The Term Structure of Bond Market Liquidity Conditional on the Economic Environment: An Analysis of Government Guaranteed Bonds[†]

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Current Version: January 2012

JEL Classification: G01, G11, G12, G13

Keywords: bond liquidity, illiquidity premia, term structure of liquidity, liquidity risk, regime-switching, financial crisis, flight-to-liquidity, systematic risk

[†]We thank seminar participants at the workshop Financial Markets & Risk 2010 (University Innsbruck), a workshop at Deutsche Bundesbank in September 2011, the DGF PhD Workshop 2011 as well as Ralf Elsas, Rüdiger Fahlenbrach, Alexander Kempf, Olaf Korn, Stefan Ruenzi, Christian Schlag, Erik Theissen, Nils Unger, and Josef Zechner for helpful comments and suggestions.

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Abstract

This paper analyzes the term structure of illiquidity premia as the difference between the zero coupon yield curves of two homogeneous bond classes that differ only in their liquidity: German government bonds and bonds of the Kreditanstalt für Wiederaufbau (KfW) which are guaranteed by the German government. We show that characteristics of the term structure of illiquidity premia depend strongly on the financial and economic situation. In crisis times, illiquidity premia are higher with the largest increase for short-term maturities. Moreover, the reaction of illiquidity premia to changes in fundamentals is only significant (and also significantly stronger) in crisis times: Premia of all maturities depend on the inventory risk of market makers supplying liquidity. Additionally, short-term premia are highly sensitive to liquidity preferences consistent with the flight-to-quality/liquidity hypothesis. This suggests that calibrating risk management models in normal times underestimates the systematic component of illiquidity risk.

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1 Introduction

It is consensus in the literature that a large part of the yield spread compensates investors for the illiquidity of a bond (see e.g. Longstaff, Mithal, and Neis (2005)). Do illiquidity premia behave substantially differently in different economic periods? Or are they driven by the same economic determinants independent of the economic environment? To what extent do such drivers affect different parts of the term structure of illiquidity premia in a different manner? A deep understanding of the characteristics of illiquidity premia is of key importance given both the enormous and rapidly increasing size of bond markets (the outstanding volume of corporate and government bonds sums up to 16.1 tr. EUR in April 2011 only within the Eurozone) and their role in the economy as a major source of financing.

The majority of empirical studies of illiquidity premia in bond markets analyzes average effects both with respect to the economic environment and the bonds' maturity. Thus, they are silent on the above issues. Moreover, many studies suffer from the problem that empirically disentangling risk premia due to illiquidity from other systematic factors such as default risk is a tedious task and often suspect to strong assumptions.

To contribute to a more comprehensive understanding of the nature of illiquidity premia we study their term structures with Hamilton's regime-switching approach. We focus on two completely homogeneous bond market segments that differ only in their liquidity: German government bonds (BUNDs) and explicitly government guaranteed bonds issued by the Kreditanstalt für Wiederaufbau (KfW). While constituting major bond market segments in the Eurozone, an important advantage of our data is that we do not need to disentangle liquidity from credit risk. In this clean environment, we are able to isolate the term structures of illiquidity premia and to study their drivers conditional on the state of the economy based on 15 years of data from 1996 to 2010.

Three main results emerge from the analysis. First, the regime-switching approach applied on an autoregressive model of illiquidity premia of different maturities identifies the 1998 bailout of Long Term Capital Management (LTCM), the period after the burst of the dot-com bubble as well as the financial crisis starting in summer 2007 as liquidity stress periods in the European bond market.

Second, we find that the term structure of illiquidity premia varies over time and is strongly dependent on the general financial and economic situation. In normal times, the average illiquidity premium measured as the extra yield to maturity of an illiquid KfW bond compared to the liquid BUND is around 15 bps. This premium nearly doubles in times of stress for all maturity segments but the increase is most prevalent at the short end. Thus, term structures of illiquidity premia in times of crisis are often strongly downward sloping.

Third, we find that none of our economic drivers plays a major role in explaining premia in normal times. In contrast, factors accounting for the degree of bond illiquidity and factors accounting for preferences for liquidity are important in periods of stress. While option-implied interest rate volatilities which proxy for the degree of illiquidity have significant explanatory power for all maturity segments, preferences for liquidity drive the short end only. This finding is consistent with the clearly higher short-term premia and can be explained with flight-to-liquidity periods that coincidence with stress periods. An increased demand for short-term and highly liquid securities within these periods leads first to a strongly increased level of illiquidity premia and second to a stronger influence of effects stemming from liquidity demand.

Overall, the regime-switching nature of the term structure of illiquidity premia documented in this study goes well together with the theoretical insights of Brunnermeier and Pedersen (2009) that the impact of changes in fundamentals on illiquidity is significantly stronger when funding is scarce and the system is in stress. This implies that calibrating e.g. risk management models in normal times, where illiquidity premia are largely invariant to changes in fundamentals, heavily underestimates the systematic component of liquidity risk. Moreover, preferences for liquidity appear only in short-term premia making it worthwhile to incorporate term structure effects.

Despite the importance of these relationships, empirical research in this area has been limited. Empirical evidence that liquidity effects in bond markets are conditional on the state of the economy include the important work of Brunnermeier (2009), Dick-Nielsen, Feldhütter, and Lando (2012), and Acharya, Amihud, and Bharath (2010). While Brunnermeier (2009) and Dick-Nielsen, Feldhütter, and Lando (2012) clearly document a different behavior of liquidity during crisis and non-crisis times, Acharya, Amihud, and Bharath (2010) explicitly identify two liquidity regimes and document the regime-dependent importance of liquidity betas. Given their focus on corporate bonds, all these studies suffer from the separation between liquidity and credit risk. Moreover, the term structure of illiquidity premia is outside the scope of these papers.

Some recent research focuses on the term structure of illiquidity premia. Longstaff (2004) studies yield differences between Treasuries and Refcorp bonds. He finds a positive influence of changes in the amount of funds held in money market mutual funds on short-term illiquidity premia, whereas medium-term premia increase when consumer confidence declines. Both effects can be interpreted in the way that an increased wariness to bear risk increases the premium of holding an illiquid bond. Kempf, Korn, and Uhrig-Homburg (2011) estimate the term structure of illiquidity premia for German Pfandbriefe. They find a positive influence of short- and long-term liquidation needs on the respective illiquidity premia – the former proxied by asset market volatilities, the latter directly linked to a deteriorating economic outlook. Both studies do not pursue a conditional approach, rather they analyze 'average' effects.

Finally, Goyenko, Subrahmanyam, and Ukhov (2011) study the term structure of Treasury market illiquidity. Although their focus is on bond market trading cost measured via bid-ask spreads instead of yield differentials, their findings of increasing illiquidity in recessions across all maturities with the increase being especially pronounced for short-term bonds is pretty much consistent with our insights on the premium side.

The remainder of the paper is structured as follows: In Section 2, we provide details of the data set and introduce our approach to extract the term structure of illiquidity premia. The section also presents the evolution of illiquidity premia and different shapes of the term structure over time. In Section 3, we first motivate our conditional approach from an empirical and theoretical perspective. Section 3.2 then applies the Markov regime-switching methodology on an autoregressive model of illiquidity premia of short-, medium-, and long-term maturities. This procedure yields different shapes of the term structure during crisis and non-crisis times. Section 3.3 identifies economic drivers influencing the term structure of illiquidity premia conditional on the crisis and non-crisis regime. Section 4 performs several robustness checks. Section 5 concludes the paper.

2 Illiquidity premia

We extract the illiquidity premium by estimating the zero coupon yield curves of two bond market segments that differ only in their liquidity: highly liquid German government bonds (BUNDs) and more illiquid bonds guaranteed by the German government but issued by the Kreditanstalt für Wiederaufbau (KfW). We interpret the difference between both yield curves as the term structure of illiquidity premia.

2.1 Data

In this section, we describe our data set and illustrate that both KfW bonds and BUNDs are virtually identical in all relevant characteristics except their liquidity. Thus, the KfW-BUND spread can be attributed to liquidity differences. The Kreditanstalt für Wiederaufbau (KfW) is a promotional bank owned by the German government and federal states. All KfW bonds are explicitly guaranteed by the German government and thus bear effectively the same default risk as government bonds. We only include those bonds that are well comparable to BUNDs: plain vanilla fixed coupon bonds with annual coupon payments that are exchange-tradeable and denominated in Euro. Table 1 gives an overview of all bonds in our sample. Average coupons and time to maturities are on the same order of magnitude. BUNDs and KfWs are comparable in their tax treatment and are both accepted by the European Central Bank (ECB) as collateral for repo transactions.[1] Both KfWs and BUNDs are zero weighted in determining capital requirements within the Basel regulations. In contrast, the two segments differ in their liquidity due to the about eleven times higher outstanding total volume and the more than three times higher average issue size of BUNDs compared to KfW bonds. For a verification, that illiquidity premia are driven by liquidity differences between the two segments, we analyze the relationship between iliquidity premia and bid-ask spread differences as well as a possible influence stemming from a convenience yield for BUNDs in Section 2.3.[2]

[Insert Table 1 about here.]

[1] The ECB divides securities in liquidity categories. KfWs are in the second highest category, whereas BUNDs are in the highest. This leads to small additional haircuts for KfWs of up to 2% for the longest maturities. KfWs and BUNDs are also both accepted by the Federal Reserve for discount window loans with the same margin haircuts. The Bank of England accepts KfWs only as 'wider collateral' which can be used for long-term open market operations and the discount window facility, whereas BUNDs are accepted for all monetary policy operations.

[2] See also Schwarz (2010) and Monfort and Renne (2010) who use the KfW-BUND spread of distinctive bonds or time to maturity buckets as a measure of illiquidity.

Our data set consists of weekly closing prices for BUNDs and KfW bonds from the Frankfurt Stock exchange from February 14th, 1996 to September 29th, 2010. Since BUNDs and KfW bonds are traded mainly over the counter and on electronic trading platforms like MTS, we test whether there are discrepancies between our exchange data and data from other platforms like MTS (for which we have one month of data) or Bloomberg Consensus quotes and do not find any fundamental difference.

2.2 Term Structure Estimation

We estimate the term structure of zero coupon yields of BUNDs and KfW bonds using the Nelson and Siegel (1987) approach. Within this approach, the entire term structure information at time t is condensed in four parameters ($\beta_{0,t}$, $\beta_{1,t}$, $\beta_{2,t}$, τ_t). The zero bond yield of bond class $i \in \{BUND, KfW\}$ at time t for time to maturity T is given as

$$y_t^i(T) = \beta_{0,t}^i + \beta_{1,t}^i \left(\frac{1 - e^{-\frac{T}{\tau_t^i}}}{\frac{T}{\tau_t^i}} \right) + \beta_{2,t}^i \left(\frac{1 - e^{-\frac{T}{\tau_t^i}}}{\frac{T}{\tau_t^i}} - e^{-\frac{T}{\tau_t^i}} \right). \tag{1}$$

The term structure of illiquidity premia at time t can then be calculated as

$$illiq_t(T) = y_t^{KfW}(T) - y_t^{BUND}(T). \tag{2}$$

To make the β-factors of both BUNDs and KfWs directly comparable and to restrict the shape of the term structure of illiquidity premia to shapes allowed within the Nelson and Siegel (1987) approach, we further impose $\tau_t^{BUND} = \tau_t^{KfW}$ (see for example Kempf, Korn, and Uhrig-Homburg (2011), Nelson and Siegel (1987) or Diebold and Li (2006) who restrict τ to be constant over time t).

