The Term Structure of Risk Premia: Evidence from CDS Spreads

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Abstract

This study estimates the term structure of risk premia before and during the 2007/2008 financial crisis using a new approach based on credit default swaps. Credit markets offer the unique possibility to estimate risk premia for distinct maturities, which considerably facilitates the estimation of the term structure of risk premia. We find that the risk premium term structure was flat before the crisis and downward sloping during the crisis. We provide several robustness tests to show that our results are not driven by counterparty risk, a biased measurement of the real-world default probability, or liquidity.
1 Introduction

It is important to understand the dynamics of expected returns and risk premia in asset markets. There is ample evidence that risk premia are time-varying and mean-reverting (Shiller (1981), Fama and French (1989), Cochrane (1991), Campbell and Viceira (1999)). Thus, risk premia are not only a function of time, but also a function of the maturity. This observation directly raises the question about the shape of the term structure of risk premia. For example, it has been argued that risk premia have increased significantly during the 2007/2008 financial crisis, thereby aggravating the worldwide decline in asset prices. This in turn raises a set of questions: how can this claim be scrutinized in a solid methodological framework? And if risk premia have indeed increased, was this increase uniform across maturities – i.e. a parallel shift in the term structure of risk premia – or did risk premia increase more for short-term assets than for long-term assets? While there is a wide body of research on the term structure of risk-free interest rates (Vasicek (1977), Cox, Ingersoll, and Ross (1985), Engle and Robins (1987), Kempf, Korn, and Uhrig-Homburg (2012)), little to nothing is known about the shape and the dynamics of the term structure of risk premia. Furthermore, it has been demonstrated that the time series behavior and the term structure of risk premia are important for several applications, for example for strategic asset allocation decisions and asset pricing (Balvers and Mitchell (1997), Flannery and Harjes (1997), Campbell and Viceira (1999), Barberis (2000), Lynch (2001), Li (2007), Koijen, Rodriguez, and Sbuelz (2009)).

Using historical data, risk premia over a given period are notoriously difficult to estimate with precision, even under a constant risk premium assumption. Time-varying risk premia are even more difficult to estimate because of the additional source of error (Poterba and Summers (1988)). In this paper, we propose a new methodology for estimating an implied term structure of risk premia from credit spreads which circumvents many of the drawbacks of using historically realized returns. We apply this methodology to estimate the term structure of risk premia before and during the recent financial crisis.

Our approach uses credit spreads to estimate the term structure of risk premia. Throughout this paper, the risk premium is measured as the CDS-implied market Sharpe ratio, i.e. the excess return of the market portfolio per unit of standard deviation as implied by CDS spreads. We thereby build on the link between credit spreads and risk premia in structural models of default established by Duffie and Singleton (2003),
Campello, Chen, and Zhang (2008) and Berg and Kaserer (2009). The main idea is to use physical and risk-neutral returns to estimate risk premia. Our approach is based on the simple observation that in structural models of default, the difference between physical and risk-neutral default probabilities depends on the Sharpe ratio of the companies’ assets (Duffie and Singleton (2003)). Recent research has shown that structural models are useful in modeling the behavior of credit spreads when contamination by non-credit factors is accounted for (Schaefer and Streibulaev (2008), Ericsson, Reneby, and Wang (2006), Ericsson, Kris, and Oviedo (2009), Giesecke, Longstaff, and Streibulaev (2011)).

To further support the validity of our analysis we also investigate the simple difference between physical and risk-neutral default probabilities – thereby not relying on any specific structural model calibration.

Using credit spreads has one key advantage: risk premia can be extracted for each maturity separately because credit instruments are – unlike equities – available for a variety of distinct maturities. We are therefore able to estimate a whole term structure of risk premia for each date which significantly facilitates the estimation of the term structure of risk premia. Furthermore, credit markets offer a variety of further advantages for estimating risk premia. Firstly, they provide an ex ante measurement of risk premia and therefore avoid the limitations of ex post averages of realized returns (see Campello, Chen, and Zhang (2008) for a similar argument). Secondly, real-world quantities (probability of default, recovery rate) only need to be estimated over a finite horizon, which avoids the terminal value problem usually associated with implied equity premium estimates.

We use more than 150,000 credit default swap (CDS) spreads from the constituents of the main CDS indices in the U.S. (CDX.NA.IG) and Europe (iTraxx Europe) from April 2004 to March 2009 for the standard CDS maturities of 3, 5, 7, and 10 years. We follow Berndt, Douglas, Duffie, Ferguson, and Schranz (2004) and Garlappi, Shu, and

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1. This is in particular an advantage when comparing our approach to implied cost of capital estimates based on analyst forecasts. Analyst forecasts are only available for a limited time horizon, usually up to five years, and publicly available analyst forecasts are known to be biased and affected by institutional factors (Brown (1993), Barber, Lehavy, and Trueman (2007)).

2. CDS spreads offer a variety of advantages over bond spreads. In particular, they are less contaminated by non-default components (Longstaff, Mithal, and Neis (2005)), they are standardized across companies and maturities, and evidence suggest that price discovery is faster in the CDS market than in the bond market (Blanco, Breman, and Marsh (2005)). By tracking the major indices in the US and Europe we avoid the inclusion of very illiquid reference entities which may be subject to matrix pricing. In addition, this approach results in a constant sample size and ensures that our results are not driven by the expansion of the CDS market over time.
Yan (2008) and measure real-world default probabilities using expected default frequencies from KMV. The resulting term structure of risk premia was flat before the 2007/2008 financial crisis and downward sloping during the 2007/2008 financial crisis.

The magnitude of these changes is economically significant. Applying our results to stocks and weighting the changes in the risk premium term structure by the maturity of the cash-flows, we find that the equivalent flat increase in the discount rate for stocks amounts to 2-3%. The change is larger for shorter duration stocks and smaller for longer duration stocks. Approximately half of the decline in share prices can be attributed to higher discount rates. The correlation of monthly changes in the 10-year CDS-implied market Sharpe ratio with S&P 500 returns is -0.34, so an increase in Sharpe ratios is, on average, accompanied by a decrease in equity returns. Furthermore, we find that CDS-implied Sharpe ratios exhibit a positive value but – in contrast to previous literature (Fama and French (1993)) – a negative value premium. However, our findings can be consistent with a positive value premium on equities. Expected risk premia on stocks are weighted average risk premia across all maturities. Cash-flows of value stocks occur on average earlier than cash-flows of growth stocks (Lettau and Wachter (2007)). Therefore, short-term risk premia have a greater weighting than long-term risk premia for value stocks. Given that the term structure of risk premia is, on average, downward sloping over our sample period, a positive (maturity-weighted) average value premium on equities can be consistent with a negative constant-maturity value premium.

