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Factors affecting the yields on noninvestment grade bond indices: a cointegration analysis

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Abstract

This study examines the long- and short-run dynamics of the yields on noninvestment grade indices. Utilizing cointegration techniques, the traditional yield spread model is found to be inadequate. A revised model finds a long-run relationship between noninvestment grade yields, Treasury securities, and default rates. Error correction models are formulated to model the short-run dynamics of different segments of the market. These models include a long-run equilibrium (between yields, default rates, and Treasuries), mutual fund flows, minor bond ratings, debt subordination measures, a stock index, and a January effect. Segmentation in the noninvestment grade market is also demonstrated. © 2000 Elsevier Science B.V. All rights reserved.

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1. Introduction

A great deal of research has focused upon the pricing of corporate debt (see Fisher, 1959; Silvers, 1973; and Boardman and McEnally, 1981). This research led to the segmentation of pricing factors into three broad categories: interest rate risk, default risk, and liquidity risk. However, given the constraints of sample size in comparing the characteristics of specific bonds, a second line of research devel-

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oped, which examined the effect of risk factors on bonds grouped by risk categories (comparable to an analysis using panel data). This study extends the analysis of the risk factors priced in risky debt by utilizing improved estimation methodologies, incorporating and testing the effect of interest-rate risk, including additional proxies for default and marketability risk, and by segmenting the market by bond rating category to improve the differentiation of risk across bond rating categories.

Fons (1987) and Altman and Bencivenga (1995) analyzed the market yield premium for holding risky debt (the average yield spread between risky debt securities and the risk-free security). This risk neutral type analysis calculates whether there is a net return (yield premium minus default rate) for holding risky bonds over a long period. This research has shown that as the risk of the bond increases, the market yield premium also increases, resulting in a positive risk-adjusted return.

Yield premium models are long-run models, which focus upon the default risk of holding a noninvestment grade security. Fridson and Jonsson (1995) and Garman and Fridson (1996) extend this type of analysis by focusing upon the short-run dynamics of the market by including liquidity risk measures in the analysis and by more broadly defining default risk. The authors formulated yield spread models (yield on the risky debt minus the risk-free rate as the dependent variable), which included both default risk and liquidity risk measures.

The current study extends the yield premium and yield spread models. First, the long-run relationship implied in the yield premium model is estimated using cointegration analysis. Second, the models formulated in this study combine both a long-run relationship and short-run dynamic components to determine the yield requirement for holding noninvestment grade bonds. Third, the yield spread studies have used a Merrill Lynch aggregate index of the noninvestment grade market to assess the effect of default and liquidity risk on the yield of corporate debt. The correlation analysis in Table 1 demonstrates the inherent problem of relying on a single broad default category. Risky debt is not homogeneous. Thus, to better explain the effect of default, interest-rate, and liquidity risk on corporate bonds, this study segmented the market by bond rating category. Rating category serves as a proxy for default risk, seniority position, securitization of the bond, industry specific risk, and other factors.

To better differentiate the risks of holding corporate bonds, the study has focused upon the riskiest segment of the corporate bond market, noninvestment grade debt. CS First Boston compiles indices of the high-yield market. For the current study, CS First Boston Indices were used for bonds rated BB and B by both Moody's and Standard and Poor's and an aggregate index of high-yield bonds rated Split BBB and lower. [1]

[1] The CS First Boston Composite High Yield Index (CSFB HYI) is composed of approximately 46% BB, 48% B, and 6% CCC rated bonds.

Table 1
Descriptive analysis of Corporate Bonds and Treasury Bonds
Bond Index information for Treasury Bond and Investment Grade Indices were collected from Citibase, which utilizes Moody's bond ratings. Noninvestment Grade Indices were provided by CS First Boston. All Index information is monthly from 1987(1) to 1996(7).

Bond index	Mean yield	Std. deviation	Correlation Treasury Bonds
T-Bonds	7.55	1.11	1.00
Aaa	8.52	0.96	0.973
Aa	8.75	0.98	0.968
A	8.96	1.04	0.966
Baa	9.41	1.13	0.959
Split BBB	9.59	1.33	0.903
BB	10.54	1.40	0.862
Split BB	11.70	1.91	0.810
B	12.67	2.17	0.745

Given the problem of pricing of bonds with relatively limited trading, monthly data from January 1987 till July 1996 was utilized in this study. Unfortunately, there is no reliable historical daily or weekly data on the high-yield market.

The paper is organized in the following manner. Section 2 examines previous empirical and theoretical research suggesting a number of explanatory variables, which could affect the risk of holding high yield securities. The possible explanatory variables are broken down into three broad risk categories: default risk measures, interest rate risk measures, and liquidity risk measures. Section 3 presents the cointegration analysis, which identifies long-run equilibrium relationships. Section 4 formulates error correction models using a general to specific modeling approach. Section 5 presents the overall conclusions on factors affecting the yields on high-yield bond indices.

2. Previous research on factors affecting the pricing of risky debt

2.1. Default risk measures

2.1.1. Default rates and economic indicators

Fridson and Jonsson (1995) found Moody's trailing-12-month default rate and an index of lagging economic indicators to have a statistically significant effect on yield spreads. An index of leading economic indicators was found to have no statistical significance. A note should be made about the default rate measure utilized. The market anticipates default well in advance of the actual default (see Altman, 1989), and hence some lag occurs between the market reaction and the default rate. The current study tests the significance of Moody's trailing-12-month default rates and Citibase leading and lagging economic indicators.

2.1.2. Equity index

A number of authors have demonstrated the correlation of returns on high-yield bonds to equity indices (see Bookstaber and Jacob, 1986; Ramaswami, 1991; and Shane, 1994). Such a relationship is consistent with the Black and Scholes model (Black and Scholes, 1973) of firm capital structure which, is sometimes termed a contingent claims analysis (CCA). The bondholders' payoff is the value of the bonds (on the upside) or the value of the firm on the downside. In this framework, the closer the value of the bonds is to the total firm value (high leverage), the more highly correlated changes in bond value and changes in equity value will become. The greater the positive difference in the value of the firm compared to the value of the bonds (low leverage), the more highly correlated changes in bond value and changes in risk free bond values will become. This theory is supported by the observed correlation structure of returns on investment and non-investment grade debt.

Unfortunately, an equity index is not available for firms, which have noninvestment grade debt outstanding. Instead, a correlation analysis was performed on a number of stock indices to determine the best index to utilize in the current study. [2] The Russell 2000 Index had the highest correlation across the noninvestment grade indices, and hence, was utilized in the study. In addition, the price/earnings ratio was also utilized in the study as a proxy for the cost of equity minus the growth rate. The price/earnings ratio of the Russell 2000 Index was unavailable for the historical time frame examined, so the Standard and Poor price/earnings ratio was utilized. Finally, a dummy variable (d.v.) was added to reflect the October 1987, market crash to determine its effect on the yield requirement of high-yield bonds. The stock market crash would be expected to have a positive effect on yield.

2.1.3. Subordinated debt

Black and Cox (1976) and Smith and Warner (1979) used a contingent claims analysis to theoretically prove that senior debt should be priced higher (lower yield) than similar subordinated debt. Carty and Lieberman (1996) demonstrated that the seniority of a bond has a significant impact on the bond's recovery rates in default. However, Fridson (1995) found that senior debt yields more than like-rated subordinated debt. Fridson's finding at first seems anomalous, but rating agencies factor into account a bond's potential recovery rate, which is a function of seniority, when assigning a rating. Hence, Fridson's finding suggests that bondholders, as compared to the rating agencies, place more emphasis on the potential of default as compared to the potential recovery in the case of default. Given the theoretical and empirical work in the area, it is hypothesized that a change in the

[2] We ran correlation analyses on the yields on the BB, B and CSFB Composite Indices and returns on a the S&P 500, S&P MidCap, NASDAQ Composite, Russell 1000, Russell 2000, and Russell 3000.

percentage of securities in an index, which are subordinated could have an effect on its yield. The percentage of subordinated bonds outstanding could also serve as a proxy for the average credit quality of the market. The current study uses COMPUSTAT to determine the monthly percentage of subordinated BB and B rated bonds. [3]

2.1.4. Minor rating classification

Previous research has found bond ratings significantly lag market changes in value (see Hettenhouse and Sartoris, 1976; Wansley and Terrence, 1985; and Ederington et al., 1987). However, the information content of minor rating classification (+ or − for S&P's and 1, 2, or 3 for Moody's) has not been analyzed. If the minor rating classifications are significant, a change in the percentage of the major rating category which is made up of the lowest credit quality (BB − and B −) would have an effect on the index's yield. COMPUSTAT [4] was used to determine the percentage of the major rating category segmented into the three different minor rating categories. If minor rating categories provide information, a positive correlation between the percentage of bonds in the lowest credit quality and yield would be expected.