Estimation is carried out for each week separately by minimizing the sum of squared yield differences over all bonds of both segments. Since there are always more BUNDs than KfWs in the sample, we weight the yield difference with the inverse of the number of bonds of the respective bond class to put equal weights on both segments for the estimation of a common τ. As in Schich (1997), who develops the estimation procedure for the benchmark yield curve employed by the German central bank, we exclude bonds with time to maturity less than three months since for them small errors in the price would translate to large yield errors. This procedure delivers weekly estimates of $\beta_{0,t}^{illiq} = \beta_{0,t}^{KfW} - \beta_{0,t}^{BUND}$,

$\beta_{1,t}^{illiq} = \beta_{1,t}^{KfW} - \beta_{1,t}^{BUND}$, $\beta_{2,t}^{illiq} = \beta_{2,t}^{KfW} - \beta_{2,t}^{BUND}$, and $\tau_t^{illiq} = \tau_t^{KfW} = \tau_t^{BUND}$. Resulting root mean squared errors (RMSE) of 7.3 bps for KfW bonds and 5.5 bps for BUNDs are smaller as reported in Schich (1997) for BUNDs. Fitting errors are in the same order of magnitude for all maturities and both segments. We obtain the largest RMSE of 9.2 bps for short-term KfW bonds with below 2 years time to maturity and the smallest RMSE of 3.5 bps for long-term KfW bonds with more than 8 years time to maturity.

2.3 Evolution of Premia and Shapes

To get a first impression on different shapes of the term structure, Figure 1 shows the evolution of the two, five, and eight year illiquidity premia. As can be seen from this figure, the term structure of illiquidity premia adopts different shapes over time. So for example from mid 2000 until mid 2001, an increasing term structure can be observed, whereas the end of 2006 and the beginning of 2007 are characterized by U-shaped term structures. After the collapse of Lehman Brothers, the term structure becomes strongly decreasing. The average illiquidity premium for the whole observation period and maturities from $T = 0$ to $T = 15$ years is 20.3 bps.

[Insert Figure 1 about here.]

The findings regarding the size and shape of the term structure of illiquidity premia are in line with existing literature. So e.g. Longstaff (2004) finds an illiquidity premium of 9 to 16 bps for zero coupon strips derived from six long-term Refcorp bonds from 1991 to 2001 and Koziol and Sauerbier (2007) find on average 20.4 bps for German Jumbo Pfandbriefe from 2000 to 2001. Kempf, Korn, and Uhrig-Homburg (2011) get an average illiquidity premium between 29 bps for one year and 40 bps for 15 years of time to maturity for the relatively heterogeneous segment of German Pfandbriefe from 2000 to 2007. Regarding the shape of the term structure, Amihud and Mendelson (1991) find a declining term structure of illiquidity premia for US Treasury notes compared to more liquid Treasury bills in 1987, whereas Longstaff (2004) presents evidence for a U-shaped, and Koziol and Sauerbier (2007) for a hump-shaped profile.[3] Our analysis provides an

[3] Liu, Longstaff, and Mandell (2006) and de Jong and Driessen (2006) filter premia for illiquidity risk out of swap and corporate bond spreads. Dick-Nielsen, Feldhütter, and Lando (2012) estimates a liquidity component as a fraction of corporate bond spreads. All three papers find an increasing term structure.

explanation for these conflicting outcomes as the shape of the term structure of illiquidity premia varies over time. In contrast to Kempf, Korn, and Uhrig-Homburg (2011) who also observe varying term structures of illiquidity premia for the Pfandbrief segment, our homogeneous data set of bonds from only two issuers allows us to rule out unobserved changing intra-segment liquidity differences as a possible explanation.[4]

To verify that the estimated illiquidity premia are due to liquidity differences, Figure 2 presents the evolution of the two, five, and eight year illiquidity premia over time together with the respective quoted bid-ask spread differences. We obtain proportional bid-ask spreads from all contributors providing quotations in Bloomberg and compute the average bid-ask spread for each bond and each date.[5] To calculate a time to maturity dependent measure, we estimate a linear relationship between the duration of the bond and the bid-ask spread for each segment and each date. With the estimated 'term structure of bid-ask spreads' we are able to aggregate the information from all bonds of a segment in maturity dependent bid-ask spreads similarly as for the estimation of the term structure of illiquidity premia. Bid-ask quotations are available only since 1999 for a majority of the KfW bonds, quotes for bonds of durations less than two years are only available after August 22nd, 2001 on a continuous basis. Figure 2 clearly shows the connection between bid-ask spread differences and illiquidity premia. Their unconditional correlation is 0.89 for two years, 0.85 for five, and 0.84 for eight years of time to maturity. To further rule out the possibility that the KfW-BUND yield spread is due to a convenience yield for BUNDs, we compare the average illiquidity premium with the spread between the general collateral repo rate (EUREPO) and BUND. In the time period for which EUREPO is available (since March 2002), the average difference between the day count adjusted EUREPO twelve months and the one year BUND rate is with 8.5 bps less than one fourth of the average one year illiquidity premium during this time period of 36.8 bps. Since EUREPO is an offer rate, this should be an absolute upper bound of the non-liquidity

[4]Bühler and Vonhoff (2011) also analyze the term structure of illiquidity premia using the difference between the zero coupon yield curves of US Treasuries and coupon strips as well as principal strips on these bonds. However, their term structure of illiquidity premia switches signs as e.g. coupon strips are more liquid than Treasury notes for short-term maturities but less liquid for long-term maturities. The inherent liquidity difference between maturities makes it very hard to interpret the term structure.

[5]Quotations are collected from exchanges, investment banks and other market participants when bid and ask prices are available. We are cautious to exclude calculated prices like consensus quotes and duplicate series.

related fraction of the illiquidity premium.

[Insert Figure 2 about here.]

3 Term Structure Dynamics: A Conditional Approach

In this section, we study the term structure of illiquidity premia in a regime-switching model and analyze economic determinants of illiquidity premia of different maturities conditional on the economic environment.

3.1 Motivation

As can be seen in Figure 1, the behavior of illiquidity premia seems to be heavily affected by the financial crisis. Illiquidity premia of all maturities strongly increase in summer 2007 and reach all time highs after the collapse of Lehman Brothers.[6] In the literature, we find both theoretical and empirical evidence for a regime-switching behavior of bond prices with respect to liquidity. So Acharya, Amihud, and Bharath (2010) analyze the influence of illiquidity on returns of corporate bonds of different rating classes with a Markov regime-switching model. They find that in the stress regime, prices of investment grade bonds rise and prices of speculative grade bonds fall with deteriorating liquidity (flight-to-liquidity). In contrast, there is no significant effect of illiquidity on bond returns in normal times. In another paper, Acharya and Pedersen (2005) find high innovations of illiquidity (measured with the Amihud (2002) measure) for American stocks during periods anecdotally characterized as liquidity crises (e.g. the Russian default and LTCM crisis).

There is also a small but growing body of research predicting a regime-switching behavior of liquidity from a theoretical point of view. Brunnermeier and Pedersen (2009) propose a model where they relate an asset's liquidity to the margin requirements of market makers. If funding becomes scarce e.g. due to macroeconomic shocks, margin requirements

[6]The regime-switching behavior is confirmed by looking at Augmented Dickey Fuller (ADF) tests. They reject the non-stationary hypothesis for the time period before the beginning of the financial crisis in June 2007 for our three time series of illiquidity premia, whereas if we include the whole time period, non-stationarity can only be rejected for two and five year illiquidity premia.

are increasing which affects the market makers' ability to provide liquidity. The deteriorating liquidity in turn leads to higher asset volatility, which again leads to higher margin requirements. These liquidity spirals result in multiple fragile liquidity equilibria in crisis times. In a similar spirit, Gârleanu and Pedersen (2007) propose a model relating tighter risk-management and deteriorating liquidity. They predict a multiplier effect, which explains sudden dry ups of liquidity after initial losses. Another factor could be an increased risk aversion during crisis times, which would also lead to higher illiquidity premia (see e.g. Acharya, Amihud, and Bharath (2010)).[7] All of this would lead to a higher level and a higher volatility of illiquidity premia during crisis times.

Ignoring a possible different behavior of illiquidity premia in different economic regimes can in the best case only lead to results describing 'average' characteristics. Our study is the first to allow explicitly for a regime dependent behavior of illiquidity premia and their determinants.

3.2 Crisis Identification

We use a Markov regime-switching model for crisis identification which has been first proposed by Hamilton (1989).[8] In contrast to our approach, most authors rely on exceptional events to identify crises (see e.g. Friewald, Jankowitsch, and Subrahmanyam (2009) or Chordia, Sarkar, and Subrahmanyam (2005)), but this procedure is somewhat arbitrary especially for the crisis end date (see Barrell, Davis, Karim, and Liadze (2010) for a short discussion of this topic). The main advantage of the Markov approach is the endogenous crisis identification. So essentially the data tells us, when the system is likely to be in the stress regime. To validate the crisis identification, we relate these endogenously derived stress probabilities to macroeconomic and financial sector variables typically used to identify financial crises. However, since crisis identification is a crucial step in our analysis, we check the robustness of our results using exogenously specified crisis periods in Section 4.

To analyze the different behavior of illiquidity premia conditional on crisis and non-crisis

[7] He and Krishnamurthy (2010) propose a model, in which the premium investors demand for holding risky assets increases in times of crisis when intermediaries are constrained in raising capital.

[8] See Hamilton and Baldev (2002) for an overview of applications of this methodology to economic time series.

times, we estimate a two-regime AR model for the two, five, and eight year illiquidity premia. This model is then augmented with additional explanatory variables in Section 3.3.

$$illiq_t(2) = a_{0,s}^{Short} + \sum_{i=1}^{p} \left(b_{i,s}^{Short} illiq_{t-i}(2) \right) + \epsilon_{s,t}^{Short}, \tag{3}$$

$$illiq_t(5) = a_{0,s}^{Medium} + \sum_{i=1}^{p} \left(b_{i,s}^{Medium} illiq_{t-i}(5) \right) + \epsilon_{s,t}^{Medium}, \tag{4}$$

$$illiq_t(8) = a_{0,s}^{Long} + \sum_{i=1}^{p} \left(b_{i,s}^{Long} illiq_{t-i}(8) \right) + \epsilon_{s,t}^{Long}, \tag{5}$$

where the state $s \in \{1, 2\}$ follows a homogeneous Markov chain with constant transition probabilities

$$\begin{aligned} P(s_t = 1 | s_{t-1} = 1) &= p_{1,1}, \\ P(s_t = 2 | s_{t-1} = 2) &= p_{2,2}. \end{aligned} \tag{6}$$

The vector of error terms $(\epsilon_{s,t}^{Short}, \epsilon_{s,t}^{Medium}, \epsilon_{s,t}^{Long})$ is multi-normally distributed with mean zero and variance-covariance matrix Ω_s where

$$\Omega_s = \begin{pmatrix} (\sigma_s^{Short})^2 & \rho_s^{Short,Med.} \cdot \sigma_s^{Short} \cdot \sigma_s^{Med.} & \rho_s^{Short,Long} \cdot \sigma_s^{Short} \cdot \sigma_s^{Long} \\ \rho_s^{Short,Med.} \cdot \sigma_s^{Short} \cdot \sigma_s^{Med.} & (\sigma_s^{Med.})^2 & \rho_s^{Med.,Long} \cdot \sigma_s^{Med.} \cdot \sigma_s^{Long} \\ \rho_s^{Short,Long} \cdot \sigma_s^{Short} \cdot \sigma_s^{Long} & \rho_s^{Med.,Long} \cdot \sigma_s^{Med.} \cdot \sigma_s^{Long} & (\sigma_s^{Long})^2 \end{pmatrix}. \tag{7}$$

We select this flexible variance-covariance matrix to allow for heteroscedasticity between the two regimes. Also, correlations of the error terms of different segments can be regime-switching. The model is estimated along the lines described in Hamilton (1990) and Hamilton (1994) using the expectation-maximization (EM) algorithm to maximize the log likelihood function. Standard errors are derived using White's (1982) approach to calculate an estimator of the variance-covariance matrix of the coefficients V (see also Hamilton (1994)).[9] We can test for linear hypotheses H_0: $R\alpha = r$ using the Wald chi-squared statistics $W = (R\hat{\alpha} - r)'(R\hat{V}R')^{-1}(R\hat{\alpha} - r)$ (see e.g. Hayashi (2000) or Acharya, Amihud, and Bharath (2010)). Here, R and r define the hypotheses for the parameter vector α. Under H_0, W is asymptotically χ^2 distributed with $(\#r)$ degrees of freedom, where $\#r$ is the rank of R.

