= \frac{\left(\begin{array}{c} \gamma\frac{|Q|}{2}\sigma^2 - \left(1 + r_f\right) \\ -\frac{W_0}{|Q|}(1 - k)\,b + 1 + b + r_f \end{array}\right)}{1 + r_f} \\
&= \frac{\gamma\frac{|Q|}{2}\sigma^2}{1 + r_f} + \frac{b}{1 + r_f}\,\frac{|Q| - W_0(1 - k)}{|Q|} \\
\frac{C_{-|Q|}}{-|Q|} &= \frac{\left(\begin{array}{c} \gamma\frac{-|Q|}{2}\sigma^2 - \left(1 + r_f\right) \\ +(1 - m)\left(1 + r_f\right) \end{array}\right)}{1 + r_f} = -\frac{\gamma\frac{|Q|}{2}\sigma^2}{1 + r_f} - m.
\end{aligned}
$$

Finally, the absolute bid-ask, calculated as the relative bid-ask spread for purchasing a quantity $Q = p_0$ multiplied by the price of the bond $p_0$, is

$$BA = \frac{\gamma p_0^2 \sigma^2}{1 + r_f} + m p_0 + b \frac{p_0 - W_0 (1 - k)}{1 + r_f}. \tag{10}$$

*The Option*

Since the default of a sovereign is, at least partly, a political decision, we take the approach of looking at the underlying process as merely that, rather than an endogenous choice of the "equity holders". We can think of a CDS contract as sort of an event-triggered put option written on the sovereign bond.[31] Brennan (1979) and Stapleton and Subrahmanyam (1984) show that a sufficient condition for a risk-neutral valuation of a contingent claim when the price of the underlying asset is assumed to be normally distributed is that the utility function of the representative investor be exponential (Theorem 6). Therefore, the price of a put option with strike $k$ at time $t$, if the price of the underlying $p$ is normally distributed $N(\overline{p}, \sigma_p^2)$, with $p = p_0(1 + r)$, $\overline{p} = p_0(1 + \overline{r})$, and $\sigma_p^2 = p_0^2 \sigma^2$, would be

$$\begin{aligned} CDS = & \frac{1}{1 + r_f} \int_{-\infty}^{+x} (x - p) \frac{1}{\sigma_p \sqrt{2\pi}} \\ & \cdot \exp\left( -\frac{1}{2\sigma_p^2} \left( p - \left(1 + r_f\right)(1 - m) p_0 \right)^2 \right) dp \end{aligned}$$

and, with a change of variable to $z = \frac{p - (1 + r_f)(1 - m) p_0}{\sigma_p}$

$$\begin{aligned} CDS = & \left( \frac{x - \left(1 + r_f\right)(1 - m) p_0}{1 + r_f} \right) \\ & \cdot N\left( \frac{x - \left(1 + r_f\right)(1 - m) p_0}{\sigma_p} \right) \\ & + \frac{\sigma_p}{1 + r_f} n\left( \frac{x - \left(1 + r_f\right)(1 - m) p_0}{\sigma_p} \right), \end{aligned}$$

where $N$ and $n$ are the cumulative and standard normal distribution functions, respectively.

If we consider a margin-adjusted at-the-money put option, i.e., one such that $x = \left(1 + r_f\right)(1 - m) p_0$, the CDS price formula simplifies to $CDS = \frac{\sigma p_0}{1 + r_f} n(0)$, so that the market maker extracts the volatility from the CDS market

[31] This is not literally correct, given that the CDS is triggered by an event, rather than by exercise at expiration, but is useful here as a simplification, to avoid the need to model the default intensity process

according to a simple linear approach:

$$\sigma = \left(1 + r_f\right) \frac{CDS}{p_0 n(0)}. \tag{11}$$

*Empirical Predictions*

Re-writing the absolute bid ask spread in Equation (10) as a function of the CDS price, and plugging Equation (11) into 10, we obtain the relation between the depending variable of interest, the bid-ask spread, and its determinants, the CDS spread, and the policy quantities set by clearing houses and the central bank:

$$BA(CDS) = \delta CDS^2 + m(b, CDS)\, p_0 + b\eta, \tag{12}$$

where $\frac{\gamma(1+r_f)}{n(0)^2} = \delta > 0$ and $\frac{p_0 - W_0(1-k)}{1+r_f} = \eta > 0$ and, where, realistically, we allow the margin setting decision by the clearing house to depend on both the level of the CDS spread and the borrowing rate set by the central bank.[32,33] From Equation (12), we obtain the following empirical predictions, which we discuss in Section 3.1:

**Empirical Prediction 1.** *The illiquidity of the bond market increases with credit risk.*

**Empirical Prediction 2.** *The dynamic relation between credit risk and market illiquidity shifts conditional on the level of the CDS spread.*

**Empirical Prediction 3.** *The monetary policy interventions of the central bank affect the dynamic relation between credit risk and market liquidity.*

*An Implicit Formulation*

Similar implications can be derived from Equation (10) even without the assumption that the market maker uses the simple linear approach in Equation (11). Indicating the relation between the CDS price and the return volatility that is extracted from it by $\sigma^2(CDS)$, Equation (10) becomes

$$BA = \frac{\gamma p_0^2}{1 + r_f} \sigma^2(CDS) + m(b, CDS)\, p_0 + b\eta$$

and the same empirical predictions follow.

[32]We refer to CDS price, in the theoretical section, and CDS spread, in the rest of the paper, interchangeably.

[33]The second inequality follows from the assumption that the market maker borrows the funds necessary to buy a bond by pledging the latter at the central bank. That is, $B_{I+|Q|} = b$, meaning that $(1-k)W_0 - (1 - M_{I+|Q|})(I + |Q|)) < 0$, which corresponds to the inequality, when we assume that $I = M_{I+|Q|} = 0$, and that the trade occurs for a quantity $|Q| = p_0$.

**Appendix B: Methodological Appendix**

*Threshold Analysis*

In empirical settings, a regression such as the OLS specification $y_i = \beta' x_i + e_i$, where $y_i$ is the dependent variable that is regressed on the independent variable $x_i$, is often repeated for subsamples, either as a robustness check or to verify whether the same relation applies to appropriately grouped observations. The sample split is often conducted in an exogenous fashion, thus dividing the data according to the distribution of a key variable, such as size and book-to-market quantile portfolios in a Fama-French (1993) setting. Hansen (1996, 2000) develops the asymptotic approximation of the distribution of the estimated threshold value $\hat{\gamma}$, when the sample split, based on the values of an independent variable $q_i$, can be rewritten as

$$Y = X\theta + X_\gamma \delta + e \qquad \text{where } X_\gamma = XI(q \leq \gamma)$$

or $y_i = \theta' x_i + \delta I(q_i \leq \gamma)x_i + e_i$, where $I(q_i \leq \gamma)$ equals 1 if $q_i \leq \gamma$, and 0 otherwise. He shows that, under a set of regularity conditions, which exclude time-trending and integrated variables, the model can be estimated by least squares, minimizing $SSR_n(\theta, \delta, \gamma) = (Y - X\theta - X_\gamma\delta)'(Y - X\theta - X_\gamma\delta)$.[34] Concentrating out all parameters but $\gamma$, i.e. expressing them as functions of $\gamma$, yields $S_n(\gamma) = SSR_n(\hat{\theta}(\gamma), \hat{\delta}(\gamma), \gamma) = Y'Y - Y'X^*_\gamma('X^*_\gamma{}'X^*_\gamma)^{-1}X^*_\gamma{}'Y$ with $X^*_\gamma = \left[X\ X_\gamma\right]$. The parameters $\theta$ and $\delta$ are formulated as functions of $\gamma$, and the sum of squared residuals depends exclusively on the observed variables and on $\gamma$. Thus, the value of $\gamma$ that minimizes $S_n(\gamma)$ is its least squares estimator $\hat{\gamma}$, and the estimators of the remaining parameters $\hat{\theta}(\hat{\gamma})$ and $\hat{\delta}(\hat{\gamma})$ can be calculated.

When there are $N$ observations, there are at most $N$ values of the threshold variable $q_i$, or equivalently $N$ values that the $SSR(\gamma)$ (step-)function can take. After re-ordering the values $q_i$ in $(q_{(1)}, q_{(2)}, ...q_{(N)})$, such that $q_{(j)} \leq q_{(j+1)}$, the method is implemented by

1. estimating by OLS $y_i = \theta_2' x_i + \delta I(q \leq q_{(j)})x_i + e_i$ (or equivalently, when all parameters are allowed to depend on the threshold, estimating separately $y_i = \theta_1' x_i + e_{1i}$ where $q_i \leq q_{(j)}$ and $y_i = \theta_2' x_i + e_{2i}$ where $q_i > q_{(j)}$),
2. calculating the sum of squared residuals, $SSR(q_{(j)}) = \sum e_i$ (or $= \sum e_{1i} + \sum e_{2i}$),
3. repeating 1 and 2 with $q_{(j+1)}$,

[34] A theory for the latter case was developed in Caner and Hansen (2001).

