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Estimating Credit Risk Premia

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Abstract

This paper investigates the nature of the credit risk premium adjustments in the Jarrow-Lando-Turnbull model of credit risk spreads. The adjustments relate the equivalent martingale measures to the empirical measures of unconditional transition probabilities. We provide a modified version of the risk adjustment that allows a linear partition of the credit spread into an unconditional default component, a recovery component, and the risk premium adjustment. The risk adjustments are related to conditional default risk, illiquidity risk, and other factors not related to recovery effects. The log-transform of these risk adjustments can be specified as linear regressions on a set of macroeconomic variables. Some new insights are gained pertaining to these conditional risks such as a typical upward sloping term structure and sensitivity to short-term Treasury rates and increasing forward rates. The conditional risks appear to be insensitive to market returns.

Keywords: Credit Spreads, Risk Adjustments, No-arbitrage equilibrium, Conditional Risks

JEL Codes: G120 G210 G330

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1 Introduction

Credit spread is the additional yield required by the market in holding credit-risky corporate bond relative to Treasury bond of the same maturity. The spread is typically computed as the difference between the spot rates of the corporate and the Treasury discount bonds. Understanding the economic factors that influence credit spread dynamics is an important issue in finance as the spreads determine relative corporate bond prices.

Duffee (1998) indicates that 3-month Treasury rate and also Treasury slope (30-year Treasury bond rate less 3-month bill rate) have negative impact on credit spread changes. Izvorski (1997) shows the importance of recovery ratios and survival times in determining corporate bond prices. Similar variables were used by Fons (1994) to explain credit spread term structures. Janosi, Jarrow and Yildrim (2002) show components of expected losses and liquidity discounts in debt prices and thus credit spreads. Bevan and Garzarelli (2000) relate credit spread movements to business cycle variables. Collin-Dufresne, Goldstein and Martin (2001) find credit-related economic variables such as leverage, liquidity proxies, and volatility that help to explain a part of credit spread movement. They also show S&P 500 market returns to be negatively correlated with credit spreads. Elton, Gruber, Agrawal and Mann (2001) suggest that unexplained spread movements are related to systematic risks in the economy. This idea was also contained in Pedrosa and Roll (1998). Wu and Yu (1996) consider preference factor such as risk aversion to also impact on debt yield dynamics. Yu (2002) shows empirically that firms with higher accounting disclosures tend to have lower credit spreads. Duffie and Lando (2001) provide the theory for the impact of accounting information on credit spreads.

Huang and Huang (2002) find that default risk accounts for only a small fraction of the credit spread. Other factors would include illiquidity, and call and conversion features. Collin-Dufresne, Goldstein, and Martin (2001) find that traditional model factors could explain only a small fraction of the variation in credit spreads. Thus it is interesting to examine the credit spread and its constitution.

The Jarrow, Lando, and Turnbull (1997) (JLT) model of credit risk spread is a no-arbitrage equilibrium pricing model, and is interesting because it explicitly incorporates the empirical probability of default of a corporate bond into its equilibrium pricing. Other studies such as Elton, Gruber, Agrawal, and Mann (2001) employ additional risk-neutral assumptions in order to price corporate credit risk premium, or else rely on regression models without explicitly characterizing the equilibrium.

In this paper we investigate the nature of the credit risk premium adjustments in the JLT model of credit risk spreads. The adjustments relate the equivalent martingale measures to the empirical measures of unconditional transition probabilities. We provide a modified version of the risk adjustment that allows a linear partition of the credit risk spread into an unconditional default component, a recovery component, and the risk premium adjustments. The risk adjustments are related to conditional default risk, illiquidity risk, and other factors not related to recovery effects. A log-transform of these risk adjustments can be specified as linear regressions on a set of macroeconomic variables.

This extension of the JLT model is interesting in at least two aspects. We show how to motivate the intuition behind the risk premium adjustments and relate it to credit spread studies where spreads are regression relations on macroeconomic and other explanatory variables. These adjustments are not necessarily deterministic functions.

Our modified risk adjustments also allow a parsimonious term structure of conditional default, illiquidity and other conditional risks to be estimated and tested. As we eliminate bonds that introduce convertibility complications, and also employ the JLT procedure to remove the callability effect, the other conditional risks in addition to the mostly illiquidity risk are likely to be risks related to diversification aspects of the bond in a portfolio context.

In the next section, we show the theoretical construction of the credit risk spread components in the context of the JLT model. We also show two propositions about the properties of the modified risk adjustments. These properties will be reviewed in the empirical examination in section 3. In section 3, we discuss the data and method of setting up the regression analysis. The empirical results of the regressions are then reported and discussed. Section 4 provides the conclusions.

2 Credit risk Premia

Jarrow, Lando, and Turnbull (1997) and Kijima and Komoribayashi (1997) employ a time-homogeneous finite state space Markov chain to represent the empirical unconditional transition probabilities of bond credit ratings. This is the per period transition matrix Q under the empirical probability, and is expressed as

$$Q = \begin{pmatrix} q_{11} & q_{12} & \dots & q_{1K} \\ q_{21} & q_{22} & \dots & q_{2K} \\ \vdots & \vdots & q_{ij} & \vdots \\ q_{K-1,1} & q_{K-1,2} & \dots & q_{K-1,K} \\ 0 & 0 & \dots & 1 \end{pmatrix}.$$

The element q_{ij} represents the transition probability of a corporate bond migrating from state or rating i at time t to state or rating j in the next period at time $t+1$, for each t. The transition probabilities satisfy the following three conditions. (i) $0 \leq q_{ij} \leq 1\,, \forall i, j$, (ii) $\sum_{j=1}^{K} q_{ij} = 1\,, \forall i$, and (iii) $q_{iK} < 1\,, \forall i \neq K$. State K is an absorbing state representing bond default. Condition (iii) is really a mild assumption to ensure that ex-ante no bonds would default with probability 1 in finite horizon.

For pricing of credit derivatives, including defaultable bonds, the fundamental theorem of stochastic finance under complete market no-arbitrage equilibrium states that there exists a unique equivalent martingale measure (EMM) on the transition probabilities.

Let $\tilde{Q}(t, t+1)$ be a per period transition matrix at time t under the EMM, that is generally not time-homogeneous, and equivalent to Q.

$$\tilde{Q}(t,t+1) = \begin{pmatrix} \tilde{q}_{11}(t,t+1) & \tilde{q}_{12}(t,t+1) & \dots & \tilde{q}_{1K}(t,t+1) \\ \tilde{q}_{21}(t,t+1) & \tilde{q}_{22}(t,t+1) & \dots & \tilde{q}_{2K}(t,t+1) \\ \vdots & \vdots & \tilde{q}_{ij}(t,t+1) & \vdots \\ \tilde{q}_{K-1,1}(t,t+1) & \tilde{q}_{K-1,2}(t,t+1) & \dots & \tilde{q}_{K-1,K}(t,t+1) \\ 0 & 0 & \dots & 1 \end{pmatrix}.$$

The element $\tilde{q}_{ij}(t, t+1)$ represents the EMM transition probability of a corporate bond migrating from rating i at time t to rating j in the next period at time $t+1$. The EMM transition probabilities satisfy the following conditions. (i) $0 \leq \tilde{q}_{ij}(t, t+1) \leq 1\,, \forall t,\ \forall i, j \neq K$, and (ii) $\sum_{j=1}^{K} \tilde{q}_{ij}(t, t+1) = 1\,, \forall t, i$.

Let $\tilde{Q}_i(t, t+1)$ be the i^{th} column of $K \times K$ matrix $\tilde{Q}(t, t+1)$, and $\tilde{Q}_K(T-1, T)$ be the K^{th} column of $K \times K$ matrix $\tilde{Q}(T-1, T)$. Then,

$$\tilde{q}_{iK}(t, T) = \tilde{Q}_i(t, t+1)^T \times \tilde{Q}(t+1, t+2) \times \cdots \times \tilde{Q}(T-2, T-1) \times \tilde{Q}_K(T-1, T)$$

denotes the EMM probability of default of a bond by time T, starting from rating i at time t. The superscript T in the expression denotes matrix transpose. Thus $\tilde{q}_{iK}$ is well defined. It is important to note that the default probabilities $\tilde{q}_{iK}(t, T)$ and its empirical equivalent $q_{iK}(t, T)$ are probabilities of default at any time between t and T inclusive. They are not marginal default probabilities.

In the JLT Markov model, suppose $v(t, T)$ and $p(t, T)$ are respectively the credit-risky or defaultable discount bond price and riskfree discount bond price at time t with bond maturity at time T, and suppose $\delta_i \in (0, 1)$ is the recovery rate at maturity T in the event of default for bond with rating i during $[t, T]$, then

$$v_i(t, T) = p(t, T)\left[(1 - \tilde{q}_{iK}(t, T)) + \delta_i \tilde{q}_{iK}(t, T)\right] \tag{1}$$

where the par value of bond is normalized at 1. The stylized feature of the JLT model is the assumption that recovery of fraction δ_i occurs at maturity T when default occurs at any time within $(t, T]$. This may differ from other models as in Elton et.al. (2001) where marginal default probabilities are used instead.

