Decomposing euro-area sovereign spreads: credit and liquidity risks*

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Abstract

This paper presents an arbitrage-free model of the joint dynamics of euro-area sovereign bond spreads. The latter reflect both credit and liquidity differences between national bonds. An innovative aspect of the approach lies in the modeling of intertwined credit- and liquidity-related crisis regimes. We find that a substantial share of the changes in euro-area yield differentials is liquidity-driven during the financial-crisis period. Once liquidity-pricing effects and risk premiums are filtered out of the spreads, we obtain estimates of the actual default probabilities. They are significantly lower than their risk-neutral counterparts, which is consistent with the existence of a non-diversifiable euro-area sovereign credit risk.

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Keywords: default risk, liquidity risk, term structure of interest rates, regime switching, euro-area spreads.

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1 Introduction

One of the most spectacular symptoms of the crisis that began in mid-2007 is the dramatic rise in intra euro-area government-bond yield spreads. Whereas all euro-area sovereign 10-year bond yields were contained in a range of 50 bp between 2002 and 2007, the average spreads over Germany of only two countries were lower than 50 basis points in 2011, the debt-weighted mean being of about 250 bp. Since the inception of the euro in 1999 and the resulting elimination of exchange-rate risk, intra-euro-area spreads reflect the fluctuations of compensations demanded by investors for holding essentially two kinds of risks: credit and liquidity risks.[1] The credit risk is linked to the issuer's probability of default (PD). If investors assess that the PD of some indebted country is higher than in the past, the prices of the bonds issued by this country fall because the expected loss increases. Liquidity risk arises from the potential difficulty that one may have in selling the asset before its redemption (for instance if one is required to do so in distressed market conditions, where it is difficult to find a counterpart for trade relatively quickly). In many ways, the ongoing financial crisis has illustrated why, along with credit risk, liquidity risk matters and should not be underestimated (see Brunnermeier, 2009).

Disentangling credit and liquidity effects in bond prices is important in several respects. For instance, appropriate policy actions that may be needed to address a sharp rise in spreads depend on the source of the movement: if the rise in spreads reflects poor liquidity, policy actions should aim at improving market functioning. But if it is linked to credit concerns, the solvency of the debtors should be enhanced (see Codogno, Favero and Missale, 2003). Furthermore, optimal investment decisions would benefit from such a decomposition. In particular, those medium to long-term investors who buy bonds to hold them until redemption seek to buy bonds whose price is low because of poor liquidity, since it provides them with higher long-run returns than more liquid bonds with the same credit quality (see Longstaff, 2009).

In this paper, we develop a multi-issuer no-arbitrage affine term-structure framework to model the dynamics of bond spreads, with a twofold objective: to disentangle credit and liquidity components in euro-area sovereign spreads and to identify the part of these spreads corresponding to risk premiums, defined as the part that would not be present if agents were risk-neutral. Risk premiums

[1] Indeed, an overwhelming share of the euro-area sovereign debt is denominated in euros (see Eurostat, 2011).

are demanded by risk-averse investors to be compensated for non-diversifiable –or systematic– risk, and our results are supportive of the findings of Pan and Singleton (2008), Longstaff et al. (2011) or Borri and Verdelhan (2011) who point to the systematic nature of sovereign risk.[2] The resulting risk premiums associated with sovereign credit quality implies that physical, or real-world, probabilities of default differ from their risk-neutral counterparts. Yet, the latter, derived from basic models like in Litterman and Iben (1991), are extensively used by market practitioners, who refer to them as *implied default probabilities*.[3] Our approach makes it possible to assess the deviations between the two kinds of PDs and we show that they can be substantial. In particular, these results are of significant interest in the current context where regulators want banks to model the *actual* default risk of even high-rated government bonds.[4]

In our framework, the countries are characterised by risk intensities that incorporates both credit and liquidity components. We propose an original use of regime-switching features to account for the joint dynamics of credit- and liquidity-related crises, the aim being to make the model consistent with theoretical approaches highlighting the potential interactions between these two kinds of risks.[5] Credit- and liquidity-crisis regimes are key drivers of default and illiquidity intensities, these processes being also affected by Gaussian shocks. In this framework, the spreads are combinations of the regime variables and of latent factors that follow Gaussian auto-regressive processes whose drifts depend on the regimes.

Most of the studies involving intensity modeling only consider credit risk, the underlying expected losses being associated with the default of the bond issuer. A few emprirical studies (reviewed in Section 2) make use of liquidity-related intensities but are relatively silent about their interpretation. A structural interpretation of the liquidity components of the countries' intensities is provided by Ericsson and Renault (2006) and He and Xiong (2012). In these models, the bond-

[2] Borri and Verdelhan (2011) propose a theoretical framework to investigate the implications of the investors' inability to hedge against correlated sovereign risks. The implications of their model are supported by an empirical study on emerging-country data.

[3] See e.g. Hull, Predescu and White (2005), Berd, Mashal and Wang (2003), Caceres, Guzzo and Segoviano (2010) or Berg (2009).

[4] In early 2012, the European Union introduced new rules on trading-book capital, known as Basel 2.5. This package notably requires the banks to model the default risk of all sovereign entities for the first time. This contrasts with the special status that government bonds have enjoyed since the Basel Committee for Banking Supervision (BCBS) first proposed rules on the capital treatment of market risks in 1993. As stressed by Carver (*Risk Magazine*, 2012), these changes in regulation reveal the practitioners' lack of tools to extract actual default probabilities from market prices.

[5] See e.g. Brunnermeier and Pedersen (2009) or Garleanu and Pedersen (2007).

holder can be hit by liquidity shocks. Upon the arrival of these shocks, she has to liquidate her holdings at a fractional cost. Ericsson and Renault (2006) show that these costs can be related to the ease with which one can find a buyer for the considered bonds when hit by the liquidity shock. Therefore, this modeling involves both market-liquidity aspects (the bonds differ in terms of the ease with which they can be sold, i.e. their market liquidity) and liquidity-funding aspects (the liquidity shock can be linked to bondholder's funding difficulties to cope with a liquidity shortage). Fluctuations in the probability of being hit by the liquidity shock generates comovement across the liquidity components of countries' intensities. In this paper, we assume that the countries' illiquidity intensities are driven by a single European liquidity-related factor. The identification of this factor is based on the exploitation of the term structure of the spreads between KfW (*Kreditanstalt für Wiederaufbau*), a German agency, and the *Bunds*, which are the bonds issued by the Federal Republic of Germany. Indeed, the bonds issued by KfW, guaranteed by the Federal Republic of Germany, benefit from the same credit quality than the *Bunds* but are less liquid.[6] Therefore, the KfW-*Bund* spread should be essentially liquidity-driven (see Schwarz, 2009). The results of our analysis indicate that liquidity-pricing effects significantly contribute to the dynamics of intra-euro spreads, supporting findings by Favero et al. (2010) or Manganelli and Wolswijk (2009).

The model is estimated on weekly data covering about six years (2006-2013). These data consist of sovereign-bond yields associated with eight euro-area countries. Our estimation dataset is supplemented with survey-based forecasts. As evidenced by Kim and Orphanides (2012), this alleviates the small-sample bias in the persistence of the yields obtained with conventional estimation.[7] These biases alter long-horizon expectations of yields (under the historical measure) and, at the same time, term-premium estimates. Generating reliable expectations is crucial given our present goal of recovering long-term historical –or actual, or real-world– probabilities of default from bond prices.

Our study contributes to the term-structure modeling literature in three main directions. First, we develop a regime-switching affine term-structure model –RS-ATSM hereinafter– that explicitly

[6] By abuse of language, we use here the term *Bunds* for the German sovereign bonds of any maturity although this name is usually used for ten-year bonds only.

[7] This way of reducing the bias is not the only one. In particular, Jardet, Monfort and Pegoraro (2009) use a "near-cointegrated framework" specification of the factors (averaging a stationary and a cointegrated specification).

incorporates liquidity and default risks in a multi-country set up.[8] This model relies on an original modeling of credit and liquidity crises based on interrelated switching regimes. Second, we bring this model to European data, shedding light on the fluctuations of intra-euro-area sovereign spreads over the last five years. Third, we investigate the potential of this RS-ATSM to generate term structures of PDs.

The remaining of this paper is organised as follows. Section 2 reviews related literature. Section 3 presents the data and highlights stylised facts that are going to be exploited in our modeling framework. The model is developed in Section 4. Section 5 presents the estimation of the model and Section 6 examines its implications in terms of liquidity and credit pricing. Section 7 summarises the results and makes concluding remarks.

2 Related literature

There is compelling evidence that yields and spreads are affected by liquidity concerns.[9] In particular, using euro-area data, Beber, Brandt and Kavajecz (2009) provide evidence of a nontrivial role in the dynamics of sovereign bond spreads, especially for low credit risk countries and during times of heightened market uncertainty.[10] In recent studies, some authors develop ATSM to breakdown several kinds of spreads into different components, including liquidity-related ones. These approaches are based on the assumption that there exists commonality amongst the liquidity components of prices of different bonds.[11] For instance, Liu, Longstaff and Mandell (2006) use a five-factor affine framework to jointly model Treasury, repo and swap term structures. One of their factors is related to the pricing of the Treasury-securities liquidity and another factor reflects default risk.[12] Feldhütter and Lando (2008) develop a six-factor model for Treasury bonds, corporate

[8] Geyer, Kossmeier and Pichler (2004) present a multi-country ATSM. However, their model does not explicitly accommodate liquidity-pricing effects.

[9] See, e.g., Longstaff (2004), Landschoot (2004), Chen, Lesmond and Wei (2007), Covitz and Downing (2007), Kempf, Korn and Uhrig-Homburg (2012) or Acharya and Pedersen (2005).

[10] Such a behaviour is captured in a theoretical framework by Vayanos (2004)[78].

[11] See e.g. Chordia and Subrahmanyam (2000), Fontaine and Garcia (2012), Feldhütter and Lando, (2008), Longstaff, Mithal and Neis (2005), Liu, Longstaff and Mandell (2006) or Dick-Nielsen, Feldhütter and Lando (2011).

[12] As noted by Feldhütter and Lando (2008), the identification of the liquidity and credit risk factors in Liu et al. relies critically on the use of the 3-month general-collateral repo rate (GC repo) as a short-term riskfree rate and of the 3-month LIBOR as a credit-risky rate. Liu et al. define the liquidity factor as the spread between the 3-month GC repo and the 3-month Treasury-bill yield (and is therefore observable). In each yield, their liquidity component is the share of the yield that is explained by this factor.

bonds and swap rates that makes it possible to decompose swap spreads into three components: a convenience yield from holding Treasuries, a credit-element associated with the underlying LIBOR rate, and a factor specific to the swap market. They find that the convenience yield is by far the largest component of spreads. Longstaff, Mithal and Neis (2005) use information in credit default swaps –in addition to bond prices– to obtain measures of the nondefault components in corporate spreads. They find that the nondefault component is time-varying and strongly related to measures of bond-specific illiquidity as well as to macroeconomic measures of bond-market liquidity.

To the best of our knowledge, the present paper is the first to explicitly incorporate liquidity-pricing effects in a no-arbitrage multi-country set-up. Despite the importance of sovereign credit risk in the financial markets, relatively little research proposing models of the joint dynamics of sovereign yields has appeared in the literature. Notable recent exceptions include Pan and Singleton (2008), Longstaff et al. (2011) or Ang and Longstaff (2011). These contributions point to an important degree of commonality across sovereign credit risk. More precisely, they show that the risk premiums included in sovereign credit spreads are substantial and covary importantly with financial measures of global risk. According to Longstaff et al., an important source of commonality in sovereign credit spreads may be their sensitivity to the funding needs of major investors in the sovereign credit markets. This view implies that a better understanding of sovereign yield spreads requires models in which both credit and liquidity risks are explicitly taken into account. Such a model is presented in Section 4.

Our paper also extends the literature that considers the introduction of regime-switching in ATSM. This literature is based on a strong evidence of regime switching in the dynamics of interest rates (see Hamilton, 1988, Aït-Sahalia, 1996, Ang and Bekaert, 2002 or Davies, 2004 for spreads). Implied shifts in the interest-rate dynamics present a systematic risk to investors. The pricing of such a risk has already been empirically investigated within default-free ATSM incorporating Markov-switching (see Monfort and Pegoraro, 2007, Ang Bekaert and Wei, 2008 or Dai, Singleton and Yang, 2007). Building on the approaches introduced by Duffie and Singleton (1999) or Duffee (1999) to deal with credit risk in ATSM, Monfort and Renne (2013) explore the potential of Markov-switching in credit ATSM models.[13] In the present paper, we propose an original use of

[13] Whereas Duffie and Singleton (1999) and Duffee (1999) present continuous-time credit ATSMs. Gourieroux, Monfort and Polimenis, 2006 and Monfort and Renne (2013) consider discrete-time frameworks.

regime-switching features to model interactions between credit-related and liquidity-related stress periods and their impact on bond pricing.

3 Data and stylised facts

3.1 Overview

The data are weekly (end of weeks), and cover the period from 7 July 2006 to 15 February 2013 (346 dates). We consider the yield curves of eight euro-area countries: Austria, Belgium, Finland, France, Germany, Italy, the Netherlands and Spain. We exclude from the analysis those countries that were placed under EU-IMF programs during that period, namely Greece, Ireland and Portugal (in April 2010 for Greece, in November 2010 for Ireland and in May 2011 for Portugal). The choice of removing these countries from the analysis stems from the facts that (a) the three EU-IMF programs cover important shares of the total estimation period and that (b) these programs coincide with severe impairments of associated sovereign-debt markets, notably illustrated by a fall in primary-market activity.[14]

Our empirical work necessitates a riskfree rate, the choice of which is not straightforward. While many studies investigating euro-area spreads consider the German yields as being purely riskfree, this choice is at odds with the levels of the German CDS that have been observed over the past years: according to Bloomberg data, 10-year German CDS have often been above 50 basis points since 2010 onwards (with peaks that were close to 150 basis points in Summer 2012). By contrast, as argued below (in 3.2.1), German Bunds are, and have remained, highly liquid during the crisis. Hence, assuming that the liquidity-related component of the German Bunds is nil (contrary to its credit-related part), we construct proxies of the riskfree rates by subtracting German CDS from maturity-matched German bonds.[15]

[14] These impairments are illustrated by bid-ask spreads on government bonds. Based on bond prices extracted from the Thomson Reuters tick history database, the bid-ask spreads on 10-year bond issued by Greece, Ireland and Portugal were on average above 200 bp in 2011 (i.e. 2% of the face value, or 3% to 4% of the bond value) while they were lower than 40 bps for other euro-area countries (i.e. lower than about 0.5% of the face value). In addition, anecdotal evidence and discussion with traders cast doubts over the reliability of quoted prices for these bonds from 2010 to 2012.

[15] The resulting (coupon) yield curve is then bootstrappead to get zero coupon yields. The arbitrage strategy underlying the construction of such a riskfree rate, as well as its limitations, is discussed in Blanco, Brennan and Marsh (2005)). An alternative would be to take the swap rates as riskfree proxies, which should result in the same riskfree rates in some instances (see Duffie, 1999); however the so-called CDS-bond basis (the deviation between (a)

In the next subsection, we introduce and examine the so-called KfW-Bund spread, that is going to be key in our estimation strategy. The subsequent subsection (3.3) highlights some features of euro-area sovereign spreads.