We conduct several robustness tests to support our results. In particular, we provide evidence that our results are not driven by counterparty risk, a biased measurement of the real-world default probability, or liquidity. Plain vanilla CDS contracts are usually subject to collateral requirements, and therefore the effect of counterparty risk on CDS spreads is extremely small (Gandhi, Longstaff, and Arora (2011)). In addition, we show that taking into account counterparty risk even strengthens our results and leads to even higher risk premia during the crisis and an even steeper downward sloping term structure of risk premia. Concerning the estimate of the real-world default probability, we conducted two robustness tests. Firstly, we use an alternative measurement for the default probability based on a hazard rate specification. Secondly, we set the short-term real-world default probabilities for the first 3 years as high as the default rates observed during the worst 3 years of the last century (which includes the experience of the 1930s), while still using our contemporaneous estimates for the long-term default probabilities. This
drives down the risk premium portion of short-term CDS spreads while leaving long-term risk premium estimates unchanged, thereby flattening the risk premium term structure as far as one could imagine. Both robustness tests confirm our results of an increase in risk premia during the crisis and a downward sloping risk premium term structure during the crisis. The literature provides mixed results on the influence of liquidity on CDS spreads (Ericsson, Reneby, and Wang (2006), Bühler and Trapp (2008) and Bongaerts, Driessen, and De Jong (2011)). There is, however, a strong argument as to why our results are not driven by liquidity: liquidity is an inverse U-shaped function of the time-to-maturity, with the 5-year maturity being the most liquid and the 3, 7 and 10-year maturities being less liquid. Therefore, liquidity cannot be the cause for the downward sloping term structure that we observe.

Our work is related to the empirical literature on the estimation of risk premia. The literature can be classified in approaches using historically realized returns (Poterba and Summers (1988), Campbell and Viceira (1999), Dimson, Marsh, and Staunton (2008)), implied estimates (Claus and Thomas (2001), Gebhardt, Lee, and Swaminathan (2001), Fama and French (2002), Pastor, Sinha, and Swaminathan (2008), Bliss and Panigirtzoglou (2004), Aït-Sahalia and Lo (2000)) or surveys (Graham and Harvey (2008)) to determine risk premia for stocks, stock options or bonds. None of these studies does, however, differentiate according to the maturity of the asset under consideration. Notable exceptions are Lettau and Wachter (2011) and van Binsbergen, Brandt, and Koijen (2010). Lettau and Wachter (2011) calibrate a parsimonious model for the pricing kernel which is able to capture well-known characteristics of the term structure of risk-free interest rates, the time series of stock returns and the cross-section of stock returns. Van Binsbergen, Brandt and Koijen (2012) use prices of short-term and long-term dividend strips and analyze differences in return characteristics between these two assets. Both van Binsbergen, Brandt, and Koijen (2010) and Lettau and Wachter (2011) use historical returns to determine risk premia and to calibrate the respective models. An inherent disadvantage of using historically realized returns is the noise associated with these estimates and therefore the large sample period which is necessary to yield significant results (van Binsbergen, Brandt, and Koijen (2010)). Our approach allows us to determine an implied term structure of risk premia, and we are therefore able to observe changes in risk premia over very short time horizons. For example, we find – in line with Lettau and Wachter (2011) and van Binsbergen, Brandt, and Koijen (2010) – that expected Sharpe ratios are, on average, higher for short-duration assets than for long-duration assets, i.e.
the term structure of risk premia is downward sloping on average. In addition, we are also able to observe the dynamics of this term structure of risk premia over time: while it was clearly downward sloping during the financial crisis, we also observe periods in which the term structure of risk premia is upward sloping.

The paper is organized as follows: section 2 presents the theoretical framework. Section 3 describes our data sources. Section 4 provides empirical results for the estimation of the term structure of risk premia. Robustness tests are shown in section 5, before section 6 concludes.

2 Model setup

This section proceeds in two steps: firstly, we present a formula for the credit risk premium which is a simple model-free proxy for CDS-implied risk premia. Secondly, we use a structural model of default to derive an expression for the CDS-implied market Sharpe ratio. The market Sharpe ratio has the advantage that it already captures asset/market-correlations and is scaled by the volatility of the market.

2.1 Credit risk premium

Inspired by Elton, Gruber, Agrawal, and Mann (2001) and Hull, Predescu, and White (2004), we start by analyzing the credit risk premium, which is a very simple and intuitive proxy for risk premia. The credit risk premium is defined as the difference between the credit spread and the annualized real-world expected loss over the maturity of a bond or CDS. It is therefore the excess return above the risk-free rate that an investor can expect to earn by holding a bond or CDS until maturity. More formally,

\[
CRP(t, \tau) = s(t, \tau) - EL^P(t, \tau),
\]

where \(CRP(t, \tau)\) denotes the credit risk premium at time \(t\) for a maturity of \(\tau\), \(s(t, \tau)\) denotes the credit spread at time \(t\) for a maturity of \(\tau\), and \(EL^P(t, \tau)\) is the annualized average real-world expected loss at time \(t\) over the next \(\tau\) years. The credit risk premium can be determined for each company separately. However, for ease of notation, we skip the company subscript \(i\). The annualized real-world expected loss can be determined as

\[
EL^P(t, \tau) = (1 - (1 - PD^P(t, \tau))^{1/\tau}) \cdot (1 - RR(t, \tau)),
\]
where $PD(t, \tau)$ denotes the cumulative real-world default probability between $t$ and $t+\tau$, and $RR(t, \tau)$ denotes the average real-world expected recovery rate between $t$ and $t+\tau$.

Although intuitive, the credit risk premium is not theoretically well motivated as a proxy for risk aversion. We therefore turn to a structural framework in the next subsection to derive a formula for the CDS-implied market Sharpe ratio.

### 2.2 CDS-implied Sharpe ratio

Similar to Duffie and Singleton (2003) and Berg and Kaserer (2009), the following estimator for the market Sharpe ratio $SR_M$ can be derived in a Merton framework with constant asset volatilities and constant Sharpe ratios:

$$
\theta_t = SR_M = \frac{\Phi^{-1}(PD^Q(t, \tau)) - \Phi^{-1}(PD^P(t, \tau))}{\sqrt{\tau}} \frac{1}{\rho_{V,M}},
$$

where $PD^Q$ and $PD^P$ denote the cumulative risk-neutral and actual default probability of the firm, $\tau$ denotes the maturity and $\rho_{V,M}$ denotes the asset/market correlation. Details are provided in appendix A.1. Formula (2) provides an estimate of the market Sharpe ratio for each company-date combination separately, i.e. for each date, we can estimate a CDS-implied market Sharpe ratio for each company in our sample.

Although this estimator is derived in a simple Merton framework, Berg (2010) has shown that the estimator is still robust in a first-passage time framework and a framework with unobservable asset values (Duffie and Lando (2001)). Huang and Huang (2003) also show that given a certain actual default probability, the risk-neutral default probability is almost the same for the main structural models of default in the literature. They analyze a model with stochastic interest rates, strategic default models and a model with mean-reverting leverage ratios. Their analysis also includes a model with time-varying asset risk premium.\footnote{The Sharpe ratio of the underlying firm’s asset value process can be estimated by omitting the $\rho_{V,M}$-term.}

\footnote{The list of these models is by far not exhausting, see for example Broadie and Detemple (2004) for an overview of option pricing models. The core idea of Huang and Huang (2003) and Berg (2010) is that changes in modeling assumptions usually have a very similar effect on both the risk-neutral and the real-world probability of default. For example, using a first-passage time framework instead of a Merton-type framework approximately doubles model-implied default probabilities due to the reflection principle. However, because both risk-neutral and the real-world probabilities of default are affected in the same way, the ratio of both will remain almost unchanged.}
Appendix A.2 shows that the Sharpe ratio estimator above is also a good estimator in a model with time-varying Sharpe ratios if the constant Sharpe ratio $SR_M$ is substituted with its real world arithmetic average:

$$SR_M(t, \tau) := \frac{1}{\tau} E^P \left[ \int_t^{t+\tau} \theta(s) ds \right] \approx \frac{\Phi^{-1}(PD^Q(t, \tau)) - \Phi^{-1}(PD^P(t, \tau))}{\sqrt{\tau}} \frac{1}{\rho_{V,M}},$$  \hspace{1cm} (3)

where $\theta(s)$ is the time-$s$ instantaneous market Sharpe ratio.