[9]White's (1982) variance-covariance matrix is still valid even if the densities of the error terms are not exactly normal.

Table 2 gives the estimation results of model (3)-(7) with $p = 3$ lags which is suggested by both Akaike's Information Criterion (AIC) and the Bayesian Information Criterion (BIC). The large and significant values of the lagged parameters for all maturities and both regimes show the high persistence of illiquidity premia over time. Figure 3 shows the average term structure of illiquidity premia in both regimes. A clear separation in the two regimes can be recognized. Whereas in the non-stress regime (regime 1), on average the extra yield to maturity of an illiquid KfW bond compared to the liquid BUND is around 15 bps, this illiquidity premium nearly doubles in the stress regime (regime 2). Additionally, the standard deviation of the innovations is between two to three times larger in the stress regime. The shape of the term structure is slightly U-shaped in both regimes, but the decreasing part is much more pronounced in the stress regime due to large short-term illiquidity premia. This result can be explained with flight-to-liquidity periods that coincidence with financial crises. Within these periods, investors seek for short-term and extremely liquid bonds as a safe haven to temporarily store their funds.[10]

The estimation of the parameters delivers the probability of the system being in the stress regime for each date in the sample. This probability is plotted in Figure 4. The stress regime can be clearly associated with economic events that might be causal for poor liquidity. So the 1998 bailout of Long Term Capital Management (LTCM), the period after the burst of the dot-com bubble as well as the financial crisis starting in summer 2007 are all identified as stress periods.

[Insert Table 2 about here.]

[Insert Figure 3 about here.]

[Insert Figure 4 about here.]

To check the reliability of our crisis identification, we formally establish the connection between the liquidity stress regime and economic and financial sector variables typically

[10] We check the validity of these results by looking at the unbiasedness of our yield curve estimates for the bond with the shortest maturity: In the stress (normal times) regime, the yield of the shortest maturity BUND is on average 7.4 (3.2) bps below the estimated BUND yield curve, whereas the respective KfW bond lies 6.5 (2.2) bps above the KfW yield curve. This suggests an even stronger flight to the highly liquid BUND with the shortest time to maturity and means that we underestimate the illiquidity premium (and the increase of the illiquidity premium in the stress regime) at the very short end of the term structure.

associated with stress periods (see e.g. Acharya, Amihud, and Bharath (2010)). Accordingly, we regress the probability of being in the stress regime on the following variables:
(i) *Recession Dummy Germany:* This dummy variable equals 1 if the observation date lies within at least two consecutive quarters of negative real GDP growth in Germany (seasonally and working-day adjusted data from Deutsche Bundesbank).
(ii) *Negative Return DAX:* Dummy variable that equals 1 if the three month return of the German stock market index DAX is below the one standard deviation bound calculated from realized five minute returns of the DAX from 1996 to September 2010 (data from Deutsche Börse).
(iii) *VDAX New:* Implied volatility of the DAX for the next 30 days calculated from options on futures on the stock market index (data from Deutsche Börse).
(iv) *ZEW German Expectation:* ZEW indicator of economic sentiment that measures the expectations of approximately 400 surveyed financial analysts regarding the future economic development (data from the Centre for European Economic Research (ZEW)).
(v) *TED Spread:* Spread between the three month USD Libor and the three month T-Bill rate. This spread measures the uncertainty in the banking system (see e.g. Brunnermeier (2009)) (data from Bloomberg).
(vi) *Credit Spread:* Spread between the yield of one year AA rated corporate bonds and BBB rated corporate bonds. It is assumed that the credit spread widens in times of economic stress (data from Bloomberg for the US bond market since European data only becomes available after 2001).
(vii) *Capitalization of the Banking System:* Yearly change in the total assets of the banking system (only commercial banks) divided by the total financial assets of households and non-financial cooperations (data is from Deutsche Bundesbank). This measure is motivated by the connection between the ability of banks to provide funding and market liquidity (see Brunnermeier and Pedersen (2009)).
(viii) *Systemic Stress*: Indicator of systemic stress in the financial system developed by Holló, Kremer, and Duca (2011) (data from the European Central Bank).

The probability of being in the stress regime P_t^{Stress} is transformed with a modified logit transformation $z_t = \log(\frac{P_t^{Stress}+0.5/T}{1-P_t^{Stress}+0.5/T})$. The modification ensures that z_t is defined for probabilities P_t^{Stress} of 0 and 1 (see e.g. Cox (1970) or Acharya, Amihud, and Bharath (2010)). The estimation results in Table 3 show a clear connection between economic variables typically associated with economic stress and the probability of being in the stress regime. The OLS estimates of z_t as a function of the economic variables are all significant in the expected direction. So a recession in Germany, the three month DAX return being below the one standard deviation bound, an increased implied volatility, large credit spreads, and TED spreads as well as a high level of systemic stress all indicate a higher probability of being in the low liquidity stress regime. On the other hand, a more positive economic outlook and an increased capitalization of the banking system lead to lower stress probabilities.[11] In the univariate regressions, the large explanatory power of systemic stress for the probability to be in the stress regime could be expected since this indicator is specifically designed to identify financial crises. Similarly, the high R^2 in the univariate regression of the credit spread indicates a very close connection between liquidity and credit risk as analyzed in Ericsson and Renault (2006) and He and Xiong (2011). With the clean separation of illiquidity from credit risk through two bond segments that are guaranteed through the same issuer, our paper sheds new light on this topic as well and confirms the hypothesis that both credit and liquidity risk are closely connected. Together with the anecdotal evidence from Figure 4, this analysis confirms the capability of our approach to identify stress periods.

[Insert Table 3 about here.]

3.3 Factors Influencing the Term Structure of Illiquidity

In this section, we analyze economic determinants of illiquidity premia of different maturities conditional on crisis and non-crisis times. We select measures based mainly on the theoretical but also empirical literature regarding the formation of illiquidity and illiquidity premia. As illiquidity premia are the extra yield investors receive for holding an illiquid bond, it is reasonable to hypothesize that they depend first on the degree of illiquidity of the respective bond and second on the preferences for liquidity of the marginal investor.

We account for the degree of illiquidity first with a measure of general market liquidity (see e.g. Acharya, Amihud, and Bharath (2010)) and second by directly looking at liq-

[11]However, the effect of the increased capitalization of the banking system only becomes significant if we control for implied volatility. A possible reason for this is that increased volatility might lead to higher capital requirements which in turn lead to an increased capitalization. So an increased capitalization is also an indicator for increased volatility which would lead to contrary impacts on the stress probability if we do not control for volatility.

uidity differences between KfW bonds and BUNDs. For the measure of general market liquidity, we follow the inventory holding cost paradigm dating back to Demsetz (1968), Stoll (1978), and Ho and Stoll (1981) and empirically validated e.g. by Benston and Hagerman (1974) and Bollen, Smith, and Whaley (2004) for the stock market and by Goyenko, Subrahmanyam, and Ukhov (2011) for the bond market. This paradigm states that an increased asset volatility leads to larger bid-ask spreads to compensate the market makers for the incurred inventory risk and in turn should lead to increased illiquidity premia (see e.g. Amihud and Mendelson (1986)). To proxy for the inventory risk of market makers, we use option implied interest rate volatility. Since coupon bonds are very similar to swaps in their cash flow structure, we use implied swaption volatilities from Bloomberg. They are derived according to Black's (1976) model from swaption prices with an option tenor of three months and a swap tenor of one, two, three, four, five, seven, and ten years. Since Black volatilities are heavily negatively correlated with the level of interest rates, but we do not want to incorporate interest rate effects, we orthogonalize the seven volatility series with the three-month Euribor.[12] We use the first principal component of the seven orthogonalized series as our measure for inventory risk.

Liquidity differences between the two bond segments are measured using the outstanding volume that is freely available for trading. For this, we construct measures of the tradeable volume of the representative two, five, and eight year KfW bond relative to the tradeable volume of the corresponding German government bond. These measures are calculated in three steps. First, we adjust for the effect that the tradeable volume and liquidity of a bond tends to decrease over its life due to the fact that the outstanding volume is absorbed by buy-and-hold investors (aging effect).[13] Ejsing and Sihvonen (2009) estimate for German government bonds that the trading volume of an issue declines by eight percent each year. Therefore, we multiply the outstanding volume of each bond with $e^{-0.08 \cdot \textit{Age of the issue}_t}$. Second, we calculate the tradeable volume of the representative bond for each of the three maturities from the tradeable volume of all outstanding bonds from each segment. More precisely, we weight the volume of each bond with the influence it has on the zero coupon yield of the respective maturity. To measure this influence, we

[12] The unconditional correlation between the seven implied volatility series and the three-month Euribor lies between -0.574 and -0.669. For the time before the introduction of the Euro, we use Deutsche Mark swaption volatilities and orthogonalize them with the three-month Fibor (Frankfurt interbank offered rate).

[13] See e.g. Warga (1992) or Ericsson and Renault (2006).

calculate the sensitivity of a small yield change of this bond on the zero coupon yield curve.[14] Third, we divide the tradeable volume of the representative KfW bond of each maturity by the tradeable volume of the respective German government bond. As can be seen in Table 4, the average tradeable volume of the representative KfW bond is about 20% of the volume of its Bund counterpart.

The selection of a measure for liquidity preferences is more difficult. On the one hand, it is obvious that future liquidation needs positively influence preferences for liquidity (see Kempf, Korn, and Uhrig-Homburg (2011)). On the other hand, it is likely that market wide risk premia also impact the extra yield investors demand for holding an illiquid bond (see e.g. Gârleanu and Pedersen (2011)). As a proxy for future liquidation needs, we select the benchmark volatility index for the German stock market (VDAX New). As Kempf, Korn, and Uhrig-Homburg (2011) argue, implied volatility is a measure of the expected amount of information flowing into the market. It is likely that trading needs increase with a growing information flow. VDAX New is calculated by Deutsche Börse from options on futures on the German stock market index (DAX). Another natural measure for future liquidation needs would be the implied interest rate volatility.[15] As implied interest rate volatility is already used to proxy for inventory risk, we cannot use it for the preferences for liquidity and it is possible that we overestimate effects stemming from inventory risk. As a further measure of financial uncertainty, which positively influences future liquidation needs, we use the spread between the unsecured three-month LIBOR rate and the U.S. Treasury bill rate (TED spread). Brunnermeier (2009) points out, that in times of higher uncertainty in the banking system, the risk of unsecured loans rises which in turn leads to higher LIBOR rates. Additionally, in times of higher uncertainty the value of first rate collateral rises pushing down T-Bill rate and widening the TED spread further.[16] As a proxy for market wide risk premia or required returns, we select the dividend yield of the German stock market index DAX. In his 2011 Presidential Address to the American Finance Association, Cochrane (2011) points out that variations in the market wide dividend yield reflect changes in risk premia rather than changes

[14] The advantage of this weighting scheme is the independence from arbitrarily selected time to maturity bucket bounds. Additionally, it minimizes the time series variation resulting from bonds changing buckets.

[15] See e.g. Vayanos (2004) who models liquidation needs of fund managers as the probability of the fund's performance falling below a threshold. This probability increases with volatility.

[16] The TED spread as a measure of uncertainty, tightness, or fear in the banking system is also employed e.g. by Nyborg and Östberg (2010) and Kucuk (2009).

in future expected returns. The dividend yield is calculated by Bloomberg under the assumption, that for all 30 constituents of the DAX, realized dividends in the year before the observation date are payed as an infinite annuity. It is available after May 7th, 1997. Unfortunately it is not possible to separate the effects of changed liquidation needs and changed risk premia since both TED spread and VDAX New, besides being measures for uncertainty, are also expected to be sensitive to an increase in risk premia. On the other hand, the dividend yield can be expected to rise in uncertain times (with high future liquidation needs) due to declining stock prices.[17] Thus, we use the first principal component of the three series as our measure for liquidity preferences.

