Expected Returns, Yield Spreads, and Asset Pricing Tests

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We construct firm-specific measures of expected equity returns using corporate bond yields, and replace standard *ex post* average returns with our expected-return measures in asset pricing tests. We find that the market beta is significantly priced in the cross section of expected returns. The expected size and value premiums are positive and countercyclical, but there is no evidence of positive expected momentum profits. (*JEL* G12, E44)

The standard asset pricing theory posits that investors demand an *ex ante* premium for acquiring risky securities (e.g., Sharpe, 1964; Lintner, 1965; Merton, 1973). Because the *ex ante* risk premium is not readily observable, empirical studies typically use *ex post* averaged stock returns as a proxy for expected stock returns. This practice is justified on the grounds that for sufficiently long horizons, the average return will "catch up and match" the expected return on equity securities. Therefore, the *ex post* average excess equity return provides an easy-to-implement, plausibly unbiased estimate of the expected equity risk premium.

Despite its popularity, the use of *ex post* return averages has significant limitations. For instance, the average realized return might not converge to the expected risk premium in finite samples. Inferences based on *ex post* returns thus depend on the properties of the particular data under examination.[1] More general difficulties associated with the use of *ex post* returns have been

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[1] For example, Lundblad (2005) and Pastor, Sinha, and Swaminathan (2007) use simulations to show that, except for very long time windows, realized returns do not converge to expected returns and often yield wrong inferences.

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recognized in the literature, but little has been done to understand their implications.[2] In his AFA presidential address, Elton (1999) observes that there are periods longer than 10 years during which stock market realized returns are on average lower than the risk-free rate (1973–1984) and periods longer than 50 years during which risky bonds on average underperform the risk-free rate (1927–1981). Based on these observations, Elton argues that

> "developing better measures of expected return and alternative ways of testing asset pricing theories that do not require using realized returns have a much higher payoff than any additional development of statistical tests that continue to rely on realized returns as a proxy for expected returns" (1200)

Because most results in the empirical asset pricing literature have been established using averaged realized returns, it is natural to ask whether extant inferences about risk-return trade-offs hold under alternative measures of expected returns.

In this paper, we construct an alternative measure of risk premium based on data from bond yield spreads and investigate whether well-known equity factors, such as market, size, book-to-market, and momentum, can explain the cross-sectional variations of expected stock returns. Motivated by Merton (1974), our basic approach recognizes that debt and equity are contingent claims written on the same productive assets and thus must share similar common risk factors. The upshot of this observation is that we can use corporate bond data to glean additional information about investors' required equity rates of returns. In what follows, we derive an analytical formula that links expected equity risk premiums and expected bond risk premiums, after adjusting bond yield spreads for default risk, rating transition risk, and the tax spreads between the corporate and the Treasury bonds.

Why use bond yield data? While relevant information regarding a firm's systematic risk is incorporated into both its stock and bond prices, the latter uniquely reveals key insights about investors' return expectations. First, bond yields are calculated in the spirit of forward-looking internal rates of return. To wit, bond yields are the expected returns if the bonds do not default and the yields do not change in the next period. Bond prices impound the probability of default, and yield spreads contain the expected risk premiums for taking default risk. Controlling for default risk, firms with higher systematic risk should have higher yield spreads, a relation that holds period-by-period in the cross section. This approach contrasts sharply with what can be gauged from realized equity returns. Equity returns reflect both cash flow shocks and discount rates

[2] Earlier studies have discussed in some detail the noisy nature of average realized returns in a number of different contexts (see, for example, Blume and Friend, 1973; Sharpe, 1978; Miller and Scholes, 1982).

shocks, and *ex post* averaging can overshadow conditional, forward-looking information.[3]

Second, the time-variation of expected returns in the equity markets often works against the convergence of average realized returns to the expected return. Consider, for example, that investors require a higher equity risk premium from a cyclical firm during economic downturns. Accordingly, the firm's equity price should fall and its discount rate should rise during recessions. The equity value of a cyclical firm indeed falls during recessions, reflecting value losses in its underlying assets. However, by averaging *ex post* its realized returns over the course of a recession, one might wrongly conclude that the cyclical firm is less risky because of its lower "expected" returns. Bond yield spreads, in contrast, increase during recessions, moving in the same direction as the discount rates.

We use our expected-return measure to study the cross section of expected returns using a sample of 1205 nonfinancial firms from January 1973 to March 1998. Our sample is restricted by the availability of the firm-level corporate bond data from the Lehman Brothers Fixed Income dataset. Our main empirical findings can be summarized as follows.

First and foremost, the market beta plays a much more important role in driving the cross-sectional variations of expected equity returns than is reported under *ex post* returns. In particular, the market beta is significantly priced even after we control for size, book-to-market, and prior returns. This finding is surprising given the well-known weak relation between the market beta and the average returns (e.g., Fama and French, 1992). Our evidence suggests that previous evidence that beta is dead might have resulted from the use of average returns as a poor proxy for expected returns.

Second, for the most part, the expected size and value premiums are significantly positive and countercyclical. This evidence is consistent with the view that book-to-market and size capture relevant dimensions of risk that are expected to be priced in equity returns (e.g., Fama and French, 1993, 1996). The countercyclical properties of the expected value premium also lend support to studies that emphasize the impact of business cycles and conditional information on the value premium (e.g., Ferson and Harvey, 1999; Lettau and Ludvigson, 2001). Our finding that the expected size premium remains significant and large at 3.61% per annum during the 1982–1998 period—after Banz's (1981) discovery—contrasts with studies under *ex post* averaged returns (e.g., Schwert, 2003).

Finally, we find no evidence of expected positive momentum profits. In fact, momentum is sometimes priced with a negative sign under our expected-return measure. This evidence is consistent with several interpretations. First, investors do not consider stocks with high prior realized returns to be riskier than stocks with low prior realized returns. Momentum is thus not a priced risk factor,

[3] As pointed out by Sharpe (1978), the CAPM only holds conditionally and expected returns might have nothing to do with future realized returns. Risk premia recovered from bond yields, in contrast, reflect conditional expectations.

consistent with behavioral models of Barberis, Shleifer, and Vishny (1998); Daniel, Hirshleifer, and Subrahmanyam (1998); and Hong and Stein (1999). Second, the distribution of expected returns can deviate from the distribution of realized returns because of incomplete information and learning. Specifically, even though *ex post* returns appear predictable to econometricians, investors can neither perceive nor exploit this predictability *ex ante* (e.g., Brav and Heaton, 2002; Lewellen and Shanken, 2002; Shanken, 2004). Third, momentum can be an empirical by-product of using average realized returns as a potentially poor proxy for expected returns. This possibility, coupled with the evidence that momentum strategies involve frequent trading in illiquid securities with high transactions costs (e.g., Lesmond, Schill, and Zhou, 2004; Korajczyk and Sadka, 2004), suggests that momentum can be an illusion of profit opportunity when, in fact, none exists.

Our approach has potential limitations that arise from the simplicity of our methodology and from constraints on the data that we have to use to operationalize our expected-return proxy. The simple contingent-claim framework in the spirit of Merton (1974) allows us to derive a conditionally linear relation between expected equity and bond excess returns. To this end, we assume that the risk-free rate is deterministic and asset volatility is at most a function of asset value for analytical tractability. Clearly, under more general conditions, the relation between expected equity and bond excess returns might not be linear. We thus emphasize that our empirical approach is only motivated by the Merton-style framework; it is not a structural test of that framework.

Further, because we must resort to existing default information to gauge the expected default loss, our constructed bond risk premiums are not entirely "*ex ante*." Fortunately, however, research has shown that yield spreads are too large to be explained by expected default losses (see, e.g., Huang and Huang, 2003). By restricting the use of historical data to the estimation of a small portion of the yield spreads, we retain crucial information on the forward-looking risk premiums embedded in bond yields. This information allows us to implement our new asset pricing tests.

Finally, because we gauge investors' expectations using bonds, our approach naturally focuses on data from bond issuers. This focus, in turn, constrains our analysis to a sample universe that is smaller than the CRSP universe. One could question whether our data engender common equity factors in the first place. We verify below that our data set is fairly consistent with the cross-sectional properties of the CRSP universe for the same period. We also note that the data restrictions we face should work against our finding of meaningful patterns. Therefore, despite potential limitations, our tests complement inferences based on average realized returns by providing new insights into the determinants of the cross section of expected returns.

Our work is related to the empirical literature that relates yield spreads to expected equity returns. Harvey (1986) was among the first to link yield spreads to consumption growth. Chen, Roll, and Ross (1986) find that loadings on the aggregate default premium are priced in the cross section of equity returns. Ferson and Harvey (1991) use the default premium as an instrument for aggregate expected excess returns (see also Keim and Stambaugh, 1986; Fama and French, 1989, 1993; Jagannathan and Wang, 1996). Our work differs because we model firm-level expected returns directly as a function of firm-level yield spreads.

Our work also adds to the recent literature that constructs alternative proxies for expected returns. Brav, Lehavy, and Michaely (2005) use financial analysts' forecasts to back out expected equity risk premiums. Graham and Harvey (2005) obtain measures of the equity market risk premium from surveying Chief Financial Officers. Blanchard (1993); Gebhardt, Lee, and Swaminathan (2001); and Fama and French (2002) use valuation models to estimate expected equity risk premiums.

Vassalou and Xing (2004) build on Merton (1974) to compute default likelihood measures for individual firms. We also extract information on equity from bonds. However, we differ because we construct alternative measures of expected returns from bond data, while Vassalou and Xing rely on average realized equity returns in their tests. Similar to Bekaert and Grenadier (2001); Bekaert, Engstrom, and Xing (2005); Bekaert, Engstrom, and Grenadier (2005); and Baele, Bekaert, and Inghelbrecht (2007), we also explore the joint determination of equity and bond pricing. However, unlike these papers, we focus on the cross section of returns. A concurrent study by Cooper and Davydenko (2004) also uses the yield spreads to estimate equity premiums, but they do not study the common equity factors or the cross section of expected returns.

The rest of the paper is organized as follows. Section 1 delineates our empirical framework for constructing expected equity excess returns. Section 2 describes our sample. Section 3 provides implementation details for constructing expected equity excess returns. Section 4 reports our main results on the time series of common equity factors and on the cross section of expected equity returns. Section 5 contains extensive robustness checks. Finally, Section 6 summarizes and interprets our results.

1. Empirical Framework

We lay out the basic idea underlying our empirical framework and then formalize it through a series of propositions. Section 3 discusses its implementation after we describe our sample in Section 2.

Our basic idea is that bond and equity risk premiums are intrinsically linked because equity and bond are contingent claims written on the same productive assets, an insight that can be traced back to Merton (1974). Building on this argument, we construct expected equity excess returns from expected bond excess returns. We back out the bond risk premium from the observable yield spreads, which are forward-looking. We then conduct asset pricing tests in which we replace realized equity returns with the constructed equity risk premium.

Let the uncertainty in the economy be represented by the vector $X_t = (x_{1t}, x_{2t}, \ldots, x_{Nt})'$ with a deterministic variance-covariance matrix. There exists a stochastic discount factor m_t that is a function of X_t. Following Merton (1974), we assume that all firms are leveraged with predetermined debt. A firm defaults if its asset value hits some lower boundary as a fraction of its initial value. With this setup both equity and bond are contingent claims on the asset value. And the relation between expected equity excess return and expected bond excess return is conditionally linear, as stated below:

Proposition 1. *Let R^i_{St} be firm i's equity return and R^i_{Bt} be its debt return. Also, let F_{it}, B_{it}, and S_{it} be its assets, debt, and equity values at time t, respectively, and let r_t be the interest rate. Under the Merton (1974) framework:*

$$E_t\left[R^i_{St}\right] - r_t = \left[\frac{\partial S_{it}}{\partial B_{it}}\frac{B_{it}}{S_{it}}\right]\left(E_t\left[R^i_{Bt}\right] - r_t\right) \tag{1}$$

Proof. See Appendix. ■

Proposition 1 is intuitive. Because both equity and debt are contingent claims written on the same productive assets, a firm's equity risk premium is naturally tied to its debt risk premium. Equation (1) formalizes this argument: the equity risk premium equals the debt risk premium multiplied by the elasticity of the equity value with respect to the bond value.

In general, the equity value and the bond value are functions of the underlying asset value, the risk-free rate, and the asset volatility. To derive Proposition 1, we assume that the risk-free rate is deterministic and the asset volatility is at most a function of asset value (e.g., Merton, 1974). As a result, the equity value and the bond value are driven only by the asset value. Our framework still allows multiple common factors, but they affect equity and bond values through the firm value. The Appendix provides further details.

Empirically, Equation (1) allows us to recover the equity risk premium from the bond risk premium without assuming average realized equity returns to be an unbiased measure of expected equity returns. This is a key departure from the extant literature. The following two propositions introduce our method of constructing expected bond risk premium, $R^i_{Bt} - r_t$, from observable bond characteristics.

