Default Risk Premia and Asset Returns*

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Abstract

This paper investigates the source for common variation in the portion of returns observed in U.S. credit markets that is not related to changes in risk-free rates or expected default losses. We extract a latent common component from firm-specific changes in default risk premia that is orthogonal to known systematic risk factors during our sample period from 2001 to 2004. Asset pricing tests using returns on Bloomberg-NASD corporate bond indices suggest that our discovered latent changes in default risk premia (DRP) factor is priced in the corporate bond market. A cross-sectional analysis of Merrill Lynch corporate bond portfolios sorted on either industry, maturity or rating supports these findings. In our tests we control for firm characteristics and find that the common variation in changes in default risk premia is not likely to be due to these. Using portfolios of put options written on the S&P 500 index and sorted on moneyness and maturity, we find that, for far-out-of-the-money options, both average returns and the beta estimate for our DRP factor increase with increasing time to maturity. The same holds true for out-of-the-money and at-the-money index put options. There is little to no evidence, however, of the DRP factor being priced in the equity markets. We develop a theoretical framework that, while the DRP factor is part of the pricing kernel, supports our empirical findings. It shows that the DRP factor captures the jump-to-default risk associated with market-wide credit events.

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1 Introduction

Recent empirical studies in financial economics by Berndt, Douglas, Duffie, Ferguson, and Schranz (2005) and Saita (2006) suggest that the compensation demanded by investors for being exposed to credit risk, above and beyond expected default losses, is substantial, and that it varies dramatically over short horizons of time. Berndt, Douglas, Duffie, Ferguson, and Schranz (2005) report that the size of these so called "default risk premia," when measured as a multiple of $1 of expected default loss, ranges from 1.5 to 4, while Saita (2006) reports a range of 1 to 3.5. In terms of variation over time, the first paper shows that default risk premia peaked in the third quarter of 2002, and dropped by roughly 50% until late 2003.

If credit markets are close to being in equilibrium most of the time, any preference-based asset pricing theory will predict that investors demand risk premia on traded assets to compensate for bearing systematic risk. While investor preferences might change over time, it is quite unlikely that they would change dramatically enough over short horizons to induce a time variation in observed default risk premia of the magnitude reported in the above-mentioned study. Alternatively, investors might demand higher compensation for being more exposed to certain systematic factors, which suddenly become more important relative to other systematic factors.

This paper studies to what extent the portion of returns observed in U.S. credit markets that is not related to changes in risk-free rates or expected default losses is a compensation for bearing systematic risk. Towards this goal, we decompose these firm-specific returns into *(i)* a part that is explained by changes in risk-free rates and changes in expected default losses, plus *(ii)* a part that is due to changes in default risk premia. Elton, Gruber, Agrawal, and Mann (2001) show that changes in expected default losses or risk-free rates do not contain information about systematic risk beyond what is already captured by the Fama-French factors (see, Fama and French (1993)).

Motivated by this observation, we focus on that portion of firm-specific returns in credit markets that is due to changes in default risk premia only. We investigate whether there is common time-series variation across the observed firm-specific changes in default risk premia, and test to what extend this common variation can be attributed to a systematic risk factor. Specifically, we aim at answering four main questions:

1. Do firm-specific changes in default risk premia exhibit common time-series vari-

ation?

2. If so, how much of this common time-series variation can be attributed to factors that are known to be priced in either equity or corporate debt markets, such as the Fama and French (1993) equity, treasury and corporate debt factors. How much of it is left unexplained? The latter is identified as the common latent component in changes in firm-specific default risk premia, after controlling for other sources of systematic risk. We will refer to it as the *changes in default risk premia factor*, or short the *DRP factor*.

3. Is the unexplained common time-series variation due to co-movement in time-varying firm characteristics or does it stem from exposure to a common source of risk?

4. If there is support for the later alternative, to what extent is our discovered DRP factor systematic?

Using data from 2001 to 2004 on default swap rates provided by the Markit Group and for Moody's KMV estimates of actual default probabilities for 108 U.S. firms in nine industry groups, we find compelling support for common time-series variation in firm-specific changes in default risk premia. While up to 42% of this co-movement can be due to exposure to other known sources of common variation, a maximum of 35% of the residual is explained by a common latent component, that is, by our DRP factor. Firm characteristics such as the weekly survival probabilities, recovery rates or leverage ratios are mostly unrelated to this common time-series variation. Finally, while we find that our DRP factor is priced in the market for corporate bonds, we find very limited support for a similar conclusion in the equity markets. The test assets employed in the asset pricing tests comprise a wide rage of equity and corporate bond portfolios, formed on various firm characteristics.

Measuring (changes in) default risk premia is not a straightforward task, in part because no pure credit-contingent claims that pay one dollar in the event of no default (survival) and zero otherwise trade in the credit market. Instead, one has to find a way to imply this information from available pricing information on actively traded credit derivatives, such as credit default swaps. This process can be cumbersome, especially because the payoff structure stipulated in these contracts can interact with the default risk itself. In this paper, we use the reduced-form approach of Berndt, Douglas, Duffie, Ferguson, and Schranz (2005) to measure default risk premia, using as pricing information credit default swap (CDS) and recovery rates obtained from Markit and estimates for actual default probabilities provided by Moody's KMV.

To study whether changes in default risk premia exhibit common time-series variation, we first compute firm-specific model-implied returns on constant-maturity credit-sensitive securities that pay one unit of account if no default occurs before maturity and zero otherwise. We then regress, firm by firm, the portion of these returns that is not due to changes in risk-free rates or expected default losses on known systematic factors and time dummies for each week in our sample period. Among the common factors that we account for are those in Fama and French (1993), including their default and term factor, as well as the momentum factor introduced in Jagadeesh and Titman (1993).) We measure our changes in default risk premia factor, at any given time, as the least-squares estimate of the contemporaneous dummy multiplier. Using the time series of the latent common factor identified in the previous step, we then apply the Fama-MacBeth methodology to test whether our DRP factor is priced in asset returns. The test assets comprise well-diversified portfolios of stocks, index options, and corporate bonds, with mean returns spanning a wide range of values. It is important to stress that the DRP factor is extracted from credit market information only. The asset pricing tests will reveal whether it is a risk factor specific to that market, or to what extent it is priced in other markets as well.

Results using Bloomberg-NASD corporate bond indices generated from actual transaction prices of actively traded issues suggest that our discovered latent factor is priced in the corporate bond market. A cross-sectional analysis of the Merrill Lynch corporate bond portfolios, sorted on industry, maturity or rating, supports these findings. For equity markets results are mixed. The DRP factor captures some of the time-series variation in the 100 Fama-French portfolios, sorted on size and book-to-market equity, even after controlling for other potential sources of common variation. The time-series loadings of these portfolios on the DRP factor, however, do not seem to align, cross-sectionally, with the average returns of these portfolio. Finally, we form portfolios using put options written on the S&P 500 index, sorted on moneyness and maturity. We find that, for far-out-of-the-money index put options both average returns and the beta estimate for our DRP factor increase with increasing time to maturity. The same holds true for out-of-the-money and at-the-money index put options.

In order to cope with the possibility that some of the co-movement in changes in default risk premia could be due to reasons other than the common variation in covariances, we also test for firm characteristics such as the firm's default probability and credit rating, the leverage ratio, and recovery rates. We find that the common variation in changes in default risk premia is not likely to be due to these firm characteristics, supporting our main theme that most of the common variation in changes in default risk premia, unaccounted for by other known sources of common variation, is due to firms' exposure to the DRP factor.

We then develop a theoretical framework in which the DRP factor arises naturally in the pricing kernel, and we show that it captures the jump-to-default risk associated with a market-wide credit events. Within this framework, we show that, unlike risk premia on corporate bonds, equity risk premia are only marginally affected by our DRP factor. This results is based on the observation that the DRP factor has a much stronger impact on the returns of assets with a non-degenerate payoff structure in the default states.

The remainder of this paper is structured as follows. Section 2 describes our data, compromised of credit default swap rates for Markit, Moody's KMV EDF estimates for actual default probabilities and other accounting and market price data. Section 3 describes our measure of model-implied returns for constant-maturity zero-coupon corporate bonds, and Section 4 presents our methodology for extracting a latent common factor from the observed firm-specific changes in default risk premia. Section 5 presents our results from the asset pricing tests, and Section 6 proposes a the theoretical framework of the relevant pricing kernel that is consistent with our empirical findings. Finally, Section 7 concludes.

2 Data

This section discusses our data sources for default swap rates, conditional default probabilities, equity and corporate bond returns, and other accounting and balance sheet information.

2.1 Credit Default Swaps

Credit default swaps (CDS) are single-name over-the-counter credit derivatives that provide bond 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.[1] 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.[2] In particular, we use default swap spreads for five-year CDS contracts with modified restructuring (MR) for U.S.-dollar denominated senior unsecured debt. The data is provided by the Markit Group. It contains daily composite CDS spreads calculated from quotes contributed by several banks and default-swap brokers for approximately 2,000 reference entities, incorporated in the U.S. and abroad. One limitation of the Markit data is that it does not report the actual trading volume. Therefore, one concern with these CDS spreads is that no trades might have been actually transacted at the quoted spreads, especially for thinly traded reference entities. In an effort to ensure that our analysis is based on quotes that are representative of actual transaction prices, we exclude from our analysis firms for which we have less than 1000 daily five-year CDS observations. We restrict ourselves to firms incorporated in the US so that our results are not confounded by cross-country differences in bankruptcy laws. As the capital structure of financial institutions is very different from non-financials, we also restrict our sample to non-financial corporate entities. Finally, to minimize market microstructure effects, we only use weekly data.

The sample of default swap rates used in this study consists of 108 entities from nine different industries, based on two-digit SIC codes. The sample period ranges from January 2001 to June 2005. The median firm in our sample has 7 contributors for the five-year CDS spread quote, and has 215 (of a maximum possible 231) valid weekly CDS observations. Figure 1 and Table 1 show the distribution of the 108 firms in our sample by median credit rating during 2001 through 2004, and across rating industries. Different from Berndt, Douglas, Duffie, Ferguson, and Schranz (2005), who use industry-specific but constant loss-given-default values, we rely on

[1]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.

[2]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).

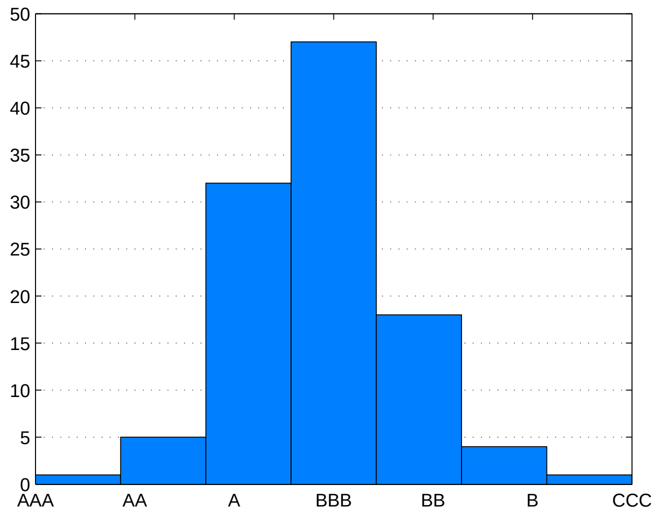

Figure 1: Distribution of firms by median credit rating during the sample period.

contemporaneous recovery rate information from Markit.[3] Table 2 shows the time series of median recovery rates by industry, for each week in our sample period.

2.2 EDF Data

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, Douglas, Duffie, Ferguson, and Schranz (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

[3]Our understanding from conversations with Markit is that the reported recovery rates are indicative of the values used by their sources when valuing CDS rates. In that sense, we treat the recovery information as risk-neutral recoveries.

Table 1: **Distribution of Firms Across Industries:** Firms are grouped into industries according to their two-digit SIC codes.

Industry Name	SIC 2-digit code	No. of Firms
Business Services	73	8
Chemicals and Allied Products	28	20
Communication	48	14
Electric, Gas and Sanitary Services	49	19
Food and Kindred Products	20	10
Industrial Machinery and Equipment	35	10
Instruments and Related Products	38	7
Paper and Allied Products	26	7
Transportation Equipment	37	13
Total	-	108

equity-market capitalization and of the liability 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 Moodys KMV EDF measure is extensively used in the financial services industry. As noted in Berndt, Douglas, Duffie, Ferguson, and Schranz (2005), 40 of the worlds 50 largest financial institutions are subscribers.

We obtain monthly one-year EDF values from Moody's KMV for the time period July 1993 through June 2005, as well as daily observations starting in 2001, for the same set of 108 firms described in Section 2.1. As discussed in Section 2.1, our CDS data only start in January 2001. In order to achieve sufficient power for our asset pricing tests we use weekly (Wednesday) observations of default swap rates, together

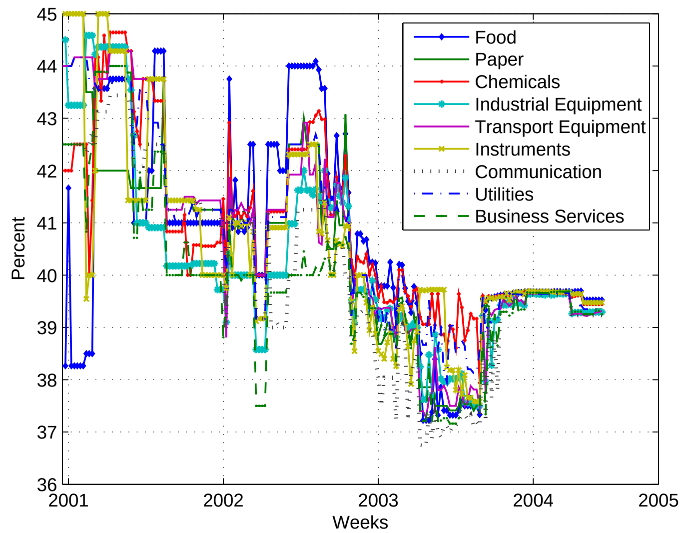

Figure 2: Time series of median recovery rate, by industry. Source: Markit

with EDF values at a weekly frequency. We list the CDS and EDF coverage for each firm in our sample in Table 2 in Appendix A.

