EUROPEAN CENTRAL BANK

EUROSYSTEM

ECB LAMFALUSSY FELLOWSHIP PROGRAMME

WORKING PAPER SERIES

NO 805 / AUGUST 2007

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In 2007 all ECB publications feature a motif taken from the €20 banknote.

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CONTENTS

Abstract

This paper investigates the determinants of the default risk premia embedded in the European credit default swap spreads. Using a modified version of the intertemporal capital asset pricing model, we show that default risk premia represent compensation for bearing exposure to systematic risk and to a new common factor capturing the proneness of the asset returns to extreme events. This new factor arises naturally because the returns on defaultable securities are more likely to have fat tails. The pricing implications of this new factor are not limited to credit markets only. We find that this common factor is priced consistently across a broad spectrum of corporate bond portfolios. In addition, our asset pricing tests also document patterns that are consistent with the so called "flight to quality" effect.

JEL No: G12, G13, G15

Keywords: credit default swap, default risk premium, European credit market, European corporate bond markets, risk factors

Non-technical Summary

This study investigates the pricing of risk in the European credit and corporate bond markets. The typical securities traded on these markets have the unique feature that their payoffs are fixed for the entire life of these securities (e.g. the coupon of a corporate bond and the credit default swap rate are fixed at the time of the issuance) as long as the reference firm or portfolio of firms do not enter events that could trigger corporate default. The mere possibility that such events might occur is typically enough to induce investors to demand a price or return compensation for bearing risk. This risk compensation is known as default risk premium. Our goal with this paper is to understand the determinants of default risk premia.

We proceed by first investigating the theoretical determinants of default risk premia. Most asset pricing models predict that default risk premia should only reflect compensation for bearing systematic risk. To get some guidance on the potential sources of systematic risk we employ a theoretical framework that builds on the discrete intertemporal capital asset pricing model of Campbell (1993). This framework predicts that risk premia, in general, and default risk premia, in particular, can be linked to two types of systematic risk: the market and the news about future discount rates. The later systematic risk is captured with the return of a zero cost portfolio which longs a riskless consol bond and shorts the riskless short rate. Whether these systematic factors are the only determinants of the risk premia of various traded securities depends heavily on the shape of the return distributions of both the systematic factors as well as the actual securities. The return distribution of our factors is close to log-normality. However, the return distribution of defaultable securities, i.e. securities with payoff structures tied to the credit quality of a reference firm or portfolio of firms, is far from being log normal.

Our theoretical framework has two predictions for default risk premia: 1) An increase in the risk premia of the two systematic factors or a higher exposure to these factors leads to higher default risk premia 2) If the return distribution of a defaultable security departs from log-normality the default risk premium of this security depends also on the extent to which its return distribution differ from log-normality. Thus, the common variation in default risk premia has essentially two components: one that captures the exposure to systematic risk factors and another that captures comovement in prices due to exacerbated sensitivity to extreme events (i.e. the return distributions of defaultable securities tend to have fat tails). We next focus on determining the relative importance of each of these components in default risk premia and, more importantly, on understanding the nature of the second component. We do so by first extracting a common component from the part of default risk premia that is not related to systematic risk and then by testing whether this common component helps price a range of test assets.

Our results are based on default risk premia extracted from the credit default swap rates of the most liquid firms on the European credit market during 2003-2006 time period. We find that, on average, the systematic component captures 21% of the time-variation in returns of defaultable assets while the other component - which we call the credit market factor, or CMF for short - captures 63% of this time-variation. More importantly, we find that CMF can help price a wide range of test assets constructed from the non-financial/industrial sector or the entire universe of traded European corporate bonds. These test assets include portfolios of corporate bonds sorted on maturity, rating, maturity/rating and sectors.

The results of this paper suggest that the prices of traded securities in both credit and corporate bond European markets reflect not only compensation for exposure to systematic risk but also compensation for exacerbated sensitivity to extreme events. These results address directly the current asset management practices in these particular markets.

1 Introduction

In the recent years, academics have been studying the behavior of the credit and corporate bond markets through the lense of various measures of default risk premia. Most of the previous studies however focused on the US markets due to data availability reasons. In this paper, we focus exclusively on understanding the behavior of the European credit markets.

Default risk premia represent compensation for bearing the risk embedded in assets whose payoffs are contingent on whether a given firm defaults in a certain period of time. Intuitively, one can think of default risk premia as the difference between the market rate of a credit default swap (CDS) and the expected loss on the same CDS contract. From a traditional asset pricing perspective, one can also think of default risk premia as the expected return on a defaultable corporate bond in excess of the risk-free rate.

The goal of this paper is to understand what drives default risk premia. To this end, we propose a theoretical framework based on the intertemporal capital asset pricing model (ICAPM) of Campbell (1993) to analyze the interaction between systematic risk and returns on zero-coupon defaultable bonds with zero recovery. We chose to analyze the default risk premia of these particular defaultable assets because of their simple payoff structure and because their market values can be inferred relatively easy from the price information of tradable, but more sophisticated, defaultable assets. It should be noted that the original framework of Campbell's ICAPM can not be applied directly here as a conditionally normal model for instantaneous returns is more likely to be mis-specified for defaultable assets than for other assets, such as stocks. Berndt, Douglas, Duffie, Ferguson and Schranz (2005) document that the instantaneous returns on zero-coupon defaultable bonds with zero recovery are more likely to follow conditionally log-normal dynamics as opposed to conditionally normal dynamics.

Our theoretical framework suggests that default risk premia arise as compensation for exposure to systematic risk and to a common factor that captures the proneness of these assets to extreme events. This common factor, which we call the credit market factor (CMF), is the common component of the deviations of the defaultable assets returns from the equivalent returns obtained under an alternative specification which assumes conditional log-normality. The model also suggests that the returns on defaultable assets are impacted by systematic risk through their covariance with two zero-cost portfolios: one that longs the market and shorts the risk-free rate and another that longs a riskless consol bond and shorts the risk-free rate.

Next, we turn to the estimation of these components of the default risk premia. First, we establish a link between the returns on defaultable zero-coupon bonds with zero recovery and the default intensities. Then, we estimate the dynamics of the actual and risk-neutral default intensities for the firms with the most liquid CDS market in Europe for the period 2003-2006. This estimation follows closely the methodology developed in Berndt et al. (2005) by exploiting the relation between the CDS spreads and the risk-neutral default probabilities. Finally, we use the main prediction of our theoretical framework to relate the returns of defaultable assets to the returns on the two zero-cost portfolios proxying for systematic risk and to identify the CMF factor.

We find that the two zero-cost portfolios can explain on average 21% of the time-variation in returns on defaultable assets while the CMF factor can explain on average 63% of the residual. These results suggest that CDS spreads of the firms in our sample incorporate compensation for bearing exposure not only to systematic risk but also to the CMF factor. To understand better the nature of the CMF factor we further investigate the pricing implications of this common factor for the corporate bond markets. We run asset pricing tests on a rich set of test assets consisting of corporate bond portfolios sorted on maturity, rating, maturity/rating and sectors. These portfolios are constructed from the non-financial/industrial sector or the entire universe of the traded European corporate bonds. Our asset pricing tests support overwhelmingly the hypothesis the CMF factor is priced in the corporate bond markets.

We also document another interesting pattern. Most of the corporate bond portfolios load negatively on the excess returns on the market. These loadings become more negative as the maturity of the assets in the portfolio increases and less negative (sometimes even positive) as the rating of the assets decreases. In the asset pricing literature this behavior is referred to as the "flight to quality" effect. As the economy goes through a recession period investors' appetite for risk decreases and they invest in safer assets with longer maturities. As the economy goes through an expansion period investors' appetite for risk increases and they invest in riskier high-yield bonds.

The results in this paper complement and extend the results of Berndt, Lookman and Obreja (2006) who find that the U.S. credit and corporate markets as well as the U.S. equity options market price a common factor that is also extracted from the CDS spreads of the U.S. firms with the most liquid CDS markets for the period 2002-2006. There are some important differences however. First, this study focuses not only on a different time period, but also on a different market. Second, the

CMF factor in this study is extracted from default risk premia that are measured as the expected excess holding returns of zero-coupon defaultable bonds with zero recovery. The default risk premia in Berndt et al. (2006) are measured in terms of the expected loss[1] in a manner similar to the one proposed by Elton, Gruber, Agrawal and Mann (2001). Finally, the theoretical framework proposed in this paper suggests that the CMF factor captures the proneness of the defaultable securities to extreme events. In the other study, the corresponding common factor is shown to capture the jump-to-default risk associated with market-wide credit events.

This study also contributes to the growing financial economics literature concerned with the measurement of the default risk premia. Noticeable contribution to this literature include Elton et al. (2001), Amato and Remolona (2005), Longstaff, Mithal and Neis (2004), Saita (2005), Berndt et al. (2005). This paper distinguishes from all these studies on several dimensions including the choice of capital markets and the methodology. We concentrate exclusively on European credit and corporate bond markets and our theoretical approach shares with the intertemporal capital asset pricing models which previously have not been adapted to accommodate returns of defaultable securities.

The remainder of this paper is structured as follows. Section 2 describes our data and presents a thorough discussion of the general terms of the credit default swap contract and an overview of the Moody's KMV EDF measure of default probability. Section 3 uses a simple measure of default risk premia to present some preliminary evidence supporting the common time-variation in European default risk premia across industries for the period 2003-2006. Section 4 measures default risk premia in terms of the actual and the risk-neutral default intensities. Section 5 present the theoretical determinants of default risk premia and constructs an expected returns-beta representation for defaultable assets. Section 6 estimates the dynamics of default intensities using the information embedded in the CDS spreads and the actual default probabilities as measured by Moody's KMV measure of default. Section 7 estimates the components of default risk premia. The following two sections investigate the nature of the CMF factor. Section 8 tests whether the time-variation in CMF is due to time-varying firm characteristics such as actual default probability, firm size or market-to-book ratio, while Section 9 tests whether the CMF factor is priced in the European corporate bond markets. Finally, Section 10 concludes.

[1] In Section 4 of this paper, we provide an extended discussion on various ways to measure default risk premia

2 Data

This section discusses our data sources for default swap rates and conditional default probabilities in Europe.

2.1 Credit Default Swaps

Credit default swaps (CDS) are single-name over-the-counter credit derivatives that provide default insurance. The payoff to the buyer of protection covers losses up to notional in the event of default of a reference entity. Default events are triggered by bankruptcy, failure to pay, or, for some CDS contracts, a debt restructuring event. The buyer of protection pays a quarterly premium, quoted as an annualized percentage of the notional value, and in return receives the payoff from the seller of protection should a credit event occur. Fueled by participation from commercial banks, insurance companies, and hedge funds, the CDS market has been doubling in size each year for the past decade, reaching $12.43 trillion in notional amount outstanding by mid-2005.[2] In this paper, we use CDS spreads instead of corporate bond yield spreads as our primitive source for prices of default risk because default swap spreads are less confounded by illiquidity, tax and various market microstructure effects that are known to have a marked effect on corporate bond yield spreads.[3]

In particular, we use default swap spreads for five-year CDS contracts for Euro-denominated senior unsecured debt. The data is provided by Credit Market Analysis (CMA) Thomson through Datastream.

