Estimating the Tax and Credit-Event Risk Components of Credit Spreads *

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November 18, 2015

*We thank Andrew Boettcher, Olena Chyruk, Murray Frank, Alex Justiniano, and Leo Melosi for comments and suggestions. Kyal Berends, Tyna Eloundou Nekoul, Scott Lee, and Harvey Stephenson provided excellent research assistance. All remaining errors are our own. The views expressed herein are those of the authors and not necessarily those of the Federal Reserve Bank of Chicago or the Federal Reserve System.

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Abstract

This paper argues that tax liabilities explain a large fraction of observed short-maturity investment-grade (IG) spreads, but credit-event premia do not. First, we extend Duffie and Lando (2001) by permitting management to issue both debt and equity. Rather than defaulting, managers of IG firms who receive bad private signals conceal this information and service existing debt via new debt issuance. Consistent with empirical observation, this strategy implies that IG firms have virtually zero credit-event risk (at least until they become "fallen angels"). Second, we provide empirical evidence that short maturity IG spreads are mostly due to taxes. By properly accounting for the tax treatment of capital gains and interest income associated with bond investments, we reconcile this finding with the previous literature which argues against a significant tax component to spreads.

1 Introduction

How did you go bankrupt?"

Two ways. Gradually, then suddenly.

Ernest Hemingway, The Sun Also Rises.

Most empirical studies of corporate bond yields find evidence of a "credit spread puzzle" in that it is difficult to explain observed spreads between corporate bond yields and Treasury yields in terms of expected losses and standard measures of risk. This credit spread puzzle is most striking for short-maturity investment-grade (IG) debt, since historical default rates for these bonds are extremely low. Several explanations for the credit spread puzzle have been suggested in the literature, including: i) illiquidity premia, ii) tax asymmetry (i.e., corporate bonds, but not Treasuries, are subject to state taxation) and iii) jump-to-default (or credit-event) premia.[1] This paper argues that taxes explain a large fraction of observed short-maturity IG spreads, but credit-event premia do not.

The notion of credit-event (or jump-to-default) risk arises naturally in reduced-form models of default (e.g., Duffie and Singleton (1997), Jarrow, Lando and Turnbull (1997)), in which default is modeled as an unpredictable jump event. If a sufficiently large premium is attributed to credit-event risk, then short-maturity spreads can be "explained" through this channel. In their seminal paper, Duffie and Lando (DL, 2001) provide an economic justification for reduced-form models. They investigate the optimal behavior of a manager of a firm that issues equity to service debt in place. DL show that if this manager receives a sufficiently bad private signal, then it will be in the best interest of shareholders for the manager to declare default rather than have them continue to service debt payments. From an outsider's information set, such a default will appear as an unexpected credit-event, which can be characterized by a default intensity process

[1] Another explanation is provided by Feldhutter and Schaefer (2015), who argue that historical default rates provide a very noisy estimate for ex-ante expected default rates, leaving the possibility that ex-ante default expectations are much higher than ex-post observations. However, they acknowledge that even their explanation cannot explain short-maturity (less than one year) IG spreads.

similar to those specified by reduced form models.

In this paper, we build on the insights of DL by allowing the manager to issue both equity *and* debt to service current liabilities. Rather than choosing to default, a manager will maximize shareholder value by concealing bad private information and instead issue new debt to service the old debt. Intuitively, because equity holders receive zero payoff in the case of bankruptcy, it is in their best interest to maintain solvency so long as they do not need to invest more funds into the firm. As such, so long as the firm has IG-status (i.e., has unused debt capacity), this firm is not subject to jump-to-default risk even if the manager receives a bad private signal. In the absence of jump-to-default risk, there is no compensation for such risk, and short-term spreads cannot be explained due to a credit-event channel. Only after debt capacity has been used up and the firm has reached "fallen-angel" status is jump-to-default possible.

To test our model, we estimate empirical default intensities as a function of the firm's rating. We show that (annualized) default rates for IG firms are tiny and insignificant at the one- to three-month horizon, and are significantly smaller than default rates at the one-year horizon. Consistent with the predictions of our model, only after a firm loses its IG status and its spreads increase significantly does default risk become non-negligible. As such, a portfolio consisting only of bonds with spreads that are commensurate with IG status is subject to virtually zero default risk and, therefore, should not command a credit-event premium. In contrast, as predicted by our model, firms with the worst ratings (C or lower) can exhibit jump-to-default behavior. This is consistent with the empirical findings of Davydenko, Strebulaev and Zhao (2012), who report a significant likelihood of a jump-to-default, accompanied by a significant loss in firm value at the default event for speculative-grade bonds.

Next, we explore the possibility that taxes explain a large fraction of short maturity IG spreads. As noted by Elton et al. (2001), corporate bonds are subject to state taxes, whereas Treasuries are not. For example, a 5% state tax coupled with an 8% coupon can explain a 40 bp spread, which would capture the majority of the 54 bp average spread of commercial paper over Treasury yields over the past several decades. However, both Longstaff, Mithal and Neis (LMN, 2005) and Chen, Lesmond and Wei (CLW, 2007) argue against a tax channel. While their approaches differ, their basic argument is that regressing yields on coupon level generates a very small regression coefficient, which they interpret as a measure of the effective state tax. We argue, however, that once the tax code is taken into account, then a low estimate for the regression coefficient is to be expected and should not be interpreted as an estimate of the state tax rate. To gain some intuition for this result, consider a firm that issues two bonds with the same maturity but with different coupon rates. The coupon on the first bond is chosen so that the bond sells for par, whereas the coupon on the second bond is higher, which in turn implies that the bond is selling for a premium. If the bond is held to maturity and there is no default, then due to this premium the investor suffers a capital loss at maturity. According to the tax code, this loss reduces the tax liability. Moreover, this loss is typically amortized over the life of the bond. Below we demonstrate that once this tax feature is accounted for, then a regression of yield on coupon level *should* produce a low regression coefficient even if there is a sizeable tax component to spreads. In that sense, the findings of LMN and CLW do not provide evidence against a significant tax component to spreads.

Interestingly, while the literature has focused on the asymmetric aspect of state taxes, below we argue that the spread component due to federal taxes is also significant, even though both corporate bonds and Treasuries are subject to federal taxation. To understand this intuition, assume that the state tax rate is zero, the federal tax rate is θ_F, and that investors require an after-tax spread of size "*spread*" to compensate them for liquidity/credit risk. This implies that the observed (i.e., pre-tax) credit spread equals $(y - r) = \left(\frac{spread}{1-\theta_F}\right)$. That is, the presence of a liquidity/credit risk component implies that observed credit spreads are amplified by the factor $\left(\frac{1}{1-\theta_F}\right)$. We demonstrate below that this amplification factor also significantly increases the component of spread due to state taxes.

Building on this intuition, we specify a simple model that links credit spreads to the level of the risk-free rate, the federal and state marginal tax rates θ_F and θ_S, and the after-tax compensation for liquidity and credit risk. The main insight of this model is that, after controlling for credit and liquidity risk, the pre-tax spread depends on the the risk-free rate only if there is a state tax impact (i.e., $\theta_S > 0$) for the marginal bond investor. Intuitively, since state taxes are paid on the entire interest payment, periods of high risk-free rates are also periods of high state taxes, implying that spreads must be higher to compensate an investor who is subject to state taxes. Casual inspection of short-maturity commercial-paper spreads and T-Bill rates in Figure 1 supports this relation.

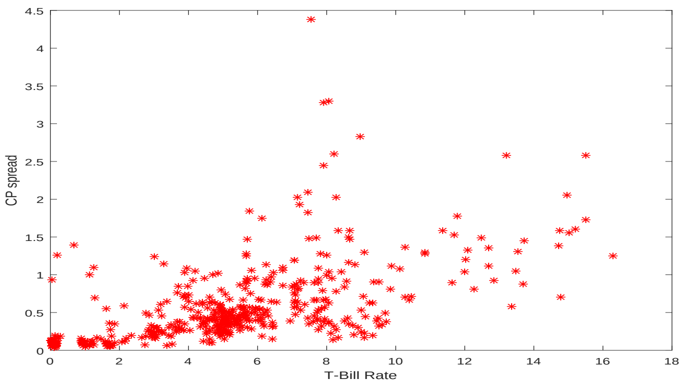

Figure 1: **Commercial Paper Spread and Treasury Bill Rate.** The figure plots the spread between three-month commercial paper and T-Bill rates (vertical axis) against the three-month T-Bill rate (horizontal axis). Sample period: April 1971 – June 2013.

We estimate the model on two series of short-maturity investment-grade rates: three-month commercial paper and certificates of deposit.[2] The risk-free rate is the three-month T-Bill rate, while the maximal individual federal tax rate gives us a measure of θ_F. To approximate the after-tax component of spreads, we use a comprehensive set of indicators of credit and liquidity risk in debt markets.

Over the 1971-2013 period, we find the marginal state tax rate θ_S to be 2%, an

[2]CD's are FDIC-insured only up to some maximum.

estimate that is significant at any conventional level. The state-tax component of credit spreads is particularly big in the 1970s and 1980s, when the T-Bill rate reaches 17% and the state-tax channel is powerful. In later years T-Bill rates decline and the state-tax component decreases as well; it vanishes during the Great Recession when the risk-free rate is at the zero lower bound. Federal taxes are an important component of spreads too, especially in the early part of the sample where top individual rates are high. Overall, federal and state taxes account for two thirds of credit spreads, while liquidity and credit risk explain only one-third. The results are robust to estimation over the post-1985 period, as well as other sub-samples.

Related Literature. In this paper, we focus on IG spreads of short maturity debt (i.e., commercial paper and certificates of deposit) for two reasons. First, the "credit spread puzzle" is most prevalent for short maturity IG bonds. Indeed, a growing literature (e.g., Chen et al. (2009), Bhamra et al. (2009) and Chen (2010)) argues that IG spreads on maturities greater than a few years can be explained by combining pricing kernels that capture time varying Sharpe ratios over the business cycle with models that match the empirically observed clustering of defaults during recessions. Second, whereas the previous literature (e.g., Sarig and Warga (1989) and Delianides and Geske (2001)) has argued against a tax component to spreads because the clientele of long-maturity corporate bond investors (e.g., pensions, 401(k) and retirement accounts) tend to have tax-advantaged status, this may be less true for investors of short maturity IG bonds. Indeed, for commercial paper and IG bonds with maturities less than 13 months, the dominant purchasers are money market mutual funds (MMMF). It is likely that the representative investor in MMMF's may not possess the same level of tax advantage status as owners of longer maturity bonds. The same applies to investors in certificates of deposit.

Our paper builds on the rather large literature that focuses on decomposing credit spreads into components of expected loss, risk premia, liquidity premia and taxes.[3] In this literature, there is considerable disagreement on the magnitude of the credit-event premium. For example, in his benchmark model, Driessen (2005) associates 31 bps with this channel. Related, he estimates the ratio of risk-neutral to actual default intensity to be $\left(\frac{\lambda^Q}{\lambda^P}\right) = 2.3$. Similar results are reported by Saita (2006) and Berndt et al (2009). However, as noted in Bai et al (2015), these authors do not identify the credit-event premia by estimating how proxies for their pricing kernel covary with bond returns, but rather attribute whatever they cannot explain through other channels to this credit risk channel. Moreover, Bai et al. (2015) show that if the pricing kernel can be proxied by returns on the corporate bond index, then even if we wish to attribute 50 bps to the combined credit-event channel and "contagion" channel (captured by movement in the corporate bond index during the credit-event), then at most 2 bps of those 50 bps can be attributed to the credit-event channel. In contrast to Bai et al (2015) who argue that equilibrium considerations imply that the risk premium $\left(\frac{\lambda^Q}{\lambda^P}\right)$ must be small, we argue that λ^P is itself very small, and that estimating λ^P using the one year historical default rate significantly exaggerates the credit-event risk faced by an investment strategy which holds only corporate bonds with IG-level spreads.

The rest of the paper is organized as follows. We revisit the Duffie and Lando (2001) model in Section 2 and generalize it in Section 3 to a setting in which the manager can issue *both* debt and equity to service debt in place. We show in a model calibration that IG firms do not jump to default and provide empirical support for the model predictions. In Section 4, we provide empirical evidence that short-maturity IG credit spreads are mostly due to taxes. We reconcile this finding with the contrasting conclusions of the previous literature by properly accounting for the tax treatment of capital gains and interest income associated with a bond investment. We conclude in Section 5, and relegate most proofs to the Appendix.

[3] See, e.g., Elton et al. (2001), Feldhütter and Schaefer (2014), LMN, CLW.

2 Duffie-Lando (2001) Revisited

The seminal paper of Duffie and Lando (DL, 2001) provides a sound economic justification for the so-called reduced-form models of default, in which default is characterized by an unpredictable intensity (or hazard rate) process. The reduced form framework provides a very different mechanism for capturing default than that suggested by the so-called structural models of default (e.g., Merton (1974), Leland (1994)), in which firm value follows a diffusion process and default is predictable. Intuitively DL argue that if debt in place can be serviced only via equity issuance, and a manager observes a sufficiently bad private signal, then she will be acting in the best interest of shareholders to declare default rather than have shareholders service the existing debt. From an outsider's perspective who does not have access to this bad signal, default will appear as an unanticipated jump event.

We argue, however, that if the manager has the ability to raise both debt and equity, then it will be in the best interest of shareholders if she were to conceal her bad private signal and issue new debt to service the debt in place. Indeed, our findings are almost the reverse of those of DL: whereas DL demonstrate that a firm which is restricted to issuing only equity might jump to default even though the true value (i.e., value of firm given all public and private information) of the firm follows a diffusion process, below we demonstrate that IG firms will never jump to default even though the "true value" of the firm can jump down to a value that would imply default in a world with complete information.

Here we investigate an economy that is very similar to that of Duffie and Lando (2001) in that if the manager is restricted to issuing only equity and if she observes a sufficiently bad private signal, it will be in the shareholders best interest for her to default. We show that from an outsider's perspective this default is characterized by an intensity process which we identify. Then in the next section we demonstrate that if we investigate the same framework, but now allow management to issue both debt and equity to service debt in place, then firms with low spreads (i.e., "investment-grade" firms) and accordingly unused debt capacity will not jump to default, at least until their debt capacity has been used up and their spreads have increased to "speculative grade" levels.