[Insert Table 4 about here.]

To analyze the impact of our economic factors on the term structure of illiquidity premia conditional on the two regimes, we augment the model (3)-(7) in the following way:

$$
\begin{aligned}
illiq_t(2) &= a_{0,s}^{Short} + \sum_{i=1}^{p} \left(b_{i,s}^{Short} illiq_{t-i}(2) \right) + a_{1,s}^{Short} invRisk_t \\
&\quad + a_{2,s}^{Short} trblVol_t(2) + a_{3,s}^{Short} liquPref_t + \epsilon_{s,t}^{Short}, \qquad (8) \\
illiq_t(5) &= a_{0,s}^{Medium} + \sum_{i=1}^{p} \left(b_{i,s}^{Medium} illiq_{t-i}(5) \right) + a_{1,s}^{Medium} invRisk_t \\
&\quad + a_{2,s}^{Medium} trblVol_t(5) + a_{3,s}^{Medium} liquPref_t + \epsilon_{s,t}^{Medium}, \qquad (9) \\
illiq_t(8) &= a_{0,s}^{Long} + \sum_{i=1}^{p} \left(b_{i,s}^{Long} illiq_{t-i}(8) \right) + a_{1,s}^{Long} invRisk_t \\
&\quad + a_{2,s}^{Long} trblVol_t(8) + a_{3,s}^{Long} liquPref_t + \epsilon_{s,t}^{Long}, \qquad (10)
\end{aligned}
$$

where again the state $s_t \in \{1, 2\}$ follows a homogeneous Markov chain as specified in (6) and the vector of error terms $(\epsilon_{s,t}^{Short}, \epsilon_{s,t}^{Medium}, \epsilon_{s,t}^{Long})$ is multi-normally distributed with mean zero and variance-covariance matrix as defined in (7). As in the Markov regime-switching AR model, we set $p = 3$. The model (8), (9), (10), (6), and (7) is again estimated using the EM algorithm to maximize the log likelihood. This procedure yields estimates for all parameters and a new time series of probabilities of being in the stress regime. These new stress probabilities are very similar to the ones derived from the estimation of model (3)-(7) (the average absolute difference between the two probability series is only

[17] For a discussion of the relationship between liquidation needs and risk premia see e.g. Vayanos (2004) in the context of fund managers.

2.6%) and they also expose the same crisis events (LTCM, burst of dot-com bubble, and financial crisis).[18]

Table 5 presents the estimation results of the model (8), (9), (10), (6), and (7). First note that conditional mean illiquidity premia as well as standard deviations and correlations of the error term are very similar to the ones estimated for the regime-switching AR model without exogenous variables (see Table 2). As before, the lagged values of the endogenous variables are highly significant for all maturities and both regimes. Regarding the explanatory variables, the results in Table 5 reveal that inventory risk proxied by implied swaption volatilities heavily influences illiquidity premia of all maturities in the crisis regime, whereas in normal times, there is no significant influence. In Table 6 we analyze the regime-switching behavior of the sensitivities. For inventory risk, the null hypothesis that the influence on the illiquidity premium is identical in both regimes can be clearly rejected for the medium- and long-term premia. The effect of a more important influence of inventory risk in crisis times can be explained by the model of Brunnermeier and Pedersen (2009). When funding is scarce, it is much more costly for market makers to put aside the additional capital required to cover potential losses on their inventory positions. Thus, they cut back liquidity provision more strongly during crisis times, when inventory risk increases.[19] To illustrate the economic significance, we look at a one standard deviation shock on inventory risk. Such a shock leads to an increase e.g. of the medium-term premium of about 2.7 bps, which is more than one tenth of the average premium in the stress regime.

[Insert Table 5 about here.]

[Insert Table 6 about here.]

Liquidity differences, proxied by the fraction of tradeable volume of the representative KfW bond compared to its BUND counterpart, only play a minor role. Although a large tradeable volume of the KfW seems to significantly decrease the five year illiquidity

[18] Regressing the logit transformed stress probabilities on the variables specified in Section 3.2 (see also Table 3) qualitatively yields the same results for all economic variables.

[19] An alternative explanation is that the sensitivity of illiquidity premia regarding a changing liquidity increases in crisis times. This is observed for the corporate bond market by Dick-Nielsen, Feldhütter, and Lando (2012) for the recent subprime crisis. Van Landschoot (2008) finds, that the effect of changes to bid-ask spreads on corporate bond spreads is higher when liquidity is low.

premium, the economic significance is weak (a one standard deviation shock has only an impact of 0.7 bps). The result that illiquidity premia do not depend on tradeable volume is also observed for the Pfandbrief market by Kempf, Korn, and Uhrig-Homburg (2011). A possible explanation the authors provide is that perceived liquidity differences do not change with each issued bond but are rather static. Another explanation could be that the value investors attribute to liquidity increases with less available liquid BUNDs due to the law of supply and demand (see Krishnamurthy and Vissing-Jorgensen (2010)). If this would be the case, there are two opposite effects from an increase of the BUND volume. First, the relative liquidity of KfW bonds decreases leading to an increase in illiquidity premia. Second, the value of liquidity decreases which should decrease illiquidity premia.

Our measure for the preferences for liquidity only has an influence on the short end of the term structure and only within the crisis regime. Looking at the economic significance, we again consider a one standard deviation shock to the preference variable which leads to an impact on the two year illiquidity premium of 3.1 bps in the crisis regime (about 10% of the average premium in this regime). The regime-switching impact of the preferences variable on the short end is also confirmed by Table 6. Again, this result can be explained with flight-to-liquidity periods. The increased demand for short-term and highly liquid securities within these periods leads first to a strongly increased level of especially short-term illiquidity premia (see Figure 3). Second, amplified by the increased wariness to bear risks within stress periods, effects stemming from liquidity demand become more important.

Overall, our results confirm the prediction of the theoretical literature (e.g. Brunnermeier and Pedersen (2009)) that the impact of changes in fundamentals on illiquidity premia is significantly stronger when funding is tight and the system is in stress. Thus, calibrating e.g. risk management models in normal times, when the influence of fundamentals on illiquidity premia is weak, strongly underestimates the contribution of illiquidity to systematic risk.

4 Robustness

In this section, we perform several robustness checks, in which we control for perceived credit risk, selling and buying pressure from abroad, and the level of interest rates. Additionally we look at the influence of the principal component analysis on our results, incorporate possible spill-over effects between different segments of the term structure, and use a different regime identification methodology.

First, we include a measure for the market wide credit spread as an additional explanatory variable to control for perceived credit risk. We use the spread between the Bloomberg index for the yield of BBB rated industrial USD bonds and the corresponding AA index. We cannot utilize credit spreads of EUR bonds since these are available only after August 2001. The credit spread indices are available with different maturities, so we add the index with the corresponding time to maturity to the respective equation for the short-, medium-, and long-term illiquidity premium in the model (8), (9), (10), (6), and (7).

The results are presented in Table 7. The parameter estimates for our explanatory variables are qualitatively and quantitatively similar and the parameters for the credit spread are not significant at the 5% level for any time to maturity or in any regime. Together with the result, that the credit spread has significant explanatory power for the probability to be in the liquidity stress regime (see Table 3), this suggests that systematic credit and liquidity risk (even though heavily correlated) are only connected over the overall state of the economy.

[Insert Table 7 about here.]

Second, we control for selling and buying pressure from abroad. There are two channels through which effects might materialize. First, it is possible that foreign investors pull out of everything that is not as well-known as government bonds in times of stress. This would lead to a negative relationship between fund flows and the KfW-BUND spread. Second, it might be the case that due to the higher awareness of foreigners for BUNDs, foreign investments into the German bond market might be directed mainly into the government segment, resulting in a positive relationship. To overcome the problem of these conflicting effects, we differentiate between foreign investments into bonds from public and non-public issuers, where Deutsche Bundesbank classifies the KfW as a non-public issuer. We use net fund flows published by Deutsche Bundesbank and deflate them with the consumer price index. The results in Table 8 confirm our findings for our main explanatory variables. Especially for medium- and long-term maturities in the stress regime the control variables are significant in the expected direction.

[Insert Table 8 about here.]

Third, we control for the influence of the level of interest rates on illiquidity premia. If the marginal investor bases his investment decision on net present value, he discounts future illiquidity premia and therefore, equilibrium premia will be proportional to $(1 + risklessRate_t)$ (see e.g. Yawitz (1977)). Additionally a lower interest rate today compared to higher interest rates in the past might increase the willingness of investors to take liquidity risk due to search for yield (see. e.g. Rajan (2006)). We control for both effects by including the yield of all outstanding debt securities of German issuers. We look at the relative yield level compared to the average yield of the previous three years. The results in Table 9 confirm our hypotheses regarding the average yield for medium- and long-term illiquidity premia in the non-stress regime. Thus, in normal times investors seem to be more willing to take liquidity risks (and therefore drive illiquidity premia down) when interest rates are low compared to the previous years.[20] Controlling for the influence of interest rates, all significances for inventory risk, liquidity differences and liquidity preferences are unchanged compared to our main analysis.

[Insert Table 9 about here.]

Fourth, we check whether our results are governed by the use of first principal components as explanatory variables. Since we cannot observe inventory risk and liquidity preferences directly, the first principal component is used in our main analysis to reduce the noise in our proxies (see e.g. Ang and Piazzesi (2003), Baker and Wurgler (2006), or Dick-Nielsen, Feldhütter, and Lando (2012)). To check whether this approach drives our findings, we use maturity dependent swaption volatilities for inventory risk[21] and the dividend yield of the DAX as the measure for liquidity preferences. The results in Table 10 confirm our main results, with the only exception, that liquidity preferences are now significant also for medium-term maturities.

[Insert Table 10 about here.]

[20] The effects are robust to using the level of interest rates relative to the average yield in the previous one and five years as well as using the level of interest rates directly.

[21] As before, we use swaptions with an option tenor of three months and a swap tenor of two, five, and eight years and orthogonalize them with the three months Euribor. As a swaption with eight years swap tenor is not directly available, we linearly interpolate its implied volatility from the seven and ten year series.

Next, we control for dynamic linkages between illiquidity premia of different maturities. Goyenko, Subrahmanyam, and Ukhov (2011) find for bid-ask spreads of US treasuries, that shocks on the liquidity of short-term bonds transmit to longer-term maturities, but they also observe transitions in the other direction. On the other hand, Kempf, Korn, and Uhrig-Homburg (2011) cannot detect any spill-over effects between different segments of the term structure for illiquidity premia of German Pfandbriefe. We control for possible spill-over effects by including the lagged values of the other variables:

$$\begin{aligned}
illiq_t(2) = {} & a_{0,s}^{Short} + \sum_{i=1}^{p} \left(b_{i,s}^{Short} illiq_{t-i}(2) \right) \\
& + \sum_{i=1}^{q} \left(c_{i,Medium,s}^{Short} illiq_{t-i}(5) + c_{i,Long,s}^{Short} illiq_{t-i}(8) \right) \\
& + a_{1,s}^{Short} invRisk_t + a_{2,s}^{Short} trblVol_t(2) + a_{3,s}^{Short} liquPref_t + \epsilon_{s,t}^{Short}, \qquad (11) \\
illiq_t(5) = {} & a_{0,s}^{Medium} + \sum_{i=1}^{p} \left(b_{i,s}^{Medium} illiq_{t-i}(5) \right) \\
& + \sum_{i=1}^{q} \left(c_{i,Short,s}^{Medium} illiq_{t-i}(2) + c_{i,Long,s}^{Medium} illiq_{t-i}(8) \right) \\
& + a_{1,s}^{Medium} invRisk_t + a_{2,s}^{Medium} trblVol_t(5) + a_{3,s}^{Medium} liquPref_t + \epsilon_{s,t}^{Medium}, \qquad (12) \\
illiq_t(8) = {} & a_{0,s}^{Long} + \sum_{i=1}^{p} \left(b_{i,s}^{Long} illiq_{t-i}(8) \right) \\
& + \sum_{i=1}^{q} \left(c_{i,Short,s}^{Long} illiq_{t-i}(2) + c_{i,Medium,s}^{Long} illiq_{t-i}(5) \right) \\
& + a_{1,s}^{Long} invRisk_t + a_{2,s}^{Long} trblVol_t(8) + a_{3,s}^{Long} liquPref_t + \epsilon_{s,t}^{Long}, \qquad (13)
\end{aligned}$$

with the specifications for the transition matrix in equation (6) and the variance-covariance matrix in (7) unchanged. We select $q = 1$ lag as suggested by the Bayesian Information Criterion.[22] The results in Table 11 confirm our results for liquidity preferences and liquidity differences between the two segments. The results for inventory risk are slightly weaker. Only the medium- and long-term premia are significant in the crisis regime. Since inventory risk influences all maturities in our main analysis, part of its impact seems to take effect over interdependencies between the segments. Regarding spill-over effects, lagged values of long-term maturities are significant for medium-term premia in both regimes.