The risk-neutral default probability $PD^Q$ can be derived from a one-period CDS spread $s$ and a recovery rate $RR$ via the standard model-free relationship (cf. Duffie and Singleton (2003))

$$s = PD^Q \cdot (1 - RR) \iff PD^Q = \frac{s}{1 - RR},$$

assuming that the timing of spread cash-flows and default cash-flows are the same, i.e. either both at the beginning or both at the end of the period. In continuous time, and denoting the CDS spread for a maturity of $\tau$ at time $t$ as $s(t, \tau)$ and the recovery rate as $RR(t, \tau)$, this yields\(^5\)

$$PD^Q(t, \tau) = 1 - e^{-s(t,\tau)/(1-RR(t,\tau))\cdot\tau}. \hspace{1cm} (4)$$

The idea of using credit spreads to estimate risk premia was first used by Campello, Chen, and Zhang (2008). Campello, Chen, and Zhang (2008) analyze the cross-sectional variation in expected returns using an instantaneous Taylor approximation of a Merton model. The elasticity of the equity value with respect to the bond value is determined with a pooled panel regression based on historical data. In the Merton model, this elasticity also depends on the maturity of the bond. While this maturity-dependence may not have been a crucial issue for the cross-sectional analysis of Campello, Chen, and Zhang (2008), it is of utmost importance for our goal of estimating a term structure of risk premia. Therefore, we explicitly model this maturity-dependence via (3).

\(^5\)Since defaults in the Merton model are assumed to occur at maturity only while spreads are paid quarterly, the formally correct relationship would $s_{\text{cum}} = PD^Q(t, \tau) \cdot (1 - RR(t, \tau))$ where $s_{\text{cum}}$ denotes the end value of all cumulative spread payments over the maturity of the CDS. Using the latter formula does not significantly change our results.
3 Data sources and descriptive statistics

The sample consists of weekly observations of 3, 5, 7 and 10-year CDS spreads from on-the-run companies of the main CDS index in the U.S., the CDX.NA.IG, from April 2004 until March 2009. The CDX.NA.IG index consists of 125 members and is rolled over every six months (at the end of March and the end of September). The index is comprised of the most liquid investment grade entities as determined by a dealer poll at each rolling date. The first series of the CDX.NA.IG was launched on 20 November 2003. Our CDS data starts in 2004, with our sample period thus beginning with the second series in April 2004. By following the major CDS index, we ensure that our results are not driven by variations in sample size. In addition, CDS included in this major index are thought of as being more liquid and frequently traded, meaning that issues about illiquidity or slow price discovery should not be a major problem.

For each company/week observation, the credit risk premium and the market Sharpe ratio was estimated based on (3) and the following input parameters: 1) CDS spreads, 2) recovery rate, 3) real-world cumulative default probabilities and 4) correlation between asset returns and market returns.

We use CDS mid spreads for 3, 5, 7 and 10-year maturities, which are by far the most liquid maturities in the CDS market. Because CDS are OTC derivatives, no spread data from exchanges or other trading platforms exist. Mayordomo, Pena, and Schwartz (2010) analyze six major sources for corporate CDS spreads and find that the CMA (Credit Markets Association) database quotes lead the price discovery process in comparison with quotes from other databases. We therefore use CMA as our source for CDS data. Together with each CDS spread, CMA provides a veracity score which indicates whether the spread is based on an actual trade, a firm bid or other sources (e.g. indicative bid, bond spread derived). A date/company combination is only included in our sample if either trades or firm bids have been reported for all maturities (3, 5, 7, 10-year) in that respective week. Therefore, our sample size is identical for all maturities.

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6Based on data provided by DTCC (http://www.dtcc.com/products/derivserv/data/index.php), CDS trade volume in constituents of the CDX.NA.IG make up more than 2/3 of the overall trade volume in single name CDS in the period from December 2009 - June 2010. Trade volume data for our period is not available, but, given the growth that the CDS market has seen since 2004, the proportion is likely to be even larger in our sample period.

7The commencement date in April 2004 also allows us to use the same sample period for Europe, where the main CDS index, the iTraxx Europe, was launched on 20 March 2004.

8We put a significant effort in ensuring that the CDS spreads we use are indeed reliable. In particular,
We use three different specifications for the recovery rate. Firstly, we use a flat recovery rate of 40% before and during the crisis, which obviously reflects an optimistic estimate for the expected recovery rate during the crisis. Secondly, we use a recovery rate of 40% before the crisis and 0% during the crisis, which reflects a conservative estimate for the period during the crisis. Thirdly, we use an approach which is based on the observation that recovery rates are negatively correlated with default rates and mean-reverting (Altman, Brady, Resti, and Sironi (2005)). Frye (2000) concludes that in a severe recession, recovery rates drop by up to 20-25 percentage points from their normal-year average. We therefore follow Altman, Brady, Resti, and Sironi (2005) and estimate the one year in advance expected recovery rates via a log/log-regression based on data from Moody’s (2004):

\[
\ln(RR_{t+1}) = -2.274 - 0.338\ln(DR_{t+1}),
\]

where \(RR_{t+1}\) is the recovery rate in year \(t\), \(DR_{t+1}\) is the default rate in year \(t\) and standard errors are given in parenthesis. To avoid a look-ahead bias, we substitute the default rate \(DR_{t+1}\) with the probability of default in year \(t\).\(^{11}\) We estimate the recovery rate for years \(t+2, t+3, \ldots, t+10\) based on the mean-reversion process

\[
RR_{t+1} = \kappa_{RR}RR_{t} + (1 - \kappa_{RR})(RR_{t} - \bar{RR})
\]

where \(\kappa_{RR} = 0.588\) and the long-term mean \(\bar{RR} = 45.5\%\) were estimated based on Moody’s (2004). The resulting average one year in advance recovery rates are 52.42% before the crisis and 33.00% during the crisis, which is a drop of roughly 20 percentage points.

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\(^9\) A recovery rate of 40% is the long-term unconditional mean recovery rate on senior unsecured bonds, cf. Moody’s (2004).

\(^{10}\) Moody’s (2004) provides data from 1982-2003, i.e. before the start of our sample period. The regression coefficients are almost identical when using more recent data, and we therefore opted against a rolling regression, using a constant specification instead.

\(^{11}\) The regression is based on default rates for all companies, including junk bond issues. We therefore do not limit the estimation of the probability of default to companies in our sample (i.e. in the respective indices). We use a time-varying recovery rate, but the same for each company at a specific date. Our estimator is almost linear in the recovery rate, therefore cross-sectional variations will cancel out.
points, and therefore consistent with the observation of Frye (2000). Longer maturity recovery rates are closer to the long-term mean of 45.50%. In the following section, we only report results based on this third methodology for the recovery rate modeling; results for the other two options are available on request. It was reassuring to see that our results are also qualitatively steadfast when using the flat 40% recovery rate assumption or when setting the recovery rate as low as zero percent during the financial crisis.\textsuperscript{12}

The actual cumulative default probability was determined based on expected default frequencies (EDFs) from Moody’s KMV. Moody’s KMV provides EDFs from 1 to 10-year maturities based on the distance-to-default measurement. To calibrate the distance-to-default to default probabilities, KMV uses its proprietary database of historical default events.\textsuperscript{13} Robustness tests were conducted and are presented in section 5.