Consider a firm that has a single project, which upon completion at random date-$\tilde{\tau}_P$ will pay out a terminal cash flow $\tilde{X}$. There are three independent sources of randomness in this model. First, the terminating cash flow $\tilde{X}$ can take on only one of two values: (X_H, X_L) with probabilities $(\pi_H, \pi_L = 1 - \pi_H)$. Second, the project completion date $\tilde{\tau}_P$ is chosen from an exponential distribution with intensity ω_P:

$$\pi(\tilde{\tau}_P \in (\tau_P, \tau_P + d\tau_P) | \tilde{\tau}_P > 0) \quad = \quad \omega_P e^{-\omega_P \tau_P} \mathbf{1}_{(\tau_P > 0)} \, d\tau_P. \tag{1}$$

Third, the manager of the firm learns the value of $\tilde{X}$ at some random date $\tilde{\tau}_M$ which is chosen from an exponential distribution with intensity ω_M:

$$\pi(\tilde{\tau}_M \in (\tau_M, \tau_M + d\tau_M) | \tilde{\tau}_M > 0) \quad = \quad \omega_M e^{-\omega_M \tau_M} \mathbf{1}_{(\tau_M > 0)} \, d\tau_M. \tag{2}$$

For convenience, we refer to the default date as $\tilde{\tau}_B$.

There are two relevant information sets in this model. The only information that outsiders have is whether or not at today's date-t the firm has defaulted $\mathbf{1}_{(\tilde{\tau}_B > t)}$ and whether or not a terminal cash flow has been paid $\mathbf{1}_{(\tilde{\tau}_P > t)}$. In contrast, the manager also knows whether or not she has received a signal $\mathbf{1}_{(\tilde{\tau}_M > t)}$, and if she has, what the true value of $\tilde{X}$ is. Note that the manager has no information advantage over outsiders at date-0.

At date-0, the firm finds itself with outstanding debt which commits the firm to making continuously-paid coupon payments of rate $c(0)$ and cumulative face value of $\frac{c(0)}{r+s}$, which is due upon project completion.[4] If default occurs before project completion $(\tilde{\tau}_B < \tilde{\tau}_P)$, then debtholders have claim to all future cash flows, which due to bankruptcy

[4]The spread s is meant to capture the fact that the bond was originally issued at par, and is not important for the results below.

costs associated with default are reduced by the fraction α. To account for tax benefits of debt, it is assumed that the government pays a fraction θ of coupon payments made.

Note that, since the firm generates no intermediate cash flows, it will need to issue either equity or debt in order to pay the interest payments owed until the project matures. In the first subsection, we restrict the manager to issuing only equity. In the second subsection, we permit the manager to issue both debt and equity.

2.1 Firm Restricted to Equity Issuance

In this section we investigate a framework that is analogous to that studied by Duffie and Lando (DL, 2000) in which the firm is restricted to issuing only equity in order to cover coupon payments due to existing debtholders (net of tax deductions) $(1-\theta)c(0)\,dt$. Similar to the findings of DL, we find that the default event is characterized by a hazard rate process $\lambda(t)$ from the viewpoint of outsiders.

We verify below that, for the calibration considered, it is in the shareholders' best interests to issue equity until either project completion or the manager receives a bad signal. As such, this model predicts that the default event is characterized as a hazard rate process from the viewpoint of outsiders, occurring at $\tilde{\tau}_{B} = \tilde{\tau}_{M}$ if the manager's signal reveals $\tilde{X} = X_{L}$. Since we specify equity to receive zero in default, we refer to the value of equity in this situation as $E_{L} = 0$. (The subscript "L" is to denote that the manager's signal was $\tilde{X} = X_{L}$).

2.1.1 Manager Receives Private Information

We solve the model via dynamic programming by working backwards through event time. Here we determine the value of equity from the viewpoint of the manager conditional upon her receiving a signal that the (eventual) project payoff will be high: $\tilde{X} = X_{H}$. Using standard Bayesian updating, it is well known from equation (1) that, conditional upon the project completion date being greater than today's date-t, the probability that the project completion date occurs in the next interval dt is

$$\pi(\tilde{\tau}_P \in (t, t+dt)|\tilde{\tau}_P > t) \;=\; \omega_P \mathbf{1}_{(\tau_P > t)}\, dt. \tag{3}$$

Thus the date-t value of the equity claim has three components: i) the after-tax liability that the firm must pay to cover coupon payments: $(1-\theta)c(0)\,dt$, ii) there is a probability $\omega_P\,dt$ that the project completes and equityholders receive a terminal dividend $\left(X_H - \frac{c(0)}{r+s}\right)$, iii) there is a probability $(1-\omega_P\,dt)$ that the project remains incomplete and therefore the equityholder still has claim to E_H. Hence, the value of equity from the manager's perspective at this time can be determined via:

$$E_H \;=\; \max\left[0,\, e^{-r\,dt}\left(-(1-\theta)c(0)\,dt + \omega_P\,dt\left(X_H - \frac{c(0)}{r+s}\right) + (1-\omega_P\,dt)E_H\right)\right]. \tag{4}$$

Intuitively, the form $\max(0,\cdot)$ captures the notion of limited liability. Simplifying, we find

$$E_H \;=\; \max\left[0,\, \left(\frac{1}{\omega_P + r}\right)\left[-(1-\theta)c(0) + \omega_P\left(X_H - \frac{c(0)}{r+s}\right)\right]\right]. \tag{5}$$

An analogous argument conditional upon the manager receiving a bad signal is

$$E_L \;=\; \max\left[0,\, \left(\frac{1}{\omega_P + r}\right)\left[-(1-\theta)c(0) + \omega_P\left(X_L - \frac{c(0)}{r+s}\right)\right]\right]. \tag{6}$$

We verify below that $E_L = 0$ and $E_H > 0$ for the parameterization we consider. That is, upon receiving bad information, the manager will choose to default, but upon receiving good information, the manager will decide to issue equity to pay for the coupon payments and avoid default.

2.1.2 Manager has Not Yet Received Information

We now move backwards to a prior time when the manager has not yet received a signal for $\tilde{X}$. Following an analogous argument, the value of equity at this time E_{NS} (the subscript NS implies "no signal") can be determined via:

$$\begin{aligned} E_{NS} \;=\; \max\Bigg[0,\, e^{-r\,dt}\Bigg[&-(1-\theta)c(0)\,dt + \pi_H\omega_P\,dt\left(X_H - \frac{c(0)}{r+s}\right) + \pi_L\omega_P\,dt\left(X_L - \frac{c(0)}{r+s}\right) \\ &+\pi_H\omega_M\,dt(E_H) + \pi_L\omega_M\,dt(E_L) + (1-(\omega_P+\omega_M)\,dt)E_{NS}\Bigg]\Bigg]. \end{aligned} \tag{7}$$

Intuitively, this equation can be understood as follows: over the following interval dt, i) with probability $\pi_H \omega_P \, dt$, the project completes and $\tilde{X} = X_H$, implying that equityholders receive a terminal payout of $\left(X_H - \frac{c(0)}{r+s}\right)$, ii) with probability $\pi_L \omega_P \, dt$, the project completes and $\tilde{X} = X_L$, implying that equityholders receive a terminal payout of $\left(X_L - \frac{c(0)}{r+s}\right)$, iii) with probability $\pi_H \omega_M \, dt$, the manager receives the signal $\tilde{X} = X_H$, implying that equityholders have a claim that is worth E_H, iv) with probability $\pi_L \omega_M \, dt$, the manager receives the signal $\tilde{X} = X_L$, implying that equityholders have a claim that is worth E_L, and v) with probability $(1 - (\omega_P + \omega_M) \, dt)$ there is no additional information, implying that equityholders still own a claim worth E_{NS}.

Setting $E_L = 0$ and simplifying, we find

$$\begin{aligned} E_{NS} &= \max \left[0, \left(\frac{1}{r + \omega_P + \omega_M} \right) \left[-(1-\theta)c(0) + \pi_H \omega_P \left(X_H - \frac{c(0)}{r+s} \right) \right. \right. \\ &\qquad \left. \left. + \pi_L \omega_P \left(X_L - \frac{c(0)}{r+s} \right) + \pi_H \omega_M \, E_H \right] \right]. \qquad (8) \end{aligned}$$

For the parameterization we study, $E_{NS} > 0$. That is, if the manager has received no information, she will choose to issue additional equity rather than declare default. Hence, default will occur only if the manager receives a bad signal before the project completion date, and it will occur exactly at the moment the manager receives this bad signal.

2.1.3 Price of Debt

From the perspective of outsiders, default intensity is a decreasing function of time because, so long as the firm remains solvent, there is an increasing probability that the manager has received a good signal regarding $\tilde{X}$ (rather than no signal at all). Indeed, consider any date $t > 0$ where the only information debtholders have is that the firm has not yet defaulted (i.e., $\tilde{\tau}_B > t$) and that the project is not yet completed (i.e., $\tilde{\tau}_P > t$). There are thus two possibilities: the manager has received no signal, or the manager has received a good signal (recall, the manager immediately defaults upon receiving bad news). From Bayes' rule we have:

$$\pi(\tilde{\tau}_M > t | \tilde{\tau}_B > t,\, \tilde{\tau}_P > t) \;\; = \;\; \frac{\big(1\big)\,(e^{-\omega_M t})}{\big(1\big)\,(e^{-\omega_M t}) + (\pi_H)\,(1 - e^{-\omega_M t})}. \tag{9}$$

where we have used the fact that $\pi(\tilde{\tau}_B > t | \tilde{\tau}_M > t,\, \tilde{\tau}_P > t) = 1$, since default will not occur before the manager receives a signal. Similarly, $\pi(\tilde{\tau}_B > t | \tilde{\tau}_M < t,\, \tilde{\tau}_P > t) = \pi_H$, since the probability that default has not occurred (and in fact, will never occur) after the signal is received is the probability that the signal is $\tilde{X} = X_H$.

The default intensity from the viewpoint of an outsider is defined as the probability of defaulting in an interval $(t + dt)$ given that it has not defaulted prior to t:

$$\lambda(t)\, dt \;\; \equiv \;\; \pi(\tilde{\tau}_B \in (t, t+dt) | \tilde{\tau}_B > t,\, \tilde{\tau}_P > t). \tag{10}$$

Within our framework, this is equivalent to

$$\begin{aligned} \lambda(t) \;\; &= \;\; \left(\frac{1}{dt}\right) \pi(\tilde{X} = X_L, \tilde{\tau}_M \in (t, t+dt) | \tilde{\tau}_B > t,\, \tilde{\tau}_P > t) \\ &= \;\; \pi_L\, \omega_M \left(\frac{e^{-\omega_M t}}{e^{-\omega_M t} + \pi_H\,(1 - e^{-\omega_M t})} \right). \end{aligned} \tag{11}$$

That is, as in DL, our framework implies that when the manager is constrained to issuing only equity to service current debt obligations, default is characterized by a hazard rate process from the viewpoint of outsiders.

The price of the bond has 3 components: i) the claim to coupon payments for all dates $t < \min(\tilde{\tau}_P,\, \tilde{\tau}_B)$, ii) the claim to principal payments at project completion date $\tilde{\tau}_P$ if $\tilde{\tau}_P < \tilde{\tau}_B$, and iii) the claim to recovery at the default event $\tilde{\tau}_B$ if $\tilde{\tau}_B < \tilde{\tau}_P$. Consistent with our argument that agents are subject to taxation, we account for the fact that losses due to default are tax-deductible.

The portion of debt due to a claim to coupons is

$$\begin{aligned} B_{coup}(\mathbf{1}_{(\tilde{\tau}_P > t)}, \mathbf{1}_{(\tilde{\tau}_B > t)}) \;\; &= \;\; \mathbf{1}_{(\tilde{\tau}_P > t)} \mathbf{1}_{(\tilde{\tau}_B > t)} \mathrm{E} \left[\int_t^{\infty} \mathbf{1}_{(\tilde{\tau}_P > T)} \mathbf{1}_{(\tilde{\tau}_B > T)} e^{-r(T-t)} c(0)\, dT \,\middle|\, \tilde{\tau}_P > t,\, \tilde{\tau}_B > t \right] \\ &= \;\; \left(\frac{\mathbf{1}_{(\tilde{\tau}_P > t)} \mathbf{1}_{(\tilde{\tau}_B > t)} c(0)}{e^{-\omega_m t} + \pi_H\,(1 - e^{-\omega_M t})} \right) \left[\frac{1 - \pi_H}{r + \omega_P + \omega_M} e^{-\omega_M t} + \frac{\pi_H}{r + \omega_P} \right]. \end{aligned} \tag{12}$$

The claim to face value $\left(\frac{c(0)}{r+s}\right)$ conditional upon $\tilde{\tau}_P < \tilde{\tau}_B$ is

$$
\begin{aligned}
B_{FV}(\mathbf{1}_{(\tilde{\tau}_P>t)}, \mathbf{1}_{(\tilde{\tau}_B>t)}) &= \left(\frac{c(0)}{r+s}\right) \mathrm{E}\left[e^{-r(\tilde{\tau}_P-t)}\mathbf{1}_{(\tilde{\tau}_P<\tilde{\tau}_B)} \,\middle|\, \tilde{\tau}_P > t,\, \tilde{\tau}_B > t\right] \mathbf{1}_{(\tilde{\tau}_P>t)}\mathbf{1}_{(\tilde{\tau}_B>t)} \quad (13)\\
&= \left(\frac{c(0)}{r+s}\right)\left(\frac{1}{\pi_H + \pi_L e^{-\omega_M t}}\right)\left[\left(\frac{\pi_H\,\omega_P}{r+\omega_P}\right) + \left(\frac{\omega_P \pi_L e^{-\omega_M t}}{r+\omega_P+\omega_M}\right)\right]\mathbf{1}_{(\tilde{\tau}_P>t)}, \mathbf{1}_{(\tilde{\tau}_B>t)}.
\end{aligned}
$$

Investors realize that default occurs only if the manager has received a bad signal $\tilde{X} = X_L$ at a time τ_M that occurs before project completion τ_P. As such, for the case $\tau_M > t$, $\tau_P > t$, the claim to recovery from an outsider's perspective satisfies

$$
\begin{aligned}
B_{recov}(\tilde{\tau}_P > t, \tilde{\tau}_M > t) &= \mathrm{E}_t^Q\left[(1-\alpha)X_L e^{-r(\tau_P-t)}\,\mathbf{1}_{(\tilde{X}=X_L)}\mathbf{1}_{(\tau_M<\tau_P)} | \tilde{\tau}_P > t, \tilde{\tau}_M > t\right]\\
&= \left(\frac{(1-\alpha)X_L\pi_L e^{-\omega_M t}}{\pi_H + \pi_L e^{-\omega_M t}}\right)\left(\frac{\omega_P}{r+\omega_P}\right)\left(\frac{\omega_M}{r+\omega_P+\omega_M}\right). \quad (14)
\end{aligned}
$$

2.1.4 Calibration

The calibration parameters are given in Table 1, and the output is given in Table 2. The parameters have been chosen so that the manager will default if she observes a bad signal (i.e., $E_L = 0$), and that she will not default if she receives a good signal (i.e., $E_H > 0$) or has not yet received a signal (i.e., $E_{NS} > 0$).