The EMM probability of default by time T, of a bond at rating i starting at time t, is therefore

$$\tilde{q}_{iK}(t, T) = \frac{p(t, T) - v_i(t, T)}{p(t, T)(1 - \delta_i)}. \tag{2}$$

The subscript i to $v(t, T)$ indicates that the defaultable corporate bond is in rating i at time t. Thus, the EMM probability of default $\tilde{q}_{iK}(t, T) \in [0, 1)$ provided $\delta_i < \frac{v_i(t,T)}{p(t,T)}$.

Both Jarrow, Lando, and Turnbull (1997) and Kijima and Komoribayashi (1997) (KK) employ a credit risk premium adjustment to relate the empirical transition probability q_{ij} to the EMM transition probability $\tilde{q}_{ij}$. Based on empirical evidence supplied by the latter, the JLT adjustment has a drawback in that estimated EMM probability may be larger than 1. Thus, we choose instead to use the KK version of JLT risk premium adjustment.

The KK model fixes

$$\tilde{q}_{ij}(t, t+1) = \pi_i(t, t+1) q_{ij} \,, \forall i, j, \,, i, j \neq K. \tag{3}$$

Using the condition that $0 \leq \tilde{q}_{iK}(t, t+1) < 1$, for $i \neq K$, the following boundaries on risk adjustment $\pi_i(t, t+1)$ are easily shown.

$$0 < \pi_i(t, t+1) \leq \frac{1}{1 - q_{iK}}.$$

In the literature on credit spread pricing, most studies have relied on regression analysis without explicitly characterizing the equilibrium probabilities. Others have employed risk-neutral assumptions to circumvent the need to grapple with EMM probabilities that are generally different from the empirical probabilities. JLT model is interesting in that it attempts to link EMM probabilities with empirical probabilities in a stylized manner. The linkage works via the risk adjustments.

Both JLT and KK posit that such risk premium adjustments are deterministic functions of time, and thus introduce non time-homogeneity into the EMM transition matrix. This is, however, a very strong assumption as we reason below.

From (1), with one period to maturity, it can be shown that

$$v_i(T-1, T) = p(T-1, T) \left[\delta_i + \pi_i(T-1, T)(1 - q_{iK})(1 - \delta_i) \right].$$

Since δ_i is a constant, the JLT and KK assumption is tantamount to $\frac{v_i(T-1,T)}{p(T-1,T)}$ being ex-ante a deterministic function of time. The natural logarithm of its inverse is the credit spread which is also a deterministic function of time under this assumption. This imposes an unnecessarily rigid condition on the credit spread dynamics.

In this paper, we relax such a strong assumption, and generalize $\pi_i(t, t+1)$ to be a stochastic process indexed by time t that is adapted to information available at time t. Indeed, we generalize the risk adjustment to cover bonds across the entire term structure. Relaxation of this assumption also allows for meaningful regression analysis to understand the economic dynamics of the adjustments. This is explained as follows.

From (2), for each $i \neq K$, at any time $t < T$,

$$1 - \sum_{j \neq K} \tilde{q}_{ij}(t,T) = \frac{p(t,T) - v_i(t,T)}{p(t,T)(1-\delta_i)}. \tag{4}$$

We introduce a modification of the risk premium adjustment in (3) to be

$$\tilde{q}_{ij}(t,T) = \pi_i(t,T) q_{ij}(t,T)\,, \forall i,j,\ , i,j \neq K, \tag{5}$$

where $q_{ij}(t,T)$ denotes the empirical probability of a bond in rating i at time t ending up in rating j by time T. Then, substituting for $\tilde{q}_{ij}(t,T)$ from (5), (4) becomes

$$\pi_i(t,T) = \frac{v_i(t,T) - \delta_i p(t,T)}{(1 - q_{iK}(t,T))p(t,T)(1-\delta_i)}. \tag{6}$$

This can be viewed as a useful generalization of the JLT or KK risk adjustments because the factors here, $\pi_i(t,T)$ are similar in form across term, i.e. T can take any value greater than, and is thus not restricted to $t+1$. It also avoids the complicated computations of adjustments in JLT or KK that together made up an adjustment of a term greater than $t+1$. A salient and useful feature of this generalization is that we can interpret a transform of the adjustments as a component of a linear partition of the credit spread.

From (1) and (5), we obtain

$$v_i(t,T) = p(t,T)\left[\delta_i + \pi_i(t,T)(1 - q_{iK}(t,T))(1-\delta_i)\right].$$

Thus the credit spread for bond with rating i and time to maturity $T-t$ is

$$(T-t)^{-1} \ln\left[\frac{p(t,T)}{v_i(t,T)}\right] = -(T-t)^{-1} \ln\left[\delta_i + \pi_i(t,T)(1 - q_{iK}(t,T))(1-\delta_i)\right]. \tag{7}$$

From (7), it is straightforward to show that the partial derivatives of the credit spread with respect to δ_i, q_{iK}, and $\pi_i(t,T)$ are respectively < 0, > 0, and < 0. Thus, reduced recovery, increased unconditional probability of default, and reduced value of $\pi_i(t,T)$ all lead to increase in credit spread, and vice-versa.

It is instructive to consider the special case when $\delta_i = 0$. Then

$$(T-t)^{-1} \ln \left[\frac{p(t,T)}{v_i(t,T)} \right] = -(T-t)^{-1} \ln \left[\pi_i(t,T)(1 - q_{iK}(t,T)) \right].$$

Thus

$$Y_i(t,T) = (T-t)^{-1} \ln \frac{1}{\pi_i(t,T)} + (T-t)^{-1} \ln \frac{1}{1 - q_{iK}(t,T)}, \tag{8}$$

where $Y_i(t,T)$ is the credit spread $(T-t)^{-1} \ln \left[\frac{p(t,T)}{v_i(t,T)} \right]$. The positivity of the risk adjustment ensures that $(T-t)^{-1} \ln \frac{1}{\pi_i(t,T)}$ exists. Thus, in the absence of a recovery factor, the credit spread is a linear combination of unconditional default component, $(T-t)^{-1} \ln \frac{1}{1-q_{iK}(t,T)}$, and a log-transform of the risk adjustments.

The interpretation of the risk adjustment or its transform $\ln \frac{1}{\pi_i(t,T)}$ is an important aspect of the contribution of this paper. From (8), it is obvious that given many existing credit spread studies where spreads are regression relations on macroeconomic and other explanatory variables, these adjustments are not necessarily deterministic functions, but are indeed similar regression relations on macroeconomic and other explanatory variables.

Suppose the risk adjustment at time t is a function of random variables $X(t)$. Then via (3), the EMM transition probabilities are functions of $X(t)$, or are conditional probabilities. (3) is explicit in terms of a relation between the EMM transition probabilities and the unconditional empirical transition probabilities. The conditional empirical transition probabilities are not modelled here or in JLT[1], but is certainly related to the conditional EMM transition probabilities. Thus the risk adjustments are related to the conditional empirical transition probabilities. The dynamics of the risk adjustments would also reflect conditional default risk, illiquidity risk, and other conditional risks not related to recovery effects.

From (6), we obtain

$$\frac{1}{\pi_i(t,T)} = (1 - q_{iK}(t,T))(1-\delta_i)\frac{p(t,T)}{v_i(t,T)}(1 - \delta_i \frac{p(t,T)}{v_i(t,T)})^{-1}.$$

[1]If the empirical transition probabilities are constant over time, i.e. not conditionally a function of $X(t)$, then strong implications such as $\pi_i(t,T) < 1$ for aggregate market risk aversion, and the interpretation of the risk adjustments being non-default related factors would be necessary.

Thus,

$$\frac{p(t,T)}{v_i(t,T)} = \frac{1}{\pi_i(t,T)} \frac{1}{(1-q_{iK}(t,T))} \frac{(1-\delta_i \frac{p(t,T)}{v_i(t,T)})}{(1-\delta_i)}.$$

Taking natural logarithms, and dividing by the term-to-maturity,

$$\begin{aligned}(T-t)^{-1}\ln\left[\frac{p(t,T)}{v_i(t,T)}\right] &= (T-t)^{-1}\ln\frac{1}{\pi_i(t,T)} + (T-t)^{-1}\ln\frac{1}{1-q_{iK}(t,T)} \\ &+ (T-t)^{-1}\ln\theta_i(t,T), \qquad (9)\end{aligned}$$

where $\theta_i(t,T) = \frac{(1-\delta_i \frac{p(t,T)}{v_i(t,T)})}{(1-\delta_i)}$ is a recovery factor that increases credit spread when recovery is reduced and vice-versa. Credit spread increases with the unconditional default component $\frac{1}{1-q_{iK}(t,T)}$ when the empirical unconditional default probability $q_{iK}(t,T)$ is increased. Credit spread also increases with the negative logarithms of the risk adjustment, $\ln\frac{1}{\pi_i(t,T)}$. We shall term this a conditional risk component. Thus, our modified version of the risk adjustments in (5) allows a linear partition of the credit spread into three components: an unconditional default component, a recovery component, and a conditional risk component. This result in (9) applies across different terms T.