2.1.5. Kuwait invasion

When Iraq invaded Kuwait, August of 1990, the effect on the world's oil supply, and thus, on the financial markets was unknown. To reflect a possible structural break caused by the Iraq invasion, a dummy variable was included for August of 1990, and lagged one period to encompass an effect in September of 1990. An additional dummy variable was included to study the effect the liberation of Kuwait had on the high-yield market.

2.2. Interest rate risk measures

Interest rate risk is the dominant factor affecting the value of investment grade bonds. However, as credit quality decreases default risk begins to dominate interest rate risk in bond valuation. By analyzing yield as compared to yield spread, the strength and significance of changes in Treasury rates on the yield of noninvestment grade bonds can be studied. Unfortunately, the historical average duration of the high-yield indices was unavailable, and hence, some of the temporal variance in each index and cross-sectional variance across indices may be due to changes in average duration as well as simple changes in Treasury rates, which cannot be controlled for in this study.

[3] The percentage of subordinated bonds in the lowest rating category was unavailable on a monthly basis for the indices utilized in this study. As a proxy, COMPUSTAT data was utilized to measure these vaiables.

[4] Ibid.

At the end of 1995, the CS First Boston Aggregate Index had an average maturity of 7.85 years and duration of 4.35 years. [5] However, since the majority of the bonds have call features the effective maturity and duration are less. To assess the appropriate Treasury yield to utilize in this study, a correlation analysis was done on 5-, 7-, and 10-year notes. The results indicate that the appropriate Treasury yield to utilize is the 10-year Treasury note.

2.3. Liquidity risk measures

2.3.1. Supply and demand factors

Warther (1995) found that mutual fund investment flow influenced stock and bond returns. The Investment Institute (ICI) tracks mutual funds by investment category including one labeled high-yield. The ICI reports data on the asset value and percentage of assets held in liquid investments on a monthly basis. Since mutual funds make up a large segment of the market, [6] the change in mutual fund flow and liquidity position of the mutual funds could have a significant effect on market yield. Fridson and Jonsson (1995) found increased fund flow into high-yield mutual funds, as a percentage, to be associated with a narrowing of the yield spread and an increase in the price of noninvestment grade securities. Further, an increase in the amount of assets held as liquid securities, as percentage of high-yield assets, was associated with an increase in yield spread and a decrease in the price of noninvestment grade securities. ICI provided mutual fund data for the current study.

2.3.2. January effect

There is a well documented January effect in bond returns (see Chang and Pinegar, 1986, 1988; Chang and Huang, 1990; Fama and French, 1993; Cooper and Shulman, 1994; and Maxwell, 1998). Chang and Pinegar (1986) concluded the evidence regarding a January effect was consistent with a tax-loss selling strategy. On the other hand, Cooper and Shulman (1994) offered a conjecture that a January effect was the result of year-end selling by portfolio managers to prevent high-yield debt from appearing on the fund's financial report. This year end selling is commonly referred to as ''window dressing'' (see Lakonishok et al., 1991). DeRosa-Farag (1996) suggested the January effect was the result of supply and demand considerations. Bond coupon payments are not evenly distributed throughout the year. Coupon payments are at their highest in December and lowest in January, and hence, DeRosa-Farag suggested the January price increase was due to an increase in fund flow in December and a decrease in fund flow in January. Fridson and Garman (1995) found no supporting evidence for a coupon-payment theory.

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[5] High Yield Handbook, 1996, (CS First Boston, New York).

[6] Chase Securities estimates high-yield mutual funds comprise 22% of the market in 1995. (DeRosa-Farag, 1996).

The year-end based theories and the coupon-payment theory provide testable hypotheses. The year-end theories and demand based theories suggest yields should be at their highest in December and lowest in January. DeRosa-Farag's demand-based theory has an additional testable hypothesis. Most coupon-bond payments are made semi-annually, and hence, the coupon-payment theory suggests that a similar effect on yields should be found in June and July. The current study tests for the significance of a January effect as well as the competing hypotheses for the causes of the anomaly.

2.3.3. Drexel Burnham bankruptcy

Drexel Burnham was the largest underwriter and market maker in high-yield bonds when it declared bankruptcy in February of 1990. Cornell (1992) found the Drexel Burnham bankruptcy had an effect on the overall marketability of high-yield bonds. To test for the significance of the marketability crisis, a Drexel Bankruptcy dummy variable was included. If the bankruptcy caused a liquidity crisis a positive effect on yield would be expected around the time of the bankruptcy. To account for the market anticipating the bankruptcy, the dummy variable was also examined over two lags.

3. Analysis of long-run equilibrium yield relationships

In this section, the variables' stationarity and order of integration are tested. The cointegration tests for long-run relationships between the simple yield spread model and our default risk adjusted yield spread model are then presented. The sample period under investigation is January 1987 to July 1996.

3.1. Stationarity and integration analysis

To avoid a potential spurious regression problem (Granger and Newbold, 1974 and Phillips, 1986), the first step in analyzing the data was to determine the stationarity or nonstationarity of the variables. The variables were plotted, the autocorrelation functions were examined, and augmented Dickey-Fuller (ADF) statistics were evaluated on the levels and the first differences of the variables. [7] As expected, most of the variables were found to be nonstationary, I(1), but the first differences of all were stationary, I(0).

3.2. Cointegration analysis and testing long-run relationships

Given the nonstationarity of a number of the variables, the traditional approach is to model a process in differences. While this is common practice, it results in a

[7] Test results are available from the authors upon request.

potential loss of information on the long-run interaction of variables. So instead of directly moving to a model utilizing differences, an analysis was performed to determine if there is a cointegrating vector. The implication of a cointegrating vector is that while the variables may be individually nonstationary a linear combination of variables is stationary (see Enders, 1995). Hence, a cointegrating vector indicates a long-run relationship between the variables.

This study considers two models to explain the long-run relationship of noninvestment grade bond yields and Treasury yields. The first model examines a yield spread model in which there is a long-term equilibrium between Treasury yields and noninvestment grade bond yields. The second model examines the more complex relationship of the default adjusted yield-premium models between Treasury yields, noninvestment grade bond yields, and default rates. Figs. 1 and 2 show graphs of the variables.

Two methodologies have been developed for testing for cointegration. The first approach was a single-equation approach developed by Engle and Granger (1987) aptly called the Engle–Granger Methodology. This approach is valid in bivariate analysis, but inadequate in a multivariate framework and systems approach. A second and more powerful system methodology was developed by Johansen (1988; 1991). The Johansen maximum likelihood procedure for a finite-order vector autoregressions (VARs) was utilized in the current study.

The first step in the Johansen methodology is to determine the appropriate lag structure to use in the VARs. The VARs include the BB, B, and CSFB bond yields, T-Bond yields, and Moody's default rates. In addition, variables, which

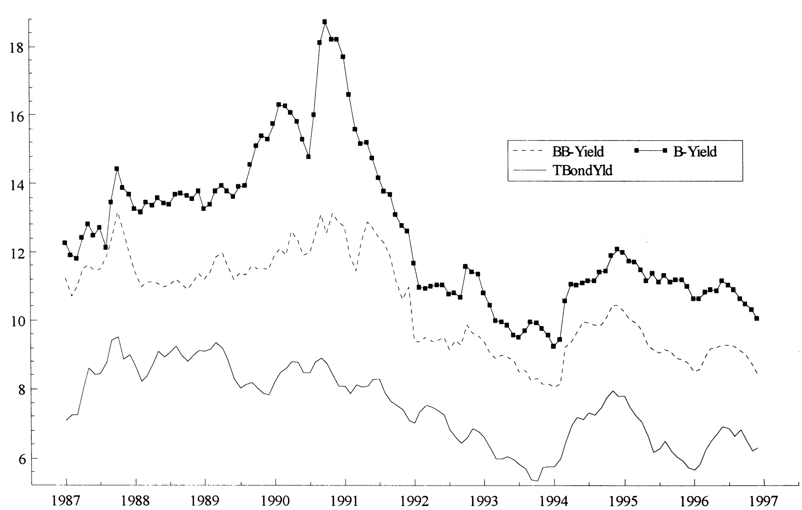

Fig. 1. Ten-year treasury, BB and B indices yields (January 1987–December 1996).

