Decomposing the Returns on European Debt[1]

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Abstract

Common variation in the prices of European corporate debt may not always be associated with a rational response to an increase in the relative importance of a macroeconomic risk factor. Building on Campbell's ICAPM framework, we show that risk premia of assets with nonlognormal return distributions represent compensation not only for exposure to macroeconomic factors but also for unexpected revisions to these assets' return distributions, such as sudden increases in the likelihood of extreme events. If such revisions happen across assets almost simultaneously, perhaps as a systemic response to a large credit event, they can induce covariation in risk premia unrelated to the time variation of the priced macroeconomic factors. Our study presents evidence from the European debt markets which supports this theory. The asset pricing tests also document patterns consistent with the "flight to quality" effect for European corporate bonds.

JEL Classifications: G12, G13, G15

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1. Introduction

This paper investigates the sources of common time-series variation in the prices of European corporate debt. Standard asset pricing models state that to the extent that CDS spreads represent compensation for exposure to systematic risk factors, common variation in default swap spreads arises mainly because the price of risk of a systematic factor increases relative to the price of risk of other systematic factors. In other words, as long as systematic risk factors are the only source of common variation in credit spreads, equity and credit markets should be well integrated. This, however, does not seem to be the case at all times. One episode of weak integration between equity and debt markets dates back to 2002, when the European high-yield market saw one quarter of its bonds default. On a par basis, nearly 90% of that year's defaults came from the telecommunications and cable sectors. Over the first quarter of 2002, European credit spreads computed as the difference between the redemption yield on the iBoxx EURO corporate bond index and the 10-year Euro-vs.-Euribor rate increased by almost 50%, whereas the Morgan Stanley's MSCI EURO stock market index remained flat for the same time period. About two years later, during the three-month period from June until August 2004, European credit spreads rose by more than 60% while European stocks climbed rather steadily, rising by more than 10%. Although one could argue that certain macroeconomic factors are the source for the common variation in credit spreads, it is at best doubtful that these would suddenly become important for the credit market but not for the equity market.

In this study, we propose an alternative explanation for the common variation in the returns of corporate bonds which is not necessarily related to systematic risk factors. We argue that for assets with nonlognormal return distributions, such as defaultable bonds, common variation in risk premia unrelated to systematic factors can arise when investors make systemic revisions to the distribution of extreme events that affect these assets' payoffs. For instance, the dramatic increase in European credit spreads during the first few months of 2002 could have resulted, at least partially, from investors realizing almost simultaneously that the payoff structure of their credit-contingent claims was more downward skewed than previously anticipated. As a result, investors in European credit markets may have repriced their credit-sensitive securities based on a revision of the cash flow distribution rather than on a revision of the discount rates.

Our argument is rooted in a theoretical result derived in the context of Campbell's (1993) intertemporal capital asset pricing model (ICAPM). Within this framework, we show that the risk premia of defaultable bonds represent compensation not only for exposure to systematic risk, but also for unexpected revisions to these bonds' non-lognormal return distributions. Such unexpected revisions can arise in the aftermath of an adverse credit-market event, such as a corporate default, when investors have the opportunity to gather more information about the distribution of extreme events that may lead to corporate default. For instance, an unexpected corporate default could present investors with the opportunity to learn that certain extreme events are more likely than previously anticipated and that the payoff distribution of the defaultable bonds they hold is more downward skewed than they previously thought.[4] If such revisions in the distribution of default events happen nearly simultaneously across the entire universe of firms, the returns of all defaultable bonds will move in the same direction. More importantly, this common movement in defaultable bond returns across firms cannot be attributed to macroeconomic risk.

The advantage of building on Campbell's ICAPM framework is that it allows us to identify likely sources of macroeconomic risk that are priced by security markets. In our context, these sources turn out to be the market factor for Europe and the long–short spread on European benchmark government bonds. At the same time, it is important to point out that Campbell's framework cannot be applied directly to defaultable bonds. The original framework was developed for securities with lognormal returns and defaultable bonds are quite unlikely to fall into this category. Indeed, Berndt (2007) provides empirical evidence that instantaneous returns on zero-coupon defaultable bonds with zero recovery are more likely to follow conditionally lognormal dynamics as opposed to conditionally normal dynamics.

The theoretical contribution of our paper is the decomposition of the common variation in the returns of defaultable bonds into a portion that is due to exposure to macroeconomic risk, proxied by the market and the long–short government term spread, plus a portion that stems from systemic revisions in the distribution of the default event across all firms.

For the remainder of our study, we investigate whether these theoretical implications have support in European corporate bond markets. Our empirical approach consists of three steps: First, we identify a latent factor which has the potential of capturing the comovement in defaultable bond returns that is due to systemic revi-

[4]The notion of investors revising the payoff distribution for their holdings in corporate debt is similar in spirit to the credit contagion model by Schönbucher (2003), who introduces the notion of frailty models to the credit-risk literature, and by Collin-Dufresne et al. (2003), who model default contagion via the updating of investors' beliefs.

sions of the distributions of the default event across firms. Second, we test whether this latent factor induces common variation in bond returns above and beyond the two macroeconomic risk factors suggested by the theory. Third, we test whether this latent factor does actually capture the systemic revision of the distributions of the default event across firms.

We now describe each of these steps in detail. Regarding the first step, to the extent that investors do revise the distributions of the default events across firms in the aftermath of adverse credit-market events, we should see price comovement not only in defaultable bonds, but also in other credit-contingent securities such as credit default swaps (CDS), over-the-counter securities that provide default insurance on debt. Based on this idea, we use CDS data to identify a latent factor which has the potential to capture the comovement in CDS spreads induced by the systemic revision of the distributions of the default event across firms. For this procedure we only use the CDS spreads of the European non-financial firms with the most liquid CDS market from January 2003 to October 2006. To extract the latent factor we first use the pricing information in the CDS market quotes to construct excess returns on zero-coupon zero-recovery defaultable bonds. Then, guided by our model, we extract our fact as the common component of the residuals of the orthogonal projections of these excess returns on the market and the long-short term spread on European benchmark government bonds. We call this latent factor the *credit market factor*, or short *CMF*. Crucial for this step is the extraction of the pricing information from the available CDS spreads. This is carried out by estimating a time-series model for risk-neutral default probabilities using credit default swap data.

This step in our empirical approach reveals that CMF captures on average between 30% and 54% of the risk-adjusted excess returns of the zero-coupon zero-recovery defaultable bonds constructed from the CDS data. This suggests that a potentially important fraction of the CDS spread captures compensation for unexpected revisions to the distribution of the default event.

The second step of our empirical analysis, we first test whether the CMF has any impact on corporate bonds. As mentioned, a revision of the default event's distribution should impact the pricing of all credit-contingent assets. This implies that whenever we observe comovements in default swap rates across the CDS market, even if they cannot be accounted for by changes in macroeconomic risk factors, we should observe a similar pattern not just in CDS-inferred prices of zero-coupon zero-recovery bonds, but the corporate bond market as a whole. Given the limited availability of firm-level pricing data for European credit and corporate bond markets, this is not a tautology, despite the fact that owners of corporate debt use CDS contracts as a natural hedge against default risk. With 55 European firms serving as the reference entities for the CDS data in our sample, the price data for these corporate bond issuers is bound to be variable and it is unlikely that the hedging effect alone will account for the comovements in the prices of these two credit-contingent asset classes. In that sense, it is important to verify that a wide range of European corporate bond returns responds contemporaneously to the CMF that we extract from CDS rates for the limited set of liquidly traded reference entities in the European credit market. In an effort to cover the whole cross section of European corporate bond returns, we consider a large variety of corporate bond portfolios sorted on one of three characteristics: credit rating, time-to-maturity and sector. The corporate bond portfolios used in this study are preconstructed by either Merrill Lynch or Lehman Brothers and span the entire universe of European corporate bonds.

If investors revise a default event's distribution, the returns on risky bonds such as those with low credit ratings or long times to maturity should reflect the impact of these revisions to a larger extent than returns on relatively safe investments in corporate debt. To illustrate this point, let us consider a simple stylized scenario where an investor holds two zero-coupon zero-recovery defaultable bonds A and B with the same time to maturity of one year. Assume that as of today, bond A pays one unit of account in states s_1 and s_2 and zero in state s_3. Bond B, on the other hand, pays one unit of account for sure. Suppose that one week from today, investors realize that state s_1 is also a default state in which neither bond will pay anything. Suppose further that each of these three states is equally likely and that discount rates remain the same throughout the year. If r denotes the weekly discount rate, the price of bond A prior to the revision is $P_0^A = \frac{2}{3}e^{-50r}$, while the price of bond B is given by $P_0^B = e^{-50r}$. After revising the distribution of the default event, these prices change to $P_1^A = \frac{1}{3}e^{-49r}$ and $P_1^B = \frac{2}{3}e^{-49r}$, respectively. The net holding returns on these bonds can be computed as $r_1^A = \log \frac{P_1^A}{P_0^A} = r \log \frac{1}{2}$ and $r_1^B = r \log \frac{2}{3}$. If there were no revisions, clearly the net return on both bonds should be r. A revision, however, is bad news for both bonds, leading to negative net returns. Thus, consistent with our intuition, the revision induces the net returns of these two bonds to move in the same direction. More importantly, the riskier bond A records a larger loss, $r - r_1^A$, than bond B since $r_1^A < r_1^B$.

Our asset pricing test results strongly support the two hypotheses: (i) Corporate bond portfolios respond contemporaneously to innovations in the CMF. (ii) This response is commensurate with the riskiness of the portfolio, as measured a priori by a set of risk characteristics. Without exception, all corporate bond portfolios load positively on the CMF, and their loadings increase with the riskiness of the portfolio.