It contains daily CDS bid/ask quotes contributed by active market participants including banks, hedge funds and active managers. CMA assures full transparency for its clients by providing a qualifier (Veracity Score) for each data point of any time-series of CDS prices. The Veracity Score indicates the liquidity or if applicable, the extent to which a value has been model-derived. We focus exclusively on firms with very liquid 5-year CDS market for the sample period between January 2003 and November 2006. The CDS contracts of these firms typically make up the iTraxx CDS Europe index of 150 most liquid non-financial 5-year CDS contracts. To mitigate optimally the tradeoff between the microstructure effects of high frequency quotes and

[2] See, for example, the International Swaps and Derivatives Association mid-2005 market survey. The CDS market is still undergoing rapid growth. The notional amount of default swaps grew by almost 48% during the first six months of 2005 to $12.43 trillion from $8.42 trillion. This represents a year-on-year growth rate of 128% from $5.44 trillion at mid-year 2004.

[3] Recent papers that analyze the contribution of non-credit factors to bond yields include Zhou (2005), Longstaff, Mithal and Neis (2004), and Ericsson and Renault (2001).

the statistical power of our tests, we focus on weekly CDS quotes. Most of the quotes have a Veracity Score of 3 or better. This indicates that the quote is associated with an actual trade or that the quote is an indication provided by a market participant. We do not consider quotes with a Veracity Score higher than 3.5. The final sample of default swap rates used in this study consists of 55 firms from eleven European countries and sixteen different industries, based on Moody's industry classification (see Table 1). A typical firm in our sample has 150 (of 196 maximum possible weekly quotes) valid weekly CDS observations. No firm in our sample has fewer than 95 weekly observations.

The fact that our sample has only 55 firms is an important caveat of this paper. The typical major concern with small samples - such as ours - is whether the sample is representative enough to support unbiased results. We believe that dispite its small size, our sample is quite diverse given that the distribution of firms in our sample spans 16 different industries. In addition, since the goal of this paper is to extract information about the compensation rewarding investors for bearing risk, we believe that this information can be extracted more precisely[4] from the quotes on the CDS contracts of those firms with very liquid 5-year CDS markets. To this extent, we are confident that the results in the paper are not biased by the size of our sample.

2.2 Actual Default Probabilities

We use the one-year Expected Default Frequency (EDF) data provided by Moody's KMV as our measure of actual default probabilities. We will discuss this measure only briefly, referring the reader to Berndt et al. (2005) for a more detailed description. The concept of the EDF measure is based on structural credit risk framework of Black and Scholes (1973) and Merton (1974). In these models, the equity of a firm is viewed as a call option on the firm's assets, with the strike price equal to the firm's liabilities. The "distance-to-default" (DD), defined as the number of standard deviations of asset growth by which its assets exceed a measure of book liabilities, is a sufficient statistic of the likelihood of default. In the current implementation of the EDF model, to the best of our knowledge, the liability measure is equal to the firms short-term book liabilities plus one half of its long-term book liabilities. Estimates of current assets and the current standard deviation of asset growth (volatility) are calibrated from historical observations of the firms equity-market capitalization and of the liability

[4]In order to extract this information we use the approach in Berndt et al. (2005) which requires relatively long time-series of prices (or quotes, in our case).

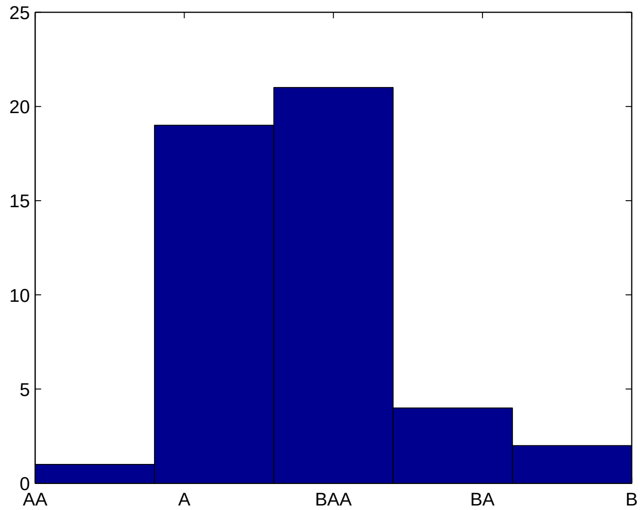

Figure 1: Distribution of firms by median credit rating during the sample period.

measure. For a detailed discussion, see, for example, Appendix A in Duffie, Saita and Wang (2005).

Crosbie and Bohn (2001) and Kealhofer (2003) provide more details on the KMV model and the fitting procedures for distance to default and EDF. Unlike the Merton model, where the likelihood of default is the inverse of the normal cumulative distribution function of DD, Moody's KMV EDF measure uses a non-parametric mapping from DD to EDF that is based on a rich history of actual defaults. Therefore, the EDF measure is somewhat less sensitive to model mis-specification. The accuracy of the EDF measure as a predictor of default, and its superior performance compared to rating-based default prediction, is documented in Bohn, Arora and Korbalev (2005). Duffie, Saita and Wang (2005) construct a more elaborate default prediction model, using distance to default as well as other covariates. Their model achieves accuracy that is only slightly higher than that of the EDF, suggesting that EDF is a useful proxy for the physical probability of default. Furthermore, the Moody's KMV EDF measure is extensively used in the financial services industry. As noted in Berndt et al. (2005), 40 of the worlds 50 largest financial institutions are subscribers.

We obtained daily one-year EDF values from Moody's KMV for the time period January 2001 through October 2006, for the same set of 55 firms described in Section 2.1. Figure 1 plots the distribution of the credit quality of the firms in our sample. As discussed in Section 2.1, our CDS data only start in January 2003. In order to achieve sufficient power for our asset pricing tests we use weekly (Wednesdays) observations of default swap rates, together with EDF values at weekly frequency.

2.3 Interest Rates, Systematic Factors and Test Assets

In Sections 3, 7 and 9 we compute expected loss for CDS contracts and realized excess returns on defaultable securities and corporate bond portfolios, and we form zero-cost portfolios to proxy for systematic risk. In all these instances we need to use information about the Euro term structure of riskless bonds. This data is obtained from Datastream from the Euro zero curves constructed relative to Euribor.[5] All the excess returns and the zero-cost portfolios are computed relative to the 1-month zero yield. Also, the discount factors used to compute the expected loss for CDS contracts in Section 3 are computed using the same Euro zero curves.

For the purpose of Sections 7 and 9 we need to compute zero-cost portfolios that are long the market portfolio and short the 1-month zero yield or long the 30-year zero yield and short the 1-month zero yield. For the later zero-cost portfolio we use the data in the Euro zero curves with the corresponding maturities. For the former zero-cost portfolio, we construct two types of market portfolios: one that incorporates the entire universe of European stocks and one the incorporates only the stocks from a specific country. To maintain consistency with the previous studies on the capital markets integration, we use whenever possible portfolios constructed from the data disseminated in the electronic version of Morgan Stanley's *Capital International Perspectives* (MSCI). For those countries where MSCI data is not available we use the local portfolios constructed by FTSE. All these portfolios are available through Datastream.[6]

Finally, for the purpose of Section 9 we need to compute realized returns on a range of test assets in excess of the 1-month zero yield. We consider the following test assets: the Merrill Lynch non-financial corporate bond portfolios sorted on rating or time-to-maturity, the Merrill Lynch AAA-, AA-, A- and BBB-rated corporate bond portfolios sorted on maturity, and the Lehman Brothers Euro-aggregate industrial corporate bond portfolios sorted on ratings, maturity or sectors. The time-series

[5]The mnemonics for the yield of a zero-coupon Euro bond with time-to-maturity of n years and m months is EMnYm. For instance the mnemonic corresponding to the maturity of 1 year and 4 months is EM01Y04.

[6]The mnemonic for the MSCI European market portfolio is MSEURIL. The mnemonics for the country-specific market portfolios are MSFRNCL (France), MSNETHL (Netherlands), MSGERML (Germany), WISWDNE (Sweden), MSITALL (Italy), MSSPANL (Spain), WIDNMKE (Denmark), WINWAYE (Norway), FTSE10E (UK), MSFINDL (Finland) and MSGDEEL (Greece).

data for all these portfolios comes from Datastream.[7]

3 Preliminary Regression Analysis

In this section we provide a preliminary analysis of the time-series properties of the European default risk premia extracted from the CDS spreads. This analysis is meant to motivate the more thorough analysis of Sections 5-9. As Section 4 will show more clearly, one way to measure the risk premia embedded in the CDS spreads is to compare the market spread with the spread obtained by setting the expected loss of the CDS to zero at the time of the issuance.[8] We denote this later spread with ELS_t. Let S_t denote the actual CDS spread. We can therefore measure the default risk premia embedded in the CDS spreads with $\log(S_t) - \log(ELS_t)$.[9]

To capture the time variation of these risk premia across industries, for a given

[7]These portfolios have respectively the following mnemonics: MLNF3AE, MLNF1AE, MLNF3BE, MLENFAE, MLENFCE, MLENFDE, MLENFGE, MLEC3AE, MLEC3EE, MLEC3GE, MLEC3KE, MLEC2CE, MLEC2GE, MLEC2JE, MLEC1CE, MLEC1GE, MLEC1JE, MLEC1KE, MLEC8CE, MLEC8GE, MLEC8JE, LHAI3AE, LHAI2AE, LHAI1AE, LHAIBAE, LHEHYBA, LHAC1YE, LHAC3YE, LHAC5YE, LHAC7YE, LHAC10E, LHEAEDE, LHEBANK, LHEBMAT, LHECAPG, LHECHEM, LHECOMM, LHACCYE, LHACNCE, LHEDMAN, LHAFBVE, LHALODE, LHAREFE, LHATLPE, LHATBCE, LHAWRSE, and LHAMNCE.

[8]Specifically, we compute this "expected loss spread" as follows: Let $D(t,n)$ denote the discount factor for the period $[t, t+n]$ and $p(t,n)$ the actual survival probability of an obligor over the same period of time. Let L denote the recovery rate (as a percent of the principal) in the event of default. Then we define the expected loss spread as the premium S that solves:

$$\sum_{n=1}^{N} D(t,n)p(t,n)S = \sum_{n=1}^{N} D(t,n)\left[p(t,n-1) - p(t,n)\right]\left[L - \frac{1}{8}S\right]$$

where N is the number of payments stipulated in the original CDS contract (N corresponds to the number of quarters, which for a 5-year contract amounts to 20).

The left-hand side in the above equation is the present value of the future payments by the buyer of the protection, while the right-hand side is the present value of the recovery in the event of default. This later quantity is not straight forward to compute as it requires information about the time of default. We chose to model it in the manner suggested by Berndt et al (2005). In particular, we assume that if the default occurs in the time period $[t+n, t+n+1]$, the protection seller returns to the buyer the fraction L of the principal, less any accrued interest. The actual survival probabilities for maturities longer than one year are estimated as simple products of the one-year survival probabilities obtained from Moody's KMV. The discount factors are computed from the term structure of Euro zero-coupon yields. We use the time-series of Euro riskless term structures relative to Euribor. For more information see Section 2.3. Following Berndt et al. (2005) we assume that L is relatively stable over the sample period at around 75%. This value value corresponds to the medium recovery rate in the US for the period 2002-2006 as documented for instance in Berndt, Lookman and Obreja (2006). We do not have data on recovery rates for the current sample consisting of European firms only.

