A Pure Test for the Elasticity of Yield Spreads

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Abstract

The correlation between interest rates and corporate bond yield spreads is a well-known feature of structural bond pricing models. Duffee (1998) argues that this correlation is weak once the effects of call options are removed from the data; a conclusion that contradicts the negative correlation expected by Longstaff and Schwartz (1995). However, Elton et al. (2001) point out that Duffee's analysis ignores the effects of the tax differential between U.S. Treasury and corporate bonds. Canadian bonds have no such tax differential, yet, after controlling for callability, we find that the correlation between interest rates and corporate bond spreads remains negligible. We also find a significant negative relationship for callable bonds with this relationship increasing with the moneyness of the call provision. These results are robust under alternate empirical specifications.

A Pure Test for the Elasticity of Yield Spreads

1. Introduction

Structural models[1] for the valuation of corporate bonds assume an asset-based default process with default occurring once the stochastic value of the firm's assets hits a default threshold. In these models the threshold is conveniently expressed as the relation between the market value of the firm's assets to its debt. One of the more notable predictions of structural models is that credit spreads –the bond yield spread between a risky bond and a near maturity riskless bond- are negatively correlated to the return on risky assets and changes in default-free interest rates, with the former generally proxied by the return on a well-diversified market index, while the later is proxied by the change in a near maturity Government bond rate.

The objective of this paper is to once again revisit the theme of the pricing of corporate credit spreads. There are two important grounds for doing so. First, despite extensive empirical examination the expected negative correlations predicted by the structural models are not necessarily present in risky bonds of all credit classes, maturities and markets (e.g. Longstaff and Schwartz, 1995; Duffee, 1998; and Collin-Dufresne et al., 2001). Second, more recent theoretical papers have questioned the underlying assumptions of the simple structural models which lead to these conclusions (e.g. Duffee, 1998; Elton et al. 2001; Campbell and Taksler, 2003; Ericsson and Renault, forthcoming).

The view that credit spreads are negatively related to asset returns is consistent with the stylized facts evident in the pricing of risky bonds by market participants: credit spreads are higher for bonds of declining credit quality; they increase when news negatively affects underlying corporate asset values and when there is a lack of liquidity. The other conclusion - of a negative relation to the riskless rate – is, however, not so apparent and has remained the subject of extensive empirical verification. *Ceteris paribus*, under risk-neutral valuation, an increase in the riskless rate implies a higher expected future value for the firm's assets relative to the default threshold, and a lower risk-neutral probability of default. This ultimately results in a lower credit spread. While Longstaff and Schwartz (1995), who pioneered this line of testing using Moody's indexed U.S. bond yields, find a strong negative relationship between yield spreads and Treasury yields, Duffee (1998) argues that since most bonds in Moody's index are callable bonds, the negative yield spread – riskless rate relation could instead be due to the negative relationship between the yield spread attributed to the call option and the riskless interest rate. Consideration of the effect embedded options in debt has on credit spreads should also be considered with other recent findings: Campbell and Taksler (2003) show that idiosyncratic firm-level volatility affects credit spreads, leading to the conclusion that firm-level volatility should also affect any embedded options in debt value, while Hackbarth et. al. (forthcoming) also argues that credit spreads should be positively correlated with the volatility of cash flows from firm assets.

While recent studies mount a strong case for the impact liquidity has on the risky yield spread (Ericsson and Renault, forthcoming; Chen et al., forthcoming), more importantly for this study, Elton et al. (2001), point out that Duffee's (1998) analysis ignores potential tax effects in his tests[2]. They study the components of yield spreads for U.S. investment-grade corporate bonds, which by definition have low default risk and so should be more sensitive to interest rate effects. In conclusion, Elton et al. warn against ignoring the tax differential when studying corporate bond yield spreads, and both Duffee (1998) and Elton et al. (2001) make a case for a pure test based on bond data with no call options and tax effects.

In this paper, we use a database of Canadian, investment-grade, corporate bond indices devoid of tax effects, since Canadian corporate and government bonds, unlike U.S. bonds, are subject to the same tax treatment. Canadian government bonds are also highly liquid and not being a reserve currency are not actively sought by international investors, notably foreign central banks and financial institutions, whose market actions have the potential to distort pricing along the term structure. This database also contains a unique provision allowing for identification of callable and noncallable indices. Yet, our results are similar to those found by Duffee (1998). The negative yield spread – riskless rate relation found in

Canadian callable bond indices is largely due to the call premium, while this relationship is negligible for noncallable bonds. This result is robust for a variety of empirical models, and still holds for an inflation-adjusted test.

The remainder of this paper is organized as follows. The next section introduces previous studies and their empirical models and describes the unique characteristics of the Canadian bond market. Section 3 outlines the data and regression methodology and presents the empirical results. In section 4, an alternative test using default rates is undertaken. Finally, the summary and conclusion are offered in section 5.

2. Previous Studies

The two factor valuation model of Longstaff and Schwartz (1995) is an extension of the closed-form solution of Merton (1974), where default is a function of the value of the firm at maturity, to a simple continuous-time valuation framework that allows for both default and interest rate risk. This structural model captures the stochastic nature of interest rates, where for simplicity the dynamics of the interest rate are explained using a simple term structure model based on Vasicek (1977). Other perspectives include the reduced-form models of Jarrow, Lando and Turnbull (1997) and Duffie and Singleton (1999) where the payoff upon default is specified exogenously, The asset level which triggers default can also be imposed endogenously by having the shareholders optimally liquidate the firm as in Leland (1994), Leland and Toft (1996) and Anderson and Sundaresan (1996). As noted by Giesecke (2006) these models do not consider the way information is revealed over time, and implicitly assume that investors can observe the inputs to the model definition of default. However, investors do not have complete information about the inputs to the model definition of default; in particular the asset value of a firm is hard to directly observe. Consequently in her model, termed a first passage default model, investors learn over time the location of the barrier, since it must lie below the observable historical low of assets to date if the firm has not already defaulted.

In the Longstaff and Schwartz (1995) model an Ordinary Least Squares (OLS) regression analysis is applied to both absolute yield spreads (the difference between the risky and riskless yields) and relative yield spreads (the ratio of the risky to riskless bond). Using Moody's corporate bond yield indices, they find a significant negative yield spread – riskless rate relationship for absolute spreads and a stronger negative relationship for relative spreads. Longstaff and Schwartz conclude that this empirical result supports their two-factor model corporate bond-pricing model.[3] However, Duffee (1998) points out the failings of this approach owing to the construction of Moody's index, which uses both callable and noncallable bonds. For callable bonds, higher interest rates imply a lower chance that the issuer will exercise the call option. Thus bondholders will demand a lower yield for this call provision, which will result in an overall decrease in the bond yield spread. To accommodate the call features present in most U.S. corporate bonds he constructs a noncallable bond index and regresses spread changes on changes both in the short yield and in a term structure slope variable, for both callable and noncallable bonds. Interestingly, he finds that the negative relationship between credit spread and interest rate is much weaker once the call option effects are removed from the data. Specifically, changes in yield spread on callable bonds are found to be strongly negatively related to changes in Treasury yields, while on the other hand, for noncallable bonds, he finds a weak, although still negative relation. Therefore the negative relationship found by Longstaff and Schwartz (1995) can be attributed to either the default premium or call premium.

In a related study using Australian Eurobond, Batten, Hogan and Jacoby (2005) find results consistent with the implications of the Longstaff and Schwartz (1995) theoretical model, with actual and relative credit spreads both negatively related to changes in the All Ordinaries Index (a proxy for the asset factor in the Longstaff and Schwartz model), and with changes in Australian Government bond yields as well (a proxy for the interest rate factor in the Longstaff and Schwartz model). Their contribution lies in providing an insight into why the relative measure tends to be statistically more significant than the alternate measure based upon the difference. They do so by introducing a simple theoretical framework that explains the effect of callability on the interest-rate factor and that by construction, relative spreads should bring about a stronger yield spread – interest rate factor relation,

Collin-Dufresne et al. (2001) identify more factors that may determine changes in credit spreads. For example, Duffee (1998) omits important factors such as the Longstaff and Schwartz (1995) asset factor in his regression analysis. Collin-Dufresne et al. (2001) combine the explanatory variables offered by Longstaff and Schwartz and Duffee, in addition to a convexity term and firm leverage variable. They conclude that credit spread changes are primarily driven by local supply/demand shocks, which are independent of both credit-risk factors and liquidity factors. Their regression model is so far the most comprehensive in the extant literature and forms the basis for a number of recent models. These accommodate time varying volatility and autocorrelation in spread changes through GARCH and ARMA specifications and more recently through the incorporation of liquidity based variables (Ericsson and Renault, forthcoming; Chen et al., forthcoming). Note that autocorrelation in spread returns may be due to illiquidity of that particular bond in secondary markets.

In this paper, we apply each of the above seminal regression models using Canadian, investment-grade, corporate bond indices. The strength of our study lies in using Canadian corporate bonds, which are devoid of the tax effect. Studies using U.S. bond data cannot avoid these tax effects. Canada's Doomsday call provision also allows identifying callable and noncallable indices. The effects of taxes and the uniqueness of the Canadian Doomsday call provision are discussed in turn.

3. The Uniqueness of Canadian Bond Data

3.1 Tax Effects

Canadian bond data enables controlling for tax effects arising from the different tax rates, which apply to U.S. corporate and Treasury bonds. Duffee (1998) notes that U.S. corporate bonds are subject to taxation at the federal, state, and local levels, while U.S. Treasury bonds are subject only to federal tax.[4] To determine the potential effect of these distortions on credit spreads, Elton et al. (2001) decompose yield spreads (calculated from estimated zero curves) for U.S. investment-grade corporate bonds into three components: default-risk premium, systematic-risk premium, and a state-tax premium. Their analysis shows that the state-tax premium is significantly more important relative to the default-risk premium (36.1 percent versus 17.8 percent of the yield spread for 10-year A-rated bonds, respectively). Since Canadian corporate and Government of Canada bonds are subject to an identical tax rate then tax effects should not play a role in the estimated relation for Canadian corporate bonds and their spreads.

Another advantage of using the Canadian bond data relates to both the level of coupons and the tax system. Constantinides and Ingersoll (1984) show that a dynamic bond trading strategy, aimed at minimizing tax liabilities, produces bond prices significantly higher than when using a buy-and-hold strategy. They attribute this difference in value to the tax-timing option. Jordan and Jordan (1991) provide strong evidence supporting the existence of a tax-timing option for U.S. Treasury bonds.