The last step of our empirical analysis answers the question of whether the CMF does capture the systemic revision of the default event's distributions across firms. To the extent that the returns on equity portfolios are more likely to be lognormally distributed, we test whether equity portfolios react to innovations in the CMF in any way. If our story is correct, we should not see any response in the equity portfolios. Indeed, the data indicate that equity portfolios respond only weakly, and sometimes even in the wrong direction, to innovations in the CMF. In summary, our results support the theory that the CMF captures a price behavior that originates from investors' systemic revisions of the distribution of the default event, across firms.

In addition to the aforementioned results, we also document another interesting pattern in the returns on defaultable debt. Most of the European 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. Similarly, as the economy goes through an expansion period, investors' appetite for risk increases and they invest in riskier high-yield bonds.

This study contributes to the growing literature concerned with the measurement of the default risk premia that includes Elton et al. (2001), Collin-Dufresne et al. (2001), Blanco et al. (2004), Longstaff et al. (2004), Driessen (2005), Amato and Remolona (2005), Berndt et al. (2005), Saita (2006) and Berndt et al. (2007). With the exception of Denzler et al. (2005), these studies all focus solely on the U.S. corporate bond market. In addition to the fact that we study European capital markets, the contribution of our work to the existing literature stems from extending the ICAPM to accommodate nonlognormal returns of defaultable securities. We compare the model's results to the data, construct a CMF and find strong empirical support for the model's theoretical implication: the CMF captures the price behavior due to investors' systemic revisions to the distribution of default events in the aftermath of adverse credit-market news.

The remainder of the paper is structured as follows. Section 2 describes in detail our source for European CDS data and pricing data on European corporate bond and equity portfolios used in this study. Section 3 describes how to compute CDS-inferred prices on zero-coupon zero-recovery defaultable bonds of various maturities and estimates a time-series model of risk-neutral default probabilities using the information embedded in the CDS spreads. Section 4 presents the theoretical determinants of default risk premia and constructs an expected-returns beta representation for the return on defaultable assets. In Section 5, we identify the common variation in the nonlognormal component of returns on zero-coupon zero-recovery defaultable bonds, and Section 6 implements a two-step methodology to disentangle the different sources of common variation in excess returns of zero-coupon zero-recovery bonds and constructs our CMF. In Section 7, we investigate the nature of the CMF in more detail and show that it captures compensation for unexpected revisions to the nonlognormal component of the return distribution of credit-sensitive assets. Finally, Section 8 implements a number of robustness checks, summarizes the results of our paper and concludes.

2. Data

This section discusses our data sources for European CDS rates, systematic factors and test assets.

2.1. Credit Default Swaps

Credit default swaps are single-name over-the-counter credit derivatives that provide default insurance. The payoff to the buyer of protection covers losses up to the notional value in the event of a default by the 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.[5] 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

[5]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.

spreads are less confounded by illiquidity, taxes and various market microstructure effects that are known to have a marked effect on corporate bond yield spreads.[6]

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 five-year CDS markets 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 nonfinancial five-year CDS contracts. To optimally mitigate the tradeoff between the microstructure effects of high-frequency quotes and 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 11 European countries and 16 different industries, based on Moody's industry classification. Detailed summary statistics are provided in Table I. A typical firm in our sample has 150 valid weekly CDS observations out of 196 maximum possible weekly quotes. The minimum number of weekly observations is 95. Figure 1 plots the distribution of the credit quality of the firms in our sample, showing a concentration of medium credit quality.

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. Despite its small size, we believe our sample is quite diverse because the firms in our sample are distributed across 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 this information can be extracted more precisely[7] from the quotes on the CDS contracts of those firms with very liquid five-year CDS markets. To this extent, we are confident that

[6]Recent papers that analyze the contribution of noncredit factors to bond yields include Zhou (2005), Longstaff et al. (2004) and Ericsson and Renault (2001).

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

the results in the paper are not biased by the size of our sample.

2.2. Interest Rates, Systematic Factors and Test Assets

Throughout our empirical analysis, we rely on information about the Euro term structure of riskless bonds. This data is obtained from the Datastream Euro zero curves constructed relative to Euribor.[8] All the excess returns and the zero-cost portfolios are computed relative to the one-month zero yield.

For the purposes of Sections 6 and 7, we need to compute zero-cost portfolios that are long the market portfolio and short the one-month zero yield or long the 30-year zero yield and short the one-month zero yield. For the latter 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 previous studies on capital markets integration, whenever possible we use portfolios constructed from the data disseminated in the electronic version of the 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.[9]

Finally, for the purposes of Section 7, we need to compute realized returns on a range of test assets in excess of the one-month zero yield. We consider the following test assets: the Merrill Lynch nonfinancial corporate bond portfolios sorted on rating and time to maturity (two separate sorts), 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 rating, maturity and sector. The time-series data for all these portfolios comes from Datastream.[10]

[8]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 one year and four months is EM01Y04.

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

[10]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,

3. Prices of Defaultable Zero-Coupon Bonds

This section describes how to compute CDS-inferred prices on zero-coupon zero-recovery defaultable bonds of various maturities. Let us take as given a probability space $(\Omega, \mathcal{F}, P)$ and information filtration $\{\mathcal{F}_t : t \geq 0\}$. In the absence of arbitrage and market frictions, there exists a stochastic discount factor, denoted by M (see, for instance, Duffie, 2001). Moreover, under mild technical conditions, Harrison and Kreps (1979) and Delbaen and Schachermayer (1999) show that there exists a "risk-neutral" probability measure Q associated with M. Note that in our setting, markets are not necessarily complete, so the stochastic discount factor and the associated risk-neutral measure may 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 $E_t\left[M_{t,T}\, Z\right] = E_t^Q\left[e^{-\int_t^T r_s\,\mathrm{d}s}\, Z\right]$, where r is the short-term interest rate and E_t^Q denotes the expectation operator with respect to Q, conditional on the information available up to and including time t. In particular, the time-t market value of a defaultable zero-coupon bond that pays one unit of account in the event that a currently surviving firm does not default before time T and zero otherwise is given by

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

where τ denotes the default time of the firm. To compute the prices in Equation (1), we rely on the reduced-form arbitrage-free approach to pricing defaultable bonds where the risk-neutral distribution of the default time τ is fully described by a risk-neutral default intensity process λ^Q. If we assume the doubly stochastic property under Q, Equation (1) reduces to

$$P(t, T-t) = E_t^Q\left[e^{-\int_t^T r_s + \lambda_s^Q\,\mathrm{d}s}\right], \tag{2}$$

subject to sufficient conditions given in Duffie (2001).

Motivated by Berndt et al. (2005), we suppose that the risk-neutral default intensity for firm i, $\lambda^{Q,i}$, statisfies

$$\mathrm{d}\log\lambda_t^{Q,i} \;\; = \;\; \kappa^Q\left(\theta^{Q,i} - \log\left(\lambda_t^{Q,i}\right)\right)\mathrm{d}t + \sigma^Q\sqrt{\rho^Q}\,\mathrm{d}B_t^c + \sigma^Q\sqrt{1-\rho^Q}\,\mathrm{d}B_t^i, \tag{3}$$

where B_t^c and B_t^i are independent standard Brownian motions with regard to the physical measure P, and $\kappa^Q, \theta^{Q,i}, \sigma^Q$ and ρ^Q are scalars to be estimated. The risk-

LHALODE, LHAREFE, LHATLPE, LHATBCE, LHAWRSE, and LHAMNCE.

neutral distribution of λ^Q is specified by assuming that

$$\mathrm{d}\log\lambda_t^{Q,i} = \tilde{\kappa}^Q\left(\tilde{\theta}^{Q,i} - \log\left(\lambda_t^{Q,i}\right)\right)\mathrm{d}t + \sigma^Q\sqrt{\rho^Q}\,\mathrm{d}B_t^{Q,c} + \sigma^Q\sqrt{1-\rho^Q}\,\mathrm{d}B_t^{Q,i},$$

where $\tilde{\kappa}^Q$ and $\tilde{\theta}^{Q,i}$ are scalars to be estimated. B_t^c and B_t^i are independent standard Brownian motions with regard to Q. Given a set of parameters $\left(\{\tilde{\theta}^{Q,i}\}, \tilde{\kappa}^Q, \sigma^Q\right)$, we can compute model-implied values for $\lambda^{Q,i}$ using data on five-year CDS rates and an assumed risk-neutral loss given default of $L^Q = 0.6$. For details, we refer the reader to Section 5.1 in Berndt et al. (2005). To improve the interpretability and the reliability of our parameter estimates, we impose the overidentifying restriction that $\theta^{Q,i}$ equals the sample mean of $\log\lambda_t^{Q,i}$ and that the ratio of $\tilde{\theta}^{Q,i}$ to the sample mean of $\log(CDS_t^i/L^Q)$ is constant within a given country. Using country-by-country maximum likelihood estimation (MLE), we obtain estimates for the parameters that govern the processes for λ^Q. The estimated values of these parameters are presented in Table II.

