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DECOMPOSING EURO AREA SOVEREIGN SPREADS: CREDIT, LIQUIDITY AND CONVENIENCE

by Marcello Pericoli* and Marco Taboga *

Abstract

The paper provides an empirical analysis of sovereign bond spreads for selected euro-area countries. Several methodologies are used to measure and assess the relative importance of three components of sovereign spreads: credit premiums, liquidity premiums and convenience yields. We find that, except for Germany, credit premiums explain most of both the level and the variability of spreads, while the other two factors play a limited role, although in several cases they are statistically significant and may become economically relevant during brief episodes of illiquidity.

JEL Classification: G12.
Keywords: sovereign spreads, liquidity premia, convenience yields.

Contents

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1 Introduction[1]

During the European sovereign debt crisis (e.g., Lane 2012, Shambaugh, Reis and Rey 2012), the yield differentials between sovereign bonds and risk-free assets – the so called sovereign spreads – increased dramatically in some euro area countries. There is a wide consensus that these increases reflected a sharp rise in the credit premia[2] required by investors as compensation for the default risk of sovereign issuers. However, some studies (e.g., Monfort and Renne 2014) have advanced the hypothesis that also increases in liquidity premia might have significantly contributed to the hike in sovereign spreads. In this paper we conduct an empirical analysis aimed at testing this hypothesis.

Assessing the relative importance of credit and liquidity premia is extremely relevant from a policy perspective, because it helps to understand what policy actions need to be undertaken in order to reduce sovereign spreads. In particular, credit premia can be reduced by increasing the perceived creditworthiness of sovereign issuers, for example, through budget discipline and structural reforms, while liquidity premia can be reduced by increasing the liquidity of sovereign bonds, which, in turn, can be achieved by fostering a better functioning of bond markets and by reducing transaction costs such as bid-ask spreads, search costs, and delays in trade execution.

Furthermore, from an investment manager's perspective, quantifying liquidity premia is important because it allows to identify the benefits of long-term strategies such as buying bonds and holding them to maturity.

We analyze 10-year sovereign bond spreads in the period 2007-2012, focusing on two euro-area countries (Italy and Spain) that were affected by significant sovereign tensions,

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[2] For the sake of simplicity, in what follows we will use the term "premium" to refer to the overall compensation required to bear a given risk, which includes both the compensation for the expected losses generated by the risk and the risk premium *stricto sensu*, that is, the additional remuneration that is required by risk-averse agents on top of the compensation for the expected losses.

but continued issuing sovereign bonds throughout the whole period[3]. During the sovereign crisis, not only the spreads on Italian and Spanish sovereign bonds experienced dramatic increases, but also the liquidity of these bonds deteriorated significantly, as indicated, for example, by a surge in their bid-ask spreads. We also analyze Germany, not only because it is representative of the group of euro area countries that were less affected by the tensions, but also because its sovereign bonds are commonly considered a risk-free benchmark, as well as a "flight-to-quality asset" (e.g., Arghyrou and Kontonikas 2012, Arce, Mayordomo and Peña 2013).

As a first step of our analysis, we discuss the identification of credit and liquidity premia and we propose several proxies for these premia. We use data from the credit default swap (CDS) market to derive proxies for the credit premia. We find evidence that raw CDS spreads might be inaccurate proxies for the credit premia, and we propose other proxies, based on risk-neutral expected default losses estimated from the whole term structure of CDS spreads. These proxies provide a better fit to sovereign spreads than raw CDS spreads. For liquidity premia, we use a measure of the degree of illiquidity of sovereign bonds (the bid-ask spread) and several estimates of the liquidity premia embedded in other financial instruments, which approximate investor's appetite for liquidity.

We perform a regression analysis where we consider several different model specifications and choices of the liquidity and credit proxies. We find that our proxies for the credit premia explain a predominant portion of the variability of Italian and Spanish sovereign spreads (up to 97 per cent), while the liquidity variables have little incremental explanatory power. Furthermore, liquidity effects are statistically significant in some specifications, but the inferences about these effects are sometimes not robust across models. As far as economic significance is concerned, the magnitude of liquidity effects is found to be on average small

[3]Note that even if Spain resorted to a financial assistance programme in 2012, the programme concerned only its financial sector, and Spain continued to have market access and to issue sovereign bonds, unlike the other countries that asked for financial assistance (Greece, Ireland and Portugal).

(across models and time). For example, the Italian sovereign spread is on average 172 basis points in our sample, and less than 10 basis points are estimated to be due to liquidity effects. However, it is not possible to rule out that these effects were large during at least part of our sample. For instance, according to one of our specifications, in November 2011 a temporary spike of the Italian bid-ask spread beyond 100 basis points is estimated to have had an impact larger than 200 basis points on the Italian sovereign spread.

In most specifications, increases in investors' appetite for liquidity are estimated to determine a reduction in Italian and Spanish sovereign spreads. We argue that this finding might be explained by the fact that, despite the deterioration in liquidity experienced on the secondary market for Italian and Spanish sovereign bonds, these bonds continued to be highly liquid assets for banks because they could be used in repo transactions (not only with other banks but also with the central bank) and could be temporarily turned into liquidity at no cost even when liquidation on the spot market became expensive. The premium associated to this possibility, which represents the economic value of pledgeability and is usually known as convenience yield, might have been higher exactly when investors attached a higher value to liquidity, that is, when their appetite for liquidity was higher. By running regressions that include a proxy for the convenience yield, we find empirical support for this hypothesis.

To our knowledge, ours is the first analysis of sovereign spreads during the sovereign debt crisis that attempts to identify convenience yields as a separate component (beyond credit and liquidity premia). We note that there seems to be no unanimous agreement on a definition of convenience yield and no standard methodologies to estimate the convenience yield on a given bond, although there are several papers that deal with the differences in convenience yield induced by repo specialness (e.g., Buraschi and Menini 2002, Cherian, Jacquier and Jarrow 2004). Nonetheless, we believe that attempting to address the issue of convenience yields is important in order to have a complete picture of the non-credit components of sovereign spreads.

We also use Bayesian vector autoregressions (VARs) to analyze the dynamic interactions among sovereign spreads and our proxies for credit and liquidity premia. The results from the VAR analysis are similar to those from the regression analysis: shocks to the credit premia explain the great majority of the variability in sovereign spreads, while shocks to the liquidity proxies explain a small fraction of variance and are often not statistically significant.

The impact of illiquidity on asset prices has been studied extensively, from both a theoretical and an empirical point of view.

Constantinides (1986) provided one of the earliest theoretical contributions and showed that illiquidity has a second-order effect on asset prices. In his intertemporal equilibrium model the primary effect of illiquidity is to reduce the frequency of trade, and this reduction induces deviations from optimal portfolio holdings that have a negligible impact on investors' utility. Subsequently, however, other scholars have proposed models where illiquidity has a first-order effect on equilibrium asset prices. For example, in Huang's (2003) model, illiquidity has a first-order effect if it is coupled with borrowing constraints. Furthermore, in principle, the impact of illiquidity should depend on how illiquidity covaries with all non-diversifiable (priced) sources of risk (as, for example, in the liquidity-adjusted capital asset pricing model of Acharya and Pedersen 2005).

From an empirical viewpoint, there is abundant evidence that illiquidity has statistically and economically significant effects on the prices of stocks (e.g., Amihud and Mendelson 1986, Pastor and Stambaugh 2003, Acharya and Pedersen 2005) and corporate bonds (e.g., Longstaff, Mithal and Neis 2005, Chen, Lesmond and Wei 2007). Instead, the evidence regarding sovereign bonds is less abundant. More importantly, it is quite mixed, especially as far as the period of the European sovereign debt crisis is concerned.

Favero, Pagano and von Thadden (2010) analyze the sovereign spreads of a number of euro area countries in the period 2002-2003. They find that the impact of illiquidity on spreads is sometimes statistically significant, but its economic magnitude is not large. Furthermore, consistently with the predictions of a model of endogenous liquidity demand, the size of the impact depends on the interaction with a variable that measures aggregate risk.