Because debt maturity is random, there is no unique way to define the yield. For a given coupon c and given face value of debt $\left(\frac{c}{r+s}\right)$, we choose what we consider to be the most logical definition of yield y by finding that value of y which equates the two sides of the equation

$$
\begin{aligned}
B(t) &= \mathrm{E}^Q\left[\int_t^{\tilde{\tau}^P} ds\, c\, e^{-y(s-t)} + \left(\frac{c}{r+s}\right)\, e^{-y(\tilde{\tau}^P-t)}\right]\\
&= \frac{c}{y} + \left(\frac{c}{r+s} - \frac{c}{y}\right)\left(\frac{\omega_P}{\omega_P + y}\right). \quad (15)
\end{aligned}
$$

Furthermore, we set the spread parameter s so that the initial bond price sells at par: $B(0) = \frac{c(0)}{r+s}$. Hence, the initial yield is $y(0) = (r+s)$.

The important take-away from this section is that, just as in DL, if the firm is restricted to issuing only equity to service existing debt payments, and if a manager receives a sufficiently bad signal, then it will be in the best interest of shareholders for the manager to default on the debt. Moreover, from an outsider's information set, default will be appear as an unpredictable point process with intensity that is characterized by equation (11).

3 Firm Permitted to Issue Equity and Debt

In this section, we generalize the insights of DL to a setting in which the manager can issue *both* debt and equity to service current debt. We demonstrate that in this case, rather than defaulting, a manager will maximize shareholder value by concealing bad private information and instead issue new debt to service the old debt. Because equity holders are assumed to receive zero in the case of bankruptcy, it is intuitive that shareholders are better off avoiding bankruptcy so long as they do not need to pay new funds into the firm via equity issuance to keep the firm solvent. As such, so long as an "investment-grade" firm with remaining debt capacity is able to issue new debt to service old debt, default will not occur. Hence, in this generalized DL framework, IG firms are not subject to jump-to-default risk due to a manager receiving a bad private signal, at least not until their debt capacity has been used up and their spreads have reached speculative-grade levels.

To simplify matters, we consider only strategies where equity issuance is never chosen before debt capacity is used up. Note that allowing for equity issuance before all debt capacity is utilized will not change the main result of this section, namely, that "investment grade" firms with unused debt capacity do not jump to default due to bad private signals.

To solve this model, we identify the process for which the cumulative coupon payment $c(t)$ evolves over time. To simplify the calculations, we assume that the face value of debt for all dates-t is equal to $\left(\frac{c(t)}{r+s}\right)$, implying that we need to keep track of only one time series, namely the cumulative coupon process $\{c(t)\}$. We assume that covenants are in place that restrict the amount of outstanding face value of debt to not exceed X_L, which in turn guarantees that default will not occur on the project completion date. We define τ^* as the date in which the face value of debt reaches X_L, and hence, the date at which all debt capacity is used up. Although we do not know a priori what the value of τ^* is, we do know that the cumulative coupon due at this date is determined implicitly via:

$$\frac{c(\tau^*)}{r+s} = X_L. \tag{16}$$

The value of τ^* will be determined below.

At any date-t prior to both τ^* and $\tilde{\tau}_P$, the firm must service the current coupon due $c(t)\,dt$ over the next interval dt. The firm accomplishes this by issuing new debt. Because government pays the fraction $\theta c(t)\,dt$ in terms of tax benefits to debt, the firm therefore must raise $(1-\theta)c(t)\,dt$. In order to do so, the new debtholders must be given claims to future cash flows whose present value is equal to what they pay.

In order to entice the new debtholder at date-v to pay $(1-\theta)c(v)\,dt$, the firm issues a new bond with coupon payment rate of $dc(v)$ and face value $\left(\frac{dc(v)}{r+s}\right)$ upon project completion. In addition, the bondholder has claim to *pari passu* recovery in the event of default. As such, by paying $(1-\theta)c(v)\,dt$ at date-v, the new corporate bond investor has claim to the following cash flows:

1) for all time intervals $(t, t+dt)$ such that $v < t < \tau^*$ and $t < \tau_P$, the bondholder will receive a coupon of size $dc(v)\,dt$.
2) for all time intervals $(t, t+dt)$ such that $v < t < \tau^*$ and $\tau_P \in (t, t+dt)$, the bondholder will receive face value $\left(\frac{dc(v)}{r+s}\right)$.
3) Assuming $\tau_P > \tau^*$, at time τ^*, investors will learn if the manager has received a bad signal prior to τ^*. If so, the firm will default immediately, and the bondholder will receive the fraction $\left(\frac{dc(v)}{c(\tau^*)}\right)$ of the recovery value of the firm.
4) for all time intervals $(t, t+dt)$ such that $\tau^* < t < \min(\tau_P, \tau_B)$, the bondholder

will receive a coupon of size $dc(v)\,dt$.

5) for all time intervals $(t, t+dt)$ such that $(\tau^* < t < \tau_{_B})$ and $\tau_{_P} \in (t, t+dt)$, the bondholder will receive face value $\left(\frac{dc(v)}{r+s}\right)$.

6) for all time intervals $(t, t+dt)$ such that $(\tau^* < t < \tau_{_P})$ and $\tau_{_B} \in (t, t+dt)$, the bondholder will receive receive the fraction $\left(\frac{dc(v)}{c(\tau^*)}\right)$ of the recovery value of the firm.

We now determine the date-v present value of these six claims:

$$\begin{aligned} B_{_{coup1}}(v < \tau_{_P}) &= \mathrm{E}^Q\left[\int_v^{\tau^*} dt\, dc(v)\, e^{-r(t-v)}\, \mathbf{1}(t < \tilde{\tau}_{_P}) | v < \tilde{\tau}_{_P}\right] \\ &= \left(\frac{dc(v)}{r+\omega_{_P}}\right)\left[1 - e^{-(r+\omega_P)(\tau^*-v)}\right] \end{aligned} \tag{17}$$

$$\begin{aligned} B_{_{FV1}}(v < \tau_{_P}) &= \mathrm{E}^Q\left[\left(\frac{dc(v)}{r+s}\right) e^{-r(\tilde{\tau}_P - v)}\mathbf{1}(\tilde{\tau}_{_P} < \tau^*) | \tilde{\tau}_{_P} > v\right] \\ &= \left(\frac{dc(v)}{r+s}\right)\left(\frac{\omega_P}{\omega_{_P}+r}\right)\left[1 - e^{-(r+\omega_P)(\tau^*-v)}\right] \end{aligned} \tag{18}$$

$$\begin{aligned} B_{_{recov,1}}(v < \tau_{_P}) &= \left(\frac{dc(v)}{c(\tau^*)}\right)\mathrm{E}^Q\left[(1-\alpha)\tilde{X}e^{-r(\tilde{\tau}_P - v)}\mathbf{1}(\tilde{X} = X_{_L})\mathbf{1}(\tilde{\tau}_{_M} < \tau^*)\mathbf{1}(\tilde{\tau}_{_P} > \tau^*) | \tilde{\tau}_{_P} > v\right] \\ &= \left(\frac{dc(v)}{c(\tau^*)}\right)(1-\alpha)\pi_{_L} X_{_L}\left(\frac{\omega_P}{\omega_{_P}+r}\right)\left[1 - e^{-\omega_M \tau^*}\right] e^{-(r+\omega_P)(\tau^*-v)} \end{aligned} \tag{19}$$

$$\begin{aligned} B_{_{coup,2}}(v < \tau_{_P}) &= \mathrm{E}^Q\left[\int_{\tau^*}^{\infty} dt\, dc(v)\, e^{-r(t-v)}\mathbf{1}(\tilde{\tau}_{_M} > t)\mathbf{1}(\tilde{X} = X_{_L})\mathbf{1}(\tilde{\tau}_{_P} > \tilde{\tau}_{_M}) | \tilde{\tau}_{_P} > v\right] \\ &+\mathrm{E}^Q\left[\int_{\tau^*}^{\infty} dt\, dc(v)\, e^{-r(t-v)}\mathbf{1}(\tilde{\tau}_{_P} > t)\left[\mathbf{1}(\tilde{X} = X_{_H}) + \mathbf{1}(\tilde{X} = X_{_L})\mathbf{1}(\tilde{\tau}_{_P} < \tilde{\tau}_{_M})\right] | \tilde{\tau}_{_P} > v\right] \\ &= dc(v)\, e^{-(\omega_P + r)(\tau^*-v)}\left[\frac{\pi_{_H}}{r+\omega_{_P}} + \frac{\pi_{_L} e^{-\omega_M \tau^*}}{r+\omega_{_P}+\omega_{_M}}\right] \end{aligned} \tag{20}$$

$$\begin{aligned} B_{_{FV,2}}(v < \tau_{_P}) &= \\ &\mathrm{E}^Q\left[e^{-r(\tau_P - v)}\left(\frac{dc(v)}{r+s}\right)\mathbf{1}(\tilde{\tau}_{_P} > \tau^*)\left[\mathbf{1}(\tilde{X} = X_{_H}) + \mathbf{1}(\tilde{X} = X_{_L})\mathbf{1}(\tilde{\tau}_{_P} < \tilde{\tau}_{_M})\right] | \tilde{\tau}_{_P} > v\right] \\ &= \left(\frac{dc(v)}{r+s}\right) e^{-(r+\omega_P)(\tau^*-v)}\left[\pi_{_H}\left(\frac{\omega_P}{r+\omega_{_P}}\right) + \pi_{_L}\, e^{-\omega_M \tau^*}\left(\frac{\omega_P}{r+\omega_{_P}+\omega_{_M}}\right)\right] \end{aligned} \tag{21}$$

$$
\begin{aligned}
B_{recov,2}(v < \tilde{\tau}_P) &= \mathrm{E}^Q\left[e^{-r(\tau_P - v)}\left(\frac{dc(v)}{c(\tau^*)}\right)(1-\alpha)\,\tilde{X}\mathbf{1}(\tilde{X} = X_L)\mathbf{1}(\tilde{\tau}_M > \tau^*)\mathbf{1}(\tilde{\tau}_P > \tilde{\tau}_M)|\tilde{\tau}_P > v\right] \\
&= \left(\frac{dc(v)}{c(\tau^*)}\right)(1-\alpha)\,\pi_L X_L e^{-\omega_M \tau^*}\left(\frac{\omega_P}{r+\omega_P}\right)\left(\frac{\omega_M}{r+\omega_P+\omega_M}\right)e^{-(r+\omega_P)(\tau^*-v)} \qquad (22)
\end{aligned}
$$

Note that if project completion date and bankruptcy date occur after date-τ^*, then the economy reverts back to the framework of the previous section where the manager will issue equity to service the coupon so long as she does not observe a bad signal. The implication is that the present value of the fourth, fifth and sixth claims are equal to the product of the values obtained in the previous section evaluated at τ^* and the date-v present value of \$1 conditional upon project completion and default occurring after date-τ^*:

$$
\begin{aligned}
V(v < \tilde{\tau}_P) &= \mathrm{E}^Q\left[e^{-r(\tau^* - v)}\,\mathbf{1}(\tilde{\tau}_P > \tau^*)\left[\mathbf{1}(\tilde{X} = X_H) + \mathbf{1}(\tilde{X} = X_L)\mathbf{1}(\tilde{\tau}_M > \tau^*)\right]|\tilde{\tau}_P > v\right] \\
&= e^{-(r+\omega_P)(\tau^*-v)}\left[\pi_H + \pi_L e^{-\omega_M \tau^*}\right]. \qquad (23)
\end{aligned}
$$

With the present value of all claims determined, we can now identify the process for the cumulative coupon payment $\{c(t)\}$. Note that the value of all six claims are proportional to $dc(t)$. Summing these six terms and setting them equal to $(1-\theta)c(t)\,dt$ identifies the deterministic process $\left(\frac{dc(t)}{dt}\right)$. To solve, we guess a value for τ^* and work backwards through time to date-0. If the implied $c(0)$ differs from the specified value, we make a new guess for τ^* until self-consistency is obtained. For the parameters specified, we find $\tau^* = 3.3$. A plot of cumulative coupon as a function of time is given in Figure (2), and credit spread as a function of time is given in Figure (3). The qualitative aspects of these figures are intuitive. By construction, cumulative coupon must be an increasing function of time, since new bondholders at date-v must be promised cash flows (which have present values that are proportional to $dc(v)$). Moreover, although not very obvious from the figure, the curve is convex upward, capturing the intuition that bonds issued at earlier dates have longer horizons for which the debt is guaranteed not to default. Moreover, as the firm raises the level of outstanding debt, it is intuitive that spreads widen, in this case, from 145 bp at $(t = 0)$ to 250 bp just before debt capacity is used up.

Hence, our framework captures the notion that default due to bad private information will not occur until debt capacity has been used up and spreads have reached speculative grade levels.

3.1 Empirical Default Intensities

Here we discuss empirical evidence consistent with the model prediction that investment grade companies do not jump to default. First, we use credit ratings to document default rates for investment- and speculative-grade firms as a function of the time horizon. Second, we conduct a similar analysis based on ratings implied by CDS contracts.

3.1.1 Evidence based on credit ratings

We collect the entire history of credit ratings given by the three main U.S. rating agencies (Moody's Investor Services, Standard & Poor's Ratings Services, and Fitch Ratings) from the Mergent database. Mergent provides ratings specific to particular bond issues, rather than an overall company rating. Hence, for a given issuer, each month we collect all ratings awarded on that month to any of its outstanding bonds and use that information to classify the company.[5] We divide individual bond ratings in three categories: investment grade (IG), higher-quality speculative grade (B), and lower-quality speculative grade (C and lower), where the last two categories, B and C, together comprise the universe of speculative grade ratings. We then assign the company to one of the three categories when the majority of the company's bond ratings are in that category. When no fresh ratings are given by any agency to the outstanding bonds of that company, we classify the firm based on ratings collected in the previous month. If no new ratings were issued the previous month, we go further back, up to 12 months. In case no new ratings are available in the entire 12 month period, we do not classify the firm. This approach mitigates the problem of classifying companies based on stale ratings.