4. An Asset Pricing Model for Assets with Non-lognormal Returns

In this section we use Campbell's (1993) discrete ICAPM 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 nonexpected-utility preferences:

$$U_t = \left\{(1-\beta)C_t^{\frac{1-\gamma}{\theta}} + \beta\left(E_t\left[U_{t+1}^{1-\gamma}\right]\right)^{\frac{1}{\theta}}\right\}^{\frac{\theta}{1-\gamma}}, \tag{4}$$

where γ is the coefficient of relative risk aversion, $\theta = \sigma\frac{1-\gamma}{\sigma-1}$ and σ is the elasticity of intertemporal substitution.[11] As Epstein and Zin (1989, 1991) show, the first-order condition of the representative agent in this economy can be stated as

$$1 = E_t\left[\left\{\beta\left(\frac{C_{t+1}}{C_t}\right)^{-\frac{1}{\sigma}}\right\}^{\theta}\left\{\frac{1}{R_{t+1}^m}\right\}^{1-\theta} R_{t+1}^i\right], \tag{5}$$

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

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 homoscedastic and lognormally distributed, the aggregate budget constraint can be exploited to substitute out consumption and to simplify the Euler equation to

$$E_t r^i_{t+1} - r^f_{t+1} = -\frac{1}{2}V_{ii} + \gamma V_{im} + (\gamma - 1)V_{ih}, \tag{6}$$

where r^* $(* = i,\ f)$ denotes log returns, $V_{ii} = \text{cov}_t\left(r^i_{t+1}, r^i_{t+1}\right)$, $V_{im} = \text{cov}_t\left(r^i_{t+1}, r^m_{t+1}\right)$ and $V_{ih} = \text{cov}_t\left(r^i_{t+1}, (E_{t+1} - E_t)\sum_{j=1}^{\infty}\rho^j r^m_{t+1+j}\right)$. 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.[12]

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

$$E_t r^i_{t+1} - r^f_{t+1} = -\frac{1}{2}V_{ii} + \gamma V_{im} + (1 - \gamma)V_{ib}, \tag{7}$$

where $V_{ib} = \text{cov}_t\left(r^i_{t+1}, r^b_{t+1}\right)$.

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

$$E_t r^i_{t+1} - r^f_{t+1} = -\frac{1}{2}V_{ii} + \left[\gamma + \beta^{b,m}_t(1 - \gamma)\right] V_{im} + (1 - \gamma)V^{\perp}_{ib}, \tag{8}$$

where $V^{\perp}_{ib} = \text{cov}_t\left(r^i_{t+1}, r^{b,\perp}_{t+1}\right)$. If we further assume that $r^{b,\perp}_{t+1}$ and the consumption growth are both jointly conditionally homoscedastic and lognormally distributed, we can apply the above relation to both r^m_{t+1} and $r^{b,\perp}_{t+1}$. Using the unconditional versions of these relations we obtain

$$\begin{aligned} \left[\gamma + \bar{\beta}^{b,m}(1 - \gamma)\right] &= \frac{E r^{m,e}_t}{V_{mm}} - \frac{1}{2} \\ 1 - \gamma &= \frac{E r^{b,\perp,e}_t}{V^{\perp}_{bb}} - \frac{1}{2}, \end{aligned} \tag{9}$$

where E denotes the unconditional expectation operator, $\bar{\beta}^{b,m} = E\beta^{b,m}_t$, $r^{m,e}_t = r^m_t - r^f_t$

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

and $r_t^{b,\perp,e} = r_t^{b,\perp} - r_t^f$. Substituting these formulas back into Equation (8) and taking expectations yields the following expected-returns beta representation.

$$Er_t^{i,e} + \frac{1}{2}V_{ii} = \beta_{im}\left[Er_t^{m,e} + \frac{1}{2}V_{mm}\right] + \beta_{ib}^{\perp}\left[Er_t^{b,\perp,e} + \frac{1}{2}V_{bb}^{\perp}\right], \tag{10}$$

where $\beta_{im} = V_{im}/V_{mm}$ and $\beta_{ib}^{\perp} = V_{ib}^{\perp}/V_{bb}^{\perp}$.[13]

The expected-returns beta representation in Equation (10) suggests that the time variation in returns is due mainly to two factors: time variation in the returns on the market portfolio in excess of the riskless short rate and time variation in the returns of a portfolio that longs a riskless console bond and shorts the riskless short rate. Close relatives of this latter 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 (10) applies to any returns that are both jointly homoscedastic and conditionally lognormally distributed with the consumption growth and the market return. However, returns on certain assets are less likely to satisfy the latter condition. For instance, Berndt et al. (2005) document that the instantaneous excess returns on defaultable zero-coupon bonds are more likely to be lognormally 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-returns beta representation model in Equation (10) can be slightly generalized to accommodate instantaneous excess returns that are not necessarily conditionally normally distributed. We describe this modified model below.

Suppose the returns on a defaultable bond r_t^D can be decomposed into two components. The first component, $r_t^{D,c}$, is jointly homoscedastic and lognormally distributed with the consumption growth and the market return. The second component, $r_t^{D,n}$, is orthogonal on the information contained on both the consumption growth and the market.[14] This latter component is going to capture the impact of the departure

[13] Notice that β_{im} and β_{ib} are in fact the conditional betas, which happen to be constant under the homoscedasticity assumption. Thus, they can be different from the unconditional betas.

[14] Here is one way to implement such a decomposition. Let $\mu = Er_t^D$ and $k_r = \operatorname{cov}(r_{t+1}, r_t)/\operatorname{cov}(r_t, r_t)$. 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,

from the conditional lognormality assumption on prices. Under these assumptions it can easily be shown that the expected-returns beta representation in Equation (10) becomes

$$Er_t^{D,c,e} + \frac{1}{2}V_{DD}^c = \beta_{Dm}^c \left[Er_t^{m,e} + \frac{1}{2}V_{mm} \right] + \beta_{Db}^{c\perp} \left[Er_t^{b,\perp,e} + \frac{1}{2}V_{bb}^{\perp} \right] + Ez_t, \tag{11}$$

where $r_t^{D,c,e} = r_t^{D,c} - r_t^f$, $V_{DD}^c = \mathrm{var}_t\left(r_{t+1}^{D,c}\right)$, $\beta_{Dm}^c = \mathrm{cov}_t\left(r_{t+1}^{D,c}, r_{t+1}^m\right)/V_{mm}$, $\beta_{Dm}^{c\perp} = \mathrm{cov}_t\left(r_{t+1}^{D,c}, r_{t+1}^{b,\perp}\right)/V_{bb}^{\perp}$, 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,[15] we can rewrite the above equation as

$$Er_t^{D,e} + \frac{1}{2}V_{DD}^c = \beta_{Dm} \left[Er_t^{m,e} + \frac{1}{2}V_{mm} \right] + \beta_{Db}^{\perp} \left[Er_t^{b,\perp,e} + \frac{1}{2}V_{bb}^{\perp} \right] + E\Delta z_t, \tag{12}$$

where $\Delta z_t = E_t r_{t+1}^{D,n} - \log E_t e^{r_{t+1}^{D,n}}$. The returns model associated with this representation can be summarized as

$$r_t^{D,e} = \alpha + \beta_{Dm} r_t^{m,e} + \beta_{Db}^{\perp} r_t^{b,\perp,e} + \Delta z_t^D + \epsilon_t^D. \tag{13}$$

where ϵ_t^D is an iid normally distributed error term with mean zero.

Thus, as do conditionally lognormal 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 lognormal returns, the returns of defaultable bonds are also moved by changes in the shape of the conditional distribution of r^D, via Δz_t. This latter source of time variation could host both a common component as well as an undiversifiable firm-specific component. More importantly, these two components affect the level of expected returns directly rather than through covariances. This follows from the fact that $r_t^{D,n}$ is orthogonal on the stochastic discount factor M. An important consequence of this result is the fact that to the extent that Δz_t hosts a common component, this component cannot be related to macroeconomic risk (i.e., the stochastic discount factor M). Of course, this argument is viable as long as our model of returns is sufficiently well specified.

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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.

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

Admittedly Campbell and Vuolteenaho (2004) show that a close version of the model in Equation (10)—the "bad beta, good beta" model—does a very good job in capturing important time-series properties of U.S. stock returns—including size and value "anomalies"—and, as such, could be a well-specified model for them. However, there is still the possibility that the same model is not well specified for European stock returns.[16] If our model is misspecified, then the common component in Δz_t could be correlated with an omitted macroeconomic factor. We take up this empirical matter in a subsequent section. Before we do so, however, we first need to identify the common component in the firm-specific time-varying terms Δz_t. The next section deals precisely with this issue.

5. Identifying the Common Variation in the Non-lognormal Component of Returns

Our methodology for identifying the common component of the terms Δz_t^i, where i is an index for firms, is reminiscent of the fixed time effects in panel regressions. It consists of two steps: In the first step, we identify the residuals from the firm-specific regressions of the bond excess returns on the market and term-spread excess returns. That is

$$r_t^{i,e} = \alpha^i + \beta_{im} r_t^{m,e} + \beta_{ib}^{\perp} r_t^{b,\perp,e} + \epsilon_t^i, \tag{14}$$

where ϵ_t^i are iid normally distributed errors with zero means. In the second step, for each time stamp s, we average these residuals across all firms represented at time s:

$$\hat{f}_s = \frac{1}{N} \sum_{i=1}^{N} \hat{\alpha}^i + \hat{\epsilon}_s^i. \tag{15}$$

The time series $\hat{f}_s$ captures the common variation of the residuals $\hat{\epsilon}_s^i$. The extent to which $\hat{f}_s$ captures *only* the common variation in these residuals depends on the number of bonds in the sample as well as the degree of correlation across residuals. For instance, if $\epsilon_s^i = \rho^i Z_s + \xi_s^i$, where Z_s and ξ_s^i are iid normal variables and if N is

[16]Several studies have addressed the issue of market integration between the U.S. and European stock markets. Some of the best known studies include: Karolyi and Stulz (2003), Bekaert and Harvey (1995), Bekaert et al. (2005), Harvey (1995), Ferson and Harvey (1993), Griffin (2002) and others.

sufficiently large, $\hat{f}_s$ becomes

$$\hat{f}_s = \frac{1}{N} \sum_{i=1}^{N} \rho^i Z_s \tag{16}$$

and therefore it captures only the common variation in the residuals.

While there are other ways to construct measures of common variation for our residuals—we construct and test a few alternatives in a later section—we choose this particular methodology for its simplicity and ease of interpretation. Intuitively, f_s picks up the fixed time effects of the residuals, which are the output of the firm-specific regressions of bond excess returns on the market and term-spread excess returns. We will refer to f_s as the CMF.

To implement this methodology, we need to compute the bond returns on the left-hand side of Equation (14). In fact, the pricing restriction in Equation (10), while exemplified so far with bond returns, holds true for any kind of returns, as long as long as they are *holding* returns.