For a particular bond rating i, by comparing across the maturity term, credit spread slopes have implications on the term structure of the risk adjustments. This is shown in Proposition 1.

Proposition 1: Given a downward sloping or flat credit spread term structure, for time $T' > T$,

$$\pi_i(t,T') > \pi_i(t,T).$$

Proof: From (7), for time $T' > T$, a downward sloping or flat credit spread term structure is represented by

$$\begin{aligned}&\ln\left[\delta_i + \pi_i(t,T')(1-q_{iK}(t,T'))(1-\delta_i)\right] \\ \geq\ &\ln\left[\delta_i + \pi_i(t,T)(1-q_{iK}(t,T))(1-\delta_i)\right]^{\frac{T'-t}{T-t}} \\ >\ &\ln\left[\delta_i + \pi_i(t,T)(1-q_{iK}(t,T))(1-\delta_i)\right].\end{aligned} \qquad (10)$$

Thus, $\pi_i(t,T')(1-q_{iK}(t,T')) > \pi_i(t,T)(1-q_{iK}(t,T))$. As $q_{iK}(t,T') > q_{iK}(t,T)$, we have $\pi_i(t,T') > \pi_i(t,T)$.

Q.E.D.

The converse situation of an upward sloping credit spread slope, however, can lead to the possibility of either cases of upward or downward sloping term structures in the risk adjustments. However, if the credit spread rises at a sufficiently fast rate, then there will be a downward sloping term structure in the risk adjustments. This is shown in Proposition 2 below.

Proposition 2: If $Y_i(t,T') > Y_i(t,T) + (T'-t)^{-1} \ln \frac{1-q_{iK}(t,T)}{1-q_{iK}(t,T')}$, for $T' > T$, then

$$\pi_i(t,T') < \pi_i(t,T).$$

A necessary condition for the latter is $Y_i(t,T') > \frac{T-t}{T'-t} Y_i(t,T)$.
Proof: Suppose

$$(T'-t)^{-1} \ln \frac{p(t,T')}{v_i(t,T')} > (T-t)^{-1} \ln \frac{p(t,T)}{v_i(t,T)} + (T'-t)^{-1} \ln \frac{1-q_{iK}(t,T)}{1-q_{iK}(t,T')}.$$

This implies

$$\ln \frac{p(t,T')}{v_i(t,T')} > \ln \frac{p(t,T)}{v_i(t,T)} + \ln \frac{1-q_{iK}(t,T)}{1-q_{iK}(t,T')}.$$

Therefore,

$$\frac{v_i(t,T')}{p(t,T')} < \frac{v_i(t,T)}{p(t,T)} \left(\frac{1-q_{iK}(t,T')}{1-q_{iK}(t,T)} \right).$$

Let $\beta = \frac{1-q_{iK}(t,T')}{1-q_{iK}(t,T)}$. Hence $0 < \beta < 1$. This implies

$$\frac{v_i(t,T')}{p(t,T')} < \frac{v_i(t,T)}{p(t,T)} \beta + \delta_i(1-\beta).$$

Hence,

$$\frac{v_i(t,T') - \delta_i p(t,T')}{(1-q_{iK}(t,T'))p(t,T')(1-\delta_i)} < \frac{v_i(t,T) - \delta_i p(t,T)}{(1-q_{iK}(t,T))p(t,T)(1-\delta_i)}.$$

From (6), it is readily seen that $\pi_i(t,T') < \pi_i(t,T)$.

The necessary condition is obtained from eliminating constants $(1-\delta_i)$ in the following:

$$\frac{v_i(t,T') - \delta_i p(t,T')}{(1-q_{iK}(t,T'))p(t,T')(1-\delta_i)} < \frac{v_i(t,T) - \delta_i p(t,T)}{(1-q_{iK}(t,T))p(t,T)(1-\delta_i)}.$$

Then,

$$\frac{v_i(t,T')}{p(t,T')} - \delta_i < \left[\frac{v_i(t,T)}{p(t,T)} - \delta_i\right]\beta < \frac{v_i(t,T)}{p(t,T)} - \delta_i,$$

from which the result is readily obtained.

Q.E.D.

In the next section, the data and empirical methods are described, and then the empirical results are reported. The time series properties of the risk adjustments $\pi_i(t,T)$ and the conditional risks will be examined.

3 Empirical Analysis

3.1 Data

The Lehman Brothers Fixed Income Database (LB)[2] is employed in this study. The LB database of US corporate bonds contains monthly information on their CUSIP, company name, issue date, maturity date, trader-quoted prices, accrued interest, annual coupon, amount outstanding, credit rating, callability, putability, and industry code, amongst others. However, for the purpose of computing spot rate curve, certain critical information are lacking in this database. It does not appear to contain details such as coupon frequency, coupon dates, and coupon change schedule. Although it indicates if a bond is callable, details of call dates and call price schedule are not provided. There is also no information about convertibility, which is a common bond feature. A sizeable number of floating bonds also appear in this database.

Although the database contains both trader-quoted prices and matrix-computed (model) prices, only trader-quoted prices are used in this study. Convertible, putable, and floating-rate bonds are eliminated. Many of the remaining bonds contain the call feature. Not all such bonds can be eliminated in order to keep a

[2]The LB database was available from the University of Houston prior to 1999, and contains corporate bond data in the period 1973 till March 1998.

sizeable sample. Bonds that are callable within one year from the price date are eliminated. This is to reduce the complicity of the call premium.[3]

In order to perform bootstrapping to find the spot rate curve, the bonds are tracked via their CUSIP in a separate Fixed Income Securities Database (FISD)[4], and their coupon frequency, coupon dates, and coupon change schedules are recorded. LB bonds that do not match those in FISD or whose detailed coupon information were not found in FISD were removed. These coupon details provide for a far more accurate bootstrapping than usual assumptions of semi-annual coupons, mid- and end-of tax year coupon dates, and zero coupon change, that were used in some earlier studies. For any remaining callable bonds, their call dates and call price schedules are recorded. Only bond data that contain complete records of such coupon and call information are included. Bonds in the LB database that have characteristics, e.g. ratings or maturity dates, that are not consistent with the records in FISD for the same CUSIP are removed to avoid data entry errors in the database.

The dataset is further restricted to bonds that are not obviously illiquid. Those with amount outstanding of less than $10,000 are eliminated. Corporate bonds with maturity longer than 10 years are not as regularly quoted, so the sample contains bonds with maturities of 10 years or less. This study also performs a pricing check such as in Elton, Gruber, Agrawal, and Mann (2001), and a small number of bonds or less than 1% of the bonds, whose quoted prices fall outside a 5% boundary of the theoretical price according to the spot rate curve, are treated as data entry errors and are eliminated. Since rating change records for speculative bonds are available in the database only after 1991, in order to perform the empirical study of investment as well as speculative grade bonds, the data sampling period from January 1992 to March 1998 is used.

The selected set of bond data is partitioned by months and for each month, it

[3]Crabble and Helwege (1994) show that such call options increased spreads by less than 10 basis points. Sarig and Warga (1989), Fons (1994), and Helwege and Turner (1999) all apply some arbitrary criteria to reduce, but not eliminate, the number of callable bonds since they form a substantial part of all corporate tradeable bonds.

[4]The FISD is provided by the LJS Global Information Services Inc. and covers the period 1987 till February 2002. Unlike the LB database, the data here are not arranged in monthly fixed intervals.

is sub-partitioned by the different corporate bond ratings according to Moody's ratings: Aaa, Aa, A, Baa, Ba, B, and Caa. The choice of Moody's follows those of the many earlier studies including JLT (1997) and KK (1998), and would enable some degree of comparison. Moody's also provides default probability data, rating transition information and recovery rate that we can use for comparison purposes. A priori there is no other reason to believe one particular agency's rating system is more appropriate than the other.

After all the various detailed checks and eliminations discussed above, of which a sizeable number of bonds that were recorded as floating and putable were discarded, we have a final sample size of 8804 bond prices, including Treasury bonds, for the bootstrapping in the sample period we are studying. Of these, the numbers for the Aaa-rating, Aa-rating, A-rating, Baa-rating, Ba-rating, B-rating, and Caa-rating are respectively 727, 1995, 2055, 1356, 890, 593, and 147. Since the numbers of B-rating and Caa-rating speculative bonds are relatively scarce, bootstrapping will not be performed for these two lower ratings.