3.2 The KfW-*Bund* spread

In this paper, a key element to identify liquidity-pricing effects is the term structure of spreads between German federal bonds and KfW agency bonds. The relevance of the KfW-*Bund* spread as a liquidity-pricing proxy is pointed out and exploited by Schwarz (2009).[16] First, we explain why both Bunds and KfW's bonds differ mainly by their liquidity. Second, we show that the KfW-Bund spread is not specific to Germany and carries European-wide liquidity-pricing effects. Third, a regression-based preliminary analysis relates these spreads to key macro-financial indicators.

3.2.1 The content of the KfW-Bund spread

KfW is a bank owned by the Federal republic of Germany and the federal states.[17] Its bonds are explicitly and fully guaranteed by the German federal government, which implies that the credit quality of KfW bonds is virtually the same as the one of the Bunds, that are their sovereign counterparts.[18] By contrast, the liquidity of KfW bonds is lower than that of Bunds. Actually, the latter are known to be the most liquid bonds in the euro-area bond market. This, added to a strong credit quality, contributes to the safe-haven status of the Bunds. As such, the Bund is known as the main interest-rate-hedging asset used by portfolio managers. The high liquidity of the Bunds is reflected by the fact that bid-ask spreads on Bunds are the tightest across euro-area

the CDS and (b) the spread between bond and swap yields) has sometimes departed from zero, especially during the crisis period (see Bai and Collin-Dufresne, 2012). Using swaps as riskfree rates results in credit parts of the German yields that are difficult to interpret and notably to reconcile with German CDS fluctuations. Notwithstanding, apart for the credit parts of Germany, Finland and the Netherlands (that tend to be smaller when swap rates are used as riskfree rates), the overall results of our study are fairly robust to the riskfree-rate choice.

[16] See also McCauley (1999), the ECB (2009) and, more recently, Ejsing, Grothe and Grothe (2012). In the same spirit, with U.S. data, Longstaff (2004) computes liquidity premiums based on the spread between Treasuries and bonds issued by Refcorp, that are guaranteed by the U.S. Treasury.

[17] 80% of KfW's capital is held by by the Federal Republic of Germany and the remaining 20% by the German federal states (Source: www.kfw.de).

[18] An understanding between the European Commission and the German Federal Ministry of Finance (1 March 2002) stated that the guarantee of the Federal Republic of Germany will continue to be available to KfW. The federal law that guarantees the credit quality of the KfW bonds may change in the future; however, the probability of such a reversal is very small. The three main rating agencies –Fitch, Standard and Poor's and Moody's– have assigned a triple-A rating to KfW. In addition, as the German federal bonds, KfW's bonds are zero-weighted under the Basel capital rules and both kinds of bonds are also identical in their tax treatment.

sovereign marketable debts. Even if KfW bonds present relatively tight bid-ask spreads, these turn out to be far larger than Bunds' ones, especially during distress periods.[19] There are several reasons why the Bunds are more liquid than KfW bonds: (a) In terms of size, measured by the total outstanding volume or issuances, German sovereign marketable debt is three to four times larger than KfW's one;[20] (b) the average issue size of a Bund is about three times larger than the average issue size of KfW's bonds (see Schuster and Uhrig-Homburg, 2012); (c) the Bunds are the only deliverable underlying in the Eurex futures contracts, which boosts demand for the German Bund compared to other euro area debt and bolsters its liquidity (see e.g. Pagano and von Thadden, 2004 and Ejsing and Sihvonen, 2009); (d) while both KfW's bonds and Bunds are accepted by the European Central Bank (ECB) as collateral for the refinancing transactions, the two kinds of bonds do not fall in the same ECB liquidity categories (KfWs' bonds are in the second highest category, the Bunds are in the highest).[21]

Consequently, the spreads between Bunds and KfW's bonds essentially correspond to the excess yield-to-maturity demanded by investors to hold less-liquid KfW bonds instead of the highly-liquid Bunds. An important aspect of using such liquidity-pricing proxies is that theit term structure is available. Indeed, this will be exploited to estimate liquidity-pricing effects along the term structure of yields.

3.2.2 The KfW-Bund spread as a European liquidity-pricing factor

Panel A of Figure 1 shows that the KfW-Bund spreads of different maturities are highly correlated. This suggests that a single factor may be adequate to model the term structure of these spreads. Moreover, it is important to check that this liquidity-pricing measure is not purely specific to Germany. To that purpose, we look at comparable liquidity-driven spreads –between government-

[19] According to Thomson Reuters Tick History data, the average bid-ask spread on the 10-year German benchmark sovereign bond was about 0.5 bp over the period 2007-2012; six times lower than the average of the bid-ask spreads associated with a benchmark bond issued by KfW. The latter is close to the average of the bid-asks spreads on sovereign bonds across euro-area countries.

[20] According to Barclays (2012), the bond debt outstanding of KfW is lower than EUR300bn, whereas the outstanding volume of tradable German-government securities is of about EUR1100bn. In 2012, the bond-issuance target for 2012 is about EUR80bn for KfW, against more than EUR250bn for the Federal Republic of Germany.

[21] This leads to small additional haircuts for KfWs of up to 2% for the longest maturities, see Governing Council of the ECB (2011).

guaranteed bonds and their sovereign counterparts– in alternative countries.[22] In France for instance, the CADES (Caisse d'amortissement de la dette sociale) issues bonds that are guaranteed by the French government. Panel B compares one of the KfW-Bund spreads with a CADES-OAT spread (OATs are French government-issued bonds) and displays spreads of government-guaranteed bank bonds –issued by the Dutch NIBC bank and the Austrian Raiffeisen Zentalbank– over their respective sovereign counterparts. This exercise points to a substantial degree of correlation among liquidity-driven spreads from different European countries, which is consistent with the literature on commonality in liquidity.[23]

3.2.3 The KfW-Bund spread and macro-financial indicators

In this subsection, we investigate the connection between the KfW-Bund spreads and other measures of economic and financial stress. Table 1 presents the results of OLS regressions of the quarterly changes in the 10-year KfW-Bund spread on several key macro-financial variables. We consider the following variables: (i) the Dow Jones Eurostoxx 50, that is the most-cited European-wide stock index, (ii) the VSTOXX, which is an implicit volatility measure based on the Eurostoxx (that can be seen as the equivalent of the American VIX for the euro area), (iii) the 3-month EURIBOR-OIS spread, that is a gauge of banks' willingness to lend on the unsecured euro-area money markets and that is known to be related to both credit and liquidity risks (see e.g. Taylor and Williams, 2009 or Filipovic and Trolle, 2012), (iv) a proxy for the short-term riskfree rate (the 3-month OIS, i.e. overnight-indexed swap), (v) an interest-rate option-implied volatility (BVOL, the 10-years-in-10-years Black volatility) and (vi) a business cycle indicator (ESI, the Economic Sentiment Index provided by the European Commission).

The results suggest that changes in the KfW-Bund spread are positively related to market volatilities (BVOL and VSTOXX) and to interbank-market risks (the EURIBOR-OIS spread), and negatively to short-term rates, stock returns and economic activity. The seventh column shows the result of the regression of the KfW-Bund spread on all variables simultaneously. Changes in the stock index and in the interbank risks appear to be the variables that are the most significantly

[22] Note that such alternative (term structures of) spreads are not available on our whole estimation period, that is why we use essentially KfW-Bund spreads to identify our liquidity factor within our econometric approach.

[23] See Footnote 11.

Tab. 1: **Preliminary analysis of KfW-*Bund* spreads**

Notes: This table presents the results of regressions of the 5-year KfW-Bund spread (in bps) on a set of macro-financial variables. All variables are in quarterly changes. The VSTOXX (implied volatility in EUROSTOXX 50 option prices), the BVOL (10-year-in-10-year interest-rate-swap Black volatility), the overnight-index swap rate (OIS, proxying the short-term riskfree rate) the ESI (business-cycle index provided by the European Commission) are expressed in percentage points, the EUROSTOXX 50 in hundreds of points and the 3-month-maturity EURIBOR-OIS spread in basis points. The data are weekly and cover the period from 7 July 2006 to 15 February 2013 (346 observations). Newey-West (12 lags) corrected standard deviations of the parameter estimates are reported in brackets. The regressions include intercepts that are not reported in the table.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VSTOXX	0.91***						0.19	
	(0.16)						(0.25)	
EUROSTX		-3.46***					-2.03***	-2.79***
		(0.55)					(0.74)	(0.36)
IBOR-OIS			0.29***				0.14**	0.18***
			(0.07)				(0.06)	(0.04)
3-mthOIS				-11.04**			0.19	
				(5.31)			(3.43)	
BVOL					2.25**		0.12	
					(0.89)		(0.57)	
ESI						-1.51***	-0.39	
						(0.47)	(0.47)	
Adj.R^2	0.35	0.4	0.27	0.09	0.12	0.18	0.49	0.49

related to the KfW-Bund spread.

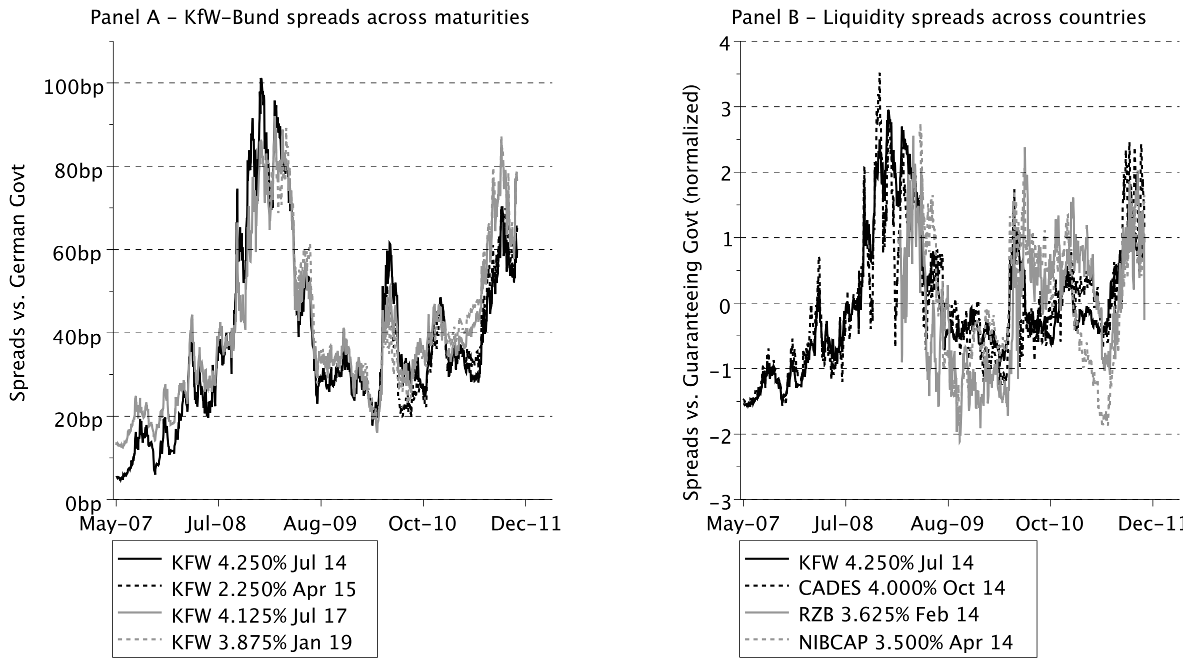
Overall, the results suggest that liquidity-pricing effects, as proxied by the KfW-Bund spread, have important connections with measures of macro-financial risks. In particular, these regressions support the view that deteriorating economic and financial situations may encourage investors to migrate to the most liquid bonds, widening the spreads between the latter and relatively less liquid bonds.

3.3 Euro-area government yields

Table 2 reports the correlations between the spreads vs. riskfree rates for different countries over the sample period. The results suggest that euro-area sovereign spreads are highly correlated across countries and across maturities (see also Favero, Pagano and von Thadden, 2010). Spreads' distributions are positively skewed and often leptokurtic. Table 3 presents a principal-component analysis of these spreads across countries. This analysis indicates that, for different maturities (2, 5 and 10 years), the first two principal components explain more than 90% of the spread variances across countries (70% for the first principal component alone), pointing towards the existence of

Fig. 1: **Differentials between government and government-guaranteed bonds**

Notes: The first plot shows the spreads between KfW bond yields and their sovereign counterparts. The second plot compares the spread between a KfW bond maturing in 2014 and its sovereign counterpart with other spreads between government-guaranteed European bonds and their respective sovereign counterparts: CADES', RZB' and NIBCAP's bonds are respectively guaranteed by the French, Austrian and Dutch governments (the spreads are demeaned and standardised). The yields come from Barclays Capital.



common sources of risk in euro-area sovereign spreads.

4 The model

We consider N countries that may default and whose bonds are not perfectly liquid. Heuristically, credit and liquidity risks reflect two kinds of possible losses for a bondholder: (a) the default of the issuer implies an early and reduced repayment of its bonds, (b) not-perfectly-liquid bonds may have to be liquidated at a discount price if the bondholder has to sell them precipitately.

Subsection 4.1 presents the notations and depicts the historical dynamics of the model variables. Subsection 4.2 introduces the stochastic discount factor and discusses the implied risk-neutral dynamics of the variables. Bond pricing is developed in Subsection 4.3.

4.1 Information, historical dynamics

At date t, the investor is provided with a new information $W_t = (r_t, z_t', \lambda_{c,t}', \lambda_{\ell,t}, d_t', \ell_t)'$ defined as follows: d_t is a N-dimensional vector of binary variables $d_t^{(n)}$, $n = 1, \ldots, N$, indicating whether

Tab. 2: **Descriptive statistics of selected spreads**

Notes: The table reports summary statistics for selected spreads (versus riskfree rates, obtained as the German yields from which we subtract German CDS). Two auto-correlations are shown (the 1-week and the 1-month auto-correlations). The underlying yields are continuously compounded and are in percentage annual terms. The lower panel of the table presents the covariances and the correlations (in italics) of the spreads. The data are weekly and cover the period from 7 July 2006 to 15 February 2012.