Given the superior liquidity of the CDS market over the corporate bond market, for instance, our intent is to use the quote data on CDS spreads to compute the left-hand side returns. Holding returns on CDS contracts are neither readily available nor easy to compute given that we only have data on the CDS premium (spread) set forth at the outset of the contract. One way to circumvent this problem is to use the pricing information of the CDS spread quotes to construct holding returns on zero-coupon bonds of various maturities and then use these holding returns in excess of the risk-free rate as our primary dependent variable in the first step of our methodology. The drawback of this approach is that returns computed in this manner will reflect not only the informational content of the CDS spread quotes, but also the assumptions used in order to extract this information. This critique applies, however, whenever returns are inferred rather than readily available. For robustness, we also consider an alternative approach which approximates holding returns with the difference between two consecutive quotes.[17] The results under this specification are presented in a later section.

The next step is to compute returns on defaultable zero-coupon bonds. Relying on the notation from Section 3, the holding return between time t and time $t+1$ for

[17]Since the quotes at time t and $t-1$ are essentially quotes on different issues (same type of CDS contract on the same entity but originated at different times), this measure of holding return will be a very coarse approximation of actual holding returns. Hence results under this specification should be interpreted with this caveat in mind.

a defaultable zero-coupon zero-recovery bond with maturity $T > t+1$ is given by

$$r_{t+1} \quad = \quad \log P(t+1, T-t-1) - \log P(t, T-t), \tag{17}$$

where $P(t, T-t)$ denotes the time-t market value of the bond. As a special case, the holding return for the period $[T-1, T]$ is computed as

$$r_T \quad = \quad -\log P(T-1, T). \tag{18}$$

It is important to notice that these returns cannot be computed directly because we do not have data on defaultable zero-coupon corporate bonds. They will be inferred from CDS rates as explained in Section 3.

We start with the formula in Equation (2). 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 homoscedastic and conditionally normally distributed) and $\lambda^{Q,n}$ is orthogonal on the information contained in the consumption growth rates and the market returns.[18] Then,

$$\begin{aligned} P(t, T-t) &= E_t\left[M_{t,T} e^{-\int_t^T \lambda_s^Q \mathrm{d}s}\right] = E\left[M_{t,T} e^{-\int_t^T \lambda_s^{Q,c}}\right] E_t\left[e^{-\int_t^T \lambda_s^{Q,n}}\right] \\ &= E_t^Q\left[e^{-\int_t^T r_s^f + \lambda_s^{Q,c}}\right] E_t\left[e^{-\int_t^T \lambda_s^{Q,n}}\right]. \end{aligned} \tag{19}$$

Since r_t and $\lambda_t^{Q,c}$ are correlated Gaussian processes, it is easy to established that

$$\log E_t^Q\left[e^{-\int_t^T r_s^f + \lambda_s^{Q,c}}\right] = A(T-t) - B(T-t) r_t^f - C(T-t)\lambda_t^{Q,c}, \tag{20}$$

where $A(T-t)$, $B(T-t)$ and $C(T-t)$ depend on $T-t$ only.[19] Thus, $\log P(t, T-t)$ can be rewritten as

$$\log P(t, T-t) = A(T-t) - B(T-t) r_t^f - C(T-t)\lambda_t^{Q,c} + \log E_t\left[e^{-\int_t^T \lambda_s^{Q,n}}\right]. \tag{21}$$

Combining, we obtain the following expression for r_{t+1}.

$$\begin{aligned} r_{t+1} = &\left[B(T-t) r_t^f - B(T-t-1) r_{t+1}^f\right] \\ &+ \left[C(T-t)\lambda_t^{Q,c} - C(T-t-1)\lambda_{t+1}^{Q,c}\right] + r_{t+1}^n, \end{aligned} \tag{22}$$

where $r_{t+1}^n = A(T-t) - A(T-t-1) + \log E_{t+1}\left[e^{-\int_{t+1}^T \lambda_s^{Q,n}}\right] - \log E_t\left[e^{-\int_t^T \lambda_s^{Q,n}}\right]$ captures

[18]See Footnote 14 for a way to construct such a decomposition

[19]See the appendix for the derivation of these coefficients.

the nonlognormal component of the returns. For a firm i, the expected-returns beta representation in Equation (11) can be restated in the following form.

$$E\left[B^i(T-t)r_t^f - B^i(T-t-1)r_{t+1}^f\right] + E\left[C^i(T-t)\lambda_t^{i,Q,c} - C^i(T-t-1)\lambda_{t+1}^{i,Q,c}\right] \\ - Er_{t+1}^f + \frac{1}{2}V_{ii}^c = \beta_{im}^c\left[Er_{t+1}^{m,e} + \frac{1}{2}V_{mm}\right] + \beta_{ib}^{c,\perp}\left[Er_{t+1}^{b,\perp,e} + \frac{1}{2}V_{bb}^{\perp}\right] + Ez_{t+1}^i \quad (23)$$

or equivalently

$$E\left[B^i(T-t)r_t^f - B^i(T-t-1)r_{t+1}^f\right] + E\left[C^i(T-t)\lambda_t^{i,Q} - C^i(T-t-1)\lambda_{t+1}^{i,Q}\right] - Er_{t+1}^f \\ + \frac{1}{2}V_{ii} = \beta_{im}\left[Er_{t+1}^{m,e} + \frac{1}{2}V_{mm}\right] + \beta_{ib}^{\perp}\left[Er_{t+1}^{b,\perp,e} + \frac{1}{2}V_{bb}^{\perp}\right] + \left[E\Delta z_{t+1}^i + \frac{1}{2}V_{ii}^n\right], \quad (24)$$

where $z_{t+1}^i = -\log E_t e^{r_{t+1}^{i,n}}$, $\Delta z_{t+1}^i = z_{t+1}^i - \left[C^i(T-t)\lambda_t^{i,Q,n} - C^i(T-t-1)\lambda_{t+1}^{i,Q,n}\right]$ and V_{ii}^n is the variance of the nonlognormal return component $r^{i,n}$. The return model in Equation (13) becomes

$$B^i(T-t)r_t^f - B^i(T-t-1)r_{t+1}^f + C^i(T-t)\lambda_t^{i,Q} - C^i(T-t-1)\lambda_{t+1}^{i,Q} \\ - r_{t+1}^f = \alpha^i + \beta_{im}r_{t+1}^{m,e} + \beta_{ib}^{\perp}r_{t+1}^{b,\perp,e} + \Delta z_{t+1}^i + \epsilon_{t+1}^i \quad (25)$$

and thus the excess returns to be used in the first step of our methodology, Equation (14), can now be computed as a function of the default intensities:

$$r_{t+1}^{i,e} = B^i(T-t)r_t^f - B^i(T-t-1)r_{t+1}^f + C^i(T-t)\lambda_t^{i,Q} - C^i(T-t-1)\lambda_{t+1}^{i,Q} - r_{t+1}^f. \quad (26)$$

Given the estimated time series for the risk-neutral default intensities from Section 3, we can now compute returns of defaultable zero-coupon bonds. In what follows, we will formally test the expected-returns beta representation derived in Section 4.

6. Time Variation in Defaultable Debt Returns

This section deals with the implementation of the two-step methodology for identifying the sources of common variation in zero-coupon zero-recovery bond excess returns of various maturities.

In order to implement the first step, we need to proxy for the two zero-cost portfolios capturing the systematic risk, namely the market portfolio and the portfolio that longs a riskless console bond paying one unit of account every period and shorts the short interest rate. For this exercise, we set the holding period to a week.

The reference entities behind the CDS contracts in our data set are from various European countries. Most of these countries are also part of the European Monetary Union[20] but there are few countries that are not (Denmark, Norway, Sweden and the UK). Since capital markets throughout Europe are more or less integrated,[21] 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 capital markets integration, whenever possible we use portfolios constructed from the data disseminated in the electronic version of the 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 2.2. Since the CDS spreads in our data set 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 2.2.

We proxy for the other source of systematic risk—captured by the portfolio which longs a riskless console bond paying one unit of account every week and shorts the short interest rate—with the term-spread portfolio, which longs the 30-year zero-coupon riskless Euro bond and shorts the one-month Euro bond. We denote the weekly returns of this portfolio with r_t^{TERM}.

The first step in extracting the common variation in the returns inferred from the CDS spreads is to extract the residuals from the following firm-specific regressions.

$$r_{t+1}^{i,e} = \alpha^i + \beta_{EMKT}^i r_{t+1}^{EMKT,e} + \beta_{CMKT}^i r_{t+1}^{CMKT,e} + \beta_{TERM}^i r_{t+1}^{TERM} + \epsilon_{t+1}^i \qquad (27)$$

The excess returns on the left-hand side are excess returns on zero-coupon zero-recovery corporate bonds and they depend on the risk-neutral default intensity and maturity. The default intensities are extracted from the CDS data as indicated in the previous section. Since these excess returns are inferred from CDS spreads data, we

[20]See Table I for more details

[21]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).

have no a priori preference for a specific maturity. To limit the potential impact of the choice of maturity on our results, we perform the same empirical exercise for six maturities: one week, one year, two years, three years, four years and five years. In addition, as mentioned in Section 4, the excess returns on the left-hand side are not lognormally distributed, in general, and thus the residuals will not be lognormally distributed either. This observation is important because the typical t-statistic might not be very informative in the context of these regressions.

Table III presents the averages of the estimated coefficients of these regressions along with their average standard errors. We present the results across the six maturities to assess the impact of the choice of maturity.