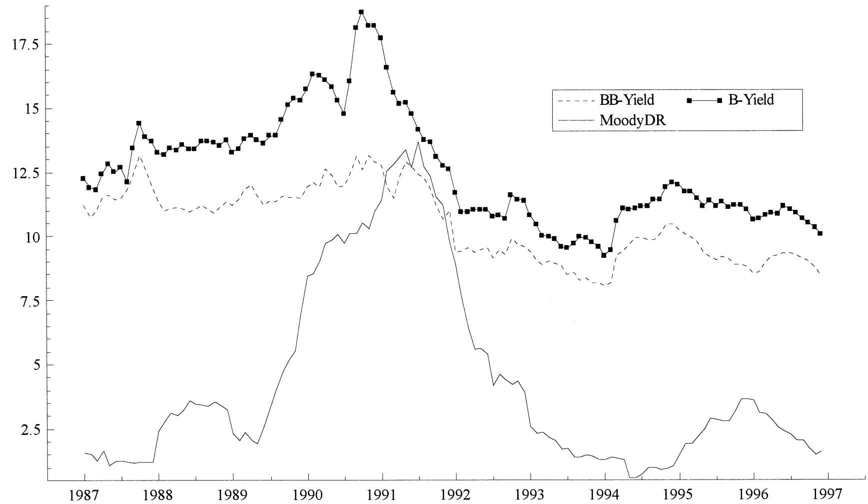

Fig. 2. BB and B indices yields and Moody's trailing-12-month default rate (January 1987–December 1996).

will be tested for significance but not considered in the cointegrating vector, were allowed to enter unrestricted. [8]

To determine the appropriate lag structure the log-likelihood, Schwartz criterion, Hannan–Quinn, and the F-statistic for model comparison were utilized. The test statistics suggested that a lag structure of two periods was appropriate.

Table 2 reports the summarized results of the Johansen procedure for the first model between the bond indices and the Treasury yields. The Johansen maximal eigenvalue (λ_{max}) for the BB Index for $r = 0$ is 8.83 and the 95% critical value is 14.1. The maximal and trace eigenvalues indicate that the null hypothesis of no cointegrating vector ($r = 0$) cannot be rejected for any of the systems.

The results in Table 2 along with the correlation analysis above demonstrates the danger of viewing a constant and instantaneous change in yields on very-risky debt to changes in the risk-free security. [9]

[8] The unrestricted variables were an index of lagging economic indicators, an index ofleading economic indicators, a Russell 2000 Index, the % subordinated debt in rating class, the % of lowest credit quality in rating class, a Iraq Invasion d.v., a 1987 Market Crash d.v., the % of high-yield mutual fund assets held in liquid securities, the change in high-yield mutual fund flow, seasonal d.v(s)., and a Drexel Bankruptcy d.v.

[9] Cointegrating vectors between investment grade indices yields and Treasury yields were found for all investment grade indices, which implies the legitimacy of a yield spread model as applied to investment grade indices.

Table 2
Cointegration analysis of high-yield indices and treasury bond yield
λ_{max} and λ^{α}_{max} are Johansen's maximal eigenvalue statistics, and λ_{trace} and λ^{α}_{trace} are Johansen's trace eigenvalue statistics. An α signifies the statistic is adjusted for degrees of freedom.

Index	BB	B	CSFB
Cointegration statistics			
Eigenvalue	0.0752	0.0578	0.04756
Null hypothesis	$r=0$	$r=0$	$r=0$
λ_{max}	8.83	6.73	5.51
λ^{α}_{max}	8.52	6.49	5.31
95% critical value	14.1	14.1	14.1
λ_{trace}	11.58	8.72	7.07
λ^{α}_{trace}	11.17	8.41	6.82
95% critical value	15.4	15.4	15.4
Standardized eigenvectors β' T-Bond yield	−1.397	−2.082	−2.605
Standardized adjustment coefficient α	−0.0950	−0.0610	−0.0340

The test of a long-term relationship between Treasury yields, default rates, and yields on very-risky debt are reported in Table 3. Variables, which will be used later to help describe the short-term dynamics of the market, were allowed to enter the VARs unrestricted. The null hypothesis of $r=0$ is that there is no cointegrating vector. The Johansen maximal eigenvalue (λ_{max}) for the BB Index for $r=0$ is 24.51 and the 95% critical value is 21.0. All the maximal and trace eigenvalue statistics strongly reject the null hypothesis that there is no cointegrating vector for all the different high-yield indices implying at least one cointegrating vector exists for each of the indices.

The null hypothesis of $r \leq 1$ is that there is one or more cointegrating vector. The Johansen maximal eigenvalue (λ_{max}) for the BB Index for $r \leq 1$ is 7.98 and the critical value is 14.1. All the maximal and trace eigenvalue statistics suggest little evidence of more than one cointegrating vector, and it was concluded that there was one cointegrating vector for each of the indices.

Table 3 includes Chi-square ($\chi^2(df)$) test statistics for the significance of Treasury Bond yields and default rates in the cointegrating vector for the different noninvestment grade bond indices. For example, the $\chi^2(1)$ for Treasury Bond Yield's significance on the BB Index is 23.751 indicating that the variable adds explanatory power at 1% to the model. The Chi-square test statistics suggest that both Treasury Bond yields and default rates are statistically significant in determining the cointegrating relationship for all the indices. [10]

[10] Cointegrating vectors were estimated for investment grade indices, Treasury yields, and default rates. The cointegrating vectors were statistically significant. However, the default rate was found to be nonsignificant in the cointegrating vector.

Table 3
Test of integrated model cointegration analysis of high-yield indices, treasury bond yields and Moody's 12-month trailing default rate

Index	BB	B	CSFB
Cointegration statistics			
Eigenvalue	0.2100	0.3958	0.2844
Null hypothesis	$r=0$	$r=0$	$r=0$
λ_{max}	24.51**	52.41***	34.80***
λ_{max}^{α}	23.10**	49.38***	32.79***
95% critical value	21.0	21.0	21.0
λ_{trace}	32.40**	61.75***	39.37***
λ_{trace}^{α}	30.53**	58.18***	37.09***
95% critical value	29.7	29.7	29.7
Null hypothesis	$r\leq 1$	$r\leq 1$	$r\leq 1$
λ_{max}	7.98	8.33	3.71
λ_{max}^{α}	7.35	7.85	3.50
95% critical value	14.1	14.1	14.1
λ_{trace}	7.89	9.34	4.57
λ_{trace}^{α}	7.44	8.80	4.31
95% critical value	15.4	15.4	15.4
Standardized eigenvectors β'			
T-Bond yield	-0.9954	-0.7936	-0.8316
Moody's default rate	-0.11502	-0.2914	-0.3458
Standardized adjustment			
Coefficient α	-0.2016	-0.0594	-0.0627

Test statistic for the significance of the variable in cointegrating vector

Variable	$\chi^2(1)$	$\chi^2(1)$	$\chi^2(1)$
T-Bond yield	23.751***	31.292***	15.291***
Moody's default rate	17.262***	32.149***	19.848***

λ_{max} and λ_{max}^{α} are Johansen's maximal eigenvalue statistics, and λ_{trace} and λ_{trace}^{α} are Johansen's trace eigenvalue statistics. An α signifies the statistic is adjusted for degrees of freedom. The following variables entered unrestricted: Constant, Seasonal d.v(s)., Drexel d.v., Iraq d.v., Mkt. Crash d.v., %MF Liq. Assets, ΔFund Assets, S&P Index, Lead Index, Lag Index, %Sub BB/B, %$-$BB/$-$B. Asterisks will be used in tables to represent statistical significance in the following manner: *** represents a statistically significant result at the 99% confidence level, ** at the 95% confidence level, and * at the 90% confidence level.

The standardized eigenvectors β' in Table 3 are the estimated cointegrating vectors for the different indices' long-term market equilibrium. The cointegrating vectors for the different indices can be written as:

$$\text{BB Index Yield} = 0.9954\ \text{T-Bond Yield} + 0.1150\ \text{Moody's Default Rate} + (\alpha\ \text{coefficient} = -0.2016) \quad (1)$$

$$\text{B Index Yield} = 0.7936 \text{ T-Bond Yield} + 0.2914 \text{ Moody's Default Rate} + (\alpha \text{ coefficient} = -0.0594) \quad (2)$$

$$\text{CSFB Index Yield} = 0.8316 \text{ T-Bond Yield} + 0.3458 \text{ Moody's Default Rate} + (\alpha \text{ coefficient} = -0.0627) \quad (3)$$

The noninvestment grade indices long-run equilibrium is essentially the risk-free rate plus a premium to reflect the increased rate of default. The relationship of the default rate to the different indices is as expected, as the bond credit quality decreases, signified by bond rating, default rate has a larger effect on yield. This result is consistent with theory.