	France		Italy		Netherlands		Spain	
	2-year	10-year	2-year	10-year	2-year	10-year	2-year	10-year
Mean	0.171	0.333	1.239	1.534	0.106	0.218	1.231	1.528
Median	0.114	0.247	0.648	0.896	0.077	0.173	0.511	0.709
Standard dev.	0.203	0.329	1.364	1.576	0.107	0.181	1.399	1.657
Skewness	2.786	1.697	1.625	1.23	1.675	0.934	1.03	0.973
Kurtosis	13.302	5.782	5.323	3.379	5.956	3.207	3.081	2.84
Auto-cor. (lag 1)	0.956	0.981	0.986	0.993	0.923	0.971	0.983	0.991
Auto-cor. (lag 4)	0.854	0.938	0.939	0.968	0.793	0.914	0.94	0.969
Correlations \ **Covariances**								
France 2-yr yd	0.041	0.06	0.226	0.235	0.014	0.027	0.173	0.18
France 10-yr yd	*0.897*	0.108	0.411	0.47	0.019	0.05	0.37	0.428
Italy 2-yr yd	*0.818*	*0.917*	1.858	2.099	0.053	0.174	1.774	2.009
Italy 10-yr yd	*0.737*	*0.907*	*0.978*	2.482	0.053	0.205	2.098	2.472
Netherlands 2-yr yd	*0.667*	*0.545*	*0.363*	*0.316*	0.011	0.015	0.029	0.029
Netherlands 10-yr yd	*0.728*	*0.836*	*0.701*	*0.717*	*0.759*	0.033	0.153	0.184
Spain 2-yr yd	*0.612*	*0.805*	*0.931*	*0.952*	*0.196*	*0.601*	1.955	2.273
Spain 10-yr yd	*0.538*	*0.785*	*0.89*	*0.947*	*0.163*	*0.613*	*0.982*	2.743

Tab. 3: **Principal component analysis of euro-area yield differentials**

Notes: This table presents results of principal-component analyses carried out on the spreads versus riskfree yields proxied by the German yields from which we subtract maturity-matching CDS. There are three analyses that correspond respectively to three maturities: 2 years, 5 years and 10 years. For each PC analysis, the table reports the eigenvalues of the covariance matrices and the proportions of variance explained by the corresponding component (designated by "Prop. of var."). The data are weekly and cover the period from 7 July 2006 to 15 February 2013. the spreads of eight countries are included in the analysis (Austria, Belgium, Finland, France, Germany, Italy, Netherlands, Spain).

Component	1	2	3	4	5	6	7	8
2-year spread								
Eigenvalue	5.68	1.44	0.37	0.3	0.1	0.05	0.04	0.02
Prop. of var.	71%	18%	4%	4%	1%	1%	0%	0%
Cumul. prop.	71%	90%	94%	97%	99%	99%	100%	100%
5-year spread								
Eigenvalue	6.68	0.85	0.23	0.11	0.07	0.03	0.02	0.01
Prop. of var.	84%	11%	3%	1%	1%	0%	0%	0%
Cumul. prop.	84%	94%	97%	98%	99%	100%	100%	100%
10-year spread								
Eigenvalue	6.52	1.08	0.19	0.11	0.05	0.02	0.02	0.01
Prop. of var.	81%	13%	2%	1%	1%	0%	0%	0%
Cumul. prop.	81%	95%	97%	99%	99%	100%	100%	100%

debtor n is in default at date t ($d_t^{(n)} = 1$) or not ($d_t^{(n)} = 0$);[24] ℓ_t is a binary variable indicating if the bondholder is affected by a liquidity shock ($\ell_t = 1$) or not ($\ell_t = 0$); the vector $\lambda_{c,t}$ is N-dimensional and its n^{th} component, $\lambda_{c,t}^{(n)}$, is a default (or credit) intensity, defined below, associated with debtor n; similarly, $\lambda_{\ell,t}$ is a liquidity-shock intensity, also defined below; z_t is a crisis-regime variable, which can take nine values denoted by $e_1, \dots, e_9$, where e_j is the 9-dimensional (selection) vector whose all components are equal to zero except the j^{th} one which is equal to one; finally r_t is the riskfree short-term rate between t and $t+1$.

4.1.1 Historical dynamics of W_t

The joint dynamics of W_t will be defined by successively specifying:

(i) the conditional distribution of $(d_t', \ell_t)'$ given $(r_t, z_t', \lambda'_{c,t}, \lambda_{\ell,t}, \underline{W}_{t-1}')'$ where $\underline{W}_{t-1} = (W_{t-1}', \dots, W_1')'$,

(ii) the conditional distribution of $(\lambda'_{c,t}, \lambda_{\ell,t})'$ given $(r_t, z_t', \underline{W}_{t-1}')'$ and

(iii) the conditional distribution of $(r_t, z_t')'$ given $\underline{W}_{t-1}$.

(i) The d_t and ℓ_t variables The variables $d_t^{(1)}, \dots, d_t^{(N)}, \ell_t$ are assumed to be independent conditional on $\mathcal{I}_t = (r_t, z_t', \lambda'_{c,t}, \lambda_{\ell,t}, \underline{W}_{t-1}')'$ and such that:[25]

$$\begin{cases} P\left(d_t^{(n)} = 1 \middle| d_{t-1}^{(n)} = 0, \mathcal{I}_t\right) & = 1 - \exp(-\lambda_{c,t}^{(n)}) \\ P\left(d_t^{(n)} = 1 \middle| d_{t-1}^{(n)} = 1, \mathcal{I}_t\right) & = 1 \\ P\left(\ell_t = 1 | \ell_{t-1} = 0, \mathcal{I}_t\right) & = 1 - \exp(-\lambda_{\ell,t}) \\ P\left(\ell_t = 1 | \ell_{t-1} = 1, \mathcal{I}_t\right) & = 1 \end{cases} \tag{1}$$

In other words, the states $d_t^{(n)} = 1$ and $\ell_t = 1$ are absorbing and $\lambda_{c,t}^{(n)}$ and $\lambda_{\ell,t}$ are, respectively, interpreted as default and liquidity-shock intensities.

(ii) The intensities At the pricing stage, it will be convenient to introduce some so-called fractional-loss intensities denoted by $\lambda_{fc,t}^{(n)}$ and $\lambda_{f\ell,t}^{(n)}$, which are derived from the intensities $\lambda_{c,t}^{(n)}$

[24] We use parenthesis to distinguish country from exponentiation in the superscript.

[25] The independence assumption appears in the "doubly stochastic" framework (see e.g. Duffie et al., 2005, Pan and Singleton, 2008 or Longstaff et al., 2011). Note however that, conditionally on the past information $\underline{W}_{t-1}$, the default events and the liquidity shocks are not independent as soon as the associated intensities are not conditionally independent.

and $\lambda_{\ell,t}$ according to:

$$\begin{cases} \exp\left(-\lambda^{(n)}_{fc,t}\right) & = \exp\left(-\lambda^{(n)}_{c,t}\right) + \zeta\left(1 - \exp\left(-\lambda^{(n)}_{c,t}\right)\right), \quad \forall n \in \{1, \dots, N\} \\ \exp\left(-\lambda^{(n)}_{f\ell,t}\right) & = \exp\left(-\lambda_{\ell,t}\right) + \theta^{(n)}\left(1 - \exp\left(-\lambda_{\ell,t}\right)\right), \quad \forall n \in \{0, 1, \dots, N\}, \end{cases} \tag{2}$$

where the superscript $n = 0$ corresponds to KfW. If the Federal republic of Germany is the first debtor ($n = 1$), the discussion in Subsection 3.2.1 leads to $\lambda^{(0)}_{fc,t} = \lambda^{(1)}_{fc,t}$.

Assuming that the default and liquidity-shock intensities are small, we consider, in the following, the linearised version of Equation (2), that is:

$$\begin{cases} \lambda^{(n)}_{fc,t} & \simeq (1 - \zeta)\lambda^{(n)}_{c,t} \\ \lambda^{(n)}_{f\ell,t} & \simeq (1 - \theta^{(n)})\lambda_{\ell,t} \end{cases}, \quad \forall n \in \{0, 1, \dots, N\}, \tag{3}$$

Let us detail the modeling context that gives rise to these fractional-loss intensities, which amounts to defining parameters ζ and $\theta^{(n)}$.[26] Consider a bond issued by debtor n with a residual maturity h at date t. We denote by $B^{(n)}_{t,h}$ the price at which the bondholder can sell this asset at date t if (a) debtor n is not in default and (b) the bondholder has not been hit by the liquidity shock at date t. Intuitively, $\lambda^{(n)}_{fc,t}$ and $\lambda^{(n)}_{f\ell,t}$ correspond to the expected losses, conditional on $\mathcal{I}_t$, associated with, respectively, the default of debtor n and the arrival of a liquidity shock. These losses are expressed as fractions of $B^{(n)}_{t,h}$. More specifically:

- If debtor n defaults between $t - 1$ and t, the bondholder receives, from the bond issuer, a fraction ζ of the price that would have prevailed otherwise at date t. In other words, in the case of default, the recovery pay-off is $\zeta B^{(n)}_{t,h}$.

- The liquidity shock is formally represented by the state $\ell_t = 1$. When hit by such a shock, the bondholder is forced to sell her bonds. The rationale behind such a shock is the possible occurrence of unexpected cash shortages (that may e.g. result from unexpected margin calls faced by the bondholder).[27] Upon the arrival of the liquidity shock ($\ell_t = 1$), the bond investor sells her bond at a fractional cost $1 - \theta^{(n)}$, that is, the proceed of the sale is then $\theta^{(n)} B^{(n)}_{t,h}$.

[26] The formal derivation of the previous equations is given in Appendix B,

[27] It may also correspond to the need to rebalance a portfolio in order to maintain a hedging or diversification strategy, or it may result from a change in capital requirements (see Ericsson and Renault, 2006).

> Appendix D provides a structural interpretation of this fractional cost that broadly follows the lines of Ericsson and Renault (2006).[28]

Knowing the $(N+1)$-dimensional vector of intensities $(\lambda'_{c,t}, \lambda_{\ell,t})'$ is equivalent to knowing the vector of fractional intensities $\lambda_t = (\lambda'_{fc,t}, \lambda^{(0)}_{f\ell,t})'$. The conditional distribution of λ_t given $(r_t, z'_t, \underline{W}'_{t-1})$ is defined by:

$$\lambda_t = \mu' z_t + \Phi \lambda_{t-1} + \Sigma \varepsilon_t \tag{4}$$

where μ is a $9 \times (N+1)$ matrix of drifts, Φ is a diagonal matrix whose diagonal entries are ρ_c (N times) and ρ_ℓ, Σ is a diagonal matrix whose diagonal entries are denoted by $(\sigma_c^{(1)}, \ldots, \sigma_c^{(N)}, \sigma_\ell)$ and ε_t is a vector of i.i.d. $\mathcal{N}(0,1)$ shocks. By abuse of notation, we may denote the entries of λ_t by $\lambda_{i,t}$ in the following.[29]

It can be seen that the marginal distribution of the $\lambda_{i,t}$s are positively skewed as soon as the μ vectors contain only positive entries. Moreover, the lower the standard deviations σ of the Gaussian shocks (in comparison with the drifts μ), the more often the $\lambda_{i,t}$'s are positive, which is important given their interpretations in terms of probabilities.

Furthermore, the instantaneous causality between z_t and λ_t implies that, conditionally on the past information $\underline{W}_{t-1}$, the variances of the $\lambda_{i,t}$'s depend on the regime variable z_{t-1}. More precisely, conditionally on $\underline{W}_{t-1}$, the distributions of the $\lambda_{i,t}$'s are some mixtures of Gaussian distributions, thereby involving a form of conditional heteroskedasticity.[30]

(iii) The regime variable The regime variable z_t is obtained by crossing two regime variables, $z_{\ell,t}$ and $z_{c,t}$, that we refer to as liquidity and credit chains, respectively. Each of these variables is valued in $\{[1,0,0]',[0,1,0]',[0,0,1]'\}$, these three values corresponding respectively to a low-stress situation, a medium-stress and a high-stress situation. The credit/liquidity state of the economy at date t is then summarised by the nine-dimensional selection vector z_t, which is the Kronecker

[28] This modeling framework is consistent with the definition of bond liquidity according to which the more liquid an asset is, the easier it is for a seller to collect high bids from potential buyers in a limited amount of time.

[29] $\lambda_{i,t} = \lambda^{(i)}_{fc,t}$ for $i \leq N$ and $\lambda_{N+1,t} = \lambda^{(0)}_{f\ell,t}$.

[30] Such a feature is discussed in Ang, Bekaert and Wei (2008).

product of $z_{\ell,t}$ and $z_{c,t}$:

$$z_t = z_{\ell,t} \otimes z_{c,t}, \tag{5}$$

The Markovian chain z_t is assumed to be exogenous. In other words, the conditional distribution of z_t given $(r_t, \underline{W}_{t-1})$ is assumed to depend on z_{t-1} only.

Importantly, there may be causal relationships between $z_{\ell,t}$ and $z_{c,t}$. For instance, we allow for the probability of a change in the liquidity state to depend on the credit regime (and vice-versa). Formally, let us denote by Π the matrix of transition probabilities, whose (i, j) entry, denoted by $\pi_{i,j}$, corresponds to $p(z_{t+1} = e_j | z_t = e_i)$. The entries of the rows of this matrix summing to one, 72 parameters are required to specify this matrix. In order to keep the model parsimonious, some constraints are introduced (Appendix C) and, eventually, 14 parameters are required to specify the matrix Π.

Finally, as in e.g. Pan and Singleton (2008) or Longstaff et al. (2011), the process r_t is assumed to be autonomous and we do not make any assumptions about its dynamics.

4.2 Stochastic discount factor and risk-neutral dynamics of W_t

We assume that the stochastic discount factor (s.d.f.) has the following expression:

$$M_{t-1,t} = \exp\left[\beta(\underline{r}_t) - \frac{1}{2}\nu_t'\nu_t + \nu_t'\varepsilon_t + (\delta z_{t-1})' z_t\right] \tag{6}$$

where $\beta(\underline{r}_t)$ is any function of $\underline{r}_t$ satisfying $E_{t-1} \exp[\beta(\underline{r}_t)] = \exp(-r_{t-1})$, δ is a 9×9 matrix and where the entries of ν_t are affine in z_t and in the corresponding entries of λ_{t-1}, that is $\nu_{i,t} = \nu_{\lambda,i}\lambda_{i,t-1} + \nu_{z,i}' z_t$, say ($\nu_{\lambda,i}$ is a scalar and $\nu_{z,i}$ is a vector). The risk-sensitivity matrix δ and vectors ν_t respectively price the regimes z_t and the (standardised) Gaussian innovations ε_t of λ_t. Since we must have $E_t(M_{t,t+1}) = \exp(-r_t)$, the entries of δ are of the form $\ln(\pi_{ij}^*/\pi_{ij})$, where the π_{ij}^* are such that $\Sigma_j \pi_{ij}^* = 1$ for any i.

Our specification of the s.d.f. has several important consequences.

First, the fact that the regime variable z_t enters the s.d.f. means that changes in credit/liquidity stress regimes are priced, in the sense that investors require specific risk premiums to carry assets that are exposed to such stresses. This feature is consistent with the typically undiversifiable nature of sovereign risk, embedded by the regime variable z_t in our framework.[31] As a consequence, the risk-neutral dynamics of z_t is not the same as its historical dynamics. Specifically, under $\mathbb{Q}$, z_t follows a time-homogenous Markovian chain whose dynamics is described by the matrix Π^* of transition probabilities $\{\pi^*_{ij}\}$ (see Monfort and Renne, 2013).

Second, denoting by $\lambda_{i,t}$ the i^{th} entry of λ_t, we have, under $\mathbb{Q}$:

$$\lambda_{i,t} = \mu_i^{*'} z_t + \rho_c^* \lambda_{i,t-1} + \sigma_i \varepsilon^*_{i,t} \tag{7}$$

where $\varepsilon^*_{i,t} \sim \mathcal{N}^{\mathbb{Q}}(0,1)$, $\mu_i^* = \mu_i + \sigma_i \nu'_{z,i}$ and $\rho_c^* = \rho_c + \sigma_i \nu_{\lambda,i}$.