For all choices of maturity, the results show that the common risk factors considered explain relatively little from the time variation of the LHS excess returns—in all instances the adjusted R^2 is around 15%. Also notable is the size of the pricing errors and their standard errors. Except for the one-week maturity, the absolute value of the pricing error increases with the maturity as well, and it ranges from 6 bps to about 21 bps. These errors are consistently different from 0, with t-statistics of at least 4. These are relatively sizable pricing errors (for weekly returns), suggesting that the return model based only on the market and the TERM factors could be misspecified. This is not surprising given that the independent variables are close to being lognormally distributed while the dependent variable is not. More importantly, from the perspective of the model derived in Section 4, these results are reassuring.

We now proceed with implementing the second step of our methodology as described in Section 5. Using the residuals extracted from the first step and Equation (15), we construct estimates for the CMF at every point in time.

The next to last column of Table III reports the average of the fraction of the pricing error explained by the CMF. We notice that the CMF captures as much as 53.84%, on average, of the time variation of the pricing errors of the risk-adjusted excess returns of zero-coupon zero-recovery bonds with a maturity of five years. This fraction decreases to 29.83% as we decrease the maturity of the bonds to one year. In the special case, when the maturity is one week, the CMF captures 35.28% of the pricing error. The last column of Table III also reports the increase in the R^2 of the regression of the CDS-inferred excess bond returns on the two macroeconomic factor *and* the CMF. We notice that by accounting for the CMF, the R^2 of the regression increases by at least 4 times.

While the change in the average magnitude from one week to one year might seem a bit puzzling, we should keep in mind that the returns of the zero-coupon zero-recovery bonds with time to maturity of one week are special because the payoff of these bonds is 1. This is unique to the time to maturity of one week, because the holding period throughout this study is fixed at one week as well.

According to the model developed in Section 4, the CMF captures the common variation of the nonlognormal component of CDS-inferred returns. The same model tells us that this factor impacts expected returns level *directly* rather than through covariances. If our model is correct, this result suggests that the CMF picks up common variation in the nonlognormal component of CDS-inferred returns that does not originate from exposure to the common risk factors endogenous to the model. Such a scenario is not completely implausible and here is why.

The spreads of a new CDS contract or the value of a corporate bond at some point in time depend on the investors' assessment of the distribution of the default event at that time. This dependency is particularly strong for zero-coupon zero-recovery bonds given that the uncertainty in the payoff structure of these assets reduces only to the uncertainty about the occurrence of the default event. In this context, the occurrence of a new credit event gives investors an opportunity to learn more about (update) the distribution of firm-specific extreme events leading to default. In particular, if the observed credit event reveals that certain extreme events are more likely than previously thought, investors would react rationally by updating the distribution of the default event across the entire universe of credit-contingent assets—including corporate bonds and CDS contracts. This systemic reaction could lead to commonality in the price (or return) behavior of these assets inducing common variation. It is important to notice that such commonality in price behavior might not be related to changes in the underlying macroeconomic risk factors. That is because the update on the investor side is essentially a revision of the distribution of payoffs and this revision might not lead to changes in discount rates.

To illustrate this argument, consider the period surrounding the U.S. corporate scandals in 2001–2002. The default of Enron and Worldcom in 2001 uncovered that certain "value-enhancing" accounting practices were more commonly employed across firms then previously thought. The downward correction in firm value demanded by rational investors in the aftermath of these credit events led mechanically to an upward correction in the likelihood of the default event (because firms were now closer to their default boundary). These revisions led to a drop in the corporate bond prices and, at the same time, induced bondholders to scramble for insurance—driving the CDS spreads up.

Thus, one can argue that the CMF is capturing the systemic revision of the default event's distribution across firms, following the occurrence of a seemingly unrelated corporate default. The next section investigates whether this theory is supported by the data.

7. Supporting Evidence

The CMF is extracted from the CDS data of the firms with the most liquid CDS market during our sample period. As documented in the previous section, this factor is able to capture a large fraction of the time variation in the returns of zero-coupon zero-recovery corporate bonds of the firms in our sample. The asset pricing model developed in Section 4 suggests that the common variation in these returns reflects exposure to macroeconomic risk factors only to the extent that this common variation originates from exposure to two sources of systematic risk: the market and the term spread. By construction, the CMF captures common variation beyond whatever can be explained by these sources of macroeconomic risk. One possible explanation suggested by our model is that the CMF captures the systemic response of the investors who act upon observed corporate defaults by revising their assessment of the default event's distribution. In this section, we investigate whether this theory has support in the data.

The revision of the distribution of the default event should impact the pricing of all credit-contingent assets. This means that, to the extent that we observe comovement in spreads across the CDS market, we should observe a similar phenomenon in the corporate bond market as well. This is not necessarily a tautology, despite the fact that CDS contracts are a natural hedge for corporate bonds. When the reference entities behind corporate bonds and CDS contracts are different firms with different characteristics, the hedging effect is not likely to be the source of the comovement in the prices of these two classes of credit-contingent assets. To ensure that the reference entities behind the corporate bonds used in our tests are different enough from the firms in our CDS sample, we consider a variety of corporate bond portfolios sorted on three different characteristics: rating, time to maturity and sector. These portfolios are preconstructed by either Merrill Lynch or Lehman Brothers and they focus on either the entire universe of European corporate bonds or on the nonfinancial/industrial sectors.[22].

Riskier bonds—such as those with lower ratings or longer time to maturity—are

[22]For more information on these portfolios see Section 2.2.

more likely to display skewness and fat tails in their return distributions. Our theory implies that if investors revise the default event's distribution in the aftermath of a corporate default, the returns of the riskier bonds should reflect the impact of these revisions to a larger extent.

To illustrate this point consider the following stylized example. Suppose an investor holds two zero-coupon bonds with the same maturity of one year. As of right now, Bond A pays 1 in states s_1 and s_2 and 0 in state s_3. Bond B pays 1 for sure. Suppose that one week later the investor realizes that s_1 could lead to default and so both bonds will now pay 0 in that state. Suppose further that each of these three states is equally likely and that discount rates remain the same throughout the year. Let r denote the weekly discount rate. Prior to the revision, the price of bond A is $P_0^A = \frac{2}{3}e^{-50r}$, while the price of bond B is $P_0^B = e^{-50r}$. After the revision, these prices change to $P_1^A = \frac{1}{3}e^{-49r}$ and $P_1^B = \frac{2}{3}e^{-49r}$, respectively. The net holding returns on these bonds can be computed as $r_1^A = \log \frac{P_1^A}{P_0^A} = r \log \frac{1}{2}$ and $r_1^B = r \log \frac{2}{3}$, respectively. If there were no revision, the net return on both bonds should be r. However, the revision is bad news for both bonds and their prices drop, leading to negative net returns. Thus, consistent with our theory, the revision induces the net returns of these two bonds to move in the same direction. More importantly, the riskier bond, bond A, records a larger loss: $r - r_1^A > r - r_1^B$

We now present evidence supporting the dual hypothesis that systemic revisions to a default event's distribution—as captured by the CMF—lead to comovement in corporate bond returns and that these revisions have more impact on the riskier bonds.

We run time-series regressions,

$$r_{t+1}^{i,e} = \alpha^i + \beta_{EMKT}^i r_{t+1}^{EMKT,e} + \beta_{TERM}^i r_{t+1}^{TERM} + \beta_{CMF}^i CMF_{t+1} + \epsilon_{t+1}^i, \tag{28}$$

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

Tables IV–X report the estimated coefficients for the corporate bond portfolios we use as test assets. The CMF used in these tests is extracted from defaultable zero-coupon bonds maturing in five years. The results for the other five choices of time to maturity are illustrated in Figures 2–5.

The results reported in these tables overwhelmingly support our dual hypothesis. All portfolios load positively on the CMF and these loadings trend in the direction suggested by the characteristic used in constructing these portfolios. For instance, the Merrill Lynch portfolios sorted on credit quality (rating) load heavier on the CMF when the rating is lower. Similarly, the Merrill Lynch portfolios within a rating class, sorted further on maturity, load heavier on the CMF when the time to maturity is larger. The same pattern can be observed for the Merrill Lynch portfolios of nonfinancials sorted on maturity. These patterns are further supported by all the Lehman Brothers portfolios sorted on either rating or time to maturity. Noticeable here is the loading of the high-yield portfolio which is almost 3 times higher than the loading of the BAA-rated corporate bonds and more than 80 times higher than the loading of the AAA-rated corporate bonds. These results transgress the choice for time to maturity when constructing the CMF. Figures 2–6 show that most of the patterns continue to hold when the time to maturity varies from one week to five years.

To increase the power of our test, we can run the previous time-series regressions as pooled time-series regressions. Specifically, for each group of portfolios—Lehman Brothers sorted on rating, maturity or sector and Merrill Lynch sorted on rating, maturity or both—we run the following pooled regression.

$$r_{t+1}^{i,e} = \alpha + \beta_{EMKT} r_{t+1}^{EMKT,e} + \beta_{TERM} r_{t+1}^{TERM} + \beta_{CMF}\, CMF_{t+1} + \epsilon_{t+1}^{i}, \quad (29)$$

where i is an index for corporate bond portfolios in a given group.

Table XI presents the results. In all instances the loading on the CMF is always positive and significant at a 5% level (after correcting for lags using the Newey–West procedure). Once again, this provides support for the importance of the CMF in explaining the time-variation of corporate bond returns.

While we expect the CMF to have an impact on corporate bond portfolios—the returns of these portfolios are more likely to be nonlognormally distributed—we also expect that the CMF will have no impact on equity portfolios—the returns of these portfolios are more likely to be lognormally distributed. To test if this is the case, we run pooled time-series regressions similar to the ones in Equation (29) for country-specific equity portfolios sorted on sector. These portfolios are based on the price-level sector indexes for each country, available from Datastream. The results are presented in Table XII. For all countries considered, our results support the hypothesis that the CMF has no impact on the returns of the equity portfolios. This result reassures us once again that the CMF is likely to capture a price behavior specific to assets with nonlognormal returns only.

Tables IV–X also reveal 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 flight to quality effect. As the economy goes through an expansion, investors' appetite for risk increases and they are 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 high-rating long-term corporate bonds.

