The Determinants of Credit Spread Changes

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ABSTRACT

Using dealer's quotes and transactions prices on straight industrial bonds, we investigate the determinants of credit spread changes. Variables that should in theory determine credit spread changes have rather limited explanatory power. Further, the residuals from this regression are highly cross-correlated, and principal components analysis implies they are mostly driven by a single common factor. Although we consider several macro-economic and financial variables as candidate proxies, we cannot explain this common systematic component. Our results suggest that monthly credit spread changes are principally driven by local supply/demand shocks that are independent of both credit-risk factors and standard proxies for liquidity.

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The relation between stock and bond returns has been widely studied at the aggregate level (see, for example, Campbell and Ammer (1993), Keim and Stambaugh (1986), Fama and French (1989), and Fama and French (1993)). Recently, a few studies have investigated that relation at both the individual firm level (see, for example, Kwan (1996)) and portfolio level (see, for example, Blume, Keim and Patel (1991), and Cornell and Green (1991)). These studies focus on corporate bond returns, or yield changes. The main conclusions of these papers are: (1) high-grade bonds behave like Treasury bonds, and (2) low-grade bonds are more sensitive to stock returns.

The implications of these studies may be limited in many situations of interest, however. For example, hedge funds often take highly levered positions in corporate bonds while hedging away interest rate risk by shorting treasuries. As a consequence, their portfolios become extremely sensitive to changes in credit spreads rather than changes in bond yields. The distinction between changes in credit spreads and changes in corporate yields is significant: while an adjusted R^2 of 60 percent is obtained when regressing high-grade bond yield changes on Treasury yield changes and stock returns (see Kwan (1996)) we find that the R^2 falls to five percent when the dependent variable is credit spread changes. Hence, while much is known about yield changes, we have very limited knowledge about the determinants of credit spread changes.

Below, we investigate the determinants of credit spread changes. From a contingent-claims, or no-arbitrage standpoint, credit spreads obtain for two fundamental reasons: 1) there is a risk of default, and 2) in the event of default, the bondholder receives only a portion of the promised payments. Thus, we examine how changes in credit spreads respond to proxies for both changes in the probability of future default and for changes in the recovery rate.

Separately, recent empirical studies find that the corporate bond market tends to have relatively high transactions costs and low volume.[1] These findings suggest looking beyond the pure contingent-claims viewpoint when searching for the determinants of credit spread changes, since one might expect to observe a liquidity premium. Thus, we also examine the extent to which credit spread changes can be explained by proxies for liquidity changes.

Our results are, in summary: although we consider numerous proxies that should measure both changes in default probability and changes in recovery rate, regression analysis can only explain about 25 percent of the observed credit spread changes. We find, however, that the residuals from these regressions are highly cross-correlated, and principal components analysis implies that they are mostly driven by a single common factor. An important implication of this finding is that if any explanatory variables have been omitted, they are likely not firm-specific. We therefore re-run the regression, but this time include several liquidity, macroeconomic, and financial variables as candidate proxies for this factor. We cannot, however, find any set of variables that can explain the bulk of this common systematic factor.

Our findings suggest that the dominant component of monthly credit spread changes in the corporate bond market is driven by local supply/demand shocks that are independent of both changes in credit-risk and typical measures of liquidity. We note that a similar, but significantly smaller effect has been documented in the mortgage backed (Ginnie Mae) securities market by Boudoukh, Richardson, Stanton, and Whitelaw (1997), who find that a 3-factor model explains over 90 percent of Ginnie Mae yields, but that the remaining variation apparently cannot be explained by the changes in the yield curve.[2] In contrast, our multiple-factor model explains only about one-quarter of the variation in credit spreads, with most of the remainder attributable to a single systematic factor. Similarly, Duffie and Singleton (1999) find that both credit-risk and liquidity factors are necessary to explain innovations in U.S. swap rates. However, when analyzing the residuals they are unable to find explanatory factors. They conclude that swap market-specific supply/demand shocks drive the unexplained changes in swap rates.

Existing literature on credit spread changes is limited.[3] Pedrosa and Roll (1998) document considerable co-movement of credit spread changes among index portfolios of bonds from various industry, quality, and maturity groups. Note that this result by itself is not surprising, since theory predicts that all credit spreads *should* be affected by aggregate variables such as changes in the interest rate, changes in business climate, changes in market volatility, etc. The particularly surprising aspect of our results is that, after controlling for these aggregate determinants, the systematic movement of credit spread changes still remains, and indeed, is the dominant factor. Brown (2000) investigates credit spread innovations at the portfolio level. Although the focus of his paper differs from ours, he also finds considerable evidence that a large portion of credit spread changes is due to non-credit risk factors.

The rest of the paper is organized as follows. In Section I, we examine the theoretical determinants of credit spread changes from a contingent-claims framework. In Section II, we discuss the data and define the proxies used. In Section III, we analyze our results. In Section IV, we provide evidence for the robustness of our results on several fronts. First, we repeat the analysis using transactions (rather than quotes) data to obtain credit spread changes. Second, we consider a host of new explanatory variables that proxy for changes in liquidity and other macro-economic effects. Finally, we perform a regression analysis on simulated data to demonstrate that our empirical findings are not being driven by the econometric techniques used. We conclude in Section V.

I. Theoretical Determinants of Credit Spread Changes

So-called structural models of default provide an intuitive framework for identifying the determinants of credit spread changes.[4] These models build on the original insights of Black and Scholes (1973), who demonstrate that equity and debt can be valued using contingent-claims analysis. Introduced by Merton (1974) and further investigated by, among others, Black and Cox (1976), Leland (1994), Longstaff and Schwartz (1995), Bryis and de Varenne (1997), and Collin-Dufresne and Goldstein (2000), structural models posit some firm value process, and assume that default is triggered when the firm value falls below some threshold. This default threshold is a function of the amount of debt outstanding. In structural models, holding a debt claim is thus analogous to holding a similar risk-free debt claim and having sold to equity holders an option to put the firm at the value of the risk-free claim.[5]

Mathematically, contingent-claims pricing is most readily accomplished by pricing derivatives under the so-called risk-neutral measure, where all traded securities have an expected return equal to the risk-free rate (see Cox and Ross (1976) and Harrison and Kreps (1979)). In particular, the value of the debt claim is determined by computing its expected (under the risk-neutral measure) future cash flows discounted at the risk-free rate.

As the credit spread $CS(t)$ is uniquely defined through: (1) the price of a debt claim, (2) this debt claim's contractual cash flows, and (3) the (appropriate) risk-free rate, we can write $CS(t) = CS(V_t, r_t, \{X_t\})$, where V is firm value, r is the spot rate, and $\{X_t\}$ represents all of the other "state variables" needed to specify the model.[6] Since credit spreads are uniquely determined given the current values of the state variables, it follows that credit spread *changes* are determined by changes in these state variables. Hence, structural models generate predictions for what the theoretical determinants of credit spread changes should be, and moreover offer a prediction for whether changes in these variables should be positively or negatively correlated with changes in credit spreads. We discuss these proposed determinants individually.

1. Changes in the Spot Rate

   As pointed out by Longstaff and Schwartz (1995), the static effect of a higher spot rate is to increase the risk-neutral drift of the firm value process. A higher drift reduces the incidence of default, and in turn, reduces the credit spreads. This prediction is borne out in their data. Further evidence is provided by Duffee (1998), who uses a sample restricted to non-callable bonds and finds a significant, albeit weaker, negative relationship between changes in credit spreads and interest rates.

2. Changes in Slope of Yield Curve

   Although the spot rate is the only interest-rate-sensitive factor that appears in the firm value process, the spot rate process itself may depend upon other factors as well.[7] For example, Litterman and Scheinkman (1991) find that the two most important factors driving the term structure of interest rates are the level and slope of the term structure. If an increase in the slope of the Treasury curve increases the expected future short rate, then by the same argument as above, it should also lead to a decrease in credit spreads.

   From a different perspective, a decrease in yield curve slope may imply a weakening economy. It is reasonable to believe that the expected recovery rate might decrease in times of recession.[8] Once again, theory predicts that an increase in the Treasury yield curve slope will create a decrease in credit spreads.

3. Changes in Leverage

   Within the structural framework, default is triggered when the leverage ratio approaches unity. Hence, it is clear that credit spreads are expected to increase with leverage. Likewise, credit spreads should be a decreasing function of the firm's return on equity, all else equal.

4. Changes in Volatility

   The contingent-claims approach implies that the debt claim has features similar to a short position in a put option. Since option values increase with volatility, it follows that this model predicts credit spreads should increase with volatility. This prediction is intuitive: increased volatility increases the probability of default.

5. Changes in the Probability or Magnitude of a Downward Jump in Firm Value

   Implied volatility smiles in observed option prices suggest that markets account for the probability of large negative jumps in firm value. Thus, increases in either the probability or the magnitude of a negative jump should increase credit spreads.

6. Changes in the Business Climate

   Even if the probability of default remains constant for a firm, changes in credit spreads can occur due to changes in the expected recovery rate. The expected recovery rate in turn should be a function of the overall business climate.[9]

II. Data

Our first objective is to investigate how well the variables identified above explain observed changes in credit spreads. Here, we discuss the data used for estimating both credit spreads and proxies for the explanatory variables.

1. Credit Spreads

   The corporate bond data are obtained from Lehman Brothers via the Fixed Income (or Warga) Database. We use only quotes on non-callable, non-puttable debt of industrial firms; quotes are discarded whenever a bond has less than four years to maturity. Monthly observations are used for the period July 1988 through December 1997. Only observations with actual quotes are used, since it has been shown by Sarig and Warga (1989) that matrix prices are problematic.[10]

   To determine the credit spread, CS_t^i, for bond i at month t, we use the Benchmark Treasury rates from Datastream for maturities of 3, 5, 7, 10, and 30 years, and then use a linear interpolation scheme to estimate the entire yield curve. Credit spreads are then defined as the difference between the yield of bond-i and the associated yield of the Treasury curve at the same maturity.