Beber, Brandt and Kavajecz (2009) also analyze euro-area sovereign spreads in a pre-crisis period (2003-2004). They find that the bulk of sovereign spreads is explained by differences in credit quality, but liquidity can play a nontrivial role in times of heightened market uncertainty.

Note that in the pre-crisis periods analyzed by Favero, Pagano and von Thadden (2010) and Beber, Brandt and Kavajecz (2009) sovereign spreads were of a different order of magnitude compared to the crisis period. For example, in the period analyzed by Beber, Brandt and Kavajecz (2009) the average sovereign spread is less than 4 basis points for Italy, as compared to an average of 172 basis points in our sample period (from 2007 to 2012).

Monfort and Renne (2014) propose a no-arbitrage term-structure model to jointly price the sovereign bonds issued by a number of euro area countries and by KfW, a German agency whose bonds have the same credit risk of German sovereign bonds because they are guaranteed by the Federal Government, but are much less liquid than German government bonds. In their model, the short-term yield spread of a given issuer is a linear function of two latent variables: one that is country-specific and one that is common to all issuers but the German government. Using data from July 2006 to February 2013, they find that the common variable explains a significant portion of the variability in the term-structures of sovereign spreads. Based on the imposed exclusion restriction that the common variable affects less liquid KfW bonds but not German bonds, they suggest that this variable might be a proxy for the premium demanded by investors as compensation for illiquidity. To compare their results with ours and with others reported in the literature, we note that they estimate that the liquidity premium is on average 45 basis points out of 168 (for Italy, on the 5-year maturity).

Wagenvoort and Zwart (2014) also analyze the sovereign yields of a number of euro area countries in a period (from February 2006 to December 2011) that overlaps with the period of the sovereign crisis. They set up a factor model where risk-free rates, credit premia and liquidity premia are latent and they propose a novel technique (called Longitudinal Factor Analysis) to estimate these latent factors. They impose several identifying restrictions on factor loadings (among which proportionality between liquidity premia and transaction costs) and estimate that liquidity premia were on average small during their sample period (less than 3 basis points on average for Italy on the 10-year maturity).

Both Monfort and Renne (2014) and Wagenvoort and Zwart (2014) use latent variable models. The fact that they reach very different conclusions about the magnitude of liquidity premia suggests that the results from their class of models might be sensitive to the assumptions made in order to identify the premia.

In light of the heterogeneity of the results found in the literature, we advocate the necessity of taking an approach that is – as much as possible – free of assumptions. For this reason, we propose a broad range of proxies for credit and liquidity premia and we use several models to estimate their impact on sovereign spreads.

The rest of the paper is organized as follows: Section 2 discusses the data and our choice of the proxies for credit and liquidity premia; Section 3 presents some descriptive statistics; Section 4 contains the regression analysis; Section 5 discusses the Bayesian VARs; Section 6 concludes.

2 The data

We employ data on three countries (Germany, Italy and Spain) from January 2007 to November 2012. Our choice of the end date is motivated by the fact that after November 2012 sovereign CDS trading restrictions came into effect in the European Union. As we use also data from the CDS market to identify credit premia, we prefer to avoid using prices that could have seriously been distorted by these restrictions (see, e.g., Banerjee and Graveline 2014). As far as the start date is concerned, we performed some robustness checks on our empirical analyses, by moving the start date forward (to the beginnings of 2008 and 2009), and we found that the main results from the analyses do not change significantly.

Although all our data are available at a daily frequency, we sample it at a weekly frequency, in order to reduce distortions due to market microstructure effects[4].

Details of the data are reported in the next sub-sections.

2.1 Government bond yields

We focus our attention on the 10-year maturity. In order to avoid coupon biases (e.g., Livingston and Jain 1982, Papageorgiou and Skinner 2006) and to ensure comparability across markets, we use 10-year zero-coupon spot rates obtained by fitting a term structure model to government bond prices. For each country, the term structure at a given date is estimated with the prices of all fixed-coupon bonds that were traded at that date, excluding those whose residual maturity was less than 30 days. As in Fisher, Nychka and Zervos (1995), we fit a smoothing spline to the curve of instantaneous forward rates, by minimizing mean squared pricing errors, and we obtain zero-coupon spot rates by integrating the interpolated forward curve.

2.2 Risk free rates

We use the 10-year Overnight Indexed Swap (OIS) rate as a proxy for the 10-year risk-free rate. OIS are swap contracts where one counterparty receives a variable payment indexed to the interest rate on overnight unsecured interbank deposits between prime banks, and the other counterparty receives the fixed OIS rate. Because overnight interbank deposits

[4] In particular, we want to reduce at a minimum spurious autocorrelations due to potentially stale CDS quotes and to the fact that the closing prices of instruments traded on different markets might be recorded at different daytimes.

between prime banks are considered virtually risk free, OIS rates are deemed a very good proxy for long-term risk-free rates (e.g., Morini 2009, Mercurio 2009, Ejsing, Grothe and Grothe 2012, Taboga 2013). Given the timing of the payments of OIS contracts, a simple recursive calculation can be used to extract a term structure of zero-coupon spot rates from the term-structure of OIS rates, so as to allow for a straightforward comparison with zero-coupon government bond rates (see previous subsection).

2.3 Proxies for credit premia

The credit premium is the part of the yield of a bond that compensates the bond holder for bearing credit risk, which includes not only the risk that the issuer of the bond will not honor its obligations, but also the risk that the price of the bond will fluctuate in response to unpredictable variations in the credit quality of the issuer or in the credit premia that investors will require in the future.

Our proxies for credit premia are based on CDS spreads. On a first approximation, we could think that the spread on a 10-year CDS written on a given 10-year bond is equal to the credit premium on that bond. The underlying argument runs as follows: if we buy a defaultable bond and we protect ourselves from credit risk via a CDS, we are *de facto* holding a default-free bond; therefore, if there are no arbitrages, the net yield of this protected position must not embed any credit premium; this implies that the additional yield we receive on the defaultable bond (the credit premium) must be equal to the CDS premium we pay as insurance against default. The fallacy in this argument is that the cash flows of a protected position (defaultable bond plus CDS protection) are not independent of credit events. As a matter of fact, when the issuer defaults and the CDS is triggered, the protection buyer delivers the defaulted bond and receives its face value. This represents an accelerated re-payment of principal (and deprivation of further coupons), and its possibility generates uncertainty about the cash flows of the protected position. Such uncertainty is not faced by the buyer of a default-free bond, whose cash flows do not depend on credit events. There are also other reasons why the cash flows of a protected position have a residual dependency on credit events (see Hull, Predescu and White 2004). For these reasons, that is, for the residual credit risk faced by the holder of a protected defaultable bond, there can be a difference between the CDS spread and the credit premium on the defaultable bond (the difference can be due also to factors that are not related to credit risk, for example, CDS liquidity premia; see the discussion below). There is ample empirical evidence that this difference, known as "CDS-bond basis", can be substantial and can vary through time (e.g., Blanco, Brennan and Marsh 2005).

In this paper, we use the 10-year CDS spread as one of our proxies for the credit premium on a 10-year bond. However, in order to take into account the fact, explained above, that the CDS spread is an imperfect measure of the credit premium on a bond, we also use another proxy: the yearly rate of expected default losses on the 10-year bond. In particular, we use Hull and White's (2000, henceforth HW) procedure to extract an implied risk-neutral survival function S_T from the whole term structure of CDS spreads, and we compute the yearly expected loss rate HW_T for a given maturity T as

$$HW_T = \left(S_T^{-1/T} - 1\right)(1 - R)$$

where S_T is the probability of not observing a default up to year T, and R is the fractional recovery rate, that is, the fraction of value of the bond that is recovered in case of default (Duffie and Singleton 1999, Duffie, Schroder and Skidas 1996). HW_T is the product of two terms. The first term $(S_T^{-1/T} - 1)$ is a constant hazard rate that can be interpreted as the average probability[5] of observing a default within a year. The second term $(1 - R)$ is a thinning term (Lando 1998) that allows to convert hazard rates into expected loss rates. In

[5] This is a risk-neutral probability.

the computation of both the survival function and the loss rate, the recovery rate is assumed to be equal to 40%. This is a standard assumption, and estimates of expected loss rates have been shown to be fairly insensitive to different assumptions, because changes in R tend to be compensated by changes in S_T (see, e.g., Hull and White 2003).