8. Discussion and Conclusion

To ensure the robustness of our findings, we investigate two alternative ways of extracting the CMF. We then investigate whether the results of the previous section remain valid.

The first alternative computes the returns of the zero-coupon zero-recovery defaultable bonds used on the left-hand side of the regression in Equation (14) simply as the difference between two consecutive five-year CDS market rates. This approach has the advantage of being nonparametric (and model free). However, it only provides a very coarse approximation of the actual returns.

The second alternative proposes a different way to extract the CMF. Essentially, the extracted value for the CMF at time t equals the loading on the time dummy at time t, δ_t in the following pooled regression.

$$r_t^{i,e} = \alpha^i + \beta_{im} r_t^{m,e} + \beta_{ib}^{\perp} r_t^{b,\perp,e} + \sum_t \delta_t + \epsilon_t^i, \tag{30}$$

where $\delta_t = 1$ if the time stamp is t, and 0 otherwise. In both cases, our unreported results support qualitatively and, sometimes, quantitatively the results reported in the previous section.

In summary, the CDS market is one of the largest and most liquid markets and comparable in many respects with the equity market. Yet, there are times when these two markets seem to move very differently on an aggregate level. The question is, Why? If common variation in these markets arises exclusively as a consequence of exposure to the same macroeconomic risk, these markets should move in sync. Yet that does not seem to happen all the time. In this paper, we try to understand the sources of common time-series variation in the premiums of the CDS contracts.

We use a slightly modified version of the Campbell's ICAPM to characterize the risk premia of the assets with and without lognormal returns. According to the model, the common variation in the returns of assets with lognormal returns can only arise from exposure to two macroeconomic risk factors: the market and the term spread—the spread between the long and the short ends of the term structure of interest rates. However, common variation in the returns of the assets with nonlognormal returns can also arise if investors systemically revise the distribution of the default event in the aftermath of a corporate default.

Using European CDS, corporate bond and equity data, we provide evidence in support of this theory. To the extent that investors learn from corporate defaults and update the distribution of the default event, the impact of these revisions should be particularly high for defaultable zero-coupon zero-recovery corporate bonds. Returns on these type of bonds are not readily available, but they can be inferred from CDS data, which is typically available. We construct such returns and identify a common component in these returns that captures the systemic updating on the part of the investors, as suggested by the theory. We call this component the CMF.

Our tests concentrate around corporate bond and equity portfolios. To overcome the potential hedging bias, we consider a large variety of bond portfolios sorted on rating, time to maturity and sector. We find that corporate bond portfolios respond to innovations in the two macroeconomic risk factors, but they also respond positively to innovations in the CMF. All our equity portfolios—presorted on sector and country—show little or no response to innovations in the CMF. This is consistent with our theory since equity returns are more likely to be lognormally distributed and should only respond to innovations in the macroeconomic risk factors.

The model and the evidence provided in this paper seem to suggest that the sources of common variation for a particular market do not necessarily have to be associated with macroeconomic risk factors. This point has been made previously by Daniel and Titman (1997) for equity markets. We expand the focus of this point to the credit and corporate bond markets.

A. Derivation of the Coefficients for Equation (20)

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.

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

Then, for any $t < T$, we have

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

with the initial conditions $A(0) = B(0) = C(0) = 0$. Notice that under the decomposition suggested in Footnote (14), k_λ can be computed as follows:

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

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Tables

Table I: Distribution of firms across industries and countries

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

Firms are grouped into industries according to the Moody's industry classification.

Table II: Estimation of the risk-neutral default intensities

	κ^Q	σ^Q	ρ^Q	mean(θ^Q)	$\tilde{\kappa}^Q$	mean($\tilde{\kappa}^Q\,\tilde{\theta}^Q$)	no firms
Denmark	0.154	0.444	–	3.624	−0.123	−0.220	1
Finland	0.290	0.410	0.309	3.870	0.003	0.242	5
France	1.363	1.455	0.683	2.959	0.158	0.402	13
Germany	1.764	1.274	0.591	2.962	0.224	0.722	10
Greece	4.533	1.043	–	3.182	0.186	0.768	1
Italy	1.086	1.872	0.767	3.257	0.306	0.611	2
Netherlands	3.339	0.938	0.902	2.848	−0.206	−0.854	4
Norway	2.640	0.257	–	2.758	−0.159	−0.214	1
Spain	2.512	0.510	0.287	3.070	0.015	0.325	2
Sweden	0.673	1.098	0.220	3.246	0.289	1.401	5
UK	1.143	0.849	0.357	3.530	0.182	0.903	11

Summary statistics for the country-by-country parameter estimates describing the dynamics of the risk-neutral default intensities in Equation (4).

Table III: Estimates for the return model in Equation (27)

α	β_{EMKT}	β_{CMKT}	β_{TERM}	$Perc(C)$	$Perc(S)$
0.0000	−0.0005	0.0001	0.1607	35.28	404.04
(3.3160)	(−0.0101)	(0.0077)	(2.6410)		
−0.0006	0.0045	0.0028	1.1120	29.83	405.40
(−4.3896)	(0.7568)	(0.1908)	(3.7908)		
−0.0010	0.0064	0.0031	1.9146	39.16	460.73
(−4.8849)	(0.6559)	(0.1730)	(4.1127)		
−0.0014	0.0073	0.0030	2.6169	45.85	497.29
(−5.1307)	(0.5875)	(0.1595)	(4.2915)		
−0.0018	0.0080	0.0027	3.2480	50.50	521.64
(−5.2826)	(0.5376)	(0.1498)	(4.4050)		
−0.0021	0.0086	0.0024	3.8203	53.84	538.83
(−5.3858)	(0.4992)	(0.1426)	(4.4827)		

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 weeks between January 2003 and October 2006, 197 weeks. The left-hand side excess returns correspond to defaultable bonds with the following times to expiration: one week, one year, two years, three years, four years and five 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 maturity. The CMF at time t corresponds to the slope coefficient of the dummy controlling for time t. The $Perc(C)$ column reports the fraction of the pricing error, obtained from the first step, explained by the CMF. The $Perc(S)$ column reports the increase in adjusted R^2 when the CMF is added to the independent variables in the regression of step 1. The t-statistics are reported in parentheses. The reported values for the estimates are averages across firms of the corresponding firm-specific estimates.

Table IV: The Merrill Lynch nonfinancial corporate bond portfolios sorted on rating

α	β_{EMKT}	β_{TERM}	β_{CMF}	adj R^2	$E[R]$
0.0005	−0.0595	0.3449	0.2627	0.1995	0.0006
(0.5557)	(−4.3825)	(0.1998)	(0.9530)		
0.0003	−0.0449	1.3751	0.3485	0.0999	0.0007
(0.2899)	(−3.2363)	(0.6576)	(1.1975)		
0.0003	−0.0182	2.0377	0.4504	0.0458	0.0009
(0.2616)	(−1.1118)	(1.0694)	(1.6468)		

This table reports the results of the time-series regressions of the excess realized returns of three Merrill Lynch nonfinancial 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 between January 2003 and October 2006, 197 weeks. The CMF is extracted from returns on defaultable zero-coupon bonds maturing in five years according to the model in Equation (27). The first line corresponds to the higher-rating portfolio. The Newey–West t-statistics (adjusting for 3 lags) are reported in parentheses.

Table V: AAA and AA Merrill Lynch corporate bond portfolios sorted on maturity

α	β_{EMKT}	β_{TERM}	β_{CMF}	adj R^2	$E[R]$
	AAA Portfolios Sorted on Maturity				
0.0001	−0.0228	0.6924	0.1827	0.1864	0.0005
(0.1321)	(−4.9212)	(0.9058)	(1.7263)		
0.0003	−0.0561	0.8716	0.3009	0.1882	0.0005
(0.3379)	(−4.6869)	(0.4915)	(1.1967)		
0.0007	−0.0759	0.6578	0.3428	0.1754	0.0007
(0.4863)	(−4.3608)	(0.2709)	(0.9914)		
0.0012	−0.1068	−0.1108	0.2672	0.1413	0.0010
(0.5301)	(−4.0024)	(−0.0278)	(0.4943)		
	AA Portfolios Sorted on Maturity				
0.0002	−0.0405	0.9782	0.2699	0.1756	0.0005
(0.2564)	(−4.6479)	(0.7125)	(1.4026)		
0.0006	−0.0735	0.9079	0.3674	0.1689	0.0007
(0.4663)	(−4.2552)	(0.3741)	(1.0345)		
0.0011	−0.0852	0.4482	0.3962	0.1472	0.0009
(0.6194)	(−4.0652)	(0.1486)	(0.9308)		

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 between January 2003 and October 2006, 197 weeks. The CMF is extracted from returns on defaultable zero-coupon bonds maturing in five years according to the model in Equation (27). The first line in each of the two panels corresponds to the lower-maturity portfolio. The t–statistics are reported in parentheses.

Table VI: A and BBB Merrill Lynch corporate bond portfolios sorted on maturity

α	β_{EMKT}	β_{TERM}	β_{CMF}	R^2	$E[R]$
	A Portfolios Sorted on Maturity				
0.0002	−0.0333	1.1517	0.2764	0.1416	0.0006
(0.2502)	(−3.9213)	(0.8837)	(1.4436)		
0.0007	−0.0638	1.1890	0.4313	0.1398	0.0008
(0.5141)	(−3.7003)	(0.5009)	(1.1918)		
0.0010	−0.0683	1.0885	0.4733	0.1068	0.0010
(0.5703)	(−3.1622)	(0.3642)	(1.1045)		
0.0015	−0.0607	1.6356	0.7687	0.0577	0.0012
(0.6192)	(−2.2942)	(0.4022)	(1.3974)		
	BBB Portfolios Sorted on Maturity				
0.0001	−0.0107	1.7770	0.3616	0.0546	0.0008
(0.1794)	(−1.1048)	(1.4401)	(1.9255)		
0.0004	−0.0341	2.4134	0.5214	0.0568	0.0009
(0.2502)	(−1.5129)	(1.0277)	(1.3538)		
0.0008	−0.0236	2.5044	0.6602	0.0359	0.0012
(0.4681)	(−0.9869)	(0.8136)	(1.5680)		

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 between January 2003 and October 2006, 197 weeks. The CMF is extracted from returns on defaultable zero-coupon bonds maturing in five years according to the model in Equation (27). The first line in each of the two panels corresponds to the lower-maturity portfolio. The Newey–West t-statistics (adjusted for 3 lags) are reported in parentheses.