Prisman, Roberts, and Tian (1996) demonstrate that, given the different tax treatment for bonds in Canada, the tax-timing option in Canada is unlikely to have an economic value to bond traders. In addition, they show that when the range of coupon rates in the portfolio contracts, the value of the tax-timing option will be even lower. Historically, SCM (Scotia Capital Markets) imposed constraints on the range of coupon rates permitted for corporate bonds to be included in its indices.[5] These constraints were designed to eliminate the coupon level effect from the yield spread of the included bonds over the yield on Government of Canada bonds, so that this spread is as close as possible to the true yield spread. Given these constraints, and the tight range of coupon rates at the present, the value for the tax-timing option for these indices is expected to be even lower, and its impact on the estimated relation minimal.

Tax effects may also explain the divergence in empirical results for credit spreads where the risky bond is a Eurobond, or similar offshore security. For example, Wagner, Hogan and Batten (2005) adopt the Longstaff and Schwartz (1995) two-factor regression, extended for correlated spread changes and heteroskedasticity to investigate Deutschemark-denominated Eurobonds. They find that while changes in spreads are significantly negatively related to the term structure level, contrary to structural theory, the proxy for the asset value does not yield a significant negative contribution, while for long bond maturities there is a significant positive relation The authors interpret this result as being consistent with (i) portfolio rebalancing or substitution effects (where long maturity high quality bonds are traded for stocks), and (ii) a market risk premium on corporate bonds, which dominates default risk. This later conclusion is consistent with Elton et al. (2001) where a time-varying premium for bearing risk in capital markets may yield a systematic risk premium for corporate bonds with the sensitivity of credit spreads to systematic risk factors tending to increase for longer maturities and lower credit quality.

However, these authors do not consider the different tax treatment that applies to domestically issued government bonds as opposed to Eurobonds from the international investor viewpoint. In most countries Eurobonds are able to be issued by resident corporations in such a way that they avoid with-holding tax provisions otherwise payable to their international investors. These provisions normally allow for the exemption to the payment of with-holding tax when the issue is “publicly available” and “widely distributed”. The legal test for this varies for each country but normally relies upon whether the issue is listed on an exchange and the face value of each specific bond. Low face values enable the bond to be widely circulated and therefore “widely distributed”. These same exemptions are not necessarily available to foreign investors who buy domestic corporate and government bonds. Consequently the overall effect on credit spread pricing is not clear-cut since it would depend on the extent that foreign investors were involved in the domestic government bond market purchases, although the opportunities for distortions in pricing are evident.

In addition, credit spreads on corporate bonds may also be compromised by the ability for some local firms to credit enhance their international as well as domestic debt through the use of dedicated cash-flow from exports. Specifically, Durbin and Ng (2005) investigate emerging markets for “sovereign ceiling effects,” which says that a firm is no more creditworthy than its government. Contrary to expectations they find evidence which refutes the presence of a sovereign ceiling, for local firms that have substantial export earnings and/or a close relationship with either a foreign firm or government.

3.2 The Canadian Doomsday Call Provision

To control for callability, Duffee (1998) suggests stratifying data by forming two portfolios for each rating category, one consisting of only callable bonds, and the other of noncallable bonds. However, there may be a selection bias in callability related to risk differences between callable and noncallable bonds. For example, Bodie and Taggart (1978) claim that firms with more growth opportunities are more likely to issue callable bonds. If the firm's expectations of growth are confirmed, its shareholders can avoid sharing this good fortune with bondholders by simply having the firm's managers call the entire bond issue. Berk, Green, and Naik (1999) show that the systematic risk of asset returns is related to the firm's growth opportunities. This implies that stratifying bond data based on callability may create two risk classes within each rating category.

Canadian corporate bonds have a feature that allows for mitigating the potential selection bias associated with callability. Most Canadian corporate bonds issued from 1987, carry a call provision called the “doomsday” call provision. This call provision makes it possible to control for callability for some bonds, facilitating the study of a set of corporate bonds broader than just noncallable bonds. Similar to the U.S. make-whole call option, a doomsday call provision sets the call price at the maximum of the par value of the bond, or the value of the bond calculated based on the yield on a Government of Canada bond (with a matching maturity) plus a spread (the doomsday spread).[6] Thus, a make-whole call price is calculated in the same manner as that of the doomsday call, but the bond's remaining cash flows are discounted with the yield of comparable maturity treasury security plus a contractually specified "make-whole premium".

Below, we demonstrate that BBB-rated Canadian bonds are always traded with yield spreads much wider than the doomsday spread set in their call provision. Thus, the exercise of the doomsday call for BBB-rated bonds will rarely cause financial damage to the bondholders, and these bonds may be considered economically noncallable.[7]

To substantiate this feature of Canadian corporate bonds, we collect doomsday spreads for all bonds carrying the doomsday call for each month during the 01:1993-12:1999 period, from the *Financial Post* Corporate Bond Record. Since in this study we focus on SCM's long-term bond indices, we limit our sample to data corresponding to bonds with a maturity greater than ten years. We stratify our data by credit rating, and for every month, we calculate the average and standard deviation of doomsday spreads across bonds within each rating category. Since data for the AAA bonds are unavailable for most of the sample period, we obtain statistics only for AA, A, and BBB rated bonds.

To determine the moneyness of the doomsday call provision, one must compare the doomsday spread to the yield spread. In Figure 1, for each rating category, we plot the yield spread on the corresponding long-term index (S), the average doomsday spread (μ), the average doomsday spread plus one standard deviation ($\mu+\sigma$), and the average doomsday spread plus two standard deviations ($\mu+2\sigma$).[8] In Panel A of Figure 1, we see that for AA-rated bonds the yield spread is lower than the average doomsday spread in a large number of cases, and it is almost always lower than the doomsday spread plus one standard deviation. Similarly, in Panel B of Figure 1 we see that the yield spread of A-rated bonds is often lower than the average doomsday spread, and in a significant number of cases it is lower than the doomsday spread plus one standard deviation.

Based on this 7-year sample it is clear that the probability of the doomsday call being in the money for AA- and A-rated bonds is substantial, with that of AA-rated bonds being significantly larger. In Panel C of Figure 1 we see that the yield spread of BBB-rated bonds is always significantly higher than the average doomsday spread plus two standard deviations. Thus, based on our sample, we conclude that the probability of the doomsday call being in the money for a BBB-rated bond is virtually zero, and thus BBB-rated bonds are economically noncallable.

Insert Figure 1 here

Mann and Powers (2003) study U.S. bonds with make-whole call provisions. They analyze 318 bonds carrying this provision, issued between October 1995 and September 1999. Similar to our findings for Canadian doomsday bonds, Mann and Powers show that at issuance the make-whole call option for U.S. bonds is out of the money. Although this is true for all ratings, lower-rated bonds are deeper out of the money.

Note that the history of the make-whole call provision in the U.S. is significantly shorter relative to that of the doomsday call in Canada (first U.S. issue is from 1995 while first Canadian issue is from 1987). If the make-whole or doomsday call option is initially issued out of the money, the bond yield spread has to shrink over time for the option to move in the money. In other words, the bond credit rating has to ameliorate or the overall market conditions have to change significantly. Given the longer history of this option in our Canadian data, we indeed find that the doomsday option for our higher-rated Canadian bonds is sometimes in the money.

The presence of the doomsday call has important implications for studying the yield spread - riskless rate relation. For a standard call provision carried by most U.S. corporate bonds, lower riskless rates imply a higher probability of the issuer calling the bond. For the doomsday call of BBB-rated bonds, the effect of lower riskless rates is defused by the call price floating upwards. Thus, the existence of the doomsday call provides a useful instrument for the isolation of the effect of default risk and its significance.

This result allows us to control for callability for the long-term BBB-rated SCM bond index, during the later period of our 25-year sample. To determine the duration of this period, for each month during the 01:1993-12:1999 7-year period we count the number of corporate bonds issued with a doomsday call provision, the number of bonds issued with a standard call provision, and the number of noncallable bonds, as reported in the *Financial Post* Corporate Bond Record. In Figure 2 we plot the proportion of each of the above categories calculated with respect to the total number of bonds within each month. As most Canadian corporate bonds issued after 1986 either are noncallable or carry a doomsday call, in Figure 2 we see that the proportion of doomsday-call bonds increases considerably during the 7-year period, from 22.82 percent in 01:1993 to 47.48 percent in 12:1999. At the same time, the proportion of bonds carrying a standard call provision decreases dramatically, from 41.03 percent in 01:1993 to 8.25 percent in 12:1999. The proportion of noncallable bonds fluctuates between 34.46 percent and 47.29 percent.

Insert Figure 2 here

The above proportions are calculated for the entire Canadian corporate bond universe as covered by the *Financial Post* Corporate Bond Record. One can expect the proportion of long-term doomsday bonds, with over ten years to maturity, to be much higher compared to that of medium-term and short-term bonds. Since most Canadian corporate bonds are issued with 10 to 20 years to maturity, it is more likely that the number of newly issued short-term and mid-term bonds in SCM's mid-term and short-term indices is dominated by the number of seasoned bonds. Thus, for those indices, it is more likely that most bonds are those originally issued prior to 1987 with the standard call provision.

CANSIM reports that as of 01:1987, the weighted average maturity for SCM's long-term AA, A, and BBB indices, is 14.61 years, 14.80 years, and 13.34 years, respectively.

Note that these maturities represent the maturity of the last Canadian corporate bonds issued with a standard call provision prior to 1987. Thus, four years or so later, the "average'" bond, issued originally with a standard call provision, is expected to become a medium-term bond, with maturity below ten years. Since these are averages, to be safe, we feel that it is reasonable to wait eight years instead, and to assert that starting 01:1995, the vast majority of bonds included in SCM's long-term bond indices carry a doomsday call rather than a standard call provision.

Following the above discussion, we conclude that SCM's long-term bond indices are the indices suitable for our study, and thus we discard the mid-term and short-term indices. Furthermore, we conclude that the 01:1995-07:2001 period, in which these long-term indices consist mainly of bonds carrying a doomsday call, is an adequate estimation period to control for the callability of the BBB-rated index.

In summary, analyzing yield spreads using the SCM Canadian corporate bond indices has the advantage of controlling for callability and effects arising from taxation. Such an analysis provides a clearer picture of the role of the default-risk adjustment in measuring the sensitivity of investment-grade bond yield spreads to changes in the riskless rate.

4. Data and Method

4.1. Data

Our sample is based on month-end yield-to-maturity data from the Scotia Capital Markets (SCM) investment-grade Canadian corporate bond indices, reported by Statistics Canada (CANSIM). This index does not enable an accurate assessment to be made of outstandings or trading volume for use as a proxy for liquidity. The SCM corporate bond indices are stratified into four different investment-grade rating categories: AAA, AA, A, and BBB. In this study, we use SCM's long-term corporate bond indices (motivated in section 1 for callability control), reported for the 08:1976-07:2001 25-year period. Data for the AAA index are available only until March 1993, as no bonds fit this category for the later period.

A long-term *i*-rated corporate bond index, *i*=AAA, AA, A, BBB, consists of all bonds in SCM's *i*-rated corporate bond universe, with remaining terms to maturity greater than 10 years.