Note that CDS premia might embed components unrelated to the credit risk of the underlying bonds. For example, they might include counterparty risk premia to compensate the parties of the CDS contract for the risk that their counterparties will not live up to their contractual obligations. In this paper we ignore counterparty risk on CDS, based on the existing evidence that counterparty credit risk premia not only have been historically negligible, but have been practically vanishing in recent years thanks to the widespread practice of collateralizing CDS contracts (Arora, Gandhi and Longstaff 2012).

CDS spreads might also embed liquidity risk premia related to the illiquidity of CDS contracts. Because CDS are derivative contracts in zero net supply, it is in principle not obvious whether the liquidity premium is positive or negative, that is, whether it is demanded by buyers or sellers of CDS protection. The empirical evidence on this issue is mixed (see, e.g., Chen, Cheng and Wu 2005, Bühler and Trapp 2009, Bongaerts, de Jong and Driessen 2011, Tang and Yan 2012, Badaoui, Cathcart and El-Jahel 2013). Furthermore, there seems to be no widespread agreement on the theoretical foundations and the methodologies to measure CDS liquidity premia. For these reasons, we do not deal with the issue of CDS liquidity premia in our baseline specifications; we do, however, recognize that it could have an impact on our results and we propose a way to deal with it in our robustness checks.

2.4 Proxies for liquidity premia

The liquidity premium is the part of the yield of a bond that compensates the bond holder for the transaction costs she might incur in case she needs to sell the bond before its expiration[6] (e.g., Constantinides 1986, Amihud and Mendelson 1986, Vayanos and Vila 1999, Jang, Koo, Liu and Loewenstein 2007); it also compensates the bond holder for the risk that the price of the bond will fluctuate in response to unpredictable variations in the transaction costs or in the liquidity premium that investors will require in the future (as, for example, in Duffie and Singleton 1999, Liu, Longstaff and Mandell 2006, Feldhütter and Lando 2008). The transaction costs can be direct costs such as broker fees and bid-ask spreads, or indirect costs arising from delay and search associated with trade execution (e.g., Duffie Garleanu and Pedersen 2002, Acharya and Pedersen 2005, Duffie Garleanu and Pedersen 2005). The magnitude of the transaction costs determines how liquid a bond is: the higher the costs, the more illiquid the bond.

From a conceptual viewpoint, it can be helpful to think of the liquidity premium on a bond as the product of two factors: the quantity of liquidity risk, that is, a quantification of the risks for which the premium represents a compensation, and the price per unit of liquidity risk, determined by market forces such as investors' demand for liquid assets. Of course, in reality things can be more complicated, and there can be a multiplicity of risks and associated prices, as, for example, in the liquidity-adjusted capital asset pricing model of Acharya and Pedersen (2005), where the liquidity premium on an asset is determined not only by the expected transaction costs specific to the asset, but also by the covariance of the transaction costs with the returns and the liquidity of the market portfolio.

While there is no readily available measure of the liquidity premium required by investors to hold the government bonds of a given country, the literature has identified some variables

[6] Jang, Koo, Liu and Loewenstein (2007) define the liquidity premium as the maximum expected return an investor is willing to exchange for zero transaction cost.

that, on the basis of theoretical considerations, should be correlated with the premium. To understand the role of these variables, we can think of them as belonging to either one of two categories:

1. variables that measure the quantity of liquidity risk faced by the holders of the bond. The underlying argument is that there should be positive correlation between the quantity of liquidity risk and the liquidity premium even if there is time variation in the proportionality coefficient between the two, that is, even if the price of liquidity risk changes through time.

2. variables that reflect how liquidity risk is priced by investors (measures of investors' appetite for liquidity). These variables are typically computed as the yield spread between two securities that have identical or almost identical cash flows and levels of credit risk, but different degrees of liquidity. The spread represents the compensation required by investors for the relative illiquidity of one of the two securities. Changes in the spread can reflect changes either in the relative illiquidity of the two securities or in the price of liquidity risk. If the latter is similar across different assets, as suggested by the evidence on the existence of systematic components in liquidity premia (e.g., Lee 2011) and by the fact that liquidity premia are correlated with global funding conditions (e.g., Nagel 2012), then the spread should be correlated with the liquidity premia on other assets.

The first variable we use in our analysis is the bid-ask spread[7] on government bonds, a widely used measure of illiquidity that falls within category 1) above. This variable is the

[7] For each sovereign issuer and at each point in time, we average the bid-ask spreads on the three most recently issued 10-year bonds. For each bond, the spread is expressed in relative terms:

$$\frac{1}{2}\frac{P_A - P_B}{P_A + P_B}$$

where P_B and P_A are the bid and ask prices of the bond.

only country-specific measure of illiquidity we have in our dataset. This should not represent a limitation, as documented by Favero, Pagano and von Thadden (2010), who analyze also other measures of bond illiquidity and find that the bid-ask spread is the most significant one.

The other variables we use fall within category 2) and measure investors' appetite for liquidity. They are:

1. the yield spread between KfW bonds and German government bonds[8] (e.g., Schwarz 2010, Monfort and Renne 2014, Ejsing, Grothe and Grothe 2012). KfW bonds are agency bonds guaranteed by the German government. They are characterized by the same credit risk of government bonds, but they are less liquid.

2. the yield spread between off-the-run and on-the-run government bonds[9] (e.g., Fontaine and Garcia 2012). On-the-run bonds are bonds that have been recently issued. Usually, they are traded more frequently and they are more liquid than bonds of comparable maturity that have been issued less recently (the so-called off-the-run bonds).

3. the yield spread between AAA corporate bonds and AAA government bonds[10] (e.g., Bao, Pan and Wang 2011, de Jong and Driessen 2012). Despite having a similar level of credit risk, as indicated by the AAA rating (and the associated historical default records), corporate bonds are usually much less liquid than government bonds.

4. the simple average of the above three variables (denoted by LIQ in what follows). It is highly correlated (coefficient higher than 99 per cent) with the first principal component

[8] We compute the yield spread by fitting two zero-coupon yield curves (one for KfW bonds and one for government bonds) with the methodology described in Section 1.1, and by computing the difference between the two curves at the 10-year maturity.

[9] We compute it by averaging, across the three countries in our sample, the yield differential between the most recently issued 10-year bond and the 10-year bond that was issued just before the most recent.

[10] We compute it as the option-adjusted spread on the Bank of America Merrill Lynch 7-10 Year AAA Euro Corporate Index.

(computed from the correlation matrix of the variables), which explains 68 per cent of the variance.

3 Descriptive statistics

Table 1 reports some descriptive statistics (mean, standard deviation, quantiles) of the variables described above and used in our empirical analysis. Plots of some of these variables can be found in Figure 1 (sovereign spreads and proxies for credit premia) and Figure 2 (bid-ask spreads and proxies for liquidity premia). We do not provide detailed comments on the temporal evolution of the variables (but see, e.g., De Santis 2013), and we concentrate only on few relevant aspects.

First of all, we note that the sovereign spread, that is, the difference between the zero-coupon and the OIS 10-year rates, is on average positive also for Germany, contrary to the widespread belief that German yields can be used as proxies for risk-free rates. The dynamics of the German spread (Figure 1), as computed in this paper, are similar to those estimated by Wagenvoort and Zwart (2014), who propose a new factor analysis technique to extract the risk-free rate – a latent variable in their model – from a panel of euro area sovereign bond yields; their measure of the German sovereign spread, like ours, has been on average positive since the beginning of 2009.

The second fact we would like to stress concerns bid-ask spreads. While these spreads are on average not large in our sample (6, 20 and 33 basis points for Germany, Italy and Spain respectively), they are characterized by a significant variability (Figure 2). In particular, there are several periods in which the Italian and Spanish bid-ask spreads are close to 100 basis points[11]. These figures suggest that Italian and Spanish bonds have experienced significant illiquidity during at least some parts of our sample period.