Table VII: The Merrill Lynch nonfinancial corporate bond portfolios sorted on maturity

α	β_{EMKT}	β_{TERM}	β_{CMF}	adj R^2	$E[R]$
0.0000	−0.0145	1.2130	0.2207	0.0909	0.0006
(0.01947)	(−2.1089)	(1.5237)	(1.8600)		
0.0002	−0.0328	1.8151	0.3889	0.0994	0.0007
(0.1957)	(−2.8268)	(1.0831)	(1.6571)		
0.0004	−0.0493	1.7211	0.4422	0.0927	0.0008
(0.3192)	(−2.5540)	(0.7296)	(1.2075)		
0.0014	−0.0597	2.1434	0.7049	0.0474	0.0015
(0.5520)	(−1.6539)	(0.5077)	(1.2781)		

This table reports the results of the time-series regressions of the excess realized returns of four Merrill Lynch nonfinancial 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 between January 2003 and October 2006, 197 weeks. The CMF is extracted from returns on defaultable zero-coupon bonds maturing in five years according to the model in Equation (27). The first line corresponds to the lower-maturity portfolio. The Newey–West t-statistics (adjusted for 3 lags) are reported in parentheses.

Table VIII: The Lehman Brothers Euro-aggregate industrial corporate bond portfolios sorted on rating

α	β_{EMKT}	β_{TERM}	β_{CMF}	R^2	$E[R]$
−0.0007	−0.0550	0.7849	0.1591	0.1627	−0.0003
(−0.7199)	(−3.8647)	(0.4435)	(0.5781)		
−0.0002	−0.0605	0.1914	0.2726	0.1377	−0.0003
(−0.1897)	(−4.2683)	(0.0889)	(0.8825)		
−0.0003	−0.0496	0.4817	0.2871	0.0914	−0.0002
(−0.2315)	(−3.0092)	(0.2214)	(1.0274)		
−0.0004	−0.0370	1.3619	0.4976	0.0708	−0.0003
(−0.3551)	(−2.7172)	(0.6662)	(1.6431)		
0.0015	0.0554	3.1904	1.3851	0.0618	0.0011
(0.5592)	(1.9091)	(0.9239)	(1.7404)		

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 between January 2003 and October 2006, 197 weeks. The CMF is extracted from returns on defaultable zero-coupon bonds maturing in five years according to the model in Equation (27). The first line corresponds to the higher-rating portfolio. The t-statistics are reported in parentheses.

Table IX: The Lehman Brothers Euro-aggregate corporate bond portfolios sorted on maturity

α	β_{EMKT}	β_{TERM}	β_{CMF}	adj R^2	$E[R]$
−0.0006	−0.0188	1.0872	0.2645	0.1186	−0.0001
(−1.1988)	(−3.4722)	(1.1677)	(2.1299)		
−0.0003	−0.0385	0.8861	0.3908	0.1066	−0.0002
(−0.3513)	(−3.4332)	(0.5141)	(1.6555)		
−0.0008	−0.0693	2.1092	0.3910	0.1417	−0.0002
(−0.5271)	(−3.6313)	(0.8244)	(1.0410)		
0.0000	−0.0763	0.0900	0.4878	0.1089	−0.0005
(0.0072)	(−3.2838)	(0.0279)	(1.1081)		
0.0012	−0.1013	2.6832	1.1192	0.0794	0.0007
(0.3786)	(−2.9269)	(0.5325)	(1.5013)		

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 between January 2003 and October 2006, 197 weeks. The CMF is extracted from returns on defaultable zero-coupon bonds maturing in five years according to the model in Equation (27). The first line corresponds to the lower-maturity portfolio. The Newey–West t-statistics (adjusted for 3 lags) are reported in parentheses.

Table X: The Lehman Brothers Euro-aggregate corporate bond portfolios sorted on sector

α	β_{EMKT}	β_{TERM}	β_{CMF}	R^2	$E[R]$
	Cross-Sectional Averages of the Estimates				
-0.0004	-0.0416	1.1984	0.4301	0.0703	-0.0002
	Cross-Sectional Standard Deviations of the Estimates				
0.0004	0.0159	0.8203	0.2036	0.0337	0.0001
	Cross-Sectional Averages of the t-Statistics				
(-0.3498)	(-3.3661)	(0.4835)	(1.5052)		
	Cross-Sectional Standard Deviations of the t-Statistics				
(0.2960)	(1.1992)	(0.2949)	(0.5092)		

This table reports the results of the time-series regressions of the excess realized returns of 16 Lehman Brothers Euro-aggregate corporate bond portfolios sorted on sector (Aero/Defense, Banking, Building Materials, Capital Goods, Chemicals, Communications, Consumer Noncyclical, Consumer Cyclical, Diversified Manufacturing, Food and Beverages, Lodging, Refining, Telephone, Tobacco, Wireless and Media Noncable), on the excess market returns ($EMKT$), the spread between long and short Euro bonds ($TERM$) and the CMF between January 2003 and October 2006, 197 weeks. The CMF is extracted from returns on defaultable zero-coupon bonds maturing in five years according to the model in Equation (27). Each panel reports an average statistic across portfolios.

Table XI: Pooled regressions for corporate bond portfolios

α	β_{EMKT}	β_{TERM}	β_{CMF}	adj R^2	$E[R]$
	Merrill Lynch sorted on rating				
0.0004	−0.0409	1.2526	0.3539	0.0963	0.0003
(0.6173)	(−4.4799)	(1.1253)	(2.1702)		
	Merrill Lynch sorted on rating/maturity				
0.0006	−0.0539	1.1575	0.4014	0.0917	0.0003
(1.5949)	(−9.2750)	(1.6852)	(4.0677)		
	Merrill Lynch sorted on maturity				
0.0005	−0.0391	1.7231	0.4392	0.0535	0.0004
(0.6617)	(−3.2384)	(1.3380)	(2.3995)		
	Lehman Brothers sorted on rating				
−0.0000	−0.0293	1.2021	0.5203	0.0329	−0.0005
(−0.0502)	(−3.0176)	(1.0575)	(2.7349)		
	Lehman Brothers sorted on maturity				
−0.0001	−0.0609	1.3711	0.5306	0.0767	−0.0005
(−0.1112)	(−5.3811)	(1.0059)	(2.7049)		
	Lehman Brothers sorted on sectors				
−0.0004	−0.0416	1.1984	0.4301	0.0706	−0.0007
(−1.1990)	(−9.5535)	(2.0372)	(4.7138)		

This table reports the results of the pooled regressions of the excess realized returns of corporate bond portfolios on the excess market returns ($EMKT$), the spread between long and short Euro bonds ($TERM$) and the CMF between January 2003 and October 2006, 197 weeks. The CMF is extracted from returns on defaultable zero-coupon bonds maturing in five years according to the model in Equation (27). The first line corresponds to the lower-maturity portfolio. The Newey–West t-statistics (adjusted for 3 lags) are reported in parentheses.

Table XII: Pooled regressions for equity portfolios

α	β_{EMKT}	β_{TERM}	β_{CMKT}	β_{CMF}	adj R^2	$E[R]$
			Finland			
−0.0008	0.2686	5.4926	0.2005	−0.5073	0.1643	0.0031
(−0.8230)	(11.908)	(2.9977)	(14.839)	(−2.3018)	$n = 75$	
			France			
0.0026	0.0945	−3.6029	0.6705	−0.0862	0.3437	0.0022
(3.9095)	(1.9018)	(−3.0799)	(12.803)	(−0.5036)	$n = 113$	
			Germany			
0.0018	0.0799	−1.3920	0.6187	0.0080	0.3352	0.0023
(2.4525)	(1.6718)	(−1.0053)	(16.535)	(0.0403)	$n = 112$	
			Netherlands			
0.0023	0.3318	−2.4220	0.3678	0.2469	0.2984	0.0019
(2.6127)	(9.1078)	(−1.5138)	(10.7255)	(1.0287)	$n = 76$	
			Sweden			
0.0017	0.0631	−2.6421	0.6434	−0.3013	0.4751	0.0028
(2.2894)	(2.3808)	(−1.8971)	(32.703)	(−1.7799)	$n = 69$	
			UK			
0.0021	−0.0095	−1.9132	0.6924	0.2251	0.2987	0.0018
(2.7902)	(−0.5649)	(−1.4125)	(30.818)	(1.2429)	$n = 75$	

This table reports the results of the pooled regressions of the excess realized returns of country-specific equity portfolios sorted on sectors on the excess market returns ($EMKT$), the spread between long and short Euro bonds ($TERM$), the country-specific excess market return ($CMKT$) and the CMF between January 2003 and October 2006, 197 weeks. The CMF is extracted from returns on defaultable zero-coupon bonds maturing in five years according to the model in Equation (27). The first line corresponds to the lower-maturity portfolio. The Newey–West t-statistics (adjusted for 3 lags) are reported in parentheses.