[5] We exclude ratings on government agencies' bonds (e.g., U.S. Treasury, U.S. and foreign agencies, municipalities).

While the Mergent database contains ratings going back to the early part of the 20$^{\text{th}}$ century, ratings are limited to a very small number of debt issues through the mid 1980s. Hence, here we focus on the sample period from 1985 to 2012. Figure 4 shows the percentage of firms in each of the three rating categories: investment grade (the IG category), higher-quality speculative grade (the B category), and lower-quality speculative grade (the C category). The figure reflects the changing nature of debt markets and the evolution of the rating agencies' services. Prior to the 1980s, rating agencies were mostly focusing on blue chip industrial firms. This is consistent with a preponderance of IG ratings in the early part of the sample period. Over time, financial disintermediation and capital markets development allowed a broader variety of firms to raise funds in the bond market. Along the way, rating agencies expanded their coverage of lower-quality issues. These changes are reflected in an increased proportion of speculative-grade firms. Higher-quality speculative issues display an increasing trend through the 1990s. The proportion of lower-quality ratings remains mostly stable over the sample period, but increases during recessions; for instance, the percentage of firms in the C category peaks in 1991, 2001, and 2009.

To document default rates among rated firms, we obtain the entire history of bankruptcy filings starting from 1985 (also available through the Mergent database). Each month, after we classify firms in the three rating categories, we identify those that filed for bankruptcy over the next 12 months. We record the number of months that have elapsed between the month of the classification and the bankruptcy date. Then we count bankruptcies that occurred within the first, second, and third month of the classification (0-1, 1-2, and 2-3 months), the second, third, and fourth quarters (3-6, 6-9, and 10-12 months), and the entire year (0-12 months). For ease of comparison across periods of different length, we annualize all count variables.[6]

Table 3, Panel A, shows the average annualized default counts over the 1985-2012

[6]We multiply the count variables for the 0-1, 1-2, and 2-3 periods by 12, and those for the quarterly periods by 4.

sample period, while Panel B reports the average annualized default rates over the same period. Heteroskedasticity- and autocorrelation-robust (Newey-West) standard errors are in parentheses. It is evident that defaults by IG firms are extremely rare. Panel A shows that on average only 1.41 IG firms file for bankruptcy within a year out of 1,282 firms that are classified investment grade on average each year; the corresponding average one-year default rate is 0.11% (Panel B). These findings are consistent with Culp, Nozawa and Veronesi (2015).

A closer look at default rates at horizons from one to twelve months provides empirical support for our model. In particular, the annualized default rate is only 0.06% over the first month. This rate increases steadily at longer horizons and peaks at 0.16% in the fourth quarter since the month of the IG classification. This pattern supports the model's prediction that firms that maintain IG-status are able to avoid bankruptcy.

We observe a similar pattern for higher-quality speculative-grade firms. The average default rate for B firms is 0.78% at the one-year horizon (Panel B). At the one-month horizon, such rate drops to 0.22% and it increases more than fivefold to 1.11% in the fourth quarter since the month of the B classification. These results again suggest that firms that are perceived as relatively safe are subject to minimal jump-to-default risk.

In contrast, lower-quality speculative-grade firms (the C and lower rating category) exhibit the opposite pattern. The default rate exceeds 15.25% at the short one-month horizon, and progressively declines to 7.68% at the four-quarter horizon. This evidence is consistent with our model, which predicts that if a firm can survive beyond the point when it is most likely to default (τ^*), then spreads will narrow in the future. Panels C and D in Table 3 confirm that these findings are robust to the choice of a shorter sample period, from 2001 to 2012.

3.1.2 Evidence based on CDS data

In the previous section, we have classified firms based on credit ratings that are up to 12 months old. Such ratings might not fully reflect the information available to market participants at the time of the classification. Hence, here we propose an alternative classification of firms into the same three rating categories that is based on CDS data.

CDS contracts provide insurance in case of credit events that affect the value of a reference entity (such as the bond issued by a company that files for bankruptcy). Hence, CDS premia reflect market participants' assessment of default risk for the company that issues the reference bond. The CDS market is generally liquid. Thus, CDS contract are a useful source of real-time information about a company's credit worthiness.

To translate CDS premia into a proxy for a company's credit rating, we compare the cost of insuring bonds issued by that company with that of insuring portfolios of investment-grade and high-yield bonds (the CDX-IG and CDX-HY indices constructed by Markit Financial Information Services). Each month from 2001 to 2012 we aggregate daily five-year CDS premia from the Markit database into an average monthly CDS premium.[7] Similarly, we compute monthly averages of daily five-year CDX-IG and CDX-HY premia. If the CDS premia on a firm's bonds do not exceed the CDX-IG index by more than 100 basis points (bps), then we classify that firm as investment grade (the IG category). We use the 100 bps threshold to avoid excluding creditworthy companies whose CDS premia lie slightly above the CDX-IG level, i.e., the average IG premium. In unreported results, we find the analysis to be robust to the choice of the threshold value. In contrast, when CDS premia on a firm's bonds exceed the CDX-HY premium we classify that firm as lower-quality speculative grade (the C category). Finally, if CDS premia lie in between the IG and C thresholds, then we classify the company as higher-quality speculative grade (the B category).

Figure 5 shows the proportion of firms in each category based on CDS-implied ratings. The trading of CDS contracts on IG companies is predominant throughout the sample period, especially in the early 2000s when the CDS market was in its infancy and trading concentrated in high-quality big names. Over time, the proportion of CDS contracts on

[7] Prior to analysis, we exclude CDS contracts written on bonds issued by Government and sovereign entities.

IG firms fluctuates around a downward trend, with drops in 2001-2002 and 2008-2009 at the depth of two recessions, and peaks during the subsequent recovery periods.

The proportion of CDS contracts on higher-quality speculative-grade firms generally increases throughout the sample. Further, CDS trading in the B category exhibits peaks during recessions and declines in the expansions that follow, a pattern that is the direct opposite of the fluctuations in the proportion of IG CDS contracts. This is consistent with both (1) a reshuffling in CDS trading across rating categories over the business cycle and (2) an increase in CDS premia for IG firms during recessions combined with a decline in CDS premia of B firms during expansions that shift firms from one rating category to the other.

Finally, Figure 5 shows that little high-yield trading takes place in the early years of the CDS market. That changes over time, with a proportion of CDS contracts on low-grade speculative firms steadily increasing over the sample period.

Similar to Section 3.1.1, each month we identify firms that filed for bankruptcy within the following 12 months. Table 4, Panel A, shows the average annualized counts of bankruptcies that occurred within the first, second, and third month of the classification (0-1, 1-2, and 2-3 months), the second, third, and fourth quarters (3-6, 6-9, and 10-12 months), and the entire year (0-12 months). Panel B reports the average annualized default rates over the same 2001-2012 sample period. Heteroskedasticity- and autocorrelation-robust (Newey-West) standard errors are in parentheses.

The evidence based on CDS data provides an even more compelling case that investment-grade firms do not jump to default. At horizons from one to three months, IG default rates are virtually zero at 0.01%, and the point estimates are statistically insignificant. Default rates climb sixfold over the next three quarters, but remain extremely low at 0.06%. Default rates for firms in the B category share a similar pattern, except for being higher. From month one to three, the average annualized default rate ranges from 0.44% to 0.48%, with t-ratios that are close to three. Average default ratios increase to 0.62% by the end of the year. In contrast, firms in the C category are most likely to default in the first month, with an average default rate of 3.87%. This estimate declines steadily to 3.27% in the fourth quarter since the firm's classification.

Taken together, these results show that defaults by investment grade firms do not come as a total surprise to market participants. Even for higher-quality speculative grade firms there is limited support for jumps to default. It is mostly the lowest-rated firms that file for bankruptcy unexpectedly to market participants, as predicted by our model.

To further validate these conclusions, we hone in on the very few firms that filed for bankruptcy *and* had IG status for at least one month during the year prior to their bankruptcy. We only find 12 such firms over the entire 2001-2012 period, which shows that such instances are extremely rare. Figure 6 shows the difference between the average CDS premium on those firms and the CDX-IG index. Such spread is very small 12 months prior to bankruptcy and then increases quickly in the ensuing months. This evidence suggests that an investment policy that (1) holds bonds issued by firms in the IG category and (2) unwinds these positions when the firm falls out of IG status face virtually zero default risk.

4 The Tax-Component of Corporate Spreads

In this section, we investigate whether a significant component of corporate yield spreads can be attributed to the fact that corporate bonds are subject to state and local taxes, whereas Treasuries are not. Elton et al (2001) point out that, for example, an 8% coupon combined with a 5% state tax rate would imply a 40 bps spread due to the differential state tax treatment, which would imply that taxes explain a large fraction of investment-grade spreads, especially at shorter maturities. Central to this calculation is the identification of the appropriate state tax rate for the "marginal" agent in the market. For example, Sarig and Warga (1989) and Delianides and Geske (2001) suggest that the dominant investors in the corporate bond market may be tax-exempt, such as pensions, 401(k) and retirement accounts. Clearly, identifying the appropriate tax rate of the marginal investor is an open empirical question.

Given the potential significance of this channel, both Longstaff, Mithal and Neis (LMN, 2005) and Chen, Lesmond and Wei (CLW, 2007) attempt to quantify the tax component of corporate spreads. The approach of LMN is to first identify the default component of spreads from CDS premia, which in turn allows them to identify the non-default component. This non-default component is then regressed on coupon (and proxies for liquidity) to estimate the tax-component. Somewhat similarly, CLW regress spreads on "distance-to-default" factors such as leverage and volatility that are meant to capture default risk, on coupon rate and on liquidity factors. Both papers report a low estimate for the regression coefficient on the coupon variable, and conclude that the component of spreads due to differential taxes is small.

Here we argue, however, that even if there is a sizeable component of spreads due to taxes, we would anticipate finding a low regression coefficient attributable to coupon level (as found by LMN and CLW) due to the tax code. This is because, controlling for default risk, as we raise the coupon level from zero to par to even higher levels, the bond price goes from being at a discount to being at a premium, and this capital gain or loss is taxable. Moreover, this capital gain or loss is amortized over the life of the bond. When this tax feature is accounted for, we find that the yield regressed on coupon *should* be associated with a low regression coefficient even if there is a sizeable tax component in credit spreads. In that sense, the findings of LMN and CLW are in fact consistent with a significant tax component in credit spreads.

Indeed, as we demonstrate below, while theoretically there should be a positive relation between coupon and yield (holding credit risk and maturity constant), this relation is actually due to credit risk and not due to taxes![8] This is because in practice

[8]Consistent with this prediction, note that LMN report in their Figure 5 that the non-default component of spreads increases with default risk (proxied by credit rating).

recovery rates are typically based on face value of the bond and not on its coupon level. Hence, upon default, a high-coupon bond loses a larger fraction of its value compared to a low-coupon bond. As such, rational pricing of the corporate bond will imply a lower price/higher yield for bonds with higher coupons.

To demonstrate these effects, we consider a simplified version of LMN. In particular, we assume that the risk-free rate r, the risk-neutral default intensity λ^Q, and the liquidity process γ are all constant. Consider a risky bond with face value of \$1 which matures at date-T and that annually pays out a pre-tax coupon of C. The date-0 price of this risky bond is specified as $P_0(C,T)$. Define θ as the tax rate. Then, accounting for amortization, the promised after-tax cash flows are:

$$\begin{aligned} t \in (1, T-1): &\qquad C(1-\theta) - a\left[1 - P_0(T,C)\right]\left(\tfrac{\theta}{T}\right) \\ T \qquad : &\qquad 1 + C(1-\theta) - a\left[1 - P_0(T,C)\right]\left(\tfrac{\theta}{T}\right). \end{aligned} \tag{24}$$

Here, the parameter a is a dummy variable which equals 1 if the agent is subject to taxation, and zero otherwise. This will allow us to demonstrate the impact of accounting for the amortization of bonds purchased not at par.

Define $\tilde{\tau}$ as the (random) default time of the corporate bond. The price of the corporate bond is a sum of three terms: i) claim to amortization-adjusted, after-tax coupon prior to default over life of bond, ii) claim to face value at maturity conditional upon no default, iii) claim to recovery subject to default. Because we are considering the possibility that the agent is subject to taxation, for consistency we assume that losses due to default are tax-deductible. As such, the bond price satisfies:

$$\begin{aligned} P_0(C,T) &= \sum_{s=1}^{T} \mathrm{E}_0^Q \left[e^{-s(r+\gamma)} \left[C(1-\theta) - a\left[1 - P_0(T,C)\right]\left(\frac{\theta}{T}\right)\right] \mathbf{1}_{(\tilde{\tau}>s)} \right] + \mathrm{E}_0^Q \left[e^{-T(r+\gamma)}\, \mathbf{1}_{(\tilde{\tau}>s)} \right] \\ &\quad + \left[\left[(1-w) + a\left(P_0(C,T) - (1-w)\right)\theta\right] \int_0^T \mathbf{d1}_{(s>\tilde{\tau})}\, e^{-s(r+\gamma)} \right] \\ &= \sum_{s=1}^{T} \left[e^{-s(r+\lambda^Q+\gamma)} \left[C(1-\theta) - a\left[1 - P_0(T,C)\right]\left(\frac{\theta}{T}\right)\right]\right] + \left[e^{-T(r+\lambda^Q+\gamma)} \right] \end{aligned}$$

$$+\left[\left[(1-w)+a\theta\left(P_0(C,T)-(1-w)\right)\right]\int_0^T ds\,\lambda^Q\,e^{-s(r+\lambda^Q+\gamma)}\right]. \quad (25)$$

Again, we have used the dummy variable a in order for us to turn on an off this term. Also, we have defined w as the loss given default.