Figure 3 shows the median five-year CDS spreads and one-year EDFs across all firms in our sample. Both, CDS spreads and EDF vary considerably through our sample period, peaking during the credit crunch of late 2002 to early 2003.

2.3 Return and Accounting Data

We obtain data on the Fama-French portfolios and factors from Ken French's website. We also use return information for the investment-grade and high-yield Bloomberg-NASD corporate bond indices. These can be downloaded from the NASD website at http://www.nasdbondinfo.com. In addition, total index returns for the Merrill Lynch corporate bond portfolios are from Datastream. Finally, data used to compute firm-specific distance-to-default measures is from COMPUSTAT, and prices of common equity and riskless bond returns are from CRSP.

3 Measuring Returns on Risky Debt

We first describe how we estimate the excess reward investors in the corporate bond market demand for taking on credit risk, after accounting for expected default losses.

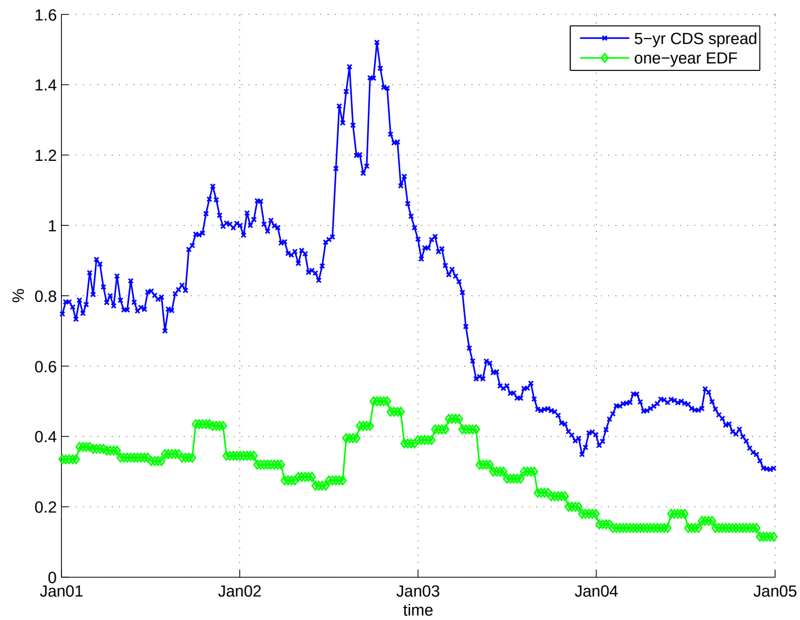

Figure 3: Time series of median five-year CDS rates and median one-year EDFs (in percent) across firms for our sample period. Sources: Markit and Moody's KMV.

Our approach to measuring actual and risk-neutral default probabilities is similar in spirit to Berndt, Douglas, Duffie, Ferguson, and Schranz (2005). The main difference is that instead of using industry-specific but constant loss-given-default values as in their study, we rely on contemporaneous recovery rate information from Markit.[4] Figure 2 in Section 2.1 shows the time series of median recovery rates by industry.

Given a probability space $(\Omega, \mathcal{F}, P)$ and information filtration $\{\mathcal{F}_t : t \geq 0\}$, the default intensity of a firm is the instantaneous mean arrival rate of default, conditional on all current information. More precisely, we suppose that default for a given firm occurs at the first event time of a (non-explosive) counting process N with intensity process λ^P, relative to a given probability space $(\Omega, \mathcal{F}, P)$ and information filtration $\{\mathcal{F}_t : t \geq 0\}$ satisfying the usual conditions. In this case, so long as the firm survives, we say that its default intensity at time t is λ_t^P. Under mild technical conditions, this implies that, 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. We also adopt the relatively standard simplifying doubly-stochastic, or Cox-process, assumption, under which the conditional probability at time t, for a

[4]Our understanding from conversations with Markit is that the reported recovery rates are indicative of the values used by their sources when valuing CDS rates. In that sense, we treat the recovery information as risk-neutral recoveries.

currently surviving obligor, that the obligor survives for some time h, is

$$p(t,h) \quad = \quad E_t\left(e^{-\int_t^{t+h} \lambda_s^P \, ds}\right). \tag{1}$$

Here, E_t denotes expectation conditional on information available up to and including time t.

Under the absence of arbitrage and market frictions, and under mild technical conditions, there exists a "risk-neutral" probability measure, also known as an "equivalent martingale" measure, as shown by Harrison and Kreps (1979) and Delbaen and Schachermayer (1999). In our setting, markets should not be assumed to be complete, so the martingale measure is not unique. This pricing approach nevertheless allows us, under its conditions, to express the price at time t of a security paying some amount, say Z, at some bounded stopping time $\tau > t$, as

$$S_t = E_t^Q\left(e^{-\int_t^{\tau} r_s \, ds}\, Z\right), \tag{2}$$

where r is the short-term interest-rate process[5] and E_t^Q denotes expectation conditional on information available up to and including time t with respect to an equivalent martingale measure Q, that we fix. One may view (2) as the definition of such a measure Q. The idea is that the actual (or physical) measure P and the risk-neutral measure Q differ by an adjustment for risk premia.

For a given firm, Elton, Gruber, Agrawal, and Mann (2001) measure returns in corporate bonds by comparing prices for constant maturity zero-coupon bonds at which they would trade under no-arbitrage assumptions when actually issued by the firm. Prices of these fictive debt securities can be derived, for example, from firm-specific time-series information on corporate bonds or credit default swaps. In (2) we derive the price of a risky bond at time t that pays one unit of account if the firm does survive until time $t+h$, for some h, as $P_{t,h} = E_t^Q \exp(-\int_t^{t+h} r_s + \lambda_s^Q \, ds)$. One length-$h$ time period earlier, that price was $P_{t-h,h} = E_{t-h}^Q \exp(-\int_{t-h}^{t} r_s + \lambda_s^Q \, ds)$. Thus, the *realized return* R_t for constant-maturity h-period zero-coupon bonds issued

[5]Here, r is a progressively measurable process with $\int_0^t |r_s| \, ds < \infty$ for all t, such that there exists a "money-market" trading strategy, allowing investment at any time t of one unit of account, with continual re-investment until any future time T, with a final value of $e^{\int_t^T r_s \, ds}$.

by the firm is given by

$$R_t \;=\; \frac{P_{t,h}}{P_{t-h,h}} = \frac{E_t^Q\left(e^{-\int_t^{t+h} r_s+\lambda_s^Q\,ds}\right)}{E_t^Q\left(e^{-\int_{t-h}^{t} r_s+\lambda_s^Q\,ds}\right)}.$$

Under the simplifying assumption that over small time intervals h the short rate r_t stays relatively flat,[6] the last equation reduces to

$$R_t \;=\; e^{-(r_t-r_{t-h})h}\,\frac{p^Q(t,h)}{p^Q(t-h,h)}, \tag{3}$$

where $p^Q(t,h)$ is the risk-neutral survival probability defined by

$$p^Q(t,h) \;=\; E_t^Q\left(e^{-\int_t^{t+h}\lambda_s^Q\,ds}\right).$$

Empirical studies of predictability of changes in credit spreads as measured by (3) have shown that structural model variables that should in theory have large explanatory power perform rather poorly (see, for example, Collin-Dufresne, Goldstein, and Martin (2001)), and that changes in expected default losses do not contain information about systematic risk beyond the information already captured by Fama-French factors (see, for example, Elton, Gruber, Agrawal, and Mann (2001)). Based on the later observation, we will focus on that portion R_t^u of the return R_t in (3) that is not due to changes in expected default losses or changes in risk-free rates. In particular, if investors where risk-neutral, the realized return on the constant-maturity h-period zero-coupon bonds would equal $R_t^P = e^{-(r_t-r_{t-h})h}\,\frac{p(t,h)}{p(t-h,h)}$, for h small. Subtracting R_t^P from (3) yields

$$R_t^u \;=\; e^{-(r_t-r_{t-h})h}\left[\frac{p^Q(t,h)}{p^Q(t-h,h)} - \frac{p(t,h)}{p(t-h,h)}\right]. \tag{4}$$

For small intervals h, R_t^u will be roughly equal to $-[(\lambda_t^Q-\lambda_{t-h}^Q)-(\lambda_t^P-\lambda_{t-h}^P)]h$.[7] $(\lambda_t^Q-\lambda_t^P)$ measures the difference between instantaneous risk-neutral and actual default probabilities, and can be interpreted as a measure of default risk premia. One may think, therefore, of the annualized unexplained returns R_t^u/h as minus the changes in default risk premia

If constant-maturity zero-coupon bonds as described above were actively traded,

[6]For the time series estimation in Section 4 and the asset pricing tests in Section 5 we will use a time interval h of one week.

[7]For the time series estimation in Section 4 and the asset pricing tests in Section 5 we will use a time interval h of one week.

we could observe prices $P_{t,h}$ directly, and it would be possible to compute returns on corporate debt using (3). As this is not the case, however, we proceed by estimating time-series models for λ^P and λ^Q from different sources, and then compute model-implied actual and risk-neutral survival probabilities $p(t,h)$ and $p^Q(t,h)$. We follow Berndt, Douglas, Duffie, Ferguson, and Schranz (2005) and identify the default intensity λ^P under the physical measure from the information contained in the Moody's KMV EDFs, while λ^Q will be estimated from default swap data. Details on the choice of the time-series models for λ^P and λ^Q and our estimation technique are discussed in Section 4.

Afterwards, we will extract the latent common component, that is, our default risk premia factor, from these firm-specific unexplained returns R_t^u. We describe our approach in Section 4. One may think of our DRP factor as a measure of realized excess returns of a common risk factor embedded in the default events across firms.

4 Extracting the Default Risk Premia Factor

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

$$d\log(\lambda_t^P) \quad = \quad \kappa(\theta - \log(\lambda_t^P))\,dt + \sigma\,dB_t, \tag{5}$$

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 majority of the 108 firms in our sample, we have 144 months of one-year EDF observations, from July 1993 to June 2005.

Given the log-autoregressive form (5) of the default intensity, in general there is no closed-form solution available for the one-year EDF, $1 - p(t,1)$, from (1). 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

$$d\log \lambda_t^Q = \kappa^Q(\theta^Q - \log(\lambda_t^Q))\, dt + \sigma^Q\, dB_t^Q, \qquad (6)$$

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

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

where $\tilde{\kappa}^Q$ and $\tilde{\theta}^Q$ are constants and $\tilde{B}_t^Q$ 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, Douglas, Duffie, Ferguson, and Schranz (2005).

Using maximum likelihood estimation (MLE), we obtain firm-by-firm estimates for the parameters that govern the processes for λ^P and λ^Q. These parameter estimates are listed in Table 3 and 4 in Appendix B. The estimation techniques employed here are similar to those used in Berndt, Douglas, Duffie, Ferguson, and Schranz (2005), except for the fact that Markit provides us with information on contemporaneous recovery rates, for each firm and each date, that we will use in place of an assumption of constant risk-neutral recovery in the event of default. As can be seen in Figure 2 in Section 2.1, median (risk-neutral) recovery rates range between 37% and 45%.

Figure 4 plots the time series of the median differences between estimated instantaneous risk-neutral default probabilities, λ_t^Q, and estimates for instantaneous actual default probabilities, λ_t^P, for each industry identified by its two-digit SIC code. Interpreting $\lambda_t^Q - \lambda_t^P$ as a measure of default risk premia in the corporate bond market, we find that it peaked quite dramatically for the Communication industry in the third quarter of 2002, and that it surged for both the Utilities and the Paper sector later that year.

Using our estimates for λ_t^Q and λ_t^P, we can now compute estimates for the unexplained part R_t^u of realized returns on constant-maturity zero-coupon bonds as given in (4) in Section 3. We will denote by F_t^D the time-t level of the latent common component to be extracted from firm-specific unexplained returns R_t^u. We refer to F^D as the DRP factor. As explained in Section 3, F^D captures the common variation in changes in default risk premia, and one may think of it as a measure of realized excess returns of a common risk factor embedded in the default events across firms.

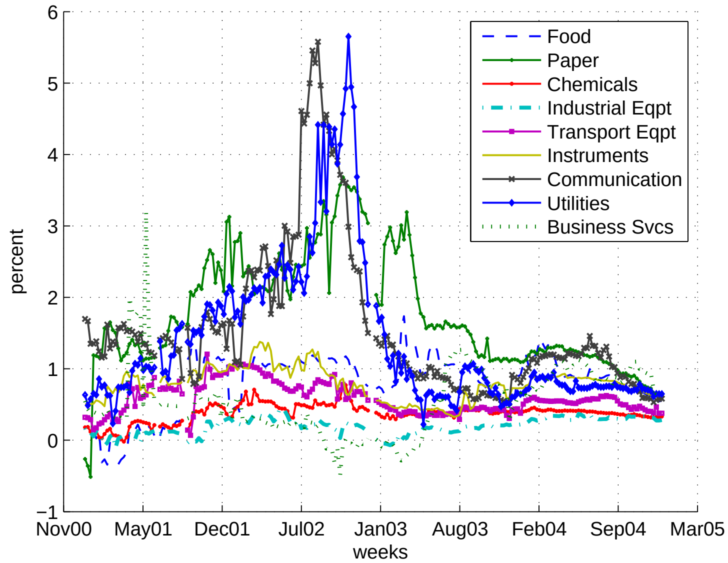

Figure 4: Time series of median default risk premia, $\lambda_t^Q - \lambda_t^P$, by industry.