[11] During a short episode of political turmoil in 2011, the Italian bid-ask spread spiked to 150 basis points.

4 Regression analysis

The analysis in this section is based on the following linear decomposition (e.g., Reinhart and Sack 2002):

$$R_{i,t} = R_t^f + C_{i,t} + L_{i,t} \tag{1}$$

where time periods and countries are indexed by the subscripts t and i respectively, $R_{i,t}$ is the zero-coupon sovereign yield, R_t^f is the risk-free rate, $C_{i,t}$ is the credit risk premium and $L_{i,t}$ is the liquidity risk premium.

Because the OIS rate is considered an excellent proxy for the risk-free rate (see Subsection 2.2), but the proxies we have for the credit and liquidity premia are at best imperfect, we estimate the following empirical counterpart of (1):

$$R_{i,t} - OIS_t = \alpha_0 + \alpha_1 c_{i,t} + \alpha_2 l_{i,t} + \varepsilon_{i,t}$$

where $c_{i,t}$ and $l_{i,t}$ are proxies for $C_{i,t}$ and $L_{i,t}$ respectively, α_i are regression coefficients, and $\varepsilon_{i,t}$ is a zero-mean error term orthogonal to the regressors. The regression coefficients and the error term can be thought of as the result of substituting C_t^i and $L_{i,t}$ in equation (1) with their linear projections on c_t^i and $l_{i,t}$.

The results from the regression analysis are reported below, and are subdivided by country, to facilitate exposition.

For each country, we estimate 24 different specifications of the regression model, obtained by combining the following alternative choices:

- in some models, we use the CDS spread as a proxy for the credit premium, while in other models we use the yearly expected credit loss rate HW;
- in some models, we impose the constraint $\alpha_1 = 1$, which is equivalent to assuming that our proxies for the credit premium are unbiased, while in other models we leave α_1 free

to vary;

- in some models, we include only the bid-ask spread (our proxy for illiquidity) or only a proxy for the price of liquidity risk (see Subsection 2.4), while in other models we include both;

- finally, we use different proxies for the price of liquidity risk (or investors' appetite for liquidity); we report the results obtained with the Kfw-Bund spread and with the LIQ indicator; the results obtained with the other proxies are not meaningfully different from those we report.

All of the statistics in this section are based on heteroskedasticity and autocorrelation consistent estimates of the standard errors (we use Newey-West estimators with automatic selection of the bandwidth).

4.1 Results for Italy

Results for Italy are reported in Table 2.

As far as the proxies for the credit premium are concerned, both the CDS spread and the HW loss rate have a positive coefficient and are significant at all conventional levels of confidence. Moreover, they alone explain most of the variability in sovereign yields, with R^2 of up to 97%. In all the specifications, the HW loss rate provides a better fit to the data than the CDS spread. Furthermore, imposing the constraint $\alpha_1 = 1$ causes a noticeable decrease in R^2 when the CDS spread is used, but not when the HW loss rate is used. As a matter of fact, in the latter case the restriction $\alpha_1 = 1$ is almost never rejected by formal statistical tests, while in the former case it is rejected in the majority of specifications. These results suggest that the raw CDS spread is probably too rough a measure of the credit premium (see the theoretical reasons illustrated in Subsection 2.3).

As far as the liquidity variables are concerned, we find that their explanatory power is limited. By adding the liquidity variables, we observe modest increases in R^2 (with respect to the models in which the only explanatory variable is a proxy for the credit premium).

In the majority of specifications, the bid-ask spread has a positive and statistically significant coefficient. This is consistent with the expected effect: the higher transaction costs are, the more illiquid a bond is, and the higher its liquidity premium and its spread are. The magnitude of the effect is estimated to be about 10 basis points (by taking an average across time and across specifications), as compared to an average value of more than 170 points for the sovereign spread. Moreover, the additional explanatory power of the bid-ask spread, in terms of increase in R^2, is never higher than 1 per cent. However, it is not possible to rule out that increases in the bid-ask spread had an economically significant impact on the sovereign spread during at least part of our sample. For instance, according to one of our specifications, in November 2011 a momentary spike of the bid-ask spread beyond 100 basis points is estimated to have had an impact larger than 200 basis points on the Italian sovereign spread.

The Kfw-Bund spread is not significant in half of the specifications where it is included. Furthermore, when it is significant, its sign changes depending on the specification.

Finally, the LIQ indicator has a negative sign in all the specifications and it is significant in some of them. The interpretation would be that Italian bonds tend to trade at a premium when investors value liquidity more highly. This might seem counterintuitive. However, take into consideration the fact that Italian sovereign bonds can be used by banks in repo transactions, not only with other banks, but also with the European Central Bank. Thus, for a category of investors (banks), these bonds can be temporarily turned into liquidity with a repo at no cost even when liquidation on the spot market becomes expensive. The premium associated to this possibility (usually known as convenience yield) might have been higher exactly when investors attached a higher value to liquidity, that is, when LIQ was higher.

This hypothesis is further discussed and tested in Section 4.4.

In the last column of Table 2 we report the standard errors of the regressions. These are on average around 0.30, which means that the part of the sovereign spread which is not explained by our variables is on average about 30 basis points. In the majority of our regressions, adding the liquidity variables reduces the standard error by less than 2 basis points.

As a robustness check, we also add to our regressions interaction terms obtained by multiplying our proxy for illiquidity (the bid-ask spread) by the proxies for the price of liquidity (LIQ and the KfW-Bund spread). We find (the results are not reported in the paper) that these interaction terms are seldom statistically significant and do not increase the explanatory power of the liquidity variables.

Overall, the results from our regressions suggest that proxies for the credit risk premium explain a predominant portion of the variability of the Italian sovereign spread, while the liquidity variables have little additional explanatory power. However, the liquidity variables are in several cases statistically significant and they can become economically relevant during short episodes of illiquidity.

4.2 Results for Spain

Results for Spain are reported in Table 3. They are qualitatively similar to those obtained for Italy.

The two proxies for the credit risk premium explain a preponderant portion of the variability of the sovereign spread, and the HW loss rate provides a better fit to the data than the CDS spread.

Furthermore, the bid-ask spread has a positive and statistically significant coefficient in the majority of specifications. The Kfw-Bund spread is significant only in one case, and in that case it has a positive sign, differently from all the other specifications, where it is not significant but it has a negative sign. The LIQ indicator has a negative sign in all the specifications and it is significant in most of them.

The R^2 of the regressions is high (up to 93%), but lower than that found for Italy. In other words, a larger portion of sovereign spreads remains unexplained.

4.3 Results for Germany

Results for Germany are reported in Table 4. Differently from what we find for Italy and Spain, the R^2 of the regressions is not high (up to 16%). In particular, the proxies for credit premia are not significant in the majority of specifications and provide limited explanatory power. This reinforces the evidence, already provided by other studies (e.g., Di Cesare et al. 2012), of a disconnect between German CDSs and government bonds. The bid-ask spread is instead highly significant also for Germany, and it provides some explanatory power. The other liquidity variables have little explanatory power and the inferences about their coefficients are not robust across models.

4.4 Alternative specifications: convenience yields and CDS liquidity premia

This section deals with two issues that have been left aside in the baseline specification: convenience yields and CDS liquidity.

If a bond can be used in repo transactions or, more generally, as collateral in secured financial transactions, the holders of the bonds might be willing to accept a reduction in its yield with respect to a similar bond that cannot be used as collateral. This reduction, which we call convenience yield, represents the compensation for the economic value of pledgeability. For example, pledgeability is valuable to banks because they can use the pledgeable bond to borrow funds on the repo market at an interest rate that is usually lower than the rate they pay on the unsecured interbank lending market. To our knowledge, there are no standard methodologies to estimate the convenience yield on a given bond, although there are several papers that deal with the differences in convenience yield induced by repo specialness (e.g., Buraschi and Menini 2002, Cherian, Jacquier and Jarrow 2004). Also, there seems to be no unanimous agreement on a definition of convenience yield, and some papers adopt a definition that is much broader than ours. For example, Feldhutter and Lando (2008) include in their definition of the convenience yield of US Treasuries also liquidity components that would be considered part of the liquidity premium in our conceptual framework (see Section 2.4). Given the lack of a standard definition (and, as a consequence, of standard estimation methodologies), we have not included proxies for the convenience yield in our baseline specifications. Nonetheless, we believe that attempting to address this issue is important, also because it helps to shed some light on the findings of the previous sections.