Proposition 2. *Let Y_{it} be the yield to maturity, H_{it} be the modified duration, and G_{it} be the convexity of firm i's bond at time t. In the absence of tax differential between corporate bonds and Treasury bonds, the following relation holds for expected bond excess return and observable bond characteristics:*

$$E_t\left[R^i_{Bt}\right] - r_t = (Y_{it} - r_t) - H_{it}\frac{E_t\left[dY_{it}\right]}{dt} + \frac{1}{2}G_{it}\frac{E_t[(dY_{it})^2]}{dt}. \tag{2}$$

Proof. See Appendix. ■

Intuitively, the first term on the right-hand side of Equation (2) is the yield spread between the corporate bond and Treasury bill, which equals the expected excess return of the bond if the bond yield remains constant. The next two terms adjust for the changes in the bond yield: the first-order change is multiplied by modified duration and the second-order change is multiplied by convexity. In essence, Equation (2) provides a second-order approximation of the bond risk premium based on the yield spread.

The next challenge is to model the yield change. The existing literature is rich in models for bond yields (e.g., Merton, 1974; Longstaff and Schwartz, 1995; Collin-Dufresne and Goldstein, 2001; Huang and Huang, 2003). We do not impose a parametric model on the yield process. Instead, we focus on capturing two important empirical patterns: (i) bond value decreases in the event of default; and (ii) bond ratings generally revert to their long-run means conditional on no-default. This task is achieved with the next proposition.

Proposition 3. *Let π_{it} be the expected default probability, dY_{it}^- be the yield change conditional on default, and dY_{it}^+ be the yield change conditional on no-default. Then, the expected bond excess return is*

$$R_{Bt}^i - r_t = (Y_{it} - r_t) + \mathrm{EDL}_{it} + \mathrm{ERND}_{it}, \quad (3)$$

where EDL *denotes expected default loss rate and is defined as*

$$\mathrm{EDL}_{it} \equiv \pi_{it}(-H_{it}E_t[dY_{it}^- \mid \text{default}] + \tfrac{1}{2}G_{it}E_t[(dY_{it}^-)^2 \mid \text{default}])/dt < 0; \quad (4)$$

and ERND *denotes the expected return due to yield changes conditional on no-default, which is defined as*

$$\begin{aligned}\mathrm{ERND}_{it} \equiv (1-\pi_{it})(-H_{it}Et[dY_{it}^+ \mid \text{no default}] \\ + \tfrac{1}{2}G_{it}E_t[(dY_{it}^+)^2 \mid \text{no default}])/dt.\end{aligned} \quad (5)$$

Proof. See Appendix. ■

Finally, notice that part of the yield spread of corporate bonds over Treasury bonds arises from the fact that corporate bond investors have to pay state and local taxes while Treasury bond investors do not. Accordingly, the component in the yield spread that is related to the tax differential should be removed from the spread if one wants to obtain an accurate measure of the bond risk premium. Let C_i be the coupon payment for bond i and let τ be the effective state and local tax rate, then

$$R_{Bt}^i - r_t = (Y_{it} - r_t) + \mathrm{EDL}_{it} + \mathrm{ERND}_{it} - \mathrm{ETC}_{it}, \quad (6)$$

where ETC denotes expected tax compensation and is given by

$$\mathrm{ETC}_{it} = \left[(1 - \pi_{it}) \frac{C_i}{B_{it}} \frac{1}{dt} + \mathrm{EDL}_{it} \right] \tau. \tag{7}$$

In Equation (7), $(1 - \pi_{it}) \frac{C_i}{B_{it}} \frac{1}{dt}$ is the expected coupon rate conditional on no-default. The expected default loss rate, EDL, is also included in (7) to capture the tax refund in the event of default.

In essence, our propositions are a second-order Taylor expansion based on the risk-neutral valuation in Merton (1974). Merton models equity as a European call option on the underlying asset. The value of corporate debt, B_{it}, which has face value K and maturity T, is $B_{it} = D_t - P_{it}$, where D_t is the price of a risk-free bond and P_{it} is a put option. The yield spread can be calculated as $y_{it} = \log(K/B_{it})/T - r$, a function of F_{it}/K and volatility σ_i only. It might appear as if the systematic risk had no effect on the yield spread. Notice, however, that the firm value process follows $dF_{it}/F_{it} = \mu_i dt + \sigma_i\, d\omega_t$, where μ_i is the instantaneous expected return of firm i, determined by its covariance with the stochastic discount factor. For the same yield spread, a systematically riskier firm (with higher μ_i) will have a lower default probability, a lower expected default loss, and a higher bond risk premium. After adjusting for default risk and other components, yield spreads can then identify the cross-sectional variation of systematic risk and expected returns.

The simple contingent-claim framework used in Proposition 1 enables us to derive a conditionally linear relation between the expected equity risk premium and the expected bond risk premium. We have assumed that the risk-free rate is deterministic and asset volatility is at most a function of asset value for analytical tractability. Clearly, under more general frameworks, the relation between the expected equity and bond risk premiums might not be conditionally linear. However, so long as underlying state variables affect equity and bond only through the firm value, the relations that matter to our applied analysis still hold. In particular, all we need as a foundation for our empirical approach is the notion that the equity and the bond return processes share common risk factors stemming from the underlying firm process. Our framework operationalizes this idea in a tractable way.[4]

2. Data and Descriptive Statistics

Before we discuss implementation details of our empirical framework, we first describe our sample in this section. We gather firm-level bond data from the

[4] Chen et al. (2006) provide alternative theoretical support to the tight link between expected equity excess returns and expected bond excess returns. Using a canonical asset pricing model, Chen et al. show that the simulated aggregate yield spread, despite an idiosyncratic component, shares very similar dynamics with the expected equity risk premium because of the dominating risk premium component in the yield spreads. Moreover, Krishnamurthy and Vissing-Jorgensen (2006) report that the yield spreads predict subsequent bond excess returns.

Lehman Brothers Fixed Income dataset, which provides detailed monthly information on corporate bonds including price, yield, coupon, maturity, modified duration, and convexity. This dataset, widely used in related research (e.g., Duffee, 1998, 1999; Elton et al., 2001), is cross-sectionally fairly deep and covers a reasonably long period. Following Elton et al., we only include nonmatrix prices because these represent true market quotes. We exclude bonds with maturity of less than one year and consider both callable and noncallable bond prices to retain as many bonds as possible. Our basic results are not affected when we only use noncallable bonds (not reported). We also restrict our analysis to bonds issued by nonfinancial firms.

We combine our bond data from Lehman Brothers with CRSP monthly data to get information on firm equity market capitalization, and then merge it with COMPUSTAT to get information on firm leverage. The merged dataset includes 1205 nonfinancial firms from January 1973 to March 1998. The Treasury yields for all maturities are obtained from the Federal Reserve Board. Following Collin-Dufresne et al. (2001), we compute yield spreads as the corporate bond yields minus the Treasury yields with matching maturities.

The Lehman Brothers Fixed Income dataset is the longest panel corporate bond dataset that has been widely used for research. The sample includes both investment grade and speculative grade bonds and tracks bonds up to default or maturity. About 67% of the firms in our sample have bonds that have investment grades, while 33% of the firms have speculative grades. The median book-to-market ratio of the firms in our sample is 0.69, which is close to the median of 0.75 for the CRSP/COMPUSTAT sample. The cross-sectional distribution in book-to-market is also similar across the two samples. The 5, 25, 75, and 95 percentiles of book-to-market in our sample are 0.10, 0.41, 1.04, and 2.03, respectively. The corresponding percentiles for the CRSP/COMPUSTAT sample are 0.08, 0.38, 1.30, and 2.95. This evidence shows that, similarly to the CRSP universe, our sample includes many value firms with low ratings. However, the median size of the firms in our sample is 1184 million dollars, considerably larger than the median of the CRSP sample, which is about 52 million dollars. Further, the 5, 25, 75, and 95 percentiles of market capitalization in our sample are 52, 353, 3388, and 15,104 million dollars, respectively. These percentiles are much larger than their counterparts in the CRSP sample, which are 3, 13, 251, and 2445 million dollars.

Because the bond issuers' data are heavily populated by large firms, it is natural to ask whether our inferences can be generalized to the broad CRSP/COMPUSTAT universe. In what follows, we show that the common factors of equity returns such as size, book-to-market, and momentum are prevalent in our sample and that these factors are largely comparable to those from the CRSP/COMPUSTAT universe. Moreover, we emphasize that the restriction of data, if anything, should make it more difficult for us to find meaningful cross-sectional patterns, especially on the value premium.

Table 1
Descriptive statistics of realized equity return factors from the CRSP/COMPUSTAT sample and from the Lehman Brothers Fixed Income Dataset/CRSP/COMPUSTAT merged sample

	Panel A: Our Sample				Panel B: CRSP/COMPUSTAT Sample			
	Mean	(t-stat)	Min	Max	Mean	(t-stat)	Min	Max
MKT	0.600	(2.27)	−23.29	16.50	0.581	(2.21)	−23.13	16.05
SMB	0.254	(1.54)	−11.92	12.63	0.111	(0.70)	−9.95	10.89
HML	0.310	(1.98)	−9.77	9.72	0.472	(3.09)	−9.77	8.37
WML	0.815	(3.87)	−14.99	16.28	0.919	(4.78)	−13.83	15.21

We use our Lehman Brothers Fixed Income/CRSP/COMPUSTAT matched sample to compute common factors based on realized stock returns and compare these factors with the corresponding common factors from the CRSP/COMPUSTAT universe. Four common factors are considered: *MKT*, *SMB*, *HML*, and *WML*, representing the market factor, size factor, value factor, and momentum factor, respectively. We report the results from our sample in Panel A and those from the CRSP/COMPUSTAT sample in Panel B. The t-statistics (t-stat) testing the null hypothesis that the average return of a given factor equals zero are adjusted for heteroscedasticity and autocorrelations. The values of mean, min, and max are in percent per month. The sample is from January 1973 to March 1998, limited by the Lehman Brothers dataset.

We first use our sample to construct common equity risk factors following the traditional practice of computing *ex post* return averages. We then compare the factors in our sample with their CRSP/COMPUSTAT counterparts, which are obtained from Kenneth French's Web site. Table 1 shows that the average return of the market factor in our sample is 0.60% per month, which is close to the average return of 0.58% for the Fama-French (1993) market factor over the same sample period. It is noteworthy that both the distributional minima and maxima of the two factors show nearly perfect matches, and the two market factors have a correlation coefficient of 0.99.

Table 1 also shows that the *HML* factor from our sample has an average return of 0.31% per month (t-statistic = 1.98), which is somewhat lower than the 0.47% (t-statistic = 3.09) for the Fama-French *HML* factor. This evidence is perhaps not surprising because our bond sample contains disproportionately more large firms than small firms. However, the *HML* factors across the two samples are highly correlated with a correlation of 0.80. Further, the momentum factor in our sample has an average return of 0.82% per month (t-statistic = 3.87), which is similar to the average return of 0.92% (t-statistic = 4.78) for the momentum factor from the CRSP sample. The two momentum factors are also highly correlated with a correlation of 0.92.

Figure 1 provides further evidence on the close match between equity factors from our sample and those from the CRSP/COMPUSTAT universe. The figure plots the monthly time series of the common factors from our sample along with those from the CRSP sample. The figure shows that the common factors from our sample track those from the CRSP sample remarkably well.

We also conduct monthly Fama-MacBeth (1973) cross-sectional regressions of realized equity excess returns onto the book-to-market, log size, and past 12-month returns—skipping the most recent month as in most of the momentum literature (e.g., Jegadeesh and Titman, 1993). We use both our sample and the CRSP sample. Table 2 shows a high level of consistency across the

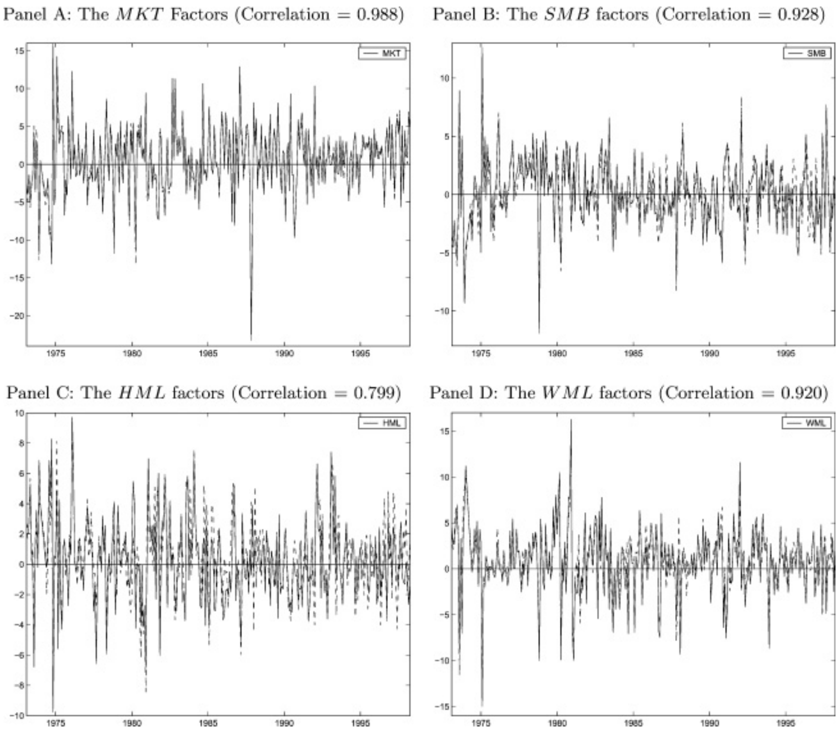

Figure 1
Comparison of Lehman Brothers/CRSP/COMPUSTAT and usual CRSP/COMPUSTAT
Comparison of realized equity return factors constructed from our Lehman Brothers/CRSP/COMPUSTAT merged sample and those constructed from the usual CRSP/COMPUSTAT merged sample. We plot the monthly time series of risk factors based on realized returns from our Lehman Brothers/CRSP COMPUSTAT merged sample against the corresponding Fama-French (1993) risk factors as well as the momentum factor from the usual CRSP/COMPUSTAT universe. The sample is from January 1973 to March 1998. The solid lines represent the Fama-French factors and the dotted lines represent our sample-based equity factors. Panel 1A: The *MKT* factors (correlation = 0.988). Panel 1B: The *SMB* factors (correlation = 0.928). Panel 1C: The *HML* factors (correlation = 0.799). Panel 1D: The *WML* factors (correlation = 0.920).

two samples. Size has significantly negative loadings with similar magnitudes across the two samples. Prior returns have significantly positive loadings, and their magnitudes are again similar across the two samples. The book-to-market ratio has a positive loading of 0.048% per month (t-statistic = 2.15) in our sample, which is lower than the loading of 0.197% (t-statistic = 4.05) in the CRSP/COMPUSTAT sample. Overall, however, the evidence suggests that our sample captures reasonably well the basic stylized facts from the CRSP universe over the same sample period.

