Interest Rates and Credit Spread Dynamics

Robert Neal
Indiana University

Douglas Rolph*
Nanyang Technological University
Nanyang Business School

Brice Dupoyet
Florida International University
College of Business Administration

Xiaoquan Jiang
Florida International University
College of Business Administration

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*Corresponding author. Nanyang Technological University, Nanyang Business School, S3-B2B-71, 50 Nanyang Avenue, Singapore 639798, Telephone: +65 6790 4806, email: drolph@ntu.edu.sg
Abstract

This paper revisits the relation between callable credit spreads and interest rates. We allow the evolution of credit spreads following shocks to government rates to depend on recent levels and changes in government rates. After conditioning on government yields, we find a negligible short-term response in callable corporate bond spreads to shocks in government rates. In contrast to existing studies, our results imply that after controlling for the prevailing interest rate environment, there is little evidence that variation in the call premium has an appreciable impact on credit spreads.

JEL Classification: G12; G18; E43

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1. Introduction

Recent empirical studies have found mixed evidence on the relation between risk-free interest rates and credit spreads, measured as the difference between corporate bond yields and equivalent maturity government bond yields. Structural models of corporate debt suggest that credit spreads should be negatively related to short maturity risk-free rates. Intuitively, an increase in the instantaneous risk-free rate shifts the risk-neutral growth rate of the firm’s assets upwards. Holding the risk of the firm fixed, this implies that credit spreads decrease with the (risk-neutral) probability of default.

In contrast, recent studies of the relation between credit spreads and interest rates find a significant relation only with callable corporate bonds (Duffee (1998), Elton et al. (2001), Davies (2008)). Tests that include only non-callable bonds find no relation (Duffee (1998), Jacoby et al. (2009)). Thus, the empirical evidence suggests that the call premium embedded in corporate bonds prices and yields drives the correlation between credit spreads and interest rates, and is an important determinant of the time variation in corporate bond spreads.

However, King (2002) finds that the call option value constitutes only around 2 percent of the par value of the average callable bond. Given the relatively small contribution of the call option feature to bond prices and yields, it seems unlikely that the time variation in the call option would be able to have such a powerful effect on the correlation between credit spreads and interest rates.

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1 Seminal papers include Merton (1974), Longstaff and Schwartz (1995), and Lehland and Toft (1994).
We revisit the relation between credit spreads and interest rates using new, powerful econometric techniques. In Lehman Brothers US corporate bond indices that include callable bonds, we find no statistically significant change in credit spreads after shocks to either short-term or long-term Treasury yields, either contemporaneously or up to three years after a shock. This is the case for corporate bond indices that include long maturity bonds and indices that include intermediate maturity bonds. As a robustness test, we also conduct the same analysis on a subset of Moody's corporate bond indices that include callable bonds. Again, we find no significant correlation between bond spreads and treasury yields.

Since we find no evidence of a significant relation in bond indices that include callable bonds, the evidence suggests that the call premium has little impact on the correlation between credit spreads and interest rates. This is important since it rules out the call premium as a factor determining the time variation in interest rates.

The key behind our results is our methodology. In contrast to the usual approach of examining credit spreads – the difference between corporate and government rates of similar maturities - we use cointegration to model the joint behavior of corporate and Treasury rates. A common approach in the literature is to regress contemporaneous credit spreads or changes in credit spreads on contemporaneous levels or changes in treasury rates. Interest rates and credit spreads, however, have a high degree of persistence. The error term of the regression is thus autocorrelated, correlated with the independent variable (interest rates) and contains information about contemporaneous interest rates. The estimates of the regression coefficients can then be inefficient, the corresponding forecasts suboptimal and the significance tests on the estimated coefficients invalid, as shown in Granger and Newbold (1974). Our approach is to let the data be
the guide by explicitly incorporating the influence of persistence through the modeling of the joint evolution of corporate and government interest rates using a cointegration framework.

We examine how credit spreads evolve after an unexpected change in government rates using a methodology that allows the path of credit spreads to depend on recent levels and changes in government rates. Numerous studies document that both levels and changes in government rates contain information about the current and future state of the macro economy. Furthermore, it is well known that credit spreads differ with economic conditions.\(^2\) By conditioning on the prevailing interest rate environment, we allow the correlation between credit spreads and interest rates to incorporate any information government rates may have about economic conditions. In contrast, most previous studies focus on the unconditional relation between credit spreads and government rates. Our results indicate that after controlling for the prevailing interest rate environment, there is little evidence that either short or long-term interest rate shocks have an appreciable impact on credit spreads.

Our results are interesting for several reasons. First, cointegration has implications for models of pricing corporate debt and credit derivatives. Cointegration supports the intuition that corporate and treasury rates are closely linked and cannot evolve in arbitrary ways. This linkage, however, is not captured in the parameterization of reduced form bond pricing models, structural models, and credit spread option-pricing models. The omission is important because Duan and Pliska (2004) show that, under reasonable conditions, ignoring cointegration will significantly bias the calculated price of spread options. Second, our finding that higher Treasury rates do not

\(^2\) For instance, see Fama and French (1989), Chan-Lau and Ivaschenko (2001, 2002), Gilchrist, Yankov and Zakrajsek (2009), and also the discussion in Mueller (2009).
have a statistically significant impact on credit spreads has implications for models that analyze credit spread dynamics. For example, the comparative statics of the capital structure models of Leland and Toft (1996) and Collin-Dufresne and Goldstein (2001), and the bond pricing models of Longstaff and Schwartz (1995), Kim, Ramaswamy, and Sundaresan (1993), and Merton (1974) all predict that higher rates will lower credit spreads. Our results suggest that there is little empirical support for this relation.

The remainder of the paper is as follows: Section 2 provides a summary of prior studies, Section 3 describes the data along with some summary statistics, Section 4 introduces the cointegration model and its associated empirical results, Section 5 presents the Generalized Impulse Response Functions methodology and its results, and Section 6 offers concluding remarks.

2. Literature review

Numerous papers in finance highlight the link between interest rates and credit spreads. From the theoretical perspective, interest rates should be linked to credit spreads through the effect that changes in interest rates have on the credit quality of firms; the impact may be either through firm-specific or macroeconomic channels. In the empirical literature, the evidence generally suggests a negative link between changes in government rates and credit spreads changes, although a number of studies question the robustness of the link. In this section, we review two strands of literature that are closely related to our study. The first strand briefly discusses the theoretical framework that leads to the articulation of factors affecting credit spreads
and interest rates in the context of structural bond models. The second strand focuses on the empirical linkage between interest rates and credit spreads, and the factors that may magnify or lessen the link.

2.1. Theoretical link

Credit spreads are correlated with interest rates through the impact of changes in interest rates on the credit conditions of corporations. Most theoretical studies consider the correlation in the context of risk-neutral valuation models of corporate debt, and focus on the effect interest rates have on the growth in firm value. The structural default models build on the original insights of Black and Scholes (1973) who demonstrate that equity and debt can be valued using contingent-claims analysis. Introduced by Merton (1974), structural models specify how firm value evolves over time and assume that default is triggered when firm value falls below some threshold. The default threshold is a function of the amount of debt outstanding. Since the (risk-neutral) growth of the firm increases with the instantaneous risk-free rate, the likelihood that the firm value falls below the face value decreases and credit risk premiums decline, thereby inducing a negative correlation between credit spreads and interest rates.

Several models also allow for a positive correlation between credit spreads and interest rates. Longstaff and Schwartz (1995) allow a positive relationship between spreads and interest rates. Longstaff and Schwartz (1995) allow a positive relationship between spreads and interest

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rates if interest rates are positively correlated with asset-value volatility or if shocks to cash flows are negatively correlated with interest rates shocks. The alternative dynamic capital structure models of Fischer, Heinkel, and Zechner (1989) and of Goldstein, Ju, and Leland (2001), the banking model of Duan, Moreau, and Sealey (1995) and the option pricing model of Lesseig and Stock (1998) all demonstrate a positive relationship in the long run.

2.2. Empirical relation between interest rates and credit spreads

Our focus lies in the link between credit spreads and interest rates, as well as other factors that may amplify or lessen the link. Applying the closed-form solution of Merton (1974), Longstaff and Schwartz (1995) provide evidence of a strong negative relation between changes in credit spreads and interest rates. In the Longstaff and Schwartz (1995) two-factor valuation 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. The negative correlation between interest rates and credit spreads persists after controlling for stock returns, which serve as a proxy for the growth rate of the firm’s assets.