Correlations between asset returns and market returns were proxied by the rolling 2-year median industry correlation between the corresponding equity returns and market returns.\textsuperscript{14} The data were taken from Datastream. Using equity instead of asset correlations is justified, as equity is a deep-in-the-money call on the company’s assets. Therefore, the delta of an equity position is approximately one and the gamma approximately zero, meaning that asset and equity correlations are almost identical.

Finally, the sample period was split into two sub-periods: 'Before Crisis' (April 2004-June 2007) and 'During Crisis' (July 2007-March 2009). Of course there is no single starting date of the 2007/2008 financial crisis\textsuperscript{15}, therefore our division of the sample pe-

\textsuperscript{12}Our recovery estimates are a combination of real-world and risk-neutral recovery rate estimates. The recovery rates reported in Moody’s (2004) are based on the market price of debt 90 days after default. Systematic risk between this date and the date of ultimate recovery is therefore accounted for, while systematic risk prior to this is not. To the extent that this risk is material, we therefore overestimate the risk-neutral recovery rate. This would lead to an upward bias in our CDS-implied Sharpe ratio estimates.

\textsuperscript{13}Therefore, using KMV EDFs to estimate risk premia does not imply any circular arguments, since the level of default probabilities does not rely on any Sharpe ratio or drift assumptions made by Moody’s KMV.

\textsuperscript{14}Another possibility is to use implied correlations, e.g. from index tranches. The particular correlation assumption used has only a minor effect on our result. Even setting the correlation as high as 1 during the financial crisis would not change our main results. Median industry correlations were taken for robustness reasons. Since the correlation enters our formula in the denominator, estimation errors result in an upwardly biased estimator for the market Sharpe ratio. Industry medians have lower standard errors than a company-by-company estimation and also allow the inclusion of companies without a 3-year equity price history. At each date, we estimated the 2-year weekly correlation between the performance index of the respective stocks in each industry and the S&P 500 index.

\textsuperscript{15}Although our data sample ends in March 2009 and the 'During Crisis' sub-period spans from July
period is – to a certain extent – arbitrary.\textsuperscript{16} As early as February 2007, HSBC announced losses of $10bn related to subprime mortgages. In April 2007, New Century Financial, one of the biggest mortgage lenders in the U.S., declared bankruptcy. The crisis accelerated in June and July 2007 when Bear Stearns had to inject $3.2bn to bail out two of its hedge funds and when Moody’s and Standard & Poor’s downgraded more than 250 subprime RMBS. The Dow Jones index peaked as late as October 2007. However, our main conclusions also hold true when choosing Q2 2007 or Q4 2007 as a starting point for the 2007/2008 financial crisis. It is not the aim to show that certain risk premia changes happened \textit{exactly} at the beginning of the crisis. The intention is rather to demonstrate that the implied risk premia gradually changed throughout the turmoil.

All in all, the study uses 22,819 weekly observations for each maturity (14,416 before the crisis, 8,403 during the crisis). The descriptive statistics for the input parameters are shown in table 1.

Insert Table 1 about here

4 Results

4.1 Credit risk premium

We begin by analyzing the question of whether the term structure of credit risk premia changed during the financial crisis. The credit risk premium – as defined in (1) – is the difference between the CDS spread and the annualized real-world expected loss, and it therefore represents the expected excess return of holding a CDS until maturity. We determine (1) for each company-date combination in our sample separately, the following results are based on simple averages. Before the financial crisis, investors could, on average, earn an expected excess return of 23 bp by buying into a 3-year maturity CDS and 64 bp for a 10-year maturity CDS (cf. table 2).\textsuperscript{17} The fact that longer maturities provide a higher excess return is in line with predictions of standard structural models. During


\textsuperscript{17}To put this into an historical perspective, Giesecke, Longstaff, and Strebulaev (2011) find an average credit risk premium of 80 bp over the 1866-2008 period.
the crisis, both 3-year and 10-year credit risk premia increased significantly. The term structure first flattened out and then reversed, with 3-year credit risk premia being as much as 30 bp higher than 10-year credit risk premia at the height of the financial crisis (cf. figure 1). The difference-in-difference between 3-year and 10-year credit risk premia is both economically and statistically highly significant (cf. Panel A of table 1). These results support the idea that the term structure of risk premia has changed during the financial crisis, with short-term risk premia increasing significantly more than long-term risk premia.

Although intuitive, these findings raise several questions. Firstly, it is not obvious whether credit risk premia are a good theoretical measurement of risk aversion. We therefore analyze the CDS-implied market Sharpe ratio in the next subsection. Secondly, these observations on credit risk premia can potentially be biased due to a biased measurement of the input parameters (such as the real-world default probability), counterparty risk, or liquidity. We discuss possible measurement errors in great detail in the robustness section (section 5). However, we can already predict intuitively that these distortions would need to be quite significant. The difference-in-difference estimate of 42 bp means that the 3-year expected loss during the financial crisis needs to be increased on average by 42 bp relative to the 10-year expected loss to yield a zero difference-in-difference estimate. Taking into account the fact that the 3-year period is also a sub-period of the 10-year period, any measurement error of the 3-year period carries over to the 10-year period. Therefore, the 3-year expected loss needs to be increased by 60 bp (= 42bp · 10/7). This translates into a 3-year expected loss of roughly 180 bp and an increase in the 3-year default probability of 2.9%, i.e. from 2.2% (cf. table 1) to 5.1%. This would imply a measurement error of more than 100% on average during the financial crisis and result in a 3-year default probability which is more than twice the default rate of the worst three years in the 1920-2010 history of Moody’s for investment grade obligors.

4.2 CDS-implied market Sharpe ratio

Using structural models of defaults allows for a more rigorous method of analyzing changes in risk premia. We use (3) to derive CDS-implied market Sharpe ratio estimates from
CDS spreads for each company-date combination in our sample. The following results are based on simple averages per date or per period. Results are depicted in figure 2 and Panel B of table 2. The CDS-implied Sharpe ratio estimates are between 30% and 60% before the 2007/2008 financial crisis. The estimates are very similar for all maturities, meaning that the term structure of Sharpe ratios is flat. The absolute magnitude is also similar, although slightly higher, compared to studies on the equity market, for example Fama and French (2002) report a realized Sharpe ratio of 44% for the period 1951 - 2000 and 31% for the period 1872 - 2000.

The time series of CDS-implied Sharpe ratios shows two significant changes: the first occurs around April 2005, when CDS-implied Sharpe ratios increase by approximately 20PP (from 30% to 50%). This is likely to be due to the downgrades of Ford and General Motors which resulted in an overall repricing in the CDS markets.\textsuperscript{18} CDS-implied Sharpe ratios increase for all maturities simultaneously and by the same amount. Both before and after April 2005, 3, 5, 7 and 10-year CDS-implied Sharpe ratios rarely deviate by more than 5PP at any date.

The second significant change occurs at the beginning of the financial crisis in July 2007. The change differs decisively from the one observed in April 2005. CDS-implied Sharpe ratios increase for all maturities. The increase is, however, much more pronounced for short-term maturities (3-year, 5-year) than for longer maturities (7-year, 10-year). As a result, the term structure of risk premia changes from being flat to being inverse, and this inverse shape is persistent throughout the financial crisis (cf. figure 2). Indeed, the inverse nature of the risk premium term structure was prevalent in every week without exception from mid 2007 until the end of our sample period. The difference-in-difference estimate of the slope before and during the crisis is 26.77\% (10-year minus 3-year slope) and 11.04\% (10-year minus 5-year slope), and both estimates are statistically significant at the 99\% level (cf. Panel B of table 2).