In addition, let F_t^S denote the vector of h-period returns on known systematic factors. Among the factors we account for are those in Fama and French (1993), including their default and term factor, and the momentum factor introduced in Jagadeesh and Titman (1993).[8] Using superscript i to indicate returns specific to firm i, for $i = 1, \ldots, N$, we run the following least-squares regression model with firm- and time-specific effects on the panel data of unexplained excess returns:

$$R_t^{u,i} \quad = \quad \alpha^i + \beta^{S,i} \cdot F_t^S + \sum_{\text{weeks } \bar{t}} \delta_{\bar{t}} 1_{\{t=\bar{t}\}} + \epsilon_t^i. \tag{7}$$

For each firm i, the errors ϵ_t^i have a sample mean of zero across time. They will absorb any variation in default risk premia that cannot be explained by linear combinations of systematic factors F^S and a single latent common component. In order to identify all unknown parameters $\{\alpha^i\}$, $\{\beta^{S,i}\}$, and $\{\delta_t\}$ we have to impose two normalizing restrictions: *(i)* the sample mean of δ_t is zero, and *(ii)* the sample correlations between δ_t and each of the systematic factors in F_t^S are zero. We measure our default risk

[8]We run all tests with and without the default factor. Results do not change substantially, and we only report them for the scenario where the default factor is included as a known systematic factor.

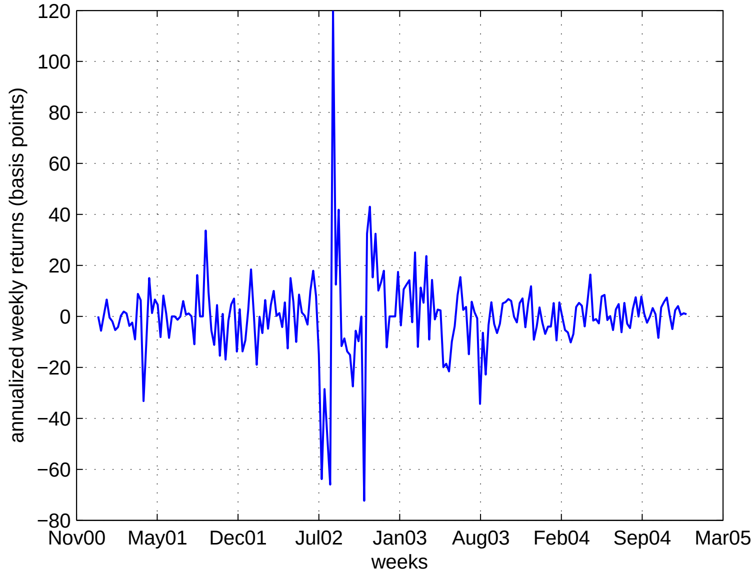

Figure 5: Time series of the latent DRP factor F^D as extracted from unexplained weekly returns on constant weekly maturity zero-coupon bonds.

premia (or DRP) factor F_t^D as

$$F_t^D = \frac{1}{N} \sum_{i=1,\ldots,N} \hat{\alpha}^i + \hat{\delta}_t,$$

where $\hat{\delta}_t$ denotes the least-squares estimate of the time-t dummy multiplier in (7). Figure 5 shows the time series of the extracted latent factor F^D. Note that most of the time variation in the DRP factor F_t^D occurs during the second part of 2002.

In Section 5 we use the time series of returns of the latent common factor, F_t^D, and employ the Fama-MacBeth methodology to test whether this factor is priced in the cross-section of corporate bond and stock returns. In what follow, DRP factor will refer to F^D extracted from returns on constant-maturity zero-coupon bonds as in (7).

5 Asset Pricing Tests

In this section we investigate whether the common variation in changes in default risk premia induced by our DRP factor, F_t^D, is systematic in nature. In other words, we test whether this factor is priced in the cross-section of stock and corporate bond returns. Our asset pricing test is a variant of the Fama-MacBeth methodology (see Fama and MacBeth (1973)). Specifically, we consider a set of test assets and investigate whether their loadings on known systematic factors and our DRP factor have explanatory power for the cross-section of average returns. Among the factors we account for are those in Fama and French (1993), including their default and term factor, and the momentum factor introduced in Jagadeesh and Titman (1993).

The ideal test assets should have two important features: *(i)* they should span the entire spectrum the capital markets, and *(ii)* the test assets should exhibit a high degree of variation in average returns. The first condition is important in defining the generality of our test, while the second feature ensures that the cross-section of expected returns is sufficiently rich. As a compromise between meeting these conditions and data availability, we decided on a set of test assets which consists of the 100 Fama-French portfolios formed on size and book-to-market equity, the two Bloomberg-NASD investment-grade and high-yield corporate bond portfolios, and several Merrill Lynch corporate bond portfolios, sorted by rating (7 portfolios ranging from AAA to C), by maturity (6 portfolios ranging from 1 to 3 years (1-3Y) to more than 15 years (15Y+), and by industry (30 high-yield portfolios, and 4 investment-grade portfolios).

Due to the fact that the last three Fama-French equity portfolio have missing data for the first part of the sample, we drop them from our set of test assets and, for the rest of the paper, we run our asset pricing test on 97 equity portfolios and 49 corporate bond portfolios.

Our asset pricing tests proceed in two stages. In the first step we determine the loadings of each of the test assets on explanatory factors such as the Fama and French's stock market factors RMO, SMB and HML, the momentum factor UMD, the corporate debt market factor DEF, the treasury bond market factor $TERM$, as well as our default risk premia factor DRP. (The DRP factor is equal to F^D as defined in (8).) More formally, we estimate the linear model

$$\begin{aligned} R^i(t) - RF(t) &= \alpha^i + \beta^i_{RMO} RMO(t) + \beta^i_{SMB} SMB(t) + \beta^i_{HML} HML(t) \\ &+ \beta^i_{UMD} UMD(t) + \beta^i_{DEF} DEF(t) + \beta^i_{TERM} TERM(t) \\ &+ \beta^i_{DRP} DRP(t) + \epsilon^i(t) \end{aligned} \tag{8}$$

for each test asset i. Here, $R^i(t)$ denotes the return on asset i over the time period $[t-h, t]$, where h is one week, and $RF(t)$ measures the risk-free rate, compounded weekly from the Fama-French T-bill daily returns.

Second, we investigate the explanatory power of the estimated loadings for the cross-section of returns of the test assets. Specifically, for every week t, we run a cross-sectional regression

$$
\begin{aligned}
R(t)^i - RF(t) &= \gamma_0(t) + \gamma^{RMO}(t)\beta^i_{RMO} + \gamma^{SMB}(t)\beta^i_{SMB} \\
&+ \gamma^{HML}(t)\beta^i_{HML} + \gamma^{UMD}(t)\beta^i_{UMD} + \gamma^{DEF}(t)\beta^i_{DEF} \\
&+ \gamma^{TERM}(t)\beta^i_{TERM} + \gamma^{DRP}(t)\beta^i_{DRP} + \epsilon^i(t). \qquad (9)
\end{aligned}
$$

Notice that, under the usual OLS assumption, the unconditional version of this regression gives the very asset pricing restriction which we investigate here. In particular, we have

$$
\begin{aligned}
E\left[R(t)^i - RF(t)\right] &= E\left[\gamma_0(t)\right] + E\left[\gamma^{RMO}(t)\right]\beta^i_{RMO} + E\left[\gamma^{SMB}(t)\right]\beta^i_{SMB} \\
&+ E\left[\gamma^{HML}(t)\right]\beta^i_{HML} + E\left[\gamma^{UMD}(t)\right]\beta^i_{UMD} + E\left[\gamma^{DEF}(t)\right]\beta^i_{DEF} \\
&+ E\left[\gamma^{TERM}(t)\right]\beta^i_{TERM} + E\left[\gamma^{DRP}(t)\right]\beta^i_{DRP}.
\end{aligned}
$$

5.1 Corporate Bonds

We test first whether the default risk premia factor DRP is priced in the market for corporate bonds. As mentioned above, whenever possible, we implement the two-step Fama-MacBeth procedure, which involves estimating regressions of type (8) and (9).

Table 5 summarizes the results of the time-series regression (8), for the two NASD corporate bond portfolios. For comparison we provide the results of a similar regression, but without the DRP factor (the standard Fama and French (1993) regression, adapted for our sample period). We notice that, after controlling for other known systematic effects, both the high-yield and the investment-grade portfolios load economically and statistically significant on our default risk premia factor. Moreover, the loading for the high-yield portfolio is several times larger than the corresponding loading for the investment-grade portfolio. Given that the average annual return on the high-yield portfolio (17.46 percent) is larger than the average annual return on the investment-grade portfolio (2.93 percent), the results of Table 3 suggest that, in the market for corporate bonds, higher expected return could be compensation for bearing default risk, as proxied by our default risk premia factor. It is of interest to note that inclusion of the DRP factor in the time series regression leads to a decrease in α by roughly 1% a year for the high-yield corporate bond portfolio. It has a much smaller effect on the intercept for investment-grade debt. So even though DRP helps explaining the unconditional mean excess returns of the two bond portfolios, the intercepts might still appear relatively high. A potential explanation can be due to the fact that none of the explanatory variables considered here captures the high illiquidity effects in the corporate bond markets, as documented for example by Driessen (2005).[9]

We further test the strength of this relation, by investigating the time-series behavior of corporate bond portfolios, sorted on specific characteristics or on sector and industry. Table 6 reports the estimates of the time-series regression (8), when the test assets are corporate bond portfolios, sorted on rating. Here, we notice that the loading on the DRP factor is highly correlated with the rating, which in turn is highly correlated with the average return of the portfolios. Thus, even after controlling for other potential systematic factors, the expected returns of the portfolios and the loadings on the DRP factor are still strongly positively related. Table 7 repeats the exercise for the case where the test assets are corporate bond portfolios sorted on time to maturity. The results here support our findings in the previous two tables, namely that higher loadings on the DRP factor translate into higher expected returns, even after controlling for other potential systematic factors.

The conjectured relation between average returns and the loadings on the default risk premia factor is further tested in Table 8 and 9, where the test assets consist of high-yield corporate bond portfolios, sorted by sector. (Table 10 repeats the analysis of industry portfolios for investment-grade debt.) Table 8 shows that, on average, the default risk premia factor is significant for the time-series variation of expected bond returns. More importantly, Table 9 shows that measures of risk based on our DRP factor (that is, the time-series loadings on the factor) are relevant for the cross-section of corporate bond returns. Once again, our results suggest that the default risk premia factor is priced in the market for corporate bonds.

5.2 Equity

We now turn our focus to the equity market. Table 11 shows that our DRP factor contributes only little to the time-series variation of the equity portfolios, while other potential systematic factors account for about 82 percent, on average. In addition, the loadings on the Fama-French factors are very significant and economically consistent with the findings of Fama and French (1993), for their sample. The loadings on our DRP factor are significant, on average, but as it is apparent from Table 12 they do not

[9]We are in the process of acquiring data which will allows us to implement some of the liquidity measures used in the literature.

matter for the cross-section of stock returns. We replicated this experiment for a wide range of test assets originating from the equity market, including the Fama-French 49 industry portfolios, the decile portfolios formed on the book-to-market equity, and others. Up to this point, none of these tests were successful in detecting a significant relation between average stock returns and the loadings on the default risk premia factor.

Overall, our results so far suggest that the common variation in firm-specific changes in default risk premia is priced in the corporate bond market. It is important to stress the fact that the risk behind the common default component DRP is completely uncorrelated with the risk behind previously known systematic factors. (Recall that our default risk premia factor defined in (8) in Section 4 is by design orthogonal to all components in F^S. This orthogonality statement holds in sample. To generalize, one needs to assume that the cross-section of firms from which we extract the default risk premia factor is sufficiently large.)

5.3 Options

In this section, we test whether our DRP factor is priced in the options market. Motivated by the findings in Coval and Shumway (2001) and Jones (2006) that a significant portion of returns on short-term out-of-the-money index put options cannot be explained by known systematic risk factors, we form test portfolios using put options written on the S&P 500 index, sorted on moneyness and maturity. Moneyness is defined as the present value of the strike price, computed using the maturity-matched risk-free rate, divided by the current value of the S&P 500 index. We then use these portfolios to examine whether the DRP factor makes a contribution to explaining the cross-sectional variation in index put options returns.

The options data is obtained from OptionMetrics. In what follows we describe in more detail how we form the options portfolios. First, we classify the options into four maturity bins, with times to maturity of 10 to 30 days, 31 to 60 days, 61 to 150 days, and more than 150 days. We don't use options with fewer than 10 days to maturity since reported prices of these options are more likely to be erroneous. Next, we split each maturity bin into three sub-bins based on moneyness. For these sub-bins, we choose the moneyness cutoff points so that within each maturity bin, the moneyness bins have approximately the same number of observations. This results in 12 portfolios sorted first on time to maturity and then on moneyness. Each week (Wednesday) t, we assign each option to a particular bin based on its maturity and moneyness as that time, and compute its returns to week $t+1$. That is, returns calculated are weekly returns. The return of any particular maturity-moneyness portfolio at time $t+1$ is then the average of the buy-and-hold returns between from week t to week $t+1$ of all options that were in this particular maturity-moneyness bin as of time t. We compute both, equally-weighted and value weighted returns. Prices are computed as the average of the best bid and best offer price on a given day. The value-weighted returns use prices as of time t for the weights. Summary statistics for the 12 put option portfolios are reported in Table 13 in the appendix.

We restrict ourselves to options with standard settlement. To eliminate prices with large errors, we only use observations that satisfy the following criteria. The trading volume is positive, both bid and offer prices are positive, offer prices are at least as high as bid prices, the open interest is positive, the sum of the mid price plus the bid-ask spread is at least as high as the intrinsic value of the option, and the reported implied volatility is at least 1%.