2. Treasury Rate Level

   We use Datastream's monthly series of 10-year Benchmark Treasury rates, r_t^{10}. To capture potential non-linear effects due to convexity, we also include the squared level of the term structure, $(r_t^{10})^2$.

3. Slope of Yield Curve

   We define the slope of the yield curve as the difference between Datastream's 10-year and 2-year Benchmark Treasury yields, $slope_t \equiv (r_t^{10} - r_t^2)$. We interpret this proxy as both an indication of expectations of future short rates, and as an indication of overall economic health.

4. Firm Leverage

   For each bond i, market values of firm equity from CRSP and book values of firm debt from COMPUSTAT are used to obtain leverage ratios, lev_t^i, which we define as

$$\frac{\text{Book Value of Debt}}{\text{Market Value of Equity + Book Value of Debt}}.$$

Since debt levels are reported quarterly, linear interpolation is used to estimate monthly debt figures. We note that previous studies of yield changes have often used the firm's equity return to proxy for changes in the firm's health, rather than changes in leverage. For robustness, we also use each firm's monthly equity return, ret_t^i, obtained from CRSP, as an explanatory variable.

5. Volatility

In theory, changes in a firm's future volatility can be extracted from changes in implied volatilities of its publicly traded options. Unfortunately, most of the firms we investigate lack publicly traded options.[11] Thus, we are forced to use the best available substitute: changes in the VIX index, VIX_t, which corresponds to a weighted average of eight implied volatilities of near-the-money options on the OEX (S&P 100) index.[12] These data are provided by the Chicago Board Options Exchange.

While use of VIX in place of firm-specific volatility assumes a strong positive correlation between the two, this assumption does not seem to affect our results significantly. Indeed, one of our main findings is that most of the credit spread innovation is unexplained, and that the residuals are highly correlated cross-sectionally. Note that if changes in individual firm volatility and market volatility are not highly correlated, then our proxy should bias our results away from finding residuals which are so systematic.

6. Jump Magnitudes and Probabilities

Changes in the probability and magnitude of a large negative jump in firm value should have a significant effect on credit spreads. This factor is rather difficult to proxy because historical occurrences of such jumps are rare enough to be of little value in predicting future probabilities and magnitude of such jumps. Therefore, we approach the problem using a forward-looking measure. In particular, we employ changes in the slope of the "smirk" of implied volatilities of options on S&P 500 futures to determine perceived changes in the probability of such jumps.

Options and futures prices were obtained from Bridge. Our proxy is constructed from at- and out-of-the money puts, and at- and in-the-money calls with the shortest maturity on the nearby S&P 500 futures contract. We first compute implied volatilities for each strike K using the standard Black and Scholes (1973) model. We then fit the linear-quadratic regression $\sigma(K) = a + bK + cK^2$, where K is the strike price. Our estimate of this slope, $jump_t$, is defined via $jump_t = [\sigma(0.9F) - \sigma(F)]$, where F is the at-the money strike price, which equals the current futures price. We choose to look at the implied volatility at $K = .9F$ because we do not want to extrapolate the quadratic regression beyond the region where actual option prices are most typically observed.

Note that if there is a non-negligible probability of large negative jumps in firm value, then the appropriate hedging tool for corporate debt may not be the firm's equity, but rather deep out-of-the-money puts on the firm's equity. Assuming large negative jumps in firm value are highly correlated with market crashes, we hope to capture systematic changes in the market's expectation of such events with this proxy. We expect that a steepening in the slope of the smirk will trigger an increase in credit spreads.

7. Changes in Business Climate

   We use monthly S&P 500 returns, $S\&P_t$, as a proxy for the overall state of the economy. The data are obtained from CRSP.

Table 1 summarizes the predicted sign of the correlation between changes in credit spreads and changes in the underlying variable.

INSERT TABLE I ABOUT HERE

III. The Empirical Test

A. Methodology

In addition to being non-callable and non-puttable, for an industrial bond i to enter our sample, it must have at least 25 monthly trader quotes CS_t^i over the period July 1988 through December 1997. These restrictions generate a final sample of 688 bonds from 261 different issuers. The average number of quotes per bond is 56. We define ΔCS_t^i as the difference in credit spreads between two consecutive quotes. Of the resulting observations ΔCS_t^i, 99.8 percent are from differences in credit spread quotes from consecutive months.

For each sample bond i at date t with credit spread CS_t^i we estimate the following regression:

$$\begin{aligned} \Delta CS_t^i \;=\;& \alpha + \beta_1^i \, \Delta lev_t^i + \beta_2^i \, \Delta r_t^{10} + \beta_3^i \, (\Delta r_t^{10})^2 \; + \beta_4^i \Delta slope_t \\ & + \beta_5^i \, \Delta VIX_t + \beta_6^i \, S\&P_t + \beta_7^i \, \Delta jump_t + \epsilon_t^i \, . \end{aligned} \tag{1}$$

For ease of analysis, each bond is assigned to a leverage group based on the firm's average leverage ratio for those months where the bond has quotes available. These groups have been chosen to broadly replicate the bottom four quintiles and top two deciles of the sample: under 15 percent, 15 up to 25 percent, 25 up to 35 percent, 35 up to 45 percent, 45 up to 55 percent, and 55 percent or more. In Table II, summary statistics of the distribution of coefficient estimates are presented.[13]

In Panels II and III of Table II we present our findings for short- and long-maturity subsamples. In the short subsample, quotes are discarded whenever a bond has more than nine years to maturity, and in the long subsample, quotes are discarded whenever a bond has less than 12 years to maturity. Then, in each subsample and for each bond i still having at least 25 monthly quotes CS_t^i over the period July 1988 to December 1997, we re-estimate the regression of equation (1).

INSERT TABLE II ABOUT HERE

Previous studies of corporate bonds have often used stock returns ret_t^i rather than changes in leverage to proxy for changes in the firm's health. Further, these studies have grouped bonds by rating rather than firm leverage. For robustness, we also investigate credit spread changes using this approach. We thus estimate the following regression:

$$\begin{aligned} \Delta CS_t^i \;=\; & \alpha + \beta_1^i \, ret_t^i + \beta_2^i \, \Delta r_t^{10} + \beta_3^i \, (\Delta r_t^{10})^2 + \beta_4^i \Delta slope_t \\ & + \beta_5^i \, \Delta VIX_t + \beta_6^i \, S\&P_t + \beta_7^i \, \Delta jump_t + \epsilon_t^i \end{aligned} \tag{2}$$

In Table III, summary statistics of the distribution of coefficient estimates are presented. Each bond is assigned to a rating group based on the firm's average rating in months where the bond has quotes available. The bond rating is taken as the weaker of Moody's or S&P ratings whenever both are available. Maturity subsample results are also presented in Panels II and III of Table III.

INSERT TABLE III ABOUT HERE

The results of the regressions of equations (1) and (2) are very similar. The adjusted R^2 ranges from 19 percent to 25 percent when the sample is divided only by leverage ratios (or ratings). When the sample is further divided into bins based on maturity, a wider range of adjusted R^2, 17 percent to 34 percent, is observed. The model performs worst when explaining variation in long-term, high-leverage bonds. This result turns out to be a general feature for all of the regressions we perform.

B. Results

Most of the variables investigated in the regressions (1) and (2) have some ability to explain changes in credit spreads. Further, the signs of the estimated coefficients generally agree with theory. We summarize some of the major findings below.

1. From Tables II and III respectively, both the change in leverage Δlev_t^i and the firm equity return ret_t^i are statistically significant, with predicted sign, for most groups in the multivariate analyses. The economic significance, however, is rather weak. Indeed, the factor loading on the S&P 500 return is typically several times larger than the loading on the firm's own equity return. This is the first indication that monthly changes in firm-specific attributes are not the driving force in credit spread changes.

   Sensitivity to changes in leverage also tends to increase as leverage does, but that result is more apparent in a univariate regression framework, shown in Tables IV and V. Tables IV and V also demonstrate that the apparently weak explanatory power of firm-specific variables is not due to potential collinearity with the market return $S\&P_t$.

INSERT TABLE IV ABOUT HERE

2. Consistent with the empirical findings of Longstaff and Schwartz (1995) and Duffee (1998), we find that an increase in the risk-free rate lowers the credit spread for all bonds. Furthermore, the sensitivity to interest rates increases monotonically across both leverage and rating groups. Once again, this finding can be explained by noting that an increase in drift decreases the risk-neutral probability of default, and that the closer firms are to the default threshold, the more sensitive they are to this change.

INSERT TABLE V ABOUT HERE

3. Overall, convexity and slope of the term structure are not very significant either statistically or economically. Interestingly, in the short- and long-maturity subsamples, the coefficients on convexity and slope tend to be of opposite sign.

4. The change in VIX is statistically significant. As seen in Panel II of Tables II and III, it appears to have its greatest economic impact for short maturity bonds' credit spreads. However, some of these results are clouded by collinearity between S&P 500 returns and changes in the VIX index (sample correlation $-.52$).

   To investigate further, we perform univariate regressions of credit spread changes on changes in VIX, and find strong economic significance throughout. Exploring this relation more closely,

Table VI demonstrates that credit spreads respond asymmetrically to changes in implied volatility: increases in implied volatility dramatically impact credit spreads, whereas decreases do not. This asymmetry is reminiscent of the findings of Bekaert and Wu (2000) for stock returns.

INSERT TABLE VI ABOUT HERE

5. The return of the S&P 500 is extremely significant both economically and statistically. Estimated coefficients have about the same magnitude for all groups. As expected, it has a negative impact. A return of one percent for the S&P 500 is associated with a credit spread decrease of about 1.6 basis points.

6. The change in the steepness of the S&P 500 smirk, $\Delta jump_t$, is statistically and economically significant. The sign, as expected, indicates that an increase in the market's expected probability of a negative jump (as revealed by an increase in out-of-the-money put prices) triggers an increase in credit spreads. The latter behavior is relatively homogeneous across all bond groups.[14]

7. The average RMSE is 14 basis points across all bonds. The average serial correlation of residuals is -0.2, and the average Durbin Watson statistic is 2.36, suggesting serial correlation is not affecting our results.