[9]We use the log specification rather than the simple difference of the two measures because the relation between the CDS spreads and the EDF rates (the actual default probabilities) is more likely

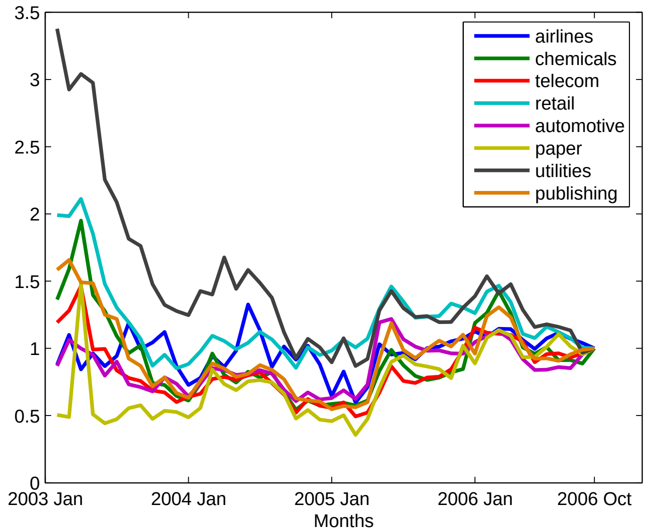

Figure 2: The time-series variation in default risk premia across industries

level of credit worthiness, we run the following panel regression:

$$\log(S_t^i) - \log(ELS_t^i) = \alpha + \beta \log EDF_t^i + \sum_m \sum_p \delta_m^p d_t^i(m,p) + \epsilon_t^i \qquad (1)$$

where $d_t^i(m,p)$ is a dummy variable which equals 1 if week t is in month m and firm i is in industry p.

Figure 2 plots the time-series variation of the monthly estimates for $\exp \delta_m^p$, relative to the last month in our sample.[10] The plot shows that for a given level of credit worthiness, there is substantial variation in risk premia over time. More importantly, the risk premia of the firms in different industries seem to move together. This co-movement is typically indicative of exposure to common risk factors or to the fact that the firms in our sample have similar characteristics (in the spirit of Daniel and Titman (1997)). To the extent that the firms in our sample are different enough from

to be multiplicative rather than linear, as the following regressions show:

$$\begin{aligned} S_t^i &= \underset{(53.6031)}{52.1577} + \underset{(62.1077)}{0.4993}\, EDF_t^i + \epsilon_t^i \\ \log S_t^i &= \underset{(208.9279)}{3.1715} + \underset{(59.8403)}{0.2777} \log EDF_t^i + \epsilon_t^i \end{aligned}$$

[10]Specifically, we plot $\exp\left(\delta_m^p - \delta_{m_0}^p\right)$, where m_0 is the last month in our sample.

each other, the goal of this paper is to disentangle how much of the co-movement in these firms' risk premia is due to "likely" systematic factors[11], and how much is due to other, potentially new, priced factors.

4 Measuring Default Risk Premia

This section describes different ways of measuring default risk premia and it provides a simple characterization of default risk premia in terms of default intensities.

Given a probability space $(\Omega, \mathcal{F}, P)$ and information filtration $\{\mathcal{F}_t : t \geq 0\}$, the default intensity of a firm λ_t^P is the instantaneous mean arrival rate of default, conditional on all current information. Intuitively, conditional on survival to time t and all information available at time t, the probability of default between times t and $t + \Delta$ is approximately $\lambda_t^P \Delta$ for small Δ. In this setting the conditional probability of surviving between t and $T > t$ can be expressed as:

$$p(t, T-t) = E_t \left[\mathbf{1}_{\{\tau > T\}} \right] = E_t \left[e^{-\int_t^T \lambda_s^P \, ds} \right]. \tag{2}$$

where τ denotes the default time and E_t denotes the expectation operator conditional on the information available up to and including time t.

Under the absence of arbitrage and market frictions there exists a stochastic discount factor, M.[12] Moreover, under mild technical conditions, Harrison and Kreps (1979) and Delbaen and Schchermayer (1999) show that there exists a "risk-neutral" probability measure associated with M. Let Q denote this measure. Note that in our setting, markets are not necessarily complete, so the stochastic discount factor and the associated risk-neutral measure might not be unique. This pricing approach nevertheless allows us to express the price at time t of a security paying Z at time $T > t$, as $E_t \left[M_{t,T} Z \right] = E_t^Q \left[e^{-\int_t^T r_s \, ds} Z \right]$, where r is the short-term interest rate and E_t^Q denotes the expectation operator conditional on the information available up to and including time t, with respect to the equivalent martingale measure Q. In particular, the market value of a defaultable zero-coupon bond that pays one unit of account in the event that a given firm does not default before time T and 0 otherwise is given by:

$$P(t, T-t) = E_t \left[M_{t,T} \mathbf{1}_{\{\tau > T\}} \right] = E_t^Q \left[e^{-\int_t^T r_s \, ds} \mathbf{1}_{\{\tau^Q > T\}} \right], \tag{3}$$

[11] Section 5 will provide some theoretical guideness in determining the systematic factors that should affect capital markets.

[12] See for instance Duffie (2001).

where τ^Q denotes the default time of the firm under the measure Q.

We can make the simplifying assumption that the default time τ^Q can be fully described by a doubly-stochastic exponentially-distributed random variable with intensity λ^Q. In this case the above formula reduces to:

$$P(t, T-t) = E_t^Q \left[e^{-\int_t^T r_s + \lambda_s^Q ds} \right] \tag{4}$$

If investors are risk-neutral (i.e. $M_{t,T}$ degenerates to $e^{-\int_t^T r_s ds}$) or the default event of this firm is idiosyncratic, the market value of such a defaultable bond should only reflect the expected loss, namely:

$$P^L(t, T-t) = E_t \left[M_{t,T} \right] E_t \left[\mathbf{1}_{\{\tau > T\}} \right] = E_t \left[e^{-\int_t^T r_s ds} \right] p(t, T-t) \tag{5}$$

If investors are not risk neutral or if the default event is not diversifiable then the market value of the defaultable bond reflects a risk adjustment relative to the expected loss. This risk adjustment is given by:

$$P^L(t, T-t) - P(t, T-t) = -cov_t \left[M_{t,T}, \mathbf{1}_{\{\tau > T\}} \right] \tag{6}$$

This measure of risk compensation is particularly appealing since both terms in the left-hand side of this equation can be computed relatively easy, once the dynamics of the default intensities are known. However, for the purpose of this paper, we are more interested in relating the dynamics of the default intensities to a more traditional measure of risk, namely the risk premium. Let $R_{t+1} = P(t+1, T-t-1)/P(t, T-t)$ denote the gross holding return on the defaultable zero-coupon bond. Then the risk premium of this asset is defined through the Euler equation as follows:

$$E_t R_{t+1} - R_{t+1}^f = -R_{t+1}^f cov_t \left[M_{t,t+1}, R_{t+1} \right] \tag{7}$$

where R^f denotes the gross return on the risk-free bond. Unless we make relatively strong assumptions about the dynamics of the default intensities,[13] neither side of the above equation are easy to relate to the dynamics of the default intensities. Nevertheless, for the rest of this section we present a special case which allows us to establish this relation in a relatively straight-forward manner.

When, $t = T - 1$, it can be easily shown that the risk premium and the risk

[13]For instance, if the default intensities follow Gaussian processes under both the physical and the risk-neutral measure, the risk premium can be computed in closed form.

adjustment relative to expected loss are identical:[14]

$$\frac{P_t^L - P_t}{P_t} = \frac{E_t R_{t+1} - R_{t+1}^f}{R_{t+1}^f} \tag{9}$$

Moreover, when the length dt of the time interval $[t-1,t]$ is sufficiently small, the left-hand side becomes:

$$e^{\left[\lambda_t^Q - \lambda_t^P\right]dt} - 1 \approx \left[\lambda_t^Q - \lambda_t^P\right] dt. \tag{10}$$

Thus, for defaultable bonds with very short maturities, the risk premium per unit of time equals the difference between the risk-neutral and the actual default intensity times the gross return on the risk-free rate. However, this need not be the case if either $t < T-1$ or the time to maturity of the defaultable bonds is large. The next section, presents a simple way to deal with the potentially complex relation between risk premia and default intensities, for the general case. It also addresses the more general question of the likely determinants of the risk premia on defaultable bonds.

5 Theoretical Determinants of Default Risk Premia

In this section we use the discrete intertemporal capital asset pricing model of Campbell (1993) to identify likely sources of macroeconomic risk and to understand the impact of these sources of risk on the prices of defaultable bonds.

Suppose the economy is populated with identical agents with non-expected-utility preferences of the following form:

$$U_t = \left\{(1-\beta)C_t^{\frac{1-\gamma}{\theta}} + \beta\left(E_t\left[U_{t+1}^{1-\gamma}\right]\right)^{\frac{1}{\theta}}\right\}^{\frac{\theta}{1-\gamma}} \tag{11}$$

where γ is the coefficient of relative risk aversion, σ is the elasticity of intertemporal

[14] This identity can be stated in a slightly more general version, for any $t < T$:

$$\frac{P_t^L - P_t}{P_t} = \frac{E_t R_{t,T} - R_{t,T}^f}{R_{t,T}^f} \tag{8}$$

where $R_{t,T}$ is the holding return between t and T, while $R_{t,T}^f$ is the yield of a riskless zero-coupon bond that matures at T. This is merely a consequence of the fact that for zero-coupon bonds (riskless or defaultable), $R_{t,T} = R_{t,t+1}R_{t+1,T}$ and thus the Euler equation holds at larger horizons.

substitution and $\theta = \sigma\frac{1-\gamma}{\sigma-1}$.[15] As Epstein and Zin (1989, 1991) show, the first order condition of the representative agent in this economy can be stated as:

$$1 = E_t \left[\left\{ \beta \left(\frac{C_{t+1}}{C_t} \right)^{-\frac{1}{\sigma}} \right\}^{\theta} \left\{ \frac{1}{R^m_{t+1}} \right\}^{1-\theta} R^i_{t+1} \right] \tag{12}$$

where C is the aggregate consumption, R^m_{t+1} is the return on the market portfolio and R^i_{t+1} is the return on a security i.

Campbell (1993) shows that under the assumption that asset returns and consumption growth are jointly conditionally homoskedastic and log-normally distributed the aggregate budget constraint can be exploited to substitute out consumption and to simplify the Euler equation to:

$$E_t r^i_{t+1} - r^f_{t+1} = -\frac{1}{2}V_{ii} + \gamma V_{im} + (\gamma - 1)V_{ih} \tag{13}$$

where r^* $(* = i,\, f)$ denotes log returns, $V_{ii} = Cov_t(r^i_{t+1}, r^i_{t+1})$, $V_{im} = Cov_t(r^i_{t+1}, r^m_{t+1})$ and $V_{ih} = Cov_t(r^i_{t+1}, (E_{t+1} - E_t)\sum_{j=1}^{\infty} \rho^j r^m_{t+1+j})$. The second argument of the last covariate captures the news about the future returns on the market. ρ is the steady-state ratio of invested wealth to total wealth.[16]

Furthermore, if r^b_{t+1} denotes the return on a riskless consol bond that pays one unit of account every period, Campbell (1993) shows that the above equation can be further simplified to:

$$E_t r^i_{t+1} - r^f_{t+1} = -\frac{1}{2}V_{ii} + \gamma V_{im} + (1 - \gamma)V_{ib} \tag{14}$$

where $V_{ib} = Cov_t(r^i_{t+1}, r^b_{t+1})$.

Let $r^{b,\perp}_{t+1} = r^b_{t+1} - \beta^{b,m}_t r^m_{t+1}$ with $\beta^{b,m}_t = \frac{cov_t(r^b_{t+1}, r^m_{t+1})}{V_{mm}}$. Substituting r^b_{t+1} in the above equation yields:

$$E_t r^i_{t+1} - r^f_{t+1} = -\frac{1}{2}V_{ii} + \left[\gamma + \beta^{b,m}_t(1-\gamma)\right] V_{im} + (1-\gamma)V^{\perp}_{ib} \tag{15}$$

where $V^{\perp}_{ib} = Cov_t(r^i_{t+1}, r^{b,\perp}_{t+1})$. If we further assume that $r^{b,\perp}_{t+1}$ and the consumption growth are both jointly conditionally homoskedastik and log-normally distributed, we can apply the above relation to both r^m_{t+1} and $r^{b,\perp}_{t+1}$. Using the unconditional versions

[15] For more details on the parameters see Campbell (1993).