3. Empirical Implementation

We present details of implementing our empirical framework developed in Section 1. In particular, we describe the steps used to calculate expected

Table 2
The cross-sectional variation of average realized returns in the Lehman Brothers Fixed Income Dataset/CRSP/COMPUSTAT merged sample and in the CRSP/COMPUSTAT sample

log(*ME*)	*BE*/*ME*	Past returns	R^2
Panel A: Our Lehman Brothers/CRSP/COMPUSTAT matched sample			
−0.101	0.048	0.012	2.07%
(−2.23)	(2.15)	(6.93)	
Panel B: CRSP/COMPUSTAT sample			
−0.127	0.197	0.008	1.27%
(−2.24)	(4.05)	(4.61)	

We conduct monthly Fama-MacBeth (1973) cross-sectional regressions of realized stock returns on firm characteristics using our matched Lehman Brothers/CRSP/COMPUSTAT sample. We also report corresponding results from the merged CRSP/COMPUSTAT sample for comparison. The regressors include book-to-market equity (*BE*/*ME*), log firm size (log(*ME*)), and past 12-month returns (skipping the most recent month). All the t-statistics are adjusted for heteroscedasticity and autocorrelations via GMM. Panel A reports the results for the Lehman Brothers/CRSP/COMPUSTAT matched sample from January 1973 to March 1998. Panel B reports the results for the CRSP/COMPUSTAT sample covering the same period. The slope coefficients are monthly in percent. The R^2s are the time-series median of the cross-sectional regression R^2s.

default loss rates, EDL_{it}, no-default yields, ERND_{it}, and the expected tax compensation, ETC_{it}. Following Collin-Dufresne et al. (2001), we calculate yield spreads, $R^i_{Bt} - r_t$, as the corporate bond yields from Lehman Brothers minus the Treasury bond yields from Federal Reserve Board with matching maturities. We also discuss the methods used to construct expected equity excess returns from these components.

3.1 Expected default loss rates

The expected default loss rate equals the default probability times the actual default loss rate. Moody's publishes information on annual default rates sorted by bond rating since 1970, and we use these data to construct expected default probabilities. The literature on default risk typically only uses the unconditional average default probability for each rating and ignores the time variation in expected default probabilities (e.g., Elton et al., 2001; Huang and Huang, 2003). Different from those papers, our approach is designed to capture time variation in default probability. To do so, we use the three-year moving average default probability from year $t-2$ to t as the one-year expected default probability for year t.[5] For the case of Baa and lower grade bonds, if the expected default probability in a given year is zero, we replace it with the lowest positive expected default probability in the sample for that rating. Doing so ensures that even in occasions of no actual default in three consecutive years, investors still anticipate positive default probabilities.

[5] The choice of a three-year window is based on the observation that there are many two-year but few three-year windows without default. While we want to keep the number of years in the window as small as possible, we also want to ensure that expected default probabilities are not literally zero. We have experimented with alternative ways of capturing the time-varying one-year expected default probabilities: (i) the average one-year default probability from year $t-3$ to $t-1$; (ii) the actual default probability itself at year t; (iii) the average default probability from year t to $t+2$; and (iv) the average default probability from year $t+1$ to $t+4$. Results from these alternative approaches (available from the authors) have no bearing on our conclusions.

Table 3 reports the constructed expected default probabilities from 1973 to 1998. With only very few exceptions, expected default probabilities decrease with bond ratings. More importantly, those default probabilities are typically higher during recessions than during expansions, highlighting the systematic nature of corporate defaults. For example, in the 1990–91 recession, the expected default probability of B3 bonds exceeds 25%, compared to only 5–8% during the late 1990s expansion.

To construct the expected default loss rate, EDL_{it}, we still need default loss rates. Following Elton et al. (2001), we use the recovery rate estimates provided by Altman and Kishore (1998). Their recovery rates for bonds rated by S&P are: 68.34% (for AAA bonds), 59.59% (AA), 60.63% (A), 49.42% (BBB), 39.05% (BB), 37.54% (B), and 38.02% (CCC). As in Elton et al., we assume the equivalence between ratings by Moody's and S&P (e.g., Aaa = AAA, ..., Baa = BBB, ..., Caa = CCC), and apply the same recovery rates.

3.2 Expected returns due to yield changes conditional on no-default

To calculate ERND_{it}, we need to calculate the expected yield changes conditional on no-default. We first show evidence on the mean-reversion of default probabilities, and then discuss our procedure of constructing expected yield changes based on the bond data.

Empirically, if a bond does not default, its default probability reverts to a long-term mean. In Table 4 we report one-year conditional default probabilities from 1 to 20 years, conditional on no-default in the previous year. The default probabilities are constructed using the one-year default transition matrices provided by Moody's and S&P Corporation. The first row of Table 4 shows that the default probability for Aaa bonds in the first year is zero, a pattern consistent with that reported in the first three columns of Table 3. Table 4 reports positive default probabilities for Aaa bonds starting from the second year. This is also consistent with the previous table, as some Aaa bonds can be downgraded and lower rated bonds have positive default probabilities. More importantly it is clear from Table 4 that, conditional on no-default, annual default probabilities increase over the years for bonds with an initially high rating, but they decrease for bonds with an initially low rating. For example, at year one, the one-year ahead default probability for Caa bonds is 22.28%. The one-year default probability then goes down to 19.28% in the second year and to 16.43% in the third year. Since mean-reverting default probabilities imply mean-reverting yields, high-quality bonds can have positive yield spreads even though their one-year default rates are close to zero.

Table 5 provides further evidence on the mean reversion of yield spreads. On an annual basis, we pool all bonds belonging to the same Moody's rating category in the Lehman Brothers dataset and study the changes in cumulative average ratings and yield spreads over the following three years. We assign numeric codes, from one to seven, to bonds rated from Aaa to Caa, with a lower number corresponding to a better rating. Table 5 shows that the ratings

Table 3
Three-year moving average annual default probability

Year	Aaa	Aa1	Aa2	Aa3	A1	A2	A3	Baa1	Baa2	Baa3	Ba1	Ba2	Ba3	B1	B2	B3
1973	0	0	0	0	0	0	0	0.343	0.343	0.343	0.430	0.430	0.430	4.920	4.920	4.920
1974	0	0	0	0	0	0	0	0.343	0.343	0.343	0.430	0.430	0.430	5.937	5.937	5.937
1975	0	0	0	0	0	0	0	0.343	0.343	0.343	0.633	0.633	0.633	5.547	5.547	5.547
1976	0	0	0	0	0	0	0	0.280	0.280	0.280	0.833	0.833	0.833	4.633	4.790	5.093
1977	0	0	0	0	0	0	0	0.280	0.280	0.280	0.867	0.867	0.867	3.427	3.583	3.887
1978	0	0	0	0	0	0	0	0.280	0.280	0.280	0.887	0.887	0.887	3.240	3.397	3.700
1979	0	0	0	0	0	0	0	0.280	0.280	0.280	0.707	0.707	0.707	3.240	3.397	3.700
1980	0	0	0	0	0	0	0	0.280	0.280	0.280	0.673	0.673	0.673	3.833	3.990	4.293
1981	0	0	0	0	0	0	0	0.280	0.280	0.280	0.450	0.450	0.450	3.527	3.683	3.987
1982	0	0	0	0	0.090	0.090	0.090	0.290	0.290	0.290	1.213	1.213	1.213	3.987	3.987	3.987
1983	0	0	0	0	0.090	0.090	0.090	0.290	0.290	0.290	1.213	1.213	1.940	2.643	5.633	8.270
1984	0	0	0	0	0.090	0.090	0.090	0.290	0.290	0.550	1.457	1.607	1.940	3.093	10.387	7.740
1985	0	0	0	0	0	0	0	0.280	0.280	0.540	0.673	1.223	2.270	3.750	12.053	11.557
1986	0	0	0	0	0	0	0	0.280	0.280	2.053	0.823	1.480	2.547	5.943	14.277	10.943
1987	0	0	0	0	0	0	0	0.280	0.280	1.793	1.680	1.260	3.387	5.640	9.460	13.433
1988	0	0	0	0	0	0	0	0.280	0.280	1.793	1.680	0.860	2.993	5.627	9.290	12.053
1989	0	0	0	0.467	0	0	0	0.280	0.453	0.543	1.650	1.067	3.417	5.170	6.493	13.213
1990	0	0	0	0.467	0	0	0	0.280	0.453	0.543	1.297	1.690	3.740	6.390	12.423	19.400
1991	0	0	0	0.467	0	0	0	0.440	0.453	0.543	1.507	1.690	6.173	6.957	14.370	25.633
1992	0	0	0	0	0	0	0	0.440	0.280	0.280	1.387	1.227	4.850	5.220	12.123	27.297
1993	0	0	0	0	0	0	0	0.440	0.280	0.280	0.767	0.430	3.793	3.463	6.413	21.480
1994	0	0	0	0	0	0	0	0.280	0.280	0.280	0.557	0.430	0.693	2.083	3.387	14.690
1995	0	0	0	0	0	0	0	0.280	0.280	0.280	0.557	0.430	1.020	3.190	4.993	7.877
1996	0	0	0	0	0	0	0	0.280	0.280	0.280	0.430	0.430	0.913	2.473	3.840	5.170
1997	0	0	0	0	0	0	0	0.280	0.280	0.280	0.430	0.430	0.873	2.183	3.120	4.957
1998	0	0	0	0	0	0	0	0.280	0.293	0.280	0.430	0.490	0.663	1.443	3.523	5.460

This table reports the three-year moving-average annual default rates (in percentage terms) for corporate bonds categorized by ratings, where the three-year window includes the current year and the preceding two years. The table is constructed using annual default rate data from Moody's from 1973 to 1998.

Table 4
Annual default probability conditional on no-default in the previous year

Year	Aaa	Aa	A	Baa	Ba	B	Caa
1	0	0	0.052	0.158	1.402	7.403	22.289
2	0.001	0.011	0.094	0.312	1.949	7.459	19.278
3	0.004	0.024	0.139	0.467	2.348	7.337	16.427
4	0.008	0.042	0.188	0.615	2.633	7.112	13.887
5	0.013	0.062	0.239	0.752	2.832	6.830	11.726
6	0.020	0.084	0.290	0.877	2.963	6.521	9.947
7	0.029	0.109	0.343	0.987	3.044	6.204	8.511
8	0.039	0.136	0.397	1.085	3.084	5.889	7.368
9	0.051	0.165	0.449	1.169	3.094	5.585	6.461
10	0.065	0.195	0.500	1.243	3.081	5.295	5.742
11	0.080	0.226	0.550	1.304	3.051	5.022	5.169
12	0.096	0.259	0.597	1.356	3.009	4.767	4.707
13	0.114	0.291	0.643	1.399	2.957	4.528	4.330
14	0.133	0.324	0.686	1.435	2.898	4.306	4.019
15	0.153	0.358	0.727	1.463	2.837	4.100	3.758
16	0.175	0.391	0.765	1.486	2.771	3.910	3.535
17	0.197	0.425	0.802	1.503	2.706	3.733	3.344
18	0.220	0.458	0.835	1.516	2.639	3.569	3.175
19	0.243	0.490	0.867	1.525	2.574	3.417	3.027
20	0.268	0.522	0.895	1.530	2.510	3.276	2.894

This table reports the annual default probability (in percentage terms) conditional on no-default in the preceding year. The table is constructed using the average one-year rating transition matrix of Moody's and that of S&P Corporation, reported in Table V of Elton et al. (2001). The sample is from 1973 to 1998.