Collin-Dufresne, Goldstein, and Martin (2001) find a negative relationship between interest rates and spreads after controlling for both firm and market-level determinants of default risk. However, they show that default risk can only explain 25% of the credit spread, and claim
that changes in spreads are mainly driven by local supply and demand shocks that are independent of both credit-risk factors and standard proxies for liquidity. In a related study using Australian Eurobonds, 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 asset growth 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). Avramov, Jostova and Philipov (2007) also find a negative relation between the levels or changes in spreads, and like Collin-Dufresne, Goldstein, and Martin (2001), show that the results hold after controlling for aggregate and company-level variables. In contrast to Collin-Dufresne, Goldstein, and Martin (2001), however, they argue that a majority of the variation in credit spreads is explained by standard proxies for default risk. Elton, Gruber, Agrawal and Mann (2001) provide evidence that the majority of the spread unaccounted for by taxes and expected default is related to compensation for bearing systematic risk.

Although the negative relationship between spreads and interest rates is consistent with theoretical structural models, the empirical literature finds mixed evidence. Jones, Mason and Rosenfeld (1984) and Eom, Helwege and Huang (2004) have shown that the empirical performance of these models is poor. Schaefer and Strebulaev (2008) show that the relation between credit spread and interest rate is much weaker than the theoretical models predict. Using bid and offer quotes on credit default swaps, Ericsson, Jacobs and Oviedo (2009) find that a large amount of credit spread variation cannot be explained by theoretical determinants such as interest rates. Campbell and Taksler (2003) show a negative relation between credit spreads and interest
rates after controlling for idiosyncratic firm-level risk; they also show that idiosyncratic risk can explain as much cross-sectional variation in yields as credit ratings can.

Empirical studies (Altman (1968), Wilson (1997) and Collin-Dufresne, Goldstein, and Martin (2001)) also find that default premia vary with business conditions. Wu and Zhang (2008) use a dynamic factor model to summarize the information in many observed macroeconomic and financial data series and quantify their impact on the term structure of Treasury yields and credit spreads. They find that the link between interest rates and credit spreads is at least partially determined by systematic covariation with inflation, and that shocks to inflation raise the term structure of both risk-free rates and credit spreads. Davies (2008), however, shows that the positive relation between spreads and interest rates is robust after controlling for inflation.

While the general consensus in the literature points to a negative link between credit spreads and government rates, the call feature of corporate debt has the potential to induce a source of common variation in credit spreads and interest rates that is unrelated to default risk. Duffee (1998) points out that the negative relation found by Longstaff and Schwartz (1995) owes to the construction of the 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 accept a lower yield for these call provisions, which will result in an overall decrease in the bond yield spread.

To accommodate the call feature present in most U.S. corporate bonds, Duffee (1998) 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 spreads and interest rates is much weaker once the call option effects are removed from the data. The negative relationship found by Longstaff and Schwartz (1995) thus seems attributable to either the default premium or the call premium. Consistent with this evidence, after controlling for callability, Jacoby, Liao, and Batten (2009) find that the relation between corporate spreads and interest rates is negligible using Canadian bonds data devoid of tax effects. In our study, we use Lehman corporate bond indices that include callable bonds. We find no relation between credit spreads and interest rates, which suggests that the call feature of corporate debt does not appear to play any role in terms of biasing the results.

Corporate bonds tend to trade less frequently than Treasury issues; a number of studies document the impact of liquidity on corporate bond prices and yields. Driessen (2005) provides evidence of a liquidity component in corporate bond spreads using the Duffie and Singleton (1999) reduced-form pricing approach. Chacko (2005) and Downing, Underwood and Xing (2005) explore the role of liquidity in the corporate bond market and find that liquidity risk is a priced factor in bond returns. De Jong and Driessen (2006) show that fluctuations in Treasury bond and equity market liquidity affect corporate bond returns. Chen, Lesmond and Wei (2007) find that liquidity is priced in corporate yield spreads: more illiquid bonds earn higher yield spreads and an improvement in liquidity causes a significant reduction in yield spreads, suggesting that neither the level nor the dynamics of yield spreads can be fully explained by default risk determinants. Ericsson and Renault (2006), however, find a negative relation between interest rates and credit spreads after controlling for liquidity risk in the Treasury bond market.
3. Data and summary statistics

3.1. Data description

We obtain monthly corporate bond yields from the Lehman Brothers U.S. Corporate Index and monthly constant-maturity government rates from the Federal Reserve’s H.15 release. The Lehman Brothers U.S. Corporate database begins in 1973 and our study therefore spans from February 1973 to December 2007. The Lehman Brothers Corporate Indices include all publicly traded U.S. corporate debentures and secured notes that meet prescribed maturity, liquidity, and quality guidelines. Securities with calls, puts and sinking fund provisions are included. However, the index excludes private placements, 144A securities, floating rate securities, and Eurobonds. In addition, to be included in the index, a bond must have an outstanding $250 million minimum face value; must be rated investment grade (Baa/BBB- or above) using the middle rating from Moody’s, S&P, and Fitch respectively; and must have at least one year left until final maturity. All the bonds included are SEC-registered, fully taxable issues.

The Lehman Brothers U.S. Corporate Index is also categorized into eight sub-indices based on respective credit ratings (AAA, AA, A and BAA) and maturity. Intermediate indices include bonds with less than ten years to maturity, while long-term indices have bonds with ten or more years to maturity. Our motivation for looking at intermediate and long term corporate bond rates has to do with the slope and curvature found in term structure models. Incorporating various maturities in the system may allow the model to capture some of the variation in both corporate and government rates unexplained by lagged long-term rates. We select the 10-year and the 3-
year government bonds as the Treasury counterparts to the long-term and intermediate-term corporate bonds.

3.2. Summary statistics

Table 1 contains summary statistics for corporate rates, Treasury rates, changes in corporate rates and changes in Treasury rates for both long-term and intermediate-term maturities. Over the 1973-2007 period, the 10-year government rate averaged 7.59%, AAA rates averaged 8.52%, AA rates averaged 8.75%, A rates averaged 9.03% and BAA rates averaged 9.62%. During the same period, the 3-year government rate averaged 7.07%, AAA rates averaged 7.81%, AA rates averaged 8.01%, A rates averaged 8.33% and BAA rates averaged 8.99%. The mean monthly changes in rates are close to zero for each series. The (unreported) mean long-term spreads for long-term AAA, AA, A and BAA rates over the 10-year government rate were 0.94%, 1.16%, 1.44% and 2.04% respectively, while the mean long-term spreads for intermediate-term AAA, AA, A and BAA rates over the 3-year government rate were 0.74%, 0.94%, 1.26% and 1.91% respectively.

Panel A of Table 2 presents autocorrelation coefficients for the 10-year Treasury and long-term corporate rates, while Panel B provides the autocorrelation levels for the 3-year Treasury and intermediate-term corporate rates. For long-term corporate rates, the autocorrelation coefficients for the first four lags exceed 0.95, while for intermediate-term rates, the autocorrelation coefficients for the first four lags exceed 0.93. The high degree of persistence is consistent with the presence of a unit root. Table 3 reports the augmented Dickey-Fuller and Phillips-Perron unit
root tests. Using between one and six lags, these two tests fail to reject at the 5% level the presence of a unit root for both long-term and intermediate-term corporate and government rates. In addition, the Dickey-Fuller and Phillips-Perron tests for the first differences (not reported) are significant at the 1% level. Thus, the levels of the interest rates appear nonstationary while the changes appear stationary. Overall, these results are consistent with the conclusions of a number of studies on unit roots in nominal interest rates.\(^4\)

One objection to this conclusion is that interest rates may follow a highly persistent, but stationary, time series process. The Dickey-Fuller type tests assume the underlying process is nonstationary and it is well known that the tests have low power against near unit root alternatives. To address this possibility, we also apply the unit root test derived by Kwiatkowski, Phillips, Schmidt and Shin (1992). This test derived under the null that the process is stationary has higher power for detecting stationarity than either the Phillip-Perron or Dickey-Fuller tests. Our inference is however unchanged: for each series, the null hypothesis that rates are stationary is rejected at a 1% confidence level. Following Granger and Swanson (1996) and Phillips (1998), it should also be noted that there are advantages to using cointegration techniques even for near unit-root processes.

(2003), this raises the possibility that evidence of nonstationary rates could be a consequence of fitting linear models, such as a Dickey-Fuller model, to a nonlinear process. While this remains a concern, other studies question the inference of a nonlinear drift process. Chapman and Pearson (2000) use simulations to show that the finite sample properties of the Aït-Sahalia and Stanton estimators frequently detect nonlinearities even when none are present. Jones (2003) shows that the evidence of nonlinearity and mean reversion is strongly related to the sampling interval. These effects are evident with daily observations, but largely disappear with monthly observations. This suggests that our analysis, which uses monthly data, should be less affected by any nonlinear reversion.