The intrinsic value is calculated as the larger of the present value of dividends plus the present value of the strike price minus the S&P index closing value, and zero. Under no-arbitrage assumptions, the price of the put option should exceed its intrinsic value. To allow for non-synchronous reporting of the value of the underlying and the option, we use a somewhat looser constraint, enforcing only that the price plus spread exceeds the intrinsic value. Following Jones (2006), we also use an implied volatility cutoff to remove options prices that are suspect.

We then conduct asset pricing tests using these 12 options portfolios. We run time-series regressions of the returns of these portfolios, in excess of the risk-free rate, on the excess returns on the market, the SMB, HML, UMD, DEF and TERM factors, as well as the volatility index factor VIX and our DRP factor. We account for VIX to capture systematic volatility risk premia. The results are shown in Table 14 in Appendix D. We find that, for far-out-off-the-money index put options (moneyness bin 1), both average returns and the beta estimate for our DRP factor increase with increasing time to maturity. The same holds true for out-of-the-money (moneyness bin 2) and at-the-money (moneyness bin 1) put options.

5.4 Test for Firm Characteristics

We conclude this section with a test based on firm characteristics. Following the argument in Daniel and Titman (1997), we study the extend to which the common component in changes in default risk premia is due to firm characteristics, which may behave very similarly, across firms, over time. Specifically, we investigate the possibility that the time variation in changes in default risk premia may be driven solely by certain firm characteristics, say $\theta(t)$. Formally, we test whether a linear model of the form

$$\begin{aligned} R^{u,i}(t) &= \alpha^i + \beta^i_{RMO} RMO(t) + \beta^i_{SMB} SMB(t) + \beta^i_{HML} HML(t) \\ &+ \beta^i_{UMD} UMD(t) + \beta^i_{DEF} DEF(t) + \beta^i_{TERM} TERM(t) \\ &+ \beta^i_{DRP} DRP(t) + \beta^i_{Char} \theta^i(t-1) + \epsilon^i(t) \end{aligned} \tag{10}$$

can be justified in this context. On the left-hand side of (10) we have the unexplained returns on constant-maturity zero-coupon bonds of firm i, $R^u_{t,i}$, as defined in (4) in Section 3, while on the right-hand side we have the usual factors plus a time-varying firm characteristics. Notice that the time-varying (or conditional) changes in default risk premia depend on the time-varying characteristic $\theta(t-1)$. Thus, if the firm characteristics move together, it would appear as if unexplained returns $R^{u,i}_t$ move together.

Table 15 reports the results of these regressions for firm characteristics such as the firm's default probability (or credit rating), the leverage ratio and the recovery rate. We notice that the common variation in changes in default risk premia is very unlikely to be due to these firm characteristics. Moreover, in each of the tests, the loading on the DRP factor is always very significant, both economically and statistically.

6 A Model Framework Explaining Our Results

In this section, we by propose a theoretical framework of corporate default that is consistent with our findings. We consider an economy with N firms, in which the fundamentals are captured by a vector of d-dimensional state variables, X_t, with dynamics given by

$$dX_t = \mu(X_t, t)dt + \Sigma(X_t, t)dW_t,$$

where $\mu(\cdot, t)$ is a d-dimensional column vector of drifts and $\Sigma(\cdot, t)$ is a $d \times d$ state-dependent volatility matrix. Here, W_t is a d-dimensional standard Brownian motion on some probabilistic space $(\Omega, \mathbb{P})$ with informational filtration $\{\mathcal{F}_t\}_{t \geq 0}$ generated by this process. The innovations dW_t in W_t describe the diffusive systematic risk in our economy.

We define $N+1$ stopping times

$$\bar{\tau}^i = \inf_{t \geq 0} \int_0^t \bar{\lambda}^{P,i}(X_s)\, ds \geq \theta^i, \text{ for } i = 0, 1, \ldots, N,$$

where, for all i, $\bar{\lambda}_t^{P,i} = \bar{\lambda}^{P,i}(X_t)$ is a non-negative random variable, and θ^i is an exponentially distributed random variable with mean one. One can interpret $\bar{\tau}^0$ as the arrival time of a market-wide credit event that affects all firms in the economy, whereas $\bar{\tau}^i$ denotes the event time of an idiosyncratic default event for firm i. Let $\{\mathcal{G}_t\}_{t \geq 0}$ denote the extended filtration, generated by the state vector X_t and the random stopping times $\{\bar{\tau}^i\}_{0 \leq i \leq N}$. Finally, let $\tilde{N}_t^i = \mathbf{1}_{\{\bar{\tau}^i < t\}}$ denote the counting processes associated with stopping time $\bar{\tau}^i$.

The following assumptions are key ingredients to the tractability of our model.

Assumption 1

(i) The joint informational content of the random variables $\{\theta^i\}_{0 \leq i \leq N}$ is independent from the information contained in $\mathcal{F}_\infty$.

(ii) The functions $\{\lambda^i\}_{0 \leq i \leq N}$ are chosen such that $N_t^i = \tilde{N}_t^i - \int_0^t \lambda^{P,i}(X_s)\, ds$ are martingales with respect to the filtration $\mathcal{G}_t$, for any $i = 0, 1, \ldots, N$.

(iii) The default event of each firm i, $i = 1, \ldots, N$, is triggered by an exogenous random variable τ^i given by

$$\tau^i = \min\left\{\bar{\tau}^0, \bar{\tau}^i\right\}. \tag{11}$$

Define $\bar{\lambda}_t^P = \bar{\lambda}_t^{P,1} + \ldots + \bar{\lambda}_t^{P,N}$, and let

$$\bar{\Gamma}_t = \frac{1}{\bar{\lambda}_t^P} \sum_{i=1}^N \bar{\lambda}_t^{P,i}\, \Gamma^i(X_t),$$

where $\Gamma_t^i = \Gamma^i(X_t)$ is the market price of jump-to-default risk associated with the idiosyncratic default event $\bar{\tau}^i$, for all i. Suppose that the relevant pricing kernel M_t for this economy is given by

$$\begin{aligned}
\frac{dM_t}{M_t} &= -r(X_t)\, dt - \Lambda(X_t)\, dW_t - \sum_{i=0}^N \Gamma^i(X_{t-})\, (dN_t^i - \bar{\lambda}_t^{P,i}\, dt) \\
&= -\left(r(X_t) + \Gamma_t^0 \bar{\lambda}_t^{P,0} + \bar{\Gamma}_t \bar{\lambda}_t^P\right) dt - \Lambda(X_t)\, dW_t - \sum_{i=0}^N \Gamma^i(X_{t-})\, dN_t^i
\end{aligned} \tag{12}$$

where $r_t = r(X_t)$ is the instantaneous risk-free interest rate, $\Lambda_t = \Lambda(X_t)$ denotes the market price of diffusive risk, and $\Gamma_t^0 = \Gamma^0(X_t)$ is the market price of jump-to-default risk associated with the market-wide credit event time $\bar{\tau}^0$.

As long as the functionals Λ, Γ^0, $\{\Gamma^i\}_{1\leq i\leq N}$ are well-behaved (e.g. bounded), the above equation admits a unique solution[10]

$$\begin{aligned} M_t &= \exp\left\{-\int_0^t r_s ds + \int_0^t \bar{\Gamma}_s \bar{\lambda}_s^P ds - \frac{1}{2}\int_0^t \Lambda_s^2 ds - \int_0^t \Lambda_s dW_s\right\} \\ &\quad \times \exp\left\{\int_0^t \Gamma_s^0 \bar{\lambda}_s^{P,0} ds\right\} \times \prod_{0\leq s<t}\left[1 - \sum_{i=0}^{N} \Gamma_s^i \Delta\tilde{N}_s^i\right], \end{aligned} \tag{13}$$

where $\Delta\tilde{N}_s^i = \tilde{N}_s^i - \tilde{N}_{s-}^i$, for all $i = 0, 1, \ldots, N$. Note that $\exp(\int_0^t r_s\, ds) M_t$ is a martingale with respect to $\mathcal{G}_t$.

Let Ω_0 denote the subset of Ω for which at least two of the counting processes $\tilde{N}_t^i$, $i = 0, 1, \ldots, N$, jump at the same time (that is, $\Delta\tilde{N}^i\, \Delta\tilde{N}^j = 1$ for some $0 \leq i \neq j \leq N$). Consider the probability space $(\Omega - \Omega_0, \mathcal{G}_t, \mathbb{P}|_{\Omega-\Omega_0})$. The following assumption allows us to generalize the results in Dai and Singleton (2003), who consider a similar environment with one firm only, within the framework of our model.

Assumption 2 *The random variables θ_i, $i = 0, 1, \ldots, N$, are approximatively exponentially distributed with mean one, on the reduced space $\Omega - \Omega_0$.*

While this assumption might seem somewhat restrictive, it is not unjustified, given that for our specifications of $\lambda^{P,i}$, the probability that two or more of the count processes $\tilde{N}^i$ will jump at the same time is very small. For the rest of the paper we will focus only on the reduced space $(\Omega - \Omega_0, \mathcal{G}_t, P|_{\Omega-\Omega_0})$. To simplify notation, we re-denote this space as the new $(\Omega, \mathcal{G}_t, P)$. Ruling our simultaneous event times implies that the actual default intensity $\lambda^{P,i}$ for firm i introduced in (5) in Section 4 can be though of as $\lambda^{P,i} = \bar{\lambda}^{P,0} + \bar{\lambda}^{P,i}$.

Under Assumption 2, (13) can be rewritten in a convenient way, by noticing that $\Delta\tilde{N}^i\, \Delta\tilde{N}^j = 0$ for any $0 \leq i \neq j \leq N$, as[11]

$$\begin{aligned} M_t &= \exp\left\{-\int_0^t r_s ds + \int_0^t \bar{\Gamma}_s \bar{\lambda}_s^P ds - \frac{1}{2}\int_0^t \Lambda_s^2 ds - \int_0^t \Lambda_s dW_s\right\} \\ &\quad \times \exp\left\{\int_0^t \Gamma_s^0 \bar{\lambda}_s^{P,0} ds\right\} \times \prod_{i=0}^{N} \prod_{0\leq s<t}\left[1 - \Gamma_s^i \Delta\tilde{N}_s^i\right]. \end{aligned}$$

[10] For details, see Protter (2005), page 84, Theorem 37.

[11] When this is the case, we have $1 - \sum_{i=0}^N \Gamma_s^i \Delta\tilde{N}_s^i = \Pi_{i=0}^N \left[1 - \Gamma_s^i \Delta\tilde{N}_s^i\right]$.

The fact that $\exp(\int_0^t r_s\,ds)M_t$ is a martingale allows us to construct an equivalent (risk-neutral) martingale measure Q on Ω from the Radon-Nikódym density $\frac{dQ}{dP}_{|t} = \exp(\int_0^t r_s\,ds)M_t$. Note that under this risk-neutral measure, $N_t^{Q,i} = \tilde{N}_t^{Q,i} - \int_0^t \bar{\lambda}^{Q,i}(X_s)\,ds$ become martingales. Here, $\tilde{N}_t^{Q,i}$ is defined as before after replacing $\bar{\lambda}^{P,i}$ by $\bar{\lambda}^{Q,i} = (1-\Gamma^i)\bar{\lambda}^{P,i}$, for all $i = 0, 1, \ldots, N$.

For for each firm i, actual and risk-neutral default probabilities for the interval from time t to $t+h$, conditioned on no default before and at time t, can be computed as $E_t^P e^{-\int_t^{t+h}\left[\bar{\lambda}_s^{P,0}+\bar{\lambda}_s^{P,i}\right]ds}$ and $E_t^Q e^{-\int_t^{t+h}\left[\bar{\lambda}_s^{Q,0}+\bar{\lambda}_s^{Q,i}\right]ds}$, respectively. In the notation of (1) and (4) of Section 3, we have, for small values of h, $p^i(t,h) = 1 - \left[\bar{\lambda}_t^{P,0} + \bar{\lambda}_t^{P,i}\right]h$ and $p^{Q,i}(t,h) = 1 - \left[\bar{\lambda}_t^{Q,0} + \bar{\lambda}_t^{Q,i}\right]h$. Hence, the unexplained return $R_t^{i,u}$ for firm i is approximately equal to

$$R_t^{i,u} \;=\; \left[\Gamma_t^0\bar{\lambda}_t^{P,0} - \Gamma_{t-h}^0\bar{\lambda}_{t-h}^{P,0}\right]h + \left[\Gamma_t^i\bar{\lambda}_t^{P,i} - \Gamma_{t-h}^i\bar{\lambda}_{t-h}^{P,i}\right]h. \tag{14}$$

Equation (14) suggests that the common component of the returns $R^{i,u}$ is to a large extent driven by the changes in the market price of jump-to-default risk associated with the market-wide default event $\bar{\tau}^0$. Thus, our default risk premia factor F^D is likely to capture the impact on returns due to this market-wide source for jump-to-default risk.

We now investigate the effect of the diffusive risk and of the jump-to-default risk on the expected returns of a firm's equity and debt claims. If the markets for both equity and debt are competitive, the pricing equation is the Euler equation. That is,

$$E_t\left[\frac{M_{t+h}}{M_t}\tilde{R}_{t+h}^i\right] \;=\; 1 \tag{15}$$

where $\tilde{R}_{t+h}^i$ denotes the gross return on either the equity or the debt of firm i.