Our yield spreads for these long-term indices are calculated with respect to the constant maturity, long-term Government of Canada index, reported by CANSIM. Following Longstaff and Schwartz (1995), to proxy firm assets' returns we use the (continuously compounded) monthly return on the Toronto Stock Exchange 300 index.[9]

Table 1 reports the summary statistics for the time series of absolute and relative yield spreads stratified by credit rating. Similar to the statistics reported by Longstaff and Schwartz (1995) for their U.S. data, the means of both absolute and relative yield spreads monotonically increase as credit quality decreases for all indices. The same is true for the standard deviation of both yield-spread measures.

***Insert Table 1 here**

4.2. Regression Methodology

Our focus is on testing the three regression models introduced in the first section, namely the Longstaff and Schwartz (1995) two-factor model, Duffee's (1998) term-structure model and Collin-Dufresne et al.'s (2001) comprehensive model. For all the models that follow, we estimate the coefficients by OLS and a combined autoregressive and GARCH (1,1) model. Financial time-series are well known to exhibit volatility clustering, a time-varying variance of the innovations. In this sense heteroskedasticity in the regression residuals may reduce the estimation efficiency but is commonly overcome through the GARCH (1,1) specification. Since previous research in this area uses OLS, we first run OLS regressions for comparison reasons. Also, using the Lagrange multiplier (LM) test and the Portmanteau Q-test to our data, we examine their properties. We find that a combined autoregressive and GARCH (1,1) model better fits our data.

The first model tested is based on Longstaff and Schwartz (1995), where the absolute yield spread is defined as the difference between the yield on the relevant SCM index and that of the constant maturity, long-term Government of Canada index. Their model for the absolute spreads is given by:

$$\Delta S = a + b\Delta Y + cI + \varepsilon \qquad (1)$$

where ΔS is the monthly change in the absolute yield spreads, ΔY is the monthly change in the constant maturity, long-term Government of Canada yield, which proxies changes in the riskless rate. I is the monthly return on the Toronto Stock Exchange 300 index, which proxies firm assets' returns.

The second model tested is the Longstaff and Schwartz (1995) model but with relative spreads instead of absolute spreads. The relative yield spread is defined as the ratio between the yield on the relevant SCM index and that of the constant maturity, long-term Government of Canada index. In Table 3, we report our result for the following Longstaff and Schwartz (1995) relative spread regression:

$$\Delta R = a + bPY + cI + \varepsilon \qquad (2)$$

where ΔR is the monthly change in relative yield spreads, and PY is the monthly percentage change in the constant-maturity long-term Government of Canada yield, which proxy's changes in interest rates. I is the monthly return on the Toronto Stock Exchange 300 index, which proxies for firm asset return.

Duffee (1998) uses a regression approach different from that of Longstaff and Schwartz (1995). He regresses spread changes on changes both in the short yield and in a term structure slope variable. Kamara (1997) presents evidence that the slope of the riskless term structure is positively related to expected economic growth. Harvey (1997) presents similar results for Canada. This finding implies a negative relation between default risk and changes in the slope of the riskless term structure. Following Duffee (1998), we estimate the following regression model for every index:

$$\Delta S = \beta_0 + \beta_1 \Delta Y_{T-bill} + \beta_2 \Delta Slope + \varepsilon \qquad (3)$$

where ΔS is the monthly change in the absolute yield spreads, ΔY_{T-bill} is the monthly change in the yield on a three-month Government of Canada treasury bill, and $\Delta Slope$ is the monthly change in the spread between the constant maturity long-term Government of Canada yield and the three-month treasury bill yield.

Collin-Dufresne et al.'s (2001) findings indicate that monthly changes in firm-specific attributes are not the driving force in credit-spread changes. Thus we also run the following regression for each index:

$$\Delta S = \beta_0 + \beta_1 \Delta Y_{LT} + \beta_2 (\Delta Y_{LT})^2 + \beta_3 \Delta Slope + \beta_4 I + \varepsilon, \qquad (4)$$

where ΔY_{LT} is the monthly change in the constant maturity long-term Government of Canada yield and the remaining variables are as defined above.

Following Fridson, Garman and Wu's (1997) prediction, we estimate the following regression for each index:

$$\Delta S = \beta_0 + \beta_1 \Delta \pi + \beta_2 \Delta Y_{LT,R} + \varepsilon \qquad (5)$$

where ΔS is the monthly change in the yield spread, $\Delta Y_{LT,R}$ is the monthly change in the real yield on the constant maturity long-term Government of Canada index, $\Delta\pi$ is the monthly change in the consumer-price index (CPI) that proxies the inflation rate.

5. Empirical Results

The results from applying the three regression models introduced in the first section, namely the Longstaff and Schwartz (1995) two-factor model, Duffee's (1998) term-structure model and Collin-Dufresne et al.'s (2001) comprehensive model are now reported.

5.1 Longstaff and Schwartz's (1995) two-factor model.

In Table 2, we report the OLS estimates for the Longstaff and Schwartz (1995) regression models. They use both absolute and relative spreads in their analysis.[10] Panel A of Table 2 outlines the estimates for our entire sample, covering the 08:1976-07:2001 25-year period. Corporate bonds carrying a standard call provision dominate the data during this sample period. Thus, due to Duffee's (1998) warning, we expect the impact of callability on the results to be significant for that period.

Insert Table 2 here

The results for regression model (1) are in agreement with the results of Longstaff and Schwartz (1995) for their U.S. bonds. The estimated coefficients for b are negative and statistically significant for all indices during the 08:1976-07:2001 estimation period. As in Longstaff and Schwartz (1995), the magnitude of the estimates of b is economically significant for the evolution of yield spreads, and the coefficient b monotonically decreases with credit quality. Also in agreement with Longstaff and Schwartz (1995), Panel A of Table 2 reports that the estimates of c are all negative and statistically significant. The coefficient c is also found to decrease monotonically with credit quality, and is economically significant. This supports the importance of Longstaff and Schwartz's (1995) asset factor, i.e., *ceteris paribus*, higher firm values result in lower probability of default and consequently lower yield spreads.

However, drawing the conclusion that default yield is negatively related to interest rate simply based on the above result is premature. Following Duffee's (1998) results, we believe that this result is due to the negative impact of callability on the estimated relationship. To test this hypothesis, we apply regression model (1) to all indices for the 01:1995-07:2001 period, in which our long indices are expected to be dominated by bonds carrying the doomsday call.

This allows us to control for callability for the BBB index. Recall that for BBB-rated bonds the doomsday call will always be out of the money, making them economically noncallable. This leaves the default term as the only factor potentially affecting the sign of b for the BBB index. Based on their findings, which are similar to our results for the entire 25-year sample period, Longstaff and Schwartz (1995) conclude that the interest-rate factor is more important for lower-rated bonds. Given that the BBB index is the lowest-rated index in our sample, one can expect the default factor, manifested by the interest-rate factor in Longstaff and Schwartz (1995), to have the strongest impact on the yield spread - riskless rate relation estimated for this index.

Panel B of Table 2 reports our results for the 01:1995-07:2001 period. Our estimated coefficients *b* for the BBB index is statistically insignificant.[11] This result implies no relation between credit spreads and the riskless rate for the economically noncallable corporate bonds rated BBB. This result indicates that the default factor, which represents Longstaff and Schwartz's interest-rate factor, is trivial for BBB-rated bonds, for which it is expected to be most important within our sample. Contrasting these results with those reported in Panel A of Table 2 for the same index, it becomes clear that what drives the negative sign of *b* is the negative impact of the callability factor.

In Figure 1 we show that the probability of a call for the doomsday call provision in the AA- and A-rated indices is significant. Thus, one should expect the doomsday call for these indices to induce a negative yield spread - riskless rate relation, as in the case of bonds carrying the standard call provision.[12] This negative relation is confirmed by the results reported in Panel B of Table 2. Our estimated coefficients *b* are negative and statistically significant for the economically callable AA- and A-rated indices during the 01:1995-07:2001 estimation period.

Figure 1 also demonstrates that the probability of a call is significantly higher for AA-rated bonds compared with A-rated bonds. Thus, one may expect the negative impact of the callability term on the sign of *b* to be greater for the AA index than for the A index. The results reported in Panel B of Table 2 support this hypothesis. Above, we saw that for the entire 25-year sample period the estimated coefficient *b* monotonically decreases with credit quality. For the 01:1995-07:2001 period this monotonicity is reversed in line with the moneyness of the doomsday call.

Finally, the estimates for the coefficient *c* for the 01:1995-07:2001 period in Panel B of

Table 2 are still all negative and statistically significant. In general, the coefficient *c* decreases with credit quality, and is similar in magnitude to that estimated for each index for the entire 25-year sample period. This clearly shows that Longstaff and Schwartz's (1995) asset factor is robust.

Next, we examine the results for relative yield spreads. Panel A of Table 3 reports the estimates for our 25-year sample, covering the 08:1976-07:2001 period. Recall that data during this period are dominated by callable bonds, carrying a standard call provision. Like Longstaff and Schwartz (1995), we find stronger support for the negative relation, i.e., higher absolute *t* statistics for *b* and regression R^2's in regression model (2).

Insert Table 3 here

Compared with the results reported in Table 2, the results reported in Panels A and Panel B of Table 3 for regression model (2), indicate that the *t* statistics for the coefficient *b* are always higher when one uses the relative spread. Also, the regression R^2's experience a significant increase when regression model (2) is applied, which suggests that this regression model introduces a negative structure into the data. Focusing on the economically noncallable BBB index during the 01:1995-07:2001 period, it is interesting to note that although we find no yield spread - riskless rate relation for absolute spreads, when relative spreads are used instead, *b* becomes statistically negative.[13]

Instead of interpreting the result to be evidence in support of the negative spread-rate relationship, we analyze the regression structure first and find that the more significant negative relation for relative spreads is born out of a mathematical definition rather than a reflection of economic relationship.

Let *R* denote relative spread, *S* denote the absolute spread, and *Y* denote the riskless rate. Then by definition: $R = \frac{Y+S}{Y}$. We find the following relationship between the sensitivity to interest rate of absolute spread and that of relative spread.

$$\frac{\partial R}{\partial Y} = \frac{\partial(\frac{Y+S}{Y})}{\partial Y} = \frac{Y(1+\frac{\partial S}{\partial Y}) - (Y+S)}{Y^2} = Y^{-1}\frac{\partial S}{\partial Y} - Y^{-2}S \qquad (6)$$

Thus, in equation (6) if the absolute spread is negatively related to the riskless rate ($\frac{\partial S}{\partial Y} < 0$), the relative spread will show a higher magnitude of negative relationship with interest rate. However, even if the absolute spread has no relationship with the riskless rate ($\frac{\partial S}{\partial Y} = 0$), we can still find a strong negative relationship due to the negative sign of $-Y^{-2}S$ in the above equation. Thus one should be cautious when analyzing the result of the relative spread regression model.