4. finding the least squares estimate of $\gamma$ as $\hat{\gamma} = \arg\min_{q_{(j)}} S(q_{(j)})$, and

5. repeating the estimation of the equations on the subsamples defined by the $\hat{\gamma}$ threshold, calculating heteroskedasticity-consistent standard errors for the parameters.

As suggested by Hansen (1999), we allow each equation to contain at least 20% of the observations, and, to minimize computing time, we search only through 0.5%-quantiles. Although Hansen (1999) presents an extension of the procedure to several thresholds, we focus in this paper on a single sample split.

To test the presence of the threshold, thus testing whether $\theta_1 = \theta_2$, the usual tests cannot be used, since $\gamma$ is not identified under the null hypothesis. This is known as the "Davies' Problem", as analyzed by Davies (1977, 1987). Hansen (1996) provides a test whose asymptotic properties can be approximated by bootstrap techniques.

To provide confidence intervals for the threshold estimate $\hat{\gamma}$, Hansen (2000) argues that no-rejection regions should be used. To test $\gamma = \gamma_0$, the likelihood ratio test can be used such that $LR(\gamma) = (SSR(\gamma) - SSR(\hat{\gamma}))/\hat{\sigma}^2$, where $\hat{\sigma}^2 = SSR(\hat{\gamma})/N$ is the estimated error variance, will be rejected if $\hat{\gamma}$ is sufficiently far from $\gamma$, i.e., the test statistic is large enough. In its homoskedastic version, the test has a non-standard pivotal distribution, such that the test is rejected at an $\alpha$-confidence level if $LR(\gamma) > -2\ln(1 - \sqrt{\alpha})$. In this paper, we choose $\alpha = 0.95$, consistent with Hansen (2000); thus, the null hypothesis is considered rejected if $LR(\gamma) >= -2\ln(1 - \sqrt{0.95}) = 7.35$. This level is plotted as a horizontal line in the plots of the test. The confidence interval for the threshold will be $[\gamma_L, \gamma_U]$, such that $LR(\gamma|\gamma < \gamma_U) > 7.35$, and $LR(\gamma|\gamma > \gamma_U) > 7.35$, or, graphically, the portion of the $x$-axis in which the plot of the test is below the 7.35 horizontal line.

*Structural Break Tests*

The Chow test is a standard break point analysis used widely in the economics literature. Based on two nested regressions, it follows an $f_{k,T-2k}$-distribution and its statistic is

$$F = \frac{(SSR_0 - SSR_1)/k}{SSR_1/(T - 2k)}.$$

where $SSR_0$ and $SSR_1$ are the sum of squared residuals of the restricted regression, $y_t = x_t'\beta + \epsilon_t$ (with $t = 1, ..., T$), and the unrestricted regression, $y_t = x_t'\beta + g_t x_t'\gamma + \epsilon_t$ , respectively. In the unrestricted regressions, the observations following the break point $t^*$, selected by the dummy variable $g_t$ (such that $g_t = 1$ if $t < t^* \leq T$ and 0 otherwise), are allowed to depend on $x_t$ through the composite parameters $\beta + \gamma$, while the previous observations depend on $x_t$ through $\beta$ only. The restriction $\gamma = 0$ thus imposes the condition that all $y_t$ depend on $x_t$ in a homogeneous fashion.[35]

A drawback of the Chow test is that the breakpoint has to be specified exogeneously. The Chow test has a null hypothesis, which is that the parameters after a specific date are equal to those that generated the data before the break date. The alternative hypothesis is that the two sets of parameters are indeed different. However, a test statistic can be calculated from the statistics resulting from the Chow test, the $F$s, to test whether a structural break took place at an *unknown* date. After the $F$-statistics have been computed for a subset of dates, e.g., all the dates in the sample except for the first and last i%, several test statistics can be calculated from them.

Andrews (1993) and Andrews and Ploberger (1994) show that the supremum and the average, respectively, of the $F$-statistics converge to a pivotal non-standard distribution, depending on the number of parameters tested and the relative number of dates tested. The test statistics that we calculate to test for a structural break at an unknown date are therefore

$$supF = \sup_t F_t$$
$$aveF = \frac{\sum_t F_t}{T},$$

where the $F_t$ are found using the Chow test estimation. We then compare the $supF$ and $aveF$ test statistics with the corresponding confidence levels, that can be found in Andrews (2003), which rectified those tabulated in Andrews (1993), and Andrews and Ploberger (1994).

[35] We exclude the first and last 10% of the observations, in order to estimate meaningful regressions.

## Tables

Table 1: ***Maturity* and *Coupon Rate* by Maturity Group and Bond Type.** This table presents the distribution of the bonds in the sample in terms of *Maturity* and *Coupon Rate*, by maturity group (Panel A) and bond type (Panel B). Maturity groups were determined by the time distance between bond maturities and the closest whole year. Our data set, obtained from the Mercato dei Titoli di Stato (MTS), consists of transactions, quotes, and orders for all 189 fixed-rate and floating Italian sovereign bonds (Buoni Ordinari del Tesoro (BOT) or Treasury bills, Certificato del Tesoro Zero-coupon (CTZ) or zero-coupon bonds, Certificati di Credito del Tesoro (CCT) or floating notes, and Buoni del Tesoro Poliennali (BTP) or fixed-income Treasury bonds) from July 1, 2010 to December 31, 2012.

Panel A

| Maturity Group | # Bonds | Coupon Rate | Maturity | MinMaturity | MaxMaturity |
|---|---|---|---|---|---|
| 0.25 | 11 | [a] | 0.26 | 0.21 | 0.27 |
| 0.50 | 38 | [a] | 0.50 | 0.36 | 0.52 |
| 1.00 | 44 | [a] | 1.00 | 0.81 | 1.02 |
| 2.00 | 13 | [b] | 2.02 | 2.01 | 2.09 |
| 3.00 | 14 | 3.38 | 2.98 | 2.93 | 3.02 |
| 5.00 | 16 | 3.86 | 5.02 | 4.92 | 5.25 |
| 6.00 | 15 | [c] | 6.71 | 5.21 | 7.01 |
| 10.00 | 21 | 4.54 | 10.44 | 10.10 | 10.52 |
| 15.00 | 7 | 4.59 | 15.71 | 15.44 | 16.00 |
| 30.00 | 10 | 5.88 | 30.88 | 30.00 | 31.79 |

Panel B

| Bond Type | N | Coupon Rate | Maturity | MinMaturity | MaxMaturity |
|---|---|---|---|---|---|
| BOT | 93 | ZCB | 0.71 | 0.21 | 1.02 |
| BTP | 68 | 4.34 | 11.12 | 2.93 | 31.80 |
| CCT | 15 | Floating | 6.71 | 5.21 | 7.01 |
| CTZ | 13 | ZCB | 2.02 | 2.00 | 2.09 |

[a] All bonds in this group are BOT, Buoni Ordinari del Tesoro (Treasury bills)
[b] All bonds in this group are CTZ, Certificati del Tesoro Zero-coupon (zero-coupon bonds)
[c] All bonds in this group are CCT, Certificati di Credito del Tesoro (floating bonds)

Table 2: **Time-series Descriptive Statistics of the Variables.** This table shows the time-series and cross-sectional distribution of various variables defined in Section 4.1, and their correlations. The sample consists of the quotes and trades from 641 days in our sample for bond market data and end-of-day quotes for the other measures. *Quoted Bonds* is the number of bonds actually quoted on each day, *Trades* is the total number of trades on the day, and *Volume* is the daily amount traded in € billion on the whole market. The liquidity measure *Bid-Ask Spread* is the difference between the best bid and the best ask. The global systemic variables are the spread between three-month Euribor and three-month German sovereign yield, the USVIX, and the Cross-Currency Basis Swap Spread CCBSS. Our bond-based data, obtained from the Mercato dei Titoli di Stato (MTS), consist of transactions, quotes, and orders for all 189 fixed-rate and floating Italian sovereign bonds (Buoni Ordinari del Tesoro (BOT) or Treasury bills, Certificato del Tesoro Zero-coupon (CTZ) or zero-coupon bonds, Certificati di Credito del Tesoro (CCT) or floating notes, and Buoni del Tesoro Poliennali (BTP) or fixed-income Treasury bonds) from July 1, 2010 to December 31, 2012. All other data were obtained from Bloomberg.