[22]Akaike's Information Criterion suggests $q = 3$. $q = 1$ is selected to remain parsimonious.

[Insert Table 11 about here.]

Finally, we control for the crisis identification mechanism by estimating two unconditional models before and after the onset of the financial crisis in June 2007. As this crisis identification ignores the LTCM and dot-com stress periods, we expect the significances for the crisis period to be weaker due to the shorter observation period and the 1997 to 2007 results to be some kind of average of crisis and non-crisis. The results in Table 12 show that this turns out to be the case for inventory risk in the crisis period, which is now only significant for medium- and long-term maturities. The positive significance of inventory risk for the period before the financial crisis confirms the results from Kempf, Korn, and Uhrig-Homburg (2011) who find for the period from 2001 to 2007 that short-term illiquidity premia for Pfandbriefe are significantly influenced by bond market volatility. The positive significance of tradeable volume for the five year illiquidity premium in the crisis is puzzling at first sight. An explanation could be that demand pressure for BUNDs in the financial crisis as studied in Krishnamurthy and Vissing-Jorgensen (2010) dominates the liquidity effect. Consistent with our previous results, the parameter for liquidity preferences of short-term maturities is only significant within the crisis. Additionally, Figure 5 shows average term structures of illiquidity premia before and after the onset of the financial crisis. This figure confirms the results in Figure 3 that illiquidity premia of all but especially short-term maturities increase sharply in stress periods.

[Insert Table 12 about here.]

[Insert Figure 5 about here.]

5 Conclusion

In this paper, we extract the term structure of illiquidity premia from the spread between two bond classes differing only in their liquidity. The availability of a data set of homogeneous bonds spanning a large time to maturity segment over a long period of time allows us to quantify the term structure of illiquidity premia without strong assumptions (e.g. regarding the separation of credit and liquidity risk). We analyze this term structure in a setting allowing for a different behavior during crisis and non-crisis times.

We find, that the term structure of illiquidity premia varies over time and is strongly dependent on the general financial and economic situation. Option implied interest rate volatilities as a measure for inventory risk influence all maturities, whereas our measure for liquidity preferences only impacts short-term maturities. The regression coefficients are regime-switching with a significant impact only in the stress regime.

Turning to the implications of our results we note, that systematic liquidity risk is prone to be underestimated. We have shown that through its regime-switching behavior, the illiquidity discount increases sharply when the general state of the economy is bad. Additionally, the sensitivity of illiquidity premia to fundamentals increases in crisis times. Ignoring one of these two channels systematically underestimates liquidity risk. From the issuer's perspective, our results show, that in times of crisis it is even more important to optimize the liquidity of an issue.[23] For an overview of possible measures to improve the liquidity see e.g. Amihud and Mendelson (2006).

[23]We have checked that issuing yields in the primary market for KfW bonds are relatively close to our yield curves estimated from secondary market prices.

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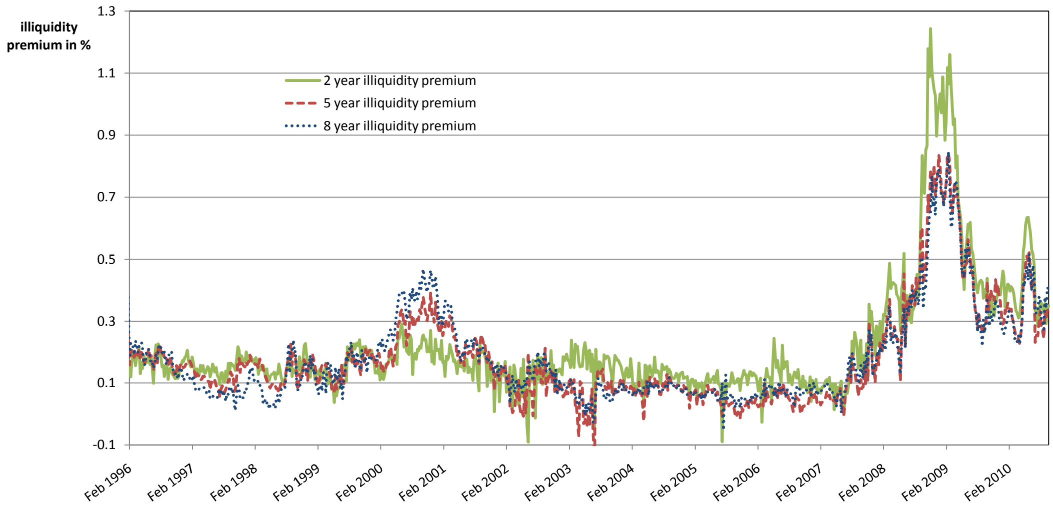

Figure 1: Illiquidity premia of different maturities
This figure shows the development of illiquidity premia over time. The solid line depicts a time to maturity of two years, the dashed line provides five years, and the dotted line eight years time to maturity. The observation period is from February 14th, 1996 to September 29th, 2010 (764 weekly observations).

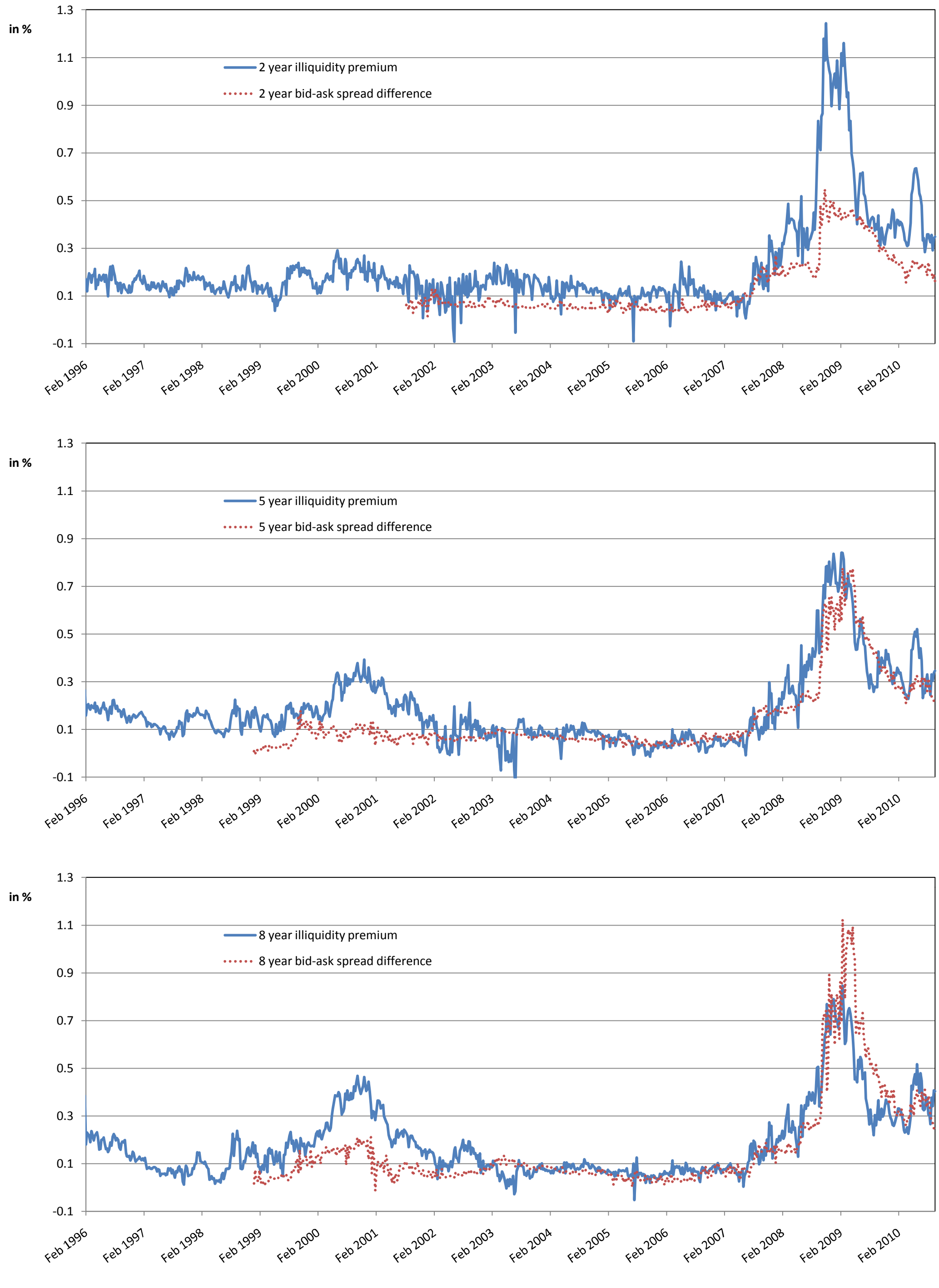

Figure 2: Illiquidity premia and quoted bid-ask spreads
This figure shows the development of illiquidity premia (solid lines) and quoted proportional bid-ask spread differences between KfW bonds and BUNDs (dotted lines) over time. The upper graph depicts a time to maturity of two years, the middle graph provides five years, and the lower graph eight years time to maturity. The observation period is from February 14th, 1996 to September 29th, 2010 (764 weekly observations).

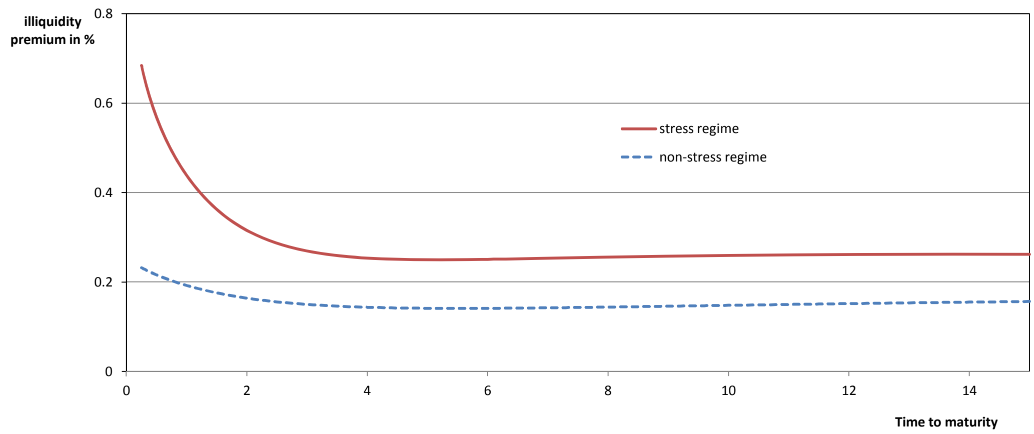

Figure 3: Shapes of the term structure of illiquidity premia in different regimes
This figure shows the shapes of the term structure of illiquidity premia in the stress regime (solid line) and in the non-stress regime (dashed line). The average term structure of illiquidity premia in one regime is calculated by weighting the term structure of each day with the probability to be in that regime on that date (see Figure 4).