To control for size and book-to-market (BM) effects both for the level and the slope of the CDS-implied term structure of Sharpe ratios, we restrict the sample to 3-year and \textsuperscript{18}Ford and General Motors were downgraded to junk status on 5 May 2005 by S&P, but this is likely to have been anticipated beforehand in the bond and CDS markets due to a negative outlook status. For example, Hull, Predescu, and White (2004) find that spread changes lead rating announcements by up to 90 days.
10-year CDS-implied Sharpe ratios and run the regression

\[ SR_M = \beta_0 + \beta_1 \cdot CrisisDummy + \beta_2 \cdot 3yrDummy + \beta_3 \cdot CrisisDummy \cdot 3yrDummy \]
\[ + \beta_4 \cdot BM + \beta_5 \cdot size + \beta_6 \cdot BM \cdot 3yrDummy + \beta_7 \cdot size \cdot 3yrDummy + \epsilon, \]

(7)

where \( CrisisDummy \) is equal to one for the period 07/2007-03/2009 and zero otherwise, and \( 3yrDummy \) is one for the 3-year CDS-implied Sharpe ratios and zero for the 10-year CDS-implied Sharpe ratios. We repeat the same regression by using 5-year and 10-year CDS-implied Sharpe ratios as a robustness check for the 10-year minus 5-year slope of the term structure of Sharpe ratios. Results for the 3-year versus 10-year CDS-implied Sharpe ratio are displayed in column (A) of table 3, while results for the 5-year versus 10-year CDS-implied Sharpe ratio are listed in column (B) of table 3. We find in both specifications that the size of our difference-in-difference estimate is almost unaffected by controlling for book-to-market and size (27.39% versus 26.77% for the 10-year minus 3-year slope and 11.21% versus 11.04% for the 10-year minus 5-year slope).

Furthermore, we find that the CDS-implied Sharpe ratio is negatively related to size and negatively related to book-to-market. While the first finding is consistent with standard literature, the second finding suggests – in contrast to previous literature – that there is a negative value premium. In contrast to the previous literature, we use a constant-maturity approach. Value stocks have a shorter duration than growth stocks. Therefore, short-term risk premia receive a greater weight for value stocks than for growth stocks when determining average implied risk premia. Given that the term structure of risk premia is, on average, downward sloping over our sample period, a positive (maturity-weighted) average value premium on equities can be consistent with a negative constant-maturity value premium.

How are the changes in CDS-implied Sharpe ratio estimates related to stock returns? If they do reflect overall risk aversion in the market, we would expect that an increase in CDS-implied market Sharpe ratios be accompanied by a decrease in stock prices (and vice versa). Indeed, the 12-month rolling correlation of monthly changes in the 10-year
CDS-implied market Sharpe ratio\(^\text{19}\) with S&P 500 returns is negative throughout our sample period (cf. figure 3). An increase in CDS-implied market Sharpe ratios is therefore accompanied by a decrease in equity returns on average. The 12-month rolling correlation varies, with few exceptions, between -0.1 and -0.6 with a correlation of -0.34 for the total sample period. This suggests that a significant part of the decline in asset prices during the financial crisis can be attributed to an increase in risk premia.

\footnote{\textsuperscript{19}We use the 10-year CDS-implied Sharpe ratios to match the longer duration of equities. However, overall results for the 5-year maturity are very similar.}

5 Robustness tests

5.1 Counterparty risk

We are comfortable that our results are not driven by counterparty risk for two reasons. Firstly, empirical studies show that counterparty risk has had only a minor impact on CDS spreads even at the height of the financial crisis. Gandhi, Longstaff, and Arora (2011) use a proprietary data set of CDS transaction prices and show that an increase in the CDS spread of a dealer by 645 basis points maps into only a one-basis-point decline in the price of credit protection. They also argue that this small impact of counterparty risk is consistent with the fact that CDS transactions are usually collateralized, thereby mitigating counterparty risk significantly.

Secondly, even if counterparty risk had a significant impact on CDS spreads, this would imply that our results are even stronger. In particular, taking into account counterparty risk results in an even larger increase of risk premia during the financial crisis and an even more pronounced downward sloping term structure of CDS-implied market Sharpe ratios. Why is this the case? Counterparty risk causes CDS spreads to be lower than they would be in a world free of counterparty risk, because in the presence of counterparty risk, CDS do not provide insurance in all states of the world. Our methodology requires CDS spreads which are free of counterparty risk, i.e. CDS which provide protection in all states of the world where the reference entity defaults. From inspection of CDS on financial institutions, counterparty risk was significantly higher during the crisis than before the crisis, and this increase was particularly pronounced for shorter maturities. Therefore, CDS-implied market Sharpe ratios would have been even higher during the
crisis and the downward slope in the term structure of CDS-implied market Sharpe ratios would have been even more pronounced if we had corrected for counterparty risk.

5.2 Real-world default probability

We perform two robustness tests for the real-world default probability: firstly, we replace the real-world default probability from Moody’s KMV with a hazard rate model. Secondly, we use the default experience of the 1930s as an extreme scenario for the real-world probability of default.\textsuperscript{20}

For the hazard rate model, we draw upon estimates from a quantitative model published by Fitch.\textsuperscript{21} Fitch provides real-world default probabilities based on a hazard rate model similar to Shumway (2001). As covariates, Fitch includes a distance-to-default measurement, financial ratios, the market performance and macro variables. Momentum is also included, which might be particularly important during the financial crisis. Fitch covers non-financials up to a maturity of 5 years and data were available to us up to December 2008. We therefore estimate the term structure of risk premia as (5-year CDS-implied Sharpe ratio) minus (3-year CDS-implied Sharpe ratio), and the sample period is reduced to the period from April 2004 to December 2008.

Results are depicted in Panel A of table 4. The 5-year CDS-implied market Sharpe ratio was higher than the 3-year CDS-implied market Sharpe ratio before the crisis but lower during the crisis, i.e. the term structure of risk premia changed from being upward to being downward sloping. The difference-in-difference estimate is 14.43% and is both economically and statistically significant. The results therefore confirm the conclusion from using EDFs from KMV as a proxy for the real-world default probability.\textsuperscript{22}

\begin{table}[h]
\centering
\caption{Real-world default probability results.}
\begin{tabular}{lcc}
\hline
Maturity & 5-year & 3-year \\
\hline
Before crisis & 14.43 & 0.00 \\
During crisis & -0.57 & -0.57 \\
\hline
\end{tabular}
\end{table}

\textsuperscript{20}We do not use agencies’ ratings, another popular measure for credit risk, because these are through-the-cycle ratings and do not provide point-in-time estimates of the real-world default probability, cf. Löffler (2004b) and Löffler (2004a).
\textsuperscript{21}For details see FitchRatings (2007).
\textsuperscript{22}In addition, we observe that the CDS-implied market Sharpe ratios resulting from using Fitch’s hazard rate model are in general lower than using expected default frequencies from KMV. This is partly due to the different sample, but more to generally higher estimates for real-world default probabilities provided by Fitch. Possible explanations are different default definitions and a different period for estimating the parameters of the models. For our purposes, we are not primarily interested in the absolute level of CDS-implied market Sharpe ratios, but only in the difference between different maturities and the periods before and during the financial crisis. Therefore, it is reassuring to observe that our results are also steadfast under such obviously different approaches for modeling the real-world probability of default.
As an extreme scenario, we next examine the default experience of the 1930s. Moody’s (2008) provides average default rates for investment grade and non-investment grade issuers on a yearly basis beginning from 1920.\footnote{Giesecke, Longstaff, and Strebulaev (2011) provide an even more extensive history of corporate default rates spanning the period from 1866-2008. For our purposes, the data provided by Moody’s (2008) has the advantage that it differentiates default rates by investment grade versus non-investment grade borrowers.} We leave the default probabilities from KMV unchanged for the period before the crisis. During the crisis, we derive 3-year default probabilities from the worst-case experience in the 1930s while leaving the 4-10-year forward default probabilities unchanged. This is a very conservative estimate and drives down both the difference between risk premia before and during the crisis and the slope of the risk premium term structure during the crisis. The highest 1-year default rate for investment grade borrowers was 1.55%, observed in 1938, while the highest 3-year cumulative default rates for investment grade borrowers were 2.62% (1933-1935) and 2.67% (1936-1938). We therefore scale the 3-year expected default frequencies from KMV so that we end up with an average 3-year cumulative default probability of 2.67% during the crisis period.