It is convenient to define the one-year discount rate $D \equiv e^{-(r+\lambda^Q+\gamma)}$. Simplifying, we find the bond price satisfies:

$$P_0(C,T) \;=\; \left(\frac{\left\{\left[C(1-\theta)-a\frac{\theta}{T}\right]D\left(\frac{1-D^T}{1-D}\right)+D^T+(1-w)(1-a\theta)\left(\frac{\lambda^Q}{r+\lambda^Q+\gamma}\right)\left(1-D^T\right)\right\}}{1-\left(\frac{a\theta}{T}\right)D\left(\frac{1-D^T}{1-D}\right)-\frac{a\theta\lambda^Q(1-D^T)}{r+\lambda^Q+\gamma}}\right). \quad (26)$$

Once the price has been determined, we use the standard industry convention to convert the price $P_0(c,T)$ into a yield y:

$$P_0(C,T) \;=\; \sum_{s=1}^{T}\frac{C}{(1+y)^s}+\frac{1}{(1+y)^T}. \quad (27)$$

To examine the sensitivity of yield on coupon, we set the after-tax risk free rate to $r = 0.04(1-0.4) = 0.024$ (and hence, the federal tax rate $\theta_F = 0.4$), the illiquidity parameter $\gamma = 0.002$, and the loss given default parameter $w = 0.6$, and then consider three cases: i) $\lambda^Q = 0.015$, $\theta = 0.05$; ii) $\lambda^Q = 0.0$, $\theta = 0.05$; iii) $\lambda^Q = 0.015$, $\theta = 0$. Note that, even though cases i) and ii) have the same tax rate and differ only through default intensities, their slope coefficients differ dramatically, and indeed are zero for case ii). This example suggests that the regression of yields on coupon are not estimating tax rates but rather default intensity. To confirm this interpretation, note that cases i) and iii) have the same default intensities but different tax rates. Indeed, the tax rate is zero for case iii). Yet, their slopes are nearly identical. In summary then, we see that if bonds are priced using after tax cash flows, then the regression coefficient of yields on coupon are capturing default risk, not the level of taxes.

4.1 Estimating the Tax Component of Short-Maturity Spreads

Let's assume that investors demand after-tax compensation in the amount of "spread" on short-maturity corporate debt to cover default and liquidity risks. If we denote the federal tax rate of the marginal bond investor by θ_F and his/her state tax-rate by θ_S, we obtain

$$y_{corp}(1-\theta_F-\theta_S) - r(1-\theta_F) \;=\; spread\,, \tag{28}$$

where y_{corp} is the yield on the corporate bond and r is the risk free Treasury yield. This equation can be rewritten in terms of the (pre-tax) credit spread:

$$y_{corp} - r \;=\; \left(\frac{r\theta_S}{1-\theta_F-\theta_S}\right) + \left(\frac{spread}{1-\theta_F-\theta_S}\right). \tag{29}$$

Equation (29) shows that, controlling for credit and liquidity risk, the pre-tax spread $(y_{corp} - r)$ depends on the the nominal interest rate r only if there is a state tax impact, i.e., $\theta_S > 0$ for the marginal bond investor. To test whether this dependence is supported by the data, we consider two measures of pre-tax short-maturity credit spreads. We obtain monthly three-month nonfinancial commercial paper (CP) and three-month secondary-market certificates of deposit (CDs) rates from the H.15 data release at the Federal Reserve Board. These series share desirable features that make them ideal for our analysis. Both instruments have short maturity and their payments are subject to state and federal personal income taxation when held by individuals outside of tax-shielded accounts. CP is typically issued by large creditworthy companies, hence credit risk is small, and it is very liquid. The FDIC guarantee makes CDs even safer than CP; since we focus on CDs traded in the secondary market, liquidity risk is limited too. We start the sample in April 1971, when the CP series first becomes available, and end it in June 2013 when the 3-month CD data series ends.[9] In both cases, we subtract the three-month U.S. Treasury bill (T-Bill) rate to construct the spread series.

Figure 7 shows the CP and CD spreads series and compares them to the level of the T-Bill rate. It is evident that the two spread measures are very close; indeed, their

[9]CD rates are an average of dealer bid rates for certificates of deposit that are actively traded in the secondary market and are issued by top-tier banks. Responses are not reported when the number of respondents is too few to be representative. Source: Haver Analytics Inc.

cross-correlation is 95%.[10] Moreover, they exhibit a great deal of co-movement with the T-Bill rate: in both cases, the correlation is around 52%, which is consistent with the presence of a state-tax channel in CP and CD spreads, as predicted by equation (29).

To quantify the effect of taxes on credit spreads, we estimate equation (29) via non-linear least squares (NLS). To this end, we need a measure for the federal tax rate, θ_F. There is considerable evidence that participation in financial markets is more common among high-income households. Hence, for each year we approximate θ_F with the marginal tax rate for households in the highest income bracket (source: www.taxfoundation.org). In contrast, we treat the state tax rate θ_S as a constant parameter to be estimated. In unreported robustness checks, we find that using a lower marginal θ_F makes the state tax channel even stronger, i.e., the estimate for θ_S remains positive and significant at all conventional levels.

Next, we identify a comprehensive set of variables that help explain credit and liquidity risk in debt markets (e.g., CLW, LMN, Collin-Dufresne and Goldstein (2001), and Krishnamurthy and Vissing-Jorgensen (20012)). The full list and the data sources are in Table 8. A preliminary data exploration shows that there is a great deal of correlation among these series. Moreover, some of them also correlate with the risk-free rate r. To remove collinearity among the explanatory variables in the regression model, we first project each liquidity/credit risk proxy onto r and, second, we extract mean-zero principal components from the panel of the projection residuals. We find that the first two principal components explain most of the variation in these variables and thus we use them to approximate the after-tax spread in equation (29), $spread = \beta_0 + \beta_1\, PC_1 + \beta_2\, PC_2$, where β_0, β_1, and β_2 are coefficients to be estimated.

We estimate the model in logarithms and report results for the full 1971-2013 sample period in Table 9, Panel A; heteroskedasticity and autocorrelation-robust (Newey-West)

[10] While CP might seem riskier than CDs, Ou, Hamilton, and Cantor (2004) note that in most cases holders of defaulted CP have ultimately been made whole. In particular, none of the defaults by issuers rated Prime-1 by Moody's at the time of issuance have resulted in any appreciable losses to investors. This could explain why the series of CP and CD rates are alike.

standard errors are in parentheses. First, in columns (1) and (3) we assume that the compensation for liquidity/credit risk is constant, i.e., $\beta_1 = \beta_2 = 0$. For both CP and CD spreads, θ_S is significant and precisely estimated with t-ratios of 8.54 and 6.13, respectively, with point estimates that correspond to a 2.8% marginal state tax rate. The β_0 coefficient is also significant and its standard errors are small; the point estimates correspond to an average compensation for liquidity/credit risk of 6 and 11 bps for CP and CD, respectively.

In columns (2) and (4) we extend the model to include the first two PCs constructed from our proxies for liquidity and credit risk. The estimates for the marginal state tax rate θ_S decline slightly to 1.90% and 2.05% and remain significant at all conventional levels (t-ratios are 6.19 and 5.41, respectively). In comparison, the liquidity/credit risk component increases to 11 and 17 bps, respectively (recall that the PCs have zero mean, hence β_0 captures the mean non-tax component even when the PCs are included). Panel B shows evidence for the post-1985 sample period. The results are similar to those for the full sample, except that the θ_S estimates increase to 3.9% and 3.6% for CP and CD, respectively. This is consistent with an increase in the number of wealthy investors entering the CP and CD market as lower T-Bill rates in the post-1985 period reduce the tax advantage of Treasuries. Finally, in unreported work we find the results to be robust to the choice of other sub-samples, e.g., periods that exclude the financial crisis and the Great Recession.

We rely on the NLS estimation results of Table 9 to decompose the pre-tax credit spread in terms of three components, each representing the portions due to (1) credit/liquidity risk, (2) federal taxes, and (3) state taxes. Due to the non-linear nature of equation (29), there is no unique way to perform such decomposition. Yet, a meaningful approach is to break down the total spread as

$$y_{corp} - r = \left[spread\right] + \left[\left(\frac{spread}{1-\theta_F}\right) - spread\right]$$

$$+\left[\left(\frac{r\theta_S}{1-\theta_F-\theta_S}\right)+\left(\frac{spread}{1-\theta_F-\theta_S}\right)-\left(\frac{spread}{1-\theta_F}\right)\right]. \tag{30}$$

Clearly, if the federal tax $\theta_F = 0$, then the second component (i.e., the federal tax component) vanishes. Similarly, if the state tax $\theta_S = 0$, then the third component (i.e., the state tax component) is zero.

Figure 8 shows the decomposition for CP (top panel) and CD spreads (bottom panel) based on the estimates in columns (2) and (4) of Table 9, Panel A. It is evident that the liquidity/credit risk portion of the spread is small and generally stable over time. It peaks in recessions (in particular, at the height of the U.S. financial crisis) and declines during periods of expansion.

The federal tax component is sizable throughout the sample period and fluctuates in tune with changes in the U.S. tax code. It is larger in the early part of the sample period, where the top federal taxes rate is as high as 70%, and declines following the Tax Reform Act of 1986. We then see an increase in the federal tax component during the years of the Clinton Administration, and a subsequent decline in connection with the Jobs and Growth Tax Relief Reconciliation Act of 2003.

Of the three, it is the state tax component that exhibits the most variation. It is very large in the mid 1970s and early 1980s, exactly at the time when T-Bill rates soared. In particular, it peaks during the Fed's monetarist policy experiment under Chairman Paul Volcker when the three-month T-Bill rate reached 17%. This is consistent with equation (29), which shows that the state-tax impact is proportional to the level of the risk-free rate. During the remaining part of the sample period the T-Bill rate has declined considerably, and so has the state-tax component in credit spreads. In particular, since the Fed policy intervention has pushed short-term rates to the zero lower bound during the Great Recession, the state-tax component in credit spreads has vanished.

Over the entire sample period, the liquidity/credit risk component is 31.49% of the total fitted CP spread; in contrast, we find that about two thirds of the total spread are due to taxes, with federal and state tax shares of 22.68% and 45.83%, respectively.

The results for CD spreads are similar, except for a slightly lower 40.31% state-tax component, and higher credit/liquidity and federal tax shares (34.08% and 25.61%, respectively). For both spread measures, the high state tax share is driven by the high level of the T-Bill rate in the 1970s and 1980s, which increases the state tax channel. As the T-Bill rate has declined considerably since the early 1990s, we recompute spread shares starting from 1993.[11] Consistent with the model predictions, we find that the average state tax share drops to 28.79% and 28.79% for CP and CD spreads, respectively. This decline is compensated, for the most part, by an increase in the liquidity/credit risk share and, to a lesser extent, an increase in the federal tax share.

5 Conclusion

We provide theoretical arguments and empirical support for notion that a large fraction of short-maturity investment-grade spreads are due to taxes but not to credit event premia. First, we investigate an economy qualitatively similar to that of Duffie and Lando (2001) and show that, if the manager is restricted to issuing only equity to service existing debt, then it will be in the best interest of shareholders for the manager to default if her private signal of firm value is sufficiently poor. However, if the manager has the ability to issue both debt and equity to service existing debt, then the manager of an IG firm receiving a bad private signal will maximize shareholder value by concealing this information and servicing existing debt via debt issuance. Consistent with empirical observation, this strategy permits low-spread firms to avoid jumping to default, at least until their debt capacity has been used up and the firm is has dropped down to "fallen angel" status with speculative-grade spreads. Since firms with IG-level spreads do not face credit-event risk, their bond yields do not command a credit-event premium. We provide empirical support for this prediction by showing that, even for those few firms which defaulted within six months of having spreads commensurate with the IG index,

[11]We use 1993 as it is the year during which Congress enacted the Omnibus Budget Reconciliation Act which, among other deficit reduction measures, increased the top individual tax rate to 39.6%.

the majority of them saw their spreads increase significantly over the months, implying that an investment strategy which holds only those bonds with IG level spreads faces virtually zero credit-event risk.

We then argue that much of the spreads of short maturity IG debt are due to taxes. We specify a simple model that links credit spreads to the level of the risk-free rate, the federal and state marginal tax rates, and the after-tax compensation for liquidity/credit risk. Across sample periods, the point estimates for the marginal state tax rate θ_S are positive and significant at any conventional level. The state-tax component of credit spreads is particularly big in the 1970s and 1980s, and decreases in later years along with the risk-free rate. Federal taxes are an important component of spreads too, especially in the early part of the sample where top individual rates are high. Overall, federal and state taxes account for two thirds of credit spreads, while liquidty/credit risk explain only a third.

A Appendix

A.1 Proof of Equation (13)

The claim to face value $\left(\frac{c(0)}{r+s}\right)$ conditional upon $\tilde{\tau}_P < \tilde{\tau}_B$ is

$$B_{FV}(\mathbf{1}_{(\tilde{\tau}_P>t)}, \mathbf{1}_{(\tilde{\tau}_B>t)}) = \left(\frac{c(0)}{r+s}\right) \mathrm{E}\left[e^{-r(\tilde{\tau}_P-t)}\mathbf{1}_{(\tilde{\tau}_P<\tilde{\tau}_B)} \middle| \tilde{\tau}_P > t,\, \tilde{\tau}_B > t\right] \mathbf{1}_{(\tilde{\tau}_P>t)}\mathbf{1}_{(\tilde{\tau}_B>t)} \quad (31)$$

It is useful to note that

$$\mathbf{1}_{(\tilde{\tau}_P<\tilde{\tau}_B)} = \mathbf{1}_{(\tilde{X}=X_H)} + \mathbf{1}_{(\tilde{X}=X_L,\, \tilde{\tau}_P<\tilde{\tau}_M)} \quad (32)$$

Hence, we have two terms. The first term is

$$\begin{aligned} B_{FV,1} &= \left(\frac{c(0)}{r+s}\right) \mathrm{E}\left[e^{-r(\tilde{\tau}_P-t)}\mathbf{1}_{(\tilde{X}=X_H)} \middle| \tilde{\tau}_P > t,\, \tilde{\tau}_B > t\right] \mathbf{1}_{(\tilde{\tau}_P>t)}\mathbf{1}_{(\tilde{\tau}_B>t)} \quad (33)\\ &= \left(\frac{c(0)}{r+s}\right) \int_0^\infty d\tau_M \int_0^\infty d\tau_P\, e^{-r(\tau_P-t)}\, \pi\left(\tilde{\tau}_P = \tau_P,\, \tilde{\tau}_M = \tau_M,\, \tilde{X} = X_H \mid \tilde{\tau}_P > t,\, \tilde{\tau}_B > t\right). \end{aligned}$$