We propose to use the 3-month Euribor-Eurepo spread (see, e.g., Cassola, Hortaçsu and Kastl 2013) as a proxy for the convenience yield. Euribor is a benchmark rate used to gauge the cost of unsecured interbank borrowing while Eurepo is an indicator of the cost of secured borrowing through repos. Thus, the spread provides a measure of how much a bank can save on an interbank loan if it pledges government bonds as collateral. Of course, this is an imperfect measure, as the convenience yield of a specific bond can depend on several factors, among which the haircut that is applied to the bond and its specialness (or its probability of going on special in the future). Furthermore, the actual savings obtained from secured borrowing can be greater than indicated by the Euribor-Eurepo spread if repos with the central bank are cheaper than repos on the interbank market (e.g., under the fixed-rate full-allotment policy implemented by the ECB since 2008). Finally, given there is no readily available indicator of the cost of unsecured borrowing on the maturity we are studying in our regressions (10-year), we choose the 3-month maturity because it is widely considered the most significant one in terms of market impact (e.g., De Socio 2013).

Table 5 reports the estimates of some of the regression models presented in the previous sections, augmented by including the Euribor-Eurepo spread among the regressors. The results are clear cut: the estimated coefficient is always significant and has the expected sign (the higher the convenience, the lower the bond yield). Furthermore, when the proxy for the convenience yield is included, the bid-ask spread remains significant in most specifications, but the LIQ variable becomes insignificant. This supports our previous conjecture that the negative and significant coefficient on LIQ found in Sections 4.1 and 4.2 could be the result of omitting a measure of the convenience yield from the regressions. In other words, our previous finding that Italian and Spanish bonds tend to trade at a premium when investors value liquidity more highly is not due to their intrinsic liquidity (how expensive it is to trade in these bonds), but to the fact that these bonds are easily pledgeable and can be used by banks in repo transactions, so that they can be temporarily turned into liquidity with a repo at no cost even when liquidation on the spot market becomes expensive. The premium associated to this possibility of temporarily turning bonds into liquidity (the convenience yield) might have been higher exactly when investors attached a higher value to liquidity, that is, when LIQ (a measure of investors' appetite for liquidity) was higher.

As far as the economic significance of the convenience yields is concerned, we find that it is not negligible and it is similar across countries: depending on the model used, convenience yields account on average for between 16 and 26 basis points of the sovereign spreads[12].

We also propose a battery of regressions (Table 5) where we try to control, albeit in an admittedly crude fashion, for potential biases introduced by CDS liquidity premia (see Section 2.4 for a discussion). We conjecture that the part of the German sovereign spread not explained by the credit and liquidity proxies is entirely due to the CDS liquidity premia that enter the regressions through the proxies of credit premia (which are calculated from CDS quotes). In particular, we use the residuals estimated from one of the regressions in Table 4

[12] Bear in mind that the effect is negative, that is, convenience yields reduce spreads.

(the last of the third block; it includes bid/ask and LIQ, and it has the coefficient on HW constrained to 1) as a proxy for the CDS liquidity premium, and we add it as an explanatory variable to the regressions for Italy and Spain. The rationale for this procedure is provided by the evidence of the presence of common factors in CDS liquidity premia (Badaoui, Cathcart and El-Jahel 2013). In most of our regressions, the coefficient on the CDS liquidity variable is statistically significant and positive, which means that the bias introduced by CDS liquidity premia has the same sign for all three countries. We also find that the coefficients on all the other variables do not change significantly when the proxy for the CDS liquidity premium is included among the regressors. This result provides some evidence that our results are not affected by CDS liquidity premia (although, as already stated, the proposed way of controlling for CDS liquidity premia is quite rough).

5 Vector autoregressions

In this section, we use vector autoregression (VAR) models to study aspects, such as the dynamic interactions among our variables, that cannot be captured in the linear regression framework used in the previous section. In particular, we use Bayesian VAR models to analyze how shocks to credit and liquidity premia impact – in terms of impulse response functions and variance decompositions – on sovereign spreads. Furthermore, we employ an agnostic identification strategy that allows to compute upper and lower bounds on the relative importance of the different premia. This might help to address possible concerns about the interpretation of the regressions in the previous sections. As a matter of fact, one could argue that gauging the importance of the different premia only from the estimates of their regression coefficients could be misleading in case some of the regressors were significantly correlated. We address this kind of concern in what follows, by showing that, even when we consider identification schemes in which the covariance between liquidity premia and credit premia can be entirely attributed to liquidity premia, the latter explain only a small fraction of the variance of sovereign spreads.

5.1 The VAR model

We use VARs to model the dynamic interactions among four variables: the sovereign spread, the HW loss rate, the bid-ask spread and the LIQ indicator. The choice of the proxies for credit and liquidity risk is based on the evidence presented in the previous section that they provide the best performance in explaining sovereign spreads.

By performing a preliminary analysis based on OLS estimates, we find that first-order models are preferred to higher-order ones according to the Bayesian Information Criterion (BIC). Therefore, we restrict our attention to first-order models

$$y_t = \mu + \rho y_{t-1} + \varepsilon_t$$

where y_t is a $K \times 1$ vector of variables, μ is a $K \times 1$ vector of constants, ρ is a $K \times K$ matrix of constants and ε_t is a $K \times 1$ vector of zero-mean multivariate normal error terms.

5.2 Estimation technique

We choose to estimate the VARs with Bayesian techniques. We prefer the latter to standard frequentist techniques for two reasons. First, we want to take parameter uncertainty into account in our impulse response analysis and in our variance decompositions, and this can be transparently accomplished with Bayesian estimation. Second, we want to rule out parameters that give rise to explosive trajectories for y_t. This can also easily be done in a Bayesian setting, by assigning zero prior probability to such parameters (more on this below).

Denote by Σ the Choleski factor of the covariance matrix of ε_t. In order to keep the analysis objective, the following uninformative and mutually independent priors are assigned to the parameters:

- μ: each entry is assigned a uniform improper prior on $(-\infty, \infty)$;
- ρ: uniform improper prior on the set

$$\left\{ \rho \in \mathbb{R}^{K \times K} : \sup_{\lambda \in \mathrm{eig}(\rho)} |\lambda| \leq 1 \right\}$$

  where $\mathrm{eig}(\rho)$ is the set of eigenvalues of ρ;
- diagonal entries of Σ: uniform improper prior on $[0, \infty)$;
- off-diagonal non-zero entries of Σ: uniform improper prior on $(-\infty, \infty)$.

An independence-chain Metropolis-Hastings algorithm is employed to simulate from the posterior distribution of the parameters. Denote by θ the vector obtained by stacking all the non-zero entries of μ, ρ and Σ. As a proposal distribution for θ, we use a multivariate normal distribution obtained from an ancillary estimation[13]. The chain is run for 105,000 iterations and the the first 5,000 draws are discarded. The lowest acceptance rate is 35 per cent (for Spain). For all countries, Raftery and Lewis' (1995) run length control diagnostic[14] indicates that the sample size is more than 200 times the minimum required size.

We use an agnostic procedure to identify the shocks. For each posterior draw of θ, we also draw a random recursive Choleski ordering of the shocks, and we use it to compute an impulse response function and a variance decomposition. Thus, we obtain distributions of impulse responses and variance decompositions that take into account the uncertainty on both the parameters and the identification of the shocks. In particular, we assign a uniform

[13]The proposal distribution is a multivariate normal distribution, whose mean is equal to the maximum likelihood estimate of θ obtained without imposing constraints on ρ, and whose covariance matrix is equal to the estimated asymptotic covariance matrix of the maximum likelihood estimator.