Results are depicted in Panel B of table 4. The CDS-implied market Sharpe ratios for the period before the crisis are the same as in table 2. During the crisis, the slope of the term structure of market Sharpe ratios is now -26.11% compared to a value of -32.73% based on unscaled EDFs from KMV. The difference-in-difference estimate for the slope of the term structure of market Sharpe ratios is now 20.15%, and therefore still positive and both economically and statistically significant.

5.3 Liquidity and market microstructure effects

There are theoretical arguments as well as several empirical studies which find only a very small impact of liquidity on CDS spreads (Longstaff, Mithal, and Neis (2005), Ericsson, Kris, and Oviedo (2009) and Bühler and Trapp (2008)) while other studies find significant liquidity effects for CDS as well (Bongaerts, Driessen, and De Jong (2011)). Abstracting from the discussion of liquidity effects on CDS in general, there is a good argument why our results, i.e. the inverse term structure of risk premia during the financial crisis, is not driven by liquidity effects: Liquidity is an inverse U-shaped function of the maturity, with 5-year CDS contracts being more liquid than the 3, 7 and 10-year maturities. This U-shaped pattern of liquidity was prevalent both before and during the financial crisis (cf. figure 4). So if liquidity has a major impact on CDS spreads, we should observe an
inverse U-shaped pattern of CDS-implied Sharpe ratios and not a downward sloping term structure as we do. In our opinion, this is the most striking evidence against liquidity effects being the driver of our results.

Insert Figure 4 about here

5.4 Further robustness tests

We have run two further robustness tests: firstly, we extend the regression (7) to control for factors that might not be adequately captured in the distance-to-default measure used by KMV. These factors include i) a forward-looking measure for volatility, the implied volatility from short-term at-the-money option prices from Bloomberg, to capture expected volatilities instead of the historical volatilities used by KMV, ii) a measure for information uncertainty, the coefficient of variation of I/B/E/S-forecasts, to control for the effect of information uncertainty on short-term default probabilities (Duffie and Lando (2001), and iii) measures for return skewness, rolling historical stock return skewness and skewness implied by I/B/E/S-forecasts, to control for a potential asymmetry in the distribution of asset returns. We also include the CDS bid-ask spread to formally control for liquidity effects. The results are both qualitatively and quantitatively very similar to the results from table 3.\textsuperscript{24}

Secondly, we replicated our results for the main European CDS index (iTraxx Europe) to see whether our findings are valid in a hold-out sample of non-U.S. borrowers. Again, we find very similar results for Europe, in particular a flat term structure of CDS-implied market Sharpe ratios before the crisis and a downward sloping term structure during the crisis.\textsuperscript{25}

6 Conclusion

This paper analyzes the term structure of risk premia before and during the 2007/2008 financial crisis. We use a novel approach based on CDS spreads and real-world default probabilities, which allows us to derive a CDS-implied term structure of risk premia. We show that this CDS-implied term structure of risk premia was flat before the 2007/2008 financial crisis and inverse during the crisis. This approach is – to the best of our knowledge

\textsuperscript{24}These results are available upon request.
\textsuperscript{25}These results are available upon request.
– the first approach which is suited to measuring an implied term structure of risk premia.

Our results suggest that risk premia had a significant effect on stock price changes during the financial crisis. Stock prices and implied risk premia have a negative correlation of -0.34 and the increase in risk premia has accounted for roughly half of the decline in stock prices. These effects are predominantly due to changes in short-term risk premia required by investors. Long-run implied risk premia were of a similar magnitude before and during the financial crisis. Our results demonstrate the economic significance that changes in risk premia had on prices during the financial crisis and that this impact was greatest for short-duration assets.
References


A Estimating Sharpe ratios from CDS Spreads

A.1 Merton framework

This section briefly reviews the procedure for estimating Sharpe ratios from CDS spreads. It is mainly based on the (theoretical) results from Berg and Kaserer (2009). The asset value $V_t$ is modeled as a geometric Brownian motion with volatility $\sigma$ and drift $\mu = \mu_V$ (actual drift) and $r$ (risk-neutral drift) respectively, i.e. $dV_t^P = \mu V_t dt + \sigma V_t dB_t^P$ and $dV_t^Q = r V_t dt + \sigma V_t dB_t^Q$, where $B_t$ denotes a standard Wiener process. The company’s debt consists of a single zero-bond and default occurs if the asset value of the company falls below the nominal value $N$ of the zero bond at maturity of the bond.

In this framework, the real-world default probability $PD^P(t, \tau)$ between $t$ and $T = t + \tau$ can be calculated as:

$$PD^P(t, \tau) = P[ V_T < N ] = \Phi \left[ \frac{\ln \frac{N}{V_t} - (\mu - \frac{1}{2} \sigma^2) \cdot \tau}{\sigma \cdot \sqrt{\tau}} \right]. \quad (8)$$

Here, $\Phi$ denotes the cumulative standard normal distribution function. The default probability under the risk-neutral measure Q can be determined accordingly as

$$PD^Q(t, \tau) = Q[ V_T < N ] = \Phi \left[ \frac{\ln \frac{N}{V_t} - (r - \frac{1}{2} \sigma^2) \cdot \tau}{\sigma \cdot \sqrt{\tau}} \right]. \quad (9)$$

Combining (8) and (9) yields a formula for the asset Sharpe ratio $SR_V$:

$$SR_V := \frac{\mu - r}{\sigma} = \frac{\Phi^{-1}(PD^Q(t, \tau)) - \Phi^{-1}(PD^P(t, \tau))}{\sqrt{\tau}}. \quad (10)$$

Using the CAPM relationship between the Sharpe ratio of the company’s assets ($SR_V$), the market Sharpe ratio ($SR_M$) and the correlation between asset returns and market returns ($\rho_{V,M}$)

$$SR_V = SR_M \cdot \rho_{V,M}$$

yields an estimator for the market Sharpe ratio:

$$\hat{\gamma}_{SR_M,Merton} := \frac{\Phi^{-1}(PD^Q(t, \tau)) - \Phi^{-1}(PD^P(t, \tau))}{\sqrt{\tau}} \frac{1}{\rho_{V,M}}. \quad (11)$$
This formula has two features which make it very convenient for estimating risk premia: firstly, it is quite robust with respect to model changes. The estimator uses the difference between (the inverse of the cumulative normal distribution of the) risk-neutral and actual default probability. This difference is barely affected by the choice of the specific structural model of default. Huang and Huang (2003) first demonstrated this robustness based on different models, including a model with jumps in the asset value process and a time-varying equity premium. Berg and Kaserer (2009) analyze a first-passage time framework and a model with unobservable asset values based on Duffie and Lando (2001) to show this robustness.