Finally, the estimates for the coefficient *c* under regression model (2) for the entire 25-year sample period (Panel A of Table 3) and for the 01:1995-07:2001 period (Panel B of Table 3) all remain negative and statistically significant. As in the analysis of absolute spreads, the last result for relative spreads demonstrates that Longstaff and Schwartz's (1995) asset factor is robust.

5.2 Duffee's (1998) Model

Table 4 outlines the estimates for the Dufee (1998), regression model (3). In some cases, we find that our SCM data set is characterized by the autoregressive nature of the OLS residuals of regression model (3). This may be a function of liquidity in the underlying Canadian bond markets, since Wagner, Hogan and Batten (2005) suggest - with respect to German Eurobonds- that dependence in spread changes may result from a low liquidity in the bond markets investigated as well a relatively illiquid corporate bond market compared to the Government bond market. In our case, the order of the autoregressive process for the residuals varies across the indices. To determine the correct autoregressive order, we apply a stepwise autoregression method. In a number of cases, where the OLS residuals follow an autoregressive process, they also exhibit non-constant volatility consistent with a GARCH (1,1) process. In those instances, we apply a maximum-likelihood estimation procedure for a combined autoregressive model and a GARCH (1,1) model. When the OLS residuals only follow an autoregressive process, we apply the Yule-Walker method. Finally, when the OLS residuals are homoskedastic and do not follow an autoregressive process, we simply use OLS to estimate regression model (3).[14]

Panel A of Table 4 outlines the estimates for our entire sample, covering the 08:1976-07:2001 25-year period. Recall that corporate bonds carrying a standard call provision dominate the data during this sample period. As expected, yield spreads are negatively related to both the three-month bill yield and the slope of the riskless term structure for all ratings. In general, both level and slope coefficients monotonically decrease with credit quality. These results agree with Duffee's (1998) results for his callable U.S. bond portfolios.

Insert Table 4 here

Panel B of Table 4 reports our results for the 01:1995-07:2001 period. For the economically callable AA- and A-rated indices, both level and slope coefficients are negative and statistically significant. Recall that BBB-rated bonds are economically noncallable during this period. Our estimated level and slope coefficients for the BBB index are both statistically insignificant. Applying regression model (3) to his U.S. noncallable bond data, Duffee still finds both level and slope coefficients to be significantly negative, although weak. He attributes this result to the coupon level effect. Our results support this assertion. Since the coupon bias is not prevalent in our SCM data, the negative sign for both slope coefficients disappears.

Recall that Figure 1 demonstrates that the probability of a call is significantly higher for AA-rated bonds compared with A-rated bonds. The results reported in Panel B of Table 4 are in line with this hypothesis. For the entire 25-year sample period, the estimated slope coefficients monotonically decrease with credit quality, whereas for the 01:1995-07:2001 period, this monotonicity is reversed in line with the moneyness of the doomsday call. Thus, applying Duffee's (1998) regression analysis to our data strengthens the hypothesis that callability drives the negative sign of both level and slope coefficients.

5.3. Collin-Dufresne et al.'s (2001) Comprehensive Regression

Next we establish if the previous results are affected by model specification by applying Collin-Dufresne et al.'s (2001) comprehensive regression (4). Since our data consists of bond indices, whereas they use firm level data, we omit some firm-specific explanatory variables used in their regression analysis. As in the previous sections, we apply a maximum-likelihood estimation procedure for a combined autoregressive model and a GARCH (1,1) model. Table 5 outlines the estimates for regression model (4). Again, in some cases, we find that our SCM data set is characterized by the autoregressive nature of the OLS residuals of regression model (4).

Insert Table 5 here

In general, we obtain results consistent with our previous regressions. Panel A of Table 5 outlines the results for our entire sample, covering the 08:1976-07:2001 25-year period, during which corporate bonds carrying a standard call provision dominate the data. We find that yield spreads are negatively related to Government of Canada yields for all bond indices, and also to the return on the Toronto Stock Exchange 300 index.

Panel B of Table 5 reports our results for the 01:1995-07:2001 period. For AA- and A-rated bonds, the yield spread is still significantly negatively related to Government of Canada yields. For BBB-rated bonds, which are economically noncallable during this period, the yield spread – Government yield relation is still insignificant. These results provide additional support for our conclusion that callability dominates the observed negative relationship between credit spread and riskless rate.[15] In the absence of an economically viable call option, the correlation between credit spread and riskless interest rate is insignificant.

5.4. Yield Spreads and Real Interest Rates

Fridson, Garman and Wu (1997) suggest that an empirical investigation of the default risk – riskless rate relation should be based on the real interest rate, rather than nominal rates that are the standard in yield spread studies. They argue that if the company's revenues are fully adjusted with respect to inflation, its effective cost of capital is related to the real riskless rate rather than the nominal rate. As the real riskless rate increases, the firm's cost of short-term debt increases as well, and its available cash flow decreases. This will hurt the firm's ability to issue new long-term debt, and expected default rates will rise.

In agreement with their hypothesis, Fridson, Garman and Wu (1997) find a significantly positive relationship between default rates and real riskless rates. In the context of the current paper, this result implies a positive yield spread – real riskless rate relation when callability is controlled for. In this section, we first examine the relationship between yield spreads and the inflation rate and real riskless rate. Then we test the most comprehensive of the above models for yield spreads, namely the Collin-Dufresne et al. (2001) model, using real, rather than nominal, Government of Canada yield. Both tests show that the yield spread of callable bonds is negatively related to the *nominal* riskless rate. However, when the real riskless rate is used instead, this significant negative correlation disappears. Since the decision to call is based on the level of nominal rates rather than real rates, the yield spread attributed to callability should be negatively related to nominal riskless rate and unrelated to the real rate. This result presents further evidence for our conclusion that call risk dominates the observed negative yield spread – nominal riskless rate relation.

5.5. Inflation, Real Interest Rate and Yield Spreads

Following Fridson, Garman and Wu's (1997) prediction Equation (5) was also estimated with the results reported in Table 6, Again, in some cases, we find that our SCM data set is characterized by the autoregressive nature of the OLS residuals of regression model (5). As in the above sections, we apply a maximum-likelihood estimation procedure for a combined autoregressive model and a GARCH (1,1) model.

Insert Table 6 here

Panel A of Table 6 reports that yield spreads are negatively related to both the inflation rate and the real government yield for all callable indices. It is interesting to note that the coefficients of both explanatory variables are very similar in magnitude. We interpret this as evidence that the influence of riskless term structure on yield spreads is due to mainly *nominal* riskless rates. Based on a Wald test we fail to reject the hypothesis that $\beta_1 = \beta_2$. Thus, we can rewrite regression (5) as:

$$\Delta S = \beta_0 + \beta_1(\Delta\pi + \Delta Y_{LT,R}) + \varepsilon \qquad (7)$$

where $(\Delta\pi + \Delta Y_{LT,R})$ is the monthly change in nominal government yield. Thus, these results are in line with the fact that the moneyness of the call provision is based on the level of nominal rates rather than real rates, which will induce a negative relationship between the yield premium attributed to callability and the nominal riskless rate. Not surprisingly, for the economically noncallable BBB-rated bonds during the 01:1995-07:2001 period the relationship is insignificant with both the inflation rate and the real rate (Panel B of Table 6). Note that for the same period, for the economically callable AA- and A-rated bonds the relationship is significantly negative with both variables.

5.6. Collin-Dufresne et al.'s (2001) Model with Real Rates

To check that the above analysis with the real interest rate is not affected by model specification, we test the most comprehensive regression model - Collin-Dufresne et al.'s (2001) - with real government rates. Table 7 outlines the results of this estimation. In Panel A, which reports the results for the entire sample, the only significant factors are the term structure slope changes and the stock index return. The interest rate factor is not significant when real rates are used. This is in line with the expectation of a negative relationship between the yield premium attributed to callability and the nominal riskless rate rather than the real rate. Panel B of Table 7 shows similar results for the 01:1995-07:2001 period.

Insert Table 7 here

We use the difference between the nominal interest rate and inflation rate as our measure for real yield in the above regressions[16]. We find that once the real interest rate is used the relationship between the bond yield spread and riskless interest rate becomes insignificant even for AA and A indices, which are mainly composed of callable bonds. This result shows that once the relationship between callability and the nominal interest rate is factored out, no relationship remains between the interest rate and credit spread.

To further test the robustness of our results, we use the return of Real Return Bonds (RRBs) issued by Government of Canada since 1991. This is a better proxy for the real yield.[17] We run the most comprehensive regression model - Collin-Dufresne et al.'s (2001) - with rate of return on RRBs. We report both OLS and AR-GARCH results in Table 8. Once again, the results are very similar to those of Table 7.

Insert Table 8 here

To sum up, irrespective of the model tested, namely the Longstaff and Schwartz (1995) two-factor model, Duffee's (1998) model, or Collin-Dufresne et al.'s (2001) comprehensive model, we find evidence the negative relationship between the yield spread and the government yield is due mainly to the effects of the call provision. Once the impact of the call option is removed, i.e. when considering BBB-rated economically noncallable bonds, the yield spread – government yield relation is not significant. When we use real government yields, we find that the yield spread – riskless rate relationship is insignificant for both callable and economically noncallable bonds.

6 A Direct Test for Default Risk

In the previous section we use several well-established regression models to estimate the yield spread – riskless rate relation for Canadian data. Because in recent year's bonds in the Canadian BBB bonds index are not exposed to callability and tax effects, it is safe to assume that their yield spread is mainly driven by credit risk. Moreover, BBB bonds serve as an excellent test case for the question of whether default risk generates a negative yield spread – riskless rate relation. This is because this risk is quite substantial for bonds of this rating category.

We find that, when callability and tax effects are controlled for, there is no significant yield spread – riskless rate relation. Thus, contrary to what previous studies conclude from their tests (see for example: Longstaff and Schwartz, 1995; and Duffee, 1998), our results cast doubt on one of the most notable predictions of structural models; that credit spreads are negatively related to the riskless rate. Recall that risk-neutral valuation is applied in the context of structural models. *Ceteris paribus*, under risk-neutral valuation an increase in the riskless rate implies a higher expected future value for the firm's assets relative to the default threshold, and a lower risk-neutral probability in default. This ultimately results in a negative credit spread – riskless rate relation.

To substantiate our contention that this prediction of structural models does not hold empirically, in this section we offer an alternative test that involves default rates. The advantage of this test is in that default rates serve as a more direct measure of credit risk. They are also clean of callability and tax effects. We obtain historical default rates from Moody's Investors Service. Moody's historical default rates are based on the credit histories of nearly 10,000 corporate and sovereign entities and over 80,000 individual debt securities since 1970.