Table 5
Evolution of ratings and yield spreads in corporate bonds

Year	Changes in	Aaa	Aa	A	Baa	Ba	B	Caa
1	Rating	0.093	0.087	0.046	0.017	0.005	−0.045	0.195
		(7.54)	(14.98)	(12.38)	(2.91)	(0.46)	(−8.28)	(−5.49)
	Yield spread	0.084	0.062	0.079	0.080	−0.024	0.556	−1.402
		(3.49)	(6.27)	(11.34)	(5.90)	(−0.75)	(7.69)	(−4.86)
2	Rating	0.175	0.181	0.088	0.047	−0.007	−0.114	−0.350
		(9.15)	(16.98)	(12.82)	(3.87)	(−0.42)	(−9.39)	(−6.04)
	Yield spread	0.177	0.153	0.160	0.104	0.213	1.075	−0.922
		(4.55)	(9.08)	(10.61)	(4.03)	(2.93)	(7.06)	(−1.98)
3	Rating	0.262	0.264	0.139	0.082	−0.019	−0.166	−0.500
		(10.19)	(18.63)	(13.79)	(4.49)	(−0.75)	(−9.40)	(−6.70)
	Yield spread	0.306	0.251	0.196	0.109	0.191	1.411	−1.461
		(6.21)	(9.05)	(11.43)	(2.92)	(1.87)	(6.75)	(−2.57)

We use Lehman Brothers Fixed Income data set to form cohorts of bonds with the same initial rating each year. Ratings from Aaa to Caa are assigned integer numbers from 1 to 7, with higher numbers indicating lower ratings. We report the average rating and yield spread changes for the same initial rating groups. Changes in yield spreads are in percent. The t-statistics adjusted for heteroscedasticity and autocorrelations via GMM are reported in parentheses. The sample is from January 1973 to March 1998.

of high-quality bonds (Aaa, Aa) indeed decline over time while their yield spreads increase. For example, the rating of Aa-rated bonds, conditional on no-default, increases by 0.087 after one year, where an increase of one indicates a full downgrade to grade A. Accordingly, the average yield spread of Aa bonds increases by 6.2 basis points. In contrast, the ratings of low-quality bonds (Caa) improve over time and their yield spreads decline.

We adopt the following three-step procedure to recover the yield change conditional on no-default, dY_{it}^{+}, from the data. First, we construct the cumulative default probability for each maturity using Table 4. For example, the conditional default probabilities for a bond initially rated Baa are 0.16% and 0.31% for the first two years, respectively. Assuming that the default rate is the same within a given year, the cumulative default probabilities are 0.16%, 0.16%, 0.47% (=0.16% + (1 − 0.16%) × 0.31%), and 0.47% for 0.5-year, 1-year, 1.5-year, and 2-year maturities, respectively.

Second, for each bond we calculate the expected cash flow, while taking into account the possibility of default. The expected cash flow for a particular coupon date before maturity is equal to

$$\text{coupon payment} \times [1 - \text{cumulative default probability} \times (1 - \text{recovery rate})].$$

We calculate the present value of the bond by discounting its expected cash flows by the corresponding Treasury yields with matching maturities.[6] After we obtain bond prices, we then calculate bond yields.

To illustrate this step, suppose that the Baa bond of the previous example has two years to maturity and the coupon rate is 8% with a face value of \$100. Also assume that the current Treasury yield, with annualized semiannual compounding, is 8% for a two-year maturity. Without default, the cash flows for the bond are \$4, \$4, \$4, and \$104 for the four half-year periods. The recovery rate for the Baa-rated bond is 49.42%. With default risk, the expected cash flows are $(1-0.16\% \times (1-49.42\%)) \times 4$, $(1-0.16\% \times (1-49.42\%)) \times 4$, $(1-0.47\% \times (1-49.42\%)) \times 4$, and $(1-0.47\% \times (1-49.42\%)) \times 104$, respectively. The present value, when we use the discount rate of 8%, is therefore \$99.77. With the promised cash flows of \$4, \$4, \$4, and \$104, and the price at \$99.77, the bond yield equals 8.12%.

Third, assume that the bond does not default within the first year. Conditional on that event, the bond maturity decreases by one year, and the second-year conditional default probability reported in Table 4 becomes the first-year default probability for the "new" bond. One can iterate the last two steps to calculate the price and yield for the new bond. Because conditional default probabilities of high-grade bonds will increase in the second year, bond prices will decrease and yields will increase, revealing a downgrading trend. Similarly, because conditional default probabilities will decrease for low-grade bonds in the second year, bond prices will increase and yields will decrease, revealing an upgrading trend. The yield difference between the last two steps can be used as a proxy for the yield change conditional on no-default within the first year. As expected, this yield change will be positive for high-grade bonds, but negative for low-grade bonds.

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[6] This is equivalent to calculating the fair price of the bond by a risk-neutral investor.

Consider again our numerical example. After one year, conditional on no-default, the new cumulative default rates will be 0.31% and 0.31% for the 0.5-year and 1-year maturities. Using our method to calculate the expected cash flows for this bond, we find the new price to be \$99.85 and the yield to be 8.17%. Thus, the bond yield will go up by five basis points due to the expected increase of default probability. The five basis points will be used as dY_{it}^{+} in calculating ERND_{it}, the expected return due to yield change conditional on no-default.

We have presented one approach to computing ERND_{it}. Different credit risk models may yield somewhat different estimates for ERND_{it}, depending on their assumption of the mean reversion of default rate. Nevertheless, we stress that ERND_{it} is on average very small (few basis points) and does not play a major role in affecting the magnitude of bond risk premium (more on this shortly).

3.3 Expected tax compensation

To calculate the expected tax compensation given by Equation (7), we follow Elton et al. (2001) and set the effective state and local tax rate to be 4% for all bonds. This completes the construction of the four components of the bond risk premium from Equation (6).

3.4 Elasticity of the equity value with respect to the bond value

From Proposition 1, expected equity excess returns can be computed as the product of expected bond excess returns and the elasticity of equity value with respect to bond value, $(\partial S_{it}/\partial B_{it})(B_{it}/S_{it})$. We calculate expected bond excess returns using the components described above. From Merton (1974), the unobservable elasticity $(\partial S_{it}/\partial B_{it})(B_{it}/S_{it})$ is a function of leverage, volatility, and the risk-free rate. We estimate this elasticity using the fitted component from regressing this ratio on leverage, stock volatility, and the risk-free rate.

Specifically, for each firm-month observation, we measure $\partial S_{it}/\partial B_{it}$ as the change in the market value of equity divided by the change in the market value of debt. We obtain the market value of debt by scaling the book value of debt by the weighted-average bond market prices. The leverage ratio, denoted LEV_{it}, is measured as the ratio of the market value of debt to the market value of equity. We measure the conditional volatility, σ_{it}, using a rolling window of 180 daily stock returns. The risk-free rate is the 30-day Treasury bill rate.

The pooled panel regression, excluding outliers, gives the following results:

$$\frac{\partial S_{it}}{\partial B_{it}}\frac{B_{it}}{S_{it}} = \underset{(9.60)}{11.08} - \underset{(-1.21)}{0.05} \times LEV_{it} + \underset{(4.11)}{133.03} \times \sigma_{it} - \underset{(-5.86)}{9.85} \times r_t + \varepsilon_{it}, \tag{8}$$

where t-statistics are reported underneath corresponding coefficients. From the t-statistics, the slopes are estimated reasonably precisely. Given the low R^2, however, our expected-return measure allows for quite a bit of noise. As we show in Section 4.2, the use of our measure in standard asset pricing tests dramatically increases cross-sectional R^2's compared to the traditional practice of using realized returns. The reason for the significant increase in predictive power is that our expected-return measure is still much more precise than the average realized equity returns. Later, in Section 5, we also conduct extensive robustness checks on the estimation of the elasticity $(\partial S_{it}/\partial B_{it})(B_{it}/S_{it})$. The results suggest that our inferences are robust to the regression specification that we use to estimate the equity-bond elasticity.

3.5 Discussion

Before presenting our results, we find it important to discuss potential limitations in our estimation strategy of expected returns. These limitations primarily arise from data restrictions that we face.

First, our calculations of expected default losses as the expected default probability times the loss rate are based on available historical data. Admittedly, this particular aspect of our approach is similar to the standard practice of using *ex post* observed information. It is well known, however, that the expected default loss is too small to explain the yield spreads. A much larger portion of the yield spreads is related to systematic risk (e.g., Huang and Huang, 2003; Chen et al., 2006).[7] By restricting the use of historical data to a small portion of the yield spreads, we retain crucial information on the forward-looking risk premiums embedded in bonds. Because we can back out a significant fraction of that information, our approach can deliver new insights on the pricing of market securities.

Second, we also need the elasticity of equity with respect to debt to construct the expected equity return [see Equation (1)]. When estimating this component of our expected-return measure, we again use historical data. We observe, however, that the equity-bond elasticity serves as a time-varying coefficient in the first-order approximation of equity return variation, and is not forward looking.[8] Therefore, even though we need to estimate this unobservable coefficient using historical data, its potential drawback is probably less severe than using average realized returns as a proxy for expected returns.

Finally, we stress that the crux of our argument is not to avoid historical data, but to move away from assuming that average realized returns converge

[7] According to Moody's annual report, the average four-year cumulative default rate for Baa-rated bonds is about 1.55% or 0.39% per annum. If the loss rate is about 51%, as in Altman and Kishore (1998), the expected annual default loss is then about $0.39\% \times 0.51\% = 0.20\%$. Therefore, a risk-neutral bondholder only requires a yield spread of about 20 basis points per annum. Adding a possible risk premium from a structural model, Huang and Huang (2003) find that the implied yield spread is about 32 basis points. In contrast, the average spread between Baa yield and Treasury bill rate is about 158 basis points per annum.

[8] Our approach is similar in spirit to the Campbell-Shiller (1988) log-linearization of the dividend yield. Campbell and Shiller take the first-order approximation of the dividend yield and assume that the partial derivative is known locally (set equal to the steady-state constant globally). The stochastic component of the dividend yield is only the linearized component. When taking expectations, Campbell and Shiller treat the partial derivative as a part of the parameter ρ outside of the expectation operator. In the actual implementation, ρ is set to equal its historical average.

Table 6
Descriptive statistics of yield spreads and expected bond excess returns by bond ratings

Panel A: Yield spreads								
Rating	Mean	SD	Min	Max	ρ_1	ρ_2	ρ_6	ρ_{12}
Aaa	0.850	0.52	0.30	3.09	0.95	0.06	−0.11	0.00
Aa	0.897	0.34	0.42	2.29	0.96	0.14	−0.06	−0.08
A	1.093	0.35	0.61	2.51	0.95	0.04	−0.12	−0.02
Baa	1.805	0.64	0.46	4.31	0.93	0.08	−0.02	−0.03
Ba	2.967	0.84	1.74	6.79	0.80	0.29	−0.03	0.09
B	5.494	2.37	2.67	18.46	0.95	0.12	−0.06	−0.01

Panel B: Expected bond excess returns								
Rating	Mean	SD	Min	Max	ρ_1	ρ_2	ρ_6	ρ_{12}
Aaa	0.464	0.49	−0.03	2.65	0.94	0.09	−0.10	0.02
Aa	0.465	0.33	0.01	1.86	0.95	0.13	−0.07	−0.09
A	0.607	0.33	0.13	2.05	0.94	0.03	−0.11	−0.04
Baa	0.930	0.57	−0.02	3.38	0.92	0.07	−0.02	−0.05
Ba	1.053	0.55	0.11	2.92	0.78	0.22	−0.07	0.04
B	2.238	1.41	0.07	10.98	0.80	0.18	0.01	−0.02

This table reports mean, standard deviation (SD), minimum (min), maximum (max), and partial autocorrelations of orders one (ρ_1), two (ρ_2), six (ρ_6), and 12 (ρ_{12}), of yield spreads (Panel A) and expected bond risk premiums (Panel B) for bonds rated from B to Aaa. The mean, SD, min, and max are reported in annualized percentage terms. The sample is from January 1973 to March 1998.

to expected returns in finite samples. Indeed, we note that the use of historical data to estimate parts of expected returns measures is common practice in the literature. Previous examples include Blanchard (1993); Gebhardt et al. (2001); Fama and French (2002); and Pastor et al. (2005). Fama and French, in particular, use the historical average dividend growth as the expected rate of capital gain and measure the equity premium as the sum of the expected rate of capital gain and the average dividend yield. As we show in Section 5, our results are robust to alternative empirical specifications we use to estimate the equity-bond elasticity. We also find that most of the results on the cross section of expected equity returns are already present in the cross section of expected bond returns, meaning that the equity-bond elasticity plays a less important role than the yield spreads in Equation (1).

4. Empirical Results

We now conduct standard asset pricing tests using our measure of expected returns. Section 4.1 uses the time series Fama and French (1993) portfolio approach, and Section 4.2 uses the Fama and MacBeth (1973) cross-sectional regression approach.

4.1 Time series analysis

As a precursor to our analysis of equity returns, Table 6 reports summary statistics of yield spreads and constructed bond risk premiums for B- through

Aaa-rated bonds. Because information on time-varying default rates for bonds rated Caa or lower is not available, we disregard these bonds, which are less than 1% of all bonds. Table 6 shows that yield spreads and expected bond excess returns increase as the bond rating decreases. The bond risk premium for Aaa-rated bonds is on average 0.46% per annum, and it goes up to 2.24% for B-rated bonds. The evidence suggests that lower-graded bonds are riskier than higher-graded bonds. Yield spreads and expected bond excess returns are also highly persistent. The first-order autocorrelation ranges from 0.78 to 0.95, but the second-order partial autocorrelation quickly drops to the 0.03–0.29 range. The evidence suggests that we only need to include a few lags to control for the impact of autocorrelations in our asset pricing tests.