Secondly, it is robust with respect to noise in the input parameters. If we look, for example, at a BBB-rated obligor with a 5-year cumulative actual default probability of 2.17%, the resulting model-based risk neutral default probability should be either 3.6% (for an asset Sharpe ratio of 10%) or 13% (for an asset Sharpe ratio of 40%). Assuming a recovery rate (RR) of 50% transforms this into a CDS spread of either 37 bp or 140 bp. This large (model-implied) effect of risk premia on CDS spreads means that the sensitivity of noise in the measurement of CDS spreads (e.g. bid/ask difference) on CDS-implied market Sharpe ratios is low.

A.2 Time-varying risk premia

If Sharpe ratios are time-varying, then the derivation of the Merton framework is no longer valid. However, it can be shown that the formula from the Merton framework still approximates the average expected Sharpe ratio if Sharpe ratios follow a mean-reverting process (e.g. CIR process) with reasonable parameters. For ease of notation, \( \sigma_V = \sigma \), \( \rho_{V,M} = \rho \) and \( W^V_t = W_t \) are used. The real-world default probability can be derived as

\[
P[V_{t+\tau} < L] = P\left[V_t \cdot e^{\int_t^{t+\tau} \sigma \theta_s ds + \int_t^{t+\tau} r_s - \frac{1}{2} \sigma^2 ds + \sigma W_t} < L\right] \\
= P\left[\sigma W_t + \rho \sigma \int_t^{t+\tau} \theta_s ds < \ln \left(\frac{L}{V_t}\right) - (r - 0.5 \sigma^2)\tau\right] \\
\approx \Phi \left[\frac{\ln \left(\frac{L}{V_t}\right) - \sigma \rho E^P \left[\int_t^{t+\tau} \theta_s ds\right] - (r - 0.5 \sigma^2)\tau}{\sigma \sqrt{\tau}}\right]
\]
In the last row, the approximation

\[ \text{Var} \left( \sigma W + \rho \sigma \int_t^{t+\tau} \theta_s ds \right) \]  

(12)

\[ = \text{Var}(\sigma W) + \rho^2 \sigma^2 \text{Var} \left( \int_t^{t+\tau} \theta_s ds \right) + 2 \text{Cov} \left( \sigma W, \int_t^{t+\tau} \theta_s ds \right) \]  

(13)

\[ \approx \text{Var}(\sigma W) \]  

(14)

is used to substitute the integral with its expected value \( E^P \left[ \int_t^{t+\tau} \theta_s ds \right] \). The approximation is justified since \( \theta_s \) is mean-reverting and the volatility of \( \rho \sigma \int_t^{t+\tau} \theta_s ds \) is 'a lot' smaller than the volatility of \( \sigma W \). In addition, the covariance term will usually be negative if it is assumed that negative equity returns go hand in hand with an increase in risk aversion.

Accordingly, the risk-neutral default probability can be calculated as

\[ Q[V_{t+\tau} < L] = \Phi \left[ \frac{\ln \left( \frac{L}{V_t} \right) - (r - 0.5 \sigma^2) \tau}{\sigma \sqrt{\tau}} \right] \]

so that

\[ \frac{\Phi^{-1}(PD^Q(t, \tau)) - \Phi^{-1}(PD^P(t, \tau))}{\sqrt{\tau}} \approx \frac{\sigma \rho E^P \left[ \int_t^{t+\tau} \theta_s ds \right]}{\rho \tau \sigma} \]

\[ \approx \frac{1}{\tau} E^P \left[ \int_t^{t+\tau} \theta_s ds \right] \]

A similar result was shown by Huang and Huang (2003) based on a specific calibration of a mean-reverting model.
Figure 1: Term structure of credit risk premia

This figure presents the credit risk premium for maturities of 3, 5, 7 and 10 years for the period from April 2004 to March 2009. The credit risk premium is defined as the credit spread minus the real-world expected loss based on (1). This figure presents simple averages across all companies in our sample at a given date. The sample consists of the intersection of the KMV database, the CDX.NA.IG on-the-run companies and the CMA CDS database (via Datastream) from April 2004 to March 2009.

Figure 2: Term structure of CDS-implied market Sharpe ratios

This figure presents the CDS-implied market Sharpe ratio for maturities of 3, 5, 7 and 10 years for the period from April 2004 to March 2009. The CDS-implied market Sharpe ratio is derived from a structural model of default based on (2). This figure presents simple averages across all companies in our sample at a given date. The sample consists of the intersection of the KMV database, the CDX.NA.IG on-the-run companies and the CMA CDS database (via Datastream) from April 2004 to March 2009.
Figure 3: Correlation between changes in the 10-year CDS-implied market Sharpe ratio and S&P 500 returns

This figure presents the rolling 12-month Pearson correlation coefficient between changes in the 10-year CDS-implied market Sharpe ratio and returns of the S&P 500 return index for the period from April 2005 to March 2009. The sample consists of the intersection of the KMV database, the CDX.NA.IG on-the-run companies and the CMA CDS database (via Datastream) from April 2004 to March 2009.

Figure 4: Bid/ask spreads by maturity before and during the crisis

Table 1: Descriptive statistics for input parameters

This table presents summary statistics for the input parameters. The sample consists of the intersection of the KMV database, the CDX.NA.IG on-the-run companies and the CMA CDS database (via Datastream) from April 2004 to March 2009. *CDS3/CDS5/CDS7/CDS10* denotes 3, 5, 7 and 10-year CDS spreads in bp. *RR1/RR3/RR5/RR7/RR10* denotes average 1, 3, 5, 7 and 10-year recovery rates. *EDF3/EDF5/EDF7/EDF10* denotes 3, 5, 7 and 10-year cumulative expected default frequencies from Moody’s KMV. $\rho$ denotes the rolling 2-year equity/market correlation based on the S&P 500 return index. All parameters are determined on a weekly basis. Averages are calculated as unweighted averages across all observations.

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<td></td>
<td>N</td>
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<td>CDS3</td>
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<td>$\rho$</td>
<td>14,416</td>
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Table 2: Credit risk premia before and during the crisis

This table presents CDS implied risk premia for maturities of 3, 5, 7 and 10 years before the crisis (April 2004 to June 2007) and during the crisis (July 2007 to March 2009). The sample consists of the intersection of the KMV database, the CDX.NA.IG on-the-run companies and the CMA CDS database (via Datastream) from April 2004 to March 2009. Panel A presents results for the credit risk premium as defined in (1). \( CRP_{3yr}/CRP_{5yr}/CRP_{7yr}/CRP_{10yr} \) denotes the 3, 5, 7, 10-year credit risk premium defined as CDS spread minus real-world expected loss. Panel B presents results for the CDS-implied market Sharpe ratio as defined in (2). \( SR_{3yr}/SR_{5yr}/SR_{7yr}/SR_{10yr} \) denote the 3, 5, 7 and 10-year CDS-implied market Sharpe ratio. All parameters are determined on a monthly basis. The total number of observations is 21,584 (3,427 observations per maturity before the crisis, 1,969 observations per maturity during the crisis). The t-statistics were determined using two-way cluster-robust standard errors (Peterson (2009)). ***, **, * denotes significance at the 1, 5 and 10 percent level respectively.