The bootstrap method in Jarrow, Lando, and Turnbull (1997) is applied to the bond data for each month and for each rating within the month. The algorithm computes the yield-to-worst if a bond is callable. Then the bond price is adjusted before the incremental spot rate is determined. The method assumes linear interpolation in the discount function. The no-arbitrage condition of decreasing discount function is also imposed in the algorithm. By utilizing the yield-to-worst adjustment as in JLT, the pricing bias introduced by residual call premia is minimized, if not removed. The bootstrapped spot rate curve for each month and each rating is also compared with spot rate curve derived using the Litzenberger and Rolfo (1984) method. The differences are negligible.

Previous research such as the JLT and KK models, and also Moody's yearly corporate bond default reports apply the same recovery rate to corporate bonds of the same seniority but of different ratings. Their reason is that a bond issue is almost always below investment grade just prior to default, so its original rating should play no role in determining its recovery rate. However, a company with high rating could also be forced into bankruptcy if its capital structure is not healthy and is thus short of cash to pay interest at a certain time. Altman and Kishore (1998) show that the recovery rates are different for the different original ratings. This finding is in agreement with extant financial theory. Elton et al.(2001) use the same set of recovery rates as in Altman and Kishore's study. In this paper, we utilize the same set of rates as in the above two studies. These recovery rates are different for different ratings. They are as follows.

Recovery Rates

Original Rating	Aaa	Aa	A	Baa	Ba	B	Caa
Recovery Rate%	68.34	59.59	60.63	49.42	39.05	37.54	38.02

Actual empirical default occurrences are noted based on records obtained from the LB and the FISD databases. Credit migration occurrences are also noted. For each month, the databases provide the numbers of occurrences of defaults and of credit rating status of bonds originating with rating i, for each i, at the beginning of the month.

Several macroeconomic variables are also employed in our analysis. Monthly data of S&P500 stock index, U.S. 3-month Treasury yields, and U.S. 30-year Treasury bond yields in the sample period January 1992 to March 1998 are collected from Datastream. The Treasury slope or the term structure of riskfree rate is calculated as the difference between the 30-year yield and the 3-month yield.

3.2 Method

For each month from January 1992 to March 1998 in the sampling period, and for each rating $i \in \{Aaa, Aa, A, Baa, Ba\}$, the zero coupon discount bond prices of various terms of 1-year, 2-year, 3-year, and so on till 10-year are computed based on the spot rates $r_i(t,T)$. The corresponding term riskfree spot rate, or spot rate on the zero coupon Treasury, is $r(t,T)$. These spot rates are obtained from the bootstrapped spot rate curves employing the JLT methodology mentioned in the last sub-section. Thus the credit risky discount bond prices and the riskfree discount bond prices are

$$v_i(t,T) = \exp(-r_i(t,T)(T-t))$$

and

$$p(t,T) = \exp(-r(t,T)(T-t)).$$

These credit risky bond prices $v_i(t,T)$ as well as riskfree bond prices $p(t,T)$ are generic to the rating class in JLT model. Some explanations are called for. The JLT model contains no bond-specific or issuer related variables. Any bond with the same rating, coupon structure, and maturity must follow the same spot rate curve, regardless of the issuing firm and other specifics of the firm. Obviously there will be some variations in spot rates even amongst bonds of the same rating and maturity if they belong to different firms with different specifics. One treatment is to price bonds from each rating class by adding an error term to the spot yield for the rating class.[5] The JLT approach, however, is to assume such within-rating-class variations are negligible, and to perform the bootstrapping algorithm (which includes smoothing) that converts all such yields within the class into a common spot rate curve that applies to all bonds in that rating class for that point in time. We adopt the JLT approach in this paper.

The focus of this paper is on the risk adjustments expressed in (6). From (6), for each month t and each credit rating $i \in \{Aaa, Aa, A, Baa, Ba\}$, the risk adjustments $\pi_i(t,T)$ of various terms of 1-year, 2-year, 3-year, and so on till 10-year are computed. Apart from the inputs of prices $v_i(t,T)$ and $p(t,T)$, and of recovery rate δ_i, the input of default probability $q_{iK}(t,T)$ is also required.

Actual numbers of defaults and of rating changes from the sample over the months from January 1992 to March 1998 are used to compute estimates of the empirical unconditional transition probabilities.[6] The JLT model assumes a constant per period empirical transition matrix or Markov Chain.

$$Q(0,1) = \begin{pmatrix} q_{11}(0,1) & q_{12}(0,1) & \dots & q_{1K}(0,1) \\ q_{21}(0,1) & q_{22}(0,1) & \dots & q_{2K}(0,1) \\ \vdots & \vdots & q_{ij}(0,1) & \vdots \\ q_{K-1,1}(0,1) & q_{K-1,2}(0,1) & \dots & q_{K-1,K}(0,1) \\ 0 & 0 & \dots & 1 \end{pmatrix}.$$

[5]This distinction is pointed out by a referee.

[6]We wish to thank the referee for pointing out the desirability of estimating the transition probabilities instead of relying on published data by agencies such as RiskMetrics.

where $q_{ij}(0,1)$ represents the unconditional transition probability of a corporate bond migrating from rating i to j over a month. Therefore, we can estimate the empirical transition probabilities using Anderson and Goodman (1957)'s maximum likelihood estimator

$$\hat{q}_{ij} = \frac{n_{ij}}{n_i},$$

where n_{ij} is the total number of i to j monthly transitions in the sample (including remaining in state i), and n_i is the total number of visits to state i at the beginning of each month.

Application of the Chapman-Kolmogorov equation allows for estimation of the 1-year unconditional transition probability matrix $\hat{Q}(1) = \hat{Q}(0,1)^{12}$. The n-year unconditional transition probability matrix is estimated as $\hat{Q}(n) = \hat{Q}(1)^n$. The estimates are consistent and asymptotically efficient. Thus the empirical probability of default $q_{iK}(t,T)$ for a bond with rating i starting at month t, and entering default by time $T = t + 12n$ or within n years, is estimated by the iK-element of $\hat{Q}(n)$.

The positivity of $\pi_i(t,T)$ allows for a suitable specification of the risk adjustment as

$$\pi_i(t,T) = \exp\left(-f(X(t))\right) , \tag{11}$$

where $f : X(t) \longrightarrow \Re$ is a regression function. Specifically, we specify a linear regression for the log-transform of the risk adjustment or the condition risk as

$$(T-t)^{-1} \ln \frac{1}{\pi_i(t,T)} = C_0 + C_1 X_1(t) + C_2 X_2(t) + C_3 X_3(t) + \tilde{\epsilon}(t) \tag{12}$$

where C_p's are regression constants, $X_p(t)$'s are exogenous explanatory variables at time t, and $\tilde{\epsilon}(t)$ is a disturbance term that has zero mean, is normally distributed, and is uncorrelated with the explanatory variables. The distributional assumption is not critical when asymptotics are considered. We allow for unknown forms of heteroskedasticity in the disturbances.

From the decomposition of the credit spread given by (9) and the discussion in section 2, the conditional risk component above reflects conditional default risk, illiquidity risk, and other risks not related to recovery effects. These conditional risks may depend on the macroeconomic variables employed as exogenous explanatory variables in the regression in (12).

The regression in (12) allows for empirical estimation and forecasting of conditional risks that are related to standard macroeconomic variables. Since the conditional risks form a linear component of the credit spread, its correlation with macroeconomic variables is consistent with existing results of correlations of such variables with credit spread itself. Regression specification (12) also allows a parsimonious term structure of conditional risks to be estimated and tested.

A note about the appropriateness of conditioning the risk adjustment, or its log-transform, on exogenous variables is in order. From (6), the dependence of $\pi_i(t,T)$ on exogenous variables is equivalent to the dependence of $\pi_i(t,T)(1 - q_{iK}(t,T))$ or $\sum_{j \neq K} \tilde{q}_{ij}$ on the exogenous variables. The latter is essentially a situation where transition matrices (EMM in this case) can be dependent on exogenous variables. Concerning this, many examples are found in the economics literature. See, for examples, some of the earlier works, Boskin and Nold (1975) and Toikka (1976), as discussed in Amemiya (1985, Chapter 11).

3.3 Results

In this section, empirical results obtained by employing the methods described in the previous section are reported and discussed. The output tables and figures are shown at the end of the paper.

Table 1 about here

In Table 1, we show the estimated empirical unconditional 1-year, 5-year, and 10-year transition probability matrices. The estimates are based on the maximum likelihood method, and are thus consistent and asymptotically efficient given that the standing assumption in JLT model is a constant per period transition probability matrix. The probability of default at any time within the interval increases as interval time lengthens, and as rating becomes lower. This feature is important for the proofs in the earlier section. Over a longer horizon or time interval, the downgrade migration probabilities increase. However for lower grade speculative bonds, there is actually increase in upgrade migration probabilities as well.

These increases in downward and upward migrations are at the expense of the probability of no change in rating transition.