We choose to apply the following Collin-Dufresne et al.'s (2001) type regression model of Equation (5) to incorporate Moody's default rates:

$$\Delta DR = \beta_0 + \beta_1 \Delta Y_{LT} + \beta_2 (\Delta Y_{LT})^2 + \beta_3 \Delta Slope + \beta_4 I + \varepsilon \qquad (8)$$

where ΔDR is the monthly change in Moody's default rate and the remaining variables are as defined above.

We initially assume that Moody's trailing twelve-month all corporate default rates, serve as a good proxy for expected future default probabilities, and therefore use these default rates in regression model (6). This sample covers the 12:1979-11:2001 period. Noting that these are *ex post* default rates, we then use Moody's monthly speculative- grade default rate forecasts (12:1999-11:2004), which may serve as a better proxy for expected default probabilities.[18]

Note that, in 2001 81.2% of the number of defaults in Moody's default universe occurred in North America (U.S. and Canada). Thus, these default rates also reflect the Canadian default experience. However, most of the North American defaults (77%) are those of U.S. companies. Thus, we also apply regression model (6) with U.S. data replacing the Canadian data to proxy the independent variables. Following Collin-Dufresne et al. (2001) we use the following proxies: the 10-year benchmark Treasury yield represents the level of the term structure; for the slope we use the spread between the 10-year and 2-year benchmark Treasury yields; and finally we use returns on the S&P 500 index.

The results are reported in Table 9. In panel A, which reports the results using historical Moody's trailing twelve-month all corporate default rates, we find that when U.S. determinants (term structure components and asset returns) are used, the stock index return coefficient is significantly negative. This is expected since higher stock returns imply higher asset returns and lower default rates. At the same time, the term structure variables are not statistically significant. When Canadian data are used instead, we find that none of the independent variables are statistically significant. This is also expected, since most of the defaults are attributed to U.S. rather than Canadian companies.

Insert Table 9 here

Recall that forecasted default rates may serve as a better proxy for expected default probabilities. In Panel B, we use Moody's monthly speculative-grade default rate forecasts and find similar results. The results for the U.S. data are similar to the results of the historical default rates, except that now all coefficients are insignificant. Surprisingly, for the Canadian data we find a positive and significant relation between default-rate forecasts and the riskless rate. This is in contrast to the negative correlation found for the callable Canadian bond indices, and the lack of a significant relation found for the noncallable indices.

The results in this section support our contention that the prediction of structural models – a negative relation between the default probabilities (and credit spreads) and the riskless rate – does not hold empirically. This is in contrast to the negative result found for yield spreads in previous studies (see for example: Longstaff and Schwartz, 1995; Duffee, 1998; and Collin-Dufresne et al., 2001). Our results are robust as, in contrast with yield spreads, default rates serve as a more direct measure of credit risk since they are free of callability and tax effects.

7. Summary and Conclusions

Longstaff and Schwartz (1995) provide a model for the valuation of corporate bonds, which account for two stochastic factors, which accommodate the effects of interest-rates and the firm's asset value. One of the most notable predictions of this and other structural models with an asset-based default process is that credit spreads are negatively related to the riskless rate. Longstaff and Schwartz test this prediction along with their predicted negative impact of the asset factor on credit spreads using Moody's indices and consider the negative relationship they find as evidence supporting their model's prediction.

Duffee (1998) argues that most of the bonds in Moody's indices are callable bonds and thus the negative yield spread – riskless rate relation could be largely due to the negative relationship between the yield premium attributed to callability and the riskless rate. Instead, Duffee uses noncallable bond portfolios and finds that the negative yield spread – riskless rate relation is much weaker once callability is controlled. Elton et al. (2001), point out that Duffee's (1998) analysis ignores the effects of state taxes on U.S. bonds, and they argue that differential taxes on corporate and government bonds have an important impact on corporate bond yield spreads.

Canadian corporate bond indices are devoid of tax effects since Canadian corporate and government bonds, unlike U.S. bonds, are subject to the same tax treatment. These indices also contain a call provision that allows for identifying callable and economically noncallable bonds. Using this Canadian bond index data, we find an insignificant yield spread – government yield relation for economically noncallable bonds. For bonds with an economically viable call option, the negative relation we find increases with the moneyness of the call option. Our tests of different models with both real interest rates and nominal interest rates, with both yield spread and real (expected) default rates, consistently show the robustness of our results. We conclude that call risk dominates the negative yield spread – government yield relation. We further conclude that, for investment-grade bonds, the role of the asset factor (manifesting default risk) in influencing the sensitivity of corporate bond yield spreads to government yields is not significant.

These results indicate that a gap remains in our understanding of the default process. Theoretically, structural models suggest that an increase in the riskless rate implies a higher expected future value for the firm's asset relative to the default threshold, and a lower risk-neutral probability of default and risk neutral credit spread. Empirically, we show that bondholders do not adjust their required default premium for an increase in the riskless rate. Our results provide support for reduced-form models that explicitly define a default hazard process and untie the relation between the firm's asset value and default probability.

Appendix

Insert Table A.1 here

Insert Table A.2 here

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Table 1
Summary Statistics for Yield Spreads in SCM Long-Term Corporate Bond Indices for the August 1976 to July 2001 Period

The yield spread is the difference between the yield on a long-term index and the yield on the constant maturity, long-term Government of Canada index. The relative spread is the ratio of the yield on a long-term index to the yield on the constant maturity, long-term Government of Canada index. Data for the AAA indices are available only until March 1993.

	No. of Observations	Mean of Credit Spread	Std. Dev. Of Credit Spread	Mean of Relative Spread	Std. Dev. Of Relative Spread
AAA	200	0.5761	0.3050	1.0559	0.0331
AA	300	0.6350	0.3184	1.0716	0.0434
A	300	0.8612	0.3533	1.0980	0.0548
BBB	300	1.5359	0.7916	1.1842	0.1266

Table 2

Regressions of Changes in Absolute Yield Spreads of SCM Long-Term Corporate Bond Indices on Changes in the Yield of the Constant Maturity Long-Term Government of Canada Index and the Return on the Toronto Stock Exchange 300 Index - OLS Estimation

Table 2 reports the results of the OLS estimation of regression model (1) in which the dependent variable is the monthly change in the absolute yield spread. This regression model is of the following form:

$$\Delta S = a + b\Delta Y + cI + \varepsilon,$$

where ΔS is the monthly change in the absolute yield spread, ΔY is the monthly change in the yield of the constant maturity, long-term Government of Canada index, and I is the monthly return on the Toronto Stock Exchange 300 index. t-values are in parentheses. Panel A reports the estimates for the entire sample, covering the 08:1976-07:2001 25-year period. Data during this sample period are dominated by corporate bonds carrying a standard call provision. Panel B outlines the results for the 01:1995-07:2001 sub-period, in which bonds carrying the doomsday call are expected to dominate all indices.
are expected to dominate all indices.

Panel A: 09:1976-07:2001				
Index	*a*	*b*	*c*	R^2
AAA	0.0057	-0.1447	-0.6796	0.10
	(0.39)	(-4.66)	(-2.21)	
AA	0.0026	-0.1538	-0.6808	0.11
	(0.24)	(-5.86)	(-3.02)	
A	0.0043	-0.1513	-0.7903	0.12
	(0.41)	(-5.99)	(-3.65)	
BBB	0.0086	-0.1892	-1.0957	0.08
	(0.53)	(-4.77)	(-3.22)	

Panel B: 01:1995-07:2001				
Index	*a*	*b*	*c*	R^2
AAA	N/A			
AA	-0.0041	-0.2106	-0.6623	0.14
	(-0.27)	(-3.12)	(-2.26)	
A	0.0015	-0.1927	-0.6160	0.13
	(0.10)	(-3.02)	(-2.22)	
BBB	0.0022	0.0442	-1.2557	0.11
	(0.10)	(0.44)	(-2.87)	

Table 3

Regressions of Changes in Relative Yield Spreads of SCM Long-Term Corporate Bond Indices on Percentage Changes in the Yield of the Constant Maturity Long-Term Government of Canada Index and the Return on the Toronto Stock Exchange 300 Index - OLS Estimation

Table 3 reports the results of the OLS estimation of regression model (2) in which the dependent variable is the monthly change in the relative yield spread. This regression model is of the following form:

$$\Delta R = a + b\Delta PY + cI + \varepsilon,$$

where ΔR is the monthly change in the relative yield spread, ΔPY is the monthly percentage change in the yield of the constant maturity, long-term Government of Canada index, and I is the monthly return on the Toronto Stock Exchange 300 index. t-values are in parentheses. Panel A reports the estimates for the entire sample, covering the 08:1976-07:2001 25-year period. Data during this sample period are dominated by corporate bonds carrying a standard call provision. Panel B outlines the results for the 01:1995-07:2001 sub-period, in which bonds carrying the doomsday call are expected to dominate all indices.

Panel A: 09:1976-07:2001				
Index	*a*	*b*	*c*	R^2
AAA	0.0007 (0.59)	-0.1927 (-6.16)	-0.0553 (-2.26)	0.16
AA	0.0005 (0.44)	-0.2304 (-7.76)	-0.0801 (-3.56)	0.17
A	0.0007 (0.68)	-0.2573 (-9.09)	-0.0889 (-4.15)	0.22
BBB	0.0013 (0.66)	-0.3778 (-6.84)	-0.1247 (-2.98)	0.14

Panel B: 01:1995-07:2001				
Index	*a*	*b*	*c*	R^2
AAA	N/A			
AA	-0.0002 (-0.08)	-0.3294 (-4.42)	-0.1266 (-2.69)	0.23
A	0.0006 (0.25)	-0.3502 (-4.83)	-0.1214 (-2.65)	0.25
BBB	0.0009 (0.23)	-0.2879 (-2.48)	-0.2223 (-3.04)	0.14

Table 4
Regressions of Changes in Absolute Yield Spreads of SCM Long-Term Corporate Bond Indices on Changes in Government of Canada Yields - AR-GARCH Estimation

Table 4 reports the results of the AR-GARCH estimation of regression model (3) in which the dependent variable is the monthly change in the absolute yield spread. This regression model is of the following form:

$$\Delta S = \beta_0 + \beta_1 \Delta Y_{T\text{-}bill} + \beta_2 \Delta Slope + \varepsilon,$$

where ΔS is the monthly change in the absolute yield spread, $\Delta Y_{T\text{-}bill}$ is the monthly change in the three-month Treasury Bill yield, and $\Delta Slope$ is the monthly change in the spread between the constant maturity long-term Government of Canada index and the three-month Treasury bill yield. *t*-values are in parentheses, *m* gives the degree of the autoregressive process as determined by the stepwise autoregression method, *p* and *q* are the GARCH(*p*,*q*) parameters, Norm. Test gives the *p*-value for the normality test for detecting misspecification of the GARCH model, and finally LM gives the *p*-value for the Lagrange multiplier test. Panel A reports the estimates for the entire sample, covering the 08:1976-07:2001 25-year period. Data during this sample period are dominated by corporate bonds carrying a standard call provision. Panel B outlines the results for the 01:1995-07:2001 sub-period, in which bonds carrying the doomsday call are expected to dominate all indices.