4.1.1 Common factors in equity returns. Our focus is on the common factors in equity returns. We define the market factor as the value-weighted average equity excess returns for all firms in our sample. The size (*SMB*) and book-to-market (*HML*) factors are constructed using the Fama and French (1993) two-by-three sorting in size and book-to-market. To construct the momentum factor, we sort stocks each month on the basis of their realized equity returns in the past 12 months into winners (top 30%), medium (middle 40%), and losers (bottom 30%) categories. We skip one month to avoid potential microstructure biases and calculate expected portfolio returns over the subsequent 12 months. The momentum factor is defined as the winner-minus-loser (*WML*) portfolio return.

Panel A of Table 7 reports descriptive statistics of the common equity factors. The expected return for the market factor is on average 3.39% per annum. This expected return is much lower than the average realized market excess return of 7.54% in the same period. However, our evidence is consistent with Fama and French (2002), who document that the expected market returns are much lower than the average realized market returns in the post war sample period. The expected size and value premiums are on average 3.81% and 1.45%, respectively, and the momentum factor earns on average a negative expected return of −0.79%. All the expected returns on average are significantly different from zero at the 1% significance level.

Panel B of Table 7 reports the correlation matrix of the expected returns of the equity factors. The expected market factor is positively correlated with the expected size factor (correlation = 0.33) and the book-to-market factor (correlation = 0.48), but it is negatively correlated with the momentum factor (correlation = −0.15). In untabulated results, we also find that the expected return of the bond market portfolio is on average 0.40% per annum (t-statistic = 9.73). Panel C reports the market regressions of *SMB*, *HML*, and *WML*. The unconditional alphas of *SMB* and *HML* are 2.85% and 0.06% per annum (t-statistics = 8.12 and 0.20, respectively). *WML* has an insignificant unconditional alpha of −0.33% (t-statistic = −1.04).

Table 7
Descriptive statistics of expected returns of common equity factors

Panel A: Summary statistics

	Mean	(t-stat)	Min	Max	ρ_1	ρ_2	ρ_6	ρ_{12}
MKT	3.391	(14.07)	0.09	15.12	0.95	0.09	0.06	−0.25
SMB	3.806	(18.10)	−0.91	16.19	0.86	−0.03	−0.11	0.05
HML	1.448	(6.82)	−4.20	13.97	0.86	0.11	0.10	−0.06
WML	−0.792	(−3.68)	−14.65	5.36	0.82	0.08	−0.01	−0.00

	Panel B: Cross correlations				Panel C: Market regressions			
	MKT	*SMB*	*HML*	*WML*	α	(t_α)	β	(t_β)
MKT	1	0.33	0.48	−0.15				
SMB		1	0.03	−0.53	2.846	(8.12)	0.27	(3.38)
HML			1	−0.14	0.056	(0.20)	0.40	(4.79)
WML				1	−0.330	(−1.04)	−0.14	(1.06)

This table reports summary statistics of expected returns of common equity factors, including *MKT* (market excess return), *SMB*, *HML*, and *WML* (the momentum factor). Panel A reports mean, t-statistic (testing the null that the mean equals zero), minimum, maximum, and partial autocorrelations of orders one (ρ_1), two (ρ_2), six (ρ_6), and 12 (ρ_{12}). Panel B reports the results of market regressions for *SMB*, *HML*, and *WML* including the intercepts (α), the slopes (β), and their t-statistics (testing the null that the coefficients are individually zero). Finally, Panel C reports the correlation matrix for these four factors. The numbers of mean, min, max, and α are in annualized percent. All cross correlations in Panel C are significant at the 1% significance level. All the t-statistics are adjusted for heteroscedasticity and autocorrelations via GMM. The sample is from January 1973 to March 1998.

4.1.2 Business cycle properties. We also investigate the cyclical properties of the expected returns of the four equity factors during the 1973–1998 period. Following the empirical business cycle literature (e.g., Stock and Watson, 1999, Table 2), Table 8 reports the cross correlations with different leads and lags for the expected returns with the cyclical component of the real industrial production index. Both the nominal index of industrial production and the inflation data are obtained from the Federal Reserve Bank of St. Louis. We follow Stock and Watson by removing from the output series its long-run growth component as well as those fluctuations that occur over periods shorter than a business cycle, which arise from temporary factors such as measurement errors. This task is achieved by passing the raw industrial production index through the Hodrick and Prescott (1997) filter. Following Hodrick and Zhang (2001), we set the monthly smooth parameter in the Hodrick-Prescott filter to be 6400.

Table 8 reports several interesting patterns. First, the expected market risk premium is negatively correlated with the business cycle. The cross correlations are mostly negative and significant across different leads and lags. This evidence suggests that the aggregate expected return is countercyclical: investors demand a higher risk premium in recessions than they do in booms. This finding speaks to the criticism voiced by Elton (1999) that *ex post* equity returns go down in recessions and thus fail to capture investors' heightened required returns from risky assets in uncertain environments.

Second, both the expected size premium and the expected value premium are significantly negatively correlated with the business cycle. Investors seem to

Table 8
Cross-correlations with cyclical component of industrial production

	Corr(r_t, y_{t+k})												
$k =$	−24	−12	−6	−3	−2	−1	0	1	2	3	6	12	24
MKT	0.28	0.05	−0.09	−0.18	−0.22	−0.26	−0.28	−0.29	−0.28	−0.27	−0.17	0.05	0.19
	(0.00)	(0.43)	(0.14)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.38)	(0.00)
SMB	0.03	0.14	−0.09	−0.29	−0.35	−0.42	−0.46	−0.49	−0.48	−0.45	−0.32	−0.04	0.32
	(0.61)	(0.02)	(0.14)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.53)	(0.00)
HML	0.03	0.00	−0.01	−0.09	−0.13	−0.15	−0.18	−0.19	−0.18	−0.16	−0.09	0.12	0.06
	(0.66)	(0.97)	(0.90)	(0.13)	(0.04)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)	(0.17)	(0.06)	(0.39)
WML	−0.19	−0.21	−0.10	0.07	0.14	0.20	0.25	0.28	0.31	0.32	0.28	0.12	−0.09
	(0.00)	(0.00)	(0.11)	(0.27)	(0.02)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.04)	(0.17)

This table reports the cross correlations of expected returns of equity factors with the cyclical component of the industrial production index, denoted corr(r_t, y_{t+k}), for different leads and lags, k. The cyclical component of the real industrial production index is estimated by passing the raw series through the Hodrick and Prescott (1997) filter. p-values of the cross-correlations are reported in parentheses. The sample is from January 1973 to March 1998.

Table 9
Regressing realized equity factor returns on their expected returns

	Panel A: *MKT*				Panel B: *SMB*			
	6-month	12-month	24-month	36-month	6-month	12-month	24-month	36-month
Slope	0.396	0.170	−0.322	0.162	0.528	0.248	−0.625	−1.115
p-value	(0.61)	(0.00)	(0.00)	(0.00)	(0.87)	(0.00)	(0.00)	(0.00)
	Panel C: *HML*				Panel D: *WML*			
	6-month	12-month	24-month	36-month	6-month	12-month	24-month	36-month
Slope	0.150	0.392	0.737	−0.150	0.003	0.004	0.010	0.011
p-value	(0.05)	(0.02)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)

This table reports predictive regressions of realized equity factor returns including market excess return, *MKT* (Panel A), *SMB* (Panel B), *HML* (Panel C), and *WML* (Panel D) on their respective constructed expected returns. We consider four different predictive horizons: (i) six-month; (ii) 12-month; (iii) 24-month; and (iv) 36-month horizons. We report the slope coefficients, and test the convergence of average realized equity returns and the constructed expected returns as follows. The null hypothesis is that the slope equals 1/2 for the regressions with the six-month horizon, the slope equals one for the 12-month horizon, two for the 24-month horizon, and three for the 36-month horizon. We report the p-values in parentheses for these tests. The p-values are computed using standard errors adjusted for heteroscedasticity and autocorrelations via GMM. The sample is from January 1973 to March 1998.

perceive small and value stocks as riskier securities than big and growth stocks, respectively, charging countercyclical risk premiums for holding these riskier assets. Finally, in contrast to the other equity factors, the expected momentum return is strongly procyclical because the cross correlations between expected momentum returns and output are positive and significant at most leads and lags.

4.1.3 Do average realized returns converge to expected returns? For the most part, the answer is no. Specifically, we run predictive regressions of realized cumulative equity factor returns on the expected factor returns constructed from yield spreads. Four different horizons are considered: 6-month, 12-month, 24-month, and 36-month. We test the hypothesis that average realized equity returns converge to expected returns by testing the null that the regression slope equals 1/2 in the six-month horizon, one in the 12-month horizon, two in the 24-month horizon, and three in the 36-month horizon.

Table 9 reports the results. From Panel A, our proxy for the expected market risk premium is closely related to the average market excess return in the short six-month horizon. The slope from regressing future realized returns on the expected return is not reliably different from 1/2 in the six-month horizon. However, the expected-return measure diverges significantly from the average realized returns over the longer horizons. The null hypotheses of convergence at the one-, two-, and three-year horizons are all rejected at the 1% significance level. We interpret this evidence as suggesting the importance of time-varying expected returns. Because expected market excess returns vary over long horizons, we should be cautious with the notion that averaged realized returns will converge to the underlying return expectations in long-horizon asset pricing tests.

Panel B of Table 9 reports a similarly poor convergence between the expected returns and the average realized returns of *SMB*. The slope from regressing realized returns on expected returns is not statistically different from the theoretical prior in the six-month horizon. But the null of convergence is strongly rejected at longer horizons. This evidence suggests an explanation for the fact that we document a significantly positive expected size premium in the 1973–1998 period, whereas studies using realized returns report a weak or even negative size premium in comparable periods.

For example, Schwert (2003) reports that the alpha of Dimensional Fund Advisors Small Company Portfolio is 0.20% per month (standard error = 0.30%). Based on the evidence that the size premium has weakened or disappeared after its discovery by Banz (1981), Schwert argues that the size anomaly is more apparent than real. Our evidence suggests that the disappearance of the size effect could result from the high volatility of realized returns. In particular, our estimate of the *SMB* alpha (0.24% per month) is well within the one-standard-error bound estimated by Schwert.

Panels C and D of Table 9 reveal a significant divergence between the expected and the average realized returns for both *HML* and *WML*. At all horizons the hypothesis that the expected risk premium is equal to the realized risk premium is rejected. Because asset pricing models are tested on the prediction of trade-offs between risk and expected return, these results highlight the importance of reexamining the standard inferences based on the *ex post* averaged returns.

4.2 Cross-sectional regressions

Using the Fama and MacBeth (1973) cross-sectional regressions, we perform both covariance- and characteristic-based tests to examine the determinants of the cross section of expected returns.

4.2.1 Covariance-based cross-sectional tests. Our covariance-based tests are conducted in two steps. In the first step, for each individual firm and month, we run a time series regression of the realized equity returns in the past 60 months (with at least 24 months of data available) on the realized return of the three Fama-French factors and the momentum factor. For comparison, we also estimate the loadings from regressing our expected-return measures on the expected equity factor risk premiums.[9]

In the second step, we run cross-sectional regressions, month by month, of firm-specific, expected equity returns on the factor loadings estimated earlier. The time series averages of the coefficients are regarded as the risk premiums associated with the loadings. We use GMM to adjust the standard errors of the coefficients for heteroscedasticity and autocorrelations. The standard errors

[9] We thank the referee for suggesting this method to us.

of the time series of the estimated coefficients are then used to compute the t-statistics.

The null hypothesis in our cross-sectional tests is the CAPM. We also use size, book-to-market, and prior returns to test for model misspecification. In doing so, we implicitly assume that our constructed risk premiums are unbiased measures of the true risk premiums. In other words, the measurement errors in the risk premiums have a mean of zero—if there were no measurement errors, then under the null hypothesis the regression R^2 should be one.

Panel A of Table 10 reports the central result of our paper. When the market beta estimated from realized returns is used alone, the slope is a positive 3.81% per annum (t-statistic = 7.45). More importantly the market beta is significantly priced even in the presence of *SMB* and *HML* loadings, although the risk premium of the market beta drops to 2.31% per annum (t-statistic = 3.48). Using the market beta estimated from expected returns reduces the slope in the univariate regression to 1.15%, but it remains significant (t-statistic = 3.59). However, the premium estimate for the market risk in the presence of the *SMB* and *HML* loadings is quantitatively similar to the estimate when the loadings are estimated from realized returns.

In untabulated results using realized returns as the dependent variable in cross-sectional regressions, we find that the market beta has an insignificant slope of 0.35% per month (t-statistic = 0.85) when used alone. The slope drops slightly to 0.33% (t-statistic = 0.69) when we control for size, book-to-market, and past returns. In contrast to these standard inferences based on realized returns as in Fama and French (1992), our tests based on expected returns show that stocks with higher market betas earn significantly higher expected excess return, consistent with standard asset pricing theories.

Panel A of Table 10 also shows that the *SMB* and *HML* betas are significantly priced in the cross section of expected returns. The magnitudes of their premium estimates are largely in line with those from the time series tests reported in Table 7. Further, the *WML* loading is insignificantly priced, also consistent with the previous evidence in Table 7.