Third, since the variables $(d_t', \ell_t)'$ do not appear in $M_{t-1,t}$, we know (see Monfort and Renne, 2013, Lemma 1) that the conditional distribution of $(d_t', \ell_t)'$ given $(r_t, z_t, \lambda_t, \underline{W}_{t-1})$ is identical in the historical and the risk-neutral worlds; in particular, the processes $\lambda_{c,t}$ and $\lambda_{\ell,t}$ also verify Equation (1) if we replace the historical probability $\mathbb{P}$ by the risk-neutral probability $\mathbb{Q}$.[32] However, it has to be stressed that while the intensities are the same processes under both measures, their $\mathbb{Q}$- and $\mathbb{P}$-dynamics are different and are defined respectively by Equation (7) and Π^* and by Equation (4) and Π. This implies in particular that the probabilities of default are different under $\mathbb{P}$ and $\mathbb{Q}$. This will be illustrated in Subsection 6.3.

Fourth, since $M_{t-1,t}$ is the product of a function of $\underline{r}_t$ by a function of the present and past values of the other processes, the process r_t remains independent of the other processes in the risk-neutral world.

4.3 Bond pricing

In this framework, the price of a defaultable and illiquid zero-coupon bond issued by country n (not in default at date t) and with residual maturity h has a price at time t that is given by (see Appendix B):

$$B^{(n)}_{t,h} = E^{\mathbb{Q}}_t \left[\exp\left(-r_t - \ldots - r_{t+h-1} - \lambda^{(n)}_{fc,t+1} - \ldots - \lambda^{(n)}_{fc,t+h} - \lambda^{(n)}_{f\ell,t+1} - \ldots - \lambda^{(n)}_{f\ell,t+h} \right) \right]. \tag{8}$$

[31] For a discussion of the systematic nature of sovereign risk, see e.g. Longstaff et al. (2011) or Borri and Verdelhan (2011).
[32] This assumption according to which default and liquidity events (d_t and ℓ_t) are not priced is reasonable when the factors already included in the s.d.f. satisfyingly capture the sovereign risk. Using non-linear latent factors in the intensity specification contributes to addressing this point.

where $E_t^{\mathbb{Q}}$ is the conditional expectation given $\underline{W}_t$ in the risk-neutral world.

Since r_t is independent from the other processes under $\mathbb{Q}$, we have:

$$\begin{aligned} B_{t,h}^{(n)} &= E_t^{\mathbb{Q}}\left[\exp\left(-r_t - \ldots - r_{t+h-1}\right)\right] \times \\ & E_t^{\mathbb{Q}}\left[\exp\left(-\lambda_{fc,t+1}^{(n)} - \ldots - \lambda_{fc,t+h}^{(n)} - \lambda_{\ell,t+1}^{(n)} - \ldots - \lambda_{\ell,t+h}^{(n)}\right)\right]. \end{aligned} \tag{9}$$

Denoting by $y_{t,h}^{(n)}$ the yield-to-maturity of this bond, we obtain:

$$\begin{aligned} y_{t,h}^{(n)} &= -\frac{1}{h}\ln(B_{t,h}^{(n)}) \\ &= r_{t,h} - \frac{1}{h}\ln\left(E_t^{\mathbb{Q}}\left[\exp\left(-\lambda_{fc,t+1}^{(n)} - \ldots - \lambda_{fc,t+h}^{(n)} - \lambda_{\ell,t+1}^{(n)} - \ldots - \lambda_{\ell,t+h}^{(n)}\right)\right]\right) \end{aligned} \tag{10}$$

where $r_{t,h}$ denotes the yield to maturity of a riskfree zero-coupon bond of residual maturity h at date t. It can be shown (see e.g. Monfort and Renne, 2013) that the vector $(z_t', \lambda_t')'$ is compound auto-regressive of order one under $\mathbb{Q}$, which implies that the second term on the right-hand side of (10) is linear in $(z_t', \lambda_t')'$.[33] Therefore, the spread between the yield associated with the defaultable bond and the riskfree bond of the same maturity is of the form:

$$y_{t,h}^{(n)} - r_{t,h} = a_h^{(n)'} z_t + b_h^{(n)'} \lambda_t \tag{11}$$

where the vectors $(a_h^{(n)'}, b_h^{(n)'})'$ are computed recursively.[34]

5 Estimation

5.1 State-space form of the model

From an econometric point of view, the model is a state-space model with Gaussian innovations and regime-switching. Equation (4) is the transition equation. The measurement equations are of two kinds: a first set of equations relates observed spreads to modeled ones; a second one relates

[33] See Darolles, Gourieroux and Jasiak (2006) or Bertholon, Monfort and Pegoraro (2008) for in-depth presentations of compound auto-regressive –or Car– processes.

[34] To compute the $(a_h^{(n)}, b_h^{(n)'})$'s, see Monfort and Renne (2013, Proposition 1) for (a) the Laplace transform of $(z_t', \lambda_t')'$and (b) the recursive algorithm that computes the multi-horizon Laplace transform of $(z_t', \lambda_t')'$ (Appendix A.5).

observed survey-based spreads' forecasts, stacked in a vector denoted by CF_t, to model-implied ones. Let us make these two sets of equations more precise.

5.1.1 Spread measurement equations

The measurement equations relate observed variables to modeled ones, up to a fitting error. Natural candidates for the observed spreads are those between the sovereign yields and their riskfree counterparts (i.e. the $y_{t,h}^{(n)} - r_{t,h}$ appearing in Equation 11). However, the most scrutinised spreads in the euro area are the ones versus Germany. As a consequence, and in order not to contaminate our estimation with uncertainties surrounding our riskfree-rate proxy (see discussion in Subsection 3.1), we decide to use the spreads versus Germany as observed spreads. It is important to note that this does not mean that we consider the German yields as riskfree; the measurement equations consistently take the German-specific risks into account. Let us put it in a formal way.

The vector of observed spreads is denoted by S_t. Denoting by $\xi_{S,t}$ a vector of i.i.d. normally-distributed pricing errors, the first set of measurement equations reads:

$$S_t = Az_t + B\lambda_t + \xi_{S,t} \tag{12}$$

where the entries of matrices A and B are respectively based on the $a_h^{(n)}$'s and the $b_h^{(n)}$'s appearing in equation (11). More precisely, consistently with the choice of the observed spreads, and replacing the German index (1) by GER, matrices A and B are constructed in the following manner:

- The rows of A and B corresponding to German spreads are respectively the a_h^{GER}'s and the b_h^{GER}'s with the appropriate maturities h.
- The rows of A and B corresponding to other debtors ($n > 1$) are of the form $a_h^{(n)} - a_h^{GER}$ and $b_h^{(n)} - b_h^{GER}$.

5.1.2 Measurement equations related to survey-based forecasts

The second set of measurement equations relates survey-based 12-month-ahead forecasts of spreads to their model-based equivalent. These equations are added to improve the estimation of the historical dynamics in a small-sample context. This approach is based on Kim and Orphanides

(2012) who show that adding such equations in the estimation alleviates the small-sample bias in the persistence of the yields.[35] Such bias typically results in too stable long-horizon expectations of yields (or spreads) and, as a consequence, overstate the variability of term premiums. Generating reliable $\mathbb{P}$-expectations is key if one wants to use the model to recover probabilities of default from bond prices. Indeed, these probabilities are based on model-based $\mathbb{P}$-forecasts of future default intensities (this is made precise in Subsection 6.3).

Due to survey-based data availability, four spreads are considered here: the yield differentials between Dutch, French, Italian and Spanish 10-year bonds and their German counterparts. These equations read:

$$CF_t^{(n)} = E_t^{\mathbb{P}}\left(y_{t+h,H}^{(n)} - y_{t+h,H}^{GER}\right) + \xi_{CF,t}^{(n)}, \quad n \in \{2,3,4,5\} \tag{13}$$

where the maturity of the considered rates is 10 years, i.e. $H = 52 \times 10$ weeks, the forecast horizon h is of 52 weeks, the $\xi_{CF,t}^{(n)}$'s are i.i.d. normally-distributed measurement errors and where the model-based forecasts $E_t(y_{H,t+h}^{(n)} - y_{H,t+h}^{GER})$ are easily derived using equation (11) and:

$$\begin{cases} E_t^{\mathbb{P}}(\lambda_{t+h}) & = \left[\mu\Pi^h + \Phi\mu\Pi^{h-1} + \ldots + \Phi^{h-1}\mu\Pi\right] z_t + \Phi^h \lambda_t \\ E_t^{\mathbb{P}}(z_{t+h}) & = \Pi^h z_t. \end{cases} \tag{14}$$

Appendix C presents and discusses different constraints that are imposed on the parameter estimates.

5.2 Estimation procedure and results

Our estimation is conducted by the maximum-likelihood method, in a single step. The likelihood function is approximated by means of the Kim's (1994) filter, that handles state-space models with Markov-switching.[36]

[35] This bias is obtained with conventional estimation. Other approaches have been proposed to deal with that problem. In particular, Jardet, Monfort and Pegoraro (2013)[54] use a "near-cointegrated framework" specification of the factors (averaging a stationary and a cointegrated specification).

[36] See also Kim and Nelson (1999). The algorithm has been slightly adapted for this application. In particular, for each iteration of the algorithm, in the updating step, we prevent the algorithm from resulting in values of the ith credit intensity $\lambda_{i,t}$ that would be below $-2\sqrt{\sigma_i^2/(1-\rho_i^2)}$. Note that $\sigma_i^2/(1-\rho_i^2)$ would be the unconditional variance of the ith unobserved factor $\lambda_{i,t}$ if μ_i was null (since the vector μ_i is positive, the unconditional mean of the ith factor is higher than zero, implying that the unconditional probability of having $\lambda_{i,t} < -2\sqrt{\sigma_i^2/(1-\rho_i^2)}$ is lower than 2.5%).

Figure 2 illustrates the ability of the model to capture spreads' fluctuations by comparing actual spreads (grey lines) to modeled ones (black solid lines). The average of the measurement-error standard deviations is about 20 basis points (across 27 time series: 3 maturities for 9 debtors). The model also proves to be able to mimic survey-based forecasts, model-implied forecasts accounting for 95% of survey-based forecasts' variances.

Tables 4 and 5 report the parameter estimates. The standard deviations of these estimates are based on the outer product of the first derivative of the likelihood function. Important differences arise in the parameters across countries. Naturally, those countries that have been characterised by the highest rises in spreads are more affected by the crises regimes. Notably, in an severe credit crisis regime, the drift of the the Italian or the Spanish credit intensities are more than 10 times larger than those of Germany, Finland or the Netherlands (see line "μ_{cc}" in Table 4). It can also be noted that the auto-regressive coefficients (the ρs) are closer to one under the risk-neutral measure than under the physical one. This suggests that credit and liquidity intensities factors are more persistent under the risk-neutral measure than under the historical one. Note that another source of persistence originates from the regime-switching features: indeed, low switching probabilities generate persistence in the processes that depend on these regimes (the $\lambda_{i,t}$'s here). However, it is difficult to draw general conclusion regarding the differences between the transition probabilities under $\mathbb{Q}$ and $\mathbb{P}$. Typically, while the severe-credit-crisis state is far more persistent under $\mathbb{Q}$ than under $\mathbb{P}$, it is the opposite situation for the severe-liquidity-crisis states.

The probabilities estimates reported in Table 5 point to the existence of causality between credit- and liquidity-stress situations. This can be seen by comparing the probabilities of switching from a given credit-stress (liquidity-stress) situation conditionally to the liquidity-stress (credit-stress) situation. For instance, under both measures, the probabilities of switching to the severe-credit-crisis are higher when the liquidity situation is deteriorated (this can be seen by comparing lines "$C_t - NL_t$" and "$C_t - L/LL_t$" in Table 5). Consider the probabilities of switching from the intermediary liquidity-stress situation (lines "$L_t - NC_t$" and "$L_t - C/CC_t$"), the probabilities of switching to the severe-liquidity-stress situation hardly depend on the credit situation (both are about 10% under $\mathbb{P}$ and close to one under $\mathbb{Q}$); however, the probability of moving to the low-liquidity-stress regime is lower in if the credit situation is in one of the crisis states (C_t or CC_t).

Overall these estimated transition probabilities point to the existence of causality between credit- and liquidity-stress situations.

The first and third panels of Figure 3 present the smoothed probabilities of being in the different crisis regimes.[37] The first period of liquidity stress is relatively short (a few weeks) and begins with the collapse of Bear Sterns in March 2008. Most of the two years following the Lehman Brothers' bankruptcy (September 2008) are identified as liquidity-crisis periods, with peaks of tension in late 2008. The premise of the so-called euro-area crisis (April 2010) and the latest development of the same crisis (from mid-2011 to mid-2012) are also estimated as being liquidity-stress periods.

Turning to the credit crises, one can broadly distinguish two long stress periods: the first half of 2009 and May 2010 to the end of the sample. Within these credit-crisis phases, a few peaks of market stress (black-shaded areas in the second panel of Figure 3) took place in May 2010, in late 2011 and in Spring 2012. The second and fourth panels in Figure 3 respectively display the estimates of the liquidity intensity $\lambda_{\ell,t}$ and of some credit intensities $\lambda_{fc,t}^{(n)}$. Looking simultaneously at the first and second (third and fourth) panels illustrates the influence of the the liquidity (credit) regimes on the liquidity-related (credit-related) intensities.

6 Spread analysis

6.1 The liquidity-related part of the spreads

In order to gauge the relative importance of the liquidity-related part of the spreads, we compute the spreads that would prevail if the default intensities were equal to zero. Figure 2 presents the resulting spreads (dotted lines) and compare them with actual (grey lines) and fitted (black solid lines) spreads.[38] The actual and fitted spreads that are reported on this figure are against Germany, except for the latter country for which we report spreads against the riskfree rate; these are the spreads that are used in the state-space measurement equations.[39] These plots show that

[37] The smoothed probabilities are obtained by applying Kim's (1994) filter. While filtered probabilities, as of date t, use only information available up to date t, smoothed probabilities exploit all sample information.

[38] Due to non-linearity effects, the sum of these counterfactual spreads and those that would be obtained, alternatively, by switching the liquidity intensities off, are not strictly equal to the complete modeled spreads. However, the differences are visually imperceptible.

[39] Again, this does not mean that we consider the German bonds are riskfree. The fact that we assume that German bonds are not credit-riskfree is taken into account in the estimation process (see 5.1.1) as well as in the plots reported in Figure2.