Panel A: 09:1976-07:2001									
	Regression Coefficients			AR and GARCH Parameters					
Index	β_0	β_1	β_2	*m*	*p*	*q*	Norm. Test	R^2	LM
AAA	-0.0027 (-0.51)	-0.0442 (-2.51)	-0.0806 (-4.25)	1, 2, 4	1	1	<0.0001	0.19	<0.0001
AA	-0.0074 (-1.82)	-0.1231 (-8.72)	-0.1893 (-12.56)	1, 2, 4	1	1	0.0005	0.18	<0.0001
A	-0.0063 (-1.13)	-0.1011 (-6.22)	-0.1514 (-8.16)	1, 2	1	1	0.0001	0.16	<0.0001
BBB	0.0025 (0.18)	-0.1244 (-3.26)	-0.1996 (-4.72)	-	1	1	<0.0001	0.09	-

Panel B: 01:1995-07:2001									
	Regression Coefficients			AR and GARCH Parameters				Goodness of Fit	
Index	β_0	β_1	β_2	*m*	*p*	*q*	Norm. Test	R^2	LM
AAA	N/A								
AA	-0.0076 (-0.48)	-0.1732 (-2.34)	-0.1751 (-2.44)	-	-	-	-	0.08	-
A	-0.0016 (-0.11)	-0.1519 (-2.17)	-0.1645 (-2.42)	-	-	-	-	0.08	-
BBB	-0.0037 (-0.15)	0.1361 (1.21)	0.0947 (0.87)	-	-	-	-	0.02	-

Table 5
Regressions of Changes in Nominal Yield Spreads of SCM Long-Term Corporate Bond Indices on Changes in Selected Nominal Determinants -
AR-GARCH Estimation

Table 5 reports the results of the AR-GARCH estimation of regression model (4) in which the dependent variable is the monthly change in the nominal yield spread. This regression model is of the following form:

$$\Delta S = \beta_0 + \beta_1 \Delta Y_{LT} + \beta_2 (\Delta Y_{LT})^2 + \beta_3 \Delta Slope + \beta_4 I + \varepsilon,$$

where ΔS is the monthly change in the nominal yield spread, ΔY_{LT} is the monthly change in the nominal yield on the constant maturity long-term Government of Canada index, $(\Delta Y_{LT})^2$ is a convexity term, $\Delta Slope$ is the monthly change in the nominal spread between the constant maturity long-term Government of Canada index and the three-month Treasury bill yield, and I is the monthly nominal return on the Toronto Stock Exchange 300 index. t-values are in parentheses, m gives the degree of the autoregressive process as determined by the stepwise autoregression method, p and q are the GARCH(p,q) parameters, Norm. Test gives the p-value for the normality test for detecting misspecification of the GARCH model, and finally LM gives the p-value for the Lagrange multiplier test. Panel A reports the estimates for the entire sample, covering the 08:1976-07:2001 25-year period. Data during this sample period are dominated by corporate bonds carrying a standard call provision. Panel B outlines the results for the 01:1995-07:2001 sub-period, in which bonds carrying the doomsday call are expected to dominate all indices.

Panel A: 09:1976-07:2001											
		Regression Coefficients				AR and GARCH Parameters					
Index	β_0	β_1	β_2	β_3	β_4	m	p	q	Norm. Test	R^2	LM
AAA	-0.0101 (-1.51)	-0.1547 (-11.06)	0.0614 (3.80)	-0.0887 (-7.40)	-0.0055 (-3.61)	1, 2, 4	1	1	<0.0001	0.22	0.0004
AA	-0.0063 (-1.47)	-0.1196 (-9.67)	0.0429 (4.40)	-0.0679 (-6.90)	-0.0050 (-4.68)	1, 2, 4	1	1	0.0018	0.23	0.0067
A	-0.0067 (-1.07)	-0.1208 (-8.86)	0.0600 (5.59)	-0.0541 (-5.06)	-0.0058 (-4.66)	1, 2	1	1	0.0021	0.20	<0.0001
BBB	0.0029 (0.21)	-0.1467 (-3.82)	0.0447 (1.58)	-0.0701 (-2.67)	-0.0109 (-3.30)	-	1	1	<0.0001	0.15	-

Panel B: 01:1995-07:2001											
		Regression Coefficients				AR and GARCH Parameters				Goodness of Fit	
Index	β_0	β_1	β_2	β_3	β_4	m	p	q	Norm. Test	R^2	LM
AAA	N/A										
AA	-0.0154 (-0.85)	-0.1676 (-2.02)	0.2605 (1.31)	-0.0041 (-0.08)	-0.0070 (-2.31)	-	-	-	-	0.16	-
A	-0.0001 (-0.00)	-0.1817 (-2.29)	0.0541 (0.29)	-0.0041 (-0.08)	-0.0063 (-2.17)	-	-	-	-	0.13	-
BBB	0.0047 (0.14)	0.0510 (0.43)	-0.0081 (-0.03)	0.0141 (0.18)	-0.0132 (-3.03)	-	-	-	-	0.16	-

Table 6
Regressions of Changes in Absolute Nominal Yield Spreads of SCM Long-Term Corporate Bond Indices on Changes in the Real Yield of the Constant Maturity Long-Term Government of Canada Index and on Changes in the Monthly Inflation Rate

Table 6 reports the results of the AR-GARCH estimation of regression model (5). This regression model is of the following form:

$$\Delta S = \beta_0 + \beta_1 \Delta Inf + \beta_2 \Delta Y_{LT} + \varepsilon,$$

where ΔS is the monthly change in the nominal yield spread, ΔY_{LT} is the monthly change in the real yield on the constant maturity long-term Government of Canada index, ΔInf is the monthly change in the inflation rate. *t*-values are in parentheses, *m* gives the degree of the autoregressive process as determined by the stepwise autoregression method, *p* and *q* are the GARCH(*p*,*q*) parameters, Norm. Test gives the *p*-value for the normality test for detecting misspecification of the GARCH model, and finally LM gives the *p*-value for the Lagrange multiplier test. Panel A reports the estimates for the entire sample, covering the 08:1976-07:2001 25-year period. Data during this sample period are dominated by corporate bonds carrying a standard call provision. Panel B outlines the results for the 01:1995-07:2001 sub-period, in which bonds carrying the doomsday call are expected to dominate all indices.

Panel A: 09:1976-07:2001									
	Regression Coefficient			AR and GARCH Parameters					
Index	β_0	β_1	β_2	*m*	*p*	*q*	Norm. Test	R^2	LM
AAA	-0.0012 (-0.15)	-0.2232 (-8.02)	-0.1801 (-8.47)	1,4	1	1	<0.0001	0.14	<.0001
AA	-0.0036 (-0.8)	-0.1542 (-6.74)	-0.1350 (-9.03)	1, 2, 4	1	1	<0.0001	0.16	<.0001
A	-0.0049 (-0.82)	-0.1527 (-5.99)	-0.1204 (-7.11)	1, 2	1	1	0.0003	0.14	<0.0001
BBB	0.0018 (-0.13)	-0.1822 (-3.22)	-0.1454 (-3.83)	-	1	1	<0.0001	0.07	-

Panel B: 01:1995-07:2001									
	Regression Coefficients			AR and GARCH Parameters				Goodness of Fit	
Index	β_0	β_1	β_2	*m*	*p*	*q*	Norm. Test	R^2	LM
AAA	N/A								
AA	-0.0079 (-0.5)	-0.2082 (-2.33)	-0.1778 (-2.62)	-	-	-	-	0.08	-
A	-0.0021 (-0.14)	-0.1916 (-2.27)	-0.1629 (-2.53)	-	-	-	-	0.08	-
BBB	-0.0045 (-0.19)	0.1266 (-0.93)	0.1143 (-1.11)	-	-	-	-	0.16	-

Table 7
Regressions of Changes in Real Yield Spreads of SCM Long-Term Corporate Bond Indices on Changes in Selected Real Determinants -
AR-GARCH Estimation

Table 7 reports the results of the AR-GARCH estimation of regression model (4) in which the dependent variable is the monthly change in the real yield spread. This regression model is of the following form:

$$\Delta S = \beta_0 + \beta_1 \Delta Y_{LT} + \beta_2(\Delta Y_{LT,R})^2 + \beta_3 \Delta Slope + \beta_4 I + \varepsilon,$$

where ΔS is the monthly change in the real yield spread, $\Delta Y_{LT,R}$ is the monthly change in the real yield on the constant maturity long-term Government of Canada index, $(\Delta Y_{LT,R})^2$ is a convexity term, $\Delta Slope$ is the monthly change in the real spread between the constant maturity long-term Government of Canada index and the three-month Treasury bill yield, and I is the monthly real return on the Toronto Stock Exchange 300 index. t-values are in parentheses, m gives the degree of the autoregressive process as determined by the stepwise autoregression method, p and q are the GARCH(p,q) parameters, Norm. Test gives the p-value for the normality test for detecting misspecification of the GARCH model, and finally LM gives the p-value for the Lagrange multiplier test. Panel A reports the estimates for the entire sample, covering the 08:1976-07:2001 25-year period. Data during this sample period are dominated by corporate bonds carrying a standard call provision. Panel B outlines the results for the 01:1995-07:2001 sub-period, in which bonds carrying the doomsday call are expected to dominate all indices.