4. **Cointegration and the unconditional relation between corporate and Treasury rates**

4.1. The cointegration model

In this section we provide a cointegration framework to analyze the relation between corporate and Treasury bond yields. Cointegration is based on the idea that while a set of variables are individually nonstationary, a linear combination of the variables might be stationary. The stationary combination arises from a long-run statistical relationship that links the cointegrated variables together. Cointegration also implies that the short-term movements of the variables will be affected by the lagged deviation from the long-run relationship between the variables. This induces mean reversion around the long-run relationship, but in the absence of economic restrictions, there is no causality in a cointegration model. An innovation in any one variable affects the other variables in the system.
An alternative view of cointegration is that two variables are cointegrated when both are driven by the same unit root process. If corporate rates can be modeled as the sum of the nonstationary Treasury rate and a risk premium, it is clear that both the Treasury and the corporate rates share a common process. Since the two rates are driven by the same stochastic trend, they cannot evolve independently and the levels of the variables will be linked together.

While cointegration is intuitively appealing, it assumes that the underlying variables are nonstationary. We impose the assumption of unit root processes not because we believe that interest rates can exhibit unbounded variation, but because it provides the distribution theory that best represents the finite sample properties of our data. For example, the asymptotic distribution theory for standard test statistics, such as t-statistics, has a discontinuity at the point when the time series stops being stationary and starts being nonstationary. Since our finite sample statistics indicate nonstationary interest rate processes, we choose to base our inference on a distribution theory that is well defined for nonstationary processes. Our view is consistent with Granger and Swanson (1996) and Phillips (1998) who show that nonstationary distribution models provide superior inference for both unit root and near-unit root processes.

To present cointegration formally, consider the vector representation $X_t = \mu_t + \varepsilon_t$, where $X_t = \{X_{1t}, X_{2t}\}$ represents two data vectors, $\mu_t = \{\mu_{1t}, \mu_{2t}\}$ represents two stochastic trends, and $\varepsilon_t = \{\varepsilon_{1t}, \varepsilon_{2t}\}$ represents two i.i.d. error terms. If there is a stationary linear combination of the two variables, then there exists a (2x2) non-zero matrix $B$ such that $B\mu_t = 0$. The test for cointegration is therefore based on the rank of $B$. In the two-variable case, there can be at most one independent linear combination of $X_{1t}$ and $X_{2t}$ that is stationary. In this case, if the rank of $B$ equals one, then the variables $(X_{1t}, X_{2t})$ are said to be cointegrated.
Assuming $\Delta X_t$ is stationary, the short-term dynamics of two cointegrated variables are captured in an error-correction model:

\[
\begin{align*}
\Delta X_{1,t} &= a_{01} + \gamma_1 (X_{1,t-1} - \lambda X_{2,t-1}) + \sum_{i=1}^{p} a_{i,11} \Delta X_{1,t-i} + \sum_{i=1}^{p} a_{i,12} \Delta X_{2,t-i} + \epsilon_{1,t} \\
\Delta X_{2,t} &= a_{02} + \gamma_2 (X_{1,t-1} - \lambda X_{2,t-1}) + \sum_{i=1}^{p} a_{i,21} \Delta X_{1,t-i} + \sum_{i=1}^{p} a_{i,22} \Delta X_{2,t-i} + \epsilon_{2,t}
\end{align*}
\]

(1)

(2)

In this model, the cointegration vector is said to be $(1,-\lambda)$, and the linear combination $X_{1,t} - \lambda X_{2,t}$ is stationary. The economic interpretation of $X_{1,t} - \lambda X_{2,t}$ is that it represents the deviation from the long-run relationship between $X_1$ and $X_2$. In the error-correction model, this deviation affects the short-term behavior of $\Delta X_t$, with the error-correction coefficients, $\gamma_1$ and $\gamma_2$ describing how quickly $X_1$ and $X_2$ respond to the deviation. Of a somewhat similar yet simpler model linking short rates to asset prices through a system of two equations, Rigobon and Sack (2004) state “This model is clearly an oversimplification of the relationship between movements in interest rates and asset prices. It imposes no structure that might arise from an asset-pricing model. However, this is also an advantage, as it allows the interaction between the variables to be fairly unrestricted.”

It is well known that the presence of cointegration between $X_1$ and $X_2$ causes the time series behavior of $X$ to differ from that of a conventional vector autoregression. Equations (1) and (2) can be written in matrix form as

\[
\Delta X_t = A_0 + \Pi X_{t-1} + A_1 \Delta X_{t-1} + ... + A_k \Delta X_{t-p} + \epsilon_t
\]

(3)

where $A_0$ is a $(1 \times 2)$ vector of intercepts and $A_1 \ldots A_k$ are $(2 \times 2)$ matrices of coefficients on lagged $\Delta X$. The important characteristic distinguishing cointegration models from VAR models is
whether $\Pi = 0$. If this restriction holds, then $\Delta X_t$ can be represented by a conventional VAR in differences. However, if the rank of $\Pi$ exceeds zero, the elements of $\Pi$ are non-zero. In this case, the series are cointegrated and the lagged $X$ variable should be included in the regression. The conventional approach, which omits the lagged levels of $X$, can generate misleading inferences because it neglects the long-run relation between the integrated variables.

The tests for cointegration involve estimating the rank of $\Pi$. For a $(n \times 1)$ vector of $I(1)$ variables, $X_t$, the cointegration model can be written as

$$ \Delta X_t = A_0 + \Pi \Delta X_{t-1} + A(L) \Delta X_{t-1} + \varepsilon_t $$

(4)

where $A(L)$ is a $p$-th order matrix polynomial in the lag operator and $\varepsilon_t$ is a vector of i.i.d. error terms. For purposes of comparing across models, we fix the number of lags (in differences) to one. We consider other forms as a robustness check where we select variable lags with Akaike’s information criterion, and our empirical results remain the same.

Johansen (1988) shows that the number of cointegrating vectors, $k$, equals the rank of $\Pi$. He provides two likelihood ratio tests for determining the rank of $\Pi$, based on the number of nonzero eigenvalues in $\Pi$. The first test, the maximal eigenvalue test, is really a sequence of tests. After sorting the estimated eigenvalues of $\Pi$ in descending order, the $k$-th statistic provides a test of the null hypothesis that the rank of $\Pi$ equals $k$ against the alternative that the rank of $\Pi$ equals $k+1$. The second test statistic, the trace statistic, is the running sum of the maximal eigenvalue statistics. The $k$-th trace statistic provides a test of the null hypothesis that the rank of $\Pi$ equals $k$. 
Π is less than or equal to \( k \) against the alternative that the rank of \( \Pi \) is strictly greater than \( k \).

Critical values for these test statistics are provided in Osterwald-Lenum (1992).

An attractive feature of the cointegration framework is that it allows one to distinguish between short-run and long-run behavior. We estimate the models with a two-stage procedure that first identifies the cointegration vector, and then includes the vector in a second-stage regression of changes in corporate rates on changes in treasury rates.

4.2. The unconditional relation

Table 4 reports tests of cointegration between corporate and government bond yields: the results indicate that corporate and government rates are cointegrated. The first set of trace statistics examines whether long-term corporate rates (Panel A) and intermediate-term rates (Panel B) share a common unit root with the corresponding government rates. All tests are statistically significant at the 5% level. Thus, we reject the null hypothesis that corporate and government yields are not cointegrated in favor of the alternative hypothesis that there is at least one cointegrating vector. The second trace statistics, however, do not support the existence of two cointegrating vectors for either the Intermediate-Term Bond series or the Long-Term Bond series as one cannot reject the null hypothesis of one cointegrating vector or less for any of the series. The results for all series are based on using two lags of the data. The lag length was determined by the Schwartz Criterion. Given the existence of cointegration between the Long-Term Bond series and the 10-year Treasury bond series as well as between the Intermediate-Term
Bond series and the 3-year Treasury bond series, Table 5 reports the corresponding cointegrating vectors.

The results in Table 4 support the view that corporate rates and Treasury bond yields are cointegrated. The cointegration evidence found in Table 4 has important implications regarding the testing of the dynamic relation between credit spreads and interest rates. The literature is generally aware of the nonstationary nature of the two series, and usually follows the standard approach of removing the deterministic linear trends by taking first differences. This method, however, ignores cointegration if it exists, and can bias the testing of the relation between credit spreads and interest rates (Campbell and Shiller (1987)).