As long as a firm is solvent, the gross return on equity claims is non-zero. We will assume a zero-recovery value to equity holders in the event of default. This implies $\tilde{R}_{t+h}^{E,i} = \tilde{R}_{t+h}^{S,i}\mathbf{1}_{\{\tau^i > t+h\}}$. Here, $\tilde{R}_{t+h}^{S,i}$ stands for the total return on firm i's equity if the company does not default prior to or at time $t+h$. With this in mind, we have

$$\begin{aligned}
E_t\left[\frac{M_{t+h}}{M_t}\tilde{R}^{E,i}_{t+h}\right] &= E_t\left[\mathcal{E}_{t+h}\prod_{i=0}^{N}\prod_{t\le s<t+h}\left[1-\Gamma^i_s\Delta\tilde{N}^i_s\right]\,\tilde{R}^{S,i}_{t+h}\,\mathbf{1}_{\{\tau^i>t+h\}}\right] \\
&= E_t\left[\mathcal{E}_{t+h}\prod_{j\neq 0,i}\prod_{t\le s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right]\,\tilde{R}^{S,i}_{t+h}\,\mathbf{1}_{\{\tau^i>t+h\}}\right] \\
&= E_t\left[\mathcal{E}_{t+h}\prod_{j\neq 0,i}\prod_{t\le s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right]\,\tilde{R}^{E,i}_{t+h}\right] \\
&\quad - E_t\left[\mathcal{E}_{t+h}\prod_{j\neq 0,i}\prod_{t\le s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right]\right]\,E_t\left[\tilde{R}^{E,i}_{t+h}\right] \\
&\quad + E_t\left[\mathcal{E}_{t+h}\prod_{j\neq 0,i}\prod_{t\le s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right]\right]\,E_t\left[\tilde{R}^{E,i}_{t+h}\right] \\
&= cov_t\left[\mathcal{E}_{t+h}\prod_{j\neq 0,i}\prod_{t\le s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right],\tilde{R}^{E,i}_{t+h}\right]+\frac{1}{q^i_tR^f_t}\,E_t\left[\tilde{R}^{E,i}_{t+h}\right], \qquad (16)
\end{aligned}$$

where and with R^f_t, $\mathcal{E}_{t+h}$ and q_t are given by

$$R^f_t = \left(E_t\left[\frac{M_{t+h}}{M_t}\right]\right)^{-1},$$

$$\begin{aligned}
\mathcal{E}_{t+h} &= \exp\left\{-\int_t^{t+h}r_s ds+\int_t^{t+h}\bar{\Gamma}_s\bar{\lambda}^P_s ds-\frac{1}{2}\int_t^{t+h}\Lambda^2_s\,ds-\int_t^{t+h}\Lambda_s\,dW_s\right\} \\
&\quad\times\exp\left\{\int_t^{t+h}\Gamma^0_s\bar{\lambda}^{P,0}_s\,ds\right\}, \qquad (17)
\end{aligned}$$

and

$$q^i_t = \frac{E_t\left[\mathcal{E}_{t+h}\Pi_j\Pi_{t\le s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right]\right]}{E_t\left[\mathcal{E}_{t+h}\Pi_{j\neq 0,i}\Pi_{t\le s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right]\right]}.$$

Note that the second equality follows from the fact that $\mathbf{1}_{\{\tau^i>t+h\}}=\mathbf{1}_{\{\bar{\tau}^0>t+h\}}\mathbf{1}_{\{\bar{\tau}^i>t+h\}}$.

The expected gross return on the firm i's equity claim can now be computed from (15) and (16) as

$$E_t\left[\tilde{R}^{E,i}_{t+h}\right] = q^i_t\left\{R^f_t-R^f_t\,cov_t\left[\mathcal{E}_{t+h}\Pi_{j\neq 0,i}\Pi_{t\le s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right],\tilde{R}^{E,i}_{t+h}\right]\right\}. \qquad (18)$$

According to (14) and (17), our DRP factor enters the return equation (18) mainly through $\mathcal{E}_{t+h}$. If the covariance term on the right-hand side is negligible, we will find no or only little evidence of the DRP factor being priced in equity markets. At the same time, the scaling term q_t^i captures the effect of the innovations in the counting process associated with the default event τ^i on the return on firm i's equity. It is driven by jump to default risk premia, that is, the ratios of risk-neutral to actual default intensities, for the market-wide and the firm-specific event times $\bar{\tau}^0$ and $\bar{\tau}^i$. We postpone a more detailed discussion of this point until Section 7.

The fact that corporate bonds yield non-zero payoffs in the event of default substantially changes the relation between bond returns and jump-to-default risk premia. The gross returns on corporate bonds, $\tilde{R}^{B,i}_{t+h}$, can be written as

$$\tilde{R}^{B,i}_{t+h} \quad = \quad \tilde{R}^{ND,i}_{t+h}\mathbf{1}_{\{\tau^i>t+h\}} + \tilde{R}^{D,i}_{t+h}\mathbf{1}_{\{t<\tau^i\leq t+h\}},$$

where $\tilde{R}^{ND,i}_{t+h}$ ($\tilde{R}^{D,i}_{t+h}$) stand for the total return on firm i's debt if the company does not (does) default prior to or at time $t+h$. We can use this fact to show that

$$\begin{aligned}
E_t\left[\frac{M_{t+h}}{M_t}\tilde{R}^{ND,i}_{t+h}\mathbf{1}_{\{\tau^i>t+h\}}\right] &= E_t\left[\mathcal{E}_{t+h}\Pi_{j\neq 0,i}\Pi_{t\leq s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right]\tilde{R}^{ND,i}_{t+h}\mathbf{1}_{\{\tau^i>t+h\}}\right] \\
&= E_t\left[\mathcal{E}_{t+h}\Pi_{j\neq 0,i}\Pi_{t\leq s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right]\tilde{R}^{B,i}_{t+h}\right] \\
&\quad - E_t\left[\mathcal{E}_{t+h}\Pi_{j\neq 0,i}\Pi_{t\leq s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right]\tilde{R}^{D,i}_{t+h}\mathbf{1}_{\{t<\tau^i\leq t+h\}}\right] \\
&= cov_t\left[\mathcal{E}_{t+h}\Pi_{j\neq 0,i}\Pi_{t\leq s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right], \tilde{R}^{B,i}_{t+h}\right] \\
&\quad + \frac{1}{q^i_t R^f_t}E_t\left[\tilde{R}^{B,i}_{t+h}\right] \\
&\quad - cov_t\left[\mathcal{E}_{t+h}\Pi_{j\neq 0,i}\Pi_{t\leq s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right], \tilde{R}^{D,i}_{t+h}\mathbf{1}_{\{t<\tau^i\leq t+h\}}\right] \\
&\quad - \frac{1}{q^i_t R^f_t}E_t\left[\tilde{R}^{D,i}_{t+h}\mathbf{1}_{\{t<\tau^i\leq t+h\}}\right].
\end{aligned}$$

This yields that the expected return on a firm i's debt claim is given by

$$\begin{aligned}
E_t\left[\tilde{R}^{B,i}_{t+h}\right] = q^i_t\Big\{ & R^f_t - R^f_t cov_t\left[\mathcal{E}_{t+h}\Pi_{j\neq 0,i}\Pi_{t\leq s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right], \tilde{R}^{B,i}_{t+h}\mathbf{1}_{\{\tau^i>t+h\}}\right] \\
& + \left(\frac{1}{q^i_t} - R^f_t\right)E_t\left[\tilde{R}^{B,i}_{t+h}\mathbf{1}_{\{t<\tau^i\leq t+h\}}\right] \\
& - R^f_t cov_t\left[\mathcal{E}_{t+h}\Pi_j\Pi_{t\leq s<t+h}\left[1-\Gamma^j_s\Delta\tilde{N}^j_s\right], \tilde{R}^{B,i}_{t+h}\mathbf{1}_{\{t<\tau^i\leq t+h\}}\right]\Big\}.
\end{aligned} \tag{19}$$

It is important to note that for corporate bonds of firm i, the DRP factor enters the return equation (19) through a new term, that is, via

$$-q_t^i R_t^f cov_t \left[\mathcal{E}_{t+h} \Pi_j \Pi_{t \leq s < t+h} \left[1 - \Gamma_s^j \Delta \tilde{N}_s^j \right], \tilde{R}_{t+h}^{B,i} \mathbf{1}_{\{t < \tau^i \leq t+h\}} \right] \Bigg\}.$$

Since the first component of the covariance term now includes the jump to default risk premia for the market-wide and the firm-specific event times $\bar{\tau}^0$ and $\bar{\tau}^i$, we expect the contribution of this term, and hence our DRP factor, to explaining corporate bond returns to be significant.

7 Discussion

Default risk premia display time-series properties that could stem from a systemic response to innovations in a common risk factor. This paper tests this hypothesis, by investigating the source for common variation in the portion of returns on credit default swaps that is not related to changes in risk-free rates or expected default losses. We extract a latent common component from firm specific changes in default risk premia that is orthogonal to known systematic risk factors during our sample period from 2001 to 2004. Asset pricing tests using returns on Bloomberg-NASD corporate bond indices generated from actual transaction prices of actively traded issues suggest that our discovered latent default risk premia factor (DRP) is priced in the corporate bond market. A cross-sectional analysis of Merrill Lynch corporate bond portfolios sorted on either industry, maturity or rating supports these findings. In our tests we control for firm characteristics such as contemporaneous default probabilities, leverage ratios, and recovery rates and we find that the common variation in changes in default risk premia is not likely to be due to these firm characteristics.

We also form portfolios using put options written on the S&P 500 index, sorted on moneyness and maturity. We find that for far-out-of-the-money index put options, both average returns and the beta estimate for our DRP factor increase with increasing time to maturity. The same holds true for out-of-the-money and at-the-money index put options. However, there is little to no evidence of the DRP factor being priced in the equity markets.

We develop a theoretical framework that shows that the DRP factor captures the jump-to-default risk associated with market-wide credit events. Within this framework, we show that, unlike risk premia on corporate bonds, equity risk premia are only marginally affected by our DRP factor. This results is based on the observation that the DRP factor has a much stronger impact on the returns of assets with a non-degenerate payoff structure in the default states. It is of interest to note that Collin-Dufresne, Goldstein, and Helwege (2003) also develop a reduced-form model where jump-to-default risk is priced. Their framework, however, can be interpreted as updating of beliefs due to an unexpected credit event. It is closely related to the concept of information-driven default contagion using frailty models introduced to the credit risk literature by Schönbucher (2003). In future work, we plan to extend the specification of the pricing kernel in 12, and of our assumptions regarding the physical and risk-neutral default intensity processes in (5) and (6), to allow for an updating of investor's beliefs upon observed default events.

As a final remark we would like to point out that the result for the expected return on firm i's equity claim in (18) suggests that equity risk premia depend on the ratios of risk-neutral to actual default intensities associated with the default event time τ^i, as captured by the scaling factor q_t^i. For the remainder of this section, let us define jump to default risk premia for firm i as $\lambda_t^{Q,i}/\lambda_t^{P,i}$. We investigate the extent to which the common component extracted from these instantaneous firm-specific jump to default risk premia, after controlling for other sources of common variation, can be associated with an alternative credit market risk factor that is more likely to be priced in equity markets.

Appendix C describes our approach to construct a new jump to default risk premia (JDRP) factor from $\{\lambda_t^{Q,i}/\lambda_t^{P,i}\}$. In order to see whether the JDRP factor is priced in equity markets we employ two sets of test assets, namely the 49 Fama-French industry portfolios and the 25 Fama-French equity portfolios sorted on size and book-to-market equity. We implement the Fama-MacBeth two-step procedure to estimate the impact of $JDRP_t$ on the average returns of these portfolios. The results of the cross-sectional regressions are reported in Table 16 and Table 17, respectively. We notice that the relation between the average excess returns on the test assets and the time-series loadings on $JDRP$ is positive and statistically significant, even after controlling for other known sources of systematic risk. These results are quite encouraging, and the subject of future investigation.