Tab. 4: **Parameter estimates (1/2)**

Notes: This table reports the estimates of the parameters defining the dynamics of the intensities under the historical and the risk-neutral measures. The superscript "*" indicates that the considered parameter is associated to the risk-neutral dynamics. It is completed by an additional table (Table 5) that presents the estimated transition probabilities under both measures. The data are weekly and span the period from 7 July 2006 to 15 February 2013 (346 observations). Standard errors are reported in parentheses below the coefficient estimates. ***, ** and * respectively denote significance at the 1%, 5% and 10% level. The historical dynamics of the elements of λ_t are of the form $\lambda_{i,t} = \mu_i^{'} z_t + \rho_i \lambda_{i,t-1} + \sigma_i \varepsilon_{i,t}$; z_t, the regime variable, is a 9-dimensional selection vector that "picks" a drift among the entries of μ_i; for each i, only three out of the nine entries of μ_i are nonzero, these three entries corresponding to severe-stress states (matrix μ is detailed in Appendix C.4). The non-zero entries of μ are given by the row "μ_c" of the table (for credit intensities) and by "μ_ℓ" (liquidity intensity). Note that the parameters $1-\theta^{(n)}$ are estimated up to a multiplicative factor (accordingly, we report $(1-\theta^{(n)})/(1-\theta^{(KfW)})$ in the table). The standard deviations σ_{perr} are the standard deviations of the (pricing) measurement errors $\xi_{S,t}$ (Equation 12).

	μ_ℓ	$1-\rho_\ell$	μ_ℓ^*	$1-\rho_\ell^*$	σ_ℓ		$1-\rho_c$	$1-\rho_c^*$
	0.12***	0.072***	0.0059***	0.00091***	0.0045***		0.008***	0.000045
	(0.0067)	(0.0032)	(0.00038)	(0.000078)	(0.000094)		(0.00036)	(0.000075)

	GER	SPA	ITA	FRA	NTH	BEL	FIN	AUS
$\frac{1-\theta^{(n)}}{1-\theta^{(KfW)}}$	0	1***	1.5***	0.58***	0.6***	1.3***	0.63***	0.96***
	–	(0.39)	(0.53)	(0.054)	(0.021)	(0.079)	(0.032)	(0.04)
μ_c	0.05***	0.59***	0.66***	0.17***	0.044***	0.26***	0.047***	0.096***
	(0.0045)	(0.042)	(0.048)	(0.0108)	(0.0042)	(0.017)	(0.0066)	(0.007)
μ_c^*	0.009***	0.011***	0.0058	0.009***	0.0108***	0.0055***	0.0105***	0.0106***
	(0.0005)	(0.0038)	(0.0043)	(0.00064)	(0.0006)	(0.00109)	(0.00079)	(0.00079)
σ_c	0.0005***	0.01***	0.0093***	0.002***	0.00109***	0.0041***	0.0012***	0.0021***
	(0)	(0.00022)	(0.00021)	(0.000046)	(0.000024)	(0.000092)	(0.000026)	(0.000048)
σ_{perr}	0.14***	0.68***	0.63***	0.14***	0.074***	0.28***	0.08***	0.15***
	(0.00048)	(0.0024)	(0.0022)	(0.00049)	(0.00026)	(0.00099)	(0.00028)	(0.00052)

Tab. 5: **Parameter estimates (2/2)**

Notes: The table presents the estimates of the probabilities of transition. These probabilities define the dynamics of the regime variable z_t. NC, C and CC respectively stand for low-stress, medium-stress and high-stress credit situation (and the same for NL, L and LL, characterising the liquidity situations). The table reports the probabilities of switching from the states indicated in the columns (prevailing at date t) to the different stress situations at date $t+1$ (columns). For instance, if, at date t, the regime is NL-C or NL-CC, then the probability of switching to L at date $t+1$ is 4.3% under the historical measure and 2.1% under the risk-neutral measure. Standard errors are reported in parentheses below the coefficient estimates. ***, ** and * respectively denote significance at the 1%, 5% and 10% significance level.

Panel A – Under $\mathbb{P}$

		NL_{t+1}	L_{t+1}	LL_{t+1}
NL_t	NC_t	0.999***	0.00106	-
		(0.0011)	(0.0011)	-
	C/CC_t	0.96***	0.043***	-
		(0.012)	(0.012)	-
L_t	NC_t	0.037***	0.85***	0.109***
		(0.002)	(0.042)	(0.042)
	C/CC_t	0.0061	0.89***	0.105***
		(0.0055)	(0.034)	(0.034)
LL_t		-	0.69***	0.31***
		-	(0.092)	(0.092)

		NC_{t+1}	C_{t+1}	CC_{t+1}
NC_t	NL_t	0.99999***	0.000048	-
		(0.00057)	(0.00057)	-
	L/LL_t	0.99***	0.0062**	-
		(0.0027)	(0.0027)	-
C_t	NL_t	0.00005	0.999***	0.0012
		(0.00033)	(0.002)	(0.002)
	L/LL_t	0.0105***	0.99***	0.0018*
		(0.0021)	(0.0022)	(0.00105)
CC_t		-	0.999***	0.00085
		-	(0.00089)	(0.00089)

Panel B – Under $\mathbb{Q}$

		NL_{t+1}	L_{t+1}	LL_{t+1}
NL_t	NC_t	0.999***	0.0015***	-
		(0.00035)	(0.00035)	-
	C/CC_t	0.98***	0.021***	-
		(0.0056)	(0.0056)	-
L_t	NC_t	0.019***	0.000049	0.98***
		(0.004)	(0.12)	(0.11)
	C/CC_t	0.000048	0.00004	0.9999***
		(0.0046)	(0.13)	(0.13)
LL_t		-	1***	0.000049
		-	(0.088)	(0.088)

		NC_{t+1}	C_{t+1}	CC_{t+1}
NC_t	NL_t	0.99999***	0.000046	-
		(0.00027)	(0.00027)	-
	L/LL_t	0.99***	0.0053***	-
		(0.00102)	(0.00102)	-
C_t	NL_t	0.0075**	0.99***	0.000049
		(0.0036)	(0.0068)	(0.0046)
	L/LL_t	0.014**	0.93***	0.054***
		(0.0058)	(0.019)	(0.014)
CC_t		-	0.037***	0.96***
		-	(0.0076)	(0.0076)

Fig. 2: **Actual vs. model-implied spreads**

Notes: These plots compare observed (light-grey solid lines) and model-implied (black solid lines) spreads. In the estimation approach, the riskfree rate is proxied by the German rates from which we subtract the maturity-matching German CDS. The spreads that are reported are those that are fitted in the estimation process (see Equation 12). For all countries but Germany, these spreads are yield differentials between 5- and 10- year zero-coupon yields and their German counterparts; note however that this does not imply that Germany is considered as credit-risk free (see 5.1.1); this is done in order to prevent the estimates from being contaminated by the uncertainty related to the riskfree-rate proxy. The dotted line is the model-implied contribution of the liquidity intensity $\lambda_{\ell,t}$ to the corresponding spreads (these contributions are computed as the spread that would prevail if the credit parts $\lambda_{c,t}^{(n)}$ of the debtor intensities were equal to zero). For Germany, the liquidity component is zero, by assumption. For the KfW-Bund spread (bottom-right plot), the fact that the dotted line and the black solid line are confounded reflects the assumption according to which this spread is essentially liquidity-driven (see Subsection 3.2.1).

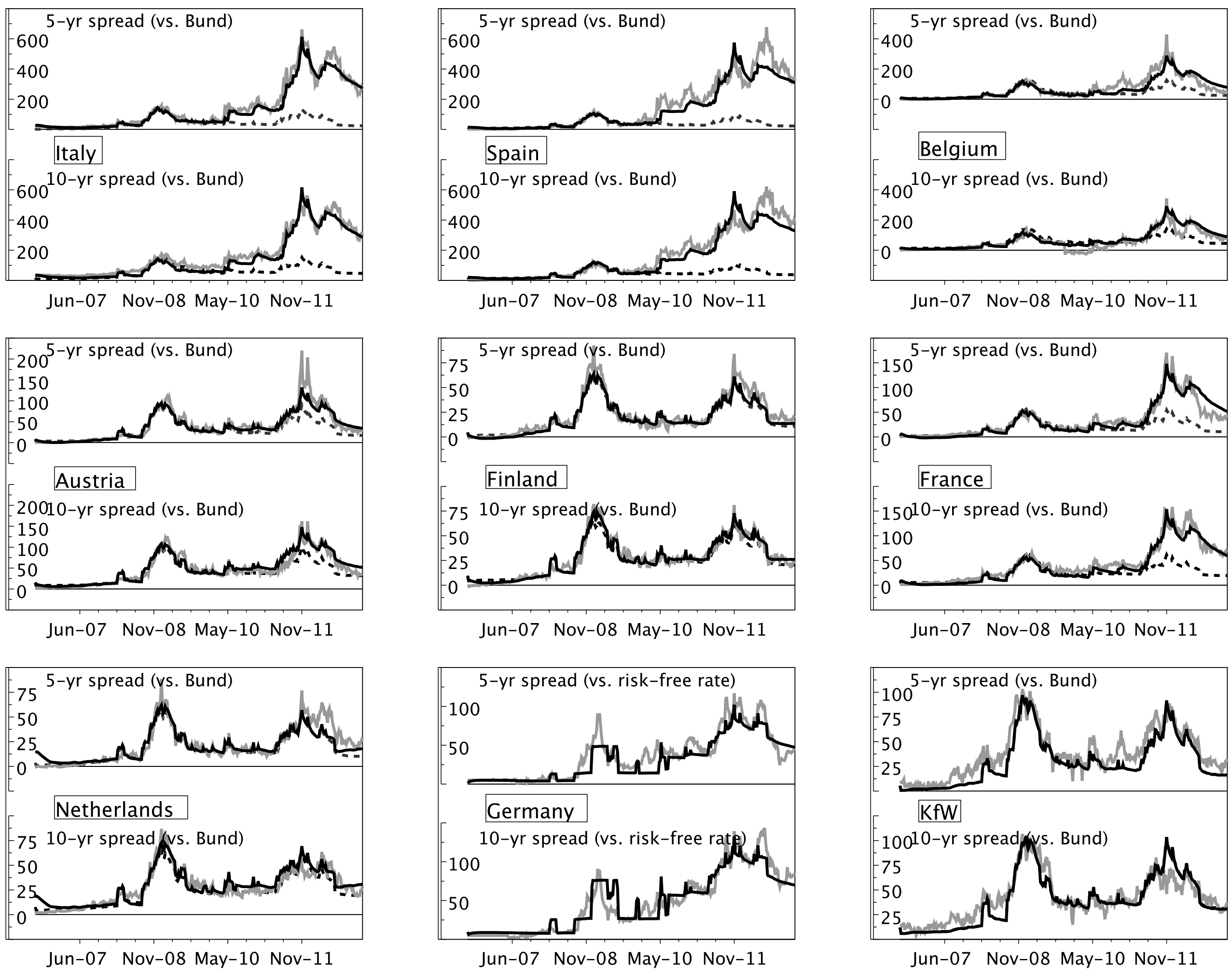
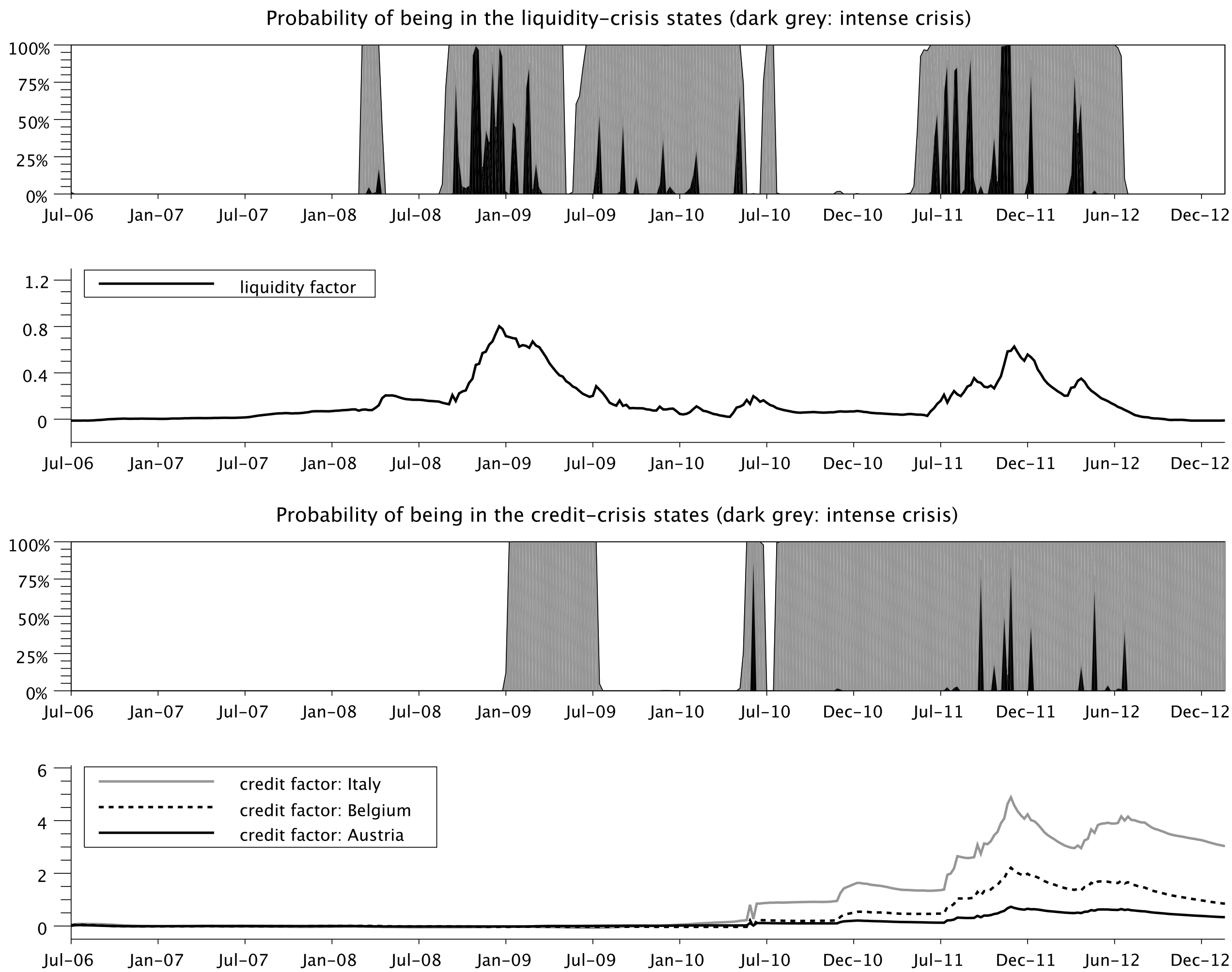


Fig. 3: **Estimated regimes and intensities**

Notes: The first and third panels display the (smoothed) probabilities of being in the different regimes (Kim's (1994) algorithm). For each kind of risk, there are three levels of stress: low (white areas), medium (grey areas) and severe (black areas). The second and fourth panels illustrate the effect of these stress episodes on some intensities (components of λ_t). The liquidity factor is $\lambda^{(0)}_{f\ell,t}$, that is the liquidty fractional-loss intensity associated with KfW. The credit factors that are displayed are some $\lambda^{(n)}_{fc,t}$s.



liquidity-related parts of the spreads turn out to account for a substantial part of the changes in spreads versus Germany, especially for the less indebted countries (the Netherlands and Finland). While the liquidity factor was explaining the main part of the spreads' fluctuations for most of the countries in the post-Lehman period, the part of the spreads explained by credit-related factors became predominant for several countries (Belgium, France, Italy and Spain) over the last two years of the sample.[40]

6.2 Credit and liquidity risk premiums

The previous subsection has illustrated the ability of our approach to decompose the spreads into a liquidity part and a credit one. Each of these two parts can be decomposed into two parts. We refer to these two components as (a) the expectation and (b) the risk-premium parts. The risk premium is the share of the spread that would not exist if investors were risk-neutral. Formally, risk premiums are the differences between the actual spreads and the ones that would be observed if the risk sensitivities (i.e. the δ and ν in the s.d.f. specification, see Equation 6) were null, in which case the risk-neutral measure would coincide with historical one.