Our transition probability estimates obtained in Table 1 may be compared with some estimates in the literature. For example, the default rate for Caa bonds over 1 year is 11.75% in our sample, but larger, 24.06% in Kijima and Komoribayashi (1997, Exhibit 1). The default rate for Caa bonds over 5 year is 43.29% in our sample, but slightly smaller, 39.5% in Barnhill, Maxwell, and Shenkman (1999, p.149, Exhibit 7-18). The latter also sorted ratings by Moody's classification. Some differences will be unavoidable as different estimates are based on different sample periods and different ratings classifications. However, what is important in our case is that the empirical estimates of the transition probabilities, including the default probabilities, are based on the same set of bonds from which the prices and rates are derived for the empirical study. This will reduce bias resulting from the use of a different sample of bonds in a different time period whose migration characteristics may not fit the ones that give rise to the pricing of the bonds.

In our sample during January 1992 till March 1998, there was not a single case of default or credit migration while a bond is under Aaa rating. Therefore the estimates of empirical unconditional transition probabilities for Aaa-rated bonds in our sample are $(1, 0, 0, 0, 0, 0, 0, 0)$. This implies, via (1), that the risk adjustment is less than 1 if $v_i(t, T) < p(t, T)$. This implication can be seen in Table 2.

In Table 2, we report summary statistics of the estimates of the risk adjustments based on (6) and of its negative log-transform or the conditional risks. They were computed using the consistent estimates of empirical default rates that were computed earlier. Based on the monthly time series of the computed risk adjustments and the conditional risks from January 1992 till March 1998, their means, standard deviations, minimums, and maximums are shown.

Table 2 about here

According to the risk adjustment constraint in (3), $\pi_i(t,T) > 0$. This is evident in Table 2. For the Aaa-rated bonds, we observe that all estimated risk adjustments are less than 1, which is a rationality implication discussed earlier for bonds in this rating. For the high grade investment bonds, for some months, the estimated risk adjustments are close to 1.

The means of the risk adjustments summarize the overall tendencies for the risk adjustments to decrease as the term increases. This result is observed across all credit ratings. There is also a tendency for the risk adjustments to decrease when the rating is lower given a fixed term. The difference in the means of the risk adjustments for speculative grade bond, such as the Ba-rated bonds, and the investment grade bonds is distinctive. For example, the Ba-rated bonds carry a range of adjustments from 0.9714 down to 0.4384, while the A-rated bonds have a range of 0.9838 down to 0.6687. Obviously, the conditional risks of the speculative bonds are distinctively higher than those of the investment grade bonds. The term structure of adjustment curves of the investment grade bonds exceeds that of the adjustment curve of the speculative bond.

The standard deviation of the risk adjustments also increases with term. However, the volatility for the conditional risk is more stable given that the measure is normalized by the term-to-maturity.

According to Proposition 1, the results do not reflect any downward sloping credit spread term structures, and instead provide evidence of increasing credit spreads at a fast enough rate according to Proposition 2. This will be confirmed with Figure 2 later.

The conditional risks, on the other hand, rise across the term structure. They are fairly stable across ratings for a given term. Supposing illiquidity risk premium is a considerable part of this conditional risk, this increasing liquidity term premium is consistent with Koziol and Sauerbier (2002) who employ an option-theoretic approach in defining liquidity.

Figures 1 & 2 about here

Figure 1 shows that the term structure of average spot rate curve is upward rising for Treasury, investment grade (Aaa, Aa, A, Baa) as well as speculative grade bonds (Ba) during this sample period. There appears to be a small hump coming at about the 9- or 10-year term for bonds with ratings Aa or lower.

Figure 2 shows the term structure of average credit spreads during the sample period. The spreads were increasing in term for all the grades down to Ba-bonds. There is a small hump at about the 9-year term for A-rated, Baa-rated, and Ba-rated bonds. The results are consistent with recent empirical research such as Helwege and Turner (1999) and theoretical research such as Collin-Dufresne and Goldstein (2001) and Duffee (1999, figure 1) that indicated upward sloping credit spreads even for speculative grade bonds. The characteristic of upward or downward sloping credit term structure has been an interesting issue, and earlier studies by Sarig and Warga (1989) and Fons (1994) indicated upward sloping credit term structures for investment grade bonds and downward sloping structures for speculative grade bonds.

We now turn to the regression results involving (12). The exogenous variables used are the S&P500 market index returns, the U.S. 3-month Treasury rate, and the Treasury slope. The computed conditional risk variable, $(T-t)^{-1}\ln\frac{1}{\pi_i(t,T)}$, are the dependent variables in the linear regression on the exogenous variables. Since we allow for unspecified heteroskedasticity in the disturbances, we apply the White's heteroskedasticity-consistent covariance matrix estimator for the inference. The method employed is least squares regression. Augmented Dickey-Fuller statistics are also reported for the dependent variable as well as for the residuals to check for stationarity or to reject the null of unit roots. The regression results are reported in Table 3.

Table 3 about here

Table 3 shows that the short-term Treasury rate (T-Rate) is significantly negatively correlated with the conditional risk variable, except for the very high grade Aaa-rated bonds. For example, in the case of the A-rated bonds, the T-rate coefficient is significantly negative (taken to be at least at the 10% level) 6 times out of 10 across the different terms, and negative 3 other times. Of the times when the coefficient is significantly negative, 5 of them occurs at 1% significance level. Very similar results occur for the cases of Aa-rated, Baa-rated, and Ba-rated bonds. Our results are consistent with past works such as Duffee (1998) and Longstaff and Schwartz (1995), and with current research such as Yamauchi (2003). They find that an increase in short-term Treasury rate reduces credit spreads.

However, since we have decomposed the credit spread into the conditional risk component for our dependent variable, the interpretation here is more interesting.

The conditional risks include conditional default risk and illiquidity risk. The results indicate that higher short-term Treasury rates signal improved economic situations, lower default probabilities, more optimistic markets, and generally reduced illiquidity risks.

Some of the T-rate coefficients on the Aaa-rated bonds are positive and significant. In this case, higher short-term Treasury rates spell higher conditional risks. We suggest that the conditional risks are also related to diversifcation aspects of the bond in a portfolio context. When the market is optimistic, the diversification benefits of very high grade bonds in stock market portfolio diminishes. The reduced diversification benefit leads to lower demand. This translates into higher conditional risk premium. It should be noted from (9) that while the conditional risk component increases with T-rate for Aaa-rated bonds, the other components could decrease with T-rate. Thus the overall credit spread could still decrease with an increase in T-rate. The latter would be consistent with existing results in the literature concerning Aaa-rated bonds.

The importance of short-term Treasury rate in explaining the credit spread could be seen as a potential contradiction to the JLT model. In the JLT model, it is assumed that the riskfree rate process is statistically independent of the default process. Nevertheless, this may not be a serious problem with the JLT model for the purpose of pricing.

The Treasury slope or equivalently forward rate shows similar impact as the short rate on the conditional risks except for some speculative Ba-rated bonds with short maturities of less or equal to 3 years. For the A-rated, and Baa-rated bonds, all the estimated Slope coefficients are negative. More than half the number of coefficient estimates in the A-rated bonds are significantly negative. The results indicate that higher Treasury slopes or increasing riskfree forward rates signal improved economic situations, lower default probabilities, more optimistic markets, and generally reduced illiquidity risks. Thus this component of credit spread is reduced.

For the short term Ba-rated speculative bonds, higher Treasury slopes signifying a more buoyant stock market could possibly introduce higher illiquidity in the speculative bonds market as funds are transferred out to the stock market. Longer term investments in these bonds may be less affected since they offer greater diversification benefits relative to shorter term bonds and other investment grade bonds.

It is interesting to note that the coefficients for market returns are generally mixed and not significant. In view of existing empirical results by Duffee (1998), Collin-Dufresne, Goldstein and Martin (2001), and others, who find market returns to be negatively correlated with credit spreads, our results could indicate that the market return impact lands itself on the other components of credit spread premium in (9), and not on conditional risks. Thus, market return could possibly impact on recovery related factors.

The constant coefficients reflect increased conditional risk premium as term increases, for all ratings, and independent of the explanatory variables. Interpreting conditional risk premium as made up of conditional default risk and illiquidity and other risks, this increasing liquidity term premium is consistent with increasing conditional default probabilities over time, and with an increasing term structure of liquidity premium. They are consistent with results in Koziol and Sauerbier (2002).

The ADF statistics show that the dependent risk variables are stationary. In the ADF test, we employ zero deterministic time trend and two lagged changes of the dependent variable. The unit root hypothesis is significantly rejected in all the cases. In 80% of these cases, rejection of unit root for a stationary alternative occurs at the 5% or 1% significance level. Similar results are obtained for the residuals in the linear regression.