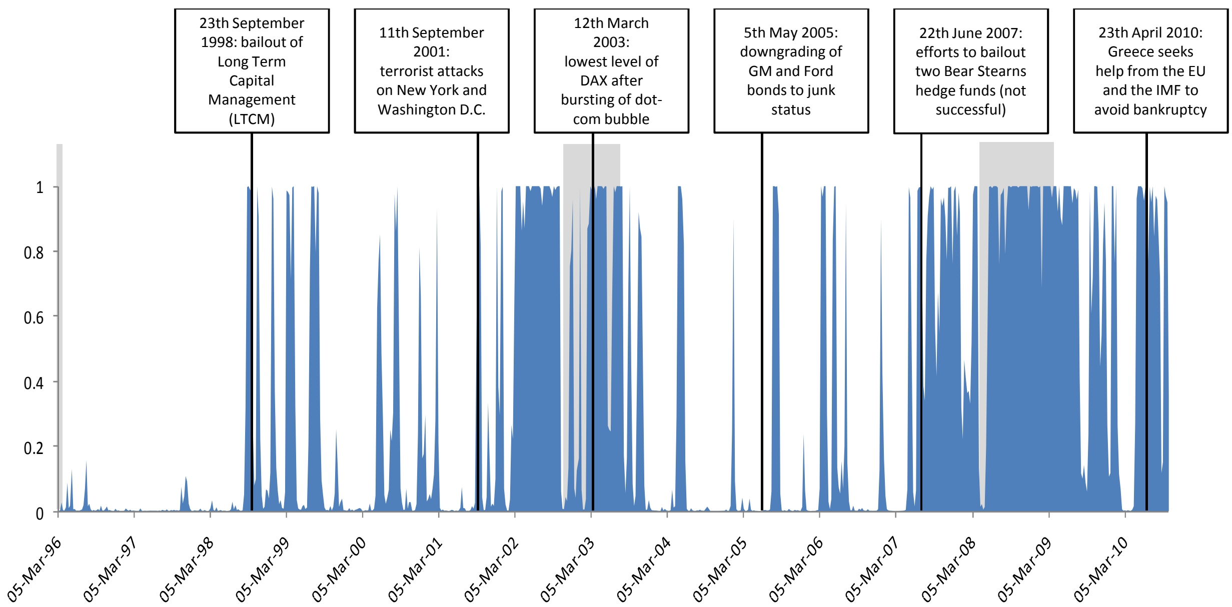

Figure 4: Endogenously derived probability to be in the stres regime
This figure shows the probability of being in the stress regime (regime 2) estimated from the Markov regime-switching model (3)-(7) with three lags. Additionally, events anecdotally linked to financial stress or low liquidity are marked. Recessions, defined as at least two consecutive quarters of negative real GDP growth in Germany (Q1 1996, Q4 2002 – Q2 2003, and Q2 2008 – Q1 2009) are shaded.

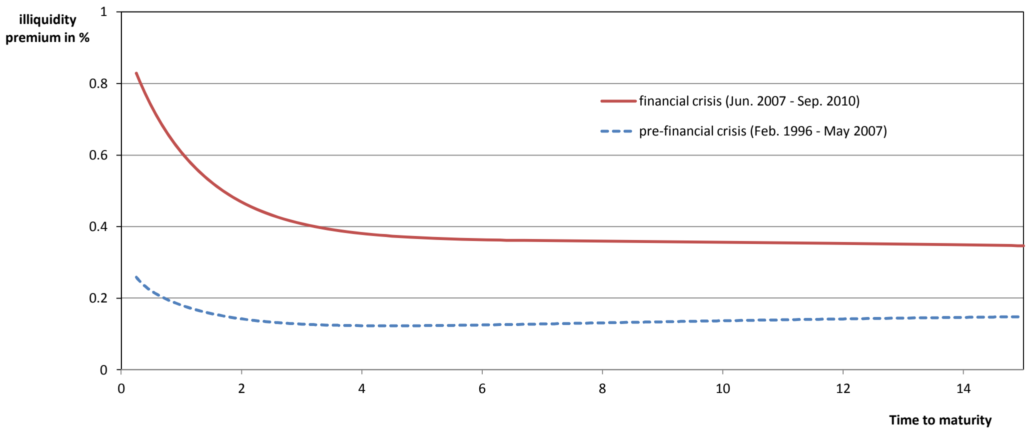

Figure 5: Robustness Check: Shapes of the term structure of illiquidity premia before and after the onset of the financial crisis
This figure shows average shapes of the term structure of illiquidity premia before the onset of the financial crisis until May 2007 (dashed line) and during the financial crisis from June 2007 until the end of the observation period (solid line). The total observation period is February 14th, 1996 to September 29th, 2010.

	Kreditanstalt für Wiederaufbau (KfW)	German government bonds (BUND)
Number of bonds	68	227
Average time to maturity at issue date (in years)	6.09	7.52
Average coupon (in %)	4.13	4.90
Average issuing volume (incl. all reopenings) (in bn EUR)	2.99	9.83
Total volume (in bn EUR)	203	2 231

Table 1: Summary statistics for KfW bonds and BUNDs
This table shows summary statistics for the bonds included in the sample. The observation period is February 14th, 1996 to September 29th, 2010.

	Regime 1 (normal times)			Regime 2 (stress regime)		
	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.
mean illiq. premium(T)	16.4 bps	14.1 bps	14.4 bps	31.6 bps	25.0 bps	25.6
σ	0.0247	0.0209	0.0201	0.0757	0.0617	0.0511
	(0.0008)	(0.0007)	(0.0006)	(0.0027)	(0.0027)	(0.0024)
correlation parameters	$\rho_1^{Short,Long}$	$\rho_1^{Short,Med.}$	$\rho_1^{Med.,Long}$	$\rho_2^{Short,Long}$	$\rho_2^{Short,Med.}$	$\rho_2^{Med.,Long}$
	0.274	0.655	0.6917	0.297	0.4907	0.7603
	(0.0464)	(0.0275)	(0.0256)	(0.0663)	(0.0462)	(0.0323)
transition probabilities		$p_{1,1}$			$p_{2,2}$	
		0.914			0.8393	
		(0.0156)			(0.0279)	
Constant	0.0072*	0.0063*	0.0056*	0.0156*	0.0162**	0.0139**
	(0.0035)	(0.0024)	(0.0025)	(0.0068)	(0.0058)	(0.0048)
$illiq_{t-1}(T)$	0.5955**	0.6675**	0.7017**	0.661**	0.5876**	0.579**
	(0.0584)	(0.0436)	(0.0523)	(0.0724)	(0.0629)	(0.0617)
$illiq_{t-2}(T)$	0.1788**	0.1488**	0.2104**	0.2349**	0.1653*	0.1903**
	(0.0556)	(0.0462)	(0.0567)	(0.0736)	(0.0691)	(0.0727)
$illiq_{t-3}(T)$	0.1754**	0.1274**	0.0362	0.0639	0.2**	0.1965**
	(0.0454)	(0.0355)	(0.048)	(0.076)	(0.0666)	(0.068)
	Log Likelihood		N	AIC	BIC	
	4976.02		761	-9876.05	-9658.18	

Table 2: Estimation results for Markov regime-switching AR model
This table shows the results of the maximum likelihood estimation of model (3)-(7) with three lags. White's (1982) standard errors are given in parentheses. *, ** indicate significance at the 5% or 1% level using the Wald chi-squared statistics $W = (R\hat{\alpha} - r)'(R\hat{V}R')^{-1}(R\hat{\alpha} - r)'$. The observation period is February 14th, 1996 to September 29th, 2010.

	(1)	(2)	(3)	(4)	(5)
Constant	-2.0021^{**}	-1.91693^{**}	-6.5536^{**}	-0.1791	-2.8417^{**}
	(0.1767)	(0.1779)	(0.4212)	(0.2194)	(0.2545)
Reces. Dummy Germ.	5.0563^{**}				
	(0.5001)				
Neg. Return DAX		4.7216^{**}			
		(0.5232)			
VDAX New			20.2211^{**}		
			(1.5225)		
ZEW German Expec.				-0.038^{**}	
				(0.0045)	
TED Spread					2.837^{**}
					(0.3661)
...					
Adjusted R^2	0.1176	0.0957	0.1875	0.085	0.0721

	(6)	(7)	(8)	(9)
Constant	-5.1469^{**}	-6.8565^{**}	-4.263^{**}	-6.2873^{**}
	(0.2494)	(0.4262)	(0.2104)	(0.5143)
Rec. Dummy Ger.				0.6463
				(0.5958)
Neg. Return DAX				0.2118
				(0.5977)
VDAX New		21.9046^{**}		8.7844^{**}
		(1.5805)		(2.1382)
ZEW German Expec.				-0.0069
				(0.0052)
TED Spread				0.6231
				(0.4345)
Credit Spread	5.6568^{**}			2.0641^{**}
	(0.3032)			(0.6251)
Capital. Bank. Syst.		-0.0833^{**}		-0.0818^{**}
		(0.023)		(0.0271)
Systemic Stress			16.1942^{**}	6.7267^{**}
			(0.8539)	(1.9347)
Adjusted R^2	0.3135	0.2003	0.3206	0.373

Table 3: Model for the probability of being in the stress regime
This table shows OLS estimates for regressions where the dependent variable is a modified logit transformation $z_t = \log(\frac{P_t^{Stress}+0.5/T}{1-P_t^{Stress}+0.5/T})$ of the probability to be in the stress regime P_t^{Stress} estimated from the Markov regime-switching model (3)-(7) with three lags (see Figure 4). This variable is regressed on different economic and fincial sector variables as described in the text. Standard errors are given in parentheses. *, ** indicate significance at the 5% or 1% level. The observation period is March 6th, 1996 to September 29th, 2010 (761 weekly observations).

	Mean	Standard Deviation	Minimum	Median	Maximum
$invRisk$	0	2.5519	-3.9135	-0.4477	14.9636
$trblVol(2)$	0.2117	0.0787	0.0838	0.2108	0.3782
$trblVol(5)$	0.2014	0.0544	0.0923	0.2051	0.3031
$trblVol(8)$	0.2321	0.0466	0.1336	0.2372	0.3328
$liquPref$	0	1.2989	-1.6635	-0.2836	7.8201

Table 4: Summary statistics of explanatory variables
This table shows summary statistics for the variables included in the analysis. $invRisk$ refers to the first principal component of orthogonalized swaption volatilities of different maturities, $trblVol(T)$ measures the tradeable volume of the representative KfW bond compared to its BUND counterpart with T years to maturity, and $liquPref$ is the first principal component of the VDAX New, the TED spread, and the dividend yield of the DAX. The observation period is May 7th, 1997 to September 29th, 2010 (700 weekly observations).