Table 6 reports summary statistics for the time series of the four resulting components for the different countries.[41] The average risk premiums are positive and substantially larger than the associated expectation parts. Such a result is consistent with the highly-systematic nature of the risk incorporated in euro-area sovereign bonds.[42] In addition, it can be seen that the relative share of the risk premiums is higher for the best-rated countries (Austria, Finland and Germany notably). This result echoes this of Huang and Huang (2012) who show, by examining U.S. corporate data, that the expectation part of spreads accounts for a smaller fraction of yield spreads for investment-grade bonds than for lower credit-quality bonds.

Credit-risk and liquidity-risk premiums are presented in Figure 4. These plots show that both

[40] The same qualitative result is obtained by Ejsing, Grothe and Grothe (2012) who consider the French-German spread only.

[41] To get, e.g., the credit-risk premiums, (a) we put the liquidity intensities equal to zero and (b) credit-risk premiums are then computed as the differentials between the model-implied spreads an the ones that are obtained when replacing the risk-neutral dynamics by the historical one (i.e. when neutralising the prices of risk). Note that the same comment as the one in Footnote 38 applies (i.e. that, owing to non-linearities, this procedure does not lead to credit- and liquidity-risk premiums that exactly sum to the total risk premiums; however, the differential is negligible).

[42] Indeed, risks that are not systematic, i.e. that can be diversified away, cannot command risk premiums.

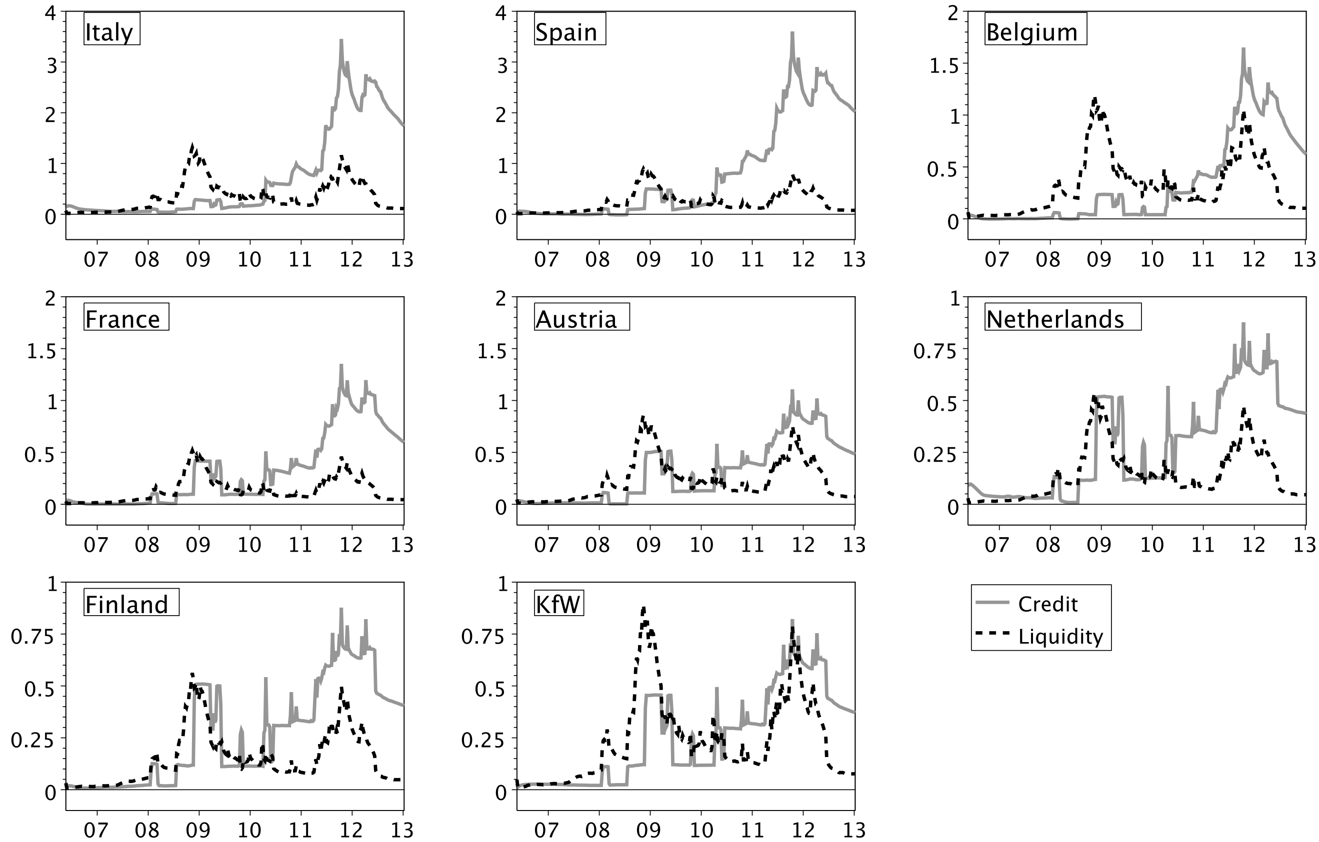
kinds of risks contribute to the fluctuations of total risk premiums. As in the case of the whole spreads (that were shown in Figure 2), while liquidity-risk premiums tended to predominate in late 2008 and 2009, credit-risk premiums appear to be the most important at the end of the sample. This is especially the case for Italy, Spain and, to a lesser extent, France. The credit-risk premiums for the Netherlands, Finland and Germany are very close.[43] Hence, the bulk of the risk premiums incorporated in the Dutch-German and the Finnish-German spread is liquidity driven.

Different studies have highlighted the relationships between international factors reflecting global uncertainty, risk appetite and sovereign credit risk. In particular, stock prices or option-implied volatility indices (e.g. VIX) are found to be particularly correlated with emerging-market sovereign spreads (see e.g. Pan and Singleton, 2008, Hilscher and Nosbusch, 2010 and Longstaff et al., 2011) or with U.S. and European ones (Ang and Longtsaff, 2011). Table 7 reports the results of regressions of the risk premiums on the macro-finance covariates that were also used in Subsection 3.2.3. The R^2s indicate that about the same shares of the expectation and the risk-premium parts are accounted for by the macroeconomic factors. Surprisingly, the coefficients associated with the VSTOXX (an European equivalent of the VIX) is negative and statistically significant.[44] This may stem from colinearity between changes in the EUROSTOXX index and those in the VSTOXX.[45] The interest-rate option-implied volatility (BVOL) are positively and significantly connected to both credit parts of the spreads for most countries. The IBOR-OIS spread is related to the credit expectation part of several countries, which supports the view that sovereign creditworthiness is tightened to the health of the European banking system (see e.g. Ejsing and Lemke, 2011, Kallestrup, Lando and Murgoci, 2012 or Alter and Schüler, 2012). This spread also appears to be related to the liquidity parts of the spreads. While this was already seen in the regression of the KfW-Bund spread on the same variables (see Table 1), these additional regressions show that it is mainly the risk-premium parts (as opposed to the expectation parts) of the bond spreads that are related to the IBOR-OIS variable. This suggests that whereas a rise in interbank risks is not likely to be accompanied by an increase in losses expected from precipitated

[43] The German credit-risk premiums appear on the KfW plot, the two debtors being charecterised by the same credit risk.
[44] Ang and Longstaff (2011) obtain a similar result.
[45] The relationships between the spread components and the VSTOXX become positive in univariate regressions, whose results are not reported here.

Fig. 4: **Credit- and liquidity-risk premiums**

Notes: These plots show the risk premiums incorporated in the spreads between 5-year sovereign yields and a 5-year riskfree rate (the riskfree rate is proxied by the 5-year German bond rate from which we subtract the 5-year German CDS). Credit and liquidity risk premiums are reported. To obtain the credit-risk premiums, we set the liquidity intensity to zero; then, the risk premiums are obtained as the yield differential between model-implied yields (using the $\mathbb{Q}$-dynamics) and the ones computed under the historical ($\mathbb{P}$) measure. The risk premiums are expressed in percentage points. Whereas the German bonds do not incorporate liquidity-risk premiums, they carry credit ones; these are the ones appearing on the KfW chart (it is assumed that KfW and German bonds have the same credit quality).



sales of sovereign bonds, it nevertheless tends to be associated with a rise in the price of such a liquidity risk.

6.3 Default probabilities

In the last part of this paper, we show how our model can be exploited to compute the default probabilities. In the spirit of Litterman and Iben (1991), various methodologies that are widely used by practitioners or market analysts end up with risk-neutral PDs (see, e.g. Chan-Lau, 2006). Our framework makes it possible to assess the potential differences that exist between the latter and their historical, or real-world, counterparts. As stated above (see Subsection 4.2), while the intensities of default are the same processes under both measures, the $\mathbb{P}$- and $\mathbb{Q}$-probabilities of default are not the same because the $\mathbb{P}$- and $\mathbb{Q}$-dynamics of these processes differ.

Tab. 6: **Descriptive statistics for spread components**

Notes: This table reports statistics for the estimated components of the 5-year spreads (versus a proxy for the 5-year zero-coupon riskfree rate). The statistics are in percentage points. The data are weekly and cover the period from 7 July 2006 to 15 February 2013 (346 observations). There are four components in the spreads because the two parts of the spreads –that respectively correspond to the pricing of the two risks (credit and liquidity)– are decomposed into two parts: the first one (the expectation one) being obtained as the spreads that would be observed if investors were risk-neutral, the second one (risk premiums) being the residual part (that reflects the excess yield demanded by the investors to be compensated for the risk carried by the considered bond). By assumption, German bonds are liquidity-risk free.

	Expectation part of the spreads						Risk premiums					
	Credit			Liquidity			Credit			Liquidity		
	Mean	Med.	SD	Mean	Med.	SD	Mean	Med.	SD	Mean	Med.	SD
Austria	0.06	0.01	0.09	0.08	0.10	0.04	0.31	0.18	0.29	0.23	0.17	0.20
Belgium	0.17	0.01	0.26	0.12	0.15	0.06	0.33	0.11	0.42	0.32	0.23	0.27
Finland	0.03	0.01	0.05	0.05	0.07	0.03	0.27	0.14	0.24	0.15	0.11	0.13
France	0.11	0.01	0.16	0.06	0.07	0.03	0.33	0.19	0.35	0.14	0.10	0.12
Germany	0.04	0.02	0.05	-	-	-	0.25	0.15	0.22	-	-	-
Italy	0.49	0.09	0.63	0.10	0.13	0.07	0.74	0.19	0.91	0.35	0.26	0.30
Netherlands	0.04	0.02	0.04	0.05	0.06	0.03	0.28	0.15	0.24	0.14	0.10	0.12
Spain	0.45	0.03	0.63	0.14	0.16	0.05	0.82	0.25	0.98	0.24	0.17	0.20

In our framework, the actual PD between time t and time $t+h$ is given by

$$\begin{aligned}
\mathbb{P}\left(d_{t+h}^{(n)}=1\middle|\underline{W}_t,d_t^{(n)}=0\right) &= E_t^{\mathbb{P}}\left(\mathbb{I}_{\left\{d_{t+h}^{(n)}=1\right\}}\middle| d_t^{(n)}=0\right)\\
&= 1-E_t^{\mathbb{P}}\left(\mathbb{I}_{\left\{d_{t+h}^{(n)}=0\right\}}\middle| d_t^{(n)}=0\right)\\
&= 1-E_t^{\mathbb{P}}\left(\exp(-\lambda_{c,t+1}^{(n)}-\ldots-\lambda_{c,t+h}^{(n)})\right)\\
&\simeq 1-E_t^{\mathbb{P}}\left(\exp\left\{\frac{1}{1-\zeta}(-\lambda_{fc,t+1}^{(n)}-\ldots-\lambda_{fc,t+h}^{(n)})\right\}\right), \qquad (15)
\end{aligned}$$

where the last approximation stems from Equation (3). Thus, up to this approximation, the survival probability, which is the expectation term on the right-hand side of Equation (15), turns out to be a multi-horizon Laplace transform of a compound auto-regressive process of order one. Then, in the same way as for the yields, a simple recursive algorithm can be used to calculate default probabilities. In the computations, we use a constant recovery rate ζ of 50%, which corresponds to the average of the recovery rates observed for sovereign defaults over the last decade (see Moody's, 2010).

Figure 5 shows the model-based 5-year probabilities of default (i.e. the probabilities that the considered countries will default in the next 5 years). One-standard-deviation bands are also reported. These standard deviations take two kinds of uncertainty into account: (1) the

Tab. 7: **Spread components and macro-finance indicators**

Notes: This table presents the results of regressions of the different components of the 5-year-maturity spreads on a set of macro-financial variables. The spreads are against a riskfree rate, proxied by the German rates from which we subtract German CDS. In our framework, spreads can be decomposed into four components (credit/liquidity and expectation/risk premium). Since the liquidity components are driven by a single factor ($\lambda_{\ell,t}$), we do not report the regressions for all the liquidity-related components, but only for KfW (the ones associated with the other debtors are approximately equal to these of KfW multiplied by the parameters reported in the row "$(1-\theta^{(n)})/(1-\theta^{(0)})$" of Table 4). Panel A (Panel B) presents the regressions of the expectation (risk premium) parts of the spreads. The VSTOXX (implied volatility in EURO STOXX 50 option prices), the BVOL (10-year-in-10-year interest-rate-swap Black volatility), the 3-month overnight index swap (OIS) and the ESI (business-cycle index) are expressed in percentage points, the EURO STOXX 50 in hundreds of points and the 3-month-maturity EURIBOR-OIS spread in basis points. The data are weekly and cover the period from 7 July 2006 to 15 February 2013 (346 observations). Newey-West (12 lags) corrected standard deviations of the parameter estimates are reported in brackets. The regressions include intercepts that are not reported in the table.