[14]The parameters of the diagnostic are set in such a way that the minimum required size allows to estimate the 2.5% quantile with an error <1% with probability 95%.

discrete distribution to all permutations of the variables where the sovereign spread occupies the last position. By assigning the last position to the innovation in the sovereign spread, we allow all the shocks that affect credit and liquidity premia to also have a contemporaneous effect on the sovereign spread. In this way, the shock to the sovereign spread has a residual role: it captures the innovations to the spread that cannot be explained by the shocks to our proxies for credit and liquidity premia.

Assigning equal probability to all possible orderings of the variables was originally proposed by Lindeman, Merenda, and Gold (1980) for estimating relative importance in regression models, and has subsequently become popular for carrying out variance decompositions (see Grömping 2007, for a review). For applications to finance of this technique, see, for example, Diebold and Yilmaz (2009) and the references in Klößner and Wagner (2013).

We also explore agnostic identification strategies based on sign restrictions (see Uhlig 2005). In particular, we use the multivariate technique based on random draws of orthonormal rotation matrices proposed by Lippi and Nobili (2012). However, the few sign restrictions we can safely impose[15] do not lead to any meaningful identification of the shocks: virtually all random rotations satisfy the restrictions and, as a result, the variance of the sovereign spread is estimated to be spread equally among the four shocks. The results from this supplementary analysis are not reported in the paper.

5.3 Results for Italy

Table 6 reports the estimated decomposition of the variance of the sovereign spread, as well as the estimated response of the sovereign spread to one-standard-deviation shocks. For both objects, we report the mean and the first and ninth deciles.

At a 4-week forecast horizon, shocks to the HW loss rate explain about 70% of the

[15] We impose that positive shocks to the credit risk premium and to the bid-ask spread have a positive impact on the sovereign spread at all horizons.

variability of the spread, while the remaining variability is explained almost entirely by residual shocks specific to the spread (i.e., shocks that affect only the spread). Shocks to the bid-ask spread and the LIQ indicator explain about 2% of the variability. The impact of these shocks has the same sign found in the regression analysis, that is, positive for the bid-ask spread and negative for LIQ, but it is not significantly different from zero.

By increasing the forecast horizon, the proportion of variance explained by the HW loss rate increases (up to 84% at a 52-week horizon), while that explained by shocks that are specific to the sovereign spread decreases. The proportion explained by the bid-ask spread and the LIQ indicator increases somewhat (up to 9% cumulatively), but the response to these variables remains not significantly different from zero.

The differences between the bounds and the central estimates of the variance decomposition are seldom greater than 10 percentage points in absolute value. Overall, even if boundary values were considered, the main qualitative deductions from the variance decomposition would not be affected. In particular, the bounds on the importance of the liquidity variables are quite tight. The interpretation is as follows: irrespective of the identification scheme that is randomly drawn, the relative importance of the liquidity variables remains limited.

In summary, the insights from the VAR analysis are similar to those already gained from the regression analysis: the proxy for the credit premium explains the great majority of the variability in the sovereign spread, while the liquidity proxies explain a small fraction of variance and are often not statistically significant.

5.4 Results for Spain

Results for Spain are analogous to those obtained for Italy. The only relevant differences are that the proportion of variance explained by shocks to the HW loss rate is smaller at short horizons (but comparable at long horizons), and shocks to the bid-ask spread explain a greater portion of the variability in the sovereign spread. Furthermore, the impact of the shocks to the bid-ask spread is significantly different from zero at short horizons.

Also in the case of Spain, insights from the VAR analysis are similar to those provided by the regression analysis.

5.5 Results for Germany

A predominant portion (91% and 76% at the 4-week and 52-week horizons respectively) of the variability in the German sovereign spread is not explained by shocks to the proxies for credit and liquidity premia. The response to shocks to the HW loss rate is estimated to be zero at all forecasting horizons (with tight confidence bounds). This provides further evidence of the disconnect between the German CDS and sovereign spreads already found in previous sections. Furthermore, the impacts of the shocks to the bid-ask spread and to the LIQ indicator are not significantly different from zero, except for the impact of the bid-ask spread at a 4-week horizon, which is positive and significant.

Overall, the results from the VAR analysis seem to confirm those from the regression analysis: all of the proxies for credit and liquidity premia provide little explanatory power for the German spread.

6 Conclusions

We have conducted an empirical analysis of sovereign bond spreads and their two main components, namely, credit and liquidity premia. We have discussed and analyzed various proxies for these components. We have provided evidence that raw CDS spreads might be inadequate for measuring the credit premia and we have proposed other proxies, also derived from CDS quotes, that provide a better fit to sovereign spreads in our sample. These proxies for the credit premia explain a predominant portion of the variability of sovereign spreads, while the proxies for the liquidity premia have little additional explanatory power, although they are in several cases statistically significant. As far as economic significance is concerned, the magnitude of liquidity effects is found to be on average small (across models and time), even if, according to some model specifications, these effects were sporadically large during part of our sample. In particular, marked increases in the illiquidity of sovereign bonds (measured by bid-ask spreads) might have contributed to exacerbate the spike in sovereign spreads observed in the second half of 2011. We have also provided evidence that increases in investors' appetite for liquidity (measured by various liquidity spreads) might have had a dampening effect on sovereign spreads. While this effect could seem counterintuitive, we argue (and we provide empirical evidence in favor of our argument) that it might have to do with the fact that, despite the deterioration in liquidity experienced on the secondary market for sovereign bonds, these bonds continued to be highly liquid assets for banks because they could be pledged as collateral in repo transactions (not only with other banks but also with the central bank) and could be temporarily turned into liquidity at no cost even when liquidation on the spot market became expensive.

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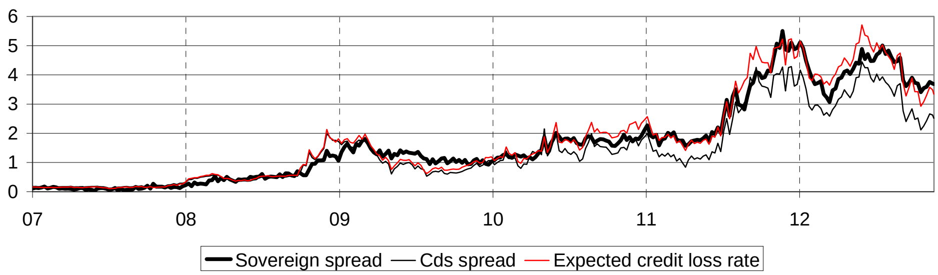
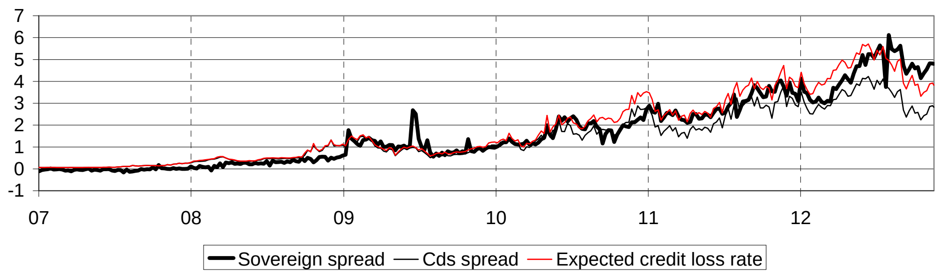
Figure 1 – Sovereign spreads and proxies for credit risk