C. Principal Components Analysis of Residuals

Overall, the variables suggested by theory are significant both economically and statistically in explaining variations in individual firms' credit spreads. However, at most they capture only around 25 percent of the variation as measured by adjusted R^2.

To better understand the nature of the remaining variation, we undertake principal components analysis on the residuals. We assign each month's residuals to one of fifteen "bins," determined by three maturity groups ($<$ 12 years, 12-18 years, $>$ 18 years), and five leverage groups: under 15 percent, 15 up to 25 percent, 25 up to 35 percent, 35 up to 45 percent, and 45 percent or over.[15] For each bin, we compute an average residual, and then extract the principal components of the covariance matrix of these residuals.

The results reveal that over 75 percent of the variation is due to the first component. Note that this first component is approximately an equally-weighted portfolio across quality and maturity groups. This result indicates that credit spread changes contain a large systematic component that lies outside of the structural model framework. Further, it implies that the low average adjusted R^2 is likely not due to noisy data, but rather to a systematic effect.

The second principal component explains an additional six percent of the remaining variation. The weights of the eigenvector are short in high-leverage debt and long in low-leverage debt. The first two principal components are displayed in Columns 3 and 4 of Table VII. Similar (unreported) results obtain when the analysis is repeated using maturity and rating bins.

INSERT TABLE VII ABOUT HERE

IV. Robustness

So far, we have only considered as regressors those factors suggested by traditional models of credit risk. If this list of factors were comprehensive, then our findings would suggest that to a large extent the corporate bond market is segmented from the equity and Treasury markets. That is, these markets would seem to be driven by different aggregate risk factors. If this conclusion holds, then using traditional models of credit risk to price and, especially, to hedge risky debt is bound to be unsuccessful. Of course, another possibility is that we have omitted important systematic explanatory variables.

In this section, we investigate the robustness of our results along several dimensions. First, we rerun the analysis of Section III.A. using transactions data. Second, we include numerous additional explanatory variables. Finally, we address the possible concern that our regression generally presumes the independent variables affect credit spread changes in a linear fashion, whereas theory predicts a non-linear relation. We perform a simulation to demonstrate that the enforced linearity of our regressions does not spuriously generate the results.

A. Transaction Prices versus Bids

Our findings in the previous section are based on dealer quotes rather than actual transaction prices. It is conceivable that the limited explanatory power that we observe, especially for the firm-specific regressors, is due to the way these bid quotes are updated by traders. In particular, these bid quotes may be slow to respond to changes in firm stock price or leverage, and thus our results may be an artifact of a "bid factor" or a "Lehman factor."[16]

There are several reasons to believe this is not the case. First, in a previous event study, Warga and Welch (1993) find that the Lehman dealer-quotes react immediately to leveraged buyouts. We also note that Lehman Brothers bears a fiduciary responsibility for the accuracy of their quotes on bonds having membership in one their bond market indices. Thus, following Elton et al. (1999), we re-run the regression (1) using only the sub-sample of quotes from bonds belonging to a Lehman index at the time of the quote. Nearly identical (unreported) results are obtained.

We further bolster support for our findings by repeating the above regressions using credit spread changes obtained from actual transactions data. Bond yields were hand-collected from the Mergent (formerly Moody's) Bond Record from January 1991 to December 1998. Of the 40 bonds so collected, 29 bonds remained after restricting the sample to those bonds having at least 25 monthly quotes and at least four years to maturity at the time of each quote. Of the bond quotes remaining in the sample, 77 percent were from actual trades (i.e., specifically labeled "sale" rather than "bid"). The results of estimating (2) on this sample are shown in Table VIII. It is interesting to note that, although the average adjusted R^2 increases somewhat, the explanatory power of the firm-specific proxy remains insignificant.

INSERT TABLE VIII ABOUT HERE

B. Additional Variables

To further substantiate our claim that a significant portion of corporate bond price innovations is driven by local supply/demand shocks that cannot be hedged using instruments from other markets, we would like to show there are no obvious systematic factors that have been omitted from the right-hand side of our regressions. While there can be no complete refutation of an omitted-variables argument, we can bolster confidence in the robustness of our findings by showing they are unchanged even after including a host of additional explanatory variables in the regressions.

B.1. Methodology

To investigate the robustness of our results, we expand our regression model in equation (1) to include additional explanatory variables. Further, we test for nonlinearities by introducing quadratic and cross-terms into the regression. In addition to the seven previous variables, we include the following independent variables:

1. Measures of Changes in Liquidity

   We construct three measures of changes in liquidity:

- First, we examine the relative frequency of quotes vs. matrix prices in the Warga database, $quote_t$. That is, for each month t, we define $quote_t$ as the log-change in the ratio of the number of quotes, q_t, to the total number of reported prices, n_t, which includes matrix prices. We interpret a higher ratio of quotes as indicative of more liquidity. Hence, the expected sign of the factor loading is negative. We note, however, that this indicator is somewhat noisy because the overall scope of the database tends to increase over time.

- The second liquidity index is more general: an estimate of changes in on-the-run minus off-the-run 30-year Treasury yields, $on{\cdot}off_t$. If liquidity worsens and the gap between these two widens, this measure decreases. Hence, we expect the factor loading to be negative.

- The third index is derived from another market of corporate transactions: an estimate of changes in the difference between yields on the 10-year swap index and 10-year Treasuries, $swap_t$. The swap index yields were obtained from Datastream. If liquidity in the swap market "dries up," it seems plausible that liquidity in the corporate bond market will dry up as well. Thus, we expect the factor loading to be positive.

2. Proxy for Firm Value Process

For robustness we include both the individual firm's equity return ret_t^i and the change in leverage Δlev_t^i as independent variables. Although they are highly correlated, it is conceivable that they provide non-redundant information.

3. Nonlinear Effects

In the previous section we included as a regressor the squared-changes in the spot rate to account for convexity issues. More generally, structural models of default predict that changes in credit spreads should be nonlinear functions of changes in leverage, interest rates, etc.[17] We therefore investigate several nonlinear terms as regressors, such as squared and cubed changes in leverage, and various cross-terms of regressors, such as $(lev_{t-1}^i \times (ret_t^i)^2)$. However, we only report the results for those variables found to have statistical significance, namely, $(\Delta r^{10})^2$ and $(\Delta r^{10})^3$.

4. SMB and HML Factors

Since the S&P 500 return was found to be an important determinant of credit spread changes, we also examine other equity return systematic factors, such as the Fama and French (1996) Small-Minus-Big, smb, and High-Minus-Low, hml, factors.

5. Economic State Variables

If there is mean-reverting behavior in spot rates, leverage, volatility, or credit spreads, then the beginning-of-month levels of those variables should contain information about the current month's change in credit spreads. We thus include the date-$(t-1)$ levels of: spot rate, r^{10}_{t-1}, firm leverage, lev^i_{t-1}, VIX index, VIX_{t-1}, and the default premium, $Spread_{t-1}$ to represent the state of the corporate bond market. The latter is measured as Datastream's BBB Index Yield minus 10-year Treasury yield.

6. Leading Effects of Stocks on Bonds

Since lagged values of equity return have been documented to have impact on changes in bond yields (e.g., Kwan (1996)), we include the one-month lagged S&P 500 return r^{SP}_{t-1} as a regressor.

B.2. Results and Analysis

Incorporating the extra variables yields the following regression:

$$\begin{aligned} \Delta CS^i_t \;=\;& \alpha + \beta^i_1 \Delta lev^i_t + \beta^i_2 \Delta r^{10}_t + \beta^i_3 (\Delta r^{10}_t)^2 + \beta^i_4 \Delta slope_t + \beta^i_5 \Delta VIX_t + \beta^i_6 S\&P_t \\ &+\; \beta^i_7 \Delta jump_t + \beta^i_8 quote_t + \beta^i_9 on\cdot off_t + \beta^i_{10} swap_t + \beta^i_{11} ret^i_t + \beta^i_{12} (\Delta r^{10}_t)^3 + \beta^i_{13} smb_t \\ &+\; \beta^i_{14} hml_t + \beta^i_{15} r^{10}_{t-1} + \beta^i_{16} lev^i_{t-1} + \beta^i_{17} VIX_{t-1} + \beta^i_{18} Spread_{t-1} + \beta^i_{19} r^{SP}_{t-1} + \epsilon^i_t \,. \end{aligned} \tag{3}$$

Due to the additional regressors, we increase to 36 the minimum number of trader quote observations a bond must have in order to qualify for the sample. As in the prior analyses, we estimate this regression on each individual corporate bond credit spread time series. We report in Table IX (Table X) the average factor loadings and associated t-statistics when the bonds are divided only by leverage (ratings). Similar results are obtained when we further divide the bins up by maturity and are omitted for conciseness.

INSERT TABLE IX ABOUT HERE

The main finding of these "kitchen-sink" regressions is that, even though the added variables do contribute somewhat to our understanding of credit spread movements, they have not explained the systematic factor which was so prominent in the earlier residuals. Indeed, although the average adjusted R^2 from equation (3) has increased to approximately 34 percent, a repetition of our principal components analysis shows that the residuals are still highly cross-correlated. The first principal component explains about 59 percent of the (now smaller) remaining variation, and the corresponding eigenvector is still roughly equally weighted in all maturity and leverage (or ratings) groups. These are reported in Columns 5 and 6 of Table VII.

Thus, the additional twelve variables have rather limited explanatory power for the systematic factor that drives credit spreads changes. Our major conclusion still holds: it appears that credit spread changes of individual bonds are mostly driven by an aggregate factor that is captured neither in existing theoretical literature, nor by the "kitchen sink" regression in equation (3). Still, several of the regression results provide interesting insights about the determinants of credit spreads. We summarize these below:

INSERT TABLE X ABOUT HERE

1. Measures of Liquidity Changes

   The factor loadings for both $quote_t$ and $on{\cdot}off_t$ have a negative sign, as predicted. However, the difference between on- and off-the-run Treasury yields is both economically and statistically more significant. The factor-loading indicates that a widening of ten basis points in $on{\cdot}off_t$ is associated with an increase of about two basis points in credit spreads. This would be consistent with posited "flight to quality" effects.