4.2.2 Characteristic-based cross-sectional tests. In Panel B of Table 10, we retain the loading on the market factor, but replace the other loadings with firm characteristics. Specifically, we use the logarithm of size, the book-to-market ratio, and the prior equity return to replace their respective factor loadings. The market beta estimated from the realized returns is again significantly priced at 2.80% per annum (t-statistic = 4.74). Size has a negative loading of −2.38%, and book-to-market has a positive loading of 1.31%, both of which are significant. Using the market beta estimated from the expected returns yields a risk premium of 1.02% per annum (t-statistic = 3.30). The pricing results on size and book-to-market are similar to those in Panel A. However, past returns are associated with significantly negative loadings, meaning that firms with

Table 10
The cross section of expected equity excess returns

Panel A: Covariance-based tests

	Model 1	Model 2	Model 3
β_{MKT}	3.865	2.310	2.076
	(7.45)	(3.40)	(3.24)
β_{SMB}		3.405	3.705
		(12.30)	(10.77)
β_{HML}		1.389	1.361
		(3.45)	(3.35)
β_{WML}			−0.185
			(−0.30)
R^2	0.03	0.13	0.16
	Model 4	Model 5	Model 6
$\widetilde{\beta}_{MKT}$	1.147	2.222	1.917
	(3.59)	(4.60)	(4.02)
$\widetilde{\beta}_{SMB}$		1.124	1.148
		(2.63)	(2.52)
$\widetilde{\beta}_{HML}$		0.575	0.508
		(2.27)	(2.08)
$\widetilde{\beta}_{WML}$			1.346
			(1.82)
R^2	0.11	0.43	0.52

Panel B: Characteristic-based tests

	Model 7	Model 8
β_{MKT}	2.081	1.742
	(4.74)	(4.49)
log(*ME*)	−2.379	−2.199
	(−14.60)	(−14.98)
BE/*ME*	1.307	1.270
	(3.96)	(4.29)
Past Returns		−6.810
		(−7.75)
R^2	0.22	0.27
	Model 9	Model 10
$\widetilde{\beta}_{MKT}$	1.021	1.107
	(3.30)	(3.55)
log(*ME*)	−1.942	−1.742
	(−7.83)	(−8.75)
BE/*ME*	0.631	0.642
	(3.32)	(3.57)
Past Returns		−5.710
		(−5.92)
R^2	0.30	0.35

We estimate monthly Fama-MacBeth (1973) cross-sectional regressions for the following 10 specifications:

$$\text{Model 1}: \quad E_t[R^i_{St} - r_t] = \gamma_{0t} + \gamma_{1t}\beta^i_{MKT,t} + \epsilon^i_t$$
$$\text{Model 2}: \quad E_t[R^i_{St} - r_t] = \gamma_{0t} + \gamma_{1t}\beta^i_{MKT,t} + \gamma_{2t}\beta^i_{SMB,t} + \gamma_{3t}\beta^i_{HML,t} + \epsilon^i_t$$
$$\text{Model 3}: \quad E_t[R^i_{St} - r_t] = \gamma_{0t} + \gamma_{1t}\beta^i_{MKT,t} + \gamma_{2t}\beta^i_{SMB,t} + \gamma_{3t}\beta^i_{HML,t} + \gamma_{4t}\beta^i_{WML,t} + \epsilon^i_t$$
$$\text{Model 4}: \quad E_t[R^i_{St} - r_t] = \gamma_{0t} + \gamma_{1t}\widetilde{\beta}^i_{MKT,t} + \epsilon^i_t$$
$$\text{Model 5}: \quad E_t[R^i_{St} - r_t] = \gamma_{0t} + \gamma_{1t}\widetilde{\beta}^i_{MKT,t} + \gamma_{2t}\widetilde{\beta}^i_{SMB,t} + \gamma_{3t}\widetilde{\beta}^i_{HML,t} + \epsilon^i_t$$
$$\text{Model 6}: \quad E_t[R^i_{St} - r_t] = \gamma_{0t} + \gamma_{1t}\widetilde{\beta}^i_{MKT,t} + \gamma_{2t}\widetilde{\beta}^i_{SMB,t} + \gamma_{3t}\widetilde{\beta}^i_{HML,t} + \gamma_{4t}\widetilde{\beta}^i_{WML,t} + \epsilon^i_t$$
$$\text{Model 7}: \quad E_t[R^i_{St} - r_t] = \gamma_{0t} + \gamma_{1t}\beta^i_{MKT,t} + \gamma_{2t}\log(ME^i_t) + \gamma_{3t}BE^i_t/ME^i_t + \epsilon^i_t$$
$$\text{Model 8}: \quad E_t[R^i_{St} - r_t] = \gamma_{0t} + \gamma_{1t}\beta^i_{MKT,t} + \gamma_{2t}\log(ME^i_t) + \gamma_{3t}BE^i_t/ME^i_t + \gamma_{4t}(\text{PastReturns})^i_t + \epsilon^i_t$$
$$\text{Model 9}: \quad E_t[R^i_{St} - r_t] = \gamma_{0t} + \gamma_{1t}\widetilde{\beta}^i_{MKT,t} + \gamma_{2t}\log(ME^i_t) + \gamma_{3t}BE^i_t/ME^i_t + \epsilon^i_t$$
$$\text{Model 10}: \quad E_t[R^i_{St} - r_t] = \gamma_{0t} + \gamma_{1t}\widetilde{\beta}^i_{MKT,t} + \gamma_{2t}\log(ME^i_t) + \gamma_{3t}BE^i_t/ME^i_t + \gamma_{4t}(\text{PastReturns})^i_t + \epsilon^i_t,$$

where $E_t[R^i_{St} - r_t]$ is the firm-level expected equity excess return for firm i at the beginning of time t, and r_t is the risk-free rate. The sample is from January 1973 to March 1998. We estimate betas in two ways. First, we estimate these betas using 60-month (or at least 24-month) rolling regressions of the realized equity excess returns on the realized excess returns of the Fama-French (1993) three factors and *WML*. These betas estimated from realized returns are denoted β_{MKT}, β_{SMB}, β_{HML}, and β_{WML}. Second, we also estimate the betas from 60-month (or at least 24-month) rolling regressions of the constructed expected excess returns on the constructed expected Fama-French three factors and *WML*. These betas are denoted $\widetilde{\beta}_{MKT}$, $\widetilde{\beta}_{SMB}$, $\widetilde{\beta}_{HML}$, and $\widetilde{\beta}_{WML}$. We also use size (log(*ME*)), book-to-market equity (*BE*/*ME*), and past 12-month returns as explanatory variables. The t-statistics reported in parentheses are adjusted for heteroscedasticity and autocorrelations via GMM. The point estimates of the intercepts and slopes are in annualized percentage terms. The R^2s are the time-series median of the cross-sectional regression R^2s.

higher prior returns earn lower expected returns than firms with lower prior returns.

Intriguingly, the median cross-sectional R^2 is 35% when all the firm characteristics are included in the regression. In contrast, Table 2 shows that the median R^2 in a similar cross-sectional regression in the CRSP/COMPUSTAT universe is only 1.27%. Therefore, using expected returns leads to a dramatic increase of explanatory power in cross-sectional regressions. The reason is likely that our expected-return measure is much more precise than the realized equity returns. We show in Section 5.5 that using expected returns to replace realized returns also yields higher R^2s, when we use portfolios (instead of individual stock returns) as testing assets in cross-sectional regressions.

5. Robustness

We now demonstrate the robustness of our conclusions along several important dimensions.

5.1 Rolling panel regressions

When estimating the elasticity of equity with respect to bond in Equation (8), we use a panel regression that includes the entire sample. The underlying rationale is that the slopes are constant over the sample period and using more data to fit the model would improve the precision of our estimations. However, one potential criticism to our approach is that it uses future information that is unknown to market investors. To check the robustness of our results along this dimension, we rerun the same panel regression, but using the data from the previous 12 months of cross sections. This procedure provides conditional estimates of the elasticity of equity with respect to debt for each month using only prior information. We then reconstruct expected equity excess returns, and conduct the same time series and cross-sectional analysis as in Section 4.

The results in Table 11 are similar to those in Tables 7 and 10. For example, the expected risk premiums for *MKT*, *SMB*, *HML*, and *WML* are 3.74%, 3.47%, 1.91%, and −0.67%, respectively, which are close to their counterparts in the benchmark estimation. The expected risk premiums for the three Fama-French (1993) factors continue to be countercyclical, whereas the momentum factor continues to be procyclical. The cross-sectional regressions are also similar to those in the benchmark estimation. In particular, the market beta continues to be significantly priced in the cross section.

5.2 Alternative specification for the equity-bond elasticity

One can argue that, although motivated from the theoretical work of Merton (1974), our baseline specification for the equity-bond elasticity in Equation (8) is too parsimonious. To this end, we perform a "kitchen-sink" regression of the elasticity by including the log market value, $\log(ME_{it})$, book-to-market, BE_{it}/ME_{it}, and prior 12-month returns, r_{it}^{12}, as additional explanatory variables. We choose these variables because they are strong cross-sectional determinants of stock returns.

Table 11
Using rolling panel regressions to estimate the equity-bond elasticity

Panel A: Summary statistics	Mean	(t-stat)	Min	Max
MKT	3.739	(10.47)	0.08	21.44
SMB	3.465	(18.06)	−1.45	12.10
HML	1.908	(7.00)	−3.83	17.53
WML	−0.647	(−2.75)	−21.14	7.73

Panel B: Correlations with cycle	−1	0	1
MKT	−0.18	−0.18	−0.16
SMB	−0.32	−0.35	−0.37
HML	−0.14	−0.16	−0.16
WML	0.14	0.20	0.22

Panel C: Covariance-based tests

	Model 1	Model 2	Model 3
β_{MKT}	3.219	1.903	1.614
	(8.45)	(3.03)	(2.54)
β_{SMB}		3.095	3.225
		(11.04)	(11.22)
β_{HML}		1.326	1.340
		(3.22)	(3.24)
β_{WML}			0.179
			(0.37)
R^2	0.03	0.12	0.15
	Model 4	**Model 5**	**Model 6**
$\widetilde{\beta}_{MKT}$	2.325	3.586	3.759
	(5.74)	(5.00)	(5.46)
$\widetilde{\beta}_{SMB}$		1.344	1.096
		(2.77)	(2.38)
$\widetilde{\beta}_{HML}$		0.344	0.427
		(1.12)	(1.37)
$\widetilde{\beta}_{WML}$			−0.150
			(−0.27)
R^2	0.09	0.51	0.55

Panel D: Characteristic-based tests

	Model 7	Model 8
β_{MKT}	1.755	1.556
	(4.65)	(4.24)
log(ME)	−2.059	−1.893
	(−15.53)	(−16.02)
BE/ME	1.406	1.288
	(4.39)	(4.72)
Past returns		−5.743
		(−7.08)
R^2	0.21	0.25
	Model 9	**Model 10**
$\widetilde{\beta}_{MKT}$	1.970	2.064
	(4.93)	(5.06)
log(ME)	−1.851	−1.660
	(−8.71)	(−9.92)
BE/ME	0.844	0.866
	(4.24)	(4.71)
Past returns		−5.576
		(−6.14)
R^2	0.26	0.32

We estimate the equity-bond elasticity using 12-month rolling panel regressions. As explained in Section 2, expected equity excess returns are calculated as the product of expected bond excess returns and the new elasticity of equity with respect to bond. Panel A reports the summary statistics for the expected returns of *MKT*, *SMB*, *HML*, and *WML* factors. Panel B reports the cross correlations of the expected factor returns with the cyclical component of industrial production. All correlations are significant at the 5% significance level. Panel C reports the monthly Fama-MacBeth (1973) cross-sectional regressions of firm-level expected excess returns on the *MKT* beta, the *SMB* beta, the *HML* beta, and the *WML* beta, separately and jointly. Panel D reports the cross-sectional regressions of expected excess returns on the *MKT* beta, size, book-to-market equity, and past 12-month returns. The model specifications are the same as in Table 10 (see the caption of that table for detailed definitions). The t-statistics reported in parentheses are adjusted for heteroscedasticity and autocorrelations via GMM. All estimates are in annualized percentage terms. The R^2s are the time-series median of the cross-sectional regression R^2s.

The expanded regression is estimated as

$$\frac{\partial S_{it}}{\partial B_{it}}\frac{B_{it}}{S_{it}} = \underset{(3.02)}{8.04} - \underset{(-0.75)}{0.04} \times LEV_{it} + \underset{(4.41)}{156.95} \times \sigma_{it} - \underset{(-5.52)}{9.73} \times r_t + \underset{(1.03)}{0.27} \times \log(ME_{it}) + \underset{(2.03)}{0.05} \times \frac{BE_{it}}{ME_{it}} + \underset{(1.98)}{0.03} \times r_{it}^{12} + \varepsilon_{it}. \quad (9)$$

Relative to our baseline specification in Equation (8), the equity volatility and the risk-free rate continue to be important, whereas the leverage ratio continues to be less important. Book-to-market and prior returns add some explanatory power in the regression, but size does not.