Given the existence of cointegration between Lehman long-term (intermediate-term) bond and Treasury bond (10-year and three-year) series, Table 5 reports the corresponding cointegrating vectors. For Lehman long-term bonds and 10-year Treasury bonds, the AAA vector is (1, -1.1054), AA vector is (1, -1.1381), A vector is (1, -1.1741), and the BAA vector is (1, -1.2508). For Lehman intermediate-term bond and 3-year Treasury bond, the AAA vector is (1, -1.0852), AA vector is (1, -1.1159), A vector is (1, -1.1599), and the BAA vector is (1, -1.2480). The p-values for both panels are less than 1%. The result in Table 5 has two interesting implications. First, since all \( \lambda \) values exceed one, a 1% increase in Treasury rates ultimately generates increases in corporate rates of more than 1%. Thus, as interest rates rise, credit spreads will eventually widen. This finding goes against the theoretical framework of Merton (1974), Longstaff and Schwartz (1995) and Leland and Toft (1996) but supports previous empirical work as reported by Morris, Neal, and Rolph (1999) and Davies (2008). Second, the lower quality bonds exhibit a greater long-run sensitivity to interest rate movements than higher
quality bonds. This is inconsistent with a commonly held view that increased credit risk will make corporate bonds less interest rate sensitive.

An alternative way to interpret the cointegrating relationship is to estimate the error-correction regressions found in equations (1) and (2). Cointegration implies that the coefficient on the error-correction term will be negative and significant, with the size of the coefficient measuring the sensitivity of corporate rates to the error-correction term. The negative sign indicates that credit spreads subsequently adjust to restore the long-run equilibrium when a deviation occurs. Using the estimated cointegrating vectors from Table 5, Table 6 presents estimates of the error-correction model. As expected, the error-correction coefficients are significantly negative. For the AAA, AA, A and BAA Long-Term Corporate rates, the error-correction coefficient terms are -0.0653, -0.0578, -0.0545 and -0.0511 respectively. For the AAA, AA, A and BAA Intermediate-Term Corporate rates, the error-correction coefficient terms are -.1042, -0.1027, -0.0892 and -0.0619 respectively. Our results indicate that cointegration plays an important role in explaining the dynamic relation between credit spreads and interest rates.

Davies (2008) also applies a Vector Error Correction Model (VECM) to examine the dynamic relation between credit spreads and interest rates. It is worthy of note that our unconditional results on both short-term and long-term relations between credit spreads and interest rates are consistent with the ones in Davies (2008), in spite of the fact that the VECM specifications are different.
5. Generalized Impulse Response Functions and the conditional relation between credit spreads and Treasury rates

The above results, consistent with Davies (2008), show the unconditional relation between credit spreads and Treasury rates. However, conditioning on the current interest rates environment is important and should not be ignored, as not doing so would be the equivalent of discarding potentially valuable information. Intuitively, if, for instance, interest rates are fairly high around the peak of a business cycle and start declining as the economy begins contracting, credit risk in aggregate is likely to go up and therefore credit spreads would be expected to increase. However, if interest rates are fairly low near, but prior to, the trough of a business cycle and decrease a bit further before the economy begins expanding again, credit risk in aggregate is likely to start declining in the near future and therefore credit spreads would be likely to decrease. Conversely, if interest rates are fairly low around the trough of a business cycle and start increasing as the economy begins expanding, credit risk in aggregate is likely to go down and therefore credit spreads would be expected to decrease. However, if interest rates are fairly high near, but prior to, the peak of a business cycle and increase a bit further before the economy begins contracting again, credit risk in aggregate is likely to start increasing in the near future and therefore credit spreads would be likely to increase. An interest rate increase or decrease thus does not necessarily have the same implication for future credit spreads depending on the current level of interest rates and the state of the economy. Using the estimates from the second-stage regression, along with the cointegrating vector, we thus construct various impulse response functions that measure the impact of a shock to government rates on corporate and Treasury rates while conditioning on prevailing interest rate conditions.
To further examine the conditional relation between corporate spread and interest rates, we apply the Generalized Impulse response Function (GIRF) of Koop et al. (1996) that corrects for the issues associated with traditional impulse response functions and allows for asymmetric responses to positive and negative interest rate shocks. When using traditional impulse response functions, however, the history of the process or the sequence of observations as well as the signs, sizes, and correlations of the shocks occurring between the initial shock and the impulse horizon can produce misleading estimates, as demonstrated by Pesaran and Potter (1997).

Moreover, traditional impulse functions analysis assumes the residuals in the VEC model are jointly normally distributed. With this assumption, and given that the VEC model is linear, the response of spreads to a negative shock is the mirror of the response of spreads to a positive shock. There may however be reason to believe that responses to shocks could be different depending on whether the shocks are positive or negative. In the GIRF framework, the residuals observed from VEC model are bootstrapped, and the standard errors reflect the empirical (not assumed) residual distribution. Separating the GIRF functions into positive and negative shocks potentially increases the power of the test to reject the null hypothesis of no response. If we segregate shocks into positive and negative cases, and we think that positive and negative shocks have the potential to mean different things in terms of the business conditions and their relation to credit risk, then we have increased the power to identify a possibly significant relation.

Finally, traditional impulse response functions’ standard errors are a ceteris paribus argument using the parameters of the model and neglect any effect of the response of the spreads to government rates shocks given current economic conditions as reflected in recent interest rates. GIRF functions construct forecasts for each period during the sample using the most recent
interest rates, and then average over the resulting forecast. Thus, there is more variability in the forecasts, leading to wider and more realistic standard errors that take the fat tails of the empirical distribution into account.

5.1. Generalized Impulse Response Functions methodology

The generalized impulse response functions are defined as the difference in month $t+i$ between the conditional expectation of the corporate bond spread (corporate yield – treasury yield) following a large shock to Treasury yields at month $t$, and the conditional expectation across all possible shocks to Treasury yields. Both expectations condition on interest rates at $t-1$.

We define “large shocks” as residuals from the government rate equation in the VECM that exceed two standard deviations of government rate residuals, computed using the full sample. We consider positive and negative shocks separately in order to control for possible asymmetric effects of positive and negative shocks to government rates on credit spreads. When positive and negative shocks are combined, all results are quantitatively similar to those reported below. The response to the shock is calculated as an average of the possible outcomes based on the present and the past, with future shocks averaged out in the expectation. The expectation is taken with respect to a conditioning information set made up of the history of the entire system that includes the past shocks to the system.

The GIRF equations can be written as:

$$GIRF(n, \varepsilon_{g,t-1}, \Omega_{t-1}) = E(y_{t+n} | \Omega_{t-1}, \varepsilon_{g,t} > 2\sigma_{g}) - E(y_{t+n} | \Omega_{t-1}) \quad \text{when } \varepsilon_{g,t} > 0$$ (5)
and as

\[
GIRF(n, \varepsilon_{g,t}, \Omega_{t-1}) = E(y_{t+n} \mid \Omega_{t-1} \mid |\varepsilon_{g,t}| > 2\sigma_e) - E(y_{t+n} \mid \Omega_{t-1}) \quad \text{when } \varepsilon_{g,t} < 0
\]  

(6)

where \(y_{t+n}\) is the corporate credit spread at time \(t+n\), \(\Omega_{t-1}\) is the time \(t-1\) information set used to produce forecasts of \(y_{t+n}\), \(\varepsilon_{g,t} > 2\sigma_e\) represents a positive shock exceeding two standard errors, and \(|\varepsilon_{g,t}| > 2\sigma_e\) represents a negative shock exceeding two standard errors.

We implement the Generalized Impulse Response Functions methodology pioneered in Koop et al. (1996), using the bootstrapped residuals from the Vector Error Correction model described in Equations (1) and (2). In explaining the procedure in detail below, we use the single government rate example where

\[
X_t = \begin{bmatrix}
X_{1t} \\
X_{2t}
\end{bmatrix} = \begin{bmatrix}
\text{corporate rate at } t \\
\text{government rate at } t
\end{bmatrix}, \quad \Lambda = \begin{bmatrix}
1 & -\lambda
\end{bmatrix}
\]  

(7)

and

\[
A_2 = \begin{bmatrix}
\gamma_1 \\
\gamma_2
\end{bmatrix}
\]  

so that \(\Pi = A_2 \otimes \Lambda\)

(8)

(where \(\otimes\) represents the Kronecker product)

Multiple government rate cases are generalizations of the two-variable case.

1. Retrieve the fitted/realized residuals from the Vector Error Correction model for each month \(t\) using actual data. In (2×2) matrix and (2×1) vector notation, for month \(t\), use real data at month \(t-1\) and month \(t-2\) and compute

\[
\Delta X_{t}^{\text{forecasted}} = A_1 \Delta X_{t-1} + A_2 \otimes \Lambda X_{t-1}
\]

where \(\Delta X_{t-1} = X_{t-1} - X_{t-2}\)

(9)
in order to retrieve the \((2\times1)\) fitted residual vector \(\varepsilon_t = \Delta X_t - \Delta X_t^{\text{forecasted}}\)

2. For each month \(t\), calculate “future” interest rates at month \(t+n\) (for \(n=0,1,\ldots,N\)) by bootstrapping the model residuals obtained in step 1.