[16] See Campbell (1993) for the exact definition.

of these relations we obtain:

$$\begin{aligned} \left[\gamma + \bar{\beta}^{b,m}(1-\gamma)\right] &= \frac{Er_t^{m,e}}{V_{mm}} - \frac{1}{2} \\ 1-\gamma &= \frac{Er_t^{b,\perp,e}}{V_{bb}^{\perp}} - \frac{1}{2} \end{aligned} \tag{16}$$

where E denotes the unconditional expectation operator, $\bar{\beta}^{b,m} = E\beta_t^{b,m}$, $r_t^{m,e} = r_t^m - r_t^f$ and $r_t^{b,\perp,e} = r_t^{b,\perp} - r_t^f$. Substituting these formulas back into (15) and taking expectations yields the following expected returns - beta representation:

$$Er_t^{i,e} + \frac{1}{2}V_{ii} = \beta_{im}\left[Er_t^{m,e} + \frac{1}{2}V_{mm}\right] + \beta_{ib}^{\perp}\left[Er_t^{b,\perp,e} + \frac{1}{2}V_{bb}^{\perp}\right] \tag{17}$$

where $\beta_{im} = V_{im}/V_{mm}$ and $\beta_{ib}^{\perp} = V_{ib}^{\perp}/V_{bb}^{\perp}$.[17]

The expected returns-beta representation in equation (17) suggests that the time variation in returns is mainly due to time variation in the returns on the market portfolio in excess of the riskless short rate and the time variation in the returns of a portfolio that longs a riskless console bond and shorts the riskless short-rate. Close relatives of this later portfolio have been previously used in the financial economic literature. One of the best known is the spread between long- and short-term treasury bonds, or TERM, for short. For the exact definition see Fama and French (1993).

The representation in equation (17) applies to any returns that are both jointly homoskedastik and conditionally log-normally distributed with the consumption growth and the market return. However, returns on certain assets are less likely to satisfy the later condition. For instance, Berndt et al. (2005) document that the instantaneous excess returns on defaultable zero-coupon bonds are more likely to be log-normally distributed rather than normally distributed (recall that the instantaneous returns are natural logs of the gross returns). Thus, the above pricing equation might not work as well for this type of returns. Under certain conditions, the expected return-beta representation model in (17) can be slightly generalized to accommodate returns that are not necessarily conditionally log-normally distributed. We describe this modified model bellow.

Suppose the returns on a defaultable bond r_t^D can be decomposed into a component, $r_t^{D,c}$, that is jointly homoskedastic and log-normally distributed with the consumption growth and the market return and another component, $r_t^{D,n}$, that is

[17]Notice that $\beta_{i,m}$ and β_{ib} are in fact the conditional betas, which happen to be constant under the homoskedasticity assumption. Thus they can be different from the unconditional betas.

orthogonal on the information contained on both the consumption-growth and the market.[18] This later component is going to capture the impact of the departure from the conditional log-normality assumption on prices. Under these assumptions it can be easily shown that the expected returns - beta representation in equation (17) becomes:

$$Er_t^{D,c,e} + \frac{1}{2}V_{DD}^c = \beta_{Dm}^c \left[Er_t^{m,e} + \frac{1}{2}V_{mm} \right] + \beta_{Db}^{c\perp} \left[Er_t^{b,\perp,e} + \frac{1}{2}V_{bb}^{\perp} \right] + Ez_t \tag{18}$$

where $r_t^{D,c,e} = r_t^{D,c} - r_t^f$, $V_{DD}^c = var_t(r_{t+1}^{D,c})$, $\beta_{Dm}^c = cov_t(r_{t+1}^{D,c}, r_{t+1}^m)/V_{mm}$, $\beta_{Dm}^{c\perp} = cov_t(r_{t+1}^{D,c}, r_{t+1}^{b,\perp})/V_{bb}^{\perp}$, and $z_t = -\log E_t e^{r_{t+1}^{D,n}}$. Making use of the fact that $r^{D,n}$ is orthogonal on the information contained in the market returns and the long-short treasury portfolio[19], we can rewrite the above as:

$$Er_t^{D,e} + \frac{1}{2}V_{DD}^c = \beta_{Dm} \left[Er_t^{m,e} + \frac{1}{2}V_{mm} \right] + \beta_{Db}^{\perp} \left[Er_t^{b,\perp,e} + \frac{1}{2}V_{bb}^{\perp} \right] + E\Delta z_t \tag{19}$$

where $\Delta z_t = E_t r_{t+1}^{D,n} - \log E_t e^{r_{t+1}^{D,n}}$.

Thus, just like conditionally log-normal returns, the returns of defaultable bonds vary over time in response to changes in excess market returns and the returns on the long-short treasury portfolio. However, unlike conditionally log-normal returns, the returns of defaultable bonds also move because of changes in the shape of the conditional distribution relative to a normal distribution (captured by Δz_t). This later source of time variation could host both a time-varying common component as well as undiversifiable firm-specific components. Both these types of components affect the level of expected returns directly rather than through covariances.

We next focus on computing the returns on defaultable zero-coupon bonds using the methodology developed in Section 4. Following the notation in Section 4, the holding returns between t and $t+1$ for a defaultable zero-coupon bond with maturity

[18]One way to implement such a decomposition is as follows: Let $\mu = Er_t^D$ and $k_r = cov(r_{t+1}, r_t)/cov(r_t, r_t)$. Define $\nu_{t+1} = \left[r_{t+1}^D - \mu - k_r(r_t^D - \mu) \right]$. Let ν_{t+1}^c denote the linear projection of ν_{t+1} onto the space generated by the consumption growth and the market return. Let $\nu_{t+1}^{\perp} = \nu_{t+1} - \nu_{t+1}^c$ denote the orthogonal residual. Since both the consumption growth and the market return are conditionally normally distributed ν_{t+1}^c will be also conditionally normally distributed. In addition since ν_{t+1} has zero mean, both ν_{t+1}^c and $\nu_{t+1}^{\perp}$ can be normalized to have zero mean. Define $r_t^{D,c}$ recursively as follows: $r_{t+1}^{D,c} - \mu = k_r(r_t^{D,c} - \mu) + \nu_{t+1}^c$, with $r_0^{D,c} = r_0^D$. Also, define $r_t^{D,n}$ recursively as follows: $r_{t+1}^{D,n} = k_r(r_t^{D,n} - 0) + \nu_{t+1}^{\perp}$, with $r_0^{D,n} = 0$. Then $r_t^D = r_t^{D,c} + r_t^{D,n}$ and $r_t^{D,c}$ and $r_t^{D,n}$ satisfy the desired properties.

[19]Campbell (1993) shows that the informational content of this portfolio overlaps with that of the market returns and the consumption growth.

$T > t+1$ is given by

$$r_{t+1} = \log P(t+1, T-t-1) - \log P(t, T-t) \tag{20}$$

where $P(t, T-t)$ is defined in Section 4. The holding returns for the period $[T-1, T]$ can be computed with

$$r_T = -\log P(T-1, T) \tag{21}$$

It is important to notice that these returns cannot be computed directly since we do not have data on defaultable zero-coupon corporate bonds. However, we can use the apparatus developed in Section 4 to compute the returns on these hypothetical assets in terms of quantities that can be measured directly or indirectly from the CDS and EDF data that we have available.

We start with the formula in equation (4). Suppose the risk-neutral default intensity $\lambda_t^Q = \lambda^{Q,c} + \lambda^{Q,n}$ such that $\lambda^{Q,c}$ and r_s are correlated Gaussian processes (in particular, they are joint homoskedastik and conditionally normally distributed) and $\lambda^{Q,n}$ is orthogonal on the information contained in the consumption growth rates and the market returns.[20]. Then,

$$\begin{aligned} P(t, T-t) &= E_t\left[M_{t,T} e^{-\int_t^T \lambda_s^Q ds}\right] = E\left[M_{t,T} e^{-\int_t^T \lambda_s^{Q,c}}\right] E_t\left[e^{-\int_t^T \lambda_s^{Q,n}}\right] \\ &= E_t^Q\left[e^{-\int_t^T r_s^f + \lambda_s^{Q,c}}\right] E_t\left[e^{-\int_t^T \lambda_s^{Q,n}}\right] \end{aligned} \tag{22}$$

Since r_t and $\lambda_t^{Q,c}$ are correlated Gaussian processes, it can be easily established that

$$\log E_t^Q\left[e^{-\int_t^T r_s^f + \lambda_s^{Q,c}}\right] = A(T-t) - B(T-t) r_t^f - C(T-t)\lambda_t^{Q,c} \tag{23}$$

where $A(T-t)$, $B(T-t)$ and $C(T-t)$ depend on $T-t$ only.[21] Thus, $\log P(t, T-t)$ can be rewritten as:

$$\log P(t, T-t) = A(T-t) - B(T-t) r_t^f - C(T-t)\lambda_t^{Q,c} + \log E_t\left[e^{-\int_t^T \lambda_s^{Q,n}}\right] \tag{24}$$

[20] See footnote (18) for a way to construct such a decomposition

[21] The coefficients $A(T-t)$, $B(T-t)$ and $C(T-t)$ can be derived in a recursive fashion as it is typically done in the affine term-structure literature. Suppose r_t^f and $\lambda_t^{Q,c}$ follow jointly Gaussian dynamics of the following form:

$$\begin{aligned} r_{t+1}^f &= k_r \bar{r}^f + (1-k_r) r_t^f + \sigma_r \xi_{t+1}^r \\ \lambda_{t+1}^{Q,c} &= k_\lambda \bar{\lambda}^{Q,c} + (1-k_\lambda)\lambda_t^{Q,c} + \sigma_\lambda \xi_{t+1}^\lambda + \sigma_{r,\lambda}\sigma_r \xi_{t+1}^r \end{aligned}$$

Combining we finally obtain the following expression for r_{t+1}:

$$\begin{aligned} r_{t+1} &= \left[B(T-t)r_t^f - B(T-t-1)r_{t+1}^f\right] \\ &+ \left[C(T-t)\lambda_t^{Q,c} - C(T-t-1)\lambda_{t+1}^{Q,c}\right] + \Delta\tilde{z}_{t+1} \\ &= \left[B(T-t)r_t^f - B(T-t-1)r_{t+1}^f\right] \\ &+ \left[C(T-t)\lambda_t^{Q} - C(T-t-1)\lambda_{t+1}^{Q}\right] + \Delta z_{t+1} \end{aligned} \tag{25}$$

where

$$\begin{aligned} \Delta\tilde{z}_t &= A(T-t) - A(T-t-1) + \log E_{t+1}\left[e^{-\int_{t+1}^{T}\lambda_s^{Q,n}}\right] - \log E_t\left[e^{-\int_t^T \lambda_s^{Q,n}}\right] \\ \Delta z_t &= \Delta\tilde{z}_{t+1} - \left[C(T-t)\lambda_t^{Q,n} - C(T-t-1)\lambda_{t+1}^{Q,n}\right]. \end{aligned}$$