   As predicted, the factor loading on the swap spread $swap_t$ is positive and statistically significant. This measure of liquidity also seems to have superior explanatory power over our other two proxies for liquidity. Still, $swap_t$ provides rather limited explanatory power for credit spread changes.

   As an example of the implications of these results, we performed a simple "out of sample" experiment. We gathered data on credit spreads, swap rates, and on-minus-off-the-run Treasury rates for late summer 1998, when the Long-Term Capital crisis severely disrupted the bond markets. During August 1998, credit spreads increased by about 34 bp for AAA and 38 bp for BBB bonds. Using our estimated coefficients on liquidity variables (swap spread and on-the-run minus off-the-run), our model can trace only about 25 percent of this variation back to changes in liquidity, mostly to the change in swap spread (which increased by 24 bp during that same month). These findings are consistent with those of Duffie and Singleton (1997), who also note that the corporate bond market is affected by forces different from those affecting the swap market.

2. Nonlinear Effects

   The cubic term in the change in interest rate is typically positive, but lacking in economic significance.

3. SMB and HML Factors

   The factor loadings on both the smb and hml factors are statistically significant for every bin, and are negative throughout. The loadings become more negative for the higher leverage bins.

4. Economic State Variables

   The coefficient on default premium levels $Spread_{t-1}$ reflects mean-reversion in credit spreads. The coefficient on the level of the risk-free rate r^{10}_{t-1} is negative and significant throughout, but this is a marginal effect. In a univariate context, reported in Table XI, the relation between changes in credit spreads and interest rate levels is uniformly positive, but there is almost no explanatory power. Finally, the coefficients on levels of leverage (lev^i_{t-1}) and VIX (VIX_{t-1}) have limited statistical significance.

INSERT TABLE XI ABOUT HERE

5. Leading Effects of Stocks on Bonds

   The coefficients on lagged S&P 500 returns are negative and are statistically significant except for higher leverage (lower rated) bonds. In terms of economic significance, the effect is smaller, roughly 30 percent of the size of the current S&P 500 return.

B.3. Additional Evidence

To further check that our observation of a systematic factor is not spurious, we repeat regression (3) with the addition of a single explanatory variable: $\Delta Spread_t$, a "market factor" for the corporate bond market which we define as the month t change in: (Datastream's BBB Index Yield minus ten-year Treasury yield). Since we have documented above a large systematic movement in credit spreads, we expect the addition of this explanatory variable to generate a very high R^2. To no surprise, the results show adjusted R^2 of over 60 percent (not reported) for the investment grade groups, and 55 percent overall.

Having included $\Delta\, Spread_t$ in the regression, we once again undertake principal components analysis of the residuals using the same methods as before. The results are telling, and are reported in Columns 7 and 8 of Table VII. The first component now accounts for only 40 percent of the (now much smaller) remaining variation, and is no longer at all equally weighted across groups. Indeed, over 63 percent of the weighting falls into a single bin.

Overall, these tests reinforce the conclusions of the previous section. In particular, there seems to exist a systematic risk factor in the corporate bond market that is independent of equity markets, swap markets, and the Treasury market and that seems to drive most of the changes in credit spreads.

C. Simulation

If the structural models of credit spreads are correct, then the change in credit spreads should be a nonlinear function of changes in maturity, leverage, and interest rates. Although our kitchen-sink regression strongly suggests that these nonlinear terms are not the cause of the relatively low R^2 obtained, here we give additional support to this claim. Further, we show that the theoretical model predicts most of the explanatory power should come from changes in firm value, in direct conflict with our findings.

Below, we construct a simulated economy generated by recently-proposed structural models of default and demonstrate that even a two-factor linear regression on this data produces a very high R^2; indeed, around 90 percent.

C.1. The Economy

The simulated economy has the following dynamics. First, under the historical measure the spot rate r_t follows the Vasicek dynamics:

$$dr_t = \kappa(\theta^P - r_t)\,dt + \sigma\,dz_1(t)\,, \tag{4}$$

where $\kappa = 0.3$, $\theta = 0.06$, $\sigma = 0.015$, $r_0 = 0.06$. In addition, to compute credit spreads we need the spot rate dynamics under the risk-neutral measure. We assume the following form:

$$dr_t = \kappa(\theta^Q - r_t)\,dt + \sigma\,dz_1^Q(t)\,, \tag{5}$$

where $\theta^Q = .09$.

We also assume firm value follows the process:

$$\frac{dV_t}{V_t} = (\mu_t - \delta)\,dt + \nu\,dz_2(t) \tag{6}$$

$$= (r_t - \delta)\,dt + \nu\,dz_2^Q(t)\,, \tag{7}$$

where $\mu_t = r_t + 0.05$, $\delta = .03$, $\nu = .2$, and $\rho = -0.2$, where ρ is defined through $dz_1(t)\, dz_2(t) = \rho\, dt$.

Given the structure above, the log-firm value $y_t \equiv \log V_t$ has the dynamics:

$$dy_t = (\mu_t - \delta - \frac{\nu^2}{2})\, dt + \nu\, dz_2(t) \quad (8)$$

$$= (r_t - \delta - \frac{\nu^2}{2})\, dt + \nu\, dz_2^Q(t)\,. \quad (9)$$

This model is consistent with both the LS model, proposed by Longstaff and Schwartz (1995), and the CG model of Collin-Dufresne and Goldstein (2000). We note, however, that the LS model assumes a constant default threshold. If this threshold is monotonic in leverage, then the LS model predicts that the expected leverage ratio decreases exponentially over time. In contrast, the CG model assumes that the log-default boundary for firm i follows the process

$$dk_t = \lambda(y_t - \nu - k_t)\, dt\,. \quad (10)$$

Defining the "log-leverage" ratio as[18]

$$\ell_t \equiv k_t - y_t\,, \quad (11)$$

its dynamics follow:

$$d\ell_t = \lambda\left(\overline{\ell} - \ell_t\right)\, dt - \nu\, dz_2(t) \quad (12)$$

$$= \lambda\left(\overline{\ell}^Q - \ell_t\right)\, dt - \nu\, dz_2^Q(t)\,, \quad (13)$$

where $\overline{\ell} \equiv -\nu + \frac{\delta + \frac{\sigma^2}{2} - \mu}{\lambda}$. That is, this model generates stationary leverage ratios. The parameters are chosen to be $\lambda = .15$, $\ell_0 = -1$, $\overline{\ell} = -1$, and $\overline{\ell}^Q \equiv -\nu + \frac{\delta + \frac{\sigma^2}{2} - r}{\lambda} = -.6$.

C.2. Data and Results

Assuming the log-leverage ratio follows this process, we first simulate 100-month sample paths for leverage and interest rates. Then, monthly credit spreads for both the LS and CG models are determined.[19] Finally, we then estimate the following regression:

$$\Delta CS_t^i = \alpha + \beta_1^i\, \Delta lev_t^i + \beta_2^i\, \Delta r_t^{10} + \epsilon_t^i\,. \quad (14)$$

The results are reported in Table XII. Several points are notable.

First, the regressions from the 100-month simulations imply that the nonlinear relationship between changes in credit spreads and changes in both interest rates and leverage ratios is not the cause of the low R^2 obtained when running regressions on actual data. Indeed, the two-factor linear regression obtains an R^2 on the order of 90 percent for both models.

Second, unreported one-factor regressions demonstrate that almost all of this explanatory power comes from the change-in-leverage factor. This result is in stark contrast to the empirical findings.[20]

Finally, the CG model exhibits less sensitivity of credit spreads to changes in firm leverage. This effect arises because in the CG model, increases in firm value are partially offset by future increases in issuances of *pari-passu* debt. This may partially explain why observed credit spreads are so insensitive to changes in leverage. Bond prices may simply reflect the fact that increases in firm value will lead to an increase in future debt issuances, and that decreases in firm value will lead to a decrease in future debt issuances.

V. Conclusion

We investigate changes in credit spreads on individual bond yields. Several surprising results are obtained.

First, we find the factors suggested by traditional models of default risk explain only about one-quarter of the variation in credit spreads as measured by the adjusted R^2. Given that the structural framework models risky debt as a derivative security which in theory can be perfectly hedged, this adjusted R^2 seems extremely low. Furthermore, principal components analysis indicates that the residuals are highly correlated, with the first principal component (which is nearly equally-weighted across all bins of bonds) capturing about 76 percent of the remaining variation. We attempt to explain this systemic factor by introducing a host of other variables as regressors. However, the added financial and economic variables provide only limited additional explanatory power.

Second, in contrast to the predictions of structural models of default, aggregate factors appear much more important than firm-specific factors in determining credit spread changes. Furthermore, changes in credit spreads are to a great extent driven by factors not associated with either the equity or Treasury markets. This has important implications for the risk-management of corporate bond portfolios.

It seems difficult to reconcile our findings with the existing models of default risk, and, in particular, with the so-called structural models, based on contingent claims analysis initiated by Merton (1974). The latter predicts a relation between credit spreads and leverage, volatility, and interest rates. Although early empirical tests of these models proved disappointing (see Jones, Mason and Rosenfeld (1984), Kim, Ramaswamy, and Sundaresan (1993)), recent extensions (e.g., Goldstein, Ju, and Leland (1998), Mella-Barral and Perraudin (1997), and Anderson and Sundaresan (1996)) have shown that introducing agency theory or dynamic capital structure decisions can help improve the fit of the *level* of the credit spread. However it seems unlikely that these extensions can generate the kind of correlation in *changes* in credit spread uncovered in our analysis.