Panel A: 09:1976-07:2001											
			Regression Coefficients			AR and GARCH Parameters					
Index	β_0	β_1	β_2	β_3	β_4	*m*	*p*	*q*	Norm. Test	R^2	LM
AAA	-0.0041	0.0019	0.0002	-0.0418	-0.0005	1, 2	1	1	<0.0001	0.15	<.0001
	(-0.55)	(0.13)	(0.02)	(-3.06)	(-0.31)						
AA	-0.0040	-0.0178	0.0052	-0.0706	-0.0040	1, 2, 4	1	1	<.0001	0.19	0.0005
	(-0.71)	(-1.36)	(0.66)	(-5.91)	(-2.90)						
A	-0.0032	0.0005	0.0044	-0.0571	-0.0034	1, 2	1	1	0.0012	0.15	<.0001
	(-0.52)	(0.04)	(0.58)	(-4.28)	(-2.81)						
BBB	-0.0022	-0.0029	0.0208	-0.0818	-0.0078	-	1	1	<0.0001	0.11	-
	(-0.15)	(-0.21)	(1.38)	(-3.10)	(-2.38)						
Panel B: 01:1995-07:2001											
			Regression Coefficients			AR and GARCH Parameters				Goodness of Fit	
Index	β_0	β_1	β_2	β_3	β_4	*m*	*p*	*q*	Norm. Test	R^2	LM
AAA	N/A										
AA	-0.0180	-0.0104	0.1060	-0.0453	-0.0044	-	-	-	-	0.16	-
	(-1.15)	(-0.29)	(2.19)	(-0.98)	(-1.49)						
A	-0.0115	-0.0130	0.0927	-0.0503	-0.0042	-	-	-	-	0.10	-
	(-0.64)	(-0.37)	(1.89)	(-1.05)	(-1.45)						
BBB	-0.0185	0.0465	0.0997	-0.0131	-0.0122	-	-	-	-	0.13	-
	(-0.68)	(0.86)	(1.33)	(-0.18)	(-2.77)						

Table 8
Regressions of Changes in Absolute Yield Spreads of SCM Long-Term Corporate Bond Indices on Changes in RRBs and Other Seclected Determinants

Table 8 reports the results of both OLS and AR-GARCH estimation of regression model (4) in which the dependent variable is the monthly change in the absolute yield spread. This regression model is of the following form:

$$\Delta S = \beta_0 + \beta_1 \Delta Y_{LT} + \beta_2 (\Delta Y_{LT})^2 + \beta_3 \Delta Slope + \beta_4 I + \varepsilon,$$

where ΔS is the monthly change in the absolute yield spread, ΔY_{LT} is the monthly yield change on the long-term Government of Canada Real Return Bonds (RBBs), $(\Delta Y_{LT})^2$ is a convexity term, $\Delta Slope$ is the monthly change in the spread between the constant maturity long-term Government of Canada index and the three-month Treasury bill yield, and I is the monthly real return on the Toronto Stock Exchange 300 index. *t*-values are in parentheses, *m* gives the degree of the autoregressive process as determined by the stepwise autoregression method, *p* and *q* are the GARCH(p,q) parameters, Norm. Test gives the *p*-value for the normality test for detecting misspecification of the GARCH model, and finally LM gives the *p*-value for the Lagrange multiplier test. Panel A reports the OLS estimates for AA, A, BBB bonds, covering the 12:1991-07:2001 11-year period. Panel B reports the AR-GARCH estimates for AA, A, BBB bonds, covering the 12:1991-07:2001 11-year period.

Panel A: 12:1991-07:2001 OLS											
Index	β_0	β_1	β_2	β_3	β_4					R^2	
AA	-0.0057	-0.0491	0.4723	-0.1030	-0.0058	-	-	-	-	0.16	
	(-0.39)	(-0.18)	(1.35)	(-3.86)	(-1.97)						
A	0.0052	-0.0200	-0.0047	-0.0856	-0.0051	-	-	-	-	0.13	
	(0.38)	(-0.23)	(-0.01)	(-3.48)	(-1.87)						
BBB	0.0219	0.0288	-0.7100	-0.2510	-0.0074	-	-	-	-	0.13	
	(0.59)	(0.11)	(-0.81)	(-3.74)	(-0.99)						
Panel B: 12:1991-07:2001 AR-GARCH											
	Regression Coefficients					AR and GARCH Parameters				Goodness of Fit	
Index	β_0	β_1	β_2	β_3	β_4	*m*	*p*	*q*	Norm. Test	R^2	LM
AA	-0.0042	0.0389	0.4067	-0.0988	-0.0051	1,2	-	-	-	0.23	0.0005
	(-0.41)	(0.41)	(1.31)	(-4.19)	(-1.86)						
A	0.0074	0.0073	-0.1152	-0.0904	-0.0047	1	-	-	-	0.17	0.0023
	(0.66)	(0.08)	(-0.38)	(-3.75)	(-1.81)						
BBB	0.0219	0.0288	-0.7100	-0.2510	-0.0074	-	-	-	-	0.18	-
	(0.59)	(0.11)	(-0.81)	(-3.74)	(-0.99)						

Table 9
Regressions of Changes in Default Rate of Corporate Bonds on Changes in Selected Nominal Determinants -
AR-GARCH Estimation

Table 9 reports the results of the AR-GARCH estimation of regression model (6) in which the dependent variable is the monthly change in the default rates from Moody's. This regression model is of the following form:

$$\Delta DR = \beta_0 + \beta_1 \Delta Y_{LT} + \beta_2 (\Delta Y_{LT})^2 + \beta_3 \Delta Slope + \beta_4 I + \varepsilon,$$

where ΔDR is the monthly change in the default rates, ΔY_{LT} is the monthly change in the nominal yield on the constant maturity long-term Government of Canada (U.S.) index, $(\Delta Y_{LT})^2$ is a convexity term, $\Delta Slope$ is the monthly change in the nominal spread between the constant maturity long-term Government of Canada (U.S.) index and the three-month Treasury bill yield, and I is the monthly nominal return on the Toronto Stock Exchange 300 index (U.S. S&P 500 index). t-values are in parentheses, m gives the degree of the autoregressive process as determined by the stepwise autoregression method, p and q are the GARCH(p,q) parameters, Norm. Test gives the p-value for the normality test for detecting misspecification of the GARCH model, and finally LM gives the p-value for the Lagrange multiplier test. Panel A reports the estimates for historical tracing default rates of all North-American corporate bonds, covering the 01:1980-12:2001 22 year period. Panel B outlines the results for the 01:2000-09:2004 period, in which forecasted default rates for only speculative bonds are used as dependent variables.

Panel A: 01:1980-12:2001											
			Regression Coefficients			AR and GARCH Parameters					
Countries	β_0	β_1	β_2	β_3	β_4	m	p	q	Norm. Test	R^2	LM
U.S.	0.0002 (1.35)	-0.0001 (-0.27)	-0.0002 (-0.36)	0.0002 (0.44)	-0.0050 (-2.72)	1,2	0	1	<0.0001	0.16	0.0070
Canada	0.0001 (0.73)	0.0000 (0.14)	0.0000 (0.16)	0.0002 (1.25)	0.0000 (0.36)	1	1	1	<0.0001	0.13	0.0050
Panel B: 01:2000-09:2004											
			Regression Coefficients			AR and GARCH Parameters				Goodness of Fit	
Countries	β_0	β_1	β_2	β_3	β_4	m	p	q	Norm. Test	R^2	LM
U.S.	-0.0009 (-1.07)	-0.0020 (-0.78)	0.0070 (1.031)	0.0040 (1.04)	0.0080 (0.64)	-	-	-	-	0.05	-
Canada	-0.0009 (-3.21)	0.0037 (2.35)	0.0240 (4.57)	0.0006 (0.73)	-0.0007 (-0.12)	-	-	-	-	0.13	-

Table A.1

Regressions of Changes in Absolute Yield Spreads of SCM Long-Term Corporate Bond Indices on Changes in the Yield of the Constant Maturity Long-Term Government of Canada Index and the Return on the Toronto Stock Exchange 300 Index - AR-GARCH Estimation

Table A.1 reports the results of the AR-GARCH estimation of regression model (1) in which the dependent variable is the monthly change in the absolute yield spread. This regression model is of the following form:

$$\Delta S = a + b\Delta Y + cI + \varepsilon,$$

where ΔS is the monthly change in the absolute yield spread, ΔY is the monthly change in the yield of the constant maturity, long-term Government of Canada index, and I is the monthly return on the Toronto Stock Exchange 300 index. t-values are in parentheses, m gives the degree of the autoregressive process as determined by the stepwise autoregression method, p and q are the GARCH(p,q) parameters, Norm. Test gives the p-value for the normality test for detecting misspecification of the GARCH model, and finally LM gives the p-value for the Lagrange multiplier test. Panel A reports the estimates for the entire sample, covering the 08:1976-07:2001 25-year period. Data during this sample period are dominated by corporate bonds carrying a standard call provision. Panel B outlines the results for the 01:1995-07:2001 sub-period, in which bonds carrying the doomsday call are expected to dominate all indices.

Panel A: 09:1976-07:2001									
	Regression Coefficients			AR and GARCH Parameters					
Index	*a*	*b*	*c*	*m*	*p*	*q*	Norm. Test	R^2	LM
AAA	0.0007 (0.12)	-0.0563 (-3.27)	-0.0302 (-0.20)	1, 2	1	1	<0.0001	0.19	0.0004
AA	0.0006 (0.12)	-0.1428 (-9.67)	-0.4287 (-4.02)	1, 2, 4	1	1	<0.0001	0.18	0.0006
A	0.0003 (0.05)	-0.1346 (-8.88)	-0.5162 (-4.26)	1, 2	1	1	<0.0001	0.19	<0.0001
BBB	0.0098 (0.81)	-0.1641 (-4.37)	-1.2119 (-3.76)	1, 2	-	-	-	0.12	-

Panel B: 01:1995-07:2001									
	Regression Coefficients			AR and GARCH Parameters				Goodness of Fit	
Index	*a*	*b*	*c*	*m*	*p*	*q*	Norm. Test	R^2	LM
AAA	N/A								
AA	-0.0041 (-0.27)	-0.2106 (-3.12)	-0.6623 (-2.26)	-	-	-	-	0.14	-
A	0.0015 (0.10)	-0.1927 (-3.02)	-0.6160 (-2.22)	-	-	-	-	0.13	-
BBB	0.0026 (0.09)	0.0578 (0.58)	-1.3416 (-3.26)	3	-	-	-	0.17	-

Table A.2

Regressions of Changes in Relative Yield Spreads of SCM Long-Term Corporate Bond Indices on Percentage Changes in the Yield of the Constant Maturity Long-Term Government of Canada Index and the Return on the Toronto Stock Exchange 300 Index - AR-GARCH Estimation

Table A.2 reports the results of the AR-GARCH estimation of regression model (2) in which the dependent variable is the monthly change in the relative yield spread. This regression model is of the following form:

$$\Delta R = a + b\Delta PY + cI + \varepsilon,$$

where ΔR is the monthly change in the relative yield spread, ΔPY is the monthly percentage change in the yield of the constant maturity, long-term Government of Canada index, and I is the monthly return on the Toronto Stock Exchange 300 index. t-values are in parentheses, m gives the degree of the autoregressive process as determined by the stepwise autoregression method, p and q are the GARCH(p,q) parameters, Norm. Test gives the p-value for the normality test for detecting misspecification of the GARCH model, and finally LM gives the p-value for the Lagrange multiplier test. Panel A reports the estimates for the entire sample, covering the 08:1976-07:2001 25-year period. Data during this sample period are dominated by corporate bonds carrying a standard call provision. Panel B outlines the results for the 01:1995-07:2001 sub-period, in which bonds carrying the doomsday call are expected to dominate all indices.