	Regime 1 (normal times)			Regime 2 (stress regime)		
	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.
mean illiq. premium(T)	16.6 bps	14.0 bps	14.4 bps	31.5 bps	24.8 bps	25.4 bps
σ	0.0239	0.0216	0.0204	0.0709	0.0595	0.0501
	(0.0019)	(0.0017)	(0.0015)	(0.0046)	(0.0037)	(0.0029)
correlation parameters	$\rho_1^{Short,Long}$	$\rho_1^{Short,Med.}$	$\rho_1^{Med.,Long}$	$\rho_2^{Short,Long}$	$\rho_2^{Short,Med.}$	$\rho_2^{Med.,Long}$
	0.2892	0.6544	0.7052	0.3001	0.4904	0.7623
	(0.0608)	(0.0429)	(0.0348)	(0.0707)	(0.0647)	(0.03)
transition probabilities		$p_{1,1}$			$p_{2,2}$	
		0.9024			0.8384	
		(0.0229)			(0.0354)	
$Constant$	0.0159**	0.0087	0.0045	0.0566**	0.0614**	0.0475*
	(0.0053)	(0.0056)	(0.0066)	(0.0143)	(0.0141)	(0.0208)
$illiq_{t-1}(T)$	0.5697**	0.6741**	0.7023**	0.539**	0.5136**	0.5248**
	(0.0558)	(0.0482)	(0.0537)	(0.0713)	(0.0667)	(0.0633)
$illiq_{t-2}(T)$	0.1829**	0.152**	0.2349**	0.1834**	0.1213	0.1477*
	(0.0522)	(0.0486)	(0.0576)	(0.0711)	(0.0685)	(0.0715)
$illiq_{t-3}(T)$	0.1772**	0.116**	0.0158	0.0499	0.1784**	0.1945**
	(0.0409)	(0.0357)	(0.0517)	(0.067)	(0.0641)	(0.0695)
$invRisk_t$	0.0027	0.0004	−0.0006	0.0078*	0.0105**	0.0094**
	(0.0014)	(0.0016)	(0.0015)	(0.0032)	(0.0029)	(0.0029)
$trblVol_t(T)$	−0.0035	0.0042	0.0095	−0.0523	−0.1373*	−0.0898
	(0.0191)	(0.0203)	(0.0274)	(0.052)	(0.0564)	(0.0823)
$liquPref_t$	0.003	0.0043	0.0031	0.0237**	0.003	−0.0015
	(0.0028)	(0.0026)	(0.0021)	(0.0072)	(0.0057)	(0.0055)
	Log Likelihood		N	AIC	BIC	
	4539.48		700	-8966.96	-8650.57	

Table 5: Estimation results for Markov regime-switching AR with exogenous variables

This table shows the results of the maximum likelihood estimation of model (8), (9), (10), (6), and (7) with $p = 3$. White's (1982) standard errors are given in parentheses. *, ** indicate significance at the 5% or 1% level using the Wald chi-squared statistics $W = (R\hat{\alpha} - r)'(R\hat{V}R')^{-1}(R\hat{\alpha} - r)'$. The observation period is May 7th, 1997 to September 29th, 2010.

	Differences between regimes		
	Short	Medium	Long
	T=2 yr.	T=5 yr.	T=8 yr.
$Constant^{s=2}(T) - Constant^{s=1}(T)$	0.0407**	0.0527**	0.043
	[6.9765]	[10.9551]	[3.6066]
$illiq_{t-1}^{s=2}(T) - illiq_{t-1}^{s=1}(T)$	−0.0307	−0.1605	−0.1775
	[0.0983]	[3.305]	[3.7979]
$illiq_{t-2}^{s=2}(T) - illiq_{t-2}^{s=1}(T)$	0.0005	−0.0307	−0.0872
	[0.0000]	[0.1172]	[0.7633]
$illiq_{t-3}^{s=2}(T) - illiq_{t-2}^{s=1}(T)$	−0.1282	0.0624	0.1787
	[2.4831]	[0.6591]	[3.5548]
$invRisk_t^{s=2} - invRisk_t^{s=1}$	0.0051	0.0101**	0.01**
	[1.9553]	[8.3308]	[8.4079]
$trblVol_t^{s=2}(T) - trblVol_t^{s=1}(T)$	−0.0488	−0.1415*	−0.0993
	[0.7402]	[5.3063]	[1.1969]
$liquPref_t^{s=2} - liquPref_t^{s=1}$	0.0207**	−0.0013	−0.0046
	[6.9825]	[0.0395]	[0.5778]

Table 6: Regime-switching behavior of economic determinants
This table shows the differences of the parameter estimates between the two regimes. The null hypothesis H_0 is, that parameter estimates are identical in both regimes. The Wald chi-squared statistics $W = (R\hat{\alpha} - r)'(R\hat{V}R')^{-1}(R\hat{\alpha} - r)'$ are given in square brackets. *, ** indicate rejection of H_0 at the 5% or 1% level. The observation period is May 7th, 1997 to September 29th, 2010.

	Regime 1 (normal times)			Regime 2 (stress regime)		
	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.
mean illiq. premium(T)	16.6 bps	14.0 bps	14.4 bps	31.5 bps	24.8 bps	25.4 bps
σ	0.0241	0.0216	0.0204	0.0711	0.0596	0.0503
	(0.0018)	(0.0017)	(0.0015)	(0.0045)	(0.0036)	(0.0029)
correlation	$\rho_1^{Short,Long}$	$\rho_1^{Short,Med.}$	$\rho_1^{Med.,Long}$	$\rho_2^{Short,Long}$	$\rho_2^{Short,Med.}$	$\rho_2^{Med.,Long}$
parameters	0.2951	0.6611	0.7052	0.3086	0.5027	0.7636
	(0.0591)	(0.0409)	(0.0335)	(0.0703)	(0.0638)	(0.0301)
transition probabilities		$p_{1,1}$			$p_{2,2}$	
		0.9035			0.8386	
		(0.0215)			(0.0322)	
$Constant$	0.0167*	0.012	0.0109	0.0525**	0.0635**	0.0463
	(0.0081)	(0.0079)	(0.0094)	(0.0143)	(0.0157)	(0.0249)
$illiq_{t-1}(T)$	0.5664**	0.672**	0.7048**	0.5402**	0.5018**	0.5197**
	(0.0506)	(0.0483)	(0.0542)	(0.0711)	(0.0672)	(0.0641)
$illiq_{t-2}(T)$	0.1847**	0.1576**	0.2405**	0.192**	0.1121	0.1407
	(0.049)	(0.0492)	(0.058)	(0.0716)	(0.0689)	(0.0726)
$illiq_{t-3}(T)$	0.1744**	0.1227**	0.0126	0.0717	0.1676**	0.1963**
	(0.0387)	(0.0339)	(0.0505)	(0.0675)	(0.0631)	(0.0696)
$invRisk_t$	0.0028	0.0008	0.0001	0.0075*	0.0108**	0.0095**
	(0.0018)	(0.0019)	(0.0019)	(0.0033)	(0.0029)	(0.0029)
$trblVol_t(T)$	−0.0086	0.0214	0.0123	0.0058	−0.1891**	−0.1021
	(0.0189)	(0.0219)	(0.0279)	(0.0567)	(0.0638)	(0.0855)
$liquPref_t$	0.0025	0.0057*	0.0042	0.0247**	0.0014	−0.0022
	(0.0028)	(0.0027)	(0.0023)	(0.0073)	(0.0059)	(0.0055)
$Credit_t(T)$	0.0014	−0.0111	−0.0103	−0.0195	0.0177	0.0077
	(0.0098)	(0.0093)	(0.0095)	(0.0141)	(0.0159)	(0.0135)
	Log Likelihood		N	AIC	BIC	
	4543.63		700	-8963.26	-8612.97	

Table 7: Robustness Check: Credit spread
This table shows the results of the maximum likelihood estimation of model (8), (9), (10), (6), and (7) with $p = 3$ including an additional variable controlling for perceived credit risk. $Credit_t(T)$ refers to the spread between the Bloomberg index for the yield of BBB rated industrial USD bonds and the corresponding AA index with T years to maturity. White's (1982) standard errors are given in parentheses. *, ** indicate significance at the 5% or 1% level using the Wald chi-squared statistics $W = (R\hat{\alpha} - r)'(R\hat{V}R')^{-1}(R\hat{\alpha} - r)'$. The observation period is May 7th, 1997 to September 29th, 2010.

	Regime 1 (normal times)			Regime 2 (stress regime)		
	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.
mean illiq. premium(T)	16.5 bps	14.0 bps	14.5 bps	31.7 bps	24.8 bps	25.3 bps
σ	0.0243	0.0218	0.0204	0.0702	0.0586	0.0496
	(0.0018)	(0.0016)	(0.0015)	(0.0046)	(0.0035)	(0.0028)
correlation parameters	$\rho_1^{Short,Long}$	$\rho_1^{Short,Med.}$	$\rho_1^{Med.,Long}$	$\rho_2^{Short,Long}$	$\rho_2^{Short,Med.}$	$\rho_2^{Med.,Long}$
	0.2958	0.6619	0.7045	0.2924	0.483	0.7605
	(0.063)	(0.0403)	(0.037)	(0.0722)	(0.0655)	(0.0302)
transition probabilities		$p_{1,1}$			$p_{2,2}$	
		0.9074			0.8442	
		(0.0204)			(0.0376)	
$Constant$	0.0149*	0.0079	0.0035	0.0746**	0.0802**	0.0595**
	(0.0059)	(0.0058)	(0.0064)	(0.0169)	(0.017)	(0.0224)
$illiq_{t-1}(T)$	0.5584**	0.6659**	0.6986**	0.5206**	0.4897**	0.5035**
	(0.0624)	(0.0531)	(0.0527)	(0.0715)	(0.0673)	(0.0654)
$illiq_{t-2}(T)$	0.1918**	0.1582**	0.2356**	0.1653*	0.1044	0.1365
	(0.0506)	(0.0486)	(0.0657)	(0.0709)	(0.0685)	(0.0758)
$illiq_{t-3}(T)$	0.1849**	0.1242**	0.0257	0.0315	0.1506*	0.1728*
	(0.0416)	(0.038)	(0.0587)	(0.0651)	(0.0629)	(0.0696)
$invRisk_t$	0.0029	0.0008	−0.0002	0.0071*	0.0093**	0.0084**
	(0.0017)	(0.0025)	(0.0024)	(0.0033)	(0.0031)	(0.0032)
$trblVol_t(T)$	−0.0032	0.0057	0.0102	−0.0522	−0.1589**	−0.0974
	(0.0191)	(0.0203)	(0.0281)	(0.0536)	(0.0578)	(0.0837)
$liquPref_t$	0.0032	0.0043	0.003	0.0273**	0.0068	0.0014
	(0.0028)	(0.0026)	(0.0022)	(0.0076)	(0.0064)	(0.0065)
$ForeignPub$	0.169	0.3006	0.3196	0.5184	1.1894*	1.1294*
	(0.1931)	(0.1829)	(0.1886)	(0.6899)	(0.5688)	(0.4949)
$ForNonPub$	0.0108	−0.0613	−0.0323	−1.8322**	−1.8724**	−1.4013*
	(0.2382)	(0.2307)	(0.229)	(0.6782)	(0.6697)	(0.6434)
	Log Likelihood		N	AIC	BIC	
	4550.2		700	-8964.39	-8580.22	

Table 8: Robustness Check: Selling and buying pressure form abroad
This table shows the results of the maximum likelihood estimation of model (8), (9), (10), (6), and (7) with $p = 3$ including two additional explanatory variables controlling for selling and buying pressure from abroad. $ForeignPub$ ($ForNonPub$) is the deflated net investments of foreigners in bonds of public (non-public) issuers from Germany in trillions of Euros (in prices of 2005). Monthly values of this variable are from Deutsche Bundesbank and used for all weeks of the respective month. White's (1982) standard errors are given in parentheses. *, ** indicate significance at the 5% or 1% level using the Wald chi-squared statistics $W = (R\hat{\alpha} - r)'(R\hat{V}R')^{-1}(R\hat{\alpha} - r)'$. The observation period is May 7th, 1997 to September 29th, 2010.