A natural explanation for our findings is segmentation of bond and equity markets. Clearly if markets are segmented and different investors trade in bonds and stocks, then prices in those markets could be driven by independent demand/supply shocks in both markets. Notwithstanding, in that case one needs to explain why these markets are segmented, and if they are, why equity and bonds do not react to the same aggregate factors.

Could imperfections in the bond market data explain our findings? The possibility cannot be precluded completely: Although we use two independent sources of data in this study, neither one reaches the standards of quality that prevail in CRSP data for the stock markets. However, our results are qualitatively consistent with those obtained from other sources, such as the high frequency FIPS data investigated by Hotchkiss and Ronen (1999).

Could imperfections in bond market institutions—e.g., transaction costs, liquidity—explain our findings? Recent studies by Schultz (1998), Chakravarty and Sarkar (1999), and Hotchkiss and Ronen (1999) conclude that the stock and bond markets are equally adept at efficiently incorporating new information into prices (i.e., “pricing efficiency”). At the same time, they also show that liquidity (as measured by trading volume and bid-ask spread) can have major effects on bond prices. So, potentially, an aggregate factor driving liquidity in the bond market could explain the common factor we are detecting. Our measures of liquidity (the spread between on- and off-the-run Treasuries, swap spreads, and the frequency of quotes vs. matrix prices in the Warga database) may simply be inadequate at capturing this factor.

Our findings appear to highlight a shortcoming of existing theoretical models of default risk. Besides interest rates, structural models of default predict that it is firm-specific factors that drive credit spreads . That is, they uniformly predict that the explanatory power of firm-specific measures (e.g., equity return, firm leverage) should swamp those of aggregate measures (e.g., market return).[21] However, we find empirically that most of the variation in credit spreads of individual bonds is explained by an aggregate factor common to all corporate bonds. Thus, our paper suggests the need for further work on the interaction between market risk and credit risk—i.e., general equilibrium models embedding default risk.[22]

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Notes

[1] At least in the period prior to 1997. See, for example, Schultz (1999), Hotchkiss and Ronen (1999), and Chakravarty and Sarkar (1999).

[2] Their finding is unexpected since Ginnie Mae securities face no default risk but may be repaid early. If prepayment is rationally grounded in interest rates, then from a contingent-claims analysis, these bonds have prices and yields completely determined by the Treasury market.

[3] However, there are many recent papers related to credit spreads. See, for example, Elton et al. (1999), Neal, Rolph, and Morris (2000), and John, Lynch, and Puri (2000).

[4] Recently, so-called "reduced-form" models of default have been proposed to provide a simple framework for estimating credit spreads. See, for example, Jarrow and Turnbull (1995), Jarrow, Lando and Turnbull (1997), and Duffie and Singleton (1999). However, as they typically abstract from the firm value process, they are much better suited to "fitting" the observed credit spreads than they are at offering insight into the fundamental determinants of credit spreads. Duffie and Lando (1997) have attempted to unite these two approaches.

[5] Equivalently when default can occur only at one time, e.g., at the maturity of the bond in the original Merton (1974) model, then, by put-call parity, holding a debt claim is equivalent to holding the total firm and having sold to the equity holders a call option on the firm with exercise price equal to the value of the outstanding risk-free debt claim.

[6] In Merton's (1974) original model no such state variables are needed. In fact, the interest rate itself is not a state variable since Merton assumes it is constant. In more general models, however, multiple state variables might be necessary to capture, for example: multiple factor models of the term structure, stochastic volatility of the firm's asset value, time-varying recovery rates, or bankruptcy costs.

[7] There is extensive literature on multi-factor models of the term structure, e.g., Duffie (1996).

[8] Fama and French (1989) find that credit spreads widen when economic conditions are weak.

[9] Altman and Kishore (1996) find that recovery rates are time-varying.

[10] Prices in the Warga database are not all quotes— in months where no bid is posted, a matrix price is recorded instead as a "best guess." Of 1,209 bonds available with at least some concurrent stock return and leverage data, 688 have at least 25 actual monthly quotes and thus ultimately qualify for our sample.

[11] Below we document very high cross-correlations in the credit spread residuals. This strongly suggests that additional firm-specific variables will have very limited ability to explain monthly changes in credit spreads. Thus, using changes in market volatility as a proxy for changes in firm volatility does not seem to be an issue.

[12] The appropriate volatility input for structural models of default is typically that associated with the volatility of (debt + equity). We expect changes in the proposed proxy to be highly correlated with changes in this volatility.

[13] Throughout this article, reported coefficient values and their associated t-statistics are computed as follows. For each of the N_j bonds within leverage or rating group j, a regression like equation (1) is performed. The reported coefficient values are averages of the resulting N_j regression estimates for the coefficient on each variable. Associated t-statistics are calculated from the cross-sectional variation over the N_j estimates for each coefficient by dividing each reported coefficient value by the standard deviation of the N_j estimates and scaling by $\sqrt{N_j}$.

[14] Again, univariate regressions (not reported) suggest that some of the explanatory power of the change in smirk may also be captured by the S&P 500 return because of collinearity between the two variables.

[15] In this section, the two groups with the highest leverage have been combined to better equalize the population of each bin.

[16] We thank the referee for pointing this out.

[17] However, the structural models predict that the sensitivities to these higher-order terms should be significantly smaller than the sensitivity to the linear terms.

[18] Note that ℓ is the log-leverage ratio only if the default threshold is identical to the level of debt outstanding.

[19] Collin-Dufresne and Goldstein (2000) note that the proposed solution of Longstaff and Schwartz (1995) serves only as an approximation to their model. Below, we use the exact solution.

[20] That most of the explanatory power comes from changes in leverage is implied in the relative size of the t-statistics in the two-factor model.

[21] Indeed, we have justified including the S&P 500 return in our regressions as a proxy for changes in expected recovery rates, even though there is limited empirical support for such a claim.

[22] See, for example, Chang and Sundaresan (1999) for first attempts in this direction.

Table I

Explanatory Variables and Expected Signs on the Coefficients of the Regression:

$$\Delta CS_t^i = \alpha + \beta_1^i \, \Delta lev_t^i + \beta_2^i \, \Delta r_t^{10} + \beta_3^i \, (\Delta r_t^{10})^2 \, + \beta_4^i \Delta slope_t + \beta_5^i \, \Delta VIX_t + \beta_6^i \, S\&P_t + \beta_7^i \, \Delta jump_t + \epsilon_t^i.$$

Variable	Description	Predicted Sign
Δlev_t^i	Change in firm leverage ratio	+
Δr_t^{10}	Change in yield on 10-year Treasury	–
$\Delta slope_t$	Change in 10-year minus 2-year Treasury yields	–
ΔVIX_t	Change in implied volatility of S&P 500	+
$S\&P_t$	Return on S&P 500	–
$\Delta jump$	Change in slope of Volatility Smirk	+

Table II

Structural Model Determinants of Credit Spread Changes by Leverage Group

For each industrial bond i having at least 25 monthly quotes CS_t^i over the period July 1988 to December 1997, we estimate the following regression: $\Delta CS_t^i = \alpha + \beta_1^i\,\Delta lev_t^i + \beta_2^i\,\Delta r_t^{10} + \beta_3^i\,(\Delta r_t^{10})^2 + \beta_4^i \Delta slope_t + \beta_5^i\,\Delta VIX_t + \beta_6^i\,S\&P_t + \beta_7^i\,\Delta jump_t + \epsilon_t^i$. Quotes are discarded whenever a bond has less than four years to maturity. Average OLS parameter estimates are reported in Panel I. Panel II shows averages for a short maturity subsample where quotes are discarded whenever a bond has more than nine years to maturity. Panel III shows averages for a long maturity subsample where quotes are discarded whenever a bond has less than 12 years to maturity. Associated t-statistics for each average appear immediately beneath.

	Leverage Groups					
	<15%	15–25%	25–35%	35–45%	45–55%	>55%
I. All Maturities						
intercept	.022	.016	.013	.013	.010	-.002
t	8.76	10.00	6.57	4.59	2.73	-0.20
Δlev_t^i	-.005	.007	.003	.004	.008	.033
	-1.74	4.89	1.86	2.02	3.35	3.75
Δr_t^{10}	-.124	-.140	-.181	-.215	-.215	-.342
	-17.84	-30.23	-18.93	-17.63	-11.93	-6.15
$(\Delta r_t^{10})^2$	-.010	-.001	.009	.048	.004	.164
	-0.54	-0.05	0.67	2.40	0.10	2.31
$\Delta slope_t$	.006	.001	-.028	.008	.004	-.033
	0.30	0.07	-2.29	0.48	0.15	-0.73
ΔVIX_t	.001	.002	.003	-.001	.005	.001
	0.82	3.44	2.85	-0.94	2.65	0.11
$S\&P_t$	-.016	-.015	-.016	-.017	-.016	-.019
	-21.00	-29.56	-22.68	-15.60	-10.65	-6.85
$\Delta jump_t$	.004	.004	.003	.002	.004	.003
	16.86	18.50	7.76	5.83	7.87	1.88
adjusted R^2	0.244	0.23	0.211	0.216	0.197	0.192
N	100	162	138	123	91	74
II. Short Maturities Only						
intercept	.023	.019	.009	.015	.006	-.008
	10.02	9.64	2.93	3.41	1.17	-0.58
Δlev_t^i	-.003	.009	.004	.003	.002	.042
	-0.77	5.00	1.51	1.14	0.76	3.04
Δr_t^{10}	-.141	-.138	-.202	-.226	-.235	-.414
	-20.65	-19.97	-11.68	-12.10	-7.68	-4.78
$(\Delta r_t^{10})^2$	-.046	-.032	-.020	.012	-.046	.165
	-2.65	-1.97	-0.89	0.37	-0.98	1.42
$\Delta slope_t$	.043	.031	-.045	.020	.031	.005
	2.15	2.87	-1.63	0.67	0.88	0.07
ΔVIX_t	.004	.004	.005	.001	.009	.002
	2.60	3.40	3.39	0.37	3.20	0.26
$S\&P_t$	-.017	-.015	-.018	-.018	-.019	-.020
	-24.03	-22.04	-14.43	-11.25	-10.53	-4.90
$\Delta jump_t$	.005	.005	.004	.002	.005	.004
	13.52	15.04	4.70	3.15	4.91	1.63
adjusted R^2	.317	.284	.264	.248	.199	.197
N	53	91	65	64	47	46
III. Long Maturities Only						
intercept	.010	.013	.006	.014	.007	.005
	1.89	3.98	3.54	4.25	1.24	1.48
Δlev_t^i	-.008	.004	.004	.002	.015	.013
	-1.68	1.39	1.90	0.78	3.32	6.22
Δr_t^{10}	-.095	-.161	-.156	-.200	-.210	-.211
	-5.86	-18.16	-12.75	-10.34	-9.93	-8.01
$(\Delta r_t^{10})^2$	.076	.057	.056	.055	.091	.143
	1.67	2.43	3.93	2.20	1.82	5.15
$\Delta slope_t$	-.029	-.028	-.035	-.019	.003	-.088
	-0.68	-2.45	-2.68	-0.89	0.07	-3.58
ΔVIX_t	-.002	.001	.003	-.001	.002	-.002
	-1.35	0.40	1.90	-0.78	0.51	-1.49
$S\&P_t$	-.014	-.015	-.012	-.017	-.013	-.017
	-14.70	-14.00	-9.87	-11.13	-4.72	-7.98
$\Delta jump_t$	.004	.004	.003	.003	.004	.002
	9.22	10.63	6.26	4.87	7.15	3.30
adjusted R^2	.205	.213	.196	.201	.216	.191
N	33	54	50	45	33	27