Note that Δz_{t+1} measures (up to a constant) the departure from the normal distribution of the conditional distribution of $\lambda_s^{Q,n}$. Given the orthogonality assumptions on $\lambda_t^{Q,c}$ and $\lambda_t^{Q,n}$, we can substitute

$$\begin{aligned} r_{t+1}^{D} &= \left[B(T-t)r_t^f - B(T-t-1)r_{t+1}^f\right] + \left[C(T-t)\lambda_t^{Q} - C(T-t-1)\lambda_{t+1}^{Q}\right] \\ r_{t+1}^{D,c} &= \left[B(T-t)r_t^f - B(T-t-1)r_{t+1}^f\right] + \left[C(T-t)\lambda_t^{Q,c} - C(T-t-1)\lambda_{t+1}^{Q,c}\right] \end{aligned}$$

in equation (15) to obtain the following expected return-beta representation for returns on defaultable bonds:

$$\begin{aligned} E[B(T-t)r_t^f - B(T-t-1)r_{t+1}^f] + E\left[C(T-t)\lambda_t^{Q} - C(T-t-1)\lambda_{t+1}^{Q}\right] - Er_{t+1}^f \\ +\frac{1}{2}V_{DD}^c = \beta_{Dm}\left[Er_{t+1}^{m,e} + \frac{1}{2}V_{mm}\right] + \beta_{Db}^{\perp}\left[Er_{t+1}^{b,\perp,e} + \frac{1}{2}V_{bb}^{\perp}\right] + E\Delta z_{t+1} \end{aligned} \tag{26}$$

Then, for any $t < T$ we have

$$\begin{aligned} A(T-t) &= A(T-t-1) - [B(T-t-1)+1]\,k_r\bar{r}^f - [C(T-t-1)+1]\,k_\lambda\bar{\lambda}^{Q,c} \\ &+ \frac{1}{2}\left[(B(T-t-1)+1) + \sigma_{r,\lambda}\left(C(T-t-1)+1\right)\right]^2\sigma_r^2 + \frac{1}{2}\left[C(T-t-1)+1\right]^2\sigma_\lambda^2 \\ B(T-t) &= [B(T-t-1)+1]\,(1-k_r) \\ C(T-t) &= [C(T-t-1)+1]\,(1-k_\lambda) \end{aligned}$$

with the intial conditions $A(0) = B(0) = C(0) = 0$. Notice that under the decomposition suggested in footnote (18), k_λ can be computed as follows:

$$1 - k_\lambda = \frac{cov\left[\lambda_t^{Q,c}, \lambda_{t+1}^{Q,c}\right]}{var\left[\lambda_t^{Q,c}\right]} = \frac{cov\left[\lambda_t^{Q}, \lambda_{t+1}^{Q}\right]}{var\left[\lambda_t^{Q}\right]}.$$

The pricing equation (26) can be tested on a panel dataset of defaultable bonds with constant maturity $d = T - t$:

$$r_{t+1}^{i,e} = \alpha^i + \beta_{im} r_{t+1}^{m,e} + \beta_{ib}^{\perp} r_{t+1}^{b,\perp,e} + \sum_s f_s 1_{\{t+1=s\}} + \epsilon_{t+1}^i \tag{27}$$

where $r_{t+1}^{i,e} = B^i(d) r_t^f - B^i(d-1) r_{t+1}^f + C^i(d)\lambda_t^{i,Q} - C^i(d-1)\lambda_{t+1}^{i,Q} - r_{t+1}^f$ is the excess realized return on firm's i defaultable zero-coupon bond maturing in $d-1$ periods, f_s captures the value of the potential common component at time s while ϵ_t^i capture the undiversifiable firm-specific component of firm i.

The following section focuses on the estimation of the default intensities from the CDS and EDF data.

6 Estimating the Default Intensities

In this section we first describe the time-series models for both actual and risk-neutral default intensities. Similar to Berndt et al. (2005), we specify a model under which the logarithm of the actual default intensities λ_t^P satisfies the Ornstein-Uhlenbeck equation

$$d \log(\lambda_t^P) \quad = \quad \kappa(\theta - \log(\lambda_t^P))\, dt + \sigma\, dB_t, \tag{28}$$

where B_t is a standard Brownian motion, and θ, κ, and σ are firm-specific constants to be estimated. The behavior for λ^P is called a Black-Karasinski model. (See Black and Karasinski (1991).) This leaves us with a three-dimensional vector $\Theta = (\theta, \kappa, \sigma)$ of unknown parameters to be estimated from available firm-by-firm EDF observations of a given firm. For the 55 firms in our sample we have daily observations of one-year EDFs, from January 2001 to October 2006. However, for the estimation procedure we only use weekly quotes (Wednesdays).

Given the log-autoregressive form (28) of the default intensity, in general there is no closed-form solution available for the one-year EDF, $1 - p(t, 1)$, from (2). We therefore rely on numerical lattice-based calculations of $p(t, 1)$, and have implemented the two-stage procedure for constructing trinomial trees proposed by Hull and White (1994).

With regard to risk-neutral default intensities, we assume that

$$d \log \lambda_t^Q \quad = \quad \kappa^Q(\theta^Q - \log(\lambda_t^Q))\, dt + \sigma^Q\, dB_t, \tag{29}$$

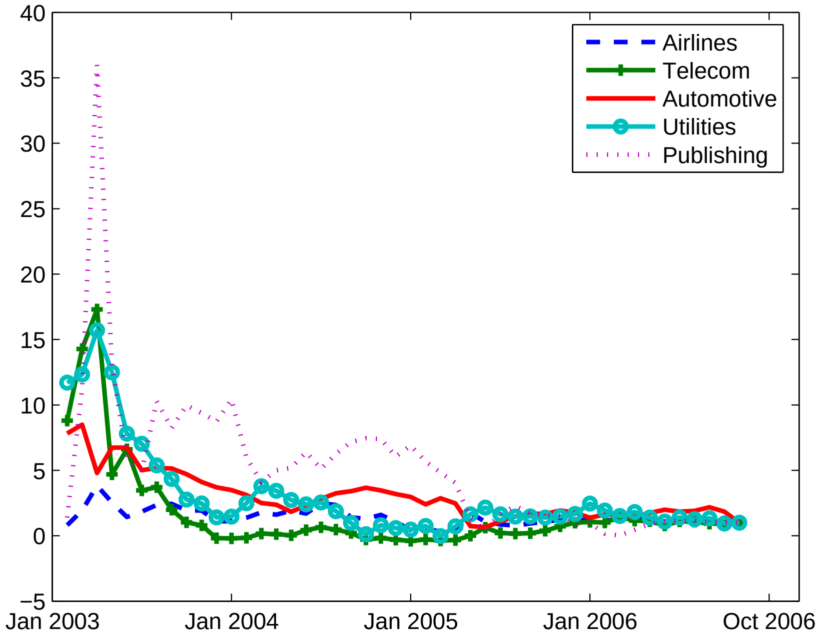

Figure 3: The time variation in $\lambda_t^Q - \lambda_t^P$ across industries.

where B_t^Q is a standard Brownian motion with regard to the physical measure P, and κ^Q, θ^Q, and σ^Q are scalars to be estimated. The risk-neutral distribution of λ^Q is specified by assuming that

$$d \log \lambda_t^Q \quad = \quad \tilde{\kappa}^Q(\tilde{\theta}^Q - \log(\lambda_t^Q))\, dt + \tilde{\sigma}^Q\, dB_t^Q,$$

where $\tilde{\kappa}^Q$ and $\tilde{\theta}^Q$ are constants and B_t^Q is a standard Brownian motion with regard to Q. Given a set of parameters $(\tilde{\theta}^Q, \tilde{\kappa}^Q, \sigma^Q)$, we can compute model-implied values for λ^Q using data on five-year CDS rates and risk-neutral loss given default. For details we refer the reader to Section 5.1 in Berndt et al. (2005). We estimate the parameters driving the dynamics of the risk-neutral default intensities under both phisycal and risk-neutral measure using the over-idendifying restriction $\kappa^Q = \tilde{\kappa}^Q$.[22]

Using maximum likelihood estimation (MLE), we obtain firm-by-firm estimates for the parameters that govern the processes for λ^P and λ^Q. The estimated values of these parameters are shown in Table 2.

As shown in Section 4, the difference between the risk neutral and the actual

[22]This over-identifying restriction improves considerably the reliability of our estimates.

default intensity can be interpreted as a measure of instantaneous risk premia, embedded in the CDS spreads. It would be informative to see whether the time-variation in these measures of risk premia resembles the patterns in Figure 2.

To see this we employ a panel regression approach similar to the one in Section 3 and extract the time-series patterns of the loadings on the dummy variables controlling for month and industry.[23]

Figure 3 plots the estimates of the slope coefficients on the dummy variables controlling for month and industry. We notice that the patters are relatively similar with the ones in Figure 2.

Given the estimated time-series for the actual and the risk-neutral default intensities, we can now compute returns of defaultable zero-coupon bonds and we can formally test the expected returns - beta representation derived in Section 5.

7 The Components of Default Risk Premia

In this section we separate the components of the default risk premia by estimating the returns model described in equation (27). We first discuss our choices for the empirical implementation of the theoretical portfolios proxying for systematic risk, namely the excess market portfolio and the portfolio that is long a riskless perpetuity and short the riskless short rate.

The reference entities behind the CDS contracts in our dataset are from various countries within Europe. Most of these countries are also part of the European Monetary Union[24] but there are few countries that are not (UK, Denmark, Norway and Sweden). Since capital markets thorughout Europe are more or less integrated,[25] we proxy for the market portfolio with both a portfolio tracking the largest stocks throughout Europe as well as local portfolios tracking the largest most liquid stocks within a specific country. To maintain consistency with the previous studies on the capital markets integration, we use whenever possible portfolios constructed from the

[23]Specifically, we run the following regression:

$$\lambda_t^{i,Q} - \lambda_t^{i,P} = \alpha + \sum_m \sum_p \delta_m^p d_t^i(m,p) + \epsilon_t^i$$

where $d_t^i(m,p)$ is a dummy variable which equals 1 if week t is in month m and if firm i is in industry p.

[24]See Table 1 for more details

[25]There is quite a bit of literature on this topic. Some of the most well known studies include Fama and French (1998), Griffin (2002), Ferson and Harvey (1993), Bekaert and Harvey (1995), and Karolyi and Stulz (2003).

data disseminated in the electronic version of Morgan Stanley's *Capital International Perspectives* (MSCI). For those countries where MSCI data is not available we use the local portfolios constructed by FTSE. For more information on these portfolios see Section 2.3. Since the CDS spreads in our dataset are reported relative to the Euro term structure it is important that the returns on these portfolios are extracted from prices reported in Euros. We denote with $r_t^{EMKT,e}$ the weekly returns on the European market portfolio in excess of the riskless short rate and with $r_t^{CMKT,e}$ the weekly returns on the local market portfolio in excess of the riskless short rate. The riskless short rate corresponds to the yield of the one-month zero-coupon Euro bond. For more information on the Euro term structure curves see Section 2.3.

The other portfolio that we have to worry about is the portfolio that longs a riskless console bond paying one unit of account every week and shorts the short interest rate. We proxy for this portfolio with a portfolio that longs the 30-years zero-coupon riskless Euro bond and shorts the 1-month Euro bond. We denote the weekly returns of this portfolio with r_t^{TERM}.