	Regime 1 (normal times)			Regime 2 (stress regime)		
	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.
mean illiq. premium(T)	16.7 bps	14.1 bps	14.6 bps	31.7 bps	24.9 bps	25.3 bps
σ	0.0242	0.0218	0.0203	0.0713	0.0598	0.0502
	(0.0015)	(0.0013)	(0.0011)	(0.0042)	(0.0035)	(0.0029)
correlation	$\rho_1^{Short,Long}$	$\rho_1^{Short,Med.}$	$\rho_1^{Med.,Long}$	$\rho_2^{Short,Long}$	$\rho_2^{Short,Med.}$	$\rho_2^{Med.,Long}$
parameters	0.3089	0.6677	0.7103	0.2996	0.4888	0.7599
	(0.0565)	(0.0385)	(0.0304)	(0.0707)	(0.0645)	(0.0305)
transition		$p_{1,1}$			$p_{2,2}$	
probabilities		0.9054			0.8392	
		(0.0183)			(0.0322)	
$Constant$	0.0255*	−0.01	−0.0244*	0.0821*	0.0804**	0.0544
	(0.0119)	(0.0095)	(0.0096)	(0.0338)	(0.0311)	(0.0294)
$illiq_{t-1}(T)$	0.562**	0.6594**	0.6801**	0.5362**	0.5134**	0.5235**
	(0.0504)	(0.0475)	(0.0525)	(0.0715)	(0.0671)	(0.0622)
$illiq_{t-2}(T)$	0.1811**	0.1491**	0.2294**	0.1807*	0.1205	0.1488*
	(0.0503)	(0.0469)	(0.0505)	(0.0704)	(0.0686)	(0.0717)
$illiq_{t-3}(T)$	0.1697**	0.1182**	0.0266	0.0476	0.1771**	0.1938**
	(0.038)	(0.0321)	(0.0457)	(0.0672)	(0.0637)	(0.0701)
$invRisk_t$	0.0036*	0.0007	−0.0007	0.0078*	0.0103**	0.0093**
	(0.0015)	(0.0015)	(0.0014)	(0.0033)	(0.0028)	(0.0028)
$trblVol_t(T)$	−0.0184	−0.0109	0.0016	−0.0622	−0.1359*	−0.073
	(0.0206)	(0.0212)	(0.029)	(0.0553)	(0.0577)	(0.0804)
$liquPref_t$	0.0024	0.0049	0.0043	0.0247**	0.0035	−0.0013
	(0.003)	(0.0028)	(0.0023)	(0.0073)	(0.0059)	(0.0056)
$AverYld$	−0.035	0.0257*	0.0354**	−0.0231	−0.0204	−0.0119
	(0.0101)	(0.0102)	(0.0104)	(0.0279)	(0.028)	(0.0227)
	Log Likelihood		N	AIC	BIC	
	4551.81		700	-8979.63	-8629.35	

Table 9: Robustness Check: Level of interest rates

This table shows the results of the maximum likelihood estimation of model (8), (9), (10), (6), and (7) with $p = 3$ including an additional explanatory variable controlling for interest rate effects. $AverYld$ is the yield of all outstanding bonds divided by the average yield of the previous three years (data from Deutsche Bundesbank). White's (1982) standard errors are given in parentheses. *, ** indicate significance at the 5% or 1% level using the Wald chi-squared statistics $W = (R\hat{\alpha} - r)'(R\hat{V}R')^{-1}(R\hat{\alpha} - r)'$. The observation period is May 7th, 1997 to September 29th, 2010.

	Regime 1 (normal times)			Regime 2 (stress regime)		
	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.
mean illiq. premium(T)	16.6 bps	14.1 bps	14.7 bps	31.7 bps	24.7 bps	25.1 bps
σ	0.0246	0.0222	0.0207	0.0724	0.0585	0.0502
	(0.0016)	(0.0015)	(0.0014)	(0.0053)	(0.0036)	(0.003)
correlation parameters	$\rho_1^{Short,Long}$	$\rho_1^{Short,Med.}$	$\rho_1^{Med.,Long}$	$\rho_2^{Short,Long}$	$\rho_2^{Short,Med.}$	$\rho_2^{Med.,Long}$
	0.304	0.669	0.7139	0.252	0.4439	0.7626
	(0.0579)	(0.0359)	(0.0319)	(0.0679)	(0.0666)	(0.0299)
transition probabilities		$p_{1,1}$			$p_{2,2}$	
		0.9056			0.8392	
		(0.0207)			(0.0347)	
$Constant$	0.0082	0.0063	0.0079	−0.0139	0.0017	0.0049
	(0.0089)	(0.0087)	(0.009)	(0.0217)	(0.0221)	(0.0293)
$illiq_{t-1}(T)$	0.5789**	0.6821**	0.7064**	0.5635**	0.4955**	0.5147**
	(0.0568)	(0.0503)	(0.0564)	(0.0726)	(0.0705)	(0.0645)
$illiq_{t-2}(T)$	0.1782**	0.164**	0.2457**	0.1762*	0.1032	0.131
	(0.0542)	(0.0469)	(0.0614)	(0.0753)	(0.067)	(0.0725)
$illiq_{t-3}(T)$	0.1717**	0.1164**	0.0117	0.0102	0.1212	0.1623*
	(0.0413)	(0.0338)	(0.0516)	(0.0733)	(0.0657)	(0.0739)
$invRisk_t(T)$	0.0005	0.0000	0.0000	0.003**	0.0027**	0.003**
	(0.0003)	(0.0007)	(0.0008)	(0.0008)	(0.0009)	(0.001)
$trblVol_t(T)$	−0.0101	0.0057	−0.0028	−0.0799	−0.1708**	−0.0292
	(0.0166)	(0.0213)	(0.0304)	(0.0561)	(0.0564)	(0.0915)
$DivYldDax_t$	0.0031	−0.0014	−0.0014	0.0336**	0.0332**	0.0153
	(0.0041)	(0.004)	(0.0041)	(0.0105)	(0.0114)	(0.0102)
	Log Likelihood		N	AIC	BIC	
	4529.64		700	-8947.28	-8630.9	

Table 10: Robustness Check: Influence of Principal Component Analysis
This table shows the results of the maximum likelihood estimation of model (8), (9), (10), (6), and (7) with $p = 3$ with $invRisk_t(T)$ instead of the first principal $invRisk_t$ and $DivYieldDax$ instead of $liquPref$. $invRisk_t(T)$ is the implied volatility of swaptions with an option tenor of three months and a swap tenor of T years orthogonalized with the three months Euribor (data from Bloomberg). $DivYldDax$ is the value weighted Dividend Yield of the members of the German stock market index DAX (data from Bloomberg). White's (1982) standard errors are given in parentheses. *, ** indicate significance at the 5% or 1% level using the Wald chi-squared statistics $W = (R\hat{\alpha} - r)'(R\hat{V}R')^{-1}(R\hat{\alpha} - r)'$. The observation period is May 7th, 1997 to September 29th, 2010.

	Regime 1 (normal times)			Regime 2 (stress regime)		
	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.
mean illiq. premium(T)	16.8 bps	14.1 bps	14.6 bps	31.4 bps	24.8 bps	25.2 bps
σ	0.024	0.0212	0.0202	0.0702	0.0572	0.0491
	(0.0018)	(0.0018)	(0.0017)	(0.0044)	(0.0033)	(0.0033)
correlation parameters	$\rho_1^{Short,Long}$	$\rho_1^{Short,Med.}$	$\rho_1^{Med.,Long}$	$\rho_2^{Short,Long}$	$\rho_2^{Short,Med.}$	$\rho_2^{Med.,Long}$
	0.2759	0.6468	0.704	0.2899	0.4849	0.7628
	(0.1032)	(0.0612)	(0.038)	(0.0718)	(0.0639)	(0.0303)
transition probabilities		$p_{1,1}$			$p_{2,2}$	
		0.8994			0.8282	
		(0.0267)			(0.0415)	
$Constant$	0.0308	0.0263	0.0195	0.0303	0.052**	0.0352
	(0.0314)	(0.0241)	(0.0326)	(0.0242)	(0.0183)	(0.0262)
$illiq_{t-1}(T)$	0.5222**	0.641**	0.7695**	0.4672**	0.3801**	0.5988**
	(0.1087)	(0.0668)	(0.08)	(0.0971)	(0.1263)	(0.0989)
$illiq_{t-2}(T)$	0.1724	0.137*	0.229**	0.1679*	0.0662	0.0925
	(0.0889)	(0.0652)	(0.0621)	(0.0786)	(0.0713)	(0.0794)
$illiq_{t-3}(T)$	0.1695*	0.1108*	0.0182	0.017	0.1033	0.155*
	(0.079)	(0.0443)	(0.0549)	(0.07)	(0.0646)	(0.0675)
$illiq_{t-1}(2)$		−0.0857	−0.0449		0.0246	0.0411
		(0.0654)	(0.0899)		(0.051)	(0.051)
$illiq_{t-1}(5)$	−0.0329		−0.0493	0.1142		0.0411
	(0.0974)		(0.0584)	(0.1578)		(0.0834)
$illiq_{t-1}(8)$	0.0428	0.0826*		0.0964	0.3199**	
	(0.0396)	(0.0383)		(0.1155)	(0.103)	
$invRisk_t$	0.0044	0.002	0.0002	0.0052	0.0074*	0.0065*
	(0.005)	(0.0036)	(0.0032)	(0.0053)	(0.0036)	(0.0032)
$trblVol_t(T)$	−0.0268	−0.0299	−0.0341	0.0142	−0.1756*	−0.0839
	(0.0722)	(0.059)	(0.0846)	(0.077)	(0.0726)	(0.1012)
$liquPref_t$	0.003	0.0051	0.0032	0.0252**	0.0032	−0.0045
	(0.0058)	(0.0038)	(0.0046)	(0.0079)	(0.0061)	(0.0054)
	Log Likelihood		N	AIC	BIC	
	4573.51		700	-9011.02	-8626.84	

Table 11: Robustness Check: Dynamic Linkages
This table shows the results of the maximum likelihood estimation of model (11), (12), (13), (6), and (7) with $p = 3$ and $q = 1$. White's (1982) standard errors are given in parentheses. *, ** indicate significance at the 5% or 1% level using the Wald chi-squared statistics $W = (R\hat{\alpha} - r)'(R\hat{V}R')^{-1}(R\hat{\alpha} - r)'$. The observation period is May 7th, 1997 to September 29th, 2010.

	Before Financial Crisis 1997 until May 2007			Financial Crisis June 2007 until Sep. 2010		
	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.	Short T=2 yr.	Medium T=5 yr.	Long T=8 yr.
mean illiq. premium(T)	14.0 bps	11.8 bps	12.8 bps	46.9 bps	36.9 bps	35.9 bps
$Constant$	0.0901**	0.0224*	0.0109	0.055	−0.1572	−0.0755
	(0.0134)	(0.0093)	(0.0089)	(0.1048)	(0.0831)	(0.0631)
$illiq_{t-1}(T)$	0.2961**	0.6735**	0.6767**	0.6358**	0.6397**	0.7007**
	(0.0505)	(0.0446)	(0.0656)	(0.084)	(0.0997)	(0.0772)
$illiq_{t-2}(T)$	0.0793	0.0404	0.1874**	0.2873**	0.1952*	0.1941
	(0.0527)	(0.0711)	(0.0669)	(0.0975)	(0.0954)	(0.1042)
$illiq_{t-3}(T)$	0.2013**	0.2138**	0.1105	−0.1169	−0.0959	−0.08
	(0.0468)	(0.0606)	(0.0704)	(0.0894)	(0.0769)	(0.0857)
$invRisk_t$	0.0082**	0.0025	−0.0001	0.0041	0.0059**	0.0058*
	(0.0021)	(0.002)	(0.0011)	(0.0029)	(0.002)	(0.0026)
$trblVol_t(T)$	−0.0983**	−0.0602*	−0.0376	−0.0222	1.0191*	0.4058
	(0.0231)	(0.0305)	(0.036)	(0.5359)	(0.4461)	(0.2483)
$liquPref_t$	0.0047	−0.0005	−0.0006	0.0228**	0.0143*	0.0075
	(0.0027)	(0.0026)	(0.0021)	(0.007)	(0.0064)	(0.0055)
adj. R^2	0.4778	0.8525	0.9305	0.9413	0.9214	0.9253
N		526			174	

Table 12: Robustness Check: Crisis Identification
This table shows the results of the estimation of two unconditional multivariate regression models. Newey and West's (1987) standard errors with five lags are given in parentheses. *, ** indicate significance at the 5% or 1% level. The total observation period is May 7th, 1997 to September 29th, 2010.