A CDS and EDF Coverage

Industry Name	Company Name	N_{CDS}	Rating	5-yr CDS spread	EDF	N_{EDF}
BUSINESS SVCS	COMPUTER ASSOCIATES INTL INC	231	BB	1.612	1.05	144
BUSINESS SVCS	COMPUTER SCIENCES CORP	194	A	0.495	0.73	144
BUSINESS SVCS	ELECTRONIC DATA SYSTEMS CORP	205	BB	1.754	1.12	144
BUSINESS SVCS	INTERPUBLIC GROUP OF COS	189	BB	1.837	1.61	144
BUSINESS SVCS	INTL BUSINESS MACHINES CORP	231	A	0.368	0.16	144
BUSINESS SVCS	LUCENT TECHNOLOGIES INC	231	B	7.08	6.43	108
BUSINESS SVCS	OMNICOM GROUP	197	BBB	0.649	0.37	144
BUSINESS SVCS	SABRE HOLDINGS CORP	160	BBB	0.663	0.63	100
CHEMICALS	ABBOTT LABORATORIES	182	AA	0.170	0.04	144
CHEMICALS	AIR PRODUCTS & CHEMICALS INC	231	A	0.270	0.07	144
CHEMICALS	AMGEN INC	171	A	0.254	0.02	144
CHEMICALS	AVON PRODUCTS	230	A	0.252	0.07	144
CHEMICALS	BRISTOL-MYERS SQUIBB CO	166	A	0.329	0.22	144
CHEMICALS	CHIRON CORP	206	BBB	0.451	0.15	144
CHEMICALS	COLGATE-PALMOLIVE CO	224	AA	0.153	0.02	144
CHEMICALS	CYTEC INDUSTRIES INC	160	BBB	0.457	0.12	132
CHEMICALS	DOW CHEMICAL	229	A	0.508	0.24	144
CHEMICALS	EASTMAN CHEMICAL CO	199	BBB	0.739	0.31	136
CHEMICALS	MERCK & CO	200	AA	0.177	0.06	144
CHEMICALS	MONSANTO CO	162	BBB	0.360	0.13	49
CHEMICALS	PFIZER INC	156	AAA	0.104	0.02	144
CHEMICALS	PPG INDUSTRIES INC	207	A	0.294	0.11	144
CHEMICALS	PRAXAIR INC	207	A	0.22	0.04	144
CHEMICALS	PROCTER & GAMBLE CO	228	AA	0.176	0.02	144
CHEMICALS	ROHM AND HAAS CO	231	BBB	0.394	0.15	144
CHEMICALS	SCHERING-PLOUGH	195	BBB	0.375	0.22	144
CHEMICALS	SHERWIN-WILLIAMS CO	215	A	0.312	0.15	144
CHEMICALS	WYETH	206	A	0.531	0.16	144
COMMUNICATION	ALLTEL CORP	154	A	0.473	0.12	144
COMMUNICATION	AT&T CORP	231	BB	1.443	0.68	144
COMMUNICATION	AT&T WIRELESS SERVICES INC	162	BBB	1.329	0.8	51
COMMUNICATION	BELLSOUTH CORP	204	A	0.423	0.13	144
COMMUNICATION	CENTURYTEL INC	189	BBB	0.988	0.22	144
COMMUNICATION	CINGULAR WIRELESS LLC	174	BBB	0.465		
COMMUNICATION	CITIZENS COMMUNICATIONS CO	211	BB	2.175	0.9	144
COMMUNICATION	CLEAR CHANNEL COMMUNICATIONS	231	BBB	1.125	0.38	144
COMMUNICATION	COMCAST CORP	220	BBB	1.061	0.38	144
COMMUNICATION	LIBERTY MEDIA CORP -SER A	212	BB	2.051	0.42	105
COMMUNICATION	NEXTEL COMMUNICATIONS INC	214	BB	3.985	3.28	144
COMMUNICATION	SBC COMMUNICATIONS INC	210	A	0.465	0.255	144
COMMUNICATION	SPRINT CORP	220	BBB	1.264	0.83	144
COMMUNICATION	VERIZON COMMUNICATIONS INC	231	A	0.713	0.23	144
ELEC., GAS & SEWAGE	AMERICAN ELEC. POWER	194	BBB	0.806	0.39	144
ELEC., GAS & SEWAGE	CINERGY CORP	210	BBB	0.575	0.22	129
ELEC., GAS & SEWAGE	CONSTELLATION ENERGY GRP INC	199	BBB	0.583	0.54	144
ELEC., GAS & SEWAGE	DOMINION RESOURCES INC	225	BBB	0.638	0.16	144
ELEC., GAS & SEWAGE	DTE ENERGY CO	194	BBB	0.741	0.42	144
ELEC., GAS & SEWAGE	DUKE ENERGY CORP	231	BBB	0.632	0.55	144
ELEC., GAS & SEWAGE	EL PASO CORP	215	CCC	5.323	6.645	144
ELEC., GAS & SEWAGE	EXELON CORP	200	BBB	0.514	0.29	144
ELEC., GAS & SEWAGE	FIRSTENERGY CORP	155	BBB	1.122	0.29	144
ELEC., GAS & SEWAGE	KINDER MORGAN ENERGY -LP	207	BBB	0.57	0.13	142
ELEC., GAS & SEWAGE	KINDER MORGAN INC	177	BBB	0.586	0.07	144
ELEC., GAS & SEWAGE	ONEOK INC	207	BBB	0.521	0.33	144
ELEC., GAS & SEWAGE	PROGRESS ENERGY INC	184	BBB	0.656	0.23	144
ELEC., GAS & SEWAGE	SEMPRA ENERGY	201	BBB	0.592	0.58	144
ELEC., GAS & SEWAGE	TECO ENERGY INC	159	BB	2.209	1.53	144
ELEC., GAS & SEWAGE	TXU CORP	231	BBB	1.157	0.56	144
ELEC., GAS & SEWAGE	WASTE MANAGEMENT INC	218	BBB	1.018	0.3	144
ELEC., GAS & SEWAGE	WILLIAMS COS INC	219	B	2.610	2.96	144
ELEC., GAS & SEWAGE	XCEL ENERGY INC	168	BBB	0.675	0.73	144
FOOD	ANHEUSER-BUSCH COS INC	187	A	0.185	0.02	144
FOOD	CAMPBELL SOUP CO	231	A	0.35	0.08	144
FOOD	COCA-COLA ENTERPRISES	231	A	0.298	–	–

Industry Name	Company Name	N_{CDS}	Rating	5-yr CDS spread	EDF	N_{EDF}
FOOD	CONAGRA FOODS INC	226	BBB	0.485	0.13	144
FOOD	GENERAL MILLS INC	214	BBB	0.513	0.03	144
FOOD	HEINZ (H J) CO	231	A	0.3	0.07	144
FOOD	KELLOGG CO	218	BBB	0.380	0.04	144
FOOD	KRAFT FOODS INC	188	A	0.479	0.02	46
FOOD	SARA LEE CORP	231	A	0.285	0.08	144
FOOD	TYSON FOODS INC -CL A	194	BBB	1.165	0.34	144
INDUSTRIAL MACH.	BAKER HUGHES INC	208	A	0.274	0.15	144
INDUSTRIAL MACH.	BLACK & DECKER CORP	225	BBB	0.438	0.38	144
INDUSTRIAL MACH.	CATERPILLAR INC	231	A	0.331	0.21	144
INDUSTRIAL MACH.	COMPAQ COMPUTER CORP	231	BBB	0.658	2.92	109
INDUSTRIAL MACH.	CUMMINS INC	227	BB	1.962	0.91	144
INDUSTRIAL MACH.	DEERE & CO	231	A	0.43	0.43	144
INDUSTRIAL MACH.	HEWLETT-PACKARD CO	219	A	0.528	1.125	144
INDUSTRIAL MACH.	PITNEY BOWES INC	231	A	0.237	0.38	144
INDUSTRIAL MACH.	SUN MICROSYSTEMS INC	219	BB	1.117	1.96	144
INDUSTRIAL MACH.	XEROX CORP	218	BB	5.093	5.8	144
INSTRUMENTS	BAXTER INTERNATIONAL INC	215	BBB	0.474	0.27	144
INSTRUMENTS	BOSTON SCIENTIFIC CORP	203	BBB	0.348	0.09	144
INSTRUMENTS	DANAHER CORP	205	A	0.32	0.07	144
INSTRUMENTS	EASTMAN KODAK CO	231	BB	1.295	0.7	144
INSTRUMENTS	MEDTRONIC INC	173	A	0.222	0.02	144
INSTRUMENTS	NORTHROP GRUMMAN CORP	223	BBB	0.586	0.16	144
INSTRUMENTS	RAYTHEON CO	231	BBB	0.97	0.6	144
PAPER	3M CO	178	AA	0.141	0.02	144
PAPER	BOWATER INC	228	BB	2.566	0.35	144
PAPER	GEORGIA-PACIFIC CORP	231	BB	2.668	0.64	144
PAPER	INTL PAPER CO	231	BBB	0.767	0.24	144
PAPER	MEADWESTVACO CORP	173	BBB	0.728	0.22	144
PAPER	SEALED AIR CORP	215	BBB	1.536	0.18	144
PAPER	TEMPLE-INLAND INC	218	BBB	1.11	0.27	144
TRANSPORTATION EQPT	ARVINMERITOR INC	230	BB	2.85	1.79	87
TRANSPORTATION EQPT	BOEING CO	231	A	0.426	0.3	144
TRANSPORTATION EQPT	BORGWARNER INC	229	BBB	0.688	0.23	140
TRANSPORTATION EQPT	DANA CORP	217	BB	4.233	1.17	144
TRANSPORTATION EQPT	DELPHI CORP	230	B	1.440	1.11	72
TRANSPORTATION EQPT	EATON CORP	215	A	0.32	0.12	144
TRANSPORTATION EQPT	FORD MOTOR CO	223	BB	2.148	0.29	144
TRANSPORTATION EQPT	GENERAL MOTORS CORP	231	BB	1.993	0.19	144
TRANSPORTATION EQPT	GOODRICH CORP	211	BBB	0.99	0.835	144
TRANSPORTATION EQPT	HONEYWELL INTERNATIONAL INC	196	A	0.372	0.29	144
TRANSPORTATION EQPT	LOCKHEED MARTIN CORP	231	BBB	0.5	0.13	144
TRANSPORTATION EQPT	UNITED TECHNOLOGIES CORP	195	A	0.217	0.14	144
TRANSPORTATION EQPT	VISTEON CORP	230	B	2.503	2.58	57

Table 2: **Firm Summary Characteristics:** N_{CDS} is the number of valid observations for the five-year CDS available for each firm. $Rating$ is the median rating. $5-yr\ CDS\ Spread$ is the median spread, in percent, of the 5-year CDS for senior unsecured debt for the firm, with a default event defined as modified restructuring. EDF is the median Moody's KMV EDF in percent. The variables N_{CDS} through EDF are computed over the period January 2001 through June 2005. N_{EDF} is the valid number of monthly observations of the EDF values for the period July 1993 - June 2005. We use this longer time period to interpolate weekly EDF values from monthly EDF values.

B Time Series Estimation of Default Intensities

Table 3: Summary statistics for firm-by-firm EDF-implied actual default intensity parameters.

	$\kappa\theta$	κ	σ
mean	1.47	0.52	1.19
median	0.90	0.36	1.07
std. dev.	2.01	0.47	0.41

Table 4: Summary statistics for firm-by-firm MLE parameter estimates, using Markit five-year CDS rates with modified-restructuring and contemporaneous recovery rates.

	θ^Q	κ^Q	σ^Q	$\tilde{\theta}^Q$	$\tilde{\kappa}^Q$
mean	4.57	0.04	0.65	6.24	-0.11
median	4.53	0.03	0.49	5.56	-0.12
std. dev.	0.94	0.09	0.43	3.73	0.10
no. firms	108				

C An Alternative Default Risk Premia Factor

This appendix describes our approach to construct the alternative jump to default risk premia (JDRP) factor, as motivated in Section 7. Let $\pi_t^i = \lambda_t^{Q,i}/\lambda_t^{P,i}$ denote the time series of jump to default risk premia for firm i. Also, let F_t^D denote the time series of the levels of the latent common component to be extracted[12]. Assume that these levels follow a following VAR process:

$$F_{t+1}^D - F_t^D = \rho(\bar{F} - F_t^D) + \sigma\xi_{t+1},$$

where $\bar{F}$ can be interpreted as the long-term level that F_t^D mean-reverts to, and ρ and σ are scalars.

[12]One should think of these levels as realized excess returns on a common risk factor embedded in the default events across firms.

Further, let π_t^D denote the conditional expectation of F_{t+1}^D, relative to the information at time t. Then,

$$\pi_t^D = (1-\rho)F_t^D + \rho\bar{F}. \tag{C.1}$$

As before, let F_t^S denote the vector of returns on other known systematic factors, and let π_t^S denote the corresponding conditional risk premia.

The process of extracting the levels of the latent component, F_t^D, is performed in several steps. First we estimate, firm by firm, the OLS model

$$\pi_t^i = \beta^i \cdot \pi_t^S + \epsilon_t^i,$$

and construct the time series of implied errors $\hat{\epsilon}_t^i = \pi_t^i - \hat{\beta}^i \cdot \pi_t^S$. Second, we model the errors ϵ_t^i in the previous regression as

$$\epsilon_t^i = \alpha^i + \gamma^i \left[\pi_t^D - \bar{F}\right] + \nu_t^i,$$

where ν_t^i are standard normal variables, independent of π_t^D and independent across firms and time.

Notice that $\alpha^i = E\epsilon_t^i$. Due to the fact that γ^i can not be simultaneously identified, we choose the normalization

$$1 = \frac{1}{N}\sum_i^N \gamma^i.$$

This normalization, together with the assumption that the errors ν_t^i are i.i.d. across firms, allows us to obtain an estimate for π_t^D, that is,

$$\pi_t^D - \bar{F} = \frac{1}{N}\sum_{i=1}^N \left(\hat{\epsilon}_t^i - \alpha^i\right). \tag{C.2}$$

We can now use (C.2) and compute the OLS estimates for γ^ias

$$\gamma^i = \frac{COVAR\left(\epsilon_t^i, \pi_t^D\right)}{VAR\left(\pi_t^D\right)},$$

where $COVAR$ and VAR are the unconditional covariance and variance operators, respectively.

Third, note that, on one hand, the following formulas hold

$$\begin{aligned} VAR_t\left[\pi_{t+1}^D\right] &= (1-\rho)^2 VAR_t\left[F_{t+1}^D\right] = (1-\rho)^2\sigma^2 \\ COVAR_t\left[\pi_{t+1}^D, \pi_{t+2}^D\right] &= (1-\rho)^3 VAR_t\left[F_{t+1}^D\right] = (1-\rho)^3\sigma^2 \end{aligned} \tag{C.3}$$

On the other hand, the left hand side of the previous equations can be expressed as

$$\begin{aligned} VAR_t\left[\pi_{t+1}^D\right] &= \mathbb{E}_t\left[\left(\pi_{t+1}^D\right)^2\right] - \left((1-\rho)^2\left[F_t^D - \bar{F}\right] + \bar{F}\right)^2 \\ COVAR_t\left[\pi_{t+1}^D, \pi_{t+2}^D\right] &= \mathbb{E}_t\left[\pi_{t+1}^D \pi_{t+2}^D\right] \\ &\quad - \left((1-\rho)^2\left[F_t^D - \bar{F}\right] + \bar{F}\right)\left((1-\rho)^3\left[F_t^D - \bar{F}\right] + \bar{F}\right) \end{aligned} \tag{C.4}$$

Applying the unconditional expectation operator in both (C.3) and (C.4), and comparing the right hand sides of these two equations while using the fact that the unconditional variance of F_t^D is $\frac{\sigma^2}{1-(1-\rho)^2}$, we obtain

$$\begin{aligned} VAR\left[\pi_{t+1}^D\right] &= \frac{(1-\rho)^2}{1-(1-\rho)^2}\sigma^2, \quad \text{and} \\ COVAR\left[\pi_{t+1}^D, \pi_{t+2}^D\right] &= \frac{(1-\rho)^3}{1-(1-\rho)^2}\sigma^2. \end{aligned}$$

This yields

$$1-\rho = \frac{COVAR\left[\pi_{t+1}^D, \pi_{t+2}^D\right]}{VAR\left[\pi_{t+1}^D\right]},$$

which together with (C.1) yields the level of the latent common factor, F_t^D, as

$$F_t^D = \bar{F} + \frac{VAR\left[\pi_{t+1}^D\right]}{COVAR\left[\pi_{t+1}^D, \pi_{t+2}^D\right]}\left[\pi_t^D - \bar{F}\right],$$

where $\pi_t^D - \bar{F}$ is computed using (C.2).