The α coefficients in Table 3 (also found under Eqs. (1)–(3)) represent the speed of adjustment to disequilibrium comparable to a mean-reversion rate. The α coefficients ranged from -0.2016 to -0.0594. The signs are as expected (a negative sign indicates as the variables move away from equilibrium there is an adjustment back towards the equilibrium relationship). The alphas suggest a slow adjustment to disequilibrium in the more volatile B and CSFB Indices.

4. Error correction models and short-run dynamics

After testing for a cointegrating vector in a system of equations, the models are then estimated as a single-equation error correction model (ECM) (for further discussion of ECMs see Enders, 1995; and Hendry, 1995). Error correction models combine the information from the short-run dynamics of the high-yield indices with the long-run relationship found in the cointegration analysis.

4.1. Reparameterizing an autoregressive distributed lag model to an ECM

The VAR suggested a two-period lag was the appropriate structure. Hence, a single-equation autoregressive distributed lag model for the BB Index is:

$$\begin{aligned} \text{BB}_t = a_0 &+ \sum_{i=1}^{2} a_{1i}\text{BB}_{t-i} + \sum_{i=0}^{2} a_{2i}\text{T-Bond}_{t-i} + \sum_{i=0}^{2} a_{3i}\text{DefaultRate}_{t-i} \\ &+ a_4\text{LeadIndex}_t + a_5\text{LagIndex}_t + a_6\Delta\text{Russell2000} \\ &+ a_7\Delta\text{S\&P P/E} + a_8\Delta\%\text{Subordinated}_t + a_9\Delta\%\text{BB}_{-t} + a_{10}\text{Iraq}_t \\ &+ a_{11}\text{KuwaitLiberation}_t + a_{12}\text{1987Crash}_t + a_{13}\%\text{MFLiqAssets}_t \\ &+ a_{14}\%\text{MFFlows}_t + a_{15}\text{Seasonals}_t \end{aligned} \quad (4)$$

Eq. (4) can be reparameterized into an ECM without a loss of generality as:

$$\begin{aligned}\Delta \mathrm{BB}_t = {} & a_0 + b_1 \Delta \mathrm{BB}_{t-1} + \sum_{i=0}^{1} b_{2i} \Delta \text{T-Bond}_{t-i} + \sum_{i=0}^{1} b_{3i} \Delta \mathrm{DefaultRate}_{t-i} \\ & + c_1 \mathrm{BB}_{t-1} + c_2 \text{T-Bond}_{t-1} + c_3 \mathrm{DefaultRate}_{t-1} \\ & + a_4 \mathrm{LeadIndex}_t + a_5 \mathrm{LagIndex}_t + a_6\, \Delta \mathrm{Russell2000}_t \\ & + a_7 \Delta \mathrm{S\&P\,P/E} + a_8 \Delta \%\mathrm{Subordinated}_t + a_9 \Delta \%\mathrm{BB}_{-t} \\ & + a_{10} \mathrm{Iraq}_t + a_{11} \mathrm{KuwaitLiberation}_t + a_{12} \mathrm{1987Crash}_t \\ & + a_{13} \%\mathrm{MFLiqAssets}_t + a_{14} \%\mathrm{MFFlows}_t + a_{15} \mathrm{Seasonals}_t \end{aligned} \tag{5}$$

Eq. (5) can then be manipulated algebraically to directly incorporate the long-run solution:

$$\begin{aligned}\Delta \mathrm{BB}_t = {} & a_0 + b_1 \Delta \mathrm{BB}_t + \sum_{i=0}^{1} b_{2i} \Delta \text{T-Bond}_{t-i} + \sum_{i=0}^{1} b_{3i} \Delta \mathrm{DefaultRate}_{t-i} \\ & + c_1 (\mathrm{BB} + \varepsilon \text{T-Bond} + \delta\, \mathrm{DefaultRate})_{t-1} + a_4 \mathrm{LeadIndex}_t \\ & + a_5 \mathrm{LagIndex}_t + a_6 \Delta \mathrm{Russell2000}_t + a_7 \Delta \mathrm{S\&P\,P/E} \\ & + a_8 \Delta \%\mathrm{Subordinated}_t + a_9 \Delta \%\mathrm{BB}_{-t} + a_{10} \mathrm{Iraq}_t \\ & + a_{11} \mathrm{KuwaitLiberation}_t + a_{12} \mathrm{1987Crash}_t + a_{13} \%\mathrm{MFLiqAssets}_t \\ & + a_{14} \%\mathrm{MFFlows}_t + a_{15} \mathrm{Seasonals}_t \end{aligned} \tag{6}$$

The reparameterized general ECM formulation found in Eq. (6) will be used as the model to discuss the single-equation results. c_1 is the feedback coefficient reflecting a long-run adjustment to disequilibrium and is comparable to the α coefficient found in Eqs. (1)–(3) from the cointegration analysis. The ε and δ coefficients are $-c_2/c_1$ and $-c_3/c_1$ from Eq. (5). Therefore, the error correction variables can be represented from the cointegration results as:

$$\mathrm{ECM}_{\mathrm{BB}} = [\mathrm{BB} - 0.9954(\text{T-Bond}) - 0.1150(\mathrm{DefaultRate})]_{t-1} \tag{7}$$

$$\mathrm{ECM}_{\mathrm{B}} = [\mathrm{B} - 0.7936(\text{T-Bond}) - 0.2914(\mathrm{DefaultRate})]_{t-1} \tag{8}$$

$$\mathrm{ECM}_{\mathrm{CSFB}} = [\mathrm{CSFB} - 0.8316(\text{T-Bond}) - 0.3458(\mathrm{DefaultRate})]_{t-1} \tag{9}$$

4.2. General to specific modeling

General ECMs are estimated, which include both the long-run solution and the variables theorized as possibly affecting the short-run dynamics of the indices. A two-period lag was found significant on the levels in the VARs. Thus, the appropriate lag structure is two periods for stationary variables. For differenced variables the appropriate lag structure is a single period. Tables 4–6 are the resulting general models for the BB, B, and CSFB High-Yield Indices.

Table 4
General error correction model for ΔBB index
The VAR suggested a two period lag for the cointegration analysis. Thus, the appropriate lag structure for differenced (Δ) variable is one lag and for stationary variables the appropriate lag structure is two periods.
R-squared = 0.747, Standard Deviation = 0.225, Durbin–Watson = 2.25, ARCH (df:7,59) = 0.890[0.520], Normality $\chi^2(2) = 9.119$ [0.011]**.

Variable	Coefficient	Standard error	t-value	t-prob.	Partial R^2
Long-run solution					
ECM (Lag 1)	−0.243	0.062	−3.933	0.000***	0.175
Interest rate risk					
ΔT-Bonds	0.217	0.143	1.520	0.133	0.031
1-Month Lag	0.125	0.150	0.831	0.409	0.009
Default rate risk					
ΔMoody's Default Rate	−0.045	0.064	−0.706	0.482	0.007
1-Month Lag	0.214	0.062	3.430	0.001***	0.139
ΔRussell 2000 stock index (ln)	0.407	0.730	0.558	0.579	0.004
1-Month Lag	−0.278	0.740	−0.375	0.709	0.002
ΔS&P Price/Earnings Ratio	−0.048	0.040	−1.194	0.237	0.019
1-Month Lag	0.051	0.036	1.442	0.154	0.028
ΔLagging Economic Indicator	0.151	0.089	1.704	0.093*	0.038
1-Month Lag	−0.037	0.093	−0.395	0.694	0.002
ΔLeading Economic Indicator	0.103	0.116	0.881	0.381	0.011
1-Month Lag	0.074	0.113	0.655	0.515	0.006
Δ%BB class subordinated	0.049	0.035	1.409	0.163	0.027
1-Month Lag	−0.043	0.034	−1.287	0.202	0.022
Δ%BB class rated minus (−)	0.048	0.030	1.586	0.117	0.033
1-Month Lag	0.059	0.030	2.000	0.049**	0.052
Iraq Invasion	−0.233	0.256	−0.908	0.367	0.011
1-Month Lag	−0.783	0.265	−2.955	0.004***	0.107
Kuwait Liberation	−0.016	0.277	−0.059	0.953	0.000
1987 Market Crash	0.326	0.350	0.931	0.355	0.012
Liquidity risk					
Drexel Bankruptcy	−0.512	0.271	−1.891	0.063*	0.047
1-Month Lag	−0.199	0.256	−0.776	0.441	0.008
2-Month Lag	0.512	0.251	2.038	0.045**	0.054
%Mutual Fund Liquid Assets	−0.058	0.043	−1.336	0.186	0.024
1-Month Lag	0.078	0.048	1.637	0.106	0.035
2-Month Lag	−0.017	0.040	−0.413	0.681	0.002
%New Mutual Fund Flow	−0.057	0.010	−5.889	0.000***	0.322
1-Month Lag	−0.004	0.010	−0.404	0.687	0.002
2-Month Lag	0.018	0.008	2.145	0.035**	0.059
January	−0.076	0.094	−0.810	0.421	0.009
June	−0.033	0.089	−0.373	0.710	0.002
July	−0.053	0.094	−0.567	0.573	0.004
December	0.029	0.100	0.290	0.773	0.001