Table III

Structural Model Determinants of Credit Spread Changes by Rating Group

For each industrial bond i having at least 25 monthly quotes CS_t^i over the period July 1988 to December 1997, we estimate the following regression: $\Delta CS_t^i = \alpha + \beta_1^i \, ret_t^i + \beta_2^i \, \Delta r_t^{10} + \beta_3^i \, (\Delta r_t^{10})^2 + \beta_4^i \Delta slope_t + \beta_5^i \, \Delta VIX_t + \beta_6^i \, S\&P_t + \beta_7^i \, \Delta jump_t + \epsilon_t^i$. Quotes are discarded whenever a bond has less than four years to maturity. Average OLS parameter estimates are reported in Panel I. Panel II shows averages for a short maturity subsample where quotes are discarded whenever a bond has more than nine years to maturity. Panel III shows averages for a long maturity subsample where quotes are discarded whenever a bond has less than 12 years to maturity. Associated t-statistics for each average appear immediately beneath.

	Rating Groups					
	AAA	AA	A	BBB	BB	B
I. All Maturities						
intercept	.021	.016	.011	.018	.009	-.033
t	2.89	8.17	10.78	9.44	1.82	-0.67
ret_t^i	.002	.000	-.001	-.002	-.003	-.018
	2.11	0.15	-2.67	-4.15	-4.58	-2.75
Δr_t^{10}	-.109	-.150	-.151	-.159	-.296	-.862
	-7.15	-17.99	-27.73	-26.03	-14.74	-4.36
$(\Delta r_t^{10})^2$	-.039	-.012	.037	-.014	.095	.568
	-0.52	-0.76	3.94	-1.02	2.15	1.19
$\Delta slope_t$	.042	.009	-.017	.027	-.060	.048
	0.55	0.70	-1.90	2.83	-1.92	0.36
ΔVIX_t	.002	.004	.002	.002	.000	-.029
	0.62	2.92	4.44	2.88	-0.11	-0.79
$S\&P_t$	-.016	-.015	-.014	-.014	-.023	-.043
	-14.36	-18.50	-37.00	-21.22	-9.82	-3.65
$\Delta jump_t$	.003	.004	.003	.003	.004	.005
	2.83	10.24	13.57	12.98	6.62	0.98
adjusted R^2	.222	.293	.234	.194	.197	.275
N	4	56	275	245	90	18
II. Short Maturities Only						
intercept	.031	.018	.014	.016	.007	-.041
	5.02	5.74	8.33	5.82	0.94	-0.70
ret_t^i	.000	.000	-.001	-.001	-.003	-.019
	-0.24	0.47	-2.72	-2.28	-2.70	-2.51
Δr_t^{10}	-.111	-.156	-.163	-.150	-.322	-.909
	-5.60	-14.39	-18.98	-14.76	-10.73	-3.86
$(\Delta r_t^{10})^2$	-.123	-.060	-.015	-.031	.040	.607
	-1.10	-2.65	-1.19	-1.89	0.65	1.05
$\Delta slope_t$	.168	.028	.001	.052	-.032	.072
	2.16	1.34	0.10	3.45	-0.67	0.44
ΔVIX_t	.006	.005	.006	.006	.001	-.038
	0.82	2.63	6.50	4.49	0.35	-0.87
$S\&P_t$	-.015	-.016	-.015	-.015	-.026	-.044
	-7.75	-18.37	-22.56	-18.76	-7.62	-3.31
$\Delta jump_t$	.002	.004	.003	.004	.005	.009
	0.97	6.99	8.46	8.85	4.60	1.51
adjusted R^2	.232	.341	.277	.235	.200	.301
N	2	34	139	120	56	15
III. Long Maturities Only						
intercept	.009	.014	.007	.015	.008	-.031
t	8.66	4.23	3.71	5.07	1.60	-2.61
ret_t^i	.004	-.001	.000	-.003	-.004	-.001
	9.38	-0.89	-1.25	-3.53	-3.65	-0.19
Δr_t^{10}	-.096	-.159	-.143	-.178	-.234	-.611
	-14.97	-10.33	-16.11	-18.05	-10.09	-5.61
$(\Delta r_t^{10})^2$	.074	.020	.078	.049	.176	.270
	2.66	0.87	4.35	2.63	3.48	2.06
$\Delta slope_t$	-.074	-.003	-.039	.000	-.083	-.197
	-3.24	-0.20	-2.72	0.02	-2.78	-0.88
ΔVIX_t	-.001	.003	.001	-.001	.000	.007
	-0.63	1.76	0.77	-1.14	0.02	0.83
$S\&P_t$	-.016	-.013	-.012	-.014	-.020	-.027
	-20.50	-6.57	-21.93	-13.22	-5.43	-2.49
$\Delta jump_t$	.004	.004	.003	.003	.004	-.003
	230.43	5.39	10.73	9.95	4.23	-1.71
adjusted R^2	.179	.265	.224	.180	.165	.302
N	2	16	114	79	28	3

Table IV

Relation Between Changes in Credit Spreads and Changes in Leverage

For each industrial bond i having at least 25 monthly quotes CS_t^i over the period July 1988 to December 1997, we estimate the following regression: $\Delta CS_t^i = \alpha + \beta_1^i\, \Delta lev_t^i + \epsilon_t^i$. Quotes are discarded whenever a bond has less than four years to maturity. Average OLS parameter estimates are reported in Panel I. Panel II shows averages for a short maturity subsample where quotes are discarded whenever a bond has more than nine years to maturity. Panel III shows averages for a long maturity subsample where quotes are discarded whenever a bond has less than 12 years to maturity. Associated t-statistics for each average appear immediately beneath.

	Leverage Groups					
	$<$15%	15–25%	25–35%	35–45%	45–55%	$>$55%
I. All Maturities						
intercept	.001	.000	-.003	-.004	-.005	.005
t	1.21	-0.01	-3.38	-2.54	-2.46	1.36
Δlev_t^i	.012	.015	.010	.011	.016	.035
	3.87	10.30	7.07	5.38	7.17	5.21
adjusted R^2	.003	.028	.011	.032	.051	.065
N	100	162	138	123	91	74
II. Short Maturities Only						
intercept	-.004	-.002	-.008	-.007	-.015	.006
	-3.24	-1.98	-5.40	-2.86	-4.40	1.03
Δlev_t^i	.016	.016	.014	.011	.013	.042
	3.45	10.03	5.19	5.27	5.55	4.25
adjusted R^2	.001	.025	.024	.033	.030	.072
N	53	91	65	64	47	46
III. Long Maturities Only						
intercept	.001	.000	-.001	.000	.003	.000
	1.06	-0.15	-1.10	-0.24	0.95	-0.12
Δlev_t^i	.006	.012	.007	.007	.021	.018
	1.60	4.10	4.47	2.47	4.20	7.66
adjusted R^2	-.008	.016	.005	.021	.084	.055
N	33	54	50	45	33	27

Table V

Relation Between Changes in Credit Spreads and Firm Equity Returns

For each industrial bond i having at least 25 monthly quotes CS_t^i over the period July 1988 to December 1997, we estimate the following regression: $\Delta CS_t^i = \alpha + \beta_1^i \, ret_t^i + \epsilon_t^i$. Quotes are discarded whenever a bond has less than four years to maturity. Average OLS parameter estimates are reported in Panel I. Panel II shows averages for a short maturity subsample where quotes are discarded whenever a bond has more than nine years to maturity. Panel III shows averages for a long maturity subsample where quotes are discarded whenever a bond has less than 12 years to maturity. Associated t-statistics for each average appear immediately beneath.