Note that

$$\pi\left(\tilde{\tau}_P=\tau_P,\,\tilde{\tau}_M=\tau_M,\,\tilde{X}=X_H\;\mid\;\tilde{\tau}_P>t,\,\tilde{\tau}_B>t\right) \tag{34}$$
$$=\frac{\pi\left(\tilde{\tau}_P>t,\,\tilde{\tau}_B>t\;\mid\;\tilde{\tau}_P=\tau_P,\,\tilde{\tau}_M=\tau_M,\,\tilde{X}=X_H\right)\pi\left(\tilde{\tau}_P=\tau_P,\,\tilde{\tau}_M=\tau_M,\,\tilde{X}=X_H\right)}{\pi\left(\tilde{\tau}_P>t,\,\tilde{\tau}_B>t\right)}$$
$$=\frac{\mathbf{1}_{(\tau_P>t)}\pi_H\,\omega_M e^{-\omega_M\tau_M}\,\omega_P e^{-\omega_P\tau_P}}{e^{-\omega_P t}\left[e^{-\omega_m t}+\pi_H\left(1-e^{-\omega_M t}\right)\right]}.$$

Hence, we find

$$B_{FV,1}\;=\;\left(\frac{c(0)}{r+s}\right)\left(\frac{\omega_P}{r+\omega_P}\right)\left(\frac{\pi_H}{\pi_H+\pi_L e^{-\omega_M t}}\right). \tag{35}$$

Similarly, the second term is

$$B_{FV,2}\;=\;\left(\frac{c(0)}{r+s}\right)\mathrm{E}\left[e^{-r(\tilde{\tau}_P-t)}\mathbf{1}_{(\tilde{X}=X_L,\,\tilde{\tau}_P<\tilde{\tau}_M)}\middle|\,\tilde{\tau}_P>t,\,\tilde{\tau}_B>t\right]\mathbf{1}_{(\tilde{\tau}_P>t)}\mathbf{1}_{(\tilde{\tau}_B>t)} \tag{36}$$
$$=\;\left(\frac{c(0)}{r+s}\right)\int_0^\infty d\tau_M\int_0^\infty d\tau_P\,e^{-r(\tau_P-t)}\,\mathbf{1}_{(\tau_P<\tau_M)}$$
$$\times\pi\left(\tilde{\tau}_P=\tau_P,\,\tilde{\tau}_M=\tau_M,\,\tilde{X}=X_L\;\mid\;\tilde{\tau}_P>t,\,\tilde{\tau}_B>t\right).$$

Note that

$$\pi\left(\tilde{\tau}_P=\tau_P,\,\tilde{\tau}_M=\tau_M,\,\tilde{X}=X_L\;\mid\;\tilde{\tau}_P>t,\,\tilde{\tau}_B>t\right) \tag{37}$$
$$=\frac{\pi\left(\tilde{\tau}_P>t,\,\tilde{\tau}_B>t\;\mid\;\tilde{\tau}_P=\tau_P,\,\tilde{\tau}_M=\tau_M,\,\tilde{X}=X_L\right)\pi\left(\tilde{\tau}_P=\tau_P,\,\tilde{\tau}_M=\tau_M,\,\tilde{X}=X_L\right)}{\pi\left(\tilde{\tau}_P>t,\,\tilde{\tau}_B>t\right)}$$
$$=\frac{\mathbf{1}_{(\tau_P>t)}\mathbf{1}_{(\tau_M>t)}\pi_L\,\omega_M e^{-\omega_M\tau_M}\,\omega_P e^{-\omega_P\tau_P}}{e^{-\omega_P t}\left[e^{-\omega_m t}+\pi_H\left(1-e^{-\omega_M t}\right)\right]}.$$

Hence, we find

$$B_{FV,2}\;=\;\left(\frac{c(0)}{r+s}\right)\left(\frac{\omega_P}{r+\omega_P+\omega_M}\right)\left(\frac{\pi_L e^{-\omega_M t}}{\pi_H+\pi_L e^{-\omega_M t}}\right). \tag{38}$$

Combining, we find

$$B_{FV}(\mathbf{1}_{(\tilde{\tau}_P>t)},\mathbf{1}_{(\tilde{\tau}_B>t)})\;=\;\left(\frac{c(0)}{r+s}\right)\left(\frac{1}{\pi_H+\pi_L e^{-\omega_M t}}\right) \tag{39}$$
$$\times\left[\left(\frac{\pi_H\,\omega_P}{r+\omega_P}\right)+\left(\frac{\omega_P\pi_L e^{-\omega_M t}}{r+\omega_P+\omega_M}\right)\right]\mathbf{1}_{(\tilde{\tau}_P>t)},\mathbf{1}_{(\tilde{\tau}_B>t)}.$$

A.2 Proof of Equation (14)

Investors realize that default occurs only if the manager has received a bad signal $\tilde{X} = X_L$ at a time τ_M that occurs before project completion τ_P. As such, for the case $\tau_M > t$, $\tau_P > t$, the claim to recovery from an outsider's perspective satisfies

$$\begin{aligned}
B_{recov}(\tilde{\tau}_P > t, \tilde{\tau}_M > t) &= \mathrm{E}_t^Q \left[(1-\alpha) X_L e^{-r(\tau_P - t)} \mathbf{1}_{(\tilde{X}=X_L)} \mathbf{1}_{(\tau_M < \tau_P)} | \tilde{\tau}_P > t, \tilde{\tau}_M > t \right] \\
&= \int_0^\infty d\tau_M \int_{\tau_M}^\infty d\tau_P \; (1-\alpha) X_L e^{-r(\tau_P - t)} \; \pi \left(\tilde{X} = X_L, \tilde{\tau}_P = \tau_P, \tilde{\tau}_M = \tau_M | \tilde{\tau}_P > t, \tilde{\tau}_M > t, \right) \\
&= \left(\frac{(1-\alpha) X_L \pi_L}{\pi_H + \pi_L e^{-\omega_M t}} \right) \int_t^\infty d\tau_M \omega_M e^{-\omega_M \tau_M} \int_{\tau_M}^\infty d\tau_P \; \omega_P \; e^{-(r+\omega_P)(\tau_P - t)} \\
&= \left(\frac{(1-\alpha) X_L \pi_L e^{-\omega_M t}}{\pi_H + \pi_L e^{-\omega_M t}} \right) \left(\frac{\omega_P}{r + \omega_P} \right) \left(\frac{\omega_M}{r + \omega_P + \omega_M} \right). \qquad (40)
\end{aligned}$$

A.3 Proof of Equation (19)

$$\begin{aligned}
B_{recov,1}(v < \tau_P) &= \left(\frac{dc(v)}{c(\tau^*)} \right) \mathrm{E}^Q \left[(1-\alpha) \tilde{X} e^{-r(\tilde{\tau}_P - v)} \mathbf{1}(\tilde{X} = X_L) \mathbf{1}(\tilde{\tau}_M < \tau^*) \mathbf{1}(\tilde{\tau}_P > \tau^*) | \tilde{\tau}_P > v \right] \\
&= \left(\frac{dc(v)}{c(\tau^*)} \right) (1-\alpha) X_L \pi_L \left(\int_0^{\tau^*} d\tau_M \, \omega_M \, e^{-\omega_M \tau_M} \right) \left(\int_{\tau^*}^\infty d\tau_P \, e^{-(r+\omega_P)(\tilde{\tau}_P - v)} \, \omega_P \right) \\
&= \left(\frac{dc(v)}{c(\tau^*)} \right) (1-\alpha) \pi_L X_L \left(\frac{\omega_P}{\omega_P + r} \right) \left[1 - e^{-\omega_M \tau^*} \right] e^{-(r+\omega_P)(\tau^* - v)}. \qquad (41)
\end{aligned}$$

A.4 Proof of Equation (20)

$$\begin{aligned}
&B_{coup,2}(v < \tau_P) \\
&\quad = \mathrm{E}^Q \left[\int_{\tau^*}^\infty dt \, dc(v) \, e^{-r(t-v)} \mathbf{1}(\tilde{\tau}_P > t) \left[\mathbf{1}(\tilde{X} = X_H) + \mathbf{1}(\tilde{X} = X_L) \mathbf{1}(\tilde{\tau}_P < \tilde{\tau}_M) \right] | \tilde{\tau}_P > v \right] \\
&\quad\quad + \mathrm{E}^Q \left[\int_{\tau^*}^\infty dt \, dc(v) \, e^{-r(t-v)} \mathbf{1}(\tilde{\tau}_M > t) \mathbf{1}(\tilde{X} = X_L) \mathbf{1}(\tilde{\tau}_P > \tilde{\tau}_M) | \tilde{\tau}_P > v \right] \\
&\quad = dc(v) \pi_H \int_{\tau^*}^\infty dt \, e^{-r(t-v)} \int_t^\infty d\tau_P \, \omega_P \, e^{-\omega_P(\tau_P - v)}
\end{aligned}$$

$$
\begin{aligned}
&+dc(v)\pi_L \int_{\tau^*}^{\infty} dt\, e^{-r(t-v)} \int_t^{\infty} d\tau_P\, \omega_P\, e^{-\omega_P(\tau_P-v)} \int_{\tau_P}^{\infty} d\tau_M\, \omega_M\, e^{-\omega_M \tau_M} \\
&+dc(v)\pi_L \int_{\tau^*}^{\infty} dt\, e^{-r(t-v)} \int_t^{\infty} d\tau_M\, \omega_M\, e^{-\omega_M \tau_M} \int_{\tau_M}^{\infty} d\tau_P\, \omega_P\, e^{-\omega_P(\tau_P-v)} \\
&= dc(v)\, e^{-(\omega_P+r)(\tau^*-v)} \left[\frac{\pi_H}{r+\omega_P} + \frac{\pi_L e^{-\omega_M \tau^*}}{r+\omega_P+\omega_M}\right]. \qquad (42)
\end{aligned}
$$

A.5 Proof of Equation (21)

$$
\begin{aligned}
&B_{FV,2}(v<\tau_P) \\
&= \mathrm{E}^Q \left[e^{-r(\tau_P-v)} \left(\frac{dc(v)}{r+s}\right) \mathbf{1}(\tilde{\tau}_P > \tau^*) \left[\mathbf{1}(\tilde{X}=X_H) + \mathbf{1}(\tilde{X}=X_L)\mathbf{1}(\tilde{\tau}_P < \tilde{\tau}_M)\right] |\tilde{\tau}_P > v \right] \\
&= \left(\frac{dc(v)}{r+s}\right) e^{-(r+\omega_P)(\tau^*-v)} \left[\pi_H \left(\frac{\omega_P}{r+\omega_P}\right) + \pi_L\, e^{-\omega_M \tau^*} \left(\frac{\omega_P}{r+\omega_P+\omega_M}\right)\right]. \qquad (43)
\end{aligned}
$$

A.6 Proof of Equation (22)

$$
\begin{aligned}
B_{recov,2}(v<\tilde{\tau}_P) &= \mathrm{E}^Q \left[e^{-r(\tau_P-v)} \left(\frac{dc(v)}{c(\tau^*)}\right) (1-\alpha)\, \tilde{X} \mathbf{1}(\tilde{X}=X_L) \mathbf{1}(\tilde{\tau}_M > \tau^*) \mathbf{1}(\tilde{\tau}_P > \tilde{\tau}_M) | \tilde{\tau}_P > v \right] \\
&= \left(\frac{dc(v)}{c(\tau^*)}\right) (1-\alpha)\, X_L \pi_L \int_{\tau^*}^{\infty} d\tau_M\, \omega_M\, e^{-\omega_M \tau_M} \int_{\tau_M}^{\infty} d\tau_P\, \omega_P\, e^{-(r+\omega_P)(\tau_P-v)} \\
&= \left(\frac{dc(v)}{c(\tau^*)}\right) (1-\alpha)\, \pi_L X_L e^{-\omega_M \tau^*} \left(\frac{\omega_P}{r+\omega_P}\right) \left(\frac{\omega_M}{r+\omega_P+\omega_M}\right) e^{-(r+\omega_P)(\tau^*-v)} \qquad (44)
\end{aligned}
$$

References

[1] J. Bai, P. Collin-Dufresne, R. Goldstein, and J. Helwege. On bounding credit-event risk premia. forthcoming, Review of Financial Studies, 2015.

[2] A. Berndt, R. Douglas, D. Duffie, M. Ferguson, and D. Schranz. Measuring default risk premia from default swap rates and EDFs. Working Paper, Stanford University, 2009.

[3] H. Bhamra, L.-A. Kuehn, and I. A. Strebulaev. The aggregate dynamics of capital structure and macroeconomic risk. *Review of Financial Studies*, 23 (12):4187–4241, 2010.

[4] H. Chen. Macroeconomic conditions and the puzzles of credit spreads and capital structure. *Journal of Finance*, 65:2171–2212, 2010.

[5] L. Chen, P. Collin-Dufresne, and R. Goldstein. On the relation between the credit spread puzzle and the equity premium puzzle. *Review of Financial Studies*, 22:3367–3409, 2009.

[6] L. Chen, D. Lesmond, and J. Wei. Corporate yield spreads and bond liquidity. *Journal of Finance*, 62:119–149, 2007.

[7] P. Collin-Dufresne, R. S. Goldstein, and J. S. Martin. The determinants of credit spread changes. *Journal of Finance*, 56:2177–2208, 2001.

[8] C. L. Culp, Y. Nozawa, and P. Veronesi. Option-based credit spreads. Working Paper, University of Chicago and Federal Reserve Board of Governors, 2015.

[9] S. Davydenko. When do firms default? a study of the default boundary. Working paper, University of Toronto, 2012.

[10] G. Delianedis and R. Geske. The components of corporate credit spreads: Default, recovery, tax, jumps, liquidity, and market factors. Working paper, UCLA, 2001.

[11] J. Driessen. Is default event risk priced in corporate bonds? *Review of Financial Studies*, 18:165–195, 2005.

[12] D. Duffie and D. Lando. Term structure of credit spreads with incomplete accounting standards. *Econometrica*, 69:633–664, 2001.

[13] D. Duffie and K. Singleton. An econometric model of the term structure of interest rate swap yields. *Journal of Finance*, 52(4):1287–1322, 1997.

[14] E. J. Elton, M. J. Gruber, D. Agrawal, and C. Mann. Explaining the rate spread on corporate bonds. *Journal of Finance*, 56:247–277, 2001.

[15] P. Feldhtter and S. Schaefer. The credit spread puzzle - myth or reality? Working paper, LBS, 2015.

[16] P. Feldhütter and S. Schaefer. The credit spread puzzle – Myth or reality? Working Paper, Lonon Business School, 2014.

[17] R. C. Green. A simple model of the taxable and tax-exempt yield curves. *Review of Financial Studies*, 6:233–264, 1993.

[18] R. Jarrow, D. Lando, and S. Turnbull. A Markov model for the term structure of credit spreads. *Review of Financial Studies*, 10:481–523, 1997.