Table 5
General error correction model for ΔB index
The VAR suggested a two period lag for the cointegration analysis. Thus, the appropriate lag structure for differenced (Δ) variable is one lag and for stationary variables the appropriate lag structure is two periods.
R-squared = 0.816, Standard Deviation = 0.240, Durbin–Watson = 1.64, ARCH (df:7,59) = 0.265[0.965], Normality $\chi^2(2) = 30.912[0.000]^{***}$.

Variable	Coefficient	Standard error	t-value	t-prob.	Partial R^2
Long-run solution					
ECM (Lag 1)	−0.137	0.044	−3.072	0.003***	0.115
Interest rate risk					
ΔT-Bonds	0.038	0.153	0.252	0.802	0.001
1-Month Lag	0.073	0.156	0.469	0.640	0.003
Default rate risk					
ΔMoody's Default Rate	0.093	0.066	1.421	0.160	0.027
1-Month Lag	0.102	0.068	1.502	0.137	0.030
ΔRussell 2000 Stock Index (ln)	−0.661	0.798	−0.828	0.410	0.009
1-Month Lag	1.130	0.775	1.458	0.149	0.028
ΔS&P Price/Earnings Ratio	−0.064	0.042	−1.513	0.135	0.030
1-Month Lag	0.001	0.040	0.025	0.980	0.000
ΔLagging Economic Indicator	0.132	0.093	1.422	0.159	0.027
1-Month Lag	0.111	0.097	1.147	0.255	0.018
ΔLeading Economic Indicator	0.100	0.130	0.771	0.443	0.008
1-Month Lag	−0.090	0.120	−0.753	0.454	0.008
Δ%B class subordinated	−0.045	0.058	−0.783	0.436	0.008
1-Month Lag	0.030	0.062	0.476	0.635	0.003
Δ%B class rated minus (−)	−0.027	0.052	−0.523	0.602	0.004
1-Month Lag	−0.040	0.054	−0.733	0.466	0.007
Iraq Invasion	0.095	0.265	0.359	0.721	0.002
1-Month Lag	−0.470	0.282	−1.670	0.099*	0.037
Kuwait Liberation	−0.076	0.300	−0.253	0.801	0.001
1987 Market Crash	−0.553	0.365	−1.514	0.134	0.031
Liquidity risk					
Drexel Bankruptcy	−0.008	0.290	−0.027	0.978	0.000
1-Month Lag	−0.077	0.274	−0.280	0.781	0.001
2-Month Lag	−0.180	0.268	−0.673	0.503	0.006
%Mutual Fund Liquid Assets	−0.029	0.045	−0.633	0.529	0.006
1-Month Lag	0.076	0.051	1.496	0.139	0.030
2-Month Lag	−0.057	0.042	−1.344	0.183	0.024
%New Mutual Fund Flow	−0.099	0.010	−9.514	0.000***	0.554
1-Month Lag	−0.014	0.011	−1.224	0.225	0.020
2-Month Lag	0.019	0.009	2.014	0.048**	0.053
January	−0.140	0.102	−1.365	0.176	0.025
June	−0.094	0.097	−0.971	0.335	0.013
July	−0.081	0.095	−0.847	0.400	0.010
December	0.038	0.111	0.343	0.732	0.002

Table 6
General error correction model for ΔCSFB index
The VAR suggested a two-period lag for the cointegration analysis. Thus, the appropriate lag structure for differenced (Δ) variable is one lag and for stationary variables the appropriate lag structure is two periods.
R-squared = 0.808, Standard Deviation = 0.233, Durbin–Watson = 1.41, ARCH (df:7, 59) = 0.300 [0.951], Normality $\chi^2(2) = 30.134$ [0.000]***.

Variable	Coefficient	Standard error	t-value	t-prob.	Partial R^2
Long-run solution					
ECM (Lag 1)	−0.068	0.035	−1.959	0.054*	0.047
Interest rate risk					
ΔT-Bonds	0.218	0.132	1.654	0.102	0.034
1-Month Lag	0.071	0.141	0.500	0.619	0.003
Default rate risk					
ΔMoody's Default Rate	0.069	0.064	1.073	0.286	0.015
1-Month Lag	0.057	0.065	0.880	0.382	0.010
ΔRussell 2000 Stock Index (ln)	−0.958	0.739	−1.297	0.199	0.021
1-Month Lag	0.988	0.725	1.362	0.177	0.023
ΔS&P Price/earnings Ratio	−0.002	0.036	−0.056	0.956	0.000
1-Month Lag	0.032	0.033	0.985	0.328	0.012
ΔLagging economic indicator	0.151	0.096	1.585	0.117	0.031
1-Month Lag	0.125	0.099	1.267	0.209	0.020
ΔLeading Economic Indicator	0.051	0.118	0.435	0.665	0.002
1-Month Lag	−0.068	0.114	−0.601	0.550	0.005
Iraq Invasion	0.106	0.252	0.423	0.674	0.002
1-Month Lag	−0.207	0.262	−0.791	0.431	0.008
Kuwait Liberation	−0.629	0.288	−2.187	0.032**	0.058
1987 Market Crash	−0.183	0.344	−0.531	0.597	0.004
Liquidity risk					
Drexel Bankruptcy	0.186	0.277	0.672	0.504	0.006
1-Month Lag	−0.211	0.258	−0.819	0.416	0.009
2-Month Lag	0.085	0.257	0.332	0.741	0.001
%Mutual Fund Liquid Assets	−0.007	0.040	−0.184	0.854	0.000
1-Month Lag	0.035	0.047	0.729	0.468	0.007
2-Month Lag	−0.029	0.040	−0.729	0.468	0.007
%New Mutual Fund Flow	−0.084	0.010	−8.490	0.000***	0.480
1-Month Lag	−0.002	0.010	−0.243	0.808	0.001
2-Month Lag	0.019	0.009	2.127	0.037**	0.055
January	−0.145	0.095	−1.518	0.133	0.029
June	−0.021	0.090	−0.231	0.818	0.001
July	−0.090	0.086	−1.049	0.298	0.014
December	0.110	0.101	1.096	0.277	0.015

The general models were then reduced to more parsimonious models. The decision criteria used to determine the final specific model was twofold. First, all variables, which were statistically significant at the 90% confidence level, were included. Second, F-tests were performed on the models to determine the significance of the loss of information from removing a variable. The differences in the statistical significance in the general and parsimonious models are in part due to the correlation of a number of the dummy variables to the other independent variables. For example, the significance of the 1987 market crash is higher if the stock index, S&P price/earnings ratio, and mutual fund flow are removed from the model. The same argument can be made about the dummy variables accounting for the Kuwait invasion and subsequent liberation. [11] The dummy variables in effect catch the additional variation. [12]

Table 7 shows the resulting parsimonious models for the BB, B, and CSFB Indices respectively. The specific models include summary test statistics to examine the properties of the models. [13]

4.3. Statistical properties of the models

Tables 4–7 contain summary test statistics at the bottom of the tables. The test statistics examine the fit of the model, the autocorrelation, heteroscedasticity, and normality of the errors. Overall the models' predicted values fit well with the actual values. The specific models had R^2 of 0.62 to 0.82 and standard errors of 0.20 to 0.25. The standard error equates to 20 to 25 basis points error. The autocorrelation and heteroscedasticity measures for a six-period lag indicate no statistically significant problem. However, there was an indication of heteroscedasticity for a single lag. An ARCH (1) model was tested but was found to add little to the model. The biggest statistical problem was the nonnormality of the errors. To assess the extent and cause of the normality problem the actual versus fitted values were examined.