Notice that our results depend additively on the free parameter $\bar{F}$, which is not identifiable in this context. Nevertheless, for the assets pricing tests we are mainly interested in covariances, and knowledge of $\bar{F}$ is not necessary.

D Asset Pricing Test Results

α	β_{RMO}	β_{SMB}	β_{HML}	β_{UMD}	β_{DEF}	β_{TERM}	β_{DRP}	R^2
			High-yield Corporate Bond Portfolio					
0.0016	0.2159	0.1658	0.0002	−0.3104	1.6699	0.1190		0.6780
(2.4653)	(7.0264)	(2.8325)	(0.0022)	(9.2244)	(7.1581)	(4.5292)		
			Investment-grade Corporate Bond Portfolio					
0.0002	0.0178	0.0417	0.0569	−0.0917	1.0242	0.2021		0.8622
(0.9384)	(1.5070)	(1.8540)	(2.0379)	(7.0726)	(11.4240)	(20.1152)		
			High-yield Corporate Bond Portfolio					
0.0014	0.2280	0.1536	0.0126	−0.3027	1.4719	0.1247	77.9450	0.7055
(2.3652)	(7.6567)	(2.7245)	(0.1819)	(−9.3351)	(6.3189)	(4.9266)	(3.1292)	
			Investment-grade Corporate Bond Portfolio					
0.0002	0.0219	0.0376	0.0610	−0.0892	0.9590	0.2041	25.3945	0.8707
(0.8087)	(1.8826)	(1.7122)	(2.2390)	(7.0484)	(10.5671)	(20.8036)	2.6183	

Table 5: **The NASD High-yield and Investment-grade Corporate Bond Portfolios: Time-series regressions** This table reports the results of the regressions of the excess realized returns of portfolios of investment and high grade corporate bonds on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$ and the default risk premia factor returns, DRP: October 2002 to December 2004, 120 weeks. Specifically, for each portfolio we estimate the following regressions (with or without the last regressor): $R(t) - RF(t) = \alpha + \beta_{RMO}RMO(t) + \beta_{SMB}SMB(t) + \beta_{HML}HML(t) + \beta_{UMD}UMD(t) + \beta_{DEF}DEF(t) + \beta_{TERM}TERM(t) + \beta_{DRP}DRP(t) + \epsilon(t)$. The first part of the table reports the results for the regression without the default risk premia factor $DRP(t)$, while the second panel reports the results for the full regression. The reported intercept, slopes and R^2 are the mean values of the intercept, slopes and R^2 of each of the 97 portfolios. The t-statistics are reported in parentheses, and they are computed as the ratio of the cross-sectional mean to the cross-sectional standard deviation.

α	β_{RMO}	β_{SMB}	β_{HML}	β_{UMD}	β_{DEF}	β_{TERM}	β_{DRP}	R^2	$E[R]$
0.0008	−0.0914	0.0388	0.0716	0.0414	−0.1939	0.2684	19.5926	0.9239	0.0011
(5.3606)	(13.4984)	(3.3710)	(6.5968)	(6.0138)	(4.4193)	(45.0936)	(4.3556)		
0.0008	−0.0698	0.0197	0.0719	0.0181	0.0404	0.2358	7.1116	0.9090	0.0010
(5.5616)	(10.8000)	(1.7989)	(6.8391)	(2.7691)	(0.9692)	(41.5626)	(1.6634)		
0.0009	−0.0710	0.0231	0.0788	0.0060	0.1230	0.2671	10.2338	0.9135	0.0012
(5.8744)	(10.0376)	(1.9277)	(6.9683)	(0.8352)	(2.6956)	(43.1547)	(2.1878)		
0.0008	−0.0187	0.0343	0.1033	−0.0726	0.8057	0.2683	19.8147	0.9310	0.0013
(5.3605)	(2.7672)	(2.9774)	(9.5216)	(10.5297)	(18.3585)	(45.0719)	(4.4042)		
0.0007	0.0824	0.0370	0.0915	−0.0962	1.0311	0.1456	33.5648	0.4586	0.0012
(1.1452)	(3.1205)	(0.7793)	(1.4354)	(3.3281)	(6.2625)	(5.9271)	(2.0815)		
0.0012	0.1332	0.0995	0.0290	−0.1934	1.1471	0.0764	44.6200	0.5903	0.0019
(2.3604)	(5.5481)	(2.3054)	(0.5006)	(7.3546)	(7.6604)	(3.4203)	(3.0426)		
0.0023	0.1413	0.3270	−0.0045	−0.2211	1.1849	0.0545	26.6474	0.4175	0.0034
(2.6687)	(3.6424)	(4.6899)	(0.0478)	(5.2041)	(4.8976)	(1.5098)	(1.1246)		

Table 6: **The Merrill Lynch Corporate Bond Portfolios, by Ratings: Time-series regressions** This table reports the results of the regressions of the excess realized returns of 7 Merrill Lynch corporate bond portfolios, by ratings, on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$ and the default risk premia factor returns, DRP: October 2002 to December 2004, 120 weeks. Specifically, for each portfolio we estimate the following regressions: $R(t) - RF(t) = \alpha + \beta_{RMO}RMO(t) + \beta_{SMB}SMB(t) + \beta_{HML}HML(t) + \beta_{UMD}UMD(t) + \beta_{DEF}DEF(t) + \beta_{TERM}TERM(t) + \beta_{DRP}DRP(t) + \epsilon(t)$. The first regression corresponds to the AAA portfolio, while the last regression corresponds to C portfolio. The t-statistics are reported in parentheses.

α	β_{RMO}	β_{SMB}	β_{HML}	β_{UMD}	β_{DEF}	β_{TERM}	β_{DRP}	R^2	$E[R]$
0.0005	−0.0176	0.0119	0.0485	−0.0152	0.1688	0.0666	8.9513	0.9239	0.0006
(3.9688)	(3.1616)	(1.2561)	(5.4426)	(2.6955)	(4.6850)	(13.6164)	(2.4235)		
0.0007	−0.0482	0.0225	0.0795	−0.0133	0.2316	0.1669	12.3774	0.9090	0.0009
(3.9138)	(6.0379)	(1.6586)	(6.2138)	(1.6389)	(4.4781)	(23.7843)	(2.3341)		
0.0009	−0.0629	0.0275	0.0907	−0.0196	0.3320	0.2445	14.6392	0.9135	0.0012
(4.5545)	(7.0838)	(1.8230)	(6.3738)	(2.1738)	(5.7720)	(31.3381)	(2.4829)		
0.0009	−0.0587	0.0252	0.1044	−0.0281	0.4455	0.3214	16.8712	0.9310	0.0013
(4.9083)	(6.9788)	(1.7626)	(7.7396)	(3.2845)	(8.1723)	(43.4650)	(3.0187)		
0.0009	−0.0935	0.0428	0.1410	−0.0111	0.4169	0.3999	22.2762	0.4586	0.0014
(4.0972)	(9.6848)	(2.6149)	(9.1411)	(1.1356)	(6.6940)	(47.3570)	(3.4904)		
0.0012	−0.0465	0.0404	0.1028	−0.0444	0.7207	0.4622	15.6869	0.5903	0.0018
(5.7287)	(4.7896)	(2.4463)	(6.9282)	(4.4780)	(11.1360)	(54.0331)	(2.4073)		

Table 7: **The Merrill Lynch Corporate Bond Portfolios, by Maturity: Time-series regressions** This table reports the results of the regressions of the excess realized returns of 6 Merrill Lynch corporate bond portfolios, by maturity, on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$ and the default risk premia factor returns, DRP: October 2002 to December 2004, 120 weeks. Specifically, for each portfolio we estimate the following regressions: $R(t) - RF(t) = \alpha + \beta_{RMO}RMO(t) + \beta_{SMB}SMB(t) + \beta_{HML}HML(t) + \beta_{UMD}UMD(t) + \beta_{DEF}DEF(t) + \beta_{TERM}TERM(t) + \beta_{DRP}DRP(t) + \epsilon(t)$. The first regression corresponds to the 1-3Y portfolio, while the last regression corresponds to 15Y+ portfolio. The t-statistics are reported in parentheses.

α	β_{RMO}	β_{SMB}	β_{HML}	β_{UMD}	β_{DEF}	β_{TERM}	β_{DRP}	R^2
0.0013	0.0678	0.1315	0.0621	-0.1479	0.6242	0.0519		0.2286
(5.5889)	(4.8076)	(5.4432)	(2.7536)	(5.8612)	(4.5641)	(7.3943)		
0.0013	0.0680	0.1309	0.0619	-0.1475	0.6214	0.0521	32.4466	0.2460
(5.5918)	(4.7993)	(5.4468)	(2.7402)	(-5.8542)	(4.5368)	(7.4066)	(5.5288)	

Table 8: **The Merrill Lynch High-yield Corporate Bond Portfolios, by Sectors: Time-series regressions** This table reports the results of the regressions of the excess realized returns of 30 Merrill Lynch high-yield corporate bond portfolios, by sector, on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$ and the default risk premia factor returns, DRP: October 2002 to December 2004, 120 weeks. Specifically, for each portfolio we estimate the following regressions (with or without the last regressor): $R(t) - RF(t) = \alpha + \beta_{RMO} RMO(t) + \beta_{SMB} SMB(t) + \beta_{HML} HML(t) + \beta_{UMD} UMD(t) + \beta_{DEF} DEF(t) + \beta_{TERM} TERM(t) + \beta_{DRP} DRP(t) + \epsilon(t)$. The first part of the table reports the results for the regression without the default risk premia factor $DRP(t)$, while the second panel reports the results for the full regression. The reported intercept, slopes and R^2 are the mean values of the intercept, slopes and R^2 of each of the portfolios. The t-statistics are reported in parentheses, and they are computed as the ratio of the cross-sectional mean to the cross-sectional standard deviation.

$E[\gamma_0]$	$E[\gamma^{RMO}]$	$E[\gamma^{SMB}]$	$E[\gamma^{HML}]$	$E[\gamma^{UMD}]$	$E[\gamma^{DEF}]$	$E[\gamma^{TERM}]$	$E[\gamma^{DRP}]$	R^2
0.0031	0.0054	0.0004	−0.0023	0.0089	−0.0000	−0.0036		0.4825
(6.3243)	(1.1861)	(0.1354)	(0.7457)	(1.4053)	(0.0031)	(1.0078)		
0.0030	0.0004	−0.0014	−0.0023	0.0058	−0.0002	−0.0045	0.0000	0.5154
(5.7571)	(0.0942)	(0.4677)	(0.7591)	(0.9891)	(0.2686)	(1.2864)	(2.3742)	

Table 9: **The Merrill Lynch High-yield Corporate Bond Portfolios, by Sectors: Cross-sectional regressions** This table reports the results of the weekly cross-sectional regressions of the excess realized returns of 30 Merrill Lynch high-yield corporate bond portfolios, by sector, on the loadings on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$ and the default factor returns, DRP: January 2001 to December 2004, 206 weeks. Specifically, we estimate the coefficients of the following regressions (with or without the last regressor): $E\left[R(t)^i - RF(t)\right] = E\left[\gamma_0(t)\right] + E\left[\gamma^{RMO}(t)\right]\beta^i_{RMO} + E\left[\gamma^{SMB}(t)\right]\beta^i_{SMB} + E\left[\gamma^{HML}(t)\right]\beta^i_{HML} + E\left[\gamma^{DRP}(t)\right]\beta^i_{DRP}$. The coefficients are estimated as the averages of the weekly coefficients in the following cross-sectional regressions: $R(t)^i - RF(t) = \gamma_0 + \gamma^{RMO}(t)\beta^i_{RMO} + \gamma^{SMB}(t)\beta^i_{SMB} + \gamma^{HML}(t)\beta^i_{HML} + \gamma^{UMD}(t)\beta^i_{UMD} + \gamma^{DEF}(t)\beta^i_{DEF} + \gamma^{TERM}(t)\beta^i_{TERM} + +\gamma^{DRP}(t)\beta^i_{DRP} + \epsilon^i(t)$. The first half of the table reports the results of the regressions without the loading on the default risk premia factor, while the second half reports the results for the full regressions. The reported intercept, slopes and R^2 are the mean values of the intercept, slopes and R^2, across weeks. The t-statistics, reported in parentheses, are the average intercepts or slopes divided by their time series standard error, after accounting for autocorrelations.

α	β_{RMO}	β_{SMB}	β_{HML}	β_{UMD}	β_{DEF}	β_{TERM}	β_{DRP}	R^2
0.0009	-0.0372	0.0116	0.0752	-0.0555	0.7227	0.2663		0.8209
(11.9017)	(1.3100)	(0.5862)	(7.0511)	(1.0687)	(1.5734)	(19.9529)		
0.0009	-0.0372	0.0117	0.0752	-0.0554	0.7227	0.2663	11.7562	0.8240
(11.9036)	(1.3096)	(0.5864)	(7.0517)	(1.0687)	(1.5734)	(19.9537)	(2.0539)	

Table 10: **The Merrill Lynch Investment-grade Corporate Bond Portfolios, by Industry: Time-series regressions** This table reports the results of the regressions of the excess realized returns of 4 Merrill Lynch investment-grade corporate bond portfolios, by industry, on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$ and the default risk premia factor returns, DRP: October 2002 to December 2004, 120 weeks. Specifically, for each portfolio we estimate the following regressions (with or without the last regressor): $R(t)-RF(t) = \alpha+\beta_{RMO}RMO(t)+\beta_{SMB}SMB(t)+\beta_{HML}HML(t)+\beta_{UMD}UMD(t)+\beta_{DEF}DEF(t)+\beta_{TERM}TERM(t)+\beta_{DRP}DRP(t)+\epsilon(t)$. The t-statistics are reported in parentheses.