Figures

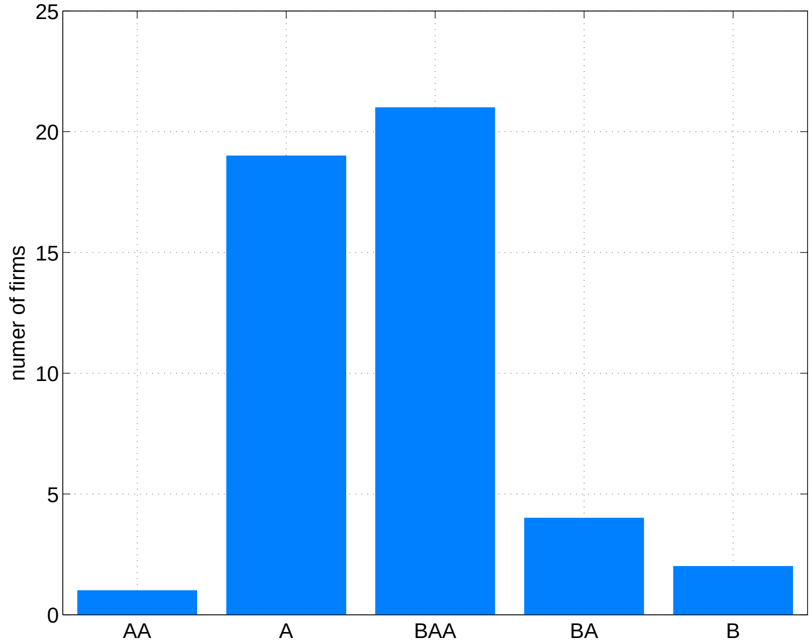

Figure 1: Distribution of firms by median credit rating during the sample period.

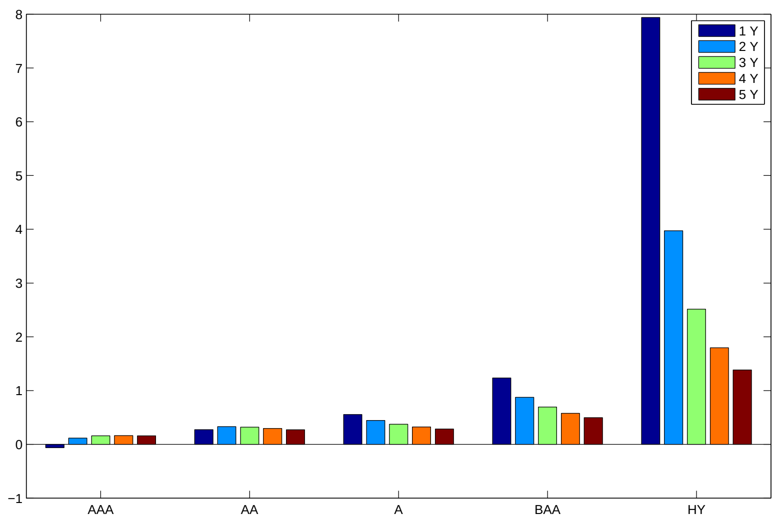

Figure 2: The estimates of the slope coefficient on the CMF extracted from returns on defaultable zero-coupon bonds with maturity varying from one to five years. These slopes are estimated from the time-series regressions of the excess realized returns of five Lehman Brothers Euro-aggregate industrial corporate bond portfolios sorted on rating (AAA, AA, A, BAA and High Yield) on the excess market returns ($EMKT$), the Euro term spread ($TERM$) and the CMF between January 2003 and October 2006, 197 weeks.

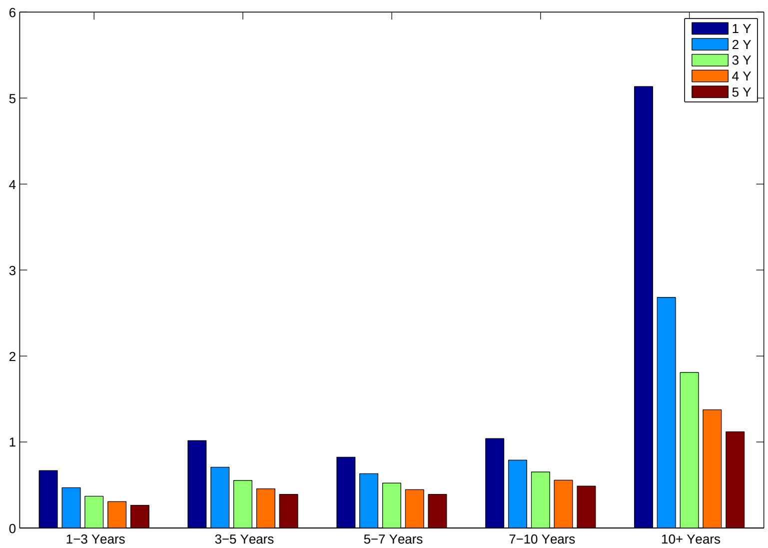

Figure 3: The estimates of the slope coefficient on the CMF extracted from returns on defaultable zero-coupon bonds with maturity varying from one to five years. These slope coefficients are estimated from the time-series regressions of the excess realized returns of five Lehman Brothers Euro-aggregate 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 Euro term spread ($TERM$) and the CMF between January 2003 and October 2006, 197 weeks.

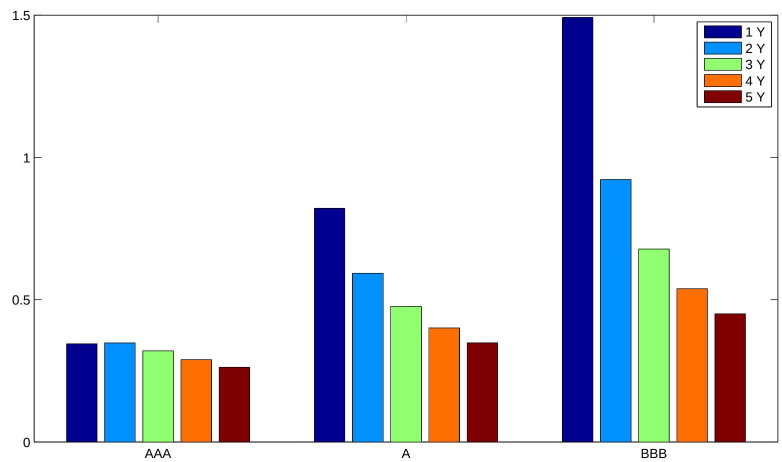

Figure 4: The estimates of the slope coefficient on the CMF extracted from returns on defaultable zero-coupon bonds with maturity varying from one to five years. These slopes are estimated from the time-series regressions of the excess realized returns of three Merrill Lynch nonfinancial corporate bond portfolios sorted on rating (AAA, A and BBB) on the excess market returns ($EMKT$), the Euro term spread ($TERM$) and the CMF between January 2003 and October 2006, 197 weeks.

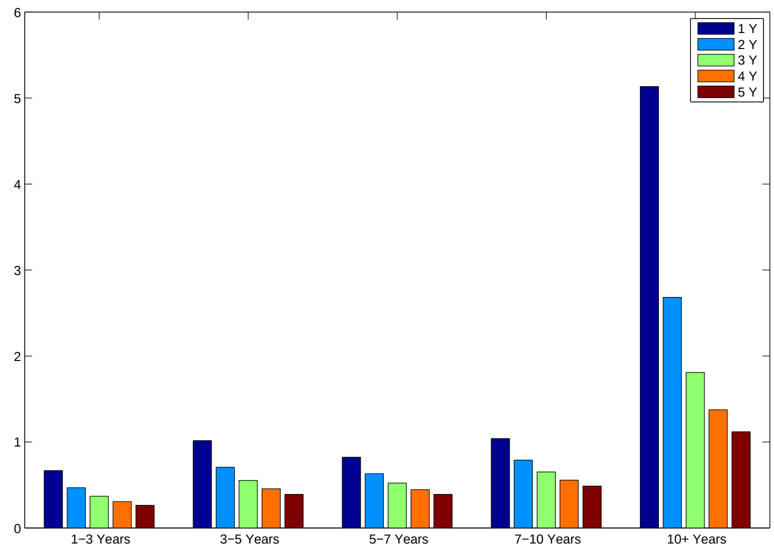

Figure 5: The estimates of the slope coefficient on the CMF extracted from returns on defaultable zero-coupon bonds with maturity varying from one to five years. These slope coefficients are estimated from the time-series regressions of the excess realized returns of four Merrill Lynch nonfinancial corporate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

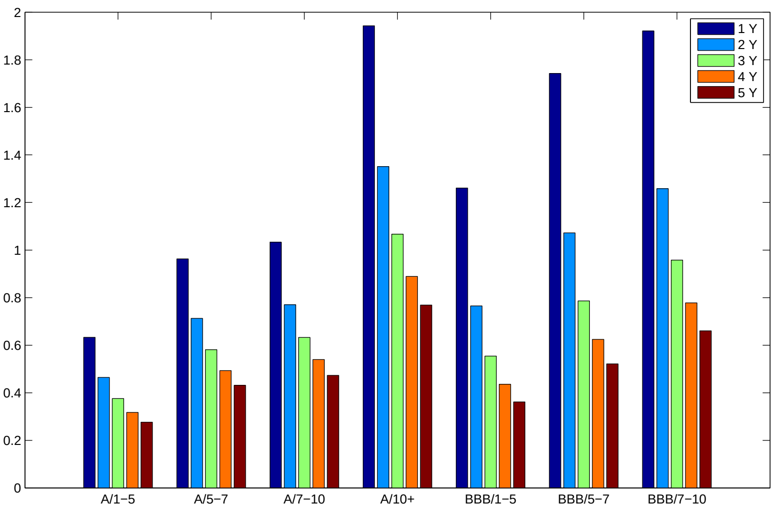

Figure 6: The estimates of the slope coefficient on the CMF extracted from returns on defaultable zero-coupon bonds with maturity varying from one to five years. These slope coefficients are estimated from the time-series regressions of the excess realized returns of four A-rated Merrill Lynch corporate 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 Euro term spread $TERM$ and the CMF between January 2003 and October 2006, 197 weeks.