| | Time Series | | | | | Unit Root Test | |
|---|---|---|---|---|---|---|---|
| **Panel A: Market Measures** | | | | | | | |
| Variable | Mean | STD | 5th Pct | Median | 95th Pct | Level | Difference |
| Quoted Bonds | 88.583 | 2.430 | 85.000 | 88.000 | 93v | | |
| Trades | 352.158 | 149.394 | 145.000 | 331.000 | 614.000 | | |
| Volume | 2.874 | 1.465 | 0.951 | 2.555 | 5.647 | | |
| **Panel B: System Variables** | | | | | | | |
| Bid-Ask Spread | 0.389 | 0.340 | 0.128 | 0.298 | 1.092 | -8.200*** | -32.597*** |
| Italian CDS | 320.748 | 137.834 | 149.356 | 302.026 | 540.147 | -1.469 | -19.922*** |
| USVIX | 21.212 | 6.302 | 15.070 | 18.970 | 34.770 | -3.951*** | -26.790*** |
| CCBSS | 44.003 | 18.915 | 21.100 | 39.900 | 79.400 | -1.613 | -25.969*** |
| Euribor-DeTBill | 0.729 | 0.357 | 0.264 | 0.629 | 1.474 | -1.750 | -31.843*** |
| **Panel C: Correlations** | | | | | | | |
| Differences\Levels | Bid-Ask Spread | Italian CDS | USVIX | CCBSS | Euribor -DeTBill | | |
| Bid-Ask Spread | 1 | 0.628 | 0.440 | 0.659 | 0.676 | | |
| Italian CDS | 0.224 | 1 | 0.318 | 0.788 | 0.589 | | |
| USVIX | 0.151 | 0.334 | 1 | 0.511 | 0.660 | | |
| CCBSS | 0.182 | 0.367 | 0.233 | 1 | 0.842 | | |
| Euribor-DeTBill | 0.049 | 0.088 | 0.050 | 0.054 | 1 | | |

Table 3: **Results for the Granger-Causality Analysis of the Italian CDS Spread and Bid-Ask Spread.** This table presents the results for the regressions of the day $t$ changes in *Bid-Ask Spread*, $\Delta \mathrm{BA}_t$, and Italian CDS spread $\Delta CDS_t$, on the lagged terms of both variables and on contemporaneous macro variable changes, in a VARX(3,0) setting as shown in Equation (4). The data have a daily frequency. The significance refers to heteroskedasticity-robust $t$-tests. Heteroskedasticity-robust $F$-test statistics and their significance are reported for the null hypothesis of $\Delta BA_t = \Delta BA_{t-1} ... = 0$ ( $BA \xrightarrow{GC} CDS$ ), and $\Delta CDS_t = \Delta CDS_{t-1} ... = 0$ ($CDS \xrightarrow{GC} BA$) respectively. We also report the contemporaneous correlation in the model residuals. Our data set consists of 641 days of trading in Italian sovereign bonds, from July 1, 2010 to December 31, 2012, and was obtained from the MTS (Mercato dei Titoli di Stato) Global Market bond trading system. The CDS spread refers to a USD-denominated, five-year CDS spread. The CDS spread and the macro variables were obtained from Bloomberg.

| Variable | $\Delta BA_t$ | $\Delta CDS_t$ |
|---|---|---|
| $\Delta BA_{t-1}$ | -0.357*** | -0.011 |
| $\Delta CDS_{t-1}$ | 0.917*** | 0.212*** |
| $\Delta BA_{t-2}$ | -0.224*** | -0.007 |
| $\Delta CDS_{t-2}$ | -0.069 | -0.091* |
| $\Delta BA_{t-3}$ | -0.174*** | -0.004 |
| $\Delta CDS_{t-3}$ | 0.117 | 0.024 |
| $\Delta EuriborDeTBill_t$ | 0.027 | 0.035 |
| $\Delta CCBSS_t$ | 0.545*** | 0.213*** |
| $\Delta USVIX_t$ | 0.334** | 0.154*** |
| Intercept | -0.001 | 0.001 |
| Adj $R^2$ | 0.180 | 0.236 |
| Granger-Causality Tests | | |
| $BA \xrightarrow{GC} CDS$ | . | 0.476 |
| $CDS \xrightarrow{GC} BA$ | 6.007*** | . |
| Residuals Correlation | | |
| $\Delta BA_t$ | 1.000 | 0.107 |
| $\Delta CDS_t$ | 0.107 | 1.000 |

* Significant at a 10% level. ** Significant at a 5% level. *** Significant at a 1% level.

Table 4: **Results for the Regression of the Bid-Ask Spread on the CDS Spread and Macro Variables.** This table presents the results for the regression of the change in the *Bid-Ask Spread* (the change in the quoted bid-ask spread) on day $t$, $\Delta BA_t$, on its lagged terms, and the change in the CDS spread on day $t$, $\Delta CDS_t$, and its lagged terms and on macro variables, using daily data. The regressions are presented in Equations (5) and (6), for Panels A and B, respectively. Parameters multiplying the identity operator $[CDS \leq (>)500]$ are reported under the $[CDS \leq (>)500]$ column. The statistical significance refers to heteroskedasticity-robust $t$-tests. The Test column reports the heteroskedasticity-robust test for the two parameters above and below the threshold being equal and distributed as chi-square (1). Our data set consists of 641 days of trading in Italian sovereign bonds, from July 1, 2010 to December 31, 2012, and was obtained from the Mercato dei Titoli di Stato (MTS) Global Market bond trading system. The CDS spread refers to a USD-denominated, five-year CDS spread and macro variables were obtained from Bloomberg.

| Variable | Panel A | Panel B | | |
|---|---|---|---|---|
| | Whole Sample | I[CDS≤500] | I[CDS>500] | Test |
| $\Delta CDS_t$ | 0.541 ** | 0.319 | 2.845*** | 11.33*** |
| $\Delta CDS_{t-1}$ | 0.794 *** | 0.983*** | -0.854* | 10.75*** |
| $\Delta BA_{t-1}$ | -0.352 *** | -0.332*** | | |
| $\Delta BA_{t-2}$ | -0.216 *** | -0.199*** | | |
| $\Delta BA_{t-3}$ | -0.167 *** | -0.164*** | | |
| $\Delta CCBSS_t$ | 0.429 *** | 0.402*** | | |
| $\Delta USVIX_t$ | 0.251 * | 0.208* | | |
| Intercept | -0.002 | -0.002 | | |
| Adj $R^2$ | 0.191 | 0.219 | | |
| N | 637 | 637 | | |

[*] Significant at a 10% level. [**] Significant at a 5% level. [***] Significant at a 1% level.

Table 5: **Results for the Regression of the Bid-Ask Spread on the CDS Spread and Macro Variables for Subsamples Based on the Structural Break.** This table presents the results for the regression of the change in the *Bid-Ask Spread* (the change in the quoted bid-ask spread) on day $t$, $\Delta\text{BA}_t$, on its lagged terms, and the change in the CDS spread on day $t$, $\Delta\text{CDS}_t$, and its lagged terms, using daily data for the *Bid-Ask Spread* and the CDS spread. The regressions are presented for Equations (6) and (5) in Panels A and B respectively. Parameters multiplying the identity operator $[CDS \leq (>)500]$ are reported under the $[CDS \leq (>)500]$ column. The statistical significance refers to heteroskedasticity-robust $t$-tests. The Test column reports the heteroskedasticity-robust test results for the two parameters above and below the threshold being equal and distributed as chi-square (1). Panel A (B) is based on the pre-(post-)structural-break sample. Our data set consists of 641 days of trading in Italian sovereign bonds, from July 1, 2010 to December 31, 2012, and was obtained from the Mercato dei Titoli di Stato (MTS) Global Market bond trading system. The CDS spread refers to a USD-denominated, five-year CDS spread and macro variables were obtained from Bloomberg.

| Variable | Panel A: 2011 | | | Panel B: 2012 |
|---|---|---|---|---|
| | I[CDS≤500] | I[CDS>500] | Test | |
| $\Delta\text{CDS}_t$ | 0.493 | 3.877*** | 16.21*** | 0.064 |
| $\Delta\text{CDS}_{t-1}$ | 1.028*** | -1.491** | 11.77*** | 0.566** |
| $\Delta\text{BA}_{t-1}$ | -0.261*** | | | -0.501*** |
| $\Delta\text{BA}_{t-2}$ | -0.183*** | | | -0.295*** |
| $\Delta\text{BA}_{t-3}$ | -0.162*** | | | -0.188*** |
| $\Delta\text{CCBSS}_t$ | 0.310* | | | 0.858*** |
| $\Delta\text{USVIX}_t$ | 0.320** | | | -0.105 |
| Intercept | 0.002 | | | -0.006 |
| Adj $R^2$ | 0.233 | | | 0.237 |
| N | 377 | | | 260 |

* Significant at a 10% level. ** Significant at a 5% level. *** Significant at a 1% level.