Panel A: 09:1976-07:2001									
	Regression Coefficients			AR and GARCH Parameters					
Index	*a*	*b*	*c*	*m*	*p*	*q*	Norm. Test	R^2	LM
AAA	0.0002 (0.28)	-0.1510 (-8.52)	-0.0193 (-1.53)	1, 2	1	1	0.0005	0.24	0.0003
AA	0.0004 (0.90)	-0.1541 (-7.89)	-0.0589 (-4.62)	1, 2	1	1	<0.0001	0.21	0.0057
A	0.0002 (0.29)	-0.2169 (-12.98)	-0.0561 (-3.91)	1, 2	1	1	<0.0001	0.26	0.0037
BBB	0.0013 (0.66)	-0.3778 (-6.84)	-0.1247 (-2.98)	-	-	-	-	0.14	-

Panel B: 01:1995-07:2001									
	Regression Coefficients			AR and GARCH Parameters				Goodness of Fit	
Index	*a*	*b*	*c*	*m*	*p*	*q*	Norm. Test	R^2	LM
AAA	N/A								
AA	-0.0002 (-0.08)	-0.3294 (-4.42)	-0.1266 (-2.69)	-	-	-	-	0.23	-
A	0.0006 (0.25)	-0.3502 (-4.83)	-0.1214 (-2.65)	-	-	-	-	0.25	-
BBB	0.0012 (0.24)	-0.2404 (-2.07)	-0.2398 (-3.49)	3	-	-	-	0.20	-

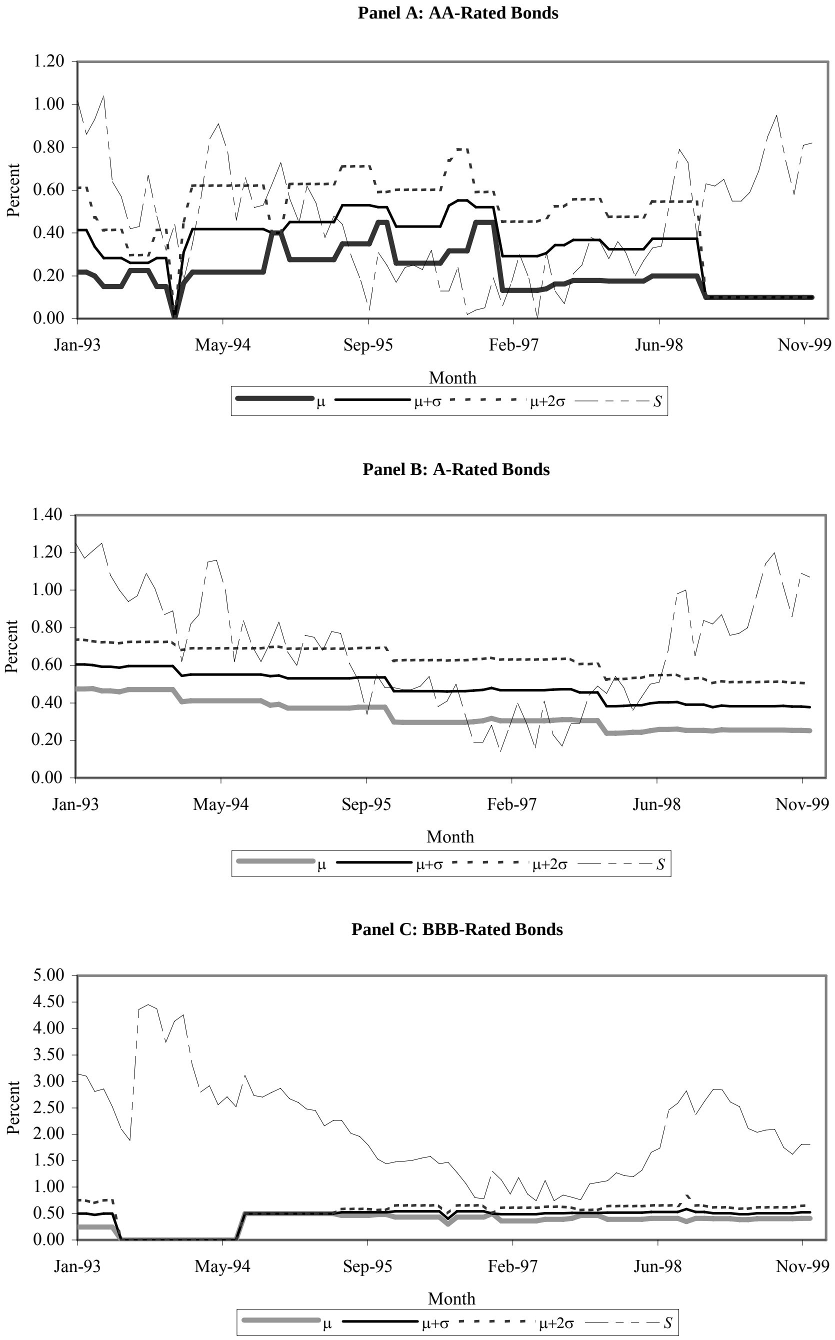

Figure 1. Moneyness of the doomsday call provision. To determine the moneyness of the doomsday call provision, we compare the doomsday spread to the yield spread. Based on a sample of doomsday spreads of long bonds for each month during the 01:1993-12:1999 period, for each rating category, we plot the yield spread on the corresponding long-term index (*S*), the average doomsday spread (μ), the average doomsday spread plus one standard deviation ($\mu+\sigma$), and the average doomsday spread plus two standard deviations ($\mu+2\sigma$).

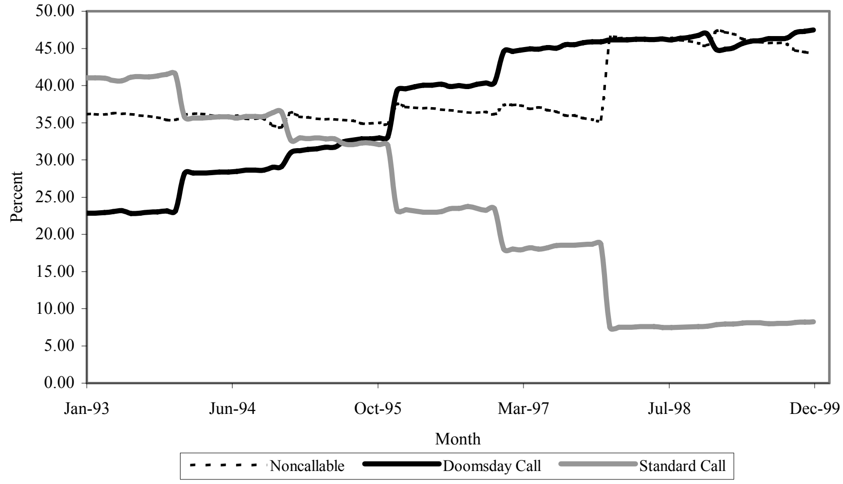

Figure 2. Distribution of Canadian Corporate Bonds Based on Callability (Percentages), 01:1993 - 12:1999. For each month during the sampled 7-year period we count the number of corporate bonds issued with a doomsday call provision, the number of bonds issued with a standard call provision, and the number of noncallable bonds, as reported in the *Financial Post* Corporate Bond Record. In the Figure we plot the proportion of each of the above categories calculated with respect to the total number of bonds within each month.

Endnotes

[1] For example see Merton, (1974) Black and Cox (1976) and Longstaff and Schwartz (1995) and Ericsson and Renault (forthcoming).

[2] Duffee (1998) notes that there are tax effects arising from the different tax rates, which apply for U.S. corporate and Treasury bonds. However, using U.S. bond data, he is unable to control for the impact of the tax differential.

[3] Batten, Hogan, and Jacoby (2005) use a sample of noncallable Australian Eurobonds and find evidence supporting the predictions of the two-factor Longstaff and Schwartz model. Nonetheless the preference by international investors for the with-holding tax exempt yields of Eurobonds may distort the pricing relationship between high credit quality Eurobonds and Government bonds.

[4] According to Duffee, following a given rise in Treasury yields, everything else being equal, the yield on a corporate bond will have to increase by a higher rate, so that the after-tax yield spread will remain unchanged. This implies that the pre-tax yield spread will widen following an increase in Treasury yields.

[5] The reader may refer to Jacoby and Roberts (2003) for a detailed description of the SCM indices.

[6] Mann and Powers (2003) refer to the make-whole call provision as "a recent innovation in the public debt markets."

[7] Note that there is still a small probability that a BBB-rated bond will be upgraded in the future and that its doomsday call will be in-the-money. However, given the small rate of upgrades, the expected value of this state is trivial.

[8] In a small number of cases, the cross-sectional standard deviation of the doomsday spread is zero, which implies: $\mu = \mu+\sigma = \mu+2\sigma$. This is due to a small number of bonds available during the given month, usually issued by the same company, all sharing the same doomsday spread. In other cases, specifically in the 06:1993-07:1994 period, the reported doomsday spread for BBB-rated bonds is zero.

[9] Note that since our sample is not stratified into different sectors, there is no need in the current study to use sector-specific stock indices as in Longstaff and Schwartz (1995).

[10] Longstaff and Schwartz (1995) estimate their regression models using OLS. As previously noted, we find that a combined autoregressive and GARCH (1,1) model fits our data best. However, the qualitative results turn out to be similar irrespective of the estimation procedure used. Thus, for the sake of comparison with Longstaff and Schwartz (1995), our focus in the discussion of regression models (1) and (2) is on the results of the OLS estimation, which are presented in the tables below. The results of the combined autoregressive and GARCH (1,1) regressions are reported in the appendix.

[11] The estimated coefficient b for the BBB index under the Yule-Walker estimation is also statistically insignificant. Thus, this result is robust.

[12] A higher probability of a call for corporate bonds carrying the doomsday call provision reduces their effective duration, or price sensitivity to changes in the riskless rate. This implies that following an upward shift in the riskless rate, the corporate bond yield will rise by a lower rate, and yield spreads will contract.

[13] Note that we obtain the same result for the BBB index when the Yule-Walker procedure is applied.

[14] In analyzing the results, note that one's conclusions are insensitive to whether one uses OLS or the estimation method we apply here.

[15] Our results are different from those in Collin-Dufresne et al. (2001). This may be due to the fact that their data, devoid of callability, still suffer from the coupon bias which may generate the observed negative relationship for noncallable bonds.

[16] We adopt the same approach as Jorion and Goetzmann (2000) Table 1.

[17] Interested readers can obtain more information on RRBs from Bank of Canada website.
[18] Moody's speculative-grade default rate forecasts are generated with a Poisson regression model, with the independent variables that proxy changes in credit quality, an aging effect (to reflect the changing nature of default risk with the time that elapsed since issuance), and macroeconomic variables. For details see Keenan, Sobehart, and Hamilton (1999). We thank David Hamilton for providing us the time series of the speculative-grade default rate forecast.