The regression adjusted R^2's are low in some cases. However, this is consistent with past work on credit spreads such as Collin-Dufresne, Goldstein, and Martin (2001) who found that traditional model factors could explain only a small fraction of the variation in credit spreads. The order of magnitudes of the adjusted R^2's is also compatible with regression results in Elton, Gruber, Agrawal and Mann (2001). The adjust R^2's are higher for many of the cases in the speculative grade bonds relative to the investment grade bonds. This could be explained by the higher proportion of the conditional risk dynamics being captured by the macroeconomic variables.

4 Conclusions

In this paper we investigate the nature of the credit risk premium adjustments in the Jarrow-Lando-Turnbull model of credit risk spreads. The adjustments relate the equivalent martingale measures to the empirical measures of unconditional transition probabilities. We provide a modified version of the risk adjustment that allows a linear partition of the credit spread into an unconditional default component, a recovery component, and the risk premium adjustment. The risk adjustments are decreasing in both term structure as well as lower credit ratings.

The negative log-transform of the risk adjustment is the conditional risk component of the credit spread. Increases in this conditional risk leads to increases in the credit spread. The conditional risks are stationary and can be specified as linear regressions on a set of macroeconomic variables.

The decomposition facilitates the interpretation of the conditional risks to be unrelated to unconditional default probabilities and recovery factors. Instead they are likely to be related to conditional default risk, illiquidity risk, and other factors including diversification aspects of the bond in a portfolio context. The constants of the regressions show an increasing term structure of conditional risk. The conditional risks are highly sensitive to short-term Treasury rate movements as well as Treasury slope changes. However, the conditional risks appear to be insensitive to market returns. The finding could suggest that the market return appears to impact on recovery related factors and not on conditional default risk or illiquidity risk premium during our sample period. Potentially interesting questions can be explored in this framework.

The sensitivity of conditional risk to short-term Treasury rate movements may suggest a contradiction to the JLT model that assumes the statistical independence of the riskfree rate and the default processes. Empirical work on the JLT model to account for possible correlation between default and short-rate processes may be a worthwhile extension of research in this area.

Analyzing components of the credit spread is an interesting advancement of research in this field. The methodology to condition transition probabilities on exogenous variables can be extended to include conditional empirical default probabilities as well, and will provide some avenues for future methodological research.

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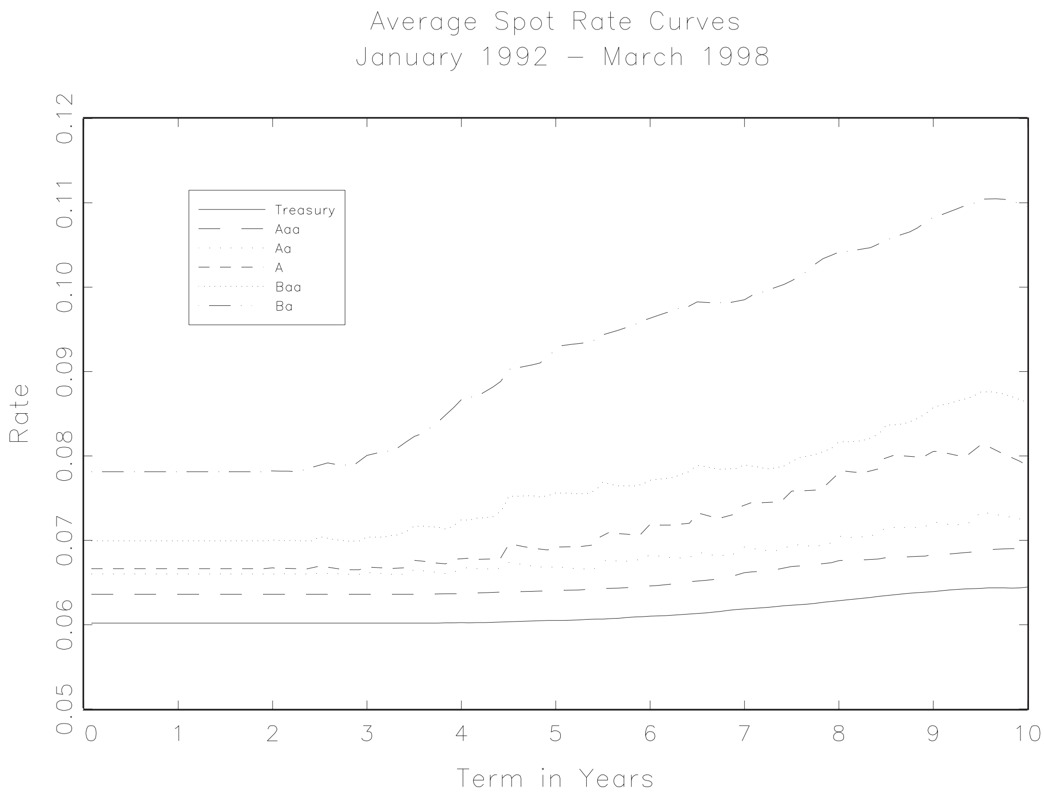
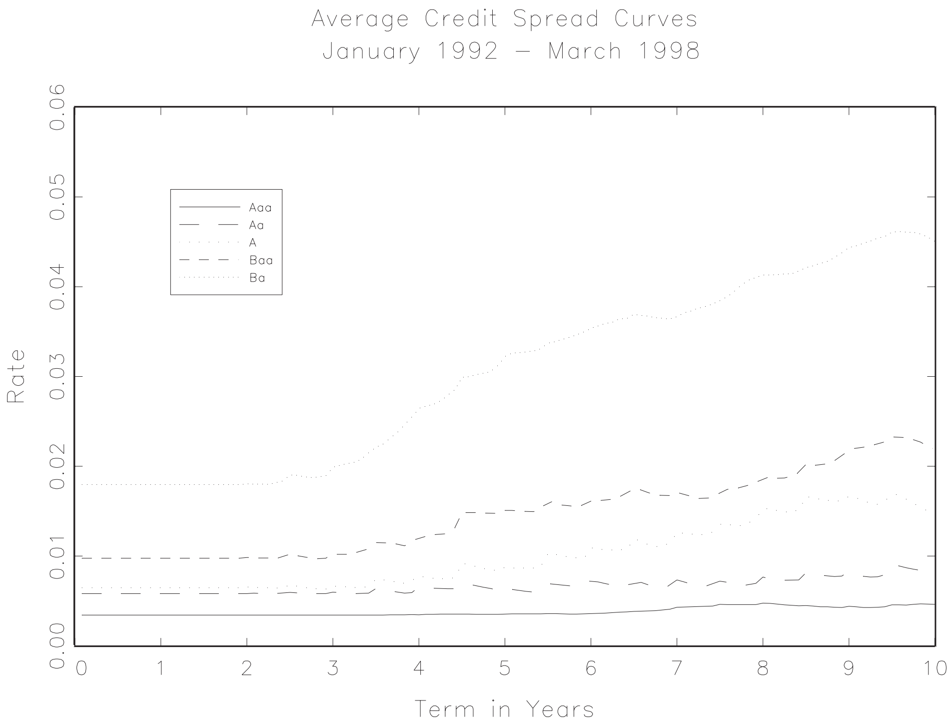
Figure 1:



Average Spot Rate Curves
January 1992 – March 1998
Treasury
Aaa
Aa
A
Baa
Ba
Rate
0.05
0.06
0.07
0.08
0.09
0.10
0.11
0.12
0
1
2
3
4
5
6
7
8
9
10
Term in Years

Figure 2:



Average Credit Spread Curves
January 1992 - March 1998
Aaa
Aa
A
Baa
Ba
Rate
0.00
0.01
0.02
0.03
0.04
0.05
0.06
0
1
2
3
4
5
6
7
8
9
10
Term in Years

Table 1: **Transition Probability Matrix.** The table shows the 1-year, 5-year, and 10-year empirical transition probability matrices of credit ratings of corporate bonds. Each entry in a row shows the probability of a bond starting with the rating in the row and ending up with the rating in the column after 1 year, 5 years, or 10 years respectively. The probabilities in each row sum to one. The numbers are maximum likelihood estimators based on actual credit migration and default data in the period in US from January 1992 to March 1998.