   - For month \(t\), use real data at month \(t-1\) and month \(t-2\) as well as residual \(\varepsilon_t\), and compute
     \[
     \Delta X_t^{\text{unconditional forecast}} = A_1 \Delta X_{t-1} + A_2 \otimes A X_{t-1} + \varepsilon_t
     \]
     and obtain
     \[
     X_t^{\text{unconditional forecast}} = X_{t-1} + \Delta X_t^{\text{unconditional forecast}}
     \]
   
   - For month \(t+1\), use \(X_t^{\text{unconditional forecast}}\), real data at month \(t-1\) and residual \(\varepsilon_{t+1}\), and compute
     \[
     \Delta X_{t+1}^{\text{unconditional forecast}} = A_1 (X_t^{\text{unconditional forecast}} - X_{t-1}) + A_2 \otimes A X_t^{\text{unconditional forecast}} + \varepsilon_{t+1}
     \]
     and obtain
     \[
     X_{t+1}^{\text{unconditional forecast}} = X_t^{\text{unconditional forecast}} + \Delta X_{t+1}^{\text{unconditional forecast}}
     \]

   - For month \(t+n, n>1\), use \(X_{t+n-1}^{\text{unconditional forecast}}\), \(X_{t+n-2}^{\text{unconditional forecast}}\) and residual \(\varepsilon_{t+n}\), and compute
     \[
     \Delta X_{t+n}^{\text{unconditional forecast}} = A_1 (X_{t+n-1}^{\text{unconditional forecast}} - X_{t+n-2}^{\text{unconditional forecast}}) + A_2 \otimes A X_{t+n-1}^{\text{unconditional forecast}} + \varepsilon_{t+n}
     \]
     and obtain
\[ X_{t+n}^{\text{unconditional forecast}} = X_{t+n-1}^{\text{unconditional forecast}} + \Delta X_{t+n}^{\text{unconditional forecast}} \] (15)

These forecasted interest rates are used to calculate the unconditional expectation of interest rate spreads.

3. For each month \( t \), calculate “future” interest rates at time \( t+n \) for \( n=0,1,...,N \) by imposing that the value of the first residual at time \( t \) be larger than two standard deviations (for positive shocks). The remaining residual series is kept exactly the same as in step 2. The only difference between step 2 and step 3 is that in step 3 the first residual is drawn from the subset of residuals that are larger than two times the residuals’ standard deviation. In other words:

- For month \( t \), use real data at month \( t-1 \) and month \( t-2 \) and residual \( \varepsilon_t \) with \( \varepsilon_t > 2\sigma_{\varepsilon} \), and compute

\[ \Delta X_t^{\text{conditional forecast}} = A_1 \Delta X_{t-1}^{\text{conditional forecast}} + A_2 \otimes \Lambda X_{t-1}^{\text{conditional forecast}} + \varepsilon_t \quad | \varepsilon_t > 2\sigma_{\varepsilon} \] (16)

and obtain

\[ X_t^{\text{conditional forecast}} = X_{t-1} + \Delta X_t^{\text{conditional forecast}} \] (17)

- For month \( t+1 \), use \( X_t^{\text{conditional forecast}} \), real data at month \( t-1 \) and residual \( \varepsilon_{t+1} \), and compute

\[ \Delta X_{t+1}^{\text{conditional forecast}} = A_1 (X_t^{\text{conditional forecast}} - X_{t-1}^{\text{conditional forecast}}) + A_2 \otimes \Lambda X_t^{\text{conditional forecast}} + \varepsilon_{t+1} \] (18)

and obtain
\[ X_{t+1}^{\text{conditional forecast}} = X_t^{\text{conditional forecast}} + \Delta X_{t+1}^{\text{conditional forecast}} \]

- For month \( t+n, n > 1 \), use \( X_{t+n-1}^{\text{conditional forecast}}, X_{t+n-2}^{\text{conditional forecast}} \) and residual \( \varepsilon_{t+n} \), and compute

\[ \Delta X_{t+n}^{\text{conditional forecast}} = A_1 (X_{t+n-1}^{\text{conditional forecast}} - X_{t+n-2}^{\text{conditional forecast}}) \\
+ A_2 \otimes \Lambda X_{t+n-1}^{\text{conditional forecast}} + \varepsilon_{t+n} \]

and obtain

\[ X_{t+n}^{\text{conditional forecast}} = X_{t+n-1}^{\text{conditional forecast}} + \Delta X_{t+n}^{\text{conditional forecast}} \]

These forecasted interest rates are used to calculate the expectation of interest rate spreads conditional on a large shock (positive or negative).

4. Record the corporate spread for both the unconditional and the conditional simulation separately, and repeat steps 2 and 3 for the chosen number of simulations (1,000 in this case).

5. For a given number of steps ahead \( n \), calculate the differences between these 1,000 simulated conditional and unconditional spreads and calculate their means, lower bounds and upper bounds (2.5% and 97.5% respectively).

6. For each number of steps ahead \( n \), compute averages of the means, lower bounds and upper bounds across all \( t \) (months).
5.2. The conditional relation between credit spreads and Treasury rates

Figures 1 through 4 plot the response of long-term corporate spreads to a two-standard-deviation shock in the 10-year treasury rate. The figures plot the average response of credit spreads, as well as the 2.5\textsuperscript{th} and 97.5\textsuperscript{th} percentiles of the GIRF functions. In order to examine the long-run response, we plot the path of credit spreads for the five years that follow the shock to the government rate. We do this for both positive and negative shocks. As mentioned earlier, segregating the GIRF functions into positive and negative shocks potentially increases the power of the test to reject the null hypothesis of no response. For inclusiveness, we also combine the positive and negative shocks, and find similar results.

As can been seen in the figures, there is little significance in the short-term reaction of the spreads to either positive or negative shocks to long government rates. After a large positive (negative) shock, the corporate spread decreases (increases) initially but subsequently reverts to near pre-shock levels. And most importantly, this temporary average deviation from initial levels is generally not statistically significantly different from zero, as shown by the width of the confidence bounds. While most empirical studies focus on the response of interest rates after a shock of one standard deviation, we focus on two-standard-deviation shocks to government rates. The choice of using a two-standard-deviation threshold is a deliberate one, made to ensure that the conclusion that government rates changes do not really have much of a statistically significant impact on credit spreads is unambiguous, since the results are obtained even when Treasury yields experience an unusually large shock. In unreported results, we also find no significant relation between credit spreads and one standard deviation shocks.
Previous studies have shown that the relation between credit spreads and interest rates is negative, particularly for callable bonds (e.g., Longstaff and Schwartz, 1995; and Duffee, 1998; Jacoby, Liao, and Batten, 2009). Based on a reduced-form VAR and traditional impulse response functions, Duffee (1998) shows that the dynamic relation between credit spreads and interest rates is significantly negative in the short run (a couple of months), but insignificant in the long run. This pattern is particularly strong for callable bonds. Jacoby, Liao, and Batten (2009) find that the relation between corporate spreads and interest rates is negligible using Canadian bonds data that exclude callability and are devoid of tax effects. Taking into account the potential cointegration between credit spreads and interest rates and applying the GIRF methodology, we find, however, that the dynamic relation is not significant in the short run, even with callable bonds. In addition, there are little differences in responses to government interest rate shocks for varying credit ratings, as can be seen by the similarities among the credit spreads graphs ranging from AAA to BAA in quality.

Figures 5 through 8 plot the response of intermediate-term corporate spreads to a two-standard-deviation shock in the 3-year treasury rate. The figures plot the average response of credit spreads, as well as the 2.5th and 97.5th percentiles of the GIRF functions. We plot the path of credit spreads for the five years that follow the shock to the government rate, and as in the long-term rate cases, the temporary average deviations from initial levels are generally not statistically significantly different from zero, as shown by the width of the confidence bounds, except for some very marginal statistical significance in the long run.

The novelty of our result comes from the fact that, unlike prior studies that test the relationship between government rates and credit spreads on an unconditional basis only, we also
test the conditional relation between government rates and credit spreads by conditioning on the prevailing interest rate environment. Both levels and changes in government rates containing information about the current and future state of the economy, it is paramount to allow the correlation between credit spreads and interest rates to incorporate information that Treasury rates might have about economic conditions.