The problem with the normality of the errors appears to be related to external shocks during the 1990 to 1992 time frame rather than a problem with the overall fit of the model. The external shocks in the 1989 to 1992 time frame included a

[11] To determine if the dummy variables were influencing our conclusions the specific models for the BB, B and CSFB Indices were reestimated excluding the dummy variables. The only change in the reestimated models was that the percentage of subordinated debt in the BB index slipped slightly below the 90% confidence level.

[12] A problem with the collinearity of the lagged change in the dependent variable was also found in the general models. To correct for this, the lagged change was dropped from the model with no loss of information.

[13] An investment grade index (A rated bonds) was also modeled. The cointegrating vector only included the A yield and the Treasury yield. [8,9] The resulting specific model for the index A included the ECM variable, the change in Treasury yields, and the change in mutual fund flow into investment grade bond funds. The change in Treasury yield was the most significant factor with a partial $R^2 = 0.833$ and the models $R^2 = 0.870$.

Table 7
Specific error correction model for ΔBB, B, and CSFB indices
The VAR suggested a two period lag for the cointegration analysis. Thus, the appropriate Lag structure for differenced (Δ) variable is one lag and for stationary variables the appropriate Lag structure is two periods.

Variable	Coefficient	Std. error	t-value	t-prob.	Partial R^2
BB index long-run solution					
ECM (Lag 1)	−0.261	0.052	−5.040	0.000***	0.203
Interest rate risk					
ΔT-Bonds	0.324	0.095	3.274	0.002***	0.097
Default rate risk					
ΔMoody's Default Rate — 1-Month Lag	0.126	0.048	3.024	0.032***	0.084
Δ% BB Class Subordinated	0.052	0.032	1.661	0.099*	0.027
Δ% BB Class Rated Minus (−)	0.053	0.027	1.941	0.055*	0.036
Liquidity risk					
Drexel Bankruptcy — 2-Month Lag	0.547	0.246	2.223	0.029**	0.047
%New Mutual Fund Flow — 2-Month Lag	−0.050	0.007	−7.411	0.000***	0.355
	0.017	0.007	3.274	0.001***	0.058
BB Model test statistics: $R^2=0.622$, Standard Deviation = 0.235, Durbin–Watson = 2.11, ARCH (df:7, 97) = 0.673[0.694], Normality $\chi^2(2)=17.231[0.000]$***					
B index long-run solution					
ECM (Lag 1)	−0.082	0.024	−6.446	0.000***	0.285
Interest rate risk					
ΔT-Bonds	0.147	0.083	2.603	0.010***	0.060

Default rate risk					
ΔMoody's Default Rate	0.167	0.045	3.477	0.001***	0.115
ΔS&P Price/Earnings Ratio	−0.057	0.024	−2.325	0.020**	0.050
Iraq Invasion —1-Month Lag	1.240	0.211	5.871	0.000***	0.245
Liquidity Risk					
%New Mutual Fund Flow	−0.083	0.006	−13.801	0.000***	0.643
January	−0.135	0.071	−1.897	0.086*	0.032
B Model test statistics: $R^2 = 0.818$, Standard Deviation = 0.201, Durbin–Watson = 1.77, ARCH (df:7,89) = 1.433[0.202], Normality $\chi^2(2) = 12.994[0.001]$ ***					
CSFB HYI long-run solution					
ECM (Lag 1)	−0.112	0.025	−4.411	0.000***	0.150
Interest rate risk					
ΔT-Bonds	0.306	0.097	3.134	0.002***	0.082
Default rate risk					
ΔMoody's Default Rate	0.140	0.053	2.647	0.009***	0.060
Iraq Invasion —1-Month Lag	0.984	0.265	3.710	0.000***	0.111
Kuwait Liberation	−0.580	0.277	−2.089	0.039**	0.038
Liquidity risk					
%New mutual fund flow	−0.085	0.007	−11.308	0.000***	0.538
2-Month Lag	0.012	0.007	1.733	0.086*	0.026
CSFB Model test statistics: $R^2 = 0.734$, Standard Deviation = 0.252, Durbin–Watson = 1.95, ARCH (df:7,88) = 1.128 [0.267], Normality $\chi^2(2) = 23.182[0.000]$***					

large number of bankruptcies, increased regulation of Savings and Loans portfolios, the bankruptcy of the largest underwriter and market maker (Drexel-Burnham), the failed UAL buyout, and the Gulf War. Given these external shocks to the market it is not surprising the model has large errors during that time period.

4.4. Formulation of specific models

The following is the formulation of the parsimonious error correction models derived from Eqs. (4)–(6) and Table 7:

$$\begin{aligned}\Delta \mathrm{BB}_t = & -0.26[(\mathrm{BB}-0.99(\text{T-Bond})-0.12(\mathrm{DefaultRate})]_{t-1}\\ & +0.32(\Delta\text{T-Bond})_t+0.13(\Delta\mathrm{DefaultRate})_{t-1}\\ & +0.05(\Delta\%\mathrm{Subordinated})+0.05(\Delta\%\mathrm{BB}-)_t\\ & +0.55(\mathrm{Drexel})_{t-2}-0.05(\%\mathrm{MutualFundFlow})_t\\ & +0.02(\%\mathrm{MutualFundFlow})_{t-2}\end{aligned} \tag{10}$$

$$\begin{aligned}\Delta \mathrm{B}_t = & -0.15[\mathrm{B}-0.79\,(\mathrm{T}-\mathrm{Bond})-0.29(\mathrm{DefaultRate})]_{t-1}\\ & +0.15(\Delta\text{T-Bond})_t+0.17(\Delta\mathrm{DefaultRate})_t\\ & -0.06(\Delta\mathrm{P/E\,Ratio})_t+1.24(\mathrm{IraqInvasion})_{t-1}\\ & -0.08(\%\mathrm{MutualFundFlow})_t-0.14(\mathrm{January})_t\end{aligned} \tag{11}$$

$$\begin{aligned}\Delta \mathrm{CSFB}_t = & -0.11[\mathrm{CSFB}-0.83(\text{T-Bond})-0.35(\mathrm{DefaultRate})]_{t-1}\\ & +0.31(\Delta\text{T-Bond})_t+0.14(\Delta\mathrm{DefaultRate})_{t-1}\\ & +0.98(\mathrm{IraqInvasion})_{t-1}-0.58(\mathrm{KuwaitLiberation})_t\\ & -0.09(\%\mathrm{MutualFundFlows})_t+0.01(\%\mathrm{MutualFundFlows})_{t-2}\end{aligned} \tag{12}$$

5. Empirical findings

In this section, conclusions are drawn from the empirical findings about the long-run properties of the model, the effect of removing mutual fund flow from the models on our conclusions, the effect of dynamic liquidity, default and interest rate measures on the indices yields, and the segmentation of the high-yield market.

5.1. Long-run properties of the models

The error correction component was statistically significant in all the models. These results in conjunction with the cointegration analysis clearly demonstrate the importance of the long-run relationship between noninvestment grade yields, Treasury yields, and default rates. The speed of adjustment response was highest for the BB Index and lowest for the CSFB Index. The value of the coefficient in the BB Index was -0.261 (t-statistic $= -5.04$). This represents a relatively rapid reversion towards the long-run relationship between the variables. The lower credit quality B Index had a lower coefficient (-0.082 with a t-statistic of (-6.45)) suggesting a slower reversion towards the equilibrium over time.

Is the long-run relationship stable or fluctuating over time? One way to assess this is to analyze the consistency of the ECM coefficient estimate over time using recursive estimation. To do this the ECM variable is estimated using three years of data and the change in the coefficient is then tracked over time as each additional time period is added to the model. Fig. 3 shows the beta estimates of the error correction coefficient over time with dashed lines representing two standard errors.