One-Year								
	Aaa	Aa	A	Baa	Ba	B	Caa	Default
Aaa	1.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
Aa	0.0000	0.9663	0.0334	0.0002	0.0001	0.0000	0.0000	0.0000
A	0.0000	0.0161	0.9640	0.0149	0.0049	0.0000	0.0000	0.0001
Baa	0.0000	0.0003	0.0343	0.9516	0.0136	0.0001	0.0000	0.0001
Ba	0.0000	0.0000	0.0039	0.0725	0.8991	0.0162	0.0078	0.0005
B	0.0000	0.0000	0.0001	0.0056	0.0191	0.9433	0.0278	0.0041
Caa	0.0000	0.0000	0.0000	0.0000	0.0004	0.0381	0.8440	0.1175
Default	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	1.0000
Five-Year								
	Aaa	Aa	A	Baa	Ba	B	Caa	Default
Aaa	1.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
Aa	0.0000	0.8473	0.1453	0.0054	0.0018	0.0001	0.0000	0.0000
A	0.0000	0.0700	0.8421	0.0659	0.0202	0.0007	0.0003	0.0005
Baa	0.0000	0.0061	0.1455	0.7929	0.0515	0.0022	0.0008	0.0006
Ba	0.0000	0.0010	0.0354	0.2685	0.5978	0.0609	0.0259	0.0101
B	0.0000	0.0001	0.0031	0.0339	0.0699	0.7574	0.0905	0.0449
Caa	0.0000	0.0000	0.0001	0.0022	0.0064	0.1227	0.4354	0.4329
Default	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	1.0000
Ten-Year								
	Aaa	Aa	A	Baa	Ba	B	Caa	Default
Aaa	1.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
Aa	0.0000	0.7282	0.2464	0.0190	0.0059	0.0003	0.0001	0.0002
A	0.0000	0.1188	0.7297	0.1137	0.0328	0.0026	0.0010	0.0014
Baa	0.0000	0.0204	0.2407	0.6524	0.0748	0.0068	0.0026	0.0023
Ba	0.0000	0.0057	0.0905	0.3780	0.3764	0.0864	0.0326	0.0304
B	0.0000	0.0006	0.0124	0.0718	0.0972	0.5892	0.1099	0.1189
Caa	0.0000	0.0000	0.0011	0.0087	0.0153	0.1468	0.2009	0.6270
Default	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	1.0000

Table 2: **Descriptive Statistics of the Credit Risk Premium Adjustments.** This table shows the mean, standard deviation, minimum, and maximum of the credit risk premium adjustments $\pi_i(t,T)$ and its transform $(T-t)^{-1}\ln\frac{1}{\pi_i(t,T)}$ for the different Aaa-rated, Aa-rated, A-rated, Baa-rated, and Ba-rated bonds. For each bond, the statistics for the factors across different terms of 1-year, 2-year, up to 10-year are shown.

		$\pi_i(t,T)$				$(T-t)^{-1}\ln\frac{1}{\pi_i(t,T)}$			
Rating	Maturity	Mean	Std.Dev.	Minimum	Maximum	Mean	Std.Dev.	Minimum	Maximum
Aaa	1-year	0.9892	0.0096	0.966	0.999	0.0109	0.0097	0.000	0.035
Aaa	2-year	0.9786	0.0191	0.932	0.999	0.0219	0.0197	0.000	0.071
Aaa	3-year	0.9679	0.0285	0.898	0.999	0.0330	0.0299	0.001	0.107
Aaa	4-year	0.9568	0.0357	0.865	0.999	0.0449	0.0380	0.001	0.145
Aaa	5-year	0.9450	0.0366	0.832	0.998	0.0573	0.0395	0.002	0.183
Aaa	6-year	0.9329	0.0465	0.741	0.998	0.0708	0.0528	0.002	0.300
Aaa	7-year	0.9068	0.0482	0.724	0.995	0.0994	0.0569	0.005	0.324
Aaa	8-year	0.8829	0.0547	0.645	0.945	0.1267	0.0679	0.056	0.439
Aaa	9-year	0.8779	0.0735	0.592	0.991	0.1341	0.0914	0.009	0.525
Aaa	10-year	0.8583	0.0754	0.668	0.997	0.1567	0.0897	0.003	0.403
Aa	1-year	0.9856	0.0093	0.967	0.999	0.0073	0.0047	0.000	0.017
Aa	2-year	0.9714	0.0184	0.935	0.999	0.0146	0.0095	0.000	0.033
Aa	3-year	0.9561	0.0289	0.897	0.999	0.0227	0.0152	0.001	0.054
Aa	4-year	0.9365	0.0373	0.847	0.985	0.0332	0.0201	0.008	0.083
Aa	5-year	0.9227	0.0409	0.822	0.979	0.0407	0.0226	0.010	0.098
Aa	6-year	0.8950	0.0555	0.779	0.975	0.0564	0.0317	0.012	0.125
Aa	7-year	0.8770	0.0562	0.755	0.944	0.0667	0.0329	0.029	0.141
Aa	8-year	0.8539	0.0596	0.716	0.935	0.0802	0.0359	0.033	0.167
Aa	9-year	0.8238	0.0637	0.675	0.930	0.0985	0.0397	0.036	0.196
Aa	10-year	0.8122	0.0571	0.669	0.906	0.1053	0.0363	0.049	0.201
A	1-year	0.9838	0.0107	0.956	0.999	0.0055	0.0036	0.000	0.015
A	2-year	0.9672	0.0217	0.913	0.999	0.0112	0.0075	0.000	0.030
A	3-year	0.9504	0.0327	0.875	0.998	0.0172	0.0115	0.001	0.045
A	4-year	0.9235	0.0504	0.709	0.980	0.0271	0.0189	0.007	0.115
A	5-year	0.8927	0.0508	0.731	0.969	0.0384	0.0194	0.011	0.104
A	6-year	0.8404	0.0663	0.676	0.950	0.0590	0.0267	0.017	0.130
A	7-year	0.7885	0.1151	0.158	0.934	0.0850	0.0736	0.023	0.614
A	8-year	0.7113	0.1171	0.153	0.877	0.1208	0.0812	0.044	0.625
A	9-year	0.6545	0.1294	0.050	0.848	0.1571	0.1365	0.055	0.999
A	10-year	0.6687	0.1386	0.021	0.850	0.1559	0.1730	0.054	1.286
Baa	1-year	0.9809	0.0100	0.962	0.999	0.0048	0.0026	0.000	0.010
Baa	2-year	0.9618	0.0201	0.924	0.999	0.0098	0.0052	0.000	0.020
Baa	3-year	0.9410	0.0316	0.858	0.999	0.0153	0.0084	0.000	0.038
Baa	4-year	0.9069	0.0424	0.759	0.982	0.0247	0.0120	0.004	0.069
Baa	5-year	0.8580	0.0601	0.606	0.947	0.0390	0.0190	0.014	0.125
Baa	6-year	0.8200	0.0723	0.499	0.916	0.0507	0.0243	0.022	0.174
Baa	7-year	0.7804	0.0873	0.330	0.902	0.0639	0.0338	0.026	0.277
Baa	8-year	0.7277	0.0915	0.340	0.870	0.0819	0.0373	0.035	0.270
Baa	9-year	0.6528	0.1073	0.166	0.775	0.1117	0.0572	0.064	0.449
Baa	10-year	0.6190	0.1327	0.124	0.812	0.1290	0.0783	0.052	0.521
Ba	1-year	0.9714	0.0183	0.919	1.000	0.0058	0.0038	0.000	0.017
Ba	2-year	0.9441	0.0358	0.843	1.000	0.0117	0.0077	0.000	0.034
Ba	3-year	0.9096	0.0520	0.771	1.002	0.0193	0.0115	-0.000	0.052
Ba	4-year	0.8431	0.0751	0.489	0.981	0.0350	0.0197	0.004	0.143
Ba	5-year	0.7684	0.1089	0.356	0.916	0.0551	0.0335	0.018	0.207
Ba	6-year	0.7001	0.1066	0.392	0.901	0.0739	0.0331	0.021	0.187
Ba	7-year	0.6430	0.0887	0.434	0.886	0.0902	0.0281	0.024	0.167
Ba	8-year	0.5591	0.1386	0.008	0.842	0.1302	0.1081	0.034	0.957
Ba	9-year	0.4889	0.1341	0.050	0.775	0.1564	0.0898	0.051	0.599
Ba	10-year	0.4384	0.1408	0.050	0.707	0.1808	0.0947	0.069	0.599

Table 3: **Relationship Between Credit Risk Premium Adjustment and Macroeconomic Variables.** This table shows the results of regressions of the credit risk premium adjustments for Aaa-rated, Aa-rated, A-rated, Baa-rated, and Ba-rated bonds on the macroeconomic variables of S&P500 stock market return, 3-month Treasury bill rate, and term structure slope of 30-year Treasury bond rate less 3-month bill rate. These variables are denoted as "Market", "T-Rate", and "Slope" respectively in the Table. The maturity indicates the maturity of the bonds in years from which the risk premium adjustment is computed. The values in parentheses are the t-values. $***$, $**$ and $*$ indicate significance at the 1%, 5% and 10% levels respectively. The standard errors are computed using White's Heteroskedasticity-Consistent estimators. We also report the Augmented Dickey-Fuller statistics of the dependent variable and regression residual for each regression. Rejection of unit root is indicated at the $***$ (1%), $**$ (5%) or $*$ (10%) level respectively.