The return model in equation (27) becomes:

$$
\begin{aligned}
r_{t+1}^{i,e} = \alpha^i + \beta_{EMKT}^i r_{t+1}^{EMKT,e} + \sum_c \beta_{CMKT}^i 1_{\{C=c\}} r_{t+1}^{CMKT,e} + \beta_{TERM}^i r_{t+1}^{TERM} \\
+ \sum_s f_s 1_{\{t+1=s\}} + \epsilon_{t+1}^i
\end{aligned}
\tag{30}
$$

where c is an index for countries and

$$
r_{t+1}^{i,e} = B^i(d) r_t^f - B^i(d-1) r_{t+1}^f + C^i(d)\lambda_t^{i,Q} - C^i(d-1)\lambda_{t+1}^{i,Q} - r_{t+1}^f \tag{31}
$$

Using the time series of estimates for the actual and the risk-neutral default intensities derived in the previous section we can compute the excess returns on the left-hand side of the above equation for various times to expiration d. Table 3 shows the averages of the estimated coefficients across firms when the time to maturity takes various values. We noticed some interesting patterns. First, except for the special case when the time to maturity is the shortest possible, namely 1 week, all the pricing errors are large, on average, and statistically significant. The systematic factors suggested by the theory capture only some of time-variation of the excess returns on the defaultable bonds. The loadings for each factor display an increasing pattern as we increase the time to maturity of the defaultable bonds. The most significant, economically and statistically, of the three factors seems to be the $TERM$ factor capturing

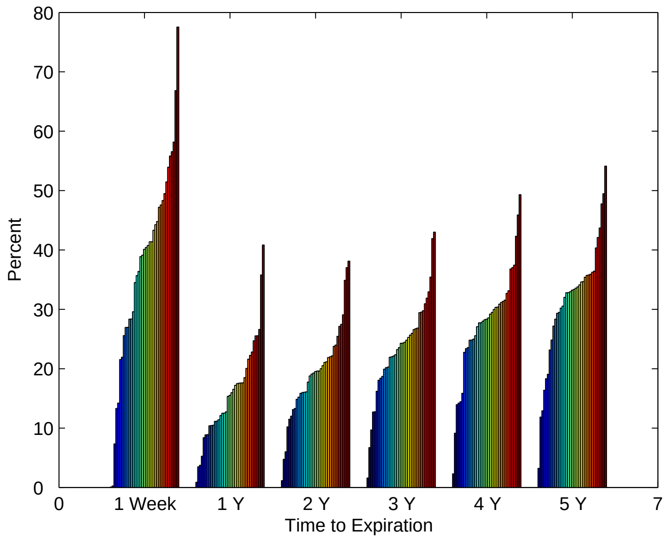

Figure 4: The variance of the CMF factor as a percentage of the variance of the pricing error obtained under the returns model in (30) with no time dummies. The CMF factor is extracted from excess zero-coupon defaultable bonds with time to maturity varying from 1 week to 5 years. Each histogram in the plot shows the firm-specific percentages of the pricing error explained by the CMF factor. Each histogram corresponds to a specific time to maturity, indicated on the horizontal axis.

the spread between the 30-year and the 1-month riskless Euro bonds. Nevertheless, these systematic factors have relatively little explanatory power, as indicated by the R^2s in the last column of Table 3.

Of interest to us, however, is whether the pricing errors obtained from the returns model in equation (30) without time dummies move together over time. The common component of these errors is captured precisely by the loadings on the time dummies, namely f_s. We will refer to this common component as the default risk premia factor, or CMF for short. We notice that CMF can capture quite a large portion of the pricing errors, as indicated by the fifth column in Table 3. Figure 4 shows more descriptively the distribution of the fraction of the pricing error captured by the common component for each firm in our sample and for each time to maturity. This plot shows that the percentage of the pricing error captured by the CMF is very high when the time to maturity is the shortest possible. These percentages are several times lower for times to maturity of 1 year or longer and they display a monotonic pattern in the direction of longer maturities.

What is behind the CMF factor? According to the pricing equation (26), CMF captures the extent to which the conditional distribution of the asset returns differs from the normal distribution. In other words, CMF captures the extent to which certain assets are more likely to have return distributions with fatter tails or non-zero skewness. These attributes are usually indicative of higher sensitivities to extreme events. This suggests that CMF captures the proneness of the defaultable securities to extreme events.

For the rest of the paper we want to understand the nature of this common component. If this common component were to show up in the pricing errors of a large cross-section of defaultable bonds than it can be easily shown that this component cannot be diversified away.[26] Since our cross-section of firms is not too large, however, there is a chance that our common component might capture some the time-variation of firm characteristic that is common for most of our firms. To rule out this possibility we need to control for various firm characteristics and see whether the common component survives. If the common component survives we can understand more about the nature of this component by investigating whether this common component is priced in large portfolios of assets. These tests are performed in the following two sections.

8 Testing for Firm Characteristics

In this section, we study whether the common component CMF arises because the firms in our sample have similar firm characteristics. This is an important step in uderstanding the nature of CMF.

To understand whether the time-variation associated with our common component is actually due to a certain firm characterisitc, we implement the following time-series regressions as suggested by Daniel and Titman (1997):

$$\begin{aligned} r_{t+1}^{i,e} &= \gamma_0^i + \gamma_\phi^i \phi_t + \gamma_{EMKT}^i r_{t+1}^{EMKT,e} + \sum_c \gamma_{CMKT}^i 1_{\{C=c\}} r_{t+1}^{CMKT,e} \\ &+ \gamma_{TERM}^i r_{t+1}^{TERM} + \gamma_{CMF}^i CMF_{t+1} + \epsilon_{t+1}^i \end{aligned} \tag{32}$$

where ϕ_t is the firm characterisitic. This regression tells us that if the common variation in conditional default risk premia (obtained by conditioning on the information

[26] This is essentially a consequence of the fact that under the assumption that the pricing errors in excess of the common component are truly firm specific and are drawn from a common distribution, the variance of a well diversified equally-weighted portfolio of defaultable bonds is given by three components: one which contains the systematic factors, $var(f_s)$ and $var(\epsilon)/N$. When N is large only the last component goes away.

at time t) is in fact due to the firm characteristic rather than the common factor then we should see γ^i_{CMF} being close to zero and statistically insignificant. Table 4 reports the average slope coefficients and their corresponding t-Statistics across firms, for three firm characterisitcs: the actual default probabilities, the firm size and the market-to-book ratio.[27] We notice that in all instances, the slope on the firm characteristic is both economically and statistically significant suggesting that part of the common variation in conditional default risk premia can be attributed to these firm characteristics. However, we are interested in understanding whether our common component arises because of the common variation in these firm characteristics. This turns out not to be the case as can be noticed from the last three columns of Table 4. In all instances, the common component remains both economically and statistically significant. These results suggest that our common component contains important information beyond what is contained in the firm characteristics considered here.[28]

9 Asset Pricing Tests

In this section we want to understand the nature of the common component CMF. We plan to do so by investigating whether this common component has any pricing implications for other classes of assets.

To see whether CMF is priced by assets other than the CDS spreads used to extract the factor, we implement simple asset pricing tests in the spirit of Fama and French (1993). We use as test assets corporate bond portfolios sorted by ratings, time-to-maturity or both, as well as corporate bond portfolios sorted by sector. These portfolios are constructed by either Merrill Lynch or Lehman Brothers and they focus on either the entire universe of European corporate bonds or on the non-financial/industrial sectors. For more information on these portfolios see Section 2.3. These portfolios are particularly attractive because they are sorted on characteristics

[27] Actual default probabilities correspond to the 1-year EDF values (see Section 2 for more details). Firm size and market-to-book ratio are constructed using the firm-level data from Datastream. Whenever the market capitalization of a firm is expressed in a currency other than Euro we convert it into Euros. Firm size is then the log of the market capitalization. Market-to-book ratio is the ratio is recovered the time-series variable PTBV in Datastream.

[28] Another firm characteristic that could be a natural candidate is the leverage ratio. However, the data for the book value of long/short liabilities is only available at annual frequencies from Compustat which means that for our sample we essentially have only three or four datapoints (2003,2004,2005, 2006 when available) for each firm. Thus the time variation in the leverage ratio is essentially driven by the time variation in the market value of equity, which we've already considered as a firm characteristic.

- such as ratings or time-to-maturity - which can be easily related to risk. In particular, the portfolios sorted on these characteristics have different exposures to risk which leads to different average returns.

The asset pricing test that we implement is a time-series regression of the following form:

$$r_{t+1}^{i,e} = \alpha^i + \beta_{EMKT}^i r_{t+1}^{EMKT,e} + \beta_{TERM}^i r_{t+1}^{TERM} + \beta_{CMF}^i CMF_{t+1} + \epsilon_{t+1}^i \qquad (33)$$

where $r^{i,e}$ is the excess return on the portfolio i used as test asset.

The null hypothesis is that the CMF factor loads up more heavily on portfolios that are more exposed to this factor. Recall that the CMF factor captures the extent to which assets are prone to extreme events. As a consequence, we should expect portfolios of corporate bonds with either higher maturities or lower ratings to load up more heavily on CMF. In addition, if the returns model in equation (33) is the true model, then we should also see small and statistically insignificant pricing errors.

Tables 5 to 11 report the estimated coefficients for various portfolios used as test assets. For these tests we use the CMF factor extracted from defaultable zero-coupon bonds with the shortest time to maturity (1 week). Throughout these tables, we notice a striking pattern. In almost all tests, the loadings on the CMF factors are positive and in many cases statistically significant. More importantly, these loadings are higher for portfolios that contain higher maturity or lower rating corporate bonds, just as the model predicted. Figures 5 to 7 show that most of the patterns uncovered when the CMF factor is extracted from defaultable zero-coupon bonds with 1 week until maturity also hold when the CMF factor is extracted from defaultable zero-coupon bonds with 1 year, 2 years, 3 years, 4 years or 5 years until maturity.

The asset pricing tests also indicate that almost all the pricing errors are statistically insignificant. The average size of these errors, in many of the tests, is not however very small, indicating that we should be careful in inferring whether the evidence in Tables 5 to 11 supports the hypothesis that the returns model in (33) is the true model.

Tables 5 to 11 also depict another interesting pattern. Most of the corporate bond portfolios load negatively on the market. These loadings become more negative as the maturity of the assets in the portfolios increases and less negative (and even positive) as the rating of the assets deteriorates. This fact seems to confirm the so called "flight to quality". As the economy goes through an expansion, investors' appetite for risk increases and they're more likely to invest in riskier assets such as high yield (lower rating) corporate bonds. As the economy goes through a recession, investors' appetite for risk turns sour and they prefer to invest in safer assets with longer maturity (such as highly-rated long-term corporate bonds).

10 Conclusion

In this paper we use quotes on CDS contracts of the firms with the most liquid CDS market in Europe to extract the components of default risk premia, measured as the average excess return of a zero-coupon defaultable bond with zero recovery. Extending the theoretical framework of Campbell's ICAPM to accommodate returns with fat-tailed distributions - such as the returns of a defaultable bond - we find that default risk premia have two major components: one associated with systematic risk and another associated with a new common factor that captures the proneness of the asset returns to extreme events. This theoretical framework yields a model of returns for defaultable securities which recognizes as main sources of time variation the returns on the market and a riskless consol bond in excess of the risk-free rate - as proxy for systematic risk - and the returns on the new common factor. To identify the new common factor we apply this model to the returns of defaultable zero-coupon bonds with zero recovery, which, despite the fact that they are not traded, we can compute using the default intensities embedded in the CDS spreads. We find that the two zero-cost portfolios proxying for systematic risk capture on average of 21% of the time-variation in the returns of the zero-coupon defaultable bonds while the new factors captures on average 63% of the residual. Moreover, we find that this new factor is also priced consistently across a broader spectrum of corporate bond portfolios, suggesting that both the European credit and the European corporate bond markets factor in the proneness of defaultable securities to extreme events. These results complement and expand the results of Berndt, Lookman and Obreja (2006) who show that a similar type of factor seems to be priced in the U.S. credit and corporate bond market, as well as the U.S. equity options market. This paper also documents a "flight to quality" effect in the European corporate bond markets.