Panel A – Expectation components (5-year maturity)

	Liquidity	Credit							
	KfW	GER	AUT	BEL	FIN	FRA	ITA	NTH	SPA
VSTOXX	-0.01	-0.04*	-0.09*	-0.27*	-0.04	-0.15*	-0.49	-0.01	-0.33
	(0.04)	(0.03)	(0.06)	(0.15)	(0.03)	(0.09)	(0.34)	(0.03)	(0.29)
EUROSTX	-0.24*	-0.07	-0.17	-0.6	-0.07	-0.3	-0.88	-0.05	-0.67
	(0.15)	(0.07)	(0.14)	(0.41)	(0.07)	(0.25)	(0.95)	(0.11)	(0.96)
IBOR-OIS	0	0.02*	0.03*	0.11*	0.01	0.06*	0.23*	0.01	0.18*
	(0.01)	(0.01)	(0.02)	(0.06)	(0.01)	(0.03)	(0.12)	(0.01)	(0.1)
3-mth OIS	-0.61	0.42	0.39	2.56	0.53	1.6	4.77	-0.16	5.41
	(0.94)	(0.47)	(0.97)	(2.77)	(0.52)	(1.8)	(6.64)	(0.57)	(5.82)
BVOL	0.02	0.34***	0.65***	1.73***	0.3***	1.24***	4.14***	0.24***	3.51***
	(0.09)	(0.09)	(0.18)	(0.54)	(0.09)	(0.32)	(1.16)	(0.09)	(0.96)
ESI	-0.11	-0.04	0.01	-0.05	-0.04	-0.09	-0.3	-0.01	-0.34
	(0.1)	(0.04)	(0.07)	(0.23)	(0.04)	(0.15)	(0.55)	(0.05)	(0.51)
Adj. R^2	29	34	31	31	30	35	31	21	28

Panel B – Risk premiums (5-year maturity)

	Liquidity	Credit							
	KfW	GER	AUT	BEL	FIN	FRA	ITA	NTH	SPA
VSTOXX	-0.18	-0.29*	-0.4**	-0.5**	-0.32*	-0.43**	-0.8*	-0.29*	-0.73*
	(0.19)	(0.15)	(0.19)	(0.23)	(0.18)	(0.18)	(0.46)	(0.18)	(0.4)
EUROSTX	-1.37*	-2.04**	-2.51**	-1.94**	-2.36**	-2.31**	-2.28*	-2.39**	-3.21**
	(0.79)	(0.86)	(1.03)	(0.81)	(0.99)	(0.91)	(1.35)	(1)	(1.55)
IBOR-OIS	0.20***	0	0.01	0.14*	-0.01	0.06	0.31*	-0.02	0.22
	(0.06)	(0.05)	(0.06)	(0.09)	(0.06)	(0.07)	(0.17)	(0.06)	(0.16)
3-mth OIS	-4.75	0.58	0.54	3.56	0.72	2.22	6.67	-0.22	7.53
	(4.05)	(4.6)	(5.53)	(5.02)	(5.36)	(5.27)	(9.68)	(5.42)	(10.25)
BVOL	0.64	0.66	1.12*	2.49***	0.64	1.89***	5.77***	0.57	5.04***
	(0.58)	(0.46)	(0.58)	(0.83)	(0.51)	(0.62)	(1.62)	(0.48)	(1.36)
ESI	-0.5	-0.21	-0.16	-0.11	-0.24	-0.25	-0.38	-0.2	-0.54
	(0.43)	(0.4)	(0.48)	(0.38)	(0.47)	(0.39)	(0.72)	(0.46)	(0.65)
Adj. R^2	50	28	30	34	27	36	34	27	33

smoothing errors that are associated with the Kim's (1994) smoothing algorithm used to estimate the intensities λ_t and (2) the uncertainty stemming from the parameters' estimation (MLE).[46]

Figure 6 presents the model-implied term-structure of PDs as of 1 January 2010 and 30 December 2011. This Figure illustrates the dramatic changes in the term-structure of PDs that took place over these 2 years. For all countries, the term-structure of the PDs is much higher and steeper in late 2011 than in early 2010.

The gap between actual PDs and their risk-neutral counterparts is significant in many cases, particularly during the most recent years (see Figure 5). Nevertheless, as stated above, risk-neutral PDs are extensively used by market practitioners and analysts. This mainly stems from the fact that risk-neutral PDs are relatively easy to compute, using basic methods inspired by the one proposed by Litterman and Iben (1991) [47]. To illustrate, Figure 7 compares the PDs estimates derived from our model with alternative estimates, as of the end of 2011. Two kinds of alternative estimates are considered: (a) PDs that are based on the Moody's credit ratings and the associated matrix of long-run credit-rating-migration probabilities and (b) risk-neutral probabilities computed by CMA Datavision (2011). This figure shows that our estimates lie somewhere between the two others.[48] In addition, it appears that our risk-neutral PDs (the triangles) are relatively close to the risk-neutral CDS-based ones computed by CMA.[49]

7 Conclusion

In this paper, we present a multi-country no-arbitrage model of the joint dynamics of euro-area sovereign spreads. At the heart of our model is an innovative approach capturing the intertwined dynamics of credit- and liquidity-related crises by a joint modeling of two kinds of switching regimes. These crises are key drivers of, respectively, credit and illiquidity intensities associated

[46] The computation of these standard errors is inspired from Hamilton (1986). It relies on the assumption that the two kinds of errors (smoothing and MLE) are independent from each other.

[47] In particular, these methods do not care about liquidity-pricing effects.

[48] Credit-rating-based PDs are extremely small (for instance: 6.10^{-6} for a AAA-rated countries, 6.10^{-4} for a A-rated countries). This reflects the fact that transition probabilities are based on past 25-year history of rating changes, during which quick sovereign downgrades were infrequent (contrary to during the current crisis period).

[49] The remaining differences between the latter two risk-neutral estimates may e.g. be attributed to (i) the measurement errors of our methodology (the $\xi_{S,t}$ in Equation 12, that are the fitting errors that can be seen in Figure 2), (ii) the absence of treatment of liquidity-pricing effects in the CMA methodology (while empirical evidence suggests that CDS contain liquidity premiums, see Buhler and Trapp, 2008) or, more generally, those reasons that are invoked to account for the non-zero CDS-bond basis (see e.g. Bai and Collin-Dufresne, 2012) .

with the different issuers. Using euro-area spread data covering the last five years, we estimate such intensities for eight euro-area countries. Interestingly, we provide evidence of causal relationships between credit- and liquidity-stress periods.

Our approach makes it possible to exhibit the part of the spreads reflecting liquidity-pricing effects. A key assumption is that the country-specific illiquidity intensities perfectly comoves, that is, that there exists a single European liquidity-pricing factor. The identification of the latter is based on the term structure of the yield differentials between the bonds issued by KfW (a German agency) and the German sovereign bonds (the *Bunds*). Indeed, KfW's liabilities are explicitly and unconditionally guaranteed by the Federal Republic of Germany. Therefore, the KfW-*Bund* spread should be essentially liquidity-driven. Our results indicate that a substantial part of intra-euro spreads is liquidity-driven.

Given some assumptions regarding the recovery process, our framework makes it possible to decompose the credit part of the spreads between actual, or real-world, probabilities of default on the one hand and risk premiums on the other hand. To that respect, our results suggest that actual PDs are significantly lower than their risk-neutral counterparts. According to these results, relying on risk-neutral PDs to assess the market participants expectations regarding future sovereign defaults would be misleading.

Fig. 5: **Default probabilities estimates (5-year horizon)**

Notes: Notes: These plots display the model-implied 5-year default probabilities computed under both measures (risk-neutral: dotted line, historical: solid line). Formally, the dotted line corresponds to the time series of $E_t^{\mathbb{Q}}(\mathbb{I}\{d_{t+5yrs}^{(n)} = 1\}|d_t^{(n)} = 0)$, where E_t denotes the expectation (under the historical measure) conditional to the information available at time t (see Section 6.3 for the computation of these default probabilities). The black solid line represents$E_t^{\mathbb{P}}(\mathbb{I}\{d_{t+5yrs}^{(n)} = 1\}|d_t^{(n)} = 0)$. Two-standard-deviation bands are reported. These standard deviations account for smoothing errors (associated to Kim's smoothing algorithm, 1994) as well as uncertainty related to the parameter estimates, following Hamilton's (1986) approach. The y-axis scales differ across countries.

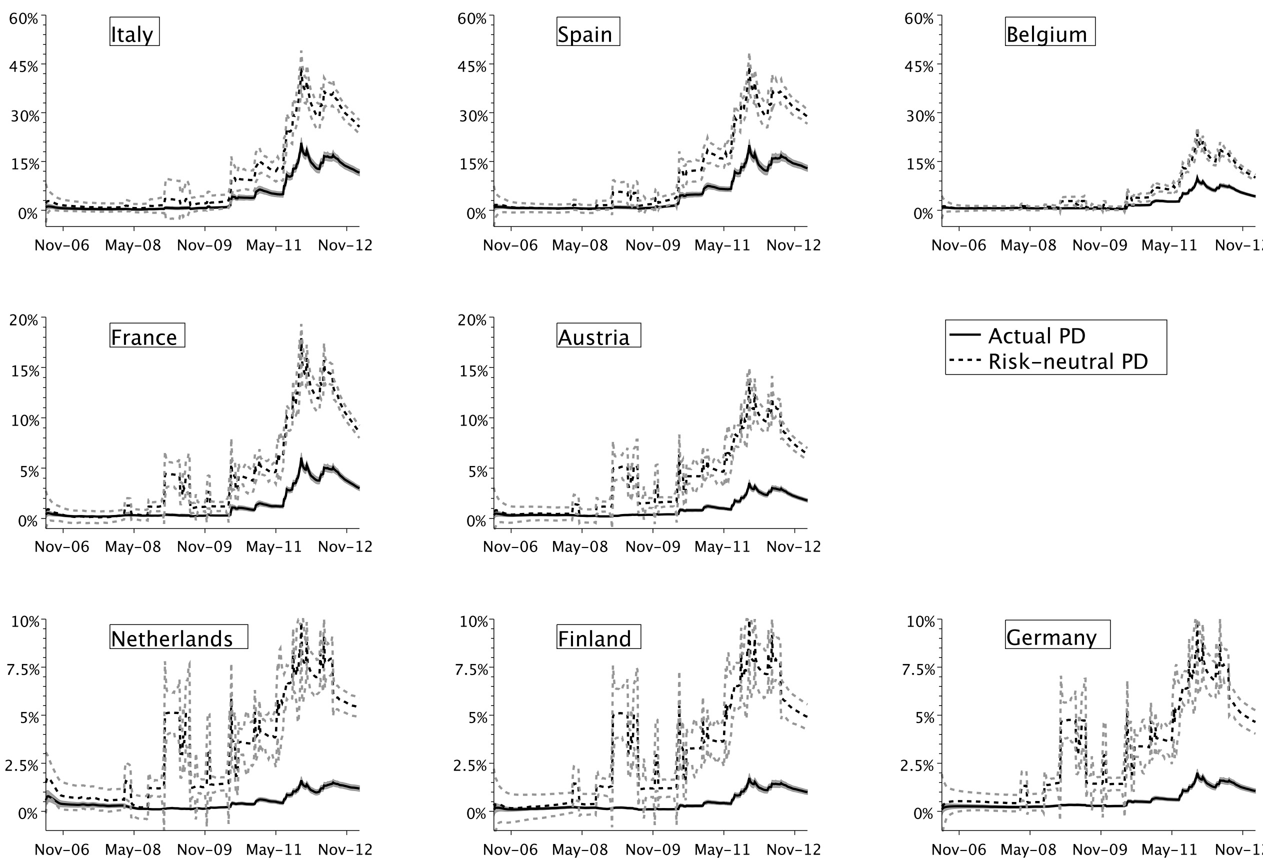


Fig. 6: **Term structure of default probabilities**

Notes: These plot display the term structure of the default probabilities for two distinct dates. Formally, for date t and debtor n, the plot shows $E_t^{\mathbb{P}}(\mathbb{I}\{d_{t+h}^{(n)} = 1\}|d_t^{(n)} = 0)$ for h between 1 month and 10 years (where $E_t^{\mathbb{P}}$ denotes the expectation –under the historical measure– conditional to the information available at time t). The grey lines delimit the ± 2 standard deviation area. These standard deviations account for smoothing errors (associated to Kim's smoothing algorithm, 1994) as well as uncertainty related to the parameter estimates, following Hamilton's (1986) approach. The y-axis scales differ across countries.

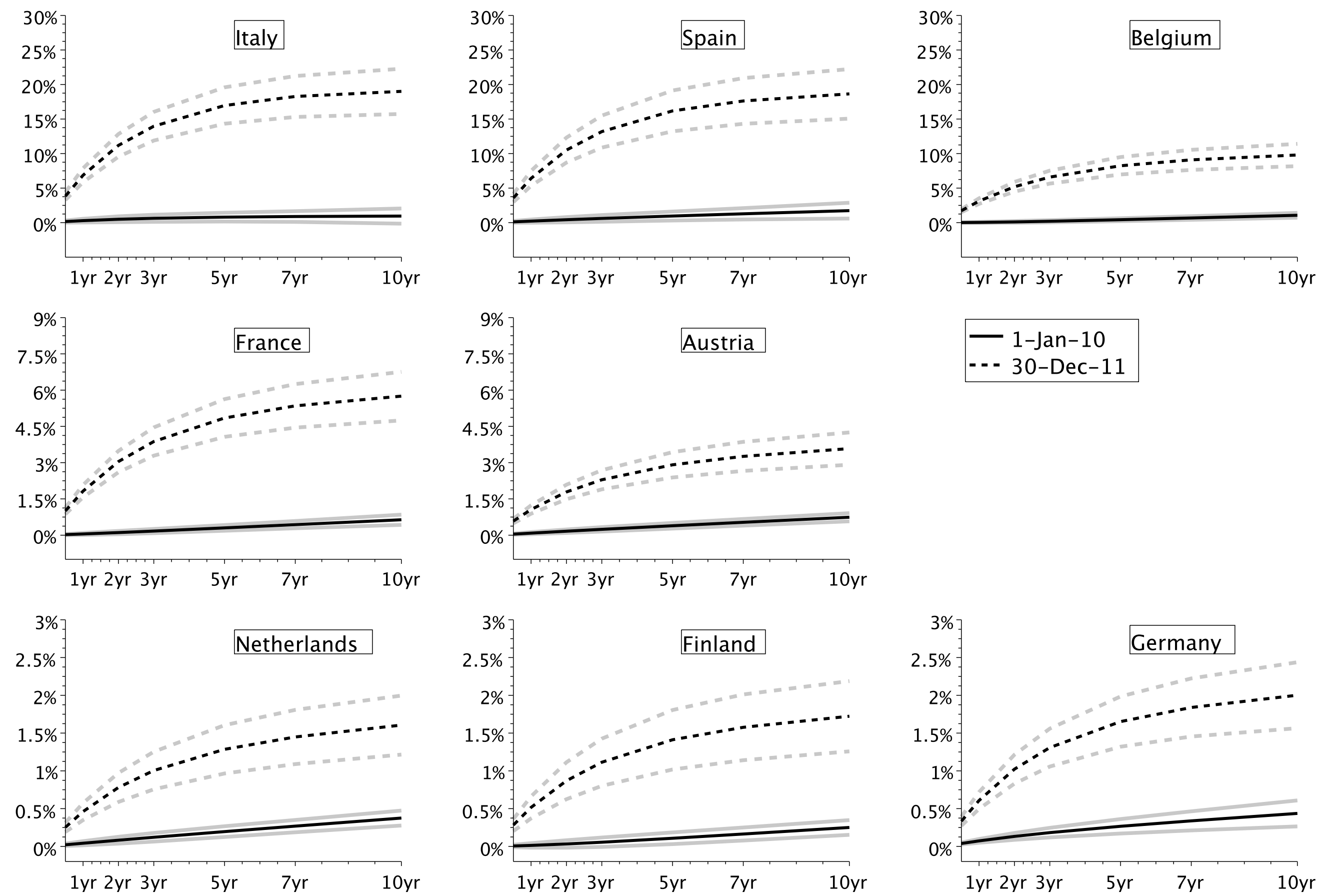
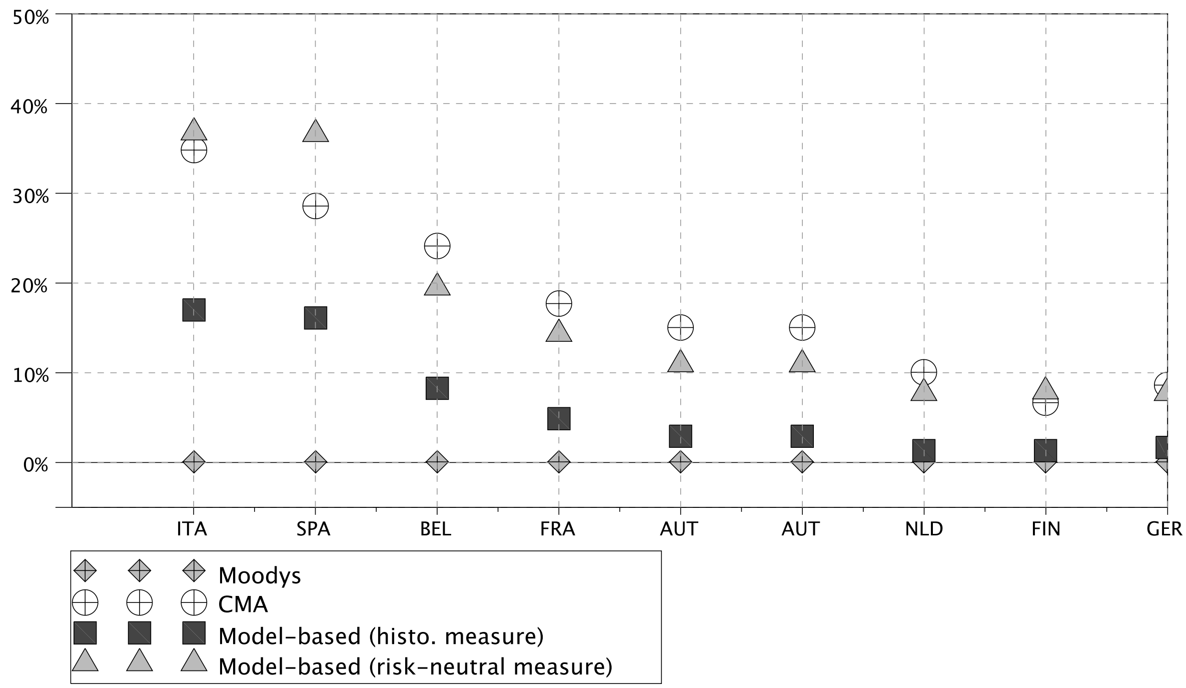