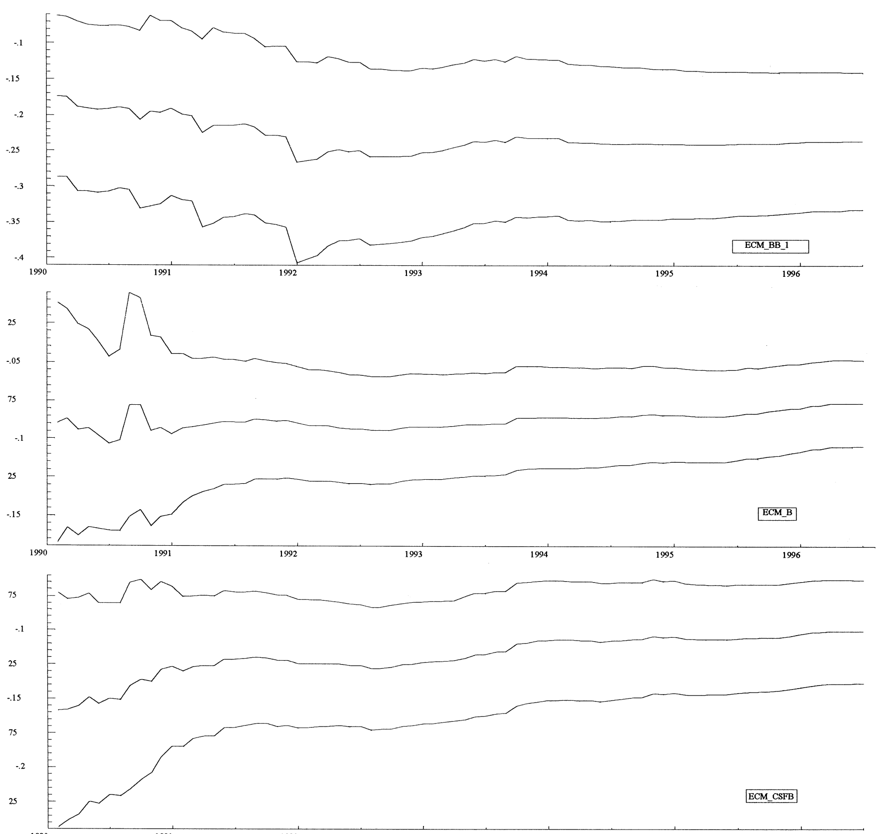

Fig. 3. Stability of Beta Estimate of the Error Correction Coefficient in the specific BB, B, and CSFB index models (the Beta Estimate is Bounded $\pm$ Two Standard Errors).

Table 8
Correlation analysis between mutual fund flow and other variables

	Correlation to mutual fund flow
Default rate measures	
ΔMoody's Default Rate	−0.19
ΔLeading Economic Indicators (natural logarithm)	0.27
ΔLagging Economic Indicators (natural logarithm)	−0.01
ΔRussell 2000 Stock Index (natural logarithm)	0.48
ΔS&P Price/Earnings Ratio	0.24
Interest rate measure	
ΔT-Bond	−0.23
Liquidity measure	
%MF liquid assets	−0.21

The results indicate a stable and consistent relationship for each of the indices. The primary exception to this conclusion is the 1990 time period when the ECM coefficient for B and CSFB Indices had large standard errors. This is not unexpected given the high volatility during that period. However, the upper recursive standard error bound is always negative. Also the recursive coefficient estimates for the full sample stay within the confidence interval at the end of the sample.

5.2. Reestimation of models excluding mutual fund flows

Given the dominance of mutual fund flow in explaining the yield on noninvestment grade indices, the significance of a number of other correlated variables may be lost. For example, mutual fund flow into the high-yield market decreased 8% in October of 1987, which coincides with the 1987 market crash. Similarly mutual fund flow decreased 9% and 10% respectively in August and September of 1990 (Iraq invaded Kuwait in 8/90), and increased by 8% in February and 7% in March when the coalition liberated Kuwait. Table 8 contains the results of a correlation analysis between mutual fund flows and other explanatory variables tested in the models. Given the correlation structure of mutual fund flows to the other explanatory variables, the models were reestimated without mutual fund flows in the models (Table 9).

The following are the formulations of the parsimonious error correction models from Table 9:

$$\begin{aligned}\Delta \mathrm{BB}_t = & -0.23\left[\left(BB - 0.99(\text{T-Bond}) - 0.12(\text{DefaultRate})\right]_{t-1}\right.\\ & + 0.54(\Delta\text{T-Bond})_t + 0.20(\Delta\text{DefaultRate})_{t-1}\\ & - 1.85(\Delta\text{Russell2000})_t + 0.08(\Delta\%\text{Subordinated})\\ & + 0.05(\Delta\%\text{BB}-)_t - 0.21(\text{January})_t \end{aligned} \tag{13}$$

$$\begin{aligned}\Delta \mathrm{B}_t = {} & -0.07[(\mathrm{B} - 0.79(\text{T-Bond}) - 0.29(\text{DefaultRate})]_{t-1} \\ & + 0.59(\Delta\text{T-Bond})_t + 0.23(\Delta\text{DefaultRate})_t \\ & - 2.26(\Delta\text{Russell2000})_t - 0.21(\Delta\text{LeadIndicators})_t \\ & + 1.71(\text{IraqInvasion})_{t-1} - 0.87(\text{KuwaitLiberation})_t \\ & - 0.27(\text{January})_t \end{aligned} \tag{14}$$

$$\begin{aligned}\Delta \mathrm{CSFB}_t = {} & -0.02[\mathrm{CSFB} - 0.83(\text{T-Bond}) - 0.35(\text{DefaultRate})]_{t-1} \\ & + 0.61(\Delta\mathrm{T} - \text{Bond})_t + 0.19(\Delta\text{DefaultRate})_t \\ & - 2.23(\Delta\text{Russell2000})_t + 1.44(\text{IraqInvasion})_{t-1} \\ & - 1.32(\text{KuwaitLiberation})_t - 0.30(\text{January})_t \end{aligned} \tag{15}$$

The results were consistent with the specific models found in Table 7 but some variables previously found not to be statistically significant entered into the new models. The change in the Russell 2000 Index entered into the BB, B, and CSFB Indices with a p-value at 1%. This is consistent with the strong correlation between mutual fund flow and changes in the stock index found in Table 9. The change in the leading economic indicator was found to have a negative effect on yield at the 90% confidence level for the B Index. However, the standard error is considerably higher for these models which have excluded mutual fund flow. This is especially true for the lower credit quality B index (see Table 10).

5.3. Dynamic liquidity risk measures affecting the yield of high-yield indices

The Percentage of Mutual Fund Flow had a statistically significant negative effect on all the indices. In fact, mutual fund flow had the highest partial R^2 of any variable in all of the models. The results also suggest the lower quality indices, B and CSFB, are more susceptible to price fluctuations as a result of fund flow. This is exemplified by the large t-statistics and partial R^2 of the fund flow coefficient in the B and CSFB Indices. The strength of the relationship of yield to fund flow could be the explanation for the seemingly contradictory finding of a higher R^2 for the more volatile B and CSFB Indices as compared to the BB Index. Curiously, fund flow lagged 2 months had a positive impact on all the indices and was statistically significant for the BB and CSFB Indices. This could indicate a seasonal component to fund flow.

The results are consistent with a January effect for all the indices. The January coefficient was negative as expected in all the general models and was statistically significant at the 10% level in the specific model of the B Index. The December coefficient was positive in all the general models. When mutual fund flow was removed from the models, a positive effect of between 21 and 30 basis points was found at the 5% level in January for all the indices.

Overall, the results regarding the January and December variables are consistent with either a tax-loss selling effect, portfolio window dressing and/or coupon payment flows. However, there was no corresponding seasonal variation during June and July which calls into question the validity of the coupon payment flow

Table 9
Specific error correction model for ΔBB, B, and CSFB indices excluding mutual fund flow
The VAR suggested a two period lag for the cointegration analysis. Thus, the appropriate lag structure for differenced (Δ) variable is one lag and for stationary variables the appropriate lag structure is two periods.