Table 6: **Descriptive Statics for Bonds Grouped by Maturity.** This table presents the time-series average of the bid-ask spread for bonds grouped by their time to maturity, the time-series average of the CDS spread with matching maturity, and the correlation between daily changes in the bid-ask and CDS spreads (contemporaneous, and with a lag). Our data set consists of 641 days of trading in Italian sovereign bonds, from July 1, 2010 to December 31, 2012, and was obtained from the Mercato dei Titoli di Stato (MTS) Global Market bond trading system. The CDS spread refers to a USD-denominated CDS spread with maturity matching the average maturity of the bond group and was obtained from the term structure of the CDS spread provided by Markit.

| Maturity Group | Bid-Ask Spread | CDS Spread | Contemporaneous Correlation | Lagged Correlation |
|---|---|---|---|---|
| 03:3-9m | 0.142 | 201.883 | 0.108 | 0.090 |
| 04:0.75-1.25y | 0.198 | 230.540 | 0.136 | 0.137 |
| 05:1.25-2y | 0.282 | 255.422 | 0.148 | 0.163 |
| 06:2-3.25y | 0.337 | 286.799 | 0.214 | 0.150 |
| 07:3.25-4.75y | 0.469 | 308.557 | 0.207 | 0.155 |
| 08:4.75-7y | 0.519 | 317.945 | 0.196 | 0.167 |
| 09:7-10y | 0.495 | 317.701 | 0.130 | 0.142 |
| 10:10-15y | 0.757 | 315.404 | 0.121 | 0.100 |
| 11:15-30y | 0.958 | 311.923 | 0.073 | 0.093 |

Table 7: **Results for the Regression of the Bid-Ask Spread on the CDS spread and Macro Variables with Maturity-Specific Coefficients.** This table presents the results for the regression of the change in the *Bid-Ask Spread* for maturity group $g$ on day $t$, $\Delta BA_{g,t}$, on its lagged terms, and the change in the CDS spread with maturity matching that of group $g$ on day $t$, $\Delta CDS_{g,t}$, and its lagged term and on macro variables, using daily data. The regressions presented in Equations (7) and (8) are used for Panels A and C, and for Panel B, respectively. Parameters multiplying the identity operator $[CDS \leq (>)500]$ are reported under the $[CDS \leq (>)500]$ column. The statistical significance refers to heteroskedasticity-robust $t$-tests. The Test column reports the heteroskedasticity-robust test for the two parameters above and below the threshold being equal and distributed as chi-square (1). Our data set consists of 641 days of trading in Italian sovereign bonds, from July 1, 2010 to December 31, 2012, and was obtained from the Mercato dei Titoli di Stato (MTS) Global Market bond trading system. The CDS spread refers to a USD-denominated CDS spread with maturity matching the average maturity of the bond group and was obtained from the term structure of the CDS spread provided by Markit.

| Variable | Panel A | Panel B:2011 | | | Panel C:2012 |
|---|---|---|---|---|---|
| | Whole Sample | I[CDS≤500] | I[CDS>500] | Test | |
| $\Delta CDS_{3,t}$ | 0.247 | 0.397* | 3.776*** | 17.62*** | -0.043 |
| $\Delta CDS_{4,t}$ | 0.301 | 0.403* | 3.751*** | 10.29*** | -0.077 |
| $\Delta CDS_{5,t}$ | 0.196 | 0.360 | 4.085*** | 21.32*** | -0.443* |
| $\Delta CDS_{6,t}$ | 0.372* | 0.422 | 2.763*** | 7.00*** | -0.052 |
| $\Delta CDS_{7,t}$ | 0.356 | 0.501 | 2.344*** | 3.85** | -0.292 |
| $\Delta CDS_{8,t}$ | 0.275 | 0.288 | 2.784*** | 6.73*** | -0.339 |
| $\Delta CDS_{9,t}$ | -0.014 | -0.146 | 2.757*** | 9.51*** | -0.421 |
| $\Delta CDS_{10,t}$ | 0.091 | 0.131 | 2.827*** | 5.57** | -0.630* |
| $\Delta CDS_{11,t}$ | -0.106 | -0.147 | 2.374** | 4.61** | -0.624 |
| $\Delta CDS_{3,t-1}$ | 0.437*** | 0.745*** | -0.227 | 2.13 | 0.032 |
| $\Delta CDS_{4,t-1}$ | 0.75*** | 1.099*** | 0.349 | 0.83 | 0.169 |
| $\Delta CDS_{5,t-1}$ | 0.941*** | 1.144*** | -0.277 | 3.61* | 0.594*** |
| $\Delta CDS_{6,t-1}$ | 0.944*** | 1.071*** | 0.078 | 2.53 | 0.745** |
| $\Delta CDS_{7,t-1}$ | 1.066*** | 1.178*** | 0.069 | 2.45 | 0.939*** |
| $\Delta CDS_{8,t-1}$ | 1.197*** | 1.521*** | -0.004 | 4.6** | 0.711** |
| $\Delta CDS_{9,t-1}$ | 0.954*** | 1.225*** | 0.291 | 1.85 | 0.395 |
| $\Delta CDS_{10,t-1}$ | 0.672*** | 1.092*** | -1.554* | 9.06*** | 0.351 |
| $\Delta CDS_{11,t-1}$ | 0.624** | 0.932** | -1.221 | 6.21** | 0.486 |
| $\Delta BA_{g,t-1}$ | -0.429*** | -0.400*** | | | -0.490*** |
| $\Delta BA_{g,t-2}$ | -0.25*** | -0.234*** | | | -0.286*** |
| $\Delta BA_{g,t-3}$ | -0.159*** | -0.168*** | | | -0.140*** |
| $\Delta CCBSS_t$ | 0.652*** | 0.515*** | | | 1.026*** |
| $\Delta USVIX_t$ | 0.315*** | 0.302*** | | | 0.142 |
| Intercept | 0.001 | 0.003 | | | -0.004 |
| Adj $R^2$ | 0.190 | 0.199 | | | 0.209 |
| N | 7007 | 4147 | | | 2860 |

$^{*}$ Significant at a 10% level. $^{**}$ Significant at a 5% level. $^{***}$ Significant at a 1% level.

## Figures

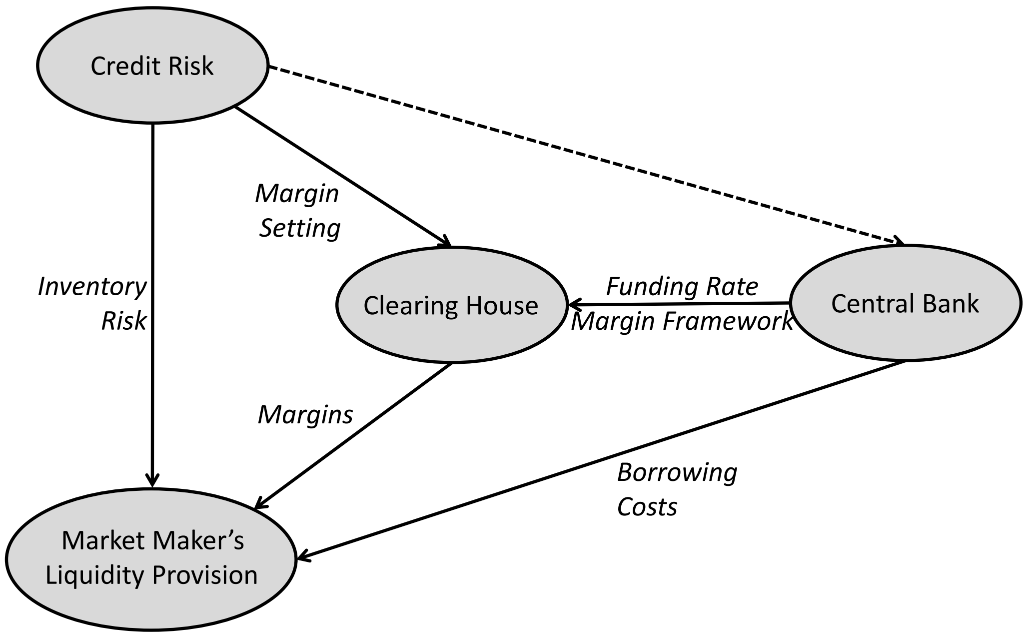



Figure 1: **The Dynamics of the Theoretical Model.** This figure shows the channels through which the players in the model are affected by credit risk and by each other.