Table 12
Using additional explanatory variables in estimating the equity-bond elasticity

Panel A: Summary statistics

	Mean	(t-stat)	Min	Max
MKT	3.580	(13.29)	0.07	16.94
SMB	3.232	(17.52)	−1.21	14.53
HML	1.591	(7.69)	−4.46	13.66
WML	−0.311	(−1.62)	−13.34	5.29

Panel B: Correlations with cycle

	−1	0	1
MKT	−0.24	−0.25	−0.26
SMB	−0.36	−0.40	−0.43
HML	−0.13	−0.15	−0.16
WML	0.19	0.24	0.27

Panel C: Covariance-based tests

	Model 1	Model 2	Model 3
β_{MKT}	3.798	2.354	2.188
	(7.79)	(3.44)	(3.44)
β_{SMB}		3.246	3.505
		(11.01)	(10.02)
β_{HML}		1.349	1.332
		(3.32)	(3.30)
β_{WML}			−0.201
			(−0.33)
R^2	0.03	0.13	0.15

	Model 4	Model 5	Model 6
$\widetilde{\beta}_{MKT}$	1.170	1.501	1.118
	(3.21)	(3.08)	(2.22)
$\widetilde{\beta}_{SMB}$		2.287	2.020
		(5.40)	(5.68)
$\widetilde{\beta}_{HML}$		0.007	−0.073
		(0.03)	(−0.29)
$\widetilde{\beta}_{WML}$			1.572
			(1.93)
R^2	0.10	0.37	0.48

Panel D: Characteristic-based tests

	Model 7	Model 8
β_{MKT}	2.076	1.764
	(5.07)	(4.67)
log(*ME*)	−2.107	−1.954
	(−13.97)	(−14.03)
BE/*ME*	1.286	1.279
	(4.02)	(4.28)
Past returns		−4.277
		(−5.16)
R^2	0.20	0.25

	Model 9	Model 10
$\widetilde{\beta}_{MKT}$	1.102	1.154
	(3.15)	(3.24)
log(*ME*)	−1.729	−1.575
	(−8.01)	(−8.98)
BE/*ME*	0.707	0.715
	(3.74)	(3.99)
Past returns		−3.637
		(−4.11)
R^2	0.28	0.31

We estimate the elasticity of equity with respect to bond by including log size, market-to-book, and past stock return, in addition to leverage, volatility, and risk-free rate in Equation (8). We then calculate the firm-specific expected equity excess returns as the bond excess returns multiplied by the new equity-bond elasticity. This table replicates the tests in Table 10 using these newly constructed expected equity returns. Panel A reports the summary statistics of the expected returns for the *MKT*, *SMB*, *HML*, and *WML* factors. Panel B reports the cross correlations of the expected factor returns with the cyclical component of industrial production. All correlations are significantly different from zero at the 10% significance level. Panel C reports the monthly Fama-MacBeth (1973) cross-sectional regressions of firm-level equity excess returns on the *MKT* beta, the *SMB* beta, the *HML* beta, and the *WML* beta, separately and jointly. Panel D reports the characteristic-based tests. The model specifications are the same as in Table 10 (see the caption of that table for detailed definitions). The t-statistics reported in parentheses are adjusted for heteroscedasticity and autocorrelations via GMM. All the estimates are in annualized percentage terms. The R^2s are the time-series median of the cross-sectional regression R^2s.

We subsequently construct expected equity excess returns using the elasticity of equity with respect to bond estimated from Equation (9). Table 12 shows that the inclusion of additional variables does not affect our basic inferences. The market beta continues to be significantly priced in the cross-sectional regressions. The three Fama-French (1993) factors continue to be countercyclical, and the expected momentum factor remains procyclical. For the most part, size and book-to-market are priced, but momentum is either not priced or priced negatively.

5.3 Liquidity

Part of the yield spreads can be due to the fact that corporate bonds are less liquid than Treasury bonds (e.g., Chen et al, 2007). The bond risk premiums can thus be overstated if we do not consider this liquidity effect. Cooper and Davydenko (2004) and Chen et al. (2006) argue that a relatively easy way to account for the liquidity effect is to use the Aaa yield (instead of the Treasury yield) to calculate the yield spreads. Because the Aaa yield over the Treasury bill rate consists of a liquidity component and a tax component (corporate bonds are taxable at state and local levels), we can recompute the bond risk premium as

$$R_{Bt}^{i} - r_t = \left(Y_{it} - \text{Aaa yield}_t\right) + \text{EDL}_{it} + \text{ERND}_{it}. \tag{10}$$

Compared to Equation (6), we have replaced the Treasury yield by the Aaa yield. In addition, we no longer need to correct for the tax component because it is contained in the Aaa yield.

We gather the time series data on long-term Aaa yields from the Federal Reserve Board. Using these data, we first calculate the long-term Aaa bond yields over long-term Treasury yields (maturity = 20 years). We assume that this spread shrinks to zero when maturity is equal to zero and interpolate the spread for maturities between 0 and 20 years. Assuming that the spread converges to a positive number does not affect our basic results (not reported). With this term structure of the Aaa yields we can then calculate the bond risk premium.

Table 13 reports our empirical results using the Aaa yields to calculate yield spreads. As shown in Panel A, using Aaa yields causes our estimate of the market risk premium to drop slightly from 3.39% in the benchmark estimation to 3.10%, suggesting a liquidity premium of 0.29% per annum for the market factor. The expected size premium drops from 3.81% to 3.23%. All other patterns are largely similar to the benchmark results. In particular, the market beta remains significantly priced in the cross section. Size and book-to-market are priced, but momentum is either not priced or priced negatively. In all, bond liquidity considerations do not materially affect our basic results.[10]

5.4 Cross section of expected bond excess returns

Proposition 1 shows that expected equity excess returns have two basic components: expected bond excess returns and the elasticity of the equity value with respect to the bond value. It is natural to ask if the cross-sectional variations in expected equity excess returns explained in the benchmark estimation are due to the variations in expected bond excess returns or the variations in the equity-bond elasticity. To this end, we perform Fama-MacBeth (1973) cross-sectional

[10] Campbell and Taksler (2003) show that firm-specific volatility can play an important role in determining bond yield spreads. It is likely that when we deduct the expected default loss from the yield spreads, we expunge this idiosyncratic component. To check robustness, we also perform tests in which we include equity volatility in the cross-section regresssions. We find that this inclusion does not affect our basic results (not reported).

Table 13
Using Aaa yield to calculate yield spreads

Panel A: Summary statistics

	Mean	(t-stat)	Min	Max
MKT	3.100	(12.05)	0.00	15.68
SMB	3.234	(16.25)	−0.69	13.19
HML	1.408	(6.52)	−3.81	14.26
WML	−0.387	(2.18)	−13.86	4.57

Panel B: Correlations with cycle

	−1	0	1
MKT	−0.16	−0.17	−0.16
SMB	−0.33	−0.36	−0.39
HML	−0.13	−0.16	−0.18
WML	0.17	0.23	0.27

Panel C: Covariance-based tests

	Model 1	Model 2	Model 3
β_{MKT}	3.562	2.402	2.119
	(7.75)	(3.44)	(3.25)
β_{SMB}		3.023	3.063
		(11.80)	(9.34)
β_{HML}		1.247	1.146
		(3.04)	(2.79)
β_{WML}			−0.387
			(−0.66)
R^2	0.03	0.13	0.15
	Model 4	Model 5	Model 6
$\widetilde{\beta}_{MKT}$	0.966	1.461	1.248
	(3.30)	(3.20)	(2.92)
$\widetilde{\beta}_{SMB}$		1.018	1.361
		(2.36)	(3.57)
$\widetilde{\beta}_{HML}$		−0.127	−0.291
		(−0.53)	(−1.18)
$\widetilde{\beta}_{WML}$			0.697
			(1.14)
R^2	0.13	0.46	0.54

Panel D: Characteristic-based tests

	Model 7	Model 8
β_{MKT}	2.068	1.788
	(4.92)	(4.62)
log(*ME*)	−1.924	−1.781
	(−13.53)	(−13.61)
BE/*ME*	1.272	1.257
	(4.02)	(4.26)
Past returns		−4.450
		(−5.64)
R^2	0.19	0.22
	Model 9	Model 10
$\widetilde{\beta}_{MKT}$	0.906	0.934
	(3.27)	(3.33)
log(*ME*)	−1.701	−1.565
	(−7.39)	(−8.10)
BE/*ME*	0.770	0.773
	(4.37)	(4.61)
Past returns		−3.313
		(−3.91)
R^2	0.30	0.33

We calculate the yield spread as the corporate bond yield minus Aaa yield with similar maturity to control for liquidity effect. We then estimate equity excess returns using the resulting yield spreads. This table replicates earlier tests in Table 10 using these newly constructed expected equity excess returns. Panel A reports the summary statistics of the equity factors of *MKT*, *SMB*, *HML*, and *WML*. Panel B reports the cross-correlations of the expected returns of these equity factors with the cyclical component of industrial production. All correlations are significant at the 1% significance level. Panel C reports the monthly Fama-MacBeth (1973) cross-sectional regressions of firm-level expected equity excess returns on the *MKT* beta, the *SMB* beta, the *HML* beta, and the *WML* beta, separately and jointly. Panel D reports the characteristic-based tests. The model specifications are the same as in Table 10 (see the footnote of that table for details). The t-statistics reported in parentheses are adjusted for heteroscedasticity and autocorrelations via GMM. All estimates are in annualized percentage terms. The R^2s are the time-series median of the cross-sectional regression R^2s.

regressions of expected bond excess returns. We find that the cross section of expected bond excess returns already captures almost all of the basic properties of the cross section of expected equity excess returns.

Table 14 reports the details. Most importantly the market beta earns a bond market risk premium of 0.41% per annum (t-statistic = 6.95). The magnitude of this risk premium estimate is similar to the time series estimate reported in Section 2. The significant risk premium subsists after we control for size, book-to-market, and prior returns in the form of either covariances or characteristics. Small firms earn higher expected bond returns than big firms, and value firms earn higher expected bond returns than growth firms. Momentum is once again either not priced or negatively priced.

Table 14
The cross section of expected bond excess returns

Panel A: Covariance-based tests

	Model 1	Model 2	Model 3
β_{MKT}	0.413	0.290	0.251
	(6.95)	(2.75)	(2.39)
β_{SMB}		0.371	0.385
		(11.41)	(10.80)
β_{HML}		0.191	0.199
		(3.32)	(3.49)
β_{WML}			−0.022
			(−0.30)
R^2	0.02	0.11	0.14
	Model 4	**Model 5**	**Model 6**
$\widetilde{\beta}_{MKT}$	0.106	0.280	0.223
	(2.55)	(4.72)	(4.01)
$\widetilde{\beta}_{SMB}$		0.071	0.045
		(1.77)	(1.03)
$\widetilde{\beta}_{HML}$		0.026	0.021
		(0.78)	(0.69)
$\widetilde{\beta}_{WML}$			0.290
			(1.74)
R^2	0.11	0.42	0.52

Panel B: Characteristic-based tests

	Model 7	Model 8
β_{MKT}	0.224	0.184
	(3.73)	(3.50)
log(*ME*)	−0.259	−0.239
	(−13.43)	(−13.77)
BE/*ME*	0.236	0.229
	(3.36)	(3.45)
Past returns		−0.752
		(−7.38)
R^2	0.22	0.27
	Model 9	**Model 10**
$\widetilde{\beta}_{MKT}$	0.092	0.102
	(2.30)	(2.55)
log(*ME*)	−0.236	−0.211
	(−6.52)	(−7.09)
BE/*ME*	0.115	0.114
	(4.50)	(4.82)
Past returns		−0.758
		(−5.17)
R^2	0.30	0.34

We estimate monthly Fama-MacBeth (1973) cross-sectional regressions for the following ten specifications:

$$\begin{aligned}
\text{Model 1}: &\quad E_t[R^i_{Bt} - r_t] = \gamma_{0t} + \gamma_{1t}\beta^i_{MKT,t} + \epsilon^i_t \\
\text{Model 2}: &\quad E_t[R^i_{Bt} - r_t] = \gamma_{0t} + \gamma_{1t}\beta^i_{MKT,t} + \gamma_{2t}\beta^i_{SMB,t} + \gamma_{3t}\beta^i_{HML,t} + \epsilon^i_t \\
\text{Model 3}: &\quad E_t[R^i_{Bt} - r_t] = \gamma_{0t} + \gamma_{1t}\beta^i_{MKT,t} + \gamma_{2t}\beta^i_{SMB,t} + \gamma_{3t}\beta^i_{HML,t} + \gamma_{4t}\beta^i_{WML,t} + \epsilon^i_t \\
\text{Model 4}: &\quad E_t[R^i_{Bt} - r_t] = \gamma_{0t} + \gamma_{1t}\widetilde{\beta}^i_{MKT,t} + \epsilon^i_t \\
\text{Model 5}: &\quad E_t[R^i_{Bt} - r_t] = \gamma_{0t} + \gamma_{1t}\widetilde{\beta}^i_{MKT,t} + \gamma_{2t}\widetilde{\beta}^i_{SMB,t} + \gamma_{3t}\widetilde{\beta}^i_{HML,t} + \epsilon^i_t \\
\text{Model 6}: &\quad E_t[R^i_{Bt} - r_t] = \gamma_{0t} + \gamma_{1t}\widetilde{\beta}^i_{MKT,t} + \gamma_{2t}\widetilde{\beta}^i_{SMB,t} + \gamma_{3t}\widetilde{\beta}^i_{HML,t} + \gamma_{4t}\widetilde{\beta}^i_{WML,t} + \epsilon^i_t \\
\text{Model 7}: &\quad E_t[R^i_{Bt} - r_t] = \gamma_{0t} + \gamma_{1t}\beta^i_{MKT,t} + \gamma_{2t}\log(ME^i_t) + \gamma_{3t}BE^i_t/ME^i_t + \epsilon^i_t \\
\text{Model 8}: &\quad E_t[R^i_{Bt} - r_t] = \gamma_{0t} + \gamma_{1t}\beta^i_{MKT,t} + \gamma_{2t}\log(ME^i_t) + \gamma_{3t}BE^i_t/ME^i_t + \gamma_{4t}(\text{PastReturns})^i_t + \epsilon^i_t \\
\text{Model 9}: &\quad E_t[R^i_{Bt} - r_t] = \gamma_{0t} + \gamma_{1t}\widetilde{\beta}^i_{MKT,t} + \gamma_{2t}\log(ME^i_t) + \gamma_{3t}BE^i_t/ME^i_t + \epsilon^i_t \\
\text{Model 10}: &\quad E_t[R^i_{Bt} - r_t] = \gamma_{0t} + \gamma_{1t}\widetilde{\beta}^i_{MKT,t} + \gamma_{2t}\log(ME^i_t) + \gamma_{3t}BE^i_t/ME^i_t + \gamma_{4t}(\text{PastReturns})^i_t + \epsilon^i_t,
\end{aligned}$$