[19] A. Krishnamurthy and A. Vissing-Jorgensen. The aggregate demand for Treasury debt. *Journal of Political Economy*, 120:233–267, 2012.

[20] H. E. Leland. Optimal capital structure, endogenous bankruptcy, and the term structure of credit spreads. *Journal of Finance*, 49:1213–1252, 1994.

[21] F. Longstaff, S. Mithal, and E. Neis. Corporate yield spreads: Default risk or liquidity? new evidence from the credit default swap market. *Journal of Finance*, 60:2213–2253, 2005.

[22] R. C. Merton. On the pricing of corporate debt: The risk structure of interest rates. *Journal of Finance*, 29:449–470, 1974.

[23] S. Ou, D. Hamilton, and R. Cantor. Short-term rating performance and corporate commercial paper defaults, 1972-2004. *Moody's Investors Service, Inc.*, page 118, 2004.

[24] L. Saita. The puzzling price of corporate default risk. Working paper, Stanford University, 2006.

[25] O. Sarig and A. Warga. Some empirical estimates of the risk structure of interest rates. *Journal of Finance*, 44:1351–1360, 1989.

[26] A. Shleifer and R. Vishny. The limits of arbitrage. *Journal of Finance*, 52:35–55, 1997.

[27] S. D. I. Strebulaev and X. Zhao. A market-based study of the cost of default. *Review of Financial Studies*, 25:3573–3609, 2012.

B Tables and Figures

Table 1: **Model Calibration Parameters**

X_H	X_L	π_H	π_L	ω_P	ω_M	$c(0)$	r	s	θ	α
140	40	0.6	0.4	0.1	0.2	1.88	0.04	0.0145	0.2	0.2

Table 2: **Output from Model**

E_H	E_L	E_{NS}	B_{coup}	B_{FV}	B_{recov}	B_{tot}	$(E_{NS} + B_{tot})$	Unleveraged Value
64.62	0.00	37.65	10.27	18.85	5.38	34.49	72.14	71.43

Table 3: **Average Defaults by Credit Rating.** For months from January 1985 to December 2012, we classify firms based on credit ratings issued within the previous 12 months by the three main rating agencies (Moody's, Standard and Poor's, and Fitch). Panel A shows the average number of firms in each rating category and, among these, the average number of firms that have defaulted in the next 12 months. Panel B reports average annualized default rates. Panels C and D show similar summary statistics for the 2001-2012 period. Heteroskedasticity- and autocorrelation-robust (Newey-West) standard errors are in parentheses.

Rating	Average Num. of Firms per Year	Annualized Defaults						
		0-1M	1-2M	2-3M	3-6M	6-9M	9-12M	0-12M
Panel A: Average annualized defaults, 1985-2012								
IG	1282	0.79	0.96	1.00	1.13	1.55	2.05	1.41
	(118)	(0.35)	(0.42)	(0.40)	(0.37)	(0.47)	(0.60)	(0.39)
B	1164	2.61	4.32	6.21	8.11	10.76	12.94	9.05
	(117)	(0.73)	(1.13)	(1.48)	(1.92)	(2.42)	(3.03)	(2.03)
C	186	28.43	27.11	24.32	21.79	17.75	14.31	20.12
	(23)	(4.84)	(4.76)	(4.15)	(3.59)	(2.96)	(2.26)	(3.24)
Panel B: Average annualized default rates, 1985-2012								
IG		0.06	0.08	0.08	0.09	0.12	0.16	0.11
		(0.03)	(0.03)	(0.03)	(0.03)	(0.03)	(0.04)	(0.03)
B		0.22	0.37	0.53	0.70	0.92	1.11	0.78
		(0.06)	(0.08)	(0.11)	(0.14)	(0.17)	(0.21)	(0.14)
C		15.25	14.54	13.05	11.69	9.52	7.68	10.79
		(1.46)	(1.48)	(1.29)	(1.13)	(0.99)	(0.82)	(0.99)

Table 3, continued

Rating	Average Num. of Firms per Year	Annualized Defaults						
		0-1M	1-2M	2-3M	3-6M	6-9M	9-12M	0-12M
Panel C: Average annualized defaults, 2001-2012								
IG	1950	1.67	2.00	2.00	1.92	2.61	3.64	2.51
	(60)	(0.75)	(0.91)	(0.83)	(0.68)	(0.85)	(1.11)	(0.65)
B	1797	4.33	7.75	10.42	12.08	15.25	17.00	12.96
	(39)	(1.29)	(2.04)	(2.50)	(2.85)	(3.55)	(4.10)	(2.87)
C	319	44.00	41.75	36.67	32.97	25.78	20.00	29.89
	(17)	(7.21)	(7.00)	(5.89)	(5.15)	(3.86)	(2.86)	(4.26)
Panel D: Average annualized default rates, 2001-2012								
IG		0.09	0.10	0.10	0.10	0.13	0.19	0.13
		(0.04)	(0.05)	(0.04)	(0.03)	(0.04)	(0.05)	(0.03)
B		0.24	0.43	0.58	0.67	0.85	0.95	0.72
		(0.07)	(0.11)	(0.14)	(0.16)	(0.20)	(0.22)	(0.16)
C		13.80	13.09	11.50	10.34	8.08	6.27	9.37
		(1.70)	(1.68)	(1.42)	(1.27)	(1.00)	(0.81)	(1.02)

Table 4: **Average Defaults by Credit Rating: Evidence based on CDS Data**. For months from January 2001 to October 2012, we classify as investment grade (IG category) those firms that have CDS contracts trading at a premium no higher than 100 basis points of the CDX-IG index. We classify as higher-quality speculative grade (B category) those firms that have CDS contracts trading at a premium in excess of 100 basis points of the CDX-IG index and lower than the premium on the CDX-HY index. Lower-quality speculative-grade firms (C category) have CDS premia in excess of the CDX-HY index. The table shows the average number of investment and speculative grade firms, among these, the average number of firms that have defaulted in the next 12 months. Panel B reports average default rates. Heteroskedasticity- and autocorrelation-robust (Newey-West) standard errors are in parentheses. Source: Markit database.

Year	Average Num. of Firms per Year	Annualized Defaults						
		0-1M	1-2M	2-3M	3-6M	6-9M	9-12M	0-12M
Panel A: Average annualized defaults, 2001-2012								
IG	1447	0.17	0.17	0.17	0.65	0.65	0.87	0.58
	(94)	(0.13)	(0.13)	(0.13)	(0.28)	(0.26)	(0.39)	(0.20)
B	404	1.77	1.94	1.86	2.11	2.65	2.51	2.28
	(48)	(0.68)	(0.66)	(0.72)	(0.77)	(0.93)	(0.90)	(0.70)
C	140	5.41	5.32	5.07	5.07	4.79	4.56	4.92
	(17)	(1.42)	(1.62)	(1.38)	(1.33)	(1.28)	(1.14)	(1.03)
Panel B: Average annualized default rates, 2001-2012								
IG		0.01	0.01	0.01	0.04	0.04	0.06	0.04
		(0.01)	(0.01)	(0.01)	(0.02)	(0.02)	(0.03)	(0.01)
B		0.44	0.48	0.46	0.52	0.66	0.62	0.57
		(0.15)	(0.15)	(0.16)	(0.18)	(0.22)	(0.21)	(0.16)
C		3.87	3.81	3.63	3.63	3.43	3.27	3.52
		(1.00)	(1.14)	(0.97)	(0.92)	(0.85)	(0.74)	(0.68)

Table 5: For the tax rate $\theta = 0.05$ and risk-neutral intensity $\lambda^Q = 0.015$, we determine the yield to maturity for risky bonds as a function of coupon (column) and maturity (row) using equations (26)-(27). The last column estimates the regression coefficient via $slope = \frac{y(10\%)-y(0\%)}{10}$. Additional parameter values are $r = 0.04$, $\gamma = 0.01$, $w = 0.5$.

	0%	2%	4%	6%	8%	10%	slope
1	6.203	6.220	6.236	6.252	6.267	6.281	0.008
2	6.177	6.203	6.228	6.251	6.273	6.294	0.012
3	6.150	6.186	6.219	6.250	6.279	6.305	0.016
4	6.122	6.169	6.211	6.249	6.284	6.316	0.019
5	6.094	6.151	6.203	6.248	6.289	6.326	0.023

Table 6: For the tax rate $\theta = 0.05$ and risk-neutral intensity $\lambda^Q = 0.0$, we determine the yield to maturity for risky bonds as a function of coupon level (column) and maturity (row) using equations (26)-(27). The last column estimates the regression coefficient via $slope = \frac{y(10\%)-y(0\%)}{10}$, which is nearly zero for all maturities. Taken together with the results from Table (5), this result emphasizes that the slope coefficient between coupon and yield in Table (5) is due to default risk, not taxes. Other parameter values are $r = 0.04$, $\gamma = 0.01$, $w = 0.5$.

	0%	2%	4%	6%	8%	10%	slope
1	5.397	5.397	5.397	5.397	5.397	5.397	0.000
2	5.397	5.397	5.397	5.397	5.397	5.397	0.000
3	5.397	5.397	5.397	5.397	5.397	5.397	0.000
4	5.398	5.398	5.397	5.397	5.397	5.396	0.000
5	5.398	5.398	5.397	5.397	5.396	5.396	0.000

Table 7: For the tax rate $\theta = 0.0$ and risk-neutral intensity $\lambda^Q = 0.015$, we determine the yield to maturity for risky bonds as a function of coupon level (column) and maturity (row) using equations (26)-(27). The last column estimates the regression coefficient via $slope = \frac{y(10\%)-y(0\%)}{10}$. Note that these coefficients are nearly identical to those in Table (5) in spite of the tax rate being zero. This emphasizes that the slope coefficient between coupon and yield in Table (5) is not due to taxes. Other parameter values are $r = 0.04$, $\gamma = 0.01$, $w = 0.5$.

	0%	2%	4%	6%	8%	10%	slope
1	5.895	5.911	5.927	5.941	5.956	5.969	0.007
2	5.871	5.896	5.920	5.942	5.963	5.982	0.011
3	5.847	5.881	5.912	5.942	5.969	5.995	0.015
4	5.821	5.865	5.905	5.942	5.975	6.006	0.019
5	5.795	5.850	5.899	5.942	5.981	6.017	0.022

Table 8: **Proxies for Credit and Liquidity Risk.** We aggregate daily variables to the monthly frequency and interpolate quarterly variables to construct monthly series.

Series	Freq.	Description
On-off-the-run spread	D	The difference between (1) the 10-year U.S. Treasury par yield computed by the staff at the Federal Reserve Board (FRB) from off-the-run Treasuries and (2) the 10-year U.S. Treasury par yield computed by the U.S. Treasury from on-the-run Treasuries and released by the FRB
TED spread	D	The difference between the (1) the rate on 3-month Eurodollar deposits and (2) the 3-month U.S. Treasury Bill rate. Both series are from the FRB.
log(D / A)	Q	Aggregate leverage, computed as the ratio between (1) Nonfinancial Corporate Business: Liabilities: Credit Market Debt and (2) Nonfinancial Corporate Business: Total Assets from the B103 table in Flow of Funds data release by the FRB
log(D / (D+E))	Q	Aggregate leverage, computed as the ratio between (1) Nonfinancial Corporate Business: Liabilities: Credit Market Debt and (2) the sum of (1) and Nonfinancial Corporate Business; Corporate Equities; Liability from the B103 table in Flow of Funds data release by the FRB
r_{Market}	M	The 12-month CRSP value-weight return inclusive of distributions
log(Vol(r_{Market}))	M	The realized volatility on the monthly CRSP value-weight return
BAA-AAA spread	M	The difference between the yields on Moody's BAA and AAA bond indices
UE	M	The seas. adjusted civilian unemployment rate from the Bureau of Labor Statistics
g_r	Q	The seas. adjusted 12-month growth rate on real gross domestic product from the Bureau of Economic Analysis
log(D / GDP)	Q	The ratio of Treasury debt outstanding to U.S. gross domestic product from Henning Bohn's web page at https://econ.ucsb.edu/ bohn

Table 9: **The Components of CP and CD Spreads**. The table shows nonlinear least squares regression results for the model

$$\log\left(y_{corp} - r\right) \;=\; \log\left(\frac{r\,\theta_S}{(1-\theta_F-\theta_S)} + \frac{\beta_0 + \sum_{i=1}^{2}\beta_i\,PC_i}{(1-\theta_F-\theta_S)}\right) + \varepsilon\,,$$

where $spread = \beta_0 + \sum_{i=1}^{2}\beta_i\,PC_i$ is the after-tax spread demanded by market participant in compensation for liquidity and credit risk. PC_1 and PC_2 are mean-zero principal components computed from a panel of variables that are proxies for credit and liquidity risk in debt markets. The full list of variables is in Table 8. In columns (1)-(2), the dependent variable is the spread between the three-month nonfinancial commercial paper and T-Bill rates, while in columns (3)-(4) it is the spread between the three-month secondary-market certificates of deposit and T-Bill rates. We estimate the model on monthly data from April 1971 to June 2013 (Panel A) and from January 1985 to June 2013 (Panel B). Heteroskedasticity- and autocorrelation-robust (Newey-West) standard errors are in parentheses.