### Panel A: CDS-implied credit risk premium

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<tr>
<td>CRP 3yr</td>
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<td>129.80</td>
<td>−106.48***</td>
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<td>133.97</td>
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<td>40.60 ***</td>
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<td>t-stat</td>
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### Panel B: CDS-implied market Sharpe ratio

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<td>SR 3yr</td>
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<td>91.47%</td>
<td>−30.63***</td>
<td>(−5.38)</td>
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<td>SR 5yr</td>
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<td>75.26%</td>
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<td>SR 7yr</td>
<td>57.62%</td>
<td>65.73%</td>
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<td>SR 10yr</td>
<td>54.88%</td>
<td>58.75%</td>
<td>−3.87%</td>
<td>(−0.91)</td>
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<tr>
<td>Difference (10yr-3yr)</td>
<td>−5.96% ***</td>
<td>−32.73%***</td>
<td>26.77%***</td>
<td>(12.40)</td>
</tr>
<tr>
<td>t-stat</td>
<td>(−4.68)</td>
<td>(−16.51)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Difference (10yr-5yr)</td>
<td>−5.48% ***</td>
<td>−16.52%***</td>
<td>11.04%***</td>
<td>(9.76)</td>
</tr>
<tr>
<td>t-stat</td>
<td>(−8.01)</td>
<td>(−16.76)</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>
We estimate the determinants of CDS-implied market Sharpe ratios. Column A provides estimates for the determinants of 3-year and 10-year CDS-implied market Sharpe ratios, while column B provides estimates for the determinants of 5-year and 10-year CDS-implied market Sharpe ratios. The sample consists of the intersection of the KMV database, the CDX.NA.IG on-the-run companies and the CMA CDS database (via Datastream) from April 2004 to March 2009. The rows Crisis*3yr (Crisis*5yr) provide estimates for the interaction term between the crisis dummy and the dummy for the 3-year maturity (5-year maturity). These rows therefore provide the difference-in-difference estimates for the change in the slope of the term structure of CDS-implied market Sharpe ratios. Crisis denotes a dummy which is equal to one for the period from July 2007 to March 2009 and zero otherwise. 3-yr-Dummy and 5-yr-Dummy denote dummy variables which are equal to one for 3-year CDS-implied market Sharpe ratios (5-year CDS-implied market Sharpe ratios) and zero otherwise. BM denotes the book-to-market ratio, log(size) the logarithm of the market capitalization in million USD. T-stats based on two-way cluster-robust standard errors are shown in parentheses. ***, **, * denotes significance at the 1, 5 and 10 percent level respectively.

<table>
<thead>
<tr>
<th>Dependent</th>
<th>(A)</th>
<th>(B)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>Crisis*3yr</td>
<td>0.2677*** (12.40)</td>
<td>0.2739*** (12.30)</td>
</tr>
<tr>
<td>Crisis*5yr</td>
<td>0.5875*** (16.19)</td>
<td>0.9235 (4.87)</td>
</tr>
<tr>
<td>Constant</td>
<td>0.0387 (0.91)</td>
<td>0.0338 (0.81)</td>
</tr>
<tr>
<td>3yr Dummy</td>
<td>0.0596*** (16.71)</td>
<td>0.0434 (0.39)</td>
</tr>
<tr>
<td>BM</td>
<td>-0.0558*** (-2.66)</td>
<td>-0.0558*** (-2.66)</td>
</tr>
<tr>
<td>log(size)</td>
<td>-0.0341* (-1.86)</td>
<td>-0.0341* (-1.86)</td>
</tr>
<tr>
<td>BM*3yr</td>
<td>-0.0214*** (-2.82)</td>
<td>0.0027 (0.24)</td>
</tr>
<tr>
<td>log(size)*3yr</td>
<td>0.0027 (0.24)</td>
<td></td>
</tr>
<tr>
<td>BM*5yr</td>
<td>0.0027 (0.24)</td>
<td></td>
</tr>
<tr>
<td>SE clustered at comp. level</td>
<td>yes</td>
<td>yes</td>
</tr>
<tr>
<td>SE clustered at time level</td>
<td>yes</td>
<td>yes</td>
</tr>
<tr>
<td>Adj. $R^2$</td>
<td>12.68%</td>
<td>15.98%</td>
</tr>
<tr>
<td>No. of observations</td>
<td>10,792</td>
<td>8,926</td>
</tr>
</tbody>
</table>
Table 4: Robustness: Estimate for the real-world default probability

This table provides results for two robustness tests using alternative proxies for the real-world default probability. Panel A provides results for using a hazard rate model by Fitch as a proxy for the real-world default probability. The sample consists of the intersection of the Fitch database (data available up to December 2008), the CDX.NA.IG on-the-run companies and the CMA CDS database (via Datastream) from April 2004 to December 2008. The total number of observations is 16,336 (2,812 observations per maturity before the crisis, 1,272 observations per maturity during the crisis). Panel B provides results for using the default experience from the 1930s as a proxy for the real-world default probability. The sample consists of the intersection of the KMV database, the CDX.NA.IG on-the-run companies and the CMA CDS database (via Datastream) from April 2004 to March 2009. The total number of observations is 21,584 (3,427 observations per maturity before the crisis, 1,969 observations per maturity during the crisis). SR3yr/SR5yr/SR7yr/SR10yr denote the 3, 5, 7 and 10-year CDS-implied market Sharpe ratio. The t-statistics were determined using two-way cluster-robust standard errors (Peterson (2009)). ***, **, * denotes significance at the 1, 5 and 10 percent level respectively.

Panel A: Fitch hazard rate model

<table>
<thead>
<tr>
<th></th>
<th></th>
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<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>SR 3yr</td>
<td>23.79% ***</td>
<td>38.90%***</td>
<td>15.10%***</td>
<td>2.93</td>
</tr>
<tr>
<td>SR 5yr</td>
<td>30.18% ***</td>
<td>30.85%***</td>
<td>0.67%</td>
<td>0.15</td>
</tr>
<tr>
<td>Difference (5-yr-3yr)</td>
<td>6.39% ***</td>
<td>-8.04%***</td>
<td>14.43%***</td>
<td>(7.37)</td>
</tr>
<tr>
<td>t-stat</td>
<td>(6.88)</td>
<td>(-4.08)</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Panel B: Default experience 1930s

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
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<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>SR 3yr</td>
<td>60.84% ***</td>
<td>83.50%***</td>
<td>22.66%***</td>
<td>(3.92)</td>
</tr>
<tr>
<td>SR 5yr</td>
<td>60.36% ***</td>
<td>71.89%***</td>
<td>11.53%**</td>
<td>(2.24)</td>
</tr>
<tr>
<td>SR 7yr</td>
<td>57.62% ***</td>
<td>63.60%***</td>
<td>5.99%</td>
<td>(1.26)</td>
</tr>
<tr>
<td>SR 10yr</td>
<td>54.88% ***</td>
<td>57.39%***</td>
<td>2.51%</td>
<td>(0.58)</td>
</tr>
<tr>
<td>Difference (10-yr-3yr)</td>
<td>-5.96% ***</td>
<td>-26.11%***</td>
<td>20.15%***</td>
<td>(9.30)</td>
</tr>
<tr>
<td>t-stat</td>
<td>(-4.68)</td>
<td>(-13.12)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Difference (10-yr-5yr)</td>
<td>-5.48% ***</td>
<td>-14.50%***</td>
<td>9.02%***</td>
<td>(7.84)</td>
</tr>
<tr>
<td>t-stat</td>
<td>(-8.01)</td>
<td>(-14.33)</td>
<td></td>
<td></td>
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