A Distribution of Firms

Industry Name	No. of Firms	Country	No. of Firms
Hotels	1	France	13
Airlines	4	Netherlands	4
Chemicals	5	UK	11
Telecom	13	Germany	10
Food/Soft Drinks	2	Sweden	5
Retail-grocery Chains	6	Italy	2
Automotives	6	Greece	1
Entertainment	1	Spain	2
Aerospace/Defence	2	Finland	5
Machinery	1	Denmark	1
Paper	3	Norway	1
Utilities	4		
Printing/Publishing	3		
Media	2		
Steel	1		
Advertising	1		
Total	55		55

Table 1: **Distribution of Firms Across Industries and Countries** Firms are grouped into industries according to the Moody's industry classification.

B Actual and Risk-neutral Default Intensities: Estimation Results

	Actual default intensities			Risk-neutral default intensities				
	$\kappa\theta$	κ	σ	θ^Q	κ^Q	σ^Q	$\tilde{\theta}^Q$	$\tilde{\kappa}^Q$
mean	1.47	0.68	1.37	3.11	0.20	1.08	5.36	0.20
median	0.83	0.33	1.20	3.29	0.17	0.95	4.18	0.17
std. dev.	1.93	0.93	0.60	1.54	0.18	0.62	3.90	0.18

Table 2: **Estimation of the actual and risk-neutral default intensities** Summary statistics for the firm-by-firm parameter estimates describing the dynamics of the actual and risk-neutral default intensities in equations (28) - (30).

C Extracting the CMF Factor: Estimation Results

α	β_{EMKT}	β_{CMKT}	β_{TERM}	$Perc(C)$	$Perc(S)$
0.0001	−0.0008	−0.0005	0.2345	62.75	20.78
1.0960	0.0117	0.00703	4.1520		
−0.0007	0.0094	0.0041	1.4754	15.93	8.32
3.9568	0.8896	0.1571	3.5443		
−0.0012	0.0114	0.0078	2.4066	19.33	8.61
4.4765	0.7916	0.1615	3.8588		
−0.0016	0.0120	0.0097	3.1751	23.51	8.75
4.7590	0.7286	0.1555	4.0645		
−0.0020	0.0124	0.0106	3.8440	27.68	8.81
4.9392	0.6837	0.1478	4.2017		
−0.0024	0.0128	0.0109	4.4399	31.50	8.83
5.0638	0.6489	0.1409	4.2980		

Table 3: **Estimates for the return model in (30)** This table reports the results of the panel regression of the excess returns of defaultable zero-coupon bonds on the excess market returns $EMKT$, the excess local market return $CMKT$, the spread between long and short Euro bonds $TERM$ and the dummies controlling for specific week between January 2003 and October 2006, 197 weeks. The left-hand side excess returns correspond to defaultable bonds with the following times to expiration: 1 week, 1 year, 2 years, 3 years, 4 years and 5 years. The first line in the table corresponds to the estimates of the returns model where the left-hand side returns correspond to corporate bonds with the shortest time to expiration. The CMF factor at time t corresponds to the slope coefficient of the dummy controlling for time t. $Perc(C)$ column reports the variance of the CMF factor as a percentage of the variance of the pricing error obtained under the returns model in (30) without time dummies, while $Perc(S)$ column reports the R^2 of the return models without time dummies. The t-statistics are reported in parentheses. The reported values for the estimates are averages across firms of the corresponding firm-specific estimates.

	γ_{ϕ}			γ_{CMF}		
Name	*Mean*	*Med*	*tStat*	*Mean*	*Med*	*tStat*
EDF	0.0001	0.0001	5.1091	0.1613	0.6424	7.3966
Size	−0.0001	−0.0002	−3.7908	0.2280	0.9964	9.1808
M/B	−0.0000	−0.0001	−2.4827	0.2477	1.1315	11.7675

Table 4: **Testing for firm characteristics:** This table reports the results of the regressions of excess returns of defaultable zero-coupon bonds on a time-varying characteristic, the excess market returns $EMKT$, the excess local market return $CMKT$, the spread between long and short Euro bonds $TERM$, and the CMF factor. Specifically, for each firm we estimate the following regressions: $r_{t+1}^{i,e} = \gamma_0^i + \gamma_{\phi}^i \phi_t + \gamma_{EMKT}^i r_{t+1}^{EMKT,e} + \sum_c 1_{\{c=C\}} \gamma_{CMKT}^i r_{t+1}^{CMKT,e} + \gamma_{TERM}^i r_{t+1}^{TERM,e} + \gamma_{CMF}^i CMF_{t+1} + \epsilon_{t+1}^i$. We consider three firm characterisitcs: the actual default probability measured by the 1-year EDF value, the firm size and the market-to-book ratio. The first column lists the name of the characteristic. The following columns report summary statistics (across firms) of the loadings on the characteristic and the default factor. The t-Statistics are the median t-Statistics across firms. The results for the loadings on the other systematic factors are not reported.

D Asset Pricing Tests: Results from the Time-series Regressions

α	β_{EMKT}	β_{TERM}	β_{CMF}	R^2	$E[R]$
−0.0002	−0.0595	0.5730	2.6583	0.2037	0.0006
(0.2011)	(6.2292)	(0.3225)	(1.5745)		
−0.0006	−0.0449	1.6727	2.9709	0.1008	0.0007
(0.5547)	(3.9228)	(0.7846)	(1.4665)		
−0.0010	−0.0182	2.4341	5.1873	0.0667	0.0009
(1.0018)	(1.6771)	(1.2074)	(2.7078)		

Table 5: **The Merrill Lynch non-financial corporate bond portfolios sorted on rating** This table reports the results of the time-series regressions of the excess realized returns of three Merrill Lynch non-financial corporate bond portfolios sorted on rating (AAA, A and BBB), on the excess market returns $EMKT$, the spread between long and short Euro bonds $TERM$ and the CMF factor between January 2003 and October 2006, 197 weeks. The CMF factor is extracted from returns on defaultable zero-coupon bonds maturing in one week, according to the model in equation (30). The first line corresponds to the higher rating portfolio. The t-statistics are reported in parentheses.

α	β_{EMKT}	β_{TERM}	β_{CMF}	R^2	$E[R]$
	AAA Portfolios Sorted on Maturity				
−0.0004	−0.0228	0.8468	1.3852	0.1849	0.0005
(1.1404)	(5.6032)	(1.1162)	(1.9214)		
−0.0005	−0.0561	1.1302	2.7471	0.1908	0.0005
(0.5571)	(5.9223	(0.6408)	(1.6391)		
−0.0003	−0.0759	0.9565	3.6126	0.1802	0.0007
(0.2325)	(5.7469)	(0.3892)	(1.5468)		
0.0004	−0.1068	0.1358	4.3918	0.1469	0.0010
(0.2207)	(5.1418)	(0.0352)	(1.1961)		
	AA Portfolios Sorted on Maturity				
−0.0005	−0.0405	1.2071	2.1288	0.1746	0.0005
(0.7774)	(5.54450)	(0.8872)	(1.6465)		
−0.0004	−0.0735	1.2273	3.7866	0.1739	0.0007
(0.3102)	(5.5830)	(0.5009)	(1.6264)		
−0.0000	−0.0852	0.7934	4.1630	0.1514	0.0009
(0.0053)	(5.1725)	(0.2590)	(1.4300)		

Table 6: **The AAA-rated and AA-rated Merrill Lynch corporate bond portfolios sorted on maturity** This table reports the results of the time-series regressions of the excess realized returns of four AAA-rated Merrill Lynch corporate bond portfolios sorted on maturity (1-3 years, 3-5 years, 5-7 years and 10+ years) and three AA-rated Merrill Lynch corporate bond portfolios sorted on maturity (1-5 years, 5-7 years, and 7-10 years), on the excess market returns $EMKT$, the spread between long and short Euro bonds $TERM$ and the CMF factor between January 2003 and October 2006, 197 weeks. The CMF factor is extracted from returns on defaultable zero-coupon bonds maturing in one week, according to the model in equation (30). The first line in each of the two panels corresponds to the lower maturity portfolio. The t-statistics are reported in parentheses.

α	β_{EMKT}	β_{TERM}	β_{CMF}	R^2	$E[R]$
	A Portfolios Sorted on Maturity				
−0.0006	−0.0333	1.3889	2.4965	0.1454	0.0006
(0.8808)	(4.7334)	(1.0610)	(2.0070)		
−0.0005	−0.0638	1.5622	4.2429	0.1452	0.0008
(0.3940)	(4.8715)	(0.6406)	(1.8309)		
−0.0003	−0.0683	1.5004	4.9212	0.1128	0.0010
(0.2072)	(4.1753)	(0.4928)	(1.7011)		
−0.0005	−0.0607	2.2813	5.3318	0.0536	0.0012
(0.2323)	(2.6691)	(0.5291)	(1.3259)		
	BBB Portfolios Sorted on Maturity				
−0.0008	−0.0107	2.0900	3.5679	0.0679	0.0008
(1.2825)	(1.4457)	(1.5199)	(2.7305)		
−0.0011	−0.0341	2.8712	5.8806	0.0718	0.0009
(0.8707)	(2.4106)	(1.0922)	(2.3542)		
−0.0010	−0.0236	3.0896	8.0863	0.0595	0.0012
(0.6922)	(1.3970)	(0.9845)	(2.7117)		

Table 7: **The A-rated and BBB-rated Merrill Lynch corporate bond portfolios sorted on maturity** This table reports the results of the time-series regressions of the excess realized returns of four A-rated Merrill Lynch corporate bond portfolios sorted on maturity (1-5 years, 5-7 years, 7-10 years and 10+ years) and three BBB-rated Merrill Lynch corporate bond portfolios sorted on maturity (1-5 years, 5-7 years, and 7-10 years), on the excess market returns $EMKT$, the spread between long and short Euro bonds $TERM$ and the CMF factor between January 2003 and October 2006, 197 weeks. The CMF factor is extracted from returns on defaultable zero-coupon bonds maturing in one week, according to the model in equation (30). The first line in each of the two panels corresponds to the lower maturity portfolio. The t-statistics are reported in parentheses.

α	β_{EMKT}	β_{TERM}	β_{CMF}	R^2	$E[R]$
-0.0006	-0.0145	1.4045	2.2386	0.1046	0.0006
(1.3937)	(3.0539)	(1.5892)	(2.6656)		
-0.0008	-0.0328	2.1475	3.3481	0.1017	0.0007
(1.0168)	(3.5472)	(1.2486)	(2.0486)		
-0.0008	-0.0493	2.1074	4.7712	0.1023	0.0008
(0.6534)	(3.6939)	(0.8491)	(2.0229)		
-0.0006	-0.0597	2.7593	7.6113	0.0551	0.0015
(0.2597)	(2.4429)	(0.6064)	(1.7602)		

Table 8: **The Merrill Lynch non-financial corporate bond portfolios sorted on maturity** This table reports the results of the time-series regressions of the excess realized returns of four Merrill Lynch non-financial corporate bond portfolios sorted on maturity (1-3 years, 3-5 years, 5-7 years and 10+ years), on the excess market returns $EMKT$, the spread between long and short Euro bonds $TERM$ and the CMF factor between January 2003 and October 2006, 197 weeks. The CMF factor is extracted from returns on defaultable zero-coupon bonds maturing in one week, according to the model in equation (30). The first line corresponds to the lower maturity portfolio. The t-statistics are reported in parentheses.