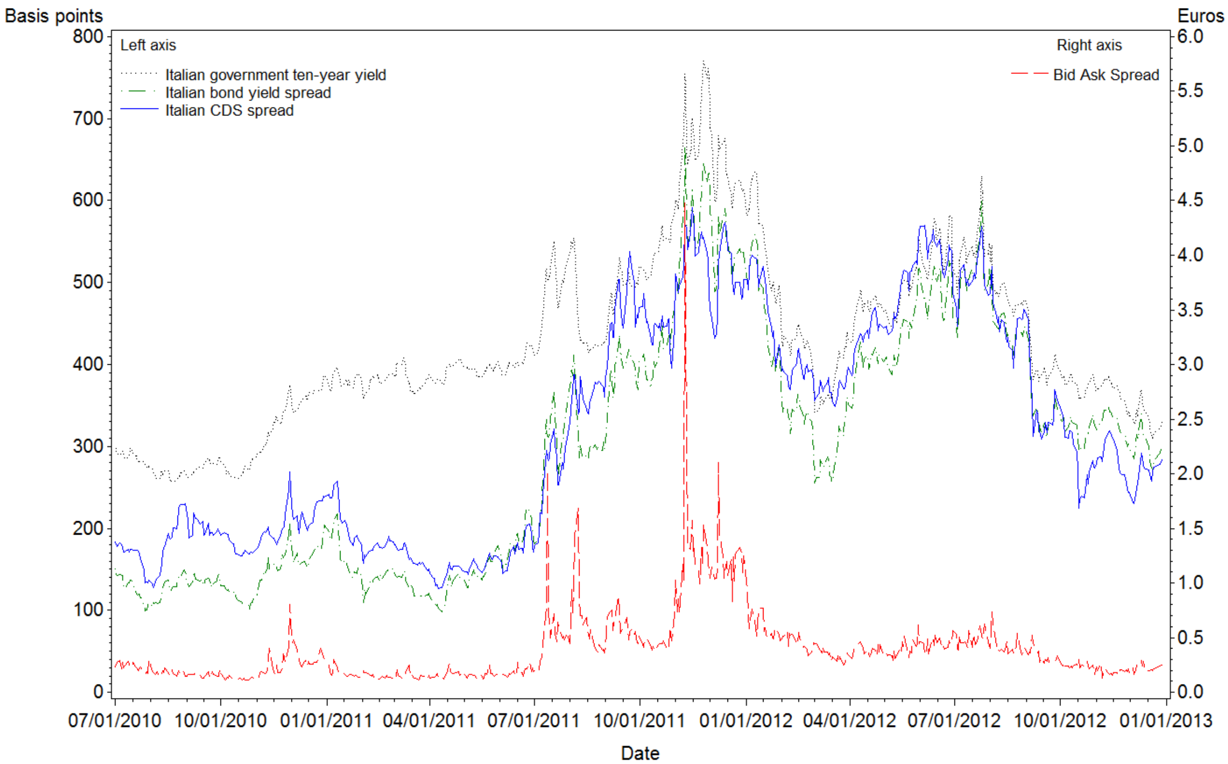
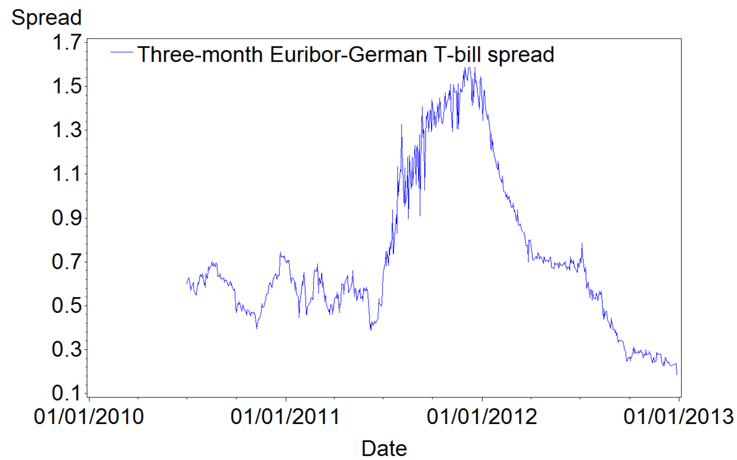
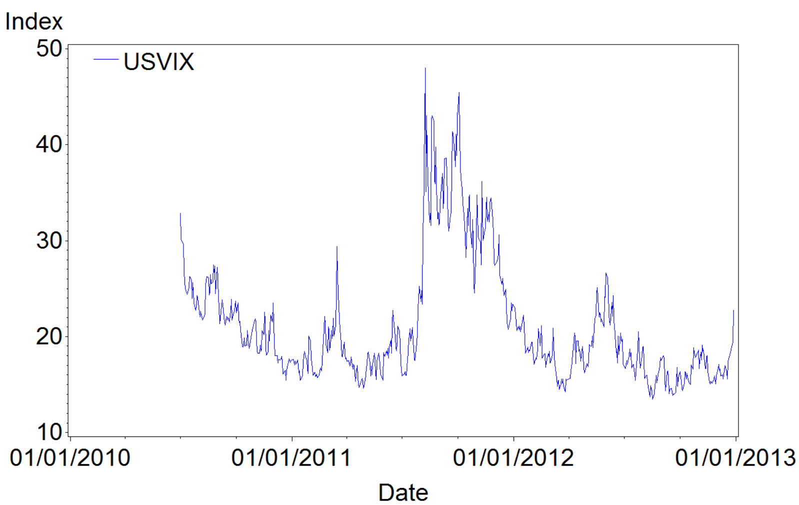


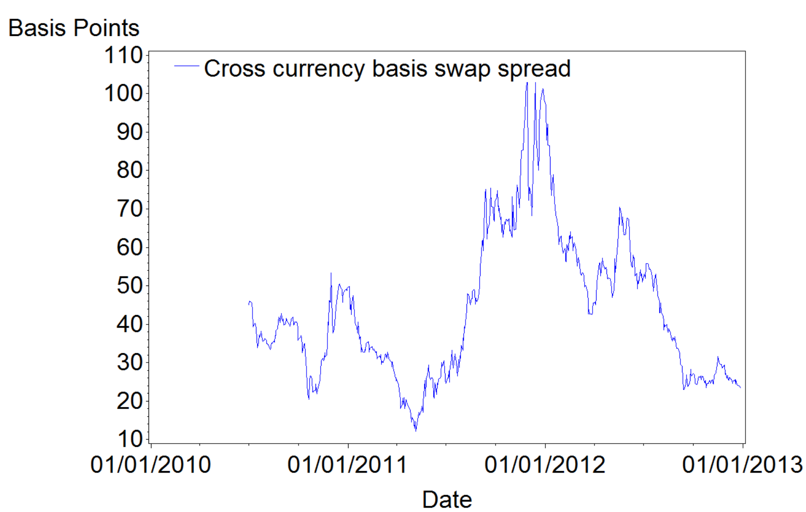

Figure 2: **Time-Series of Bond Yield, Bond Yield Spread, CDS Spread, and Bid-Ask Spread.** The bond yield spread (dotdash line, left-hand axis) is calculated between the Italian (dotted, left-hand axis) and German bonds with ten years to maturity. The CDS Spread (solid, left-hand axis) is the spread for a five-year US-denominated CDS contract. This MTS bid-ask spread (dashed, right-hand axis) is a market-wide illiquidity measure. Our data set consists of transactions, quotes, and orders for all 189 fixed-rate and floating Italian sovereign bonds (Buoni Ordinari del Tesoro (BOT) or Treasury bills, Certificato del Tesoro Zero-coupon (CTZ) or zero-coupon bonds, Certificati di Credito del Tesoro (CCT) or floating notes, and Buoni del Tesoro Poliennali (BTP) or fixed-income Treasury bonds) from July 1, 2010 to December 31, 2012. Data for the bond yield, yield spread, and CDS spread were obtained from Bloomberg.