	Rating Groups					
	AAA	AA	A	BBB	BB	B
I. All Maturities						
intercept	.007	.003	.003	.001	-.007	.022
t	3.67	4.76	6.06	1.10	-2.69	1.41
ret_t^i	-.003	-.003	-.003	-.004	-.005	-.014
	-1.97	-7.78	-14.70	-9.22	-7.39	-3.82
adjusted R^2	.004	.018	.030	.040	.047	.115
N	4	56	275	245	90	18
II. Short Maturities Only						
intercept	.009	.002	.001	-.004	-.015	.020
	2.34	2.04	1.58	-2.73	-3.76	1.10
ret_t^i	-.005	-.003	-.003	-.003	-.005	-.015
	-2.62	-5.92	-12.97	-8.82	-4.86	-3.86
adjusted R^2	.027	.019	.033	.035	.033	.116
N	2	34	139	120	56	15
III. Long Maturities Only						
intercept	.004	.003	.002	.003	.000	-.011
	21.06	3.88	2.52	2.23	0.12	-0.97
ret_t^i	-.001	-.002	-.002	-.004	-.005	-.001
	-3.96	-5.36	-8.03	-5.21	-5.66	-0.18
adjusted R^2	-.016	.004	.011	.050	.067	.079
N	2	16	114	79	28	3

Table VI

Relation Between Changes in Credit Spreads and Changes in VIX by Leverage Group

For each industrial bond i having at least 25 monthly quotes CS_t^i over the period July 1988 to December 1997, we estimate the following regression: $\Delta CS_t^i = \alpha + \beta_1^i \, \Delta VIX_t \, d_t + \beta_1^i \, \Delta VIX_t \, (1 - d_t) + \epsilon_t^i$, where $d_t = 1$ if $\Delta VIX_t > 0$, and 0 otherwise. Quotes are discarded whenever a bond has less than four years to maturity. Average OLS parameter estimates are reported in Panel I. Panel II shows averages for a short maturity subsample where quotes are discarded whenever a bond has more than nine years to maturity. Panel III shows averages for a long maturity subsample where quotes are discarded whenever a bond has less than 12 years to maturity. Associated t-statistics for each average appear immediately beneath.

	Leverage Groups					
	$<$15%	15–25%	25–35%	35–45%	45–55%	$>$55%
I. All Maturities						
intercept	-.015	-.019	-.017	-.020	-.021	-.021
t	-8.76	-8.30	-7.40	-5.75	-4.46	-3.36
positive ΔVIX_t	.014	.016	.014	.013	.016	.026
	20.27	14.58	11.54	8.49	7.72	7.55
negative ΔVIX_t	.001	.001	.003	.001	.005	.005
	1.15	0.32	2.15	0.27	2.09	1.34
adjusted R^2	.041	.048	.029	.023	.029	.030
N	100	162	138	123	91	74
II. Short Maturities Only						
intercept	-.021	-.022	-.027	-.033	-.039	-.022
	-7.99	-9.19	-5.76	-6.40	-4.71	-2.26
positive ΔVIX_t	.018	.018	.019	.019	.024	.031
	14.50	17.89	7.27	8.70	7.65	5.89
negative ΔVIX_t	.004	.004	.003	-.001	.005	.010
	1.73	2.12	1.46	-0.22	1.28	1.74
adjusted R^2	.075	.060	.046	.045	.054	.043
N	53	91	65	64	47	46
III. Long Maturities Only						
intercept	-.016	-.022	-.007	-.008	-.004	-.023
	-5.64	-4.10	-3.90	-1.51	-0.62	-4.69
positive ΔVIX_t	.011	.014	.009	.008	.008	.013
	11.69	5.50	6.84	4.00	2.68	4.58
negative ΔVIX_t	-.002	-.005	.004	.003	.007	.000
	-1.47	-1.30	3.48	0.96	2.23	-0.18
adjusted R^2	.017	.041	.011	.015	.009	.013
N	33	54	50	45	33	27

Table VII

Principal Components

For each industrial bond i having at least 25 monthly quotes CS_t^i over the period July 1988 to December 1997, we estimate equation (1): $\Delta CS_t^i = \alpha + \beta_1^i \, \Delta lev_t^i + \beta_2^i \, \Delta r_t^{10} + \beta_3^i \, (\Delta r_t^{10})^2 + \beta_4^i \Delta slope_t + \beta_5^i \, \Delta VIX_t + \beta_6^i \, S\&P_t + \beta_7^i \, \Delta jump_t + \epsilon_t^i$. For each industrial bond i having at least 36 monthly quotes CS_t^i over the period July 1988 to December 1997, we estimate equation (3): $\Delta CS_t^i = \alpha + \beta_1^i \Delta lev_t^i + \beta_2^i \, \Delta r_t^{10} + \beta_3^i \, (\Delta r_t^{10})^2 + \beta_4^i \Delta slope_t + \beta_5^i \, \Delta VIX_t + \beta_6^i \, S\&P_t + \beta_7^i \, \Delta jump_t + \beta_8^i \, quote_t + \beta_9^i \, on{\cdot}off_t + \beta_{10}^i \, swap_t + \beta_{11}^i \, ret_t^i + \beta_{12}^i (\Delta r_t^{10})^3 + \beta_{13}^i smb_t + \beta_{14}^i hml_t + \beta_{15}^i r_{t-1}^{10} + \beta_{16}^i lev_{t-1}^i + \beta_{17}^i VIX_{t-1} + \beta_{18}^i \, Spread_{t-1} + \beta_{19}^i \, r_{t-1}^{SP} + \epsilon_t^i$. Finally, for the '$\Delta BBB$' regression, we add to equation (3) changes in the BBB credit spread as reported in Datastream, and then re-run the regression. Quotes are discarded whenever a bond has less than four years to maturity. The residuals are then assigned to one of 15 analysis bins based on maturity and firm leverage. Short maturity is under 12 years; Medium maturity is 12 to 18 years; Long maturity is over 18 years. Monthly averages for each bin are calculated, and then the principal components of the resulting covariance matrix are extracted. The first two vectors for each set of residuals are reported below, along with the percent of the remaining variance associated with each vector. The adjusted R^2 and unexplained variance from each regression are reported as well.

Analysis Bins		Principal Components					
		Equation (1) Residuals		Equation (3) Residuals		ΔBBB Residuals	
Maturity	Leverage	First	Second	First	Second	First	Second
Short	Low	.23803	.11438	.24327	$-.05569$	.15353	.21257
Short	2	.24508	.12107	.25666	$-.05202$	.16936	.21077
Short	3	.27665	.04722	.26324	$-.07952$	.13979	.21893
Short	4	.30059	$-.08293$	.26757	$-.04632$	.14980	.17982
Short	High	.26998	$-.63059$	.26441	$-.01370$	.19105	.17506
Medium	Low	.23074	.28626	.25312	$-.09284$	.12572	.22903
Medium	2	.25226	.22294	.26871	$-.07669$	.14537	.21452
Medium	3	.27640	.16116	.26986	$-.10780$	.12765	.23277
Medium	4	.28481	.11761	.29077	$-.11450$	.14421	.24728
Medium	High	.25870	$-.52780$	.23424	.95794	.79434	$-.58382$
Long	Low	.23811	.23054	.25385	$-.09508$	.14877	.27150
Long	2	.22060	.13328	.21696	$-.07955$	.12553	.21473
Long	3	.23623	.11610	.23824	$-.08967$	.13327	.23880
Long	4	.25895	$-.00930$	.27148	$-.03257$	.20496	.22586
Long	High	.27196	$-.17609$	.27139	.06468	.25808	.13027
Cum. % Explained by PC		75.9	82.2	58.5	79.1	39.8	70.4
Avg. Adj. R^2 of regression		0.21		0.35		0.60	
Unexplained Variance		0.114		0.078		0.048	

Table VIII

Structural Model Determinants of Credit Spread Changes Using Transactions Data

We collected by hand from Mergent (Moody's) Bond Record a sample of 29 bonds having at least 25 monthly quotes CS_t^i over the period January, 1991, to December, 1998. For each bond i, we estimate the following regression: $\Delta CS_t^i = \alpha + \beta_1^i \, ret_t^i + \beta_2^i \, \Delta r_t^{10} + \beta_3^i \, (\Delta r_t^{10})^2 + \beta_4^i \Delta slope_t + \beta_5^i \, \Delta VIX_t + \beta_6^i \, S\&P_t + \beta_7^i \, \Delta jump_t + \epsilon_t^i$. Quotes are discarded whenever a bond has less than four years to maturity. Average OLS parameter estimates are reported in Panel I. Associated t-statistics for each average appear immediately beneath.

I. All Maturities	
intercept	-.019
t	-1.69
ret_t^i	-.001
	-0.45
Δr_t^{10}	-.809
	-19.39
$(\Delta r_t^{10})^2$	.218
	2.08
$\Delta slope_t$	.072
	0.87
ΔVIX_t	-.030
	-3.99
$S\&P_t$	-.013
	-2.36
$\Delta jump_t$	.006
	2.94
adjusted R^2	.456
N	29

Table IX

Additional Determinants of Credit Spread Changes by Leverage Group

For each industrial bond i having at least 36 monthly quotes CS_t^i over the period July 1988 to December 1997, we estimate the following regression: $\Delta CS_t^i = \alpha + \beta_1^i \Delta lev_t^i + \beta_2^i \Delta r_t^{10} + \beta_3^i (\Delta r_t^{10})^2 + \beta_4^i \Delta slope_t + \beta_5^i \Delta VIX_t + \beta_6^i S\&P_t + \beta_7^i \Delta jump_t + \beta_8^i quote_t + \beta_9^i on\cdot off_t + \beta_{10}^i swap_t + \beta_{11}^i ret_t^i + \beta_{12}^i (\Delta r_t^{10})^3 + \beta_{13}^i smb_t + \beta_{14}^i hml_t + \beta_{15}^i r_{t-1}^{10} + \beta_{16}^i lev_{t-1}^i + \beta_{17}^i VIX_{t-1} + \beta_{18}^i Spread_{t-1} + \beta_{19}^i r_{t-1}^{SP} + \epsilon_t^i$. Quotes are discarded whenever a bond has less than four years to maturity. Average OLS parameter estimates are reported in Panel I. Associated t-statistics for each average appear immediately beneath.