Panel A - Italy



Panel B - Spain



Panel C - Germany

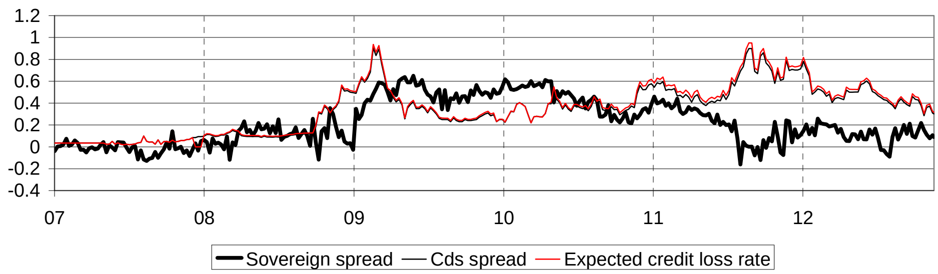
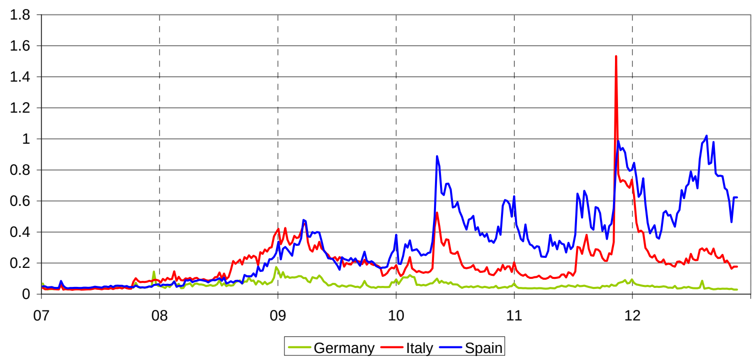


Figure 2 – Measures of liquidity and liquidity risk premia

Panel A – Bid-ask spreads



Panel B – Observable liquidity risk premia

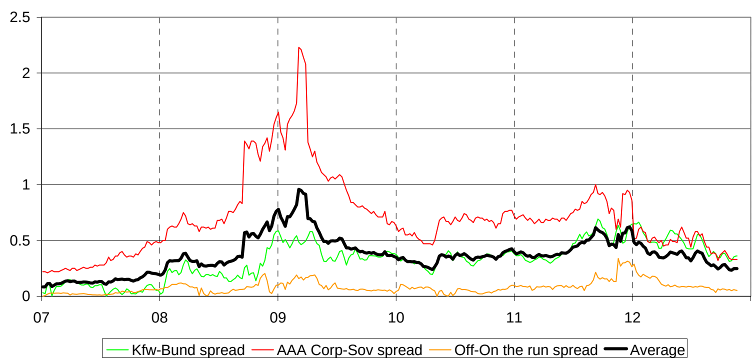


Table 1 - Descriptive statistics[16]

	Mean	St. dev.	Min	q1	q2	q3	Max
OIS	3.13	0.99	1.18	2.39	3.10	4.15	4.86
y Ger	3.36	0.97	1.25	2.59	3.53	4.19	4.86
y Ita	4.85	0.71	3.87	4.34	4.67	5.05	7.64
y Spa	4.81	0.79	3.49	4.20	4.52	5.47	7.60
y-OIS Ger	0.22	0.21	-0.16	0.04	0.19	0.41	0.65
y-OIS Ita	1.72	1.46	0.03	0.51	1.33	2.04	5.50
y-OIS Spa	1.68	1.58	-0.16	0.28	1.20	2.68	6.12
CDS Ger	0.34	0.22	0.02	0.11	0.35	0.48	0.90
CDS Ita	1.45	1.18	0.12	0.50	1.15	1.97	4.45
CDS Spa	1.45	1.15	0.06	0.47	1.15	2.38	4.22
HW Ger	0.35	0.23	0.01	0.12	0.37	0.51	0.95
HW Ita	1.79	1.54	0.11	0.53	1.40	2.37	5.71
HW Spa	1.85	1.58	0.06	0.50	1.28	3.16	5.71
b/a Ger	0.06	0.02	0.01	0.04	0.05	0.07	0.17
b/a Ita	0.20	0.15	0.03	0.10	0.18	0.25	1.53
b/a Spa	0.33	0.25	0.04	0.09	0.28	0.48	1.02
KfW	0.33	0.16	0.00	0.20	0.35	0.43	0.69
AAA	0.70	0.35	0.21	0.48	0.66	0.78	2.23
off-on	0.07	0.11	-0.03	0.00	0.05	0.10	0.45
LIQ	0.40	0.20	0.07	0.30	0.36	0.45	1.01

[16] q1, q2 and q3 are the first, second and third quartiles of the empirical distribution of the variables reported in the rows. y is the zero-coupon 10-year yield. OIS is the 10-year OIS rate. b/a is the bid-ask spread on 10-year sovereign bonds. HW is the estimate of the risk-neutral expected credit loss rate. KfW, AAA, off-on and LIQ are the liquidity indicators proposed in subsection 2.4. Ger, Ita and Spa are abbreviations for Germany, Italy and Spain respectively.

Table 2 – Baseline regressions – Italy[17]

c	CDS	HW	b/a	KfW	LIQ	R^2	s.e.
-0.03	1.21 ***					0.95	0.33
-0.03	1.21 ***		-0.01			0.95	0.33
-0.05	1.19 ***			0.13		0.95	0.33
-0.05	1.19 ***		-0.05	0.16		0.95	0.33
0.15	1.23 ***				-0.57	0.95	0.32
0.16	1.20 ***		0.51 **		-0.75 **	0.95	0.32
0.03		0.94 ***				0.97	0.25
-0.02		0.91 ***	0.52 *			0.97	0.24
-0.00		0.92 ***		0.22		0.97	0.25
-0.01		0.91 ***	0.54 **	-0.06		0.97	0.24
0.08		0.94 ***			-0.14	0.97	0.25
0.10		0.90 ***	0.83 ***		-0.45 *	0.97	0.24
0.27 ***	1 cn					0.92	0.41
0.08	1 cn		0.98 **			0.93	0.39
-0.13 *	1 cn			1.24 ***		0.94	0.36
-0.13 *	1 cn		0.17	1.14 ***		0.94	0.36
0.32	1 cn				-0.14	0.92	0.41
0.28	1 cn		1.51 ***		-0.84 *	0.94	0.37
-0.08 *		1 cn				0.97	0.27
-0.08		1 cn	-0.00			0.97	0.27
0.04		1 cn		-0.37 **		0.97	0.26
0.05		1 cn	0.46 **	-0.65 ***		0.97	0.26
0.02		1 cn			-0.25	0.97	0.27
0.01		1 cn	0.22 *		-0.35	0.97	0.26

[17] The dependent variable is y-OIS. y is the zero-coupon 10-year yield. OIS is the 10-year OIS rate. b/a is the bid-ask spread on 10-year sovereign bonds. HW is the estimate of the risk-neutral expected credit loss rate. KfW and LIQ are two of the liquidity indicators proposed in subsection 2.4. s.e. is the standard error of the regressions. When a coefficient value is constrained, it is followed by cn. The first column reports the estimates of the intercepts. *, ** and *** denote significance at the 10, 5 and 1 per cent levels respectively.

Table 3 – Baseline regressions – Spain[18]

c	CDS	HW	b/a	KfW	LIQ	R^2	s.e.
-0.23 ***	1.31 ***					0.91	0.49
-0.29 ***	1.10 ***		1.13			0.91	0.47
-0.10	1.40 ***			-0.76		0.91	0.48
-0.12	1.18 ***		1.29 *	-1.02		0.92	0.45
-0.00	1.33 ***				-0.70	0.91	0.47
-0.04	1.10 ***		1.26 *		-0.79 **	0.92	0.45
-0.11		0.96 ***				0.93	0.42
-0.19 **		0.82 ***	1.05 **			0.94	0.39
-0.08		0.98 ***		-0.17		0.93	0.42
-0.10		0.84 ***	1.19 **	-0.54		0.94	0.39
-0.06		0.97 ***			-0.14	0.93	0.42
-0.09		0.81 ***	1.15 **		-0.32	0.94	0.39
0.22	1 cn					0.86	0.60
-0.27 ***	1 cn		1.52 ***			0.91	0.47
-0.22 **	1 cn			1.35 ***		0.87	0.56
-0.16 *	1 cn		1.84 **	-0.67		0.91	0.46
0.32	1 cn				-0.25	0.86	0.60
-0.02	1 cn		1.66 ***		-0.79 **	0.92	0.45
-0.18 **		1 cn				0.93	0.42
-0.20 **		1 cn	0.08			0.93	0.42
-0.07		1 cn		-0.33		0.93	0.42
-0.05		1 cn	0.50	-0.88		0.93	0.41
-0.11		1 cn			-0.18	0.93	0.42
-0.13 *		1 cn	0.11		-0.22	0.93	0.42

[18] The dependent variable is y-OIS. y is the zero-coupon 10-year yield. OIS is the 10-year OIS rate. b/a is the bid-ask spread on 10-year sovereign bonds. HW is the estimate of the risk-neutral expected credit loss rate. KfW and LIQ are two of the liquidity indicators proposed in subsection 2.4. s.e. is the standard error of the regressions. When a coefficient value is constrained, it is followed by cn. The first column reports the estimates of the intercepts. *, ** and *** denote significance at the 10, 5 and 1 per cent levels respectively.