(a) 3-Month Euribor-German T-Bill Spreads

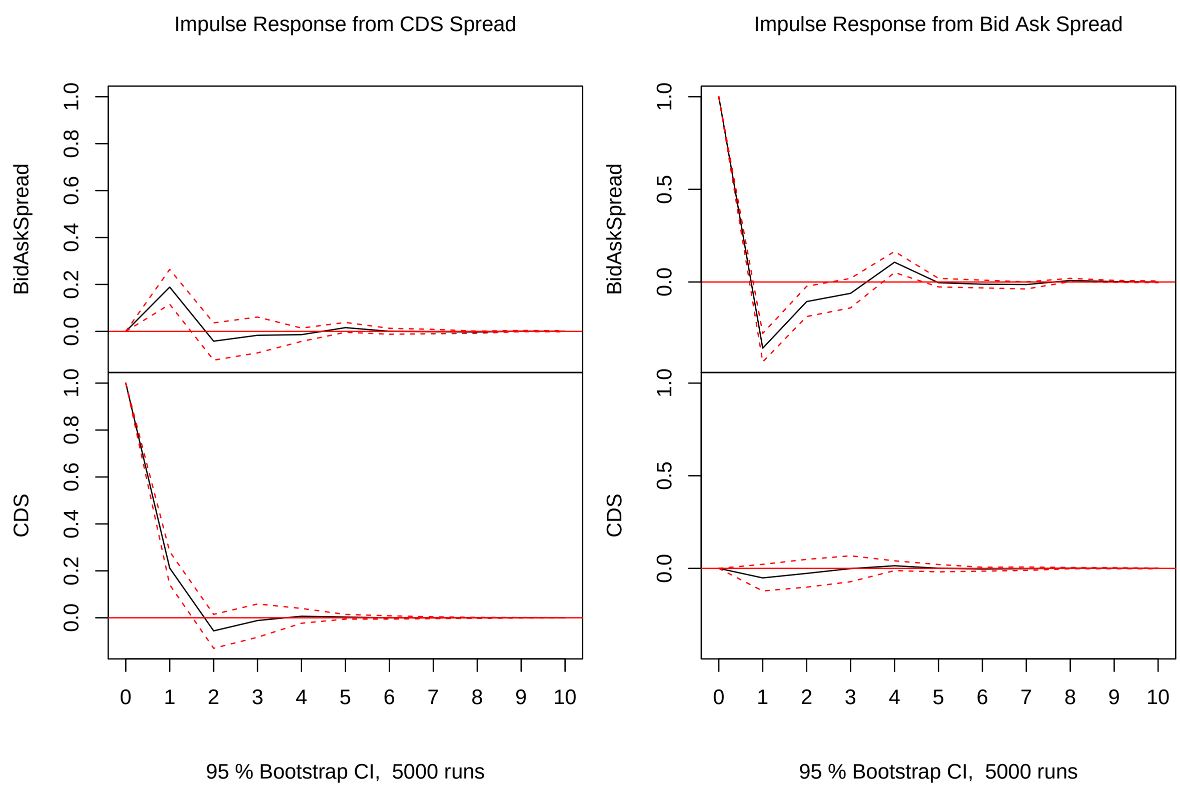



(b) USVIX Index




(c) Cross-Currency Basis Swap Spread

Figure 3: **Time-Series of Macro variables.** The time-series evolution of the global variables: the spread between the three-month Euribor and the three-month yield of the German TBill, the USVIX, and the Cross-Currency Basis Swap Spread are shown in Panels (a), (b), and (c), respectively. Global variables are described in detail in Section 4.1. Our data set was obtained from Bloomberg and covers the period




(a) Shock to CDS spread

(b) Shock to the bond market liquidity

Figure 4: **Impulse Response Functions for the VARX(3,0) System**. This graph shows the evolution of the impulse response functions (IRFs) following a shock in the CDS spread and the bond market liquidity, as measured by the Bid-Ask Spread, in Panels (a) and (b) respectively. The VARX(3,0) system that produces these IRFs is presented in Equation (4) and discussed in Section 6.1. Our data set consists of transactions, quotes, and orders for all 189 fixed-rate and floating Italian sovereign bonds from July 1, 2010 to December 31, 2012.

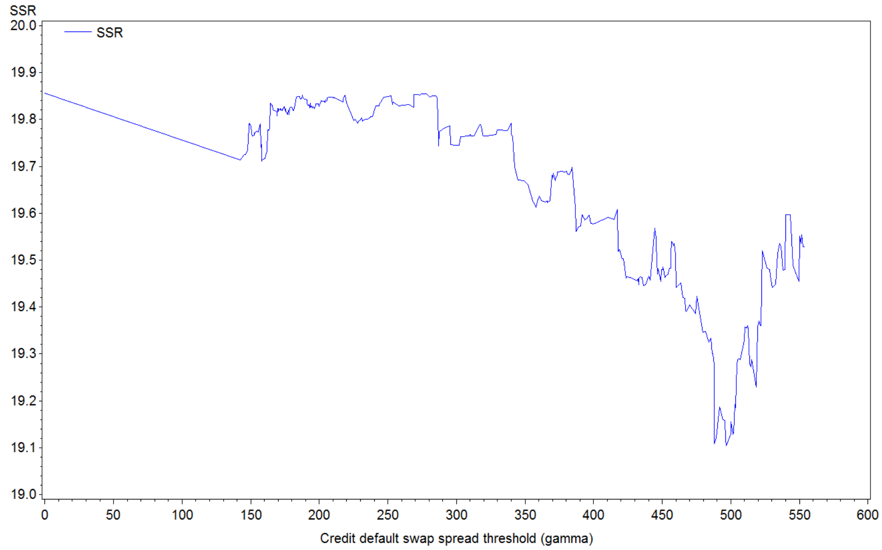



Figure 5: **Sum of Squared Residuals as $\gamma$ Changes.** The evolution of the sum of squared residuals (SSR) from Equation (6) is plotted as the threshold value $\gamma$ changes. The $\gamma$ that minimizes SSR ($\hat{\gamma}$) is the estimate for the threshold. The point at $\gamma = 0$ is the SSR for Equation (5), namely the regression with no threshold. Our data set consists of transactions, quotes, and orders for all 189 fixed-rate and floating Italian sovereign bonds, from July 1, 2010 to December 31, 2012.

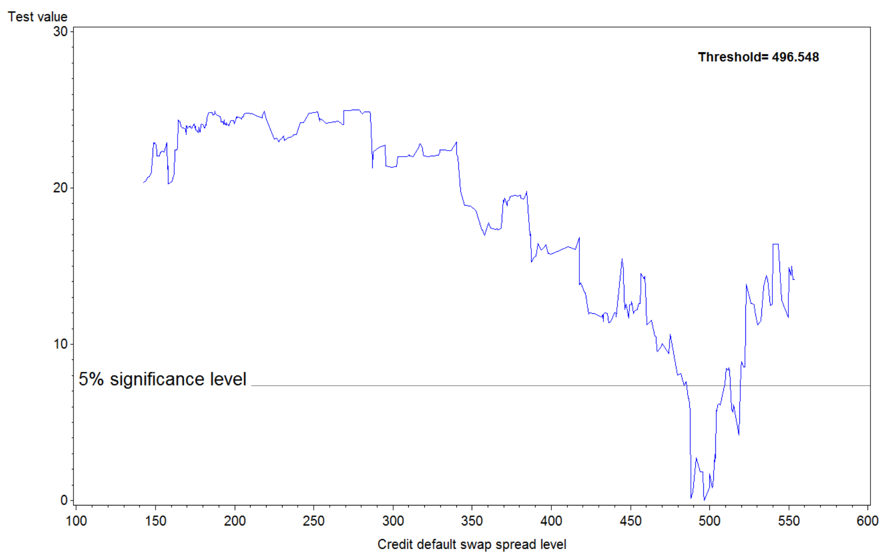



Figure 6: **Test to Determine Confidence Bands around the CDS Threshold.** The test statistic described in Appendix B is plotted here for Equation (6). The test statistic is normalized at 0 at the threshold that minimizes the sum of squared residuals. The horizontal line at 7.35 marks the 5% confidence values for the threshold. Our data set consists of transactions, quotes, and orders for all 189 fixed-rate and floating Italian sovereign bonds, from July 1, 2010 to December 31, 2012.

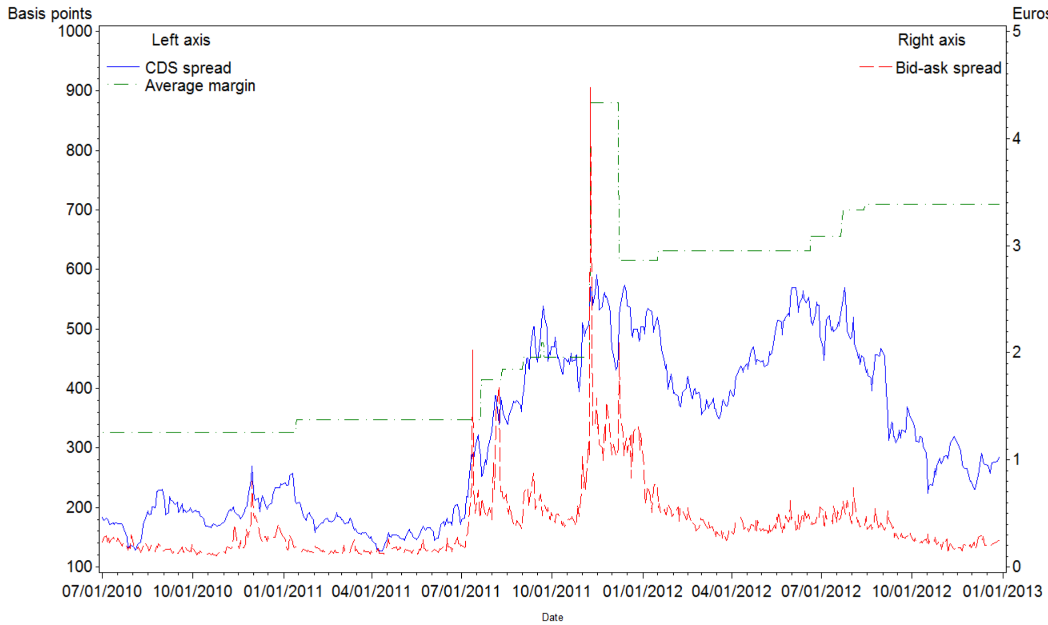



Figure 7: **Time-Series of Margins, CDS Spread, and Bid-Ask Spread.** This graph shows the time-series of the average of the margins (dotdashed, left axis) set by Cassa Compensazione e Garanzia, a clearing house, on Italian bonds, the spread of a five-year CDS contract (solid, left axis), and the liquidity of the bond market (dashed, right axis), as measured by the market-wide bid-ask spread. Our data set consists of transactions, quotes, and orders for all 189 fixed-rate and floating Italian sovereign bonds, from July 1, 2010 to December 31, 2012.