	CP Spread		CD Spread	
	(1)	(2)	(3)	(4)
	Panel A: The sample period is from April 1971 to June 2013			
θ_s	0.0280 (0.0033)	0.0190 (0.0031)	0.0277 (0.0045)	0.0205 (0.0038)
β_0	0.0006 (0.0001)	0.0011 (0.0002)	0.0011 (0.0003)	0.0017 (0.0003)
β_1		-0.0001 (0.0001)		0.0001 (0.0001)
β_2		0.0003 (0.0001)		0.0005 (0.0002)
R^2	46.89%	55.22%	39.07%	55.02%
	Panel B: The sample period is from January 1985 to June 2013			
θ_s	0.0393 (0.0037)	0.0438 (0.0038)	0.0362 (0.0057)	0.0369 (0.0062)
β_0	0.0005 (0.0001)	0.0004 (0.0001)	0.0011 (0.0003)	0.0012 (0.0003)
β_1		0.0001 (0.0001)		0.0003 (0.0002)
β_2		0.0001 (0.0001)		0.00002 (0.0001)
R^2	53.91%	59.11%	29.86%	42.99%

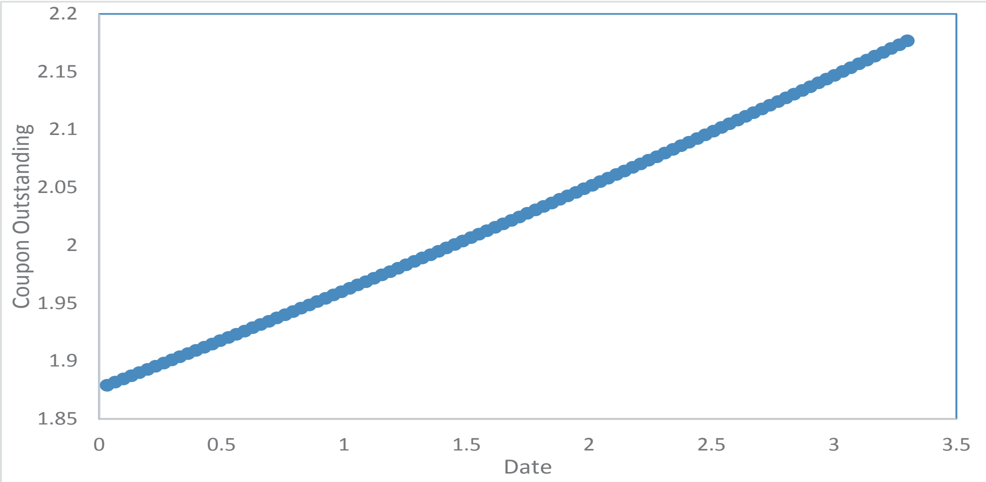

Figure 2: **Cumulative Coupon as a Function of Time.** This figure displays the cumulative coupon value as a function of time. Model parameters are as in Table 1.

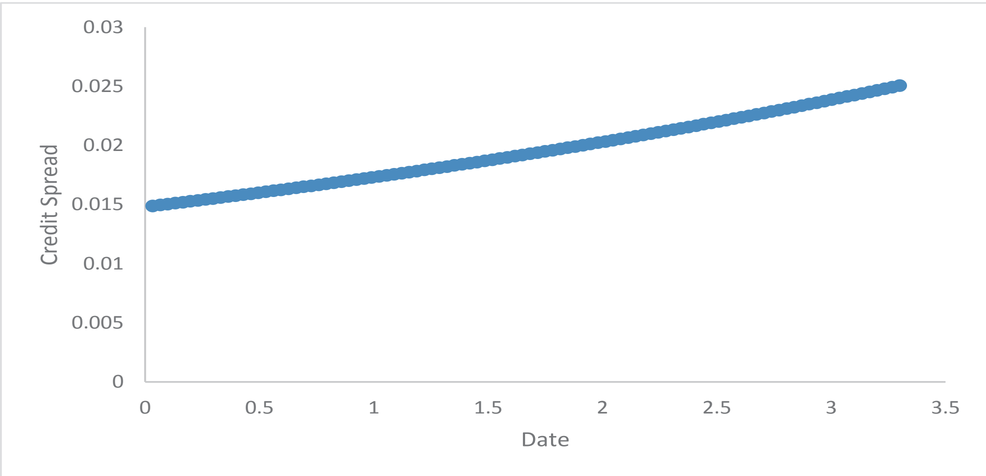

Figure 3: **Credit Spread as a Function of Time.** This figure shows how credit spreads increase over time due to the increasing level of debt. Model parameters are as in Table 1.

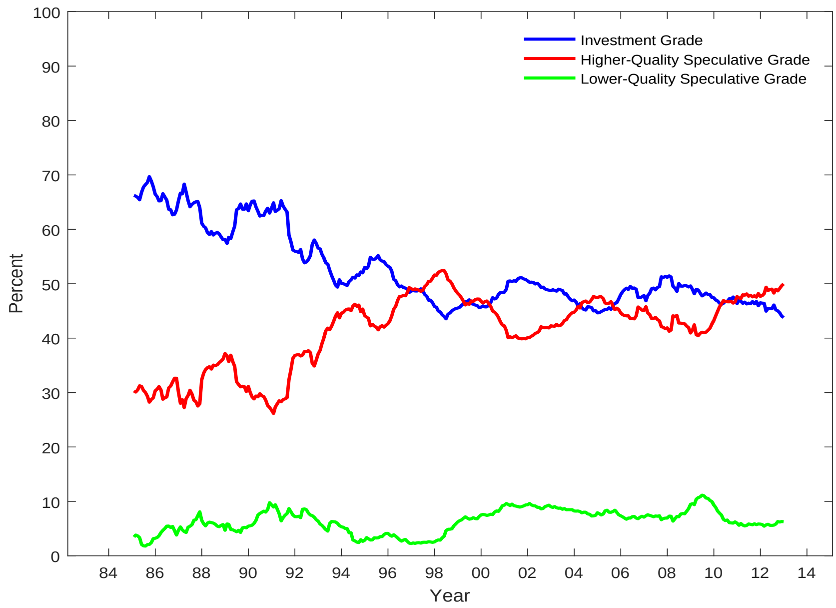

Figure 4: **Percentage of Firms by Credit Ratings.** The plots show the percentage of firms in each of the three rating categories: investment grade (the IG category), higher-quality speculative grade (the B category), and lower-quality speculative grade (the C category). The sample period goes from 1985 to 2012. Source: Mergent database.

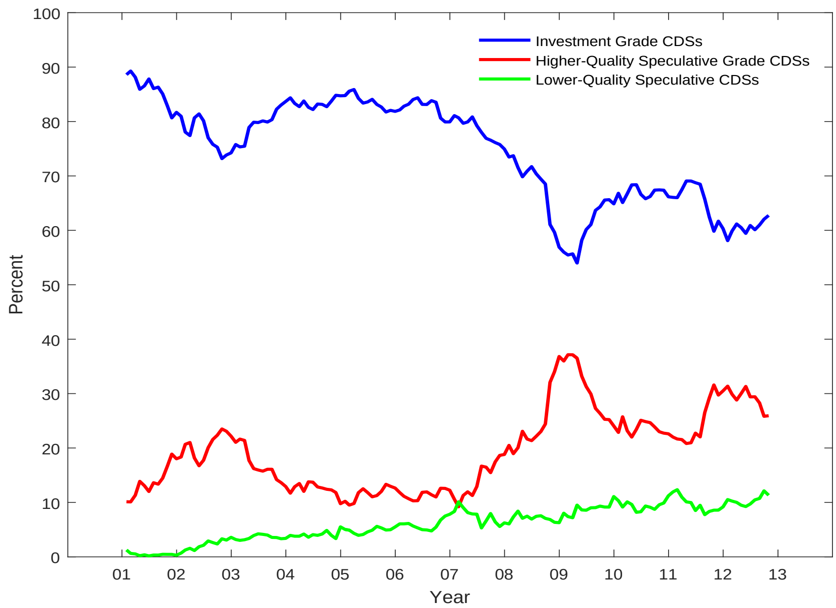

Figure 5: **Percentage of Firms by CDS-Implied Ratings.** The plots show the percentage of firms in each of the three rating categories (IG, B, and C) based on ratings implied by CDS data. The sample period goes from 2001 to 2012. Source: Markit databases.

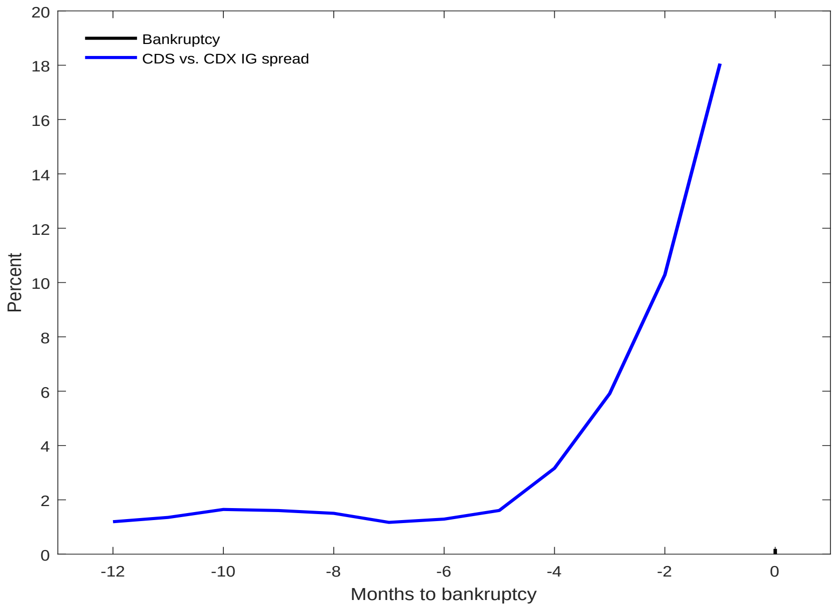

Figure 6: **Average CDS Premium on Investment Grade Firms up to Bankruptcy.** Among the firms that went bankrupt from 2001 to 2014, we classify as investment grade those that had CDS contracts trading at a premium no higher than 100 basis points of the CDX Investment Grade Index for at least one of the 12 months preceding the bankruptcy date. The plot shows the average CDS premium, in excess of the CDX Investment Grade Index, on those investment grade firms in the 12 months leading up to their bankruptcy.

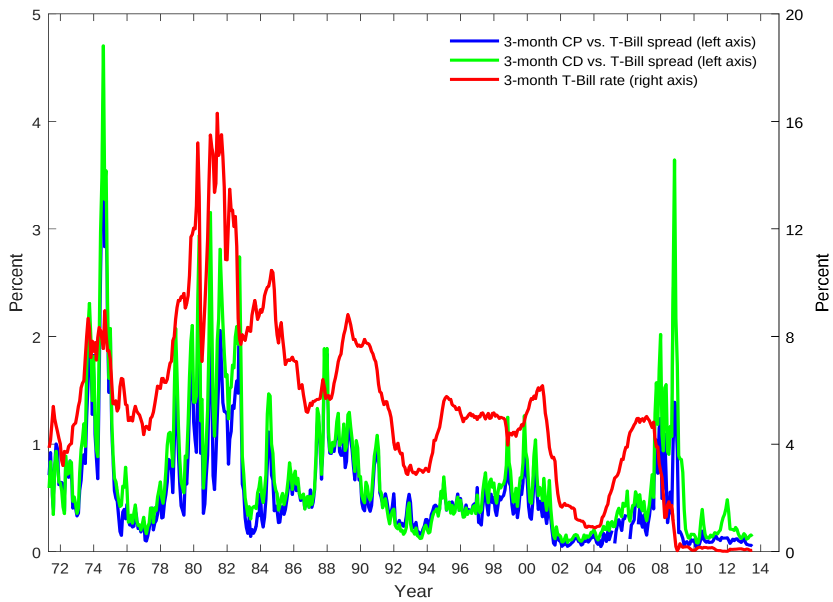

Figure 7: **Commercial Paper, Certificates of Deposit, and Treasury Bill Rates.** The blue and green lines show the spread of three-month commercial paper and three-month certificates of deposit rates vs. the three-month Treasury bill rate, respectively. The red line shows the 3-month Treasury bill rate. Sample period: from April 1971 to June 2013.

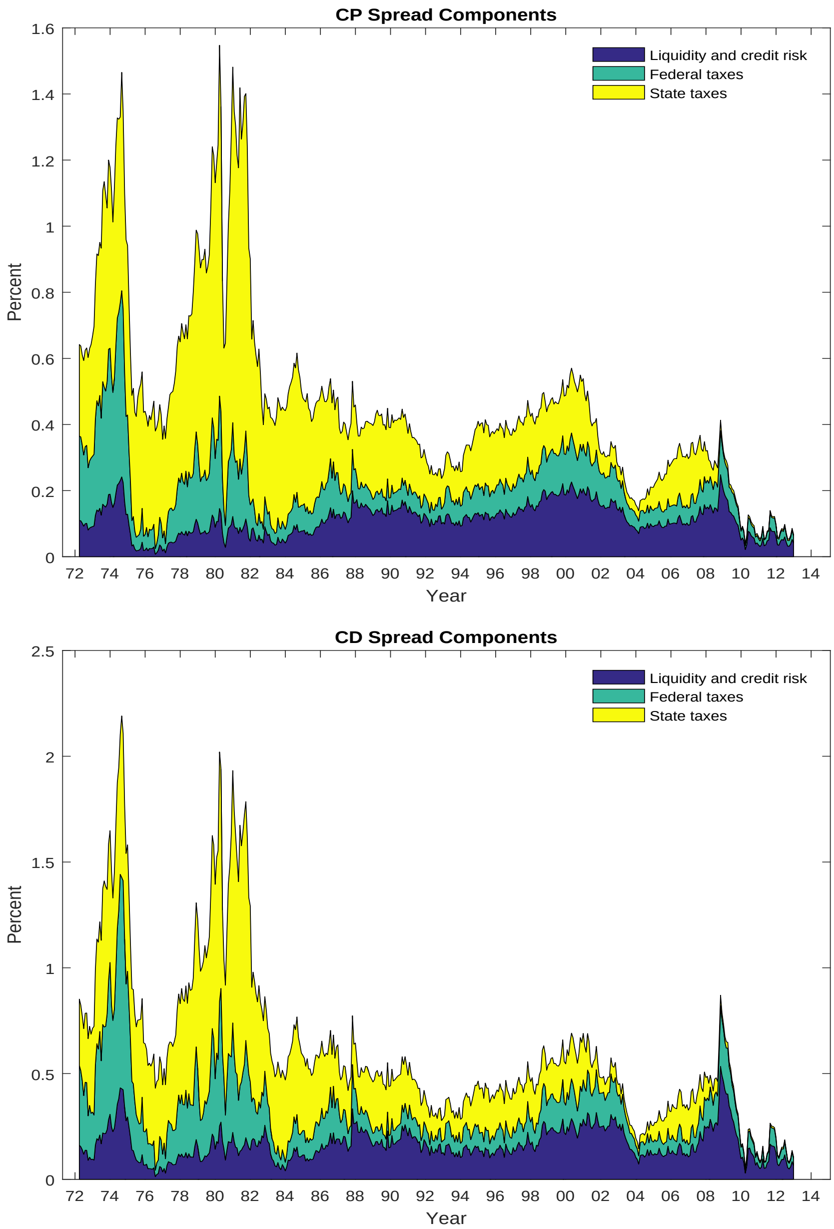

Figure 8: **The Components of CP and CD Spreads.** The area plots show the components of credit spreads due to liquidity/credit risk (blue), federal taxes (green) and state taxes (yellow). The top panel shows the decomposition of the CP spread, while the bottom panel depicts results for the CD spread. Sample period: from April 1971 to June 2013.