For completeness, we also estimate yet do not report generalized impulse responses for one-year rates and find very marginal short-run significance. As a robustness check, we also estimate the VECM that also incorporates either short (3-month) or intermediate (3-year) government rates. A number of structural models focus on the relation between the instantaneous risk-free rate and credit spreads. In the VECM that only includes long-term government rates, the 10-year Treasury rate serves as a proxy for the instantaneous risk-free rate. When we also incorporate short or intermediate government rates, we again find no significant relation between credit spreads and shocks to government rates, regardless of the maturity of the government rate. Once the prevailing interest rate environment is controlled for, it thus becomes clear that neither short-term nor long-term interest rate shocks seem to have much of a statistically significant impact on credit spreads.
6. Conclusion

The relation between credit spreads and interest rates is at most mixed. Prior studies have focused on their unconditional relation. We provide new insights on the relation between corporate spreads and interest rates. We contribute to the literature by first using cointegration to model the joint behavior of corporate and Treasury rates and then using the results of an Error Correction Model to test the conditional relation between credit spreads and interest rates. The Vector Error Correction model used along with Generalized Impulse Response Functions allows us to provide new insights into the relation between credit spreads and interest rates. When conditioning on the current interest rate environment, we find that the short-run relation between credit spreads and interest rates is no longer significant although it is negative, and that the long-run relation is weakly positive. Since the weakly positive long-run relation is obtained after an interest rate shock larger than two standard deviations, these results imply that for most shocks the long-run relation between credit spreads and interest rates will be statistically negligible.
References


Table 1

Descriptive Statistics

The statistics are based on monthly data from 1973:2 to 2007:12. In Panel A, the AAA, AA, A and BAA series are Lehman Brothers Long-Term Corporate Yields and the 10-year Treasury series is a constant maturity series from the Board of Governors. In Panel B, the AAA, AA, A and BAA series are Lehman Brothers Intermediate-Term Corporate Yields and the 3-year Treasury series is a constant maturity series from the Board of Governors.

Panel A: Lehman Long-Term Bond Series and 10-year Treasury rates

<table>
<thead>
<tr>
<th>Variables</th>
<th>Mean</th>
<th>Std. Dev.</th>
<th>Percentile (10%)</th>
<th>Percentile (90%)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Lehman AAA Long</td>
<td>8.52</td>
<td>2.33</td>
<td>5.83</td>
<td>12.21</td>
</tr>
<tr>
<td>Lehman AA Long</td>
<td>8.75</td>
<td>2.43</td>
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<td>12.52</td>
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<tr>
<td>Lehman A Long</td>
<td>9.03</td>
<td>2.47</td>
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<td>12.83</td>
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<tr>
<td>Lehman BAA Long</td>
<td>9.62</td>
<td>2.61</td>
<td>6.65</td>
<td>13.46</td>
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<tr>
<td>10-Year Treasury</td>
<td>7.59</td>
<td>2.61</td>
<td>4.50</td>
<td>11.67</td>
</tr>
<tr>
<td>Δ Lehman AAA Long</td>
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<td>0.339</td>
<td>-0.340</td>
<td>0.360</td>
</tr>
<tr>
<td>Δ Lehman AA Long</td>
<td>-0.003</td>
<td>0.331</td>
<td>-0.340</td>
<td>0.340</td>
</tr>
<tr>
<td>Δ Lehman A Long</td>
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<td>0.342</td>
<td>-0.350</td>
<td>0.350</td>
</tr>
<tr>
<td>Δ Lehman BAA Long</td>
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<td>0.372</td>
<td>-0.380</td>
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<td>Δ 10-Year Treasury</td>
<td>-0.006</td>
<td>0.372</td>
<td>-0.390</td>
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Panel B: Lehman Intermediate-Term Bond Series and 3-year Treasury rates

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<tr>
<th>Variables</th>
<th>Mean</th>
<th>Std. Dev.</th>
<th>Percentile (10%)</th>
<th>Percentile (90%)</th>
</tr>
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<tr>
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<td>2.83</td>
<td>5.21</td>
<td>12.29</td>
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<tr>
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<td>2.90</td>
<td>5.71</td>
<td>13.02</td>
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<td>3-Year Treasury</td>
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<td>0.401</td>
<td>-0.401</td>
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<td>Δ 3-Year Treasury</td>
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<td>-0.500</td>
<td>0.500</td>
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</table>
Table 2

Sample Autocorrelations

The estimates are based on monthly data from 1973:2 to 2007:12. In Panel A, the AAA, AA, A and BAA series are Lehman Brothers Long-Term Corporate Yields and the 10-year Treasury series is a constant maturity series from the Board of Governors. In Panel B, the AAA, AA, A and BAA series are Lehman Brothers Intermediate-Term Corporate Yields and the 3-year Treasury series is a constant maturity series from the Board of Governors. The Box-Ljung Q-Statistic test the null hypothesis that the series is not serially correlated. This statistic is distributed $\chi^2(n)$, where $n$ is the number of lags. The null hypothesis is rejected at a significance level of less than 0.1% for all lags.

Panel A: Lehman Long-Term Bond Series and 10-year Treasury rates

<table>
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<tr>
<th>Variables</th>
<th>Autocorrelation Coefficient</th>
<th>Q-Statistic</th>
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<tr>
<td>Lehman AAA Long</td>
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<tr>
<td>Lehman AA Long</td>
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</tr>
<tr>
<td>Lehman A Long</td>
<td>0.99</td>
<td>0.98</td>
</tr>
<tr>
<td>Lehman BAA Long</td>
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<td>0.97</td>
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<tr>
<td>10-Year Treasury</td>
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<td>0.97</td>
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Panel B: Lehman Intermediate-Term Bond Series and 3-year Treasury rates

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<td>Lehman AAA Intermediate</td>
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<td>Lehman AA Intermediate</td>
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<tr>
<td>Lehman A Intermediate</td>
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<td>0.97</td>
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<tr>
<td>Lehman BAA Intermediate</td>
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<td>0.97</td>
</tr>
<tr>
<td>3-Year Treasury</td>
<td>0.98</td>
<td>0.96</td>
</tr>
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</table>
Table 3

**Unit Root Tests for Levels of Interest Rates**

The null hypothesis for the Dickey-Fuller and the Phillips-Perron test is that the series contains a unit root. The percentage p-values (in parentheses) are approximate asymptotic p-values calculated using the method described in MacKinnon (1991). The estimates are based on monthly data from 1973:2 to 2007:12. In Panel A, the AAA, AA, A and BAA series are Lehman Brothers Long-Term Corporate Yields and the 10-year Treasury series is a constant maturity series from the Board of Governors. In Panel B, the AAA, AA, A and BAA series are Lehman Brothers Intermediate-Term Corporate Yields and the 3-year Treasury series is a constant maturity series from the Board of Governors.

Panel A: Lehman Long-Term Bond Series and 10-year Treasury rates

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<th>Variables</th>
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<th>Phillips-Perron</th>
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<td>n=1 n=2 n=3 n=4 n=5 n=6</td>
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<td>-1.34 -1.33 -1.31 -1.27 -1.27 -1.28</td>
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<tr>
<td></td>
<td>(0.58) (0.63) (0.66) (0.72) (0.66) (0.65)</td>
<td>(0.61) (0.61) (0.62) (0.64) (0.65) (0.64)</td>
</tr>
<tr>
<td>Lehman AA Long</td>
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<td>-1.35 -1.35 -1.34 -1.32 -1.32 -1.34</td>
</tr>
<tr>
<td></td>
<td>(0.55) (0.63) (0.62) (0.68) (0.61) (0.62)</td>
<td>(0.61) (0.61) (0.61) (0.62) (0.62) (0.61)</td>
</tr>
<tr>
<td>Lehman A Long</td>
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<tr>
<td></td>
<td>(0.56) (0.62) (0.61) (0.69) (0.63) (0.65)</td>
<td>(0.61) (0.61) (0.61) (0.62) (0.62) (0.62)</td>
</tr>
<tr>
<td>Lehman BAA Long</td>
<td>-1.49 -1.39 -1.39 -1.29 -1.41 -1.37</td>
<td>-1.39 -1.40 -1.40 -1.39 -1.40 -1.41</td>
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<tr>
<td></td>
<td>(0.54) (0.59) (0.59) (0.63) (0.58) (0.60)</td>
<td>(0.59) (0.58) (0.58) (0.59) (0.58) (0.58)</td>
</tr>
<tr>
<td>10-Year Treasury</td>
<td>-1.39 -1.24 -1.16 -1.17 -1.19 -1.23</td>
<td>-1.27 -1.27 -1.25 -1.23 -1.24 -1.24</td>
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<td>(0.59) (0.66) (0.69) (0.69) (0.68) (0.66)</td>
<td>(0.64) (0.64) (0.65) (0.66) (0.66) (0.66)</td>
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</tbody>
</table>