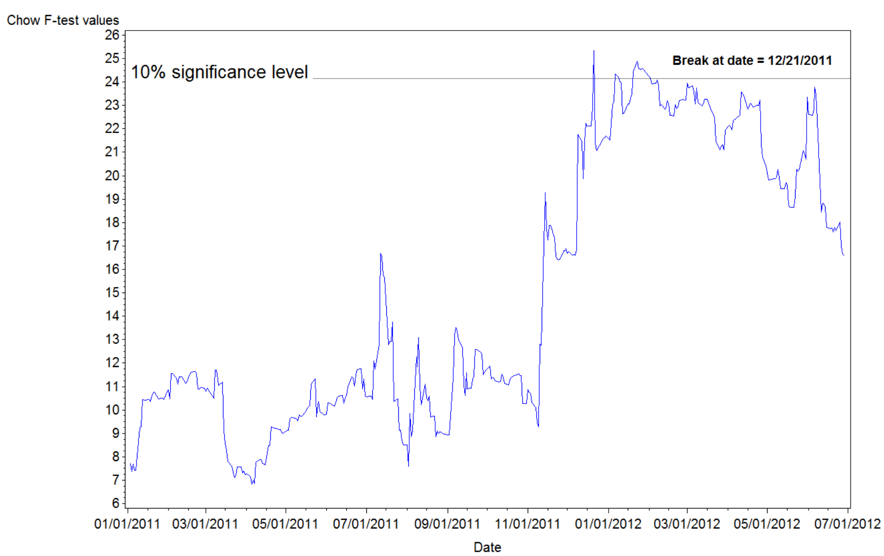
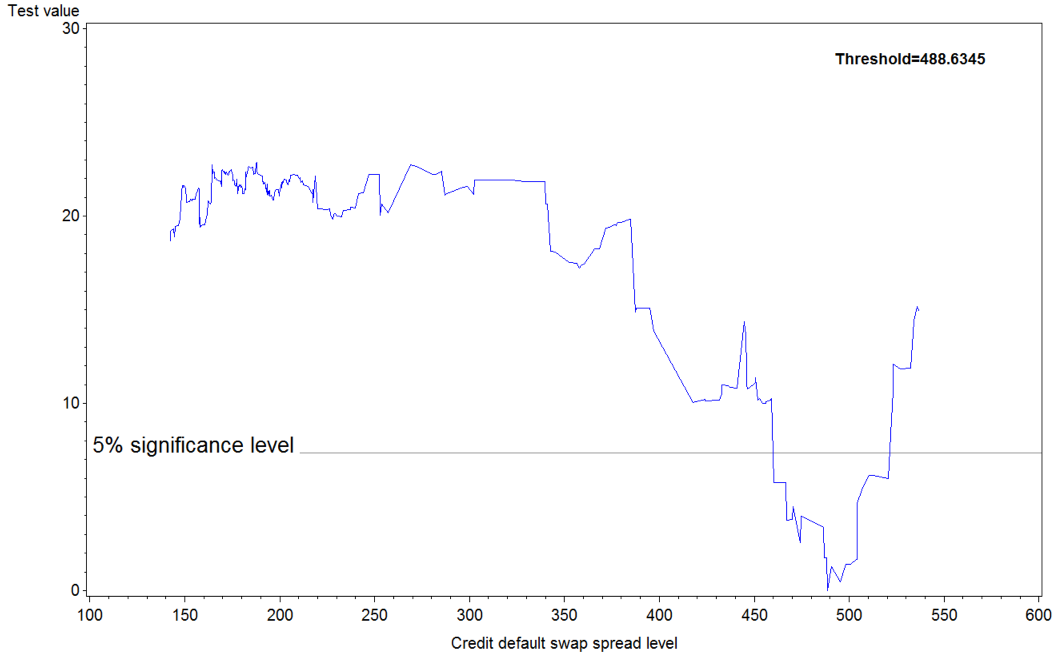
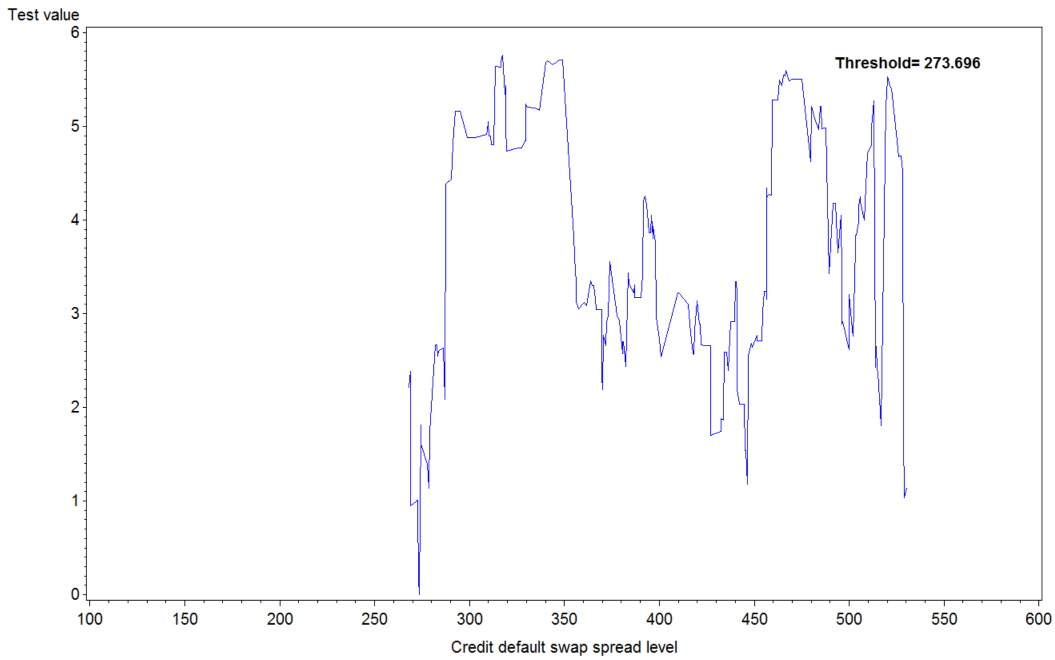



Figure 8: **Structural Break Test.** This figure shows the $F$-test results for the Chow test performed for Equation (6) for each day in our sample, excluding the first and last 20% of observations. The horizontal line marks the 10% level of significance for the largest of the $F$-test values. Our data set consists of transactions, quotes, and orders for all 189 fixed-rate and floating Italian sovereign bonds, from July 1, 2010 to December 31, 2012. The CDS data were obtained from Bloomberg.