Fig. 7: **Default probabilities estimates (5-year horizon)**

Notes: This plot displays different estimates of probabilities of default (PD) of 8 euro-area governments (as of 30 December 2011). The squares and the triangles correspond to outputs of our model. While the squares indicate "real-world" PDs (i.e. the default probabilities obtained under the physical, or historical, measure), the triangles are risk-neutral PDs. The circles indicate the PDs computed by CMA, using an industry standard model and proprietary CDS data from CMA Datavision (2011). The diamonds correspond to PDs that derive from (a) the Moody's' ratings of the countries (as of 2011Q4) and (b) the matrice of credit-rating-migration probabilities given by Moody's (2010).



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A Data sources and preliminary treatment

Sovereign zero-coupon yields are extracted from Bloomberg. Zero-coupon yields derived from KfW bond prices come from the Thomson Reuters Tick History database. The estimation dataset is completed by 12-month-ahead forecasts of 10-year sovereign yields for France, Germany, Italy, Spain and the Netherlands. These forecasts are the mean values of the respondents' forecasts by the *Consensus Economics*' expert panel. The survey is released around the middle of the month. Note that the survey implicitly targets yields-to-maturity of coupon bonds and not zero-coupon bonds. However, our zero-coupon yields remain very close to coupon yields over the estimation sample. The remaining discrepancy, of a few basis points, are attributed to the deviation between the survey-based forecasts and the model-based ones (the $\xi_t^{(n)}$'s introduced in equation 13). This monthly series is converted into a weekly one using a cubic spline.

B Pricing of defaultable bonds

For the sake of notational convenience, we drop the issuer superscript n in this Appendix.

Let us consider the price $B_{T-1,1}$, at date $T-1$, of a one-period bond issued by the debtor (before $T-1$). If the debtor is not in default at $T-1$, then:

$$B_{T-1,1} = \exp(-r_{T-1})E^{\mathbb{Q}}\left[(1-d_T+\zeta d_T)(1-\ell_T+\theta\ell_T) \mid \underline{W}_{T-1}\right]$$

($\underline{W}_{T-1}$ containing the information $d_{T-1}=0$)

$$\begin{aligned}
&= \exp(-r_{T-1})E^{\mathbb{Q}}\left[E^{\mathbb{Q}}\left((1-d_T+\zeta d_T)(1-\ell_T+\theta\ell_T) \mid w_T, \underline{W}_{T-1}\right)\middle| \underline{W}_{T-1}\right]\\
&= \exp(-r_{T-1})E^{\mathbb{Q}}\left[\left\{\exp\left(-\lambda_{c,T}^{(n)}\right)+\zeta\left(1-\exp\left(-\lambda_{c,T}^{(n)}\right)\right)\right\}\times\right.\\
&\qquad \left.\left\{\exp\left(-\lambda_{\ell,T}\right)+\theta\left(1-\exp\left(-\lambda_{\ell,T}\right)\right)\right\}\right)\left| \underline{W}_{T-1}\right].
\end{aligned}$$

The last equality is obtained by using the conditional independence of d_t and ℓ_t and the expressions of the conditional $\mathbb{Q}$-distributions of d_t and ℓ_t (that are the same as their historical counterparts, see Subsection 4.2). From that, using the expressions of the fractional-loss intensities $\lambda_{fd,T}$ and $\lambda_{f\ell,T}$ given in formula (2), it follows that:

$$\begin{aligned}
B_{T-1,1} &= \exp(-r_{T-1})E^{\mathbb{Q}}\left[\exp(-\lambda_{fc,T}-\lambda_{f\ell,T}) \mid \underline{W}_{T-1}\right]\\
&= \mathcal{B}(r_{T-1}, z_{T-1}, \lambda_{T-1}) \quad \text{(say)}
\end{aligned}$$

The fact that $E^{\mathbb{Q}}\left[\exp(-\lambda_{fc,T}-\lambda_{f\ell,T}) \mid \underline{W}_{T-1}\right]$ is a function of (z_{T-1}, λ_{T-1}) originates from the assumptions on the distribution of (z_T, λ_T) given $\underline{W}_{T-1}$, in particular the non-causality from (d_t, ℓ_t) to (z_t, λ_t) under $\mathbb{Q}$.

Let us then consider $B_{T-2,2}$, we have:

$$B_{T-2,2} = \exp(-r_{T-2})E^{\mathbb{Q}}\left[(1-d_{T-1}+\zeta d_{T-1})(1-\ell_{T-1}+\theta\ell_{T-1})\mathcal{B}(r_{T-1}, z_{T-1}, \lambda_{T-1})\middle| \underline{W}_{T-2}\right],$$

$\underline{W}_{T-2}$ containing the information $d_{T-2}=0$.

Conditioning first by $(w_{T-1}, \underline{W}_{T-2})$ and using the fact that $\mathcal{B}_{T-1}(z_{T-1}, \lambda_{T-1})$ only depends on $\mathcal{I}_{T-1}$ only (i.e. not on (d_{T-1}, ℓ_{T-1})), we get:

$$B_{T-2,2} = E^{\mathbb{Q}}\left[\exp(-r_{T-2}-\lambda_{fc,T-1}-\lambda_{f\ell,T-1})\mathcal{B}(r_{T-1}, z_{T-1}, \lambda_{T-1})\middle| \underline{W}_{T-2}\right].$$

Replacing $\mathcal{B}(r_{T-1}, z_{T-1}, \lambda_{T-1})$ by $E^{\mathbb{Q}}\left[\exp(-r_{T-1} - \lambda_{c,T} - \lambda_{\ell,T}) \mid \underline{W}_{T-1}\right]$ and using the fact that $\exp(-r_{T-2} - \lambda_{fc,T-1} - \lambda_{f\ell,T-1})$ is function of $\underline{W}_{T-1}$, we get:

$$\begin{aligned} B_{T-2,2} &= E^{\mathbb{Q}}\left[E^{\mathbb{Q}}\left(\exp(-r_{T-2} - \lambda_{fc,T-1} - \lambda_{f\ell,T-1} - r_{T-1} - \lambda_{fc,T} - \lambda_{f\ell,T}) \mid \underline{W}_{T-1}\right) \middle| \underline{W}_{T-2}\right] \\ &= E^{\mathbb{Q}}\left[\exp(-r_{T-2} - \lambda_{fc,T-1} - \lambda_{f\ell,T-1} - r_{T-1} - \lambda_{fc,T} - \lambda_{f\ell,T}) \middle| \underline{W}_{T-2}\right]. \end{aligned}$$

Applying this methodology recursively leads to equation (8).

C Parameter constraints

C.1 Parameterization of the matrix of transition probabilities Π

Matrix Π defines the dynamics of z_t where $z_{\ell,t} \otimes z_{c,t}$ (see Subsection 4.1). First, we assume that there is no instantaneous causality between $z_{c,t}$ and $z_{\ell,t}$, meaning that conditionally to z_{t-1}, $z_{c,t}$ and $z_{\ell,t}$ are independent. Second, whereas the switching probabilities of the liquidity-regime variable $z_{\ell,t}$ (respectively $z_{c,t}$) between date $t-1$ and date t may be influenced by the existence of a credit crisis (resp. liquidity crisis) at date $t-1$, it does not depend on the *distinction* between the two highest credit-stress (resp. liquidity-stress) levels. Third, the probabilities of switching from one of the two (credit or liquidity) severe crisis states to the respective low-stress regimes is zero, as well as the opposite (the first crisis levels act as intermediary regimes between the low-stress and the severe-crisis states). Fourth, the probability of remaining in the severe credit-crisis (resp. liquidity-crisis) states does not depend on the liquidity (resp. credit) state. With these restrictions, 14 parameters are required to define the matrix Π.

C.2 The size of the Gaussian shocks

The standard deviations of the Gaussian shocks ε_t entering Equation (4) are constrained to make sure that the regime variables z_t are the main sources of the spreads fluctuations. If such constraints are not imposed, most of the spread fluctuations tend to be accounted for by the Gaussian shocks. This phenomenon, that reflects that Gaussian shocks are more flexible than the discrete-numbered regimes to fit the spreads, has two undesirable implications within our framework: (i) the higher the standard deviation of the Gaussian shocks, the higher the frequency of generating/estimating negative intensities, (ii) the lower the importance of the regime variables, the less information about the relationships between liquidity- and credit-stress periods the estimation is brought to reveal. Accordingly, we constrain the parameters to be such that a limited part of the (unconditional) fluctuations of the intensities is accounted for by Gaussian shocks. Practically, we impose the following constraints on the parameter estimates: $\sigma_i/\sqrt{1-\rho_i^2} \leq 10\%\tilde{\sigma}_i$, where $\tilde{\sigma}_i$ is the sample standard deviation of the (observed) spreads associated with entity i (and where σ_i is expressed in the same unit as the spreads). This calibration implies unconditional distribution of the intensities that is consistent with mainly positive intensities. Alternative estimation (with ratios of 5% and 20%) suggest that the qualitative results presented above are fairly robust to changes in this 10% ratio.

C.3 The standard deviations of the measurment errors ξ

For a given debtor, the standard deviations of the spread pricing errors (gathered in the vector $\xi_{S,t}$, see equation 12) are assumed to be the same across maturities. However, they differ across countries, proportionally to the standard deviations of the observed spreads (the proportionality coefficient being estimated by the MLE).

C.4 Matrix μ

The entries of matrix μ are non-zero only for severe-stress regimes. Indeed, preliminary estimations, that allowed for non-zero entries associated with intermediary-stress regimes, suggested that these were not significant. Therefore, we have:

$$\mu = \left[\begin{array}{ccccccccc} 0 & 0 & \mu_c^{(1)} & 0 & 0 & \mu_c^{(1)} & 0 & 0 & \mu_c^{(1)} \\ \vdots & \vdots & \vdots & \vdots & \vdots & \vdots & \vdots & \vdots & \vdots \\ 0 & 0 & \mu_c^{(N)} & 0 & 0 & \mu_c^{(N)} & 0 & 0 & \mu_c^{(N)} \\ 0 & 0 & 0 & 0 & 0 & 0 & \mu_\ell & \mu_\ell & \mu_\ell \end{array}\right]'.$$

D Structural interpretation of the liquidity-related fractional costs $1-\theta^{(n)}$

In that Appendix, broadly following the lines of Ericsson and Renault (2006), we develop a structural interpretation of the fractional cost $1-\theta^{(n)}$. While this fractional cost is considered to be a fixed parameter in the main text, it can be seen as the expectation of a random process that may be caused but does not Granger-cause the information W_t.

Upon the arrival of the liquidity shock ($\ell_t = 1$), the investor has to exit by selling her bond holdings. This liquidation has to be done in a limited period of time, between t and $t+\varepsilon$, say (where $\varepsilon << 1$). During that period, the bond seller obtains a random number K of offers from traders. Conditionally on $\ell_t = 1$, this number is drawn from a Poisson distribution: $K \sim \mathcal{P}(\gamma^{(n)})$. The expected number of offers $\gamma^{(n)}$ reflects the liquidity of the bonds issued by debtor n (the more liquid the bond, the higher $\gamma^{(n)}$). Each offer is a random fraction ω_i ($i \in \{1,\dots,K\}$) of $B_{t,h}^{(n)}$. The ω_i's are uniformly distributed in $[0,1]$.

At $t+\varepsilon$, the bond is sold to the trader that has offered the highest price. Formally, when $\ell_t = 1$, the selling price is:

$$\max_{i\in\{1,\dots,K\}}(\omega_i)B_{t,h}^{(n)}.$$

It is easily shown that, conditional on K, the expectation of $\max_i(\omega_i)$ is equal to equal to $K/(K+1)$.[50] Therefore, $\theta^{(n)}$ –which is the unconditional expectation of $\max_i(\omega_i)$– is given by:

$$\begin{aligned} \sum_{k=1}^{\infty} e^{-\gamma^{(n)}} \frac{\left(\gamma^{(n)}\right)^k}{k!}\frac{k}{k+1} &= \left[\left(1-e^{-\gamma^{(n)}}\right)\frac{\gamma^{(n)}-1}{\gamma^{(n)}} + e^{-\gamma^{(n)}}\right] \\ &= g(\gamma^{(n)}), \qquad \text{say.} \end{aligned}$$

The function g is monotonically increasing and is valued in $[0,1]$.

[50] This is easily obtained from the fact that the cumulative distribution function of $\max(\omega_1,\dots,\omega_K)$ is $x \mapsto F(x)^K$ where $F$is the cumulative distribution function of some i.i.d. random variables ω_i.