Variable	Coefficient	Std. error	t-value	t-prob.	Partial R^2
BB index long-run solution					
ECM (Lag 1)	−0.229	0.055	−4.148	0.000***	0.146
Interest rate risk					
ΔT-Bonds	0.537	0.107	5.027	0.000***	0.200
Default rate risk					
ΔMoody's Default Rate — 1-Month Lag	0.199	0.054	3.721	0.000***	0.121
ΔRussell 2000 Stock Index (ln)	−1.851	0.474	−3.907	0.000***	0.131
Δ%BB class subordinated	0.075	0.036	2.079	0.040**	0.041
Δ%BB class rated minus (−)	0.048	0.031	1.547	0.095*	0.023
Liquidity risk					
January	−0.207	0.096	−2.163	0.033**	0.044
BB Model test statistics: R^2 = 0.499, Standard Deviation = 0.270, Durbin–Watson = 2.14, ARCH (df:7,94) = 2.472[0.023]**, Normality $\chi^2(2)$ = 10.031[0.007]***					
B index long-run solution					
ECM (Lag 1)	−0.066	0.037	−1.766	0.080*	0.030
Interest rate risk					
ΔT-Bonds	0.593	0.137	4.500	0.000***	0.168

Default rate risk					
ΔMoody’s Default Rate	0.228	0.072	3.178	0.002***	0.092
ΔRussell 2000 stock index (ln)	−2.260	0.582	−3.886	0.000***	0.131
ΔLeading Economic Indicator	−0.211	0.122	−1.740	0.085*	0.029
Iraq Invasion — 1-Month Lag	1.713	0.333	5.137	0.000***	0.209
Kuwait Liberation	−0.868	0.338	−2.569	0.012***	0.062
Liquidity Risk					
January	−0.272	0.114	−2.388	0.019**	0.054
B Model test statistics: $R^2 = 0.557$, Standard Deviation = 0.318, Durbin–Watson = 1.92, ARCH (df:7,85) = 0.654[0.709], Normality $\chi^2(2) = 5.339[0.069]$*					
CSFB HYI long-run solution					
ECM (Lag 1)	−0.022	0.030	−0.721	0.473	0.005
Interest rate risk					
ΔT-Bonds	0.606	0.125	4.851	0.000***	0.191
Default rate risk					
ΔMoody’s Default Rate	0.191	0.066	2.896	0.004***	0.077
ΔRussell 2000 stock index (ln)	−2.234	0.610	−3.659	0.004***	0.118
Iraq Invasion — 1-Month Lag	1.444	0.311	4.628	0.000***	0.176
Kuwait Liberation	−1.318	0.328	−4.017	0.000***	0.139
Liquidity risk					
January	−0.299	0.104	−2.886	0.005***	0.077
CSFB Model test statistics: $R^2 = 0.584$, Standard Deviation = 0. 303, Durbin–Watson = 1.85, ARCH (df:7,85) = 0.410 [0.893], Normality $\chi^2(2) = 6.715[0.035]$**					

Table 10
Standard error of models with and without mutual fund flow

Index	Model with MFF standard error	Model without MFF standard error	Increase in standard error (%)
BB	0.235	0.270	14.89
B	0.201	0.318	58.21
CSFB	0.252	0.303	20.24

theory. The absence of the January variable in the results found in Eqs. (10) and (12) implies the January effect is also partially due to mutual fund flow.

Unlike previous yield spread studies, no support was found for the Percentage of Mutual Fund Assets held in Liquid Securities as a factor affecting the yield of the different indices.

The Drexel Bankruptcy variable had a statistically significant positive effect on the yield of the BB index 3 months prior to the actual filing of bankruptcy. This suggests the market reacted well before the actual bankruptcy.

5.4. Dynamic default risk measures affecting the yield of high-yield indices

Moody's Default Rate was included in the long-run equilibrium models, in addition the Change in Moody's Default Rate also entered into the error correction models. This clearly indicates the impact of both long and short-run changes in the default rate on noninvestment grade yields. It also indicates that long-term expectations can be significantly influenced by current default rates.

The Change in the Russell 2000 Stock Market Index had no statistical effect on the general or specific models when mutual fund flow was included as an explanatory variable. However, the change in the Stock Market Index was statistically significant at the 1% level across all the models when the mutual fund flow variable was dropped. The Standard and Poor's Price/Earnings Ratio had a statistically significant negative impact on the yield of the B Index. This suggests the relative pricing of the stock market also impacts noninvestment grade bond yields. These results are consistent with Shane's (1994) findings that the lower the credit quality the higher the correlation between yield changes and changes in stock values.

The Change in the Percentage of Debt Outstanding Rated BB Minus had a statistically significant positive effect on the yield of the BB Index in both the general and specific models which included mutual fund flows and in the specific model which excluded mutual fund flow. The positive effect on the BB Index suggests that minor rating categories serve as an indicator of the perceived risk in the market. The results also support the informational content of bond ratings and suggest that at least a portion of the change in the risk of the market is captured by S&P's Bond Ratings in a relatively short period of time.

The Change in the Percentage of Subordinated Debt had a statistically significant positive effect on both parsimonious models of the BB Index. The direction of the coefficient is inconsistent with Fridson's (1995) findings. A number of other one time events were found to have statistically significant effects on the different models separate from any impact of the stock index. The Iraq Invasion of Kuwait and the subsequent Liberation of Kuwait both had an impact on yield. As expected, these exogenous shocks had a greater impact on the lower quality B and CSFB Indices.

A Leading Economic Indicator was found to have a statistically significant negative effect on the yield of the B index when mutual fund flows was excluded from the model. The sign of the coefficient is as expected. As the economy improves, the price of risky debt increases as there is a lower probability of default. No support was found for previous results (Garman and Fridson, 1996) suggesting a Lagging Economic Indicator had explanatory power for the pricing of the high-yield debt.

5.5. Dynamic interest rate risk measures affecting the yield of high-yield indices

The results demonstrate a long-run equilibrium between Treasury yields, default rates, and the yields on the different indices. In addition, short-run changes in T-Bond Yield also had an impact. The strength of the short-run change was greater when mutual fund flow was removed from the models. This suggests mutual fund flow is in part sensitive to changes in the T-Bond Yield, which is confirmed by the correlation analysis in Table 8.

5.6. Segmentation of the high-yield market

The results clearly indicate the dangers in viewing the noninvestment grade market as a homogeneous market. This study demonstrates the segmentation of the market in a number of ways. Different factors effect the yields of the BB and B Indices and shared explanatory variables have different relationships across the Indices. As expected the lower credit quality B Index is more sensitive to the default rate and less so to the Treasury yield in the long and short run. In general, the B Index was also more sensitive to changes in stock prices and mutual fund flow.

6. Summary

The yield premium model and yield spread model provide frameworks to analyze factors which affect the risk of holding risky debt instruments. However, neither model has properly factored into account interest-rate risk. We demonstrate using a correlation and cointegration analysis that interest rate risk is not constant or instantaneous as implied by both models. Instead a broader model which allows interest rate risk and default risk to fluctuate over time provides a better analytical framework to understand the long-term relationship between default rates, Treasury yields, and the yields on noninvestment grade indices.

The system and single-equation models found varying adjustments to disequilibrium. Thus, even though there is a long-term equilibrium, short-term dynamics can significantly alter the short-run relationship. Lower rated indices exhibit slower reversion toward equilibrium and larger short-run dynamic changes in yield.

In addition to a long-run equilibrium, short-run dynamic factors also affected the monthly yields. The dynamics of the market were explained by changes in the Default Risk Measures including the Moody's Trailing-12-Month Default Rates, Russell 2000 stock index, Percentage of Outstanding Debt in the Lowest Minor Rating Category, and the Percentage of Subordinated Debt Outstanding. No support was found for economic indicators as risk measures when Mutual Fund Flow was included in the model. However, when Mutual Fund Flow is dropped from the model, leading economic indicators were found to have a statistically significant negative effect on the yield of the B index.

The changes in Liquidity-Risk Measures including Mutual Fund Flow and Seasonal Variables had a significant effect on yield. The strongest explanatory variable across all models was the Change in Mutual Fund Flow. The results also indicate that the lower credit quality indices are more sensitive to mutual fund flow. A seasonal component in the yields consistent with a January effect was also found.

Finally, the study demonstrated the segmentation of the bond market and the dangers in viewing the market as being homogeneous with interest rate and default risk as constant. First, the long-term factors influencing the yield on investment and noninvestment grade bonds differ. We find the long-run equilibrium yield on noninvestment grade bonds is a function of Treasury yields and default rates. For investment grade bonds the long-run equilibrium is a function of only Treasury yields. Second, there is a danger in viewing the long-run equilibrium as static across noninvestment grade categories. The speed of adjustment coefficients in both the system and ECM differ across rating categories. The single B index adjusts much slower to disequilibrium. Third, dynamic factors also have different affects on the B and BB indices. Fourth, by properly modeling both the long and short run factors influencing the yield on bonds a great deal of the variation ($R^2 = 0.818$) in the most volatile segment of the bond market (B index) can be explained.

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