Table 4 – Baseline regressions – Germany[19]

c	CDS	HW	b/a	KfW	LIQ	R^2	s.e.
0.14	0.25					0.07	0.20
0.01	0.19		2.60 **			0.16	0.19
0.12	0.13			0.18		0.07	0.20
-0.00	0.11		2.57 **	0.13		0.16	0.19
0.04	-0.01				0.49 *	0.17	0.19
0.00	0.04		1.33		0.35	0.19	0.19
0.14		0.23				0.07	0.20
0.01		0.18	2.63 **			0.16	0.19
0.12		0.11		0.19		0.07	0.20
-0.00		0.11	2.59 **	0.12		0.16	0.19
0.04		-0.01			0.49 *	0.17	0.19
0.00		0.04	1.35		0.34	0.19	0.19
-0.12 *	1 cn					Neg.	0.26
-0.18 *	1 cn		1.19			Neg.	0.26
0.15	1 cn			-0.83 ***		Neg.	0.22
0.04	1 cn		2.40 **	-0.90 ***		Neg.	0.21
-0.01	1 cn				-0.28	Neg.	0.26
-0.11	1 cn		3.83 **		-0.60 **	Neg.	0.24
-0.13 *		1 cn				Neg.	0.27
-0.20 *		1 cn	1.29			Neg.	0.27
0.16		1 cn		-0.89 ***		Neg.	0.23
0.04		1 cn	2.59 **	-0.97 ***		Neg.	0.22
-0.01		1 cn			-0.32	Neg.	0.26
-0.13 *		1 cn	4.23 **		-0.67 **	Neg.	0.25

[19] The dependent variable is y-OIS. y is the zero-coupon 10-year yield. OIS is the 10-year OIS rate. b/a is the bid-ask spread on 10-year sovereign bonds. HW is the estimate of the risk-neutral expected credit loss rate. KfW and LIQ are two of the liquidity indicators proposed in subsection 2.4. s.e. is the standard error of the regressions. When a coefficient value is constrained, it is followed by a cn. The first column reports the estimates of the intercepts. *, ** and *** denote significance at the 10, 5 and 1 per cent levels respectively.

Table 5 – Augmented regressions[20]

Panel A - Italy

c	CDS	HW	b/a	LIQ	CONV	cds-LIQ	R^2	s.e.
0.19	1.21 ***		0.73 **	-0.37	-0.42 ***		0.96	0.29
0.12		0.91 ***	1.01 ***	-0.17	-0.31 ***		0.98	0.22
0.31 *	1 cn		1.74 ***	-0.50	-0.38 **		0.94	0.35
0.03		1 cn	0.41 ***	-0.07	-0.32 **		0.97	0.25
0.18 *	1.29 ***		0.62 **	-0.29	-0.51 ***	0.57 ***	0.96	0.28
0.11 **		0.96 ***	0.94 ***	-0.10	-0.38 ***	0.51 ***	0.98	0.20
0.29	1 cn		1.61 ***	-0.51	-0.35 *	-0.24	0.95	0.35
0.08		1 cn	0.77 ***	-0.05	-0.40 ***	0.66 ***	0.98	0.21

Panel B - Spain

c	CDS	HW	b/a	LIQ	CONV	cds-LIQ	R^2	s.e.
-0.00	1.14 ***		1.09 *	-0.33	-0.40 ***		0.93	0.43
-0.05		0.83 ***	1.03 **	0.11	-0.35 ***		0.94	0.38
0.01	1 cn		1.64 ***	-0.38	-0.35 **		0.93	0.44
-0.09		1 cn	0.09	0.28	-0.43 ***		0.94	0.40
-0.01	1.19 ***		1.04	-0.29	-0.43 ***	0.34	0.93	0.43
-0.07		0.87 ***	0.99 **	0.16	-0.39 ***	0.35 **	0.94	0.37
0.01	1 cn		1.69 ***	-0.38	-0.35 **	0.10	0.93	0.44
-0.10		1 cn	0.37	0.30	-0.45 ***	0.58 ***	0.94	0.38

[20] The dependent variable is y-OIS. y is the zero-coupon 10-year yield. OIS is the 10-year OIS rate. b/a is the bid-ask spread on 10-year sovereign bonds. HW is the estimate of the risk-neutral expected credit loss rate. KfW and LIQ are two of the liquidity indicators proposed in subsection 2.4. CONV and cds-LIQ are the proxies for the convenience yield and the CDS liquidity premium proposed in Section 4.4. s.e. is the standard error of the regressions. When a coefficient value is constrained, it is followed by cn. The first column reports the estimates of the intercepts. *, ** and *** denote significance at the 10, 5 and 1 per cent levels respectively.

Table 6 – Agnostic VAR[21]

	Variance decomposition				Impulse response			
Italy								
	y-OIS	HW	b/a	LIQ	y-OIS	HW	b/a	LIQ
4-week	29 (23,35)	69 (62,74)	2 (1,4)	0 (0,1)	8 (6,9)	16 (14,17)	1 (-1,3)	-1 (-2,0)
13-week	15 (8,23)	83 (74,90)	2 (0,4)	1 (0,2)	3 (0,6)	17 (14,20)	0 (-3,3)	2 (-1,4)
26-week	10 (4,18)	86 (77,93)	2 (0,4)	2 (0,7)	1 (-2,5)	15 (12,19)	0 (-3,3)	3 (-1,7)
52-week	8 (3,14)	84 (70,94)	2 (0,5)	7 (0,18)	0 (-3,4)	11 (6,17)	1 (-2,3)	4 (-1,10)
Spain								
	y-OIS	HW	b/a	LIQ	y-OIS	HW	b/a	LIQ
4-week	56 (49,63)	40 (33,47)	3 (1,7)	1 (0,2)	10 (7,13)	18 (16,19)	4 (1,7)	-1 (-3,0)
13-week	25 (18,32)	71 (62,79)	4 (1,9)	1 (0,2)	1 (-2,5)	20 (17,23)	1 (-4,6)	0 (-3,2)
26-week	15 (10,20)	80 (72,87)	4 (1,8)	1 (0,3)	1 (-3,4)	18 (14,23)	-1 (-6,4)	0 (-4,5)
52-week	10 (6,15)	82 (72,90)	4 (1,10)	4 (0,9)	1 (-2,3)	13 (7,20)	-1 (-5,3)	1 (-5,8)
Germany								
	y-OIS	HW	b/a	LIQ	y-OIS	HW	b/a	LIQ
4-week	91 (86,95)	1 (0,2)	5 (1,9)	4 (1,6)	6 (5,6)	0 (-1,0)	2 (1,3)	-1 (-1,0)
13-week	87 (80,94)	2 (0,4)	8 (2,15)	3 (1,5)	2 (1,3)	0 (-1,1)	1 (0,2)	0 (0,1)
26-week	83 (73,91)	3 (0,7)	8 (2,15)	7 (2,12)	0 (0,1)	0 (-1,1)	0 (0,1)	1 (0,2)
52-week	76 (63,87)	4 (1,10)	7 (2,14)	12 (4,23)	0 (-1,0)	0 (0,1)	0 (0,0)	1 (0,1)

[21] y is the zero-coupon 10-year yield. OIS is the 10-year OIS rate. b/a is the bid-ask spread on 10-year sovereign bonds. HW is the estimate of the risk-neutral expected credit loss rate. LIQ is the liquidity indicator proposed in subsection 2.4. Figures in parentheses are the first and ninth decile of the distributions of variance decompositions and impulse response functions. Forecasting horizons are reported in the leftmost column.

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