α	β_{RMO}	β_{SMB}	β_{HML}	β_{UMD}	β_{DEF}	β_{TERM}	β_{DRP}	R^2
0.0000	1.0246	0.3613	−0.3541	−0.5106	0.1309	−0.0150		0.8186
(0.8673)	(42.0155)	(7.6857)	(7.7704)	(41.3457)	(2.0734)	(3.6345)		
0.0000	1.0246	0.3613	−0.3541	−0.5106	0.1309	−0.0150	−11.2811	0.8201
(0.0517)	(57.4215)	(8.0426)	(7.9811)	(32.0391)	(2.7223)	(3.6662)	(3.2307)	

Table 11: **The Size and Book-to-market Equity Portfolios: Time-series regressions** This table reports the results of the regressions of the excess realized returns of 97 portfolios formed on size and book-to-market equity on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$ and the default risk premia factor returns, DRP: January 2001 to December 2004, 206 weeks. Specifically, for each portfolio we estimate the following regressions (with or without the last regressor): $R(t) - RF(t) = \alpha + \beta_{RMO}RMO(t) + \beta_{SMB}SMB(t) + \beta_{HML}HML(t) + \beta_{UMD}UMD(t) + \beta_{DEF}DEF(t) + \beta_{TERM}TERM(t) + \beta_{DRP}DRP(t) + \epsilon(t)$. The first part of the table reports the results for the regression without the default factor $DRP(t)$, while the second panel reports the results for the full regression. The reported intercept, slopes and R^2 are the mean values of the intercept, slopes and R^2 of each of the 97 portfolios. The t-statistics are reported in parentheses, and they are computed as the ratio of the cross-sectional mean to the cross-sectional standard deviation.

$E[\gamma_0]$	$E[\gamma^{RMO}]$	$E[\gamma^{SMB}]$	$E[\gamma^{HML}]$	$E[\gamma^{UMD}]$	$E[\gamma^{DEF}]$	$E[\gamma^{TERM}]$	$E[\gamma^{DRP}]$	R^2
0.0037	−0.0004	0.0017	0.0019	0.0005	0.0000	0.0044		0.3715
(3.8398)	(0.2701)	(1.8621)	(1.8223)	(0.2419)	(0.0457)	(1.4879)		
0.0037	−0.0006	0.0017	0.0019	0.0003	−0.0000	0.0045	−0.0000	0.3847
(3.8392)	(0.3568)	(1.8960)	(1.8740)	(0.1355)	(0.0145)	(1.5482)	(1.2256)	

Table 12: **The Size and Book-to-market Equity Portfolios: Cross-sectional regressions** This table reports the results of the weekly cross-sectional regressions of the excess realized returns of 97 portfolios formed on size and book-to-market equity on the loadings on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$ and the default risk premia factor returns, DRP: January 2001 to December 2004, 206 weeks. Specifically, we estimate the coefficients of the following regressions (with or without the last regressor): $E\left[R(t)^i - RF(t)\right] = E\left[\gamma_0(t)\right] + E\left[\gamma^{RMO}(t)\right]\beta^i_{RMO} + E\left[\gamma^{SMB}(t)\right]\beta^i_{SMB} + E\left[\gamma^{HML}(t)\right]\beta^i_{HML} + E\left[\gamma^{DRP}(t)\right]\beta^i_{DRP}$. The coefficient are estimated as the averages of the weekly coefficients in the following cross-sectional regressions: $R(t)^i - RF(t) = \gamma_0 + \gamma^{RMO}(t)\beta^i_{RMO} + \gamma^{SMB}(t)\beta^i_{SMB} + \gamma^{HML}(t)\beta^i_{HML} + \gamma^{UMD}(t)\beta^i_{UMD} + \gamma^{DEF}(t)\beta^i_{DEF} + \gamma^{TERM}(t)\beta^i_{TERM} + +\gamma^{DRP}(t)\beta^i_{DRP} + \epsilon^i(t)$. The first half of the table reports the results of the regressions without the loading on the default risk premia factor, while the second half reports the results for the full regressions. The reported intercept, slopes and R^2 are the mean values of the intercept, slopes and R^2, across weeks. The t-statistics, reported in parentheses, are the average intercepts or slopes divided by their time series standard error, after accounting for autocorrelations.

	Maturity Bin			
	$10 \leq T \leq 30$	$31 \leq T \leq 60$	$61 \leq T \leq 150$	$151 \leq T$
	Moneyness			
Moneyness Bin				
1	0.90	0.86	0.83	0.78
2	0.97	0.95	0.93	0.91
3	1.04	1.04	1.04	1.03
	Maturity (days)			
Moneyness Bin				
1	21.3	46.8	104.8	368.5
2	20.9	45.0	100.9	353.8
3	20.7	43.6	98.0	312.4
	Number of Valid Returns Observations			
Moneyness Bin				
1	768	1555	1295	1154
2	769	1556	1296	1155
3	768	1555	1296	1154

Table 13: **Summary Statistics for the Index Put Option Portfolios** This table summarizes the characteristics of the options in the different maturity and moneyness bins that are used to form the put option test portfolios. *Moneyness* is the present value of the strike price divided by the current value of the S&P 500 index. *Maturity (T)* is the time to option expiration in days. *Returns* are the *weekly* buy-and-hold returns. Values for moneyness, maturity and returns are value-weighted averages for the options in each bin. β_{DRP} is the factor loading on our default risk premia factor for the returns of the options portfolios.

	Maturity Bin			
	$10 \leq T \leq 30$	$31 \leq T \leq 60$	$61 \leq T \leq 150$	$151 \leq T$
	Average returns			
Moneyness Bin				
1	-30.6	-12.0	-4.85	-0.680
2	-21.4	-8.26	-3.23	0.256
3	-10.7	-6.51	-2.35	-1.35
	Loadings on the DRP Factor			
Moneyness Bin				
1	-949.587	-322.355	-130.370	-2.685
	(-1.3299)	(-0.6511)	(-0.4604)	(-0.0185)
2	-1384.944	-416.051	-233.972	-51.447
	(-1.6121)	(-1.0182)	(-1.0397)	(-0.4739)
3	-911.947	-351.320	-178.894	-104.671
	(-1.6349)	(-1.2811)	(-0.9536)	(-1.0393)
	Loadings on the VIX Factor			
Moneyness Bin				
1	3.5172	2.4808	1.6232	0.7957
	(7.6207)	(9.7348)	(11.1266)	(10.4411)
2	3.1595	1.8663	1.1921	0.5911
	(5.6895)	(8.8736)	(10.2821)	(10.5946)
3	1.0469	0.9039	0.6728	0.4164
	(2.9037)	(6.4027)	(6.9616)	(8.0548)

Table 14: **Index Put Options Portfolios: Time-series Regressions** This table reports the results of the regressions of the excess realized returns of 12 index put options portfolios sorted on moneyness and maturity on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$, the volatility index return, VIX and the default risk premia factor returns, DRP: January 2001 to December 2004, 206 weeks. Specifically, for each portfolio we estimate the following regressions: $R(t)-RF(t) = \alpha+\beta_{RMO}RMO(t)+\beta_{SMB}SMB(t)+\beta_{HML}HML(t)+\beta_{UMD}UMD(t)+\beta_{DEF}DEF(t)+\beta_{TERM}TERM(t)+\beta_{VIX}VIX(t)+\beta_{DRP}DRP(t)+\epsilon(t)$. We only report the estimates of the loadings on the VIX and DRP factors, and their corresponding t-statistics in parentheses.

$Name$	$\Delta\alpha$ Med	b_{Char} $Mean$	b_{Char} Med	b_{Char} $tStat$	b_{DRP} $Mean$	b_{DRP} Med	b_{DRP} $tStat$
λ^P	$2.1e^{-006}$	-0.0268	-0.0295	-0.7362	1.3844	0.0631	2.0675
$RecRate$	$6.1e^{-006}$	$-6.4e^{-007}$	$-1.7e^{-007}$	-0.1485	1.3806	0.0700	2.0481
$LevRatio$	$4.0e^{-006}$	$-6.6e^{-006}$	$-1.6e^{-005}$	-0.4593	1.3882	0.0625	2.0566
$Volatility$	$8.7e^{-008}$	$-1.7e^{-006}$	$-7.7e^{-006}$	-0.4480	1.3831	0.0667	2.0500

Table 15: **The Effect of the Characteristics: Time-series regressions** This table reports the results of the regressions of the changes in default risk premia of 71 firms on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$, the default factor returns, DRP, and a time-varying characteristic: January 2001 to December 2004, 209 weeks. Specifically, for each firm we estimate the following regressions: $R^u(t) = \alpha + \beta_{RMO}RMO(t) + \beta_{SMB}SMB(t) + \beta_{HML}HML(t) + \beta_{UMD}UMD(t) + \beta_{DEF}DEF(t) + \beta_{TERM}TERM(t) + \beta_{DRP}DRP(t) + \beta_{Char}\theta(t-1) + \epsilon(t)$. The first column lists the name of the characteristic, while the second column reports the change in the intercept due to the characteristic. The following columns report summary statistics (across firms) of the loadings on the characteristic and the default factor. The results for the loadings on the other systematic factors are not reported.

$E[\gamma_0]$	$E[\gamma^{RMO}]$	$E[\gamma^{SMB}]$	$E[\gamma^{HML}]$	$E[\gamma^{UMD}]$	$E[\gamma^{DEF}]$	$E[\gamma^{TERM}]$	$E[\gamma^{JDRP}]$	R^2
0.0047	-0.0000	0.0040	0.0006	0.0023	0.0008	0.0038	0.0001	0.4106
(5.0182)	(0.0000)	(2.2203)	(0.4229)	(0.8491)	(1.2516)	(0.9584)	(2.6012)	

Table 16: **The Industry Equity Portfolios: Cross-sectional regressions** This table reports the results of the weekly cross-sectional regressions of the excess realized returns of 49 industry portfolios on the loadings on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$ and the default risk premia factor returns, $JDRP$: January 2001 to December 2004, 206 weeks. Specifically, we estimate the coefficients of the following regressions (with or without the last regressor): $E[R(t)^i - RF(t)] = E[\gamma_0(t)] + E\left[\gamma^{RMO}(t)\right]\beta^i_{RMO} + E\left[\gamma^{SMB}(t)\right]\beta^i_{SMB} + E\left[\gamma^{HML}(t)\right]\beta^i_{HML} + E\left[\gamma^{JDRP}(t)\right]\beta^i_{JDRP}$. The coefficient are estimated as the averages of the weekly coefficients in the following cross-sectional regressions: $R(t)^i - RF(t) = \gamma_0 + \gamma^{RMO}(t)\beta^i_{RMO} + \gamma^{SMB}(t)\beta^i_{SMB} + \gamma^{HML}(t)\beta^i_{HML} + \gamma^{UMD}(t)\beta^i_{UMD} + \gamma^{DEF}(t)\beta^i_{DEF} + \gamma^{TERM}(t)\beta^i_{TERM} + \gamma^{JDRP}(t)\beta^i_{JDRP} + \epsilon^i(t)$. The first half of the table reports the results of the regressions without the loading on the default risk premia factor, while the second half reports the results for the full regressions. The reported intercept, slopes and R^2 are the mean values of the intercept, slopes and R^2, across weeks. The t-statistics, reported in parentheses, are the average intercepts or slopes divided by their time series standard error, after accounting for autocorrelations.

$E[\gamma_0]$	$E[\gamma^{RMO}]$	$E[\gamma^{SMB}]$	$E[\gamma^{HML}]$	$E[\gamma^{UMD}]$	$E[\gamma^{DEF}]$	$E[\gamma^{TERM}]$	$E[\gamma^{JDRP}]$	R^2
0.0018	0.0037	0.0025	0.0016	0.0059	0.0010	-0.0007	0.0001	0.6916
(1.4794)	(1.5681)	(2.8334)	(1.7347)	(2.2780)	(1.4675)	(0.1408)	(2.8336)	

Table 17: **The Size and Book-to-Market Equity Portfolios: Cross-sectional regressions** This table reports the results of the weekly cross-sectional regressions of the excess realized returns of 25 portfolios formed on size and book-to-market equity on the loadings on the stock market returns, RMO, SMB, HML, UMD, the corporate bond market return DEF, the treasury bond market return $TERM$ and the default risk premia factor returns, DRP: January 2001 to December 2004, 206 weeks. Specifically, we estimate the coefficients of the following regressions (with or without the last regressor): $E[R(t)^i - RF(t)] = E[\gamma_0(t)] + E\left[\gamma^{RMO}(t)\right]\beta^i_{RMO} + E\left[\gamma^{SMB}(t)\right]\beta^i_{SMB} + E\left[\gamma^{HML}(t)\right]\beta^i_{HML} + E\left[\gamma^{JDRP}(t)\right]\beta^i_{JDRP}$. The coefficient are estimated as the averages of the weekly coefficients in the following cross-sectional regressions: $R(t)^i - RF(t) = \gamma_0 + \gamma^{RMO}(t)\beta^i_{RMO} + \gamma^{SMB}(t)\beta^i_{SMB} + \gamma^{HML}(t)\beta^i_{HML} + \gamma^{UMD}(t)\beta^i_{UMD} + \gamma^{DEF}(t)\beta^i_{DEF} + \gamma^{TERM}(t)\beta^i_{TERM} + \gamma^{JDRP}(t)\beta^i_{JDRP} + \epsilon^i(t)$. The first half of the table reports the results of the regressions without the loading on the default risk premia factor, while the second half reports the results for the full regressions. The reported intercept, slopes and R^2 are the mean values of the intercept, slopes and R^2, across weeks. The t-statistics, reported in parentheses, are the average intercepts or slopes divided by their time series standard error, after accounting for autocorrelations.

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