α	β_{EMKT}	β_{TERM}	β_{CMF}	R^2	$E[R]$
-0.0013	-0.0550	0.9363	3.1294	0.1763	-0.0003
(1.4395)	(5.5805)	(0.5109)	(1.7968)		
-0.0010	-0.0605	0.4328	3.3011	0.1445	-0.0003
(0.9280)	(4.9913)	(0.1919)	(1.5405)		
-0.0010	-0.0497	0.7245	2.1791	0.0905	-0.0002
(0.9014)	(3.8842)	(0.3045)	(0.9638)		
-0.0017	-0.0371	1.7892	4.5144	0.0751	-0.0003
(1.5304)	(2.9211)	(0.7578)	(2.0122)		
-0.0022	0.0554	4.3796	12.5528	0.0693	0.0011
(0.9110)	(2.0726)	(0.8804)	(2.6555)		

Table 9: **The Lehman Brothers Euro-aggregate industrial corporate bond portfolios sorted on rating** This table reports the results of the time-series regressions of the excess realized returns of five Lehman Brothers Euro-aggregate industrial corporate bond portfolios sorted on rating (AAA, AA, A, BAA and High Yield), on the excess market returns $EMKT$, the spread between long and short Euro bonds $TERM$ and the CMF factor between January 2003 and October 2006, 197 weeks. The CMF factor is extracted from returns on defaultable zero-coupon bonds maturing in one week, according to the model in equation (30). The first line corresponds to the higher rating portfolio. The t-statistics are reported in parentheses.

α	β_{EMKT}	β_{TERM}	β_{CMF}	R^2	$E[R]$
−0.0013	−0.0188	1.3124	2.1769	0.1190	−0.0001
(−2.8902)	(−3.7374)	(1.3987)	(2.4415)		
−0.0013	−0.0386	1.2124	2.4885	0.0992	−0.0002
(−1.4862)	(−3.9172)	(0.6616)	(1.4290)		
−0.0018	−0.0693	2.4489	4.0009	0.1466	−0.0002
(−1.4634)	(−4.9331)	(0.9368)	(1.6107)		
−0.0013	−0.0763	0.5133	4.9687	0.1137	−0.0005
(−0.8228)	(−4.2719)	(0.1544)	(1.5727)		
−0.0019	−0.1013	3.6662	12.6694	0.0939	0.0007
(−0.6869)	(−3.3023)	(0.6424)	(2.3361)		

Table 10: **The Lehman Brothers Euro-aggregate corporate bond portfolios sorted on maturity** This table reports the results of the time-series regressions of the excess realized returns of five Lehman Brothers Euro-aggregate corporate bond portfolios sorted on maturity (1-3 years, 3-5 years, 5-7 years, 7-10 years and 10+ years), on the excess market returns $EMKT$, the spread between long and short Euro bonds $TERM$ and the CMF factor between January 2003 and October 2006, 197 weeks. The CMF factor is extracted from returns on defaultable zero-coupon bonds maturing in one week, according to the model in equation (30). The first line corresponds to the lower maturity portfolio. The t-statistics are reported in parentheses.

α	β_{EMKT}	β_{TERM}	β_{CMF}	R^2	$E[R]$
	Cross-sectional Averages of the Estimates				
-0.0012	-0.0455	1.2214	0.2250	0.1005	-0.0002
	Cross-sectional Standard Deviations of the Estimates				
0.0003	0.0147	0.8873	0.2358	0.0645	0.0001
	Cross-sectional Averages of the t-Statistics				
(1.0525)	(3.6354)	(0.4937)	(1.6493)		
	Cross-sectional Standard Deviations of the t-Statistics				
(0.3135)	(1.1553)	(0.3352)	(1.8141)		

Table 11: **The Lehman Brothers Euro-aggregate corporate bond portfolios sorted on sector** This table reports the results of the time-series regressions of the excess realized returns of sixteen Lehman Brothers Euro-aggregate corporate bond portfolios sorted on sector (Aero/Defense, Banking, Building Materials, Capital Goods, Chemicals, Communications, Consumer Non-cyclical, Consumer Cyclical, Diversified Manufacturing, Food and Beverages, Lodging, Refining, Telephone, Tobacco, Wireless and Media Non-Cable), on the excess market returns $EMKT$, the spread between long and short Euro bonds $TERM$ and the CMF factor between January 2003 and October 2006, 197 weeks. The CMF factor is extracted from returns on defaultable zero-coupon bonds maturing in one week, according to the model in equation (30). Each panel reports an average statistic across portfolios.

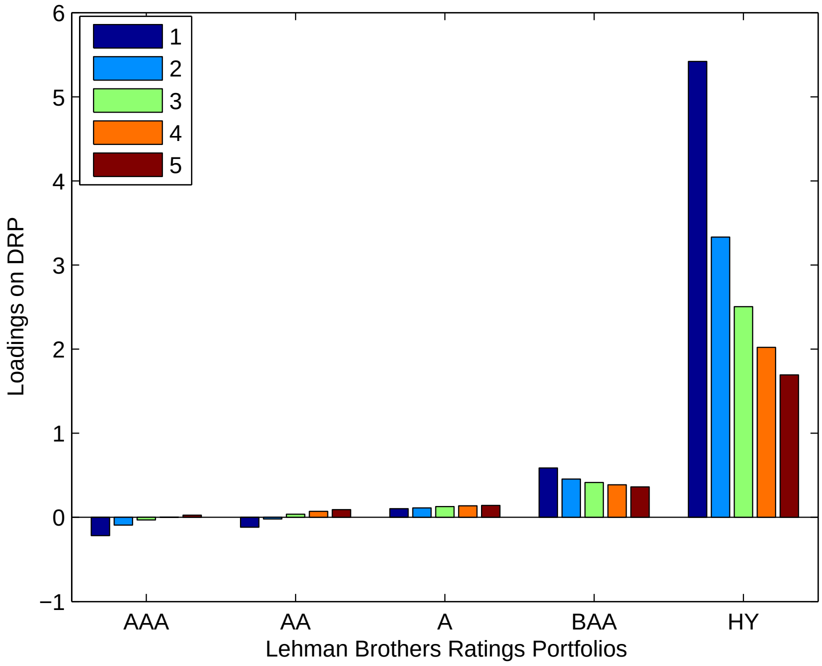

Figure 5: The estimates of the slope coefficient on the CMF factor extracted from returns on defaultable zero-coupon bonds with maturity varying from 1 to 5 years. These slopes are estimated from the time-series regressions of the excess realized returns of five Lehman Brothers Euro-aggregate industrial corporate bond portfolios sorted on rating (AAA, AA, A, BAA and High Yield), on the excess market returns $EMKT$, the spread between long and short Euro bonds $TERM$ and the CMF factor between January 2003 and October 2006, 197 weeks.

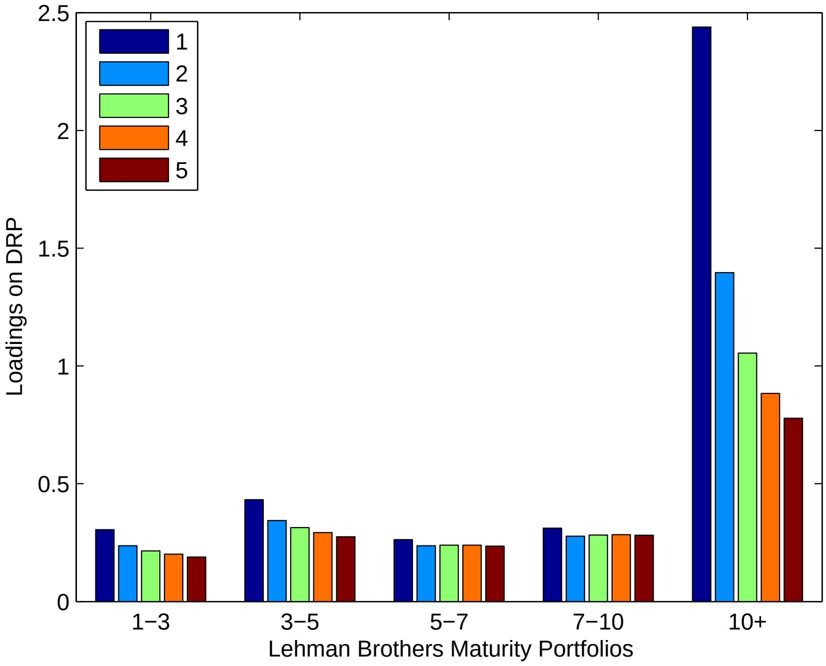

Figure 6: The estimates of the slope coefficient on the CMF factor extracted from returns on defaultable zero-coupon bonds with maturity varying from 1 to 5 years. These slope coefficients are estimated from the time-series regressions of the excess realized returns of five Lehman Brothers Euro-aggregate corporate bond portfolios sorted on maturity (1-3 years, 3-5 years, 5-7 years, 7-10 years and 10+ years), on the excess market returns $EMKT$, the spread between long and short Euro bonds $TERM$ and the CMF factor between January 2003 and October 2006, 197 weeks.

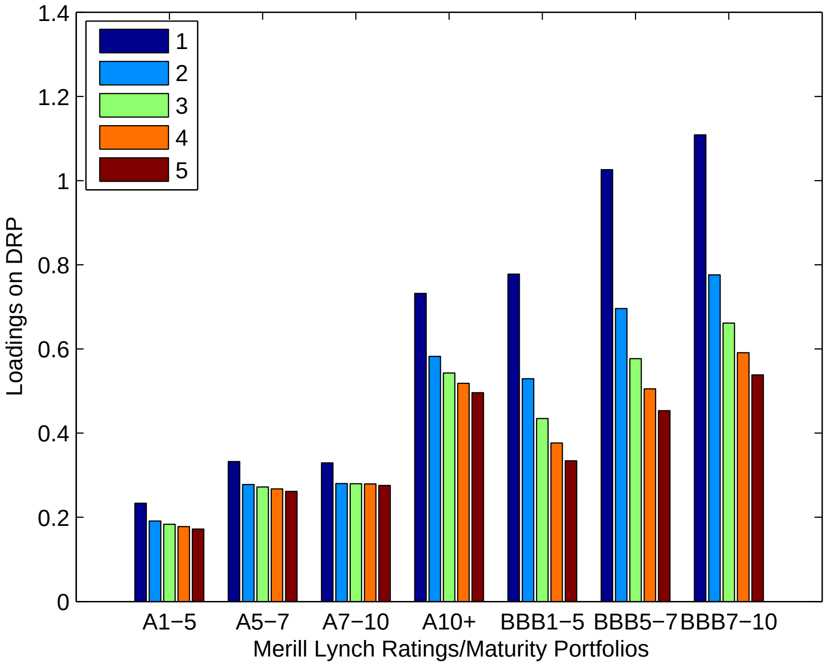

Figure 7: The estimates of the slope coefficient on the CMF Factor extracted from returns on defaultable zero-coupon bonds with maturity varying from 1 to 5 years. These slope coefficients are estimated from the time-series regressions of the excess realized returns of four A-rated Merrill Lynch corporate bond portfolios sorted on maturity (1-5 years, 5-7 years, 7-10 years and 10+ years) and three BBB-rated Merrill Lynch corporate bond portfolios sorted on maturity (1-5 years, 5-7 years, and 7-10 years), on the excess market returns $EMKT$, the spread between long and short Euro bonds $TERM$ and the CMF factor between January 2003 and October 2006, 197 weeks.

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