Maturity	Constant	Market	T-Rate	Slope	Adj-R2	ADF (dep var)	ADF (resid)
			Panel I: Aaa-rated Bonds				
1	0.0031	0.0450	0.1580	-0.0026	0.017	-3.078**	-3.298**
	(0.227)	(1.124)	(0.764)	(-0.014)			
2	0.0063	0.0907	0.3177	-0.0071	0.017	-3.085**	-3.304**
	(0.229)	(1.121)	(0.760)	(-0.018)			
3	0.0097	0.1370	0.4791	-0.0135	0.017	-3.091**	-3.312**
	(0.230)	(1.118)	(0.756)	(-0.023)			
4	-0.0016	0.1564	0.8879	0.1666	0.021	-3.473***	-3.695***
	(-0.029)	(1.037)	(1.069)	(0.223)			
5	-0.0494	0.1942	1.8399*	0.8642	0.065	-4.092***	-4.552***
	(-0.803)	(1.597)	(1.860)	(1.099)			
6	-0.1360	0.2918**	3.4080**	2.0271**	0.105	-4.113***	-4.816***
	(-1.667)	(2.398)	(2.494)	(2.124)			
7	0.0941	-0.0424	0.3832	-0.4877	-0.015	-4.521***	-4.471***
	(1.361)	(-0.219)	(0.351)	(-0.576)			
8	-0.1454	-0.0171	4.1494***	3.5458***	0.066	-2.964**	-3.284**
	(-1.579)	(-0.081)	(2.791)	(2.720)			
9	0.4242***	-0.1493	-5.2617**	-2.0896	0.104	-2.755*	-3.006**
	(3.230)	(-0.466)	(-2.521)	(-1.382)			
10	0.2022**	-0.1648	-1.8775	1.7246	0.173	-2.675*	-3.070**
	(2.336)	(-0.526)	(-1.450)	(1.302)			
			Panel II: Aa-rated Bonds				
1	0.0261***	0.0050	-0.3135***	-0.1960**	0.080	-3.078**	-3.129**
	(4.003)	(0.257)	(-3.055)	(-2.355)			
2	0.0526***	0.0100	-0.6341***	-0.3968**	0.080	-3.085**	-3.133**
	(3.999)	(0.256)	(-3.057)	(-2.358)			
3	0.0849***	0.0601	-1.0322***	-0.6803**	0.097	-3.091**	-3.062**
	(4.168)	(1.026)	(-3.248)	(-2.563)			
4	0.0921***	-0.0741	-0.8940**	-0.7278**	0.016	-3.473***	-3.800***
	(3.591)	(-0.836)	(-2.188)	(-2.267)			
5	0.1284***	0.0364	-1.3988***	-1.0368**	0.063	-4.092***	-3.734***
	(4.100)	(0.403)	(-2.800)	(-2.630)			
6	0.1421***	0.2536**	-1.3431**	-1.1750**	0.090	-4.113***	-3.562**
	(3.531)	(2.283)	(-2.062)	(-2.358)			
7	0.1140***	-0.0609	-0.4171	-1.1515**	0.060	-4.521***	-4.908***
	(3.210)	(-0.481)	(-0.751)	(-2.343)			
8	-0.0325	0.0869	1.9429***	1.0017*	0.061	-2.964**	-4.841***
	(-0.795)	(0.623)	(2.961)	(1.713)			
9	0.0545	0.1181	0.3357	1.1059*	0.029	-2.755*	-5.398***
	(1.088)	(0.786)	(0.416)	(1.731)			
10	0.1124**	0.2017	-0.2556	0.0250	-0.012	-2.675*	-4.938***
	(2.251)	(1.465)	(-0.328)	(0.039)			

continuation of Table 3.

Maturity	Constant	Market	T-Rate	Slope	Adj-R2	ADF (dep var)	ADF (resid)
Panel III: A-rated Bonds							
1	0.0194***	-0.0002	-0.2260***	-0.1536**	0.059	-3.078***	-3.300**
	(3.755)	(-0.013)	(-2.789)	(-2.286)			
2	0.0415***	0.0039	-0.4882***	-0.3463**	0.068	-3.085**	-3.114**
	(3.897)	(0.120)	(-2.943)	(-2.464)			
3	0.0598***	0.0281	-0.6955***	-0.4836**	0.060	-3.091**	-3.170**
	(3.753)	(0.583)	(-2.793)	(-2.311)			
4	0.1004***	-0.0632	-1.0419***	-1.0579**	0.072	-3.473***	-3.870***
	(3.622)	(-0.621)	(-2.763)	(-2.573)			
5	0.1187***	0.0538	-1.0760***	-1.3477***	0.210	-4.092***	-3.818***
	(4.738)	(0.632)	(-2.914)	(-3.839)			
6	0.0640**	-0.0380	0.2983	-0.7382*	0.157	-4.113***	-3.600***
	(2.118)	(-0.363)	(0.622)	(-1.821)			
7	0.1714	-0.0476	-0.7007	-2.2386	0.046	-4.521***	-1.398
	(1.476)	(-0.194)	(-0.447)	(-1.283)			
8	0.3027*	0.1216	-2.9063	-2.1531	-0.006	-2.964**	-2.224
	(1.892)	(0.482)	(-1.386)	(-0.898)			
9	0.5045**	-0.0021	-5.1135*	-4.8725	0.006	-2.755*	-3.056**
	(2.132)	(-0.005)	(-1.680)	(-1.343)			
10	0.4478	0.1335	-4.1073	-4.5276	-0.012	-2.675*	-2.487
	(1.324)	(0.251)	(-0.954)	(-0.868)			
Panel IV: Baa-rated Bonds							
1	0.0106***	-0.0054	-0.1232**	-0.0044	0.210	-3.078**	-2.600
	(3.469)	(-0.557)	(-2.566)	(-0.114)			
2	0.0227***	-0.0084	-0.2675***	-0.0301	0.193	-3.085**	-2.431
	(3.601)	(-0.428)	(-2.719)	(-0.369)			
3	0.0364***	0.0165	-0.4284***	-0.0835	0.133	-3.091**	-2.343
	(3.542)	(0.500)	(-2.747)	(-0.602)			
4	0.0600***	-0.0847	-0.5452***	-0.4014*	0.048	-3.473***	-3.926***
	(4.148)	(-1.197)	(-2.763)	(-1.828)			
5	0.1153***	-0.0163	-0.9756***	-1.3374***	0.200	-4.092***	-4.117***
	(5.594)	(-0.249)	(-3.361)	(-4.365)			
6	0.0551	0.0595	0.1876	-0.5675	0.113	-4.113***	-4.959***
	(1.262)	(0.650)	(0.257)	(-1.102)			
7	0.1096**	0.1006	-0.4610	-1.0886	0.067	-4.521***	-1.932
	(2.044)	(0.887)	(-0.596)	(-1.431)			
8	0.1851***	0.0236	-1.5373*	-1.4069	0.014	-2.964**	-2.422
	(2.849)	(0.161)	(-1.728)	(-1.489)			
9	0.1251	0.0111	0.1999	-0.9626	0.018	-2.755*	-3.122**
	(1.407)	(0.061)	(0.165)	(-0.725)			
10	0.1560	0.1227	-0.0962	-1.0393	-0.011	-2.675*	-4.114***
	(1.066)	(0.521)	(-0.044)	(-0.504)			
Panel V: Ba-rated Bonds							
1	0.0084**	0.0092	-0.1190*	0.1094**	0.403	-3.078**	-3.346**
	(2.072)	(0.868)	(-1.786)	(2.272)			
2	0.0178**	0.0203	-0.2544*	0.2073**	0.395	-3.085**	-3.298**
	(2.150)	(0.946)	(-1.879)	(2.108)			
3	0.0426***	0.0589	-0.5650**	0.0514	0.264	-3.091**	-2.938**
	(3.009)	(1.632)	(-2.631)	(0.274)			
4	0.0964***	-0.1071	-0.9271**	-0.7627	0.034	-3.473***	-3.826***
	(3.155)	(-0.930)	(-2.367)	(-1.651)			
5	0.2284***	-0.0121	-2.3547***	-2.7961***	0.258	-4.092***	-2.921*
	(4.870)	(-0.110)	(-3.718)	(-4.054)			
6	0.1005**	0.0301	0.1030	-1.3367**	0.257	-4.113***	-3.200**
	(2.607)	(0.250)	(0.187)	(-2.403)			
7	0.1751***	0.0711	-1.0627*	-1.5849***	0.157	-4.521***	-3.707***
	(4.352)	(0.627)	(-1.741)	(-2.890)			
8	0.0940	-0.7328	0.7224	0.5670	0.002	-2.964**	-4.166***
	(0.366)	(-1.021)	(0.157)	(0.200)			
9	0.2735	-0.8232	-1.3971	-1.7766	0.030	-2.755*	-4.646***
	(1.496)	(-1.365)	(-0.469)	(-0.771)			
10	0.2069	-0.5752	-0.0656	-0.6219	-0.012	-2.675*	-4.773***
	(1.070)	(-0.938)	(-0.021)	(-0.257)			