Panel B: Lehman Intermediate-Term Bond Series and 3-year Treasury rates

<table>
<thead>
<tr>
<th>Variables</th>
<th>Dickey-Fuller</th>
<th>Phillips-Perron</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>n=1 n=2 n=3 n=4 n=5 n=6</td>
<td>n=1 n=2 n=3 n=4 n=5 n=6</td>
</tr>
<tr>
<td>Lehman AAA Intermediate</td>
<td>-1.61 -1.45 -1.38 -1.27 -1.46 -1.28</td>
<td>-1.47 -1.47 -1.45 -1.42 -1.43 -1.43</td>
</tr>
<tr>
<td></td>
<td>(0.48) (0.56) (0.59) (0.64) (0.55) (0.64)</td>
<td>(0.55) (0.55) (0.55) (0.57) (0.57) (0.57)</td>
</tr>
<tr>
<td>Lehman AA Intermediate</td>
<td>-1.54 -1.42 -1.39 -1.24 -1.42 -1.26</td>
<td>-1.42 -1.43 -1.42 -1.39 -1.40 -1.40</td>
</tr>
<tr>
<td></td>
<td>(0.51) (0.57) (0.59) (0.66) (0.57) (0.65)</td>
<td>(0.57) (0.57) (0.57) (0.59) (0.58) (0.58)</td>
</tr>
<tr>
<td>Lehman A Intermediate</td>
<td>-1.52 -1.50 -1.51 -1.27 -1.46 -1.26</td>
<td>-1.42 -1.45 -1.47 -1.44 -1.45 -1.44</td>
</tr>
<tr>
<td></td>
<td>(0.53) (0.53) (0.53) (0.64) (0.55) (0.65)</td>
<td>(0.57) (0.56) (0.55) (0.56) (0.56) (0.56)</td>
</tr>
<tr>
<td>Lehman BAA Intermediate</td>
<td>-1.61 -1.61 -1.57 -1.36 -1.43 -1.25</td>
<td>-1.46 -1.51 -1.53 -1.52 -1.51 -1.48</td>
</tr>
<tr>
<td></td>
<td>(0.48) (0.48) (0.50) (0.60) (0.57) (0.65)</td>
<td>(0.55) (0.53) (0.52) (0.53) (0.53) (0.54)</td>
</tr>
<tr>
<td>3-Year Treasury</td>
<td>-1.76 -1.51 -1.40 -1.33 -1.43 -1.34</td>
<td>-1.57 -1.57 -1.53 -1.48 -1.47 -1.46</td>
</tr>
<tr>
<td></td>
<td>(0.40) (0.53) (0.58) (0.62) (0.57) (0.61)</td>
<td>(0.50) (0.50) (0.52) (0.54) (0.55) (0.55)</td>
</tr>
</tbody>
</table>
### Table 4

**Cointegration Results**

This table uses Johansen’s (1988) maximum likelihood method to estimate the rank of $\Pi$ for the corporate and government rates in the two-variable regression. In Panel A the corporate rates are the Lehman Long-Term Bond series ranging from AAA to BAA, and the government rate is the 10-year constant maturity Treasury series from the Board of Governors. In Panel B the corporate rates are the Lehman Intermediate-Term Bond series ranging from AAA to BAA, and the government rate is the 3-year constant maturity Treasury series from the Board of Governors. The estimates are based on monthly data from 1973:2 to 2007:12.

**Panel A: Lehman Long-Term Bond Series and 10-year Treasury rates**

<table>
<thead>
<tr>
<th>Variables</th>
<th>Maximum Rank</th>
<th>Trace Statistic</th>
<th>5% Critical Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Lehman AAA Long</td>
<td>0</td>
<td>18.35</td>
<td>12.53</td>
</tr>
<tr>
<td></td>
<td>1</td>
<td>0.46</td>
<td>3.84</td>
</tr>
<tr>
<td>Lehman AA Long</td>
<td>0</td>
<td>16.03</td>
<td>12.53</td>
</tr>
<tr>
<td></td>
<td>1</td>
<td>0.44</td>
<td>3.84</td>
</tr>
<tr>
<td>Lehman A Long</td>
<td>0</td>
<td>15.79</td>
<td>12.53</td>
</tr>
<tr>
<td></td>
<td>1</td>
<td>0.44</td>
<td>3.84</td>
</tr>
<tr>
<td>Lehman BAA Long</td>
<td>0</td>
<td>14.98</td>
<td>12.530</td>
</tr>
<tr>
<td></td>
<td>1</td>
<td>0.47</td>
<td>3.84</td>
</tr>
</tbody>
</table>

**Panel B: Lehman Intermediate-Term Bond Series and 3-year Treasury rates**

<table>
<thead>
<tr>
<th>Variables</th>
<th>Maximum Rank</th>
<th>Trace Statistic</th>
<th>5% Critical Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Lehman AAA Intermediate</td>
<td>0</td>
<td>33.71</td>
<td>12.53</td>
</tr>
<tr>
<td></td>
<td>1</td>
<td>0.64</td>
<td>3.84</td>
</tr>
<tr>
<td>Lehman AA Intermediate</td>
<td>0</td>
<td>34.63</td>
<td>12.53</td>
</tr>
<tr>
<td></td>
<td>1</td>
<td>0.61</td>
<td>3.84</td>
</tr>
<tr>
<td>Lehman A Intermediate</td>
<td>0</td>
<td>33.09</td>
<td>12.53</td>
</tr>
<tr>
<td></td>
<td>1</td>
<td>0.67</td>
<td>3.84</td>
</tr>
<tr>
<td>Lehman BAA Intermediate</td>
<td>0</td>
<td>24.52</td>
<td>12.530</td>
</tr>
<tr>
<td></td>
<td>1</td>
<td>0.71</td>
<td>3.84</td>
</tr>
</tbody>
</table>
Table 5

Estimates of Cointegrating Vectors

This table reports estimates of $\lambda$ in the cointegrating vector $(1, -\lambda)$ for the corporate and government rates using Johansen’s (1988) maximum likelihood method. In Panel A the corporate rates are the Lehman Long-Term Bond series ranging from AAA to BAA, and the government rate is the 10-year constant maturity Treasury series from the Board of Governors. In Panel B the corporate rates are the Lehman Intermediate-Term Bond series ranging from AAA to BAA, and the government rate is the 3-year constant maturity Treasury series from the Board of Governors. The estimates are based on monthly data from 1973:2 to 2007:12. The Johansen normalization restrictions are imposed and we report the coefficient estimates as well as the respective p-values.

Panel A: Lehman Long-Term Bond Series and 10-year Treasury rates

<table>
<thead>
<tr>
<th>Variables</th>
<th>Coefficient</th>
<th>Estimate</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Lehman AAA Long</td>
<td>1</td>
<td>1</td>
<td></td>
</tr>
<tr>
<td>$\lambda$</td>
<td>1.1054</td>
<td></td>
<td>0.00</td>
</tr>
<tr>
<td>Lehman AA Long</td>
<td>1</td>
<td>1</td>
<td></td>
</tr>
<tr>
<td>$\lambda$</td>
<td>1.1381</td>
<td></td>
<td>0.00</td>
</tr>
<tr>
<td>Lehman A Long</td>
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<td>1</td>
<td></td>
</tr>
<tr>
<td>$\lambda$</td>
<td>1.1741</td>
<td></td>
<td>0.00</td>
</tr>
<tr>
<td>Lehman BAA Long</td>
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<td>1</td>
<td></td>
</tr>
<tr>
<td>$\lambda$</td>
<td>1.2508</td>
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<td>0.00</td>
</tr>
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</table>

Panel B: Lehman Intermediate-Term Bond Series and 3-year Treasury rates

<table>
<thead>
<tr>
<th>Variables</th>
<th>Coefficient</th>
<th>Estimate</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Lehman AAA Intermediate</td>
<td>1</td>
<td>1</td>
<td></td>
</tr>
<tr>
<td>$\lambda$</td>
<td>1.0852</td>
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<td>0.00</td>
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<tr>
<td>Lehman AA Intermediate</td>
<td>1</td>
<td>1</td>
<td></td>
</tr>
<tr>
<td>$\lambda$</td>
<td>1.1159</td>
<td></td>
<td>0.00</td>
</tr>
<tr>
<td>Lehman A Intermediate</td>
<td>1</td>
<td>1</td>
<td></td>
</tr>
<tr>
<td>$\lambda$</td>
<td>1.1599</td>
<td></td>
<td>0.00</td>
</tr>
<tr>
<td>Lehman BAA Intermediate</td>
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<td>1</td>
<td></td>
</tr>
<tr>
<td>$\lambda$</td>
<td>1.2480</td>
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<td>0.00</td>
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</table>
Table 6

Estimates of the Error-Correction Model

This table reports equation (1)’s estimates along with standard errors in parentheses for the bivariate error-correction system of equations (1) and (2) in the one-lag difference case. The error-correction terms are estimated using Johansen’s (1988) maximum likelihood method, $X_1$ alternatively represents the various corporate rates, and $X_2$ alternatively represents the various government rates. In Panel A the corporate rates are the Lehman Long-Term Bond series ranging from AAA to BAA, and the government rate is the 10-year constant maturity Treasury series from the Board of Governors. In Panel B the corporate rates are the Lehman Intermediate-Term Bond series ranging from AAA to BAA, and the government rate is the 3-year constant maturity Treasury series from the Board of Governors.