	Leverage Groups					
	$<$15%	15–25%	25–35%	35–45%	45–55%	$>$55%
I. All Maturities						
intercept	.452	.324	.172	.188	-.009	-.378
t	6.66	8.90	3.37	2.97	-0.10	-2.32
Δlev_t^i	-.677	1.099	.853	1.061	-.927	-.762
	-0.96	4.13	2.06	2.20	-0.75	-0.81
Δr_t^{10}	-.146	-.145	-.176	-.250	-.301	-.418
	-14.82	-18.25	-12.25	-11.29	-8.59	-5.98
$(\Delta r_t^{10})^2$	-.129	-.129	-.060	-.045	-.075	-.114
	-3.97	-10.17	-2.38	-1.36	-2.19	-1.96
$\Delta slope_t$	.074	.079	.048	.097	.060	.051
	2.99	7.60	2.96	4.21	2.08	1.07
ΔVIX_t	.001	.002	.004	.001	.015	.019
	1.12	2.24	2.43	0.30	4.61	3.33
$S\&P_t$	-.017	-.017	-.017	-.018	-.014	-.013
	-13.93	-26.73	-15.66	-9.47	-5.62	-3.22
$\Delta jump_t$	.004	.004	.004	.002	.005	.003
	11.46	14.37	6.77	3.67	7.20	2.30
$quote_t$	-.818	-.284	-.186	-.575	1.227	.144
	-2.05	-1.71	-0.55	-1.39	2.75	0.22
$\Delta on\cdot off_t$	-.219	-.173	-.155	-.246	-.173	-.244
	-4.33	-3.49	-2.56	-2.87	-1.93	-1.59
$swap_t$	.283	.409	.444	.366	.533	.675
	8.19	16.27	14.20	5.57	7.11	7.88
ret_t^I	-.091	.141	.150	.101	-.472	-.732
	-1.42	3.35	1.65	0.80	-1.47	-2.71
$(\Delta r_t^{10})^3$	-.132	-.155	-.147	-.012	.136	.439
	-2.71	-6.35	-3.18	-0.20	1.53	1.65
smb_t	.000	-.002	-.004	-.007	-.009	-.009
	-0.26	-3.31	-3.68	-4.76	-4.29	-2.15
hml_t	-.006	-.008	-.007	-.012	-.011	-.010
	-5.77	-10.17	-6.96	-6.17	-3.67	-2.49
r_{t-1}^{10}	-.024	-.020	-.021	-.026	-.036	-.020
	-4.62	-7.44	-5.16	-5.23	-5.19	-2.27
lev_{t-1}^i	.225	.139	.225	.368	.334	.653
	1.65	2.37	3.06	3.68	3.50	3.55
VIX_{t-1}	.002	.003	.006	.009	.020	.021
	1.69	3.52	4.30	2.29	5.01	2.95
$Spread_{t-1}$	-.292	-.224	-.147	-.247	-.157	-.185
	-10.21	-12.89	-5.53	-9.17	-5.28	-3.47
r_{t-1}^{SP}	-.005	-.005	-.005	-.004	-.004	-.009
	-5.29	-9.42	-5.66	-3.15	-1.95	-2.23
adjusted R^2	.395	.348	.314	.313	.301	.306
N	75	130	112	96	73	63

Table X

Additional Determinants of Credit Spread Changes by Rating Group

For each industrial bond i having at least 36 monthly quotes CS_t^i over the period July 1988 to December 1997, we estimate the following regression: $\Delta CS_t^i = \alpha + \beta_1^i \Delta lev_t^i + \beta_2^i \, \Delta r_t^{10} + \beta_3^i \, (\Delta r_t^{10})^2 + \beta_4^i \Delta slope_t + \beta_5^i \, \Delta VIX_t + \beta_6^i \, S\&P_t + \beta_7^i \, \Delta jump_t + \beta_8^i quote_t + \beta_9^i on{\cdot}off_t + \beta_{10}^i \, ret_t^i + \beta_{11}^i (\Delta r_t^{10})^3 + \beta_{12}^i smb_t + \beta_{13}^i r_{t-1}^{10} + \beta_{14}^i lev_{t-1}^i + \beta_{15}^i VIX_{t-1} + \beta_{16}^i \, Spread_{t-1} + \beta_{17}^i \, r_{t-1}^{SP} + \beta_{18}^i \, swap_t + \epsilon_t^i$. Quotes are discarded whenever a bond has less than four years to maturity. Average OLS parameter estimates are reported in Panel I. Associated t-statistics for each average appear immediately beneath.

	Rating Groups					
	AAA	AA	A	BBB	BB	B
I. All Maturities						
intercept	.277	.333	.237	.238	-.306	-.432
	0.59	4.69	8.19	5.32	-2.21	-0.79
Δlev_t^i	.234	.835	.834	.382	-.828	-5.639
	0.10	0.85	3.38	1.29	-0.75	-0.96
Δr_t^{10}	-.108	-.152	-.149	-.202	-.419	-1.033
	-2.18	-13.75	-19.77	-17.82	-8.30	-5.22
$(\Delta r_t^{10})^2$	-.151	-.125	-.073	-.107	-.062	-.225
	-6.01	-6.74	-6.27	-6.29	-0.92	-1.06
$\Delta slope_t$	.086	.087	.063	.094	.038	-.058
	1.54	5.82	5.01	7.78	0.82	-0.48
ΔVIX_t	.001	.004	.002	.003	.019	.060
	0.18	2.61	2.44	2.09	3.45	4.08
$S\&P_t$	-.019	-.015	-.016	-.018	-.021	.011
	-21.30	-12.49	-25.31	-17.85	-5.16	1.06
$\Delta jump_t$	.005	.004	.003	.004	.005	-.002
	3.16	7.89	12.13	9.86	4.51	-1.10
$quote_t$	1.749	-1.053	-.083	-.292	1.059	-2.567
	1.66	-2.37	-0.60	-1.39	1.31	-1.03
$\Delta on{\cdot}off_t$	-.249	-.122	-.204	-.207	-.218	-.044
	-1.76	-2.05	-4.68	-4.61	-1.50	-0.11
$swap_t$	.330	.366	.392	.449	.527	.950
	2.56	10.11	22.86	13.65	4.47	4.00
ret_t^I	.046	-.001	.148	-.069	-.553	-2.026
	0.26	-0.01	3.23	-0.91	-1.80	-1.38
$(\Delta r_t^{10})^3$	-.344	-.184	-.113	-.019	.087	1.816
	-2.03	-5.17	-3.80	-0.46	0.71	1.62
smb_t	.002	.000	-.003	-.009	-.001	-.021
	1.16	-0.37	-4.90	-8.07	-0.24	-2.49
hml_t	-.005	-.006	-.006	-.014	-.010	.018
	-1.01	-5.30	-9.32	-12.05	-2.71	1.86
r_{t-1}^{10}	-.029	-.016	-.018	-.031	-.024	-.054
	-1.78	-3.42	-7.58	-10.56	-2.54	-1.61
lev_{t-1}^i	.980	.281	.160	.304	.567	.902
	5.03	1.59	3.10	5.81	3.32	1.52
VIX_{t-1}	.001	.004	.004	.006	.029	.051
	0.26	2.54	5.86	4.81	3.61	2.95
$Spread_{t-1}$	-.313	-.265	-.204	-.193	-.158	-.526
	-2.23	-7.93	-16.37	-11.04	-2.65	-4.19
r_{t-1}^{SP}	-.006	-.004	-.004	-.004	-.010	-.002
	-5.92	-4.85	-9.00	-6.34	-2.61	-0.27
adjusted R^2	.400	.421	.343	.327	.224	.352
N	4	47	233	183	69	13

Table XI

Relation Between Changes in Credit Spreads and Interest Rate Levels

For each industrial bond i having at least 25 monthly quotes CS_t^i over the period July 1988 to December 1997, we estimate the following regression: $\Delta CS_t^i = \alpha + \beta_1^i \, r_{t-1}^{10} + \epsilon_t^i$. Quotes are discarded whenever a bond has less than four years to maturity. Average OLS parameter estimates are reported in Panel I. Panel II shows averages for a short maturity subsample where quotes are discarded whenever a bond has more than nine years to maturity. Panel III shows averages for a long maturity subsample where quotes are discarded whenever a bond has less than 12 years to maturity. Associated t-statistics for each average appear immediately beneath.

	Leverage Groups					
	$<$15%	15–25%	25–35%	35–45%	45–55%	$>$55%
I. All Maturities						
intercept	-.038	-.044	-.086	-.095	-.114	-.285
t	-2.56	-3.57	-4.96	-3.67	-4.01	-2.57
r_{t-1}^{10}	.006	.006	.011	.012	.015	.040
	2.51	3.50	4.74	3.33	3.53	2.62
adjusted R^2	-.016	-.012	-.010	-.008	-.008	-.008
N	100	162	138	123	91	74
II. Short Maturities Only						
intercept	-.093	-.102	-.153	-.146	-.098	-.413
	-3.96	-5.57	-4.88	-3.28	-2.06	-2.34
r_{t-1}^{10}	.013	.014	.020	.018	.010	.058
	3.67	5.51	4.61	2.96	1.47	2.38
adjusted R^2	-.014	-.015	-.008	-.009	-.014	-.010
N	53	91	65	64	47	46
III. Long Maturities Only						
intercept	.002	.011	-.028	-.081	-.104	-.088
	0.11	0.46	-1.02	-1.30	-3.42	-2.88
r_{t-1}^{10}	.000	-.002	.003	.009	.015	.012
	-0.10	-0.51	0.89	1.19	3.34	2.71
adjusted R^2	-.014	-.012	-.010	-.003	-.008	-.011
N	33	54	50	45	33	27

Table XII

Determinants of Credit Spread Changes in Simulated Economies

For bonds simulated for 100 months in the LS and CG model economies, we estimate the following regression: $\Delta CS_t^i = \alpha + \beta_1^i \Delta lev_t^i + \beta_2^i \Delta r_t^{10} + \epsilon_t^i$. Average OLS parameter estimates are reported below. Associated t-statistics appear immediately beneath.

	Model Economy	
	LS	CG
Δlev_t^i	6.45	2.88
t	38.24	27.25
Δr_t^{10}	-.151	-.097
	-7.14	-7.35
adjusted R^2	.94	.89