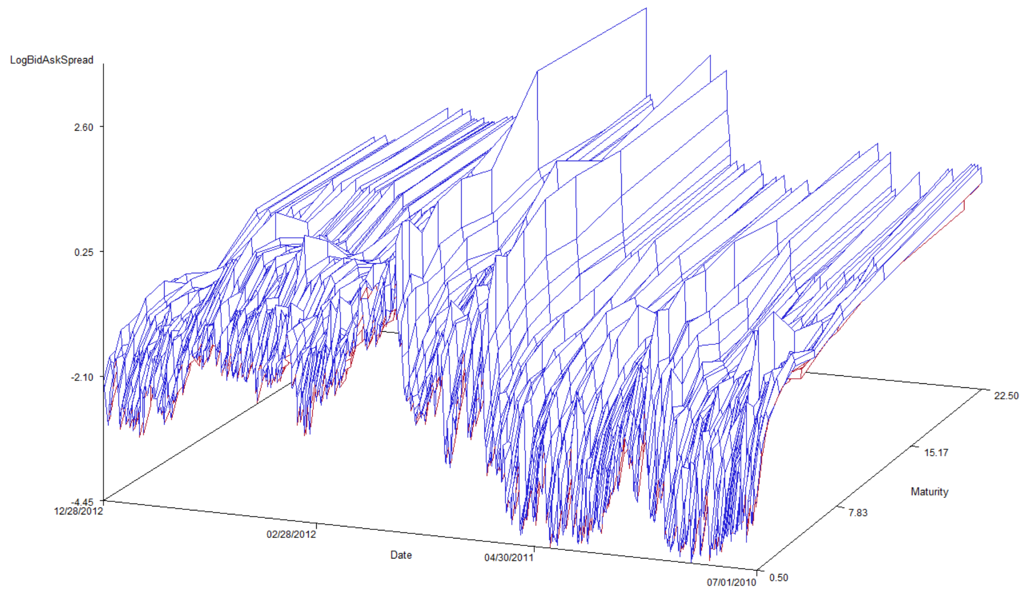



(a) Threshold Confidence Bands Determination: 2011 Sample

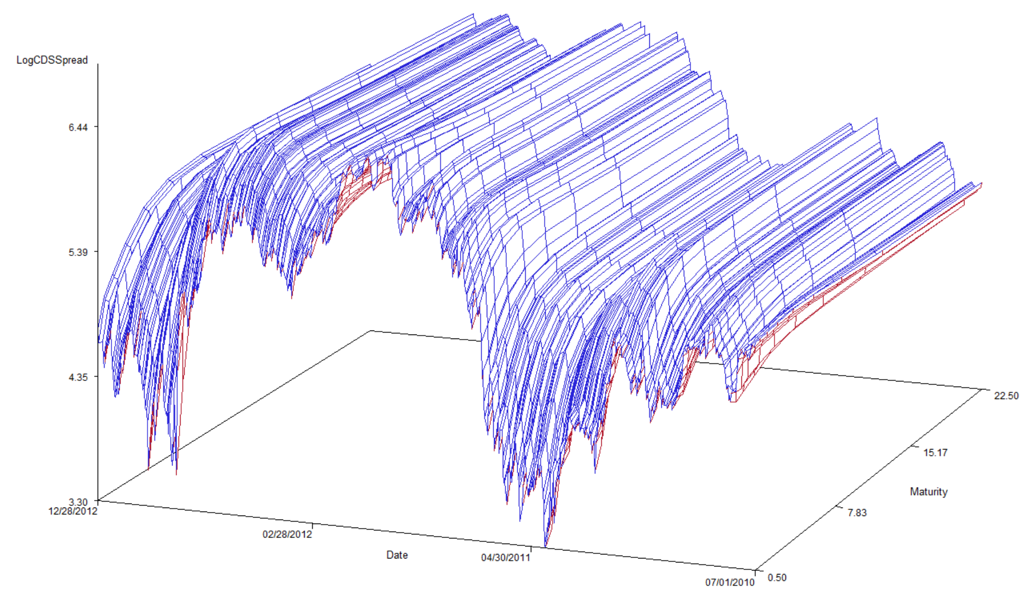



(b) Threshold Confidence Bands Determination: 2012 Sample

Figure 9: **Confidence Bands Determination for Two Subsamples.** The test statistic described in Appendix B is plotted here for Equation (6) in Panels (a) and (b) for the subsamples before and after the structural break, respectively. The test statistic is normalized at 0 at the threshold that minimizes the sum of squared residuals. The horizontal line at 7.35 marks the 5% confidence values for the threshold.




(a) Bid-Ask Spread Evolution and Maturity.




(b) CDS Spread and Maturity.

Figure 10: **Bid-Ask Spread, CDS Spread, and Maturity.** This figure shows time-series of the log of the average bid-ask spread for bonds as a function of maturity and the time-series of the log of the CDS spread for 9 maturities of the contract, in Panels (a) and (b), respectively. Our data set consists of transactions, quotes, and orders for all 189 fixed-rate and floating Italian sovereign bonds, from July 1, 2010 to December 31, 2012.

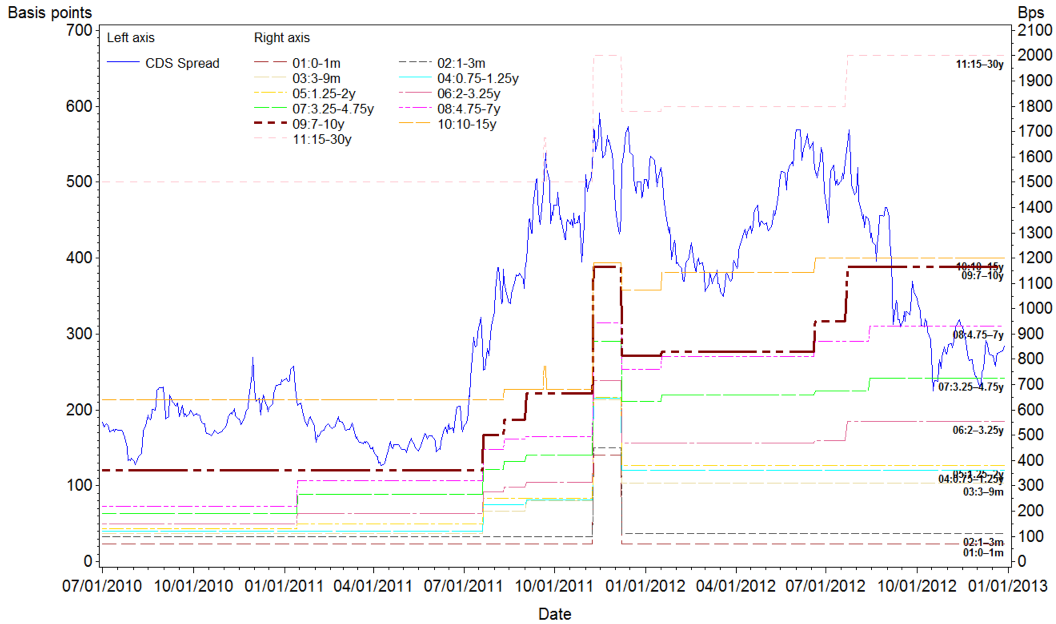



Figure 11: **CDS Spread and Margins for the Cross-Section of Italian Bonds.** This figure shows time-series of the CDS spread for a 5-year contract and the margins applied to different maturity bonds by Cassa Compensazione e Garanzia.

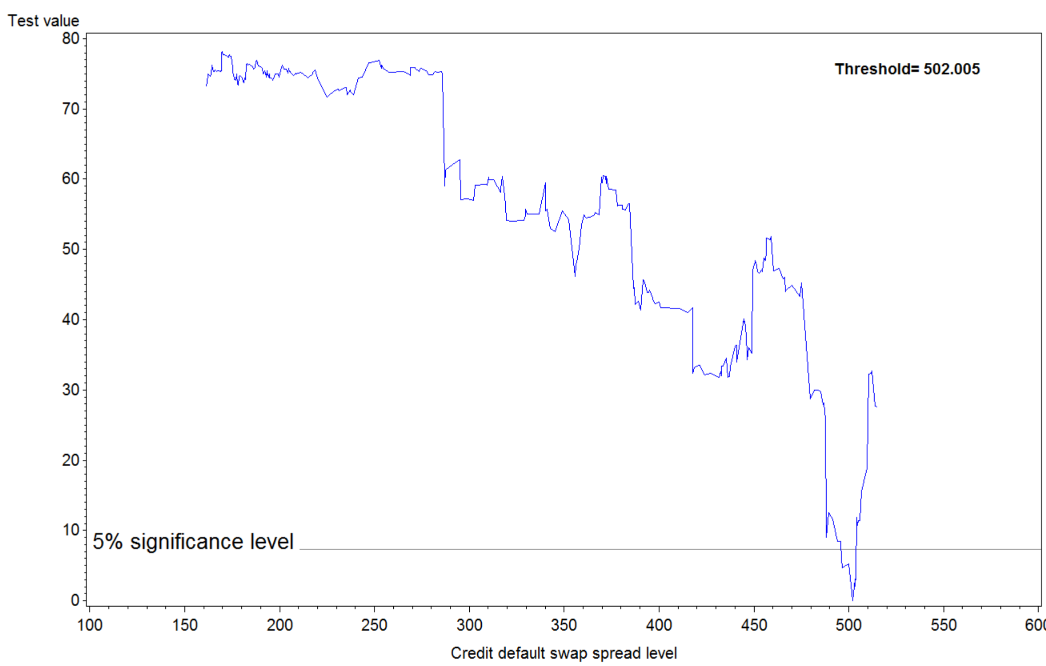



Figure 12: **Confidence Bands Determination for the Panel Analysis.** The test statistic described in Appendix B is plotted here for Equation (8). The test statistic is normalized at 0 at the threshold that minimizes the sum of squared residuals. The horizontal line at 7.35 marks the 5% confidence values for the threshold.

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