Panel A: Lehman Long-Term Bond Series and 10-year Treasury rates

<table>
<thead>
<tr>
<th>Dependent Variable</th>
<th>Independent Variables</th>
<th>Error-Correction</th>
<th>$\Delta X_{1,t-1}$</th>
<th>$\Delta X_{2,t-1} = \Delta$ 10-Year Treasury$_{t-1}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta X_{1,t} = \Delta$ Lehman AAA Long$_{t}$</td>
<td></td>
<td>-0.0653</td>
<td>-0.1262</td>
<td>0.6038</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.0166)</td>
<td>(0.0354)</td>
<td>(0.0356)</td>
</tr>
<tr>
<td>$\Delta X_{1,t} = \Delta$ Lehman AA Long$_{t}$</td>
<td></td>
<td>-0.0578</td>
<td>-0.0971</td>
<td>0.6103</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.0159)</td>
<td>(0.0348)</td>
<td>(0.0343)</td>
</tr>
<tr>
<td>$\Delta X_{1,t} = \Delta$ Lehman A Long$_{t}$</td>
<td></td>
<td>-0.0545</td>
<td>-0.1206</td>
<td>0.5931</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.0155)</td>
<td>(0.0376)</td>
<td>(0.0374)</td>
</tr>
<tr>
<td>$\Delta X_{1,t} = \Delta$ Lehman BAA Long$_{t}$</td>
<td></td>
<td>-0.0511</td>
<td>-0.0809</td>
<td>0.612</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.0143)</td>
<td>(0.0384)</td>
<td>(0.0410)</td>
</tr>
</tbody>
</table>

Panel B: Lehman Intermediate-Term Bond Series and 3-year Treasury rates

<table>
<thead>
<tr>
<th>Dependent Variable</th>
<th>Independent Variables</th>
<th>Error-Correction</th>
<th>$\Delta X_{1,t-1}$</th>
<th>$\Delta X_{2,t-1} = \Delta$ 3-Year Treasury$_{t-1}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta X_{1,t} = \Delta$ Lehman AAA Intermediate$_{t}$</td>
<td></td>
<td>-0.1042</td>
<td>-0.0702</td>
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</tr>
<tr>
<td></td>
<td></td>
<td>(0.0189)</td>
<td>(0.0303)</td>
<td>(0.0306)</td>
</tr>
<tr>
<td>$\Delta X_{1,t} = \Delta$ Lehman AA Intermediate$_{t}$</td>
<td></td>
<td>-0.1027</td>
<td>-0.0765</td>
<td>0.5599</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.0185)</td>
<td>(0.0309)</td>
<td>(0.0309)</td>
</tr>
<tr>
<td>$\Delta X_{1,t} = \Delta$ Lehman A Intermediate$_{t}$</td>
<td></td>
<td>-0.0892</td>
<td>-0.1192</td>
<td>0.5211</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.0166)</td>
<td>(0.0355)</td>
<td>(0.0334)</td>
</tr>
<tr>
<td>$\Delta X_{1,t} = \Delta$ Lehman BAA Intermediate$_{t}$</td>
<td></td>
<td>-0.0619</td>
<td>-0.0319</td>
<td>0.5111</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.0131)</td>
<td>(0.0372)</td>
<td>(0.0348)</td>
</tr>
</tbody>
</table>
Figure 1: This figure shows the response of Lehman AAA Long spreads to a positive (figure 1a) and negative (figure 1b) two-standard-deviation shock in the 10-year government rate. The graph is obtained by constructing generalized impulse response functions (GIRF) that capture both the short-term dynamics and the long-run relation between corporate and Treasury rates. The generalized impulse response functions are defined as the difference in month $t+i$ between the conditional expectation of the corporate bond spread (corporate yield – treasury yield) following a large shock to Treasury yields at month $t$, and the conditional expectation across all possible shocks to Treasury yields. Both expectations condition on interest rates at $t-1$. 
Figure 2: Response of Lehman AA Long Spread to shock in 10-year Treasury rate

Figure 2a. Figure 2b.

Figure 2: This figure shows the response of Lehman AA Long spreads to a positive (figure 2a) and negative (figure 2b) two-standard-deviation shock in the 10-year government rate. The graph is obtained by constructing generalized impulse response functions (GIRF) that capture both the short-term dynamics and the long-run relation between corporate and Treasury rates. The generalized impulse response functions are defined as the difference in month $t+i$ between the conditional expectation of the corporate bond spread (corporate yield – treasury yield) following a large shock to Treasury yields at month $t$, and the conditional expectation across all possible shocks to Treasury yields. Both expectations condition on interest rates at $t-1$. 


Figure 3: Response of Lehman A Long Spread to shock in 10-year Treasury rate

Figure 3a. Figure 3b.

Figure 3: This figure shows the response of Lehman A Long spreads to a positive (figure 3a) and negative (figure 3b) two-standard-deviation shock in the 10-year government rate. The graph is obtained by constructing generalized impulse response functions (GIRF) that capture both the short-term dynamics and the long-run relation between corporate and Treasury rates. The generalized impulse response functions are defined as the difference in month \(t+i\) between the conditional expectation of the corporate bond spread (corporate yield – treasury yield) following a large shock to Treasury yields at month \(t\), and the conditional expectation across all possible shocks to Treasury yields. Both expectations condition on interest rates at \(t-1\).
Figure 4: This figure shows the response of Lehman BAA Long spreads to a positive (figure 4a) and negative (figure 4b) two-standard-deviation shock in the 10-year government rate. The graph is obtained by constructing generalized impulse response functions (GIRF) that capture both the short-term dynamics and the long-run relation between corporate and Treasury rates. The generalized impulse response functions are defined as the difference in month $t+i$ between the conditional expectation of the corporate bond spread (corporate yield – treasury yield) following a large shock to Treasury yields at month $t$, and the conditional expectation across all possible shocks to Treasury yields. Both expectations condition on interest rates at $t-1$. 
Figure 5: Response of Lehman AAA Intermediate Spread to shock in 3-year Treasury rate

Figure 5a.  Figure 5b.

Figure 5: This figure shows the response of Lehman AAA Intermediate spreads to a positive (figure 5a) and negative (figure 5b) two-standard-deviation shock in the 3-year government rate. The graph is obtained by constructing generalized impulse response functions (GIRF) that capture both the short-term dynamics and the long-run relation between corporate and Treasury rates. The generalized impulse response functions are defined as the difference in month $t+i$ between the conditional expectation of the corporate bond spread (corporate yield – treasury yield) following a large shock to Treasury yields at month $t$, and the conditional expectation across all possible shocks to Treasury yields. Both expectations condition on interest rates at $t-I$. 
Figure 6: Response of Lehman AA Intermediate Spread to shock in 3-year Treasury rate

Figure 6a. 

Figure 6b. 

Figure 6: This figure shows the response of Lehman AA Intermediate spreads to a positive (figure 6a) and negative (figure 6b) two-standard-deviation shock in the 3-year government rate. The graph is obtained by constructing generalized impulse response functions (GIRF) that capture both the short-term dynamics and the long-run relation between corporate and Treasury rates. The generalized impulse response functions are defined as the difference in month $t+i$ between the conditional expectation of the corporate bond spread (corporate yield – treasury yield) following a large shock to Treasury yields at month $t$, and the conditional expectation across all possible shocks to Treasury yields. Both expectations condition on interest rates at $t-1$. 
Figure 7: Response of Lehman A Intermediate Spread to shock in 3-year Treasury rate

Figure 7a.  Figure 7b.

Figure 7: This figure shows the response of Lehman A Intermediate spreads to a positive (figure 7a) and negative (figure 7b) two-standard-deviation shock in the 3-year government rate. The graph is obtained by constructing generalized impulse response functions (GIRF) that capture both the short-term dynamics and the long-run relation between corporate and Treasury rates. The generalized impulse response functions are defined as the difference in month $t+i$ between the conditional expectation of the corporate bond spread (corporate yield – treasury yield) following a large shock to Treasury yields at month $t$, and the conditional expectation across all possible shocks to Treasury yields. Both expectations condition on interest rates at $t-1$. 
Figure 8: Response of Lehman BAA Intermediate Spread to shock in 3-year Treasury rate

Figure 8a.

Figure 8b.

Figure 8: This figure shows the response of Lehman BAA Intermediate spreads to a positive (figure 8a) and negative (figure 8b) two-standard-deviation shock in the 3-year government rate. The graph is obtained by constructing generalized impulse response functions (GIRF) that capture both the short-term dynamics and the long-run relation between corporate and Treasury rates. The generalized impulse response functions are defined as the difference in month $t+i$ between the conditional expectation of the corporate bond spread (corporate yield – treasury yield) following a large shock to Treasury yields at month $t$, and the conditional expectation across all possible shocks to Treasury yields. Both expectations condition on interest rates at $t-1$. 