where $E_t[R^i_{Bt} - r_t]$ is the firm-level expected bond excess return for firm i at the beginning of time t, and r_t is the risk-free rate. The sample is from January 1973 to March 1998. We estimate betas in two ways. First, we estimate these betas using 60-month (or at least 24-month) rolling regressions of the realized equity excess returns on the realized excess returns of the Fama-French (1993) three factors and *WML*. These betas estimated from realized returns are denoted β_{MKT}, β_{SMB}, β_{HML}, and β_{WML}. Second, we also estimate the betas from 60-month (or at least 24-month) rolling regressions of the constructed expected excess returns on the constructed expected Fama-French three factors and *WML*. These betas are denoted $\widetilde{\beta}_{MKT}$, $\widetilde{\beta}_{SMB}$, $\widetilde{\beta}_{HML}$, and $\widetilde{\beta}_{WML}$. We also use size (log(*ME*)), book-to-market equity (*BE*/*ME*), and past 12-month returns as explanatory variables. The t-statistics reported in parentheses are adjusted for heteroscedasticity and autocorrelations via GMM. The point estimates of the intercepts and slopes are in annualized percentage terms. The R^2s, are the time-series median of the cross-sectional regression R^2s.

The equity-bond elasticity serves to strengthen the pricing results for the market beta and size, but it plays a relatively minor role in driving the cross section of expected equity returns. In untabulated results, we perform Fama-MacBeth (1973) cross-sectional regressions of the equity-bond elasticity. Both the market beta and size play a significant role in driving the cross-sectional variations of the equity-bond elasticity. Book-to-market and prior returns, however, show mixed pricing results.

5.5 Using testing portfolios in the cross-sectional tests

Table 10 shows that using our expected-return measure as the dependent variable in cross-sectional regressions dramatically increases the R^2s relative to using realized returns. This evidence is documented using individual stock returns as testing assets. However, testing portfolios are more commonly used in asset pricing tests because the diversification in portfolios helps reduce measurement errors in factor loadings (and expected returns). It is therefore natural to ask if our expected-return measure can increase regression R^2 when we use testing portfolios. The answer is affirmative.

We consider 30 testing portfolios including 10 size, 10 book-to-market, and 10 momentum portfolios, all of which are based on one-way sorts. A similar set of testing portfolios is used in the asset pricing tests of Bansal, Dittmar, and Lundblad (2005). We opt to use this set of portfolios because some of the double-sorted portfolios such as 25 size and book-to-market portfolios often contain too few firms. (Our sample coverage in the cross section is not as broad as the CRSP/COMPUSTAT universe.)

Table 15 reports Fama-MacBeth (1973) cross-sectional regressions of the 30 testing portfolios. Panels A and B use our expected-return measures as the dependent variables, whereas Panels C and D use the realized future returns. Using expected returns increases cross-sectional R^2s relative to using realized returns, although the increases are not as dramatic as using individual stock returns as testing assets. For example, regressing expected excess returns on the Carhart (1997) four factor loadings generates a R^2 of 60%. In contrast, regressing realized excess returns on the same set of factor loadings only generates an R^2 of 38%.

Further, our basic inferences drawn from the cross section of individual stock returns are not materially affected by the use of testing portfolios. In particular, the market beta is priced using expected returns, and the risk premium estimates are mostly significant. In contrast, using realized returns in the tests generates mostly insignificant risk premium estimates.

6. Summary and Interpretation

In this paper, we construct alternative measures of expected returns using bond yield spreads. Built on the structural-form framework of Merton (1974), the rationale of our approach is straightforward. Because both equity and bond are contingent claims written on the same productive assets, equity and bond must share similar systematic risk factors. Moreover, corporate bonds are special because bond yield spreads contain forward-looking risk premiums beyond expected default losses. Although we need some past information to estimate

Table 15
Using testing portfolios to perform Fama-MacBeth (1973) cross-sectional regressions of expected equity excess returns and realized equity excess returns

Panel A: Covariance-based tests (expected excess returns)

	Model 1	Model 2	Model 3
β_{MKT}	10.576	4.077	2.653
	(8.70)	(3.44)	(2.38)
β_{SMB}		7.824	8.383
		(11.27)	(10.33)
β_{HML}		3.835	4.037
		(7.08)	(7.24)
β_{WML}			3.898
			(2.13)
R^2	0.13	0.55	0.60
	Model 4	Model 5	Model 6
$\widetilde{\beta}_{MKT}$	1.515	0.387	0.297
	(11.01)	(6.04)	(5.84)
$\widetilde{\beta}_{SMB}$		3.666	3.152
		(11.78)	(6.92)
$\widetilde{\beta}_{HML}$		−0.032	0.592
		(−0.09)	(1.81)
$\widetilde{\beta}_{WML}$			1.554
			(3.85)
R^2	0.32	0.69	0.77

Panel B: Characteristic-based tests (expected excess returns)

	Model 7	Model 8
β_{MKT}	2.922	2.493
	(5.52)	(4.08)
log(*ME*)	−2.574	−2.354
	(−11.60)	(−13.54)
BE/*ME*	0.570	0.475
	(4.89)	(4.01)
Past returns		−8.065
		(−16.13)
R^2	0.64	0.72
	Model 9	Model 10
$\widetilde{\beta}_{MKT}$	−0.066	−0.047
	(−0.65)	(−0.56)
log(*ME*)	−2.763	−2.406
	(−19.21)	(−23.12)
BE/*ME*	0.445	0.360
	(3.66)	(3.69)
Past returns		−6.821
		(−15.53)
R^2	0.64	0.72

Panel C: Covariance-based tests (realized excess returns)

	Model 11	Model 12	Model 13
β_{MKT}	0.330	0.295	0.254
	(1.86)	(1.47)	(1.11)
β_{SMB}		0.155	0.184
		(1.05)	(1.23)
β_{HML}		0.120	0.149
		(1.06)	(1.11)
β_{WML}			0.194
			(1.17)
R^2	0.05	0.27	0.38

Panel D: Characteristic-based tests (realized excess returns)

	Model 14	Model 15
β_{MKT}	0.313	0.142
	(2.21)	(1.16)
log(*ME*)	−0.125	−0.147
	(−1.87)	(−2.47)
BE/*ME*	0.080	0.068
	(1.42)	(1.61)
Past returns		1.372
		(5.40)
R^2	0.33	0.48

This table replicates tests in Table 10 using these 30 testing portfolios including 10 size, 10 book-to-market, and 10 momentum portfolios. In contrast, Table 10 uses individual stock returns as testing assets. Panel A reports the monthly Fama-MacBeth (1973) cross-sectional regressions of expected equity excess returns of testing portfolios on the *MKT* beta, the *SMB* beta, the *HML* beta, and the *WML* beta, separately and jointly. We estimate betas in two ways. First, we estimate these betas using 60-month (or at least 24-month) rolling regressions of the realized equity excess returns on the realized excess returns of the Fama-French (1993) three factors and *WML*. These betas estimated from realized returns are denoted β_{MKT}, β_{SMB}, β_{HML}, and β_{WML}. Second, we also estimate the betas from 60-month (or at least 24-month) rolling regressions of the constructed expected excess returns on the constructed expected Fama-French three factors and *WML*. These betas are denoted $\widetilde{\beta}_{MKT}$, $\widetilde{\beta}_{SMB}$, $\widetilde{\beta}_{HML}$, and $\widetilde{\beta}_{WML}$. Panel B reports the characteristic-based tests of expected equity excess returns. We use size (log(*ME*)), book-to-market equity (*BE*/*ME*), and past 12-month returns as explanatory variables. The sample is from January 1973 to March 1998. The regression specifications are the same as in Table 10. In Panels C and D, we repeat the tests in regression models 1, 2, 3, 7, and 8 in Panels A and B using realized equity excess returns to replace expected equity excess returns as the dependent variables. The t-statistics reported in parentheses are adjusted for heteroscedasticity and autocorrelations via GMM. All estimates are in annualized percentage terms. The R^2 are the time-series median of the cross-sectional regression R^2s.

expected returns, our approach largely retains the forward-looking nature of the risk premium embedded in yield spreads. Using our expected-return measures to study the cross section of returns, our analysis yields several fresh insights.

First and foremost, we document that the market beta plays a much more important role in the cross section of expected returns than previously reported. The market beta is significantly priced in our cross-sectional regressions even after we control for size, book-to-market, and prior returns. Our results contrast with those of Fama and French (1992), who show that the market beta does not have significant explanatory power in the cross section of average returns. Our evidence suggests that previous reports that beta is dead can be driven by the use of average returns as a poor proxy for expected returns. Our evidence lends support to Kothari et al. (1995), who find economically and statistically significant compensations for beta risk (about 6–9% per annum). The evidence is also consistent with Brav et al. (2005), who use analysts' forecasts to construct expected returns and find a positive and robust relation between expected returns and the market beta.

Second, we find that the expected *SMB* and *HML* returns are significantly positive and countercyclical. This evidence lends support to Fama and French (1993, 1996), who argue that size and book-to-market factors are priced risk factors. In this regard, our result differs from Brav et al. (2005), who find that high book-to-market firms are not expected to earn higher returns than low book-to-market firms. Our evidence is important because our sample contains disproportionately more large firms than the CRSP/COMPUSTAT universe. The cyclicality of the expected value premium also lends support to studies that suggest an important role of conditional information in driving the value premium. For example, Ferson and Harvey (1999) show that loadings on aggregate predictive variables provide significant cross-sectional explanatory power for stock returns even after controlling for the Fama-French factors. Lettau and Ludvigson (2001) show that value stocks correlate more strongly with consumption growth than growth stocks in bad times, when the price of risk is high.[11]

Finally, we find that momentum is either not priced or negatively priced using expected returns. This evidence is consistent with several interpretations. First, investors do not perceive expected momentum portfolios to be risky. Positive momentum profits in *ex post* returns come as a surprise to investors, consistent with mispricing stories of Barberis et al. (1998), Daniel et al. (1998), and Hong and Stein (1999). Second, the distribution of expected returns can deviate from the distribution of realized returns (e.g., Brav and Heaton, 2002; Lewellen and Shanken, 2002). Shanken (2004) argues that this deviation can affect the interpretation of asset pricing tests. Even though *ex post* returns can appear predictable to econometricians, investors in reality can neither perceive

[11] See also Jagannathan and Wang (1996); Ang and Chen (2004); Jostova and Philipov (2004); and Anderson et al. (2005). However, Lewellen and Nagel (2006) are more skeptical about the role of conditioning information.

nor exploit this predictability. Third, momentum profits can be an empirical by-product of using averaged realized returns as a poor proxy for expected returns.

It is important to acknowledge several caveats in our empirical approach. The simple contingent-claim framework à la Merton (1974) allows us to derive a conditionally linear relation between expected equity and bond excess returns. Under more general conditions, the relation might not be conditionally linear. We therefore emphasize that our empirical approach is only motivated by the Merton-style framework, as opposed to being a literal structural test of that framework. Further, we have to use existing default information to gauge expected default losses, meaning that our constructed bond excess returns are not entirely "*ex ante*." However, recent research has shown that expected default losses can only explain a relatively small portion of the yield spreads. Most of the yield spreads are likely driven by systematic risk (e.g., Huang and Huang, 2003). By restricting the use of historical data to a small portion of the yield spreads, we retain crucial information on the forward-looking risk premiums embedded in bond yields.

While we do not suggest that our expected-return measures dominate any other, using measures other than average realized returns to reexamine the standard, deep-rooted inferences in empirical asset pricing seems to be a valid experiment. Because our measure captures information imperfectly correlated with average realized returns, tests like ours are likely to provide fresh insights into the economic drivers of the cross section of expected returns.

A Appendix: Proofs

A.1 Proof of Proposition 1

The uncertainty in the economy is represented by an N-dimensional vector $X_t = (x_{1t}, x_{2t}, \dots, x_{Nt})'$ with a constant variance-covariance matrix Ω. The stochastic discount factor m_t is a function of X_t. The asset value of any firm i, F_i, is also a function of X_t. We further assume that asset volatility can be a function of asset value but is not a direct function of X_t. Barring arbitrage,

$$F_{it} = E_t\left[m_t F_{it+1}\right]$$

and we obtain

$$E_t\left[\frac{dF_{it}}{F_{it}}\right] - r_t \equiv E_t\left[R^i_{Ft}\right] - r_t = -\mathrm{cov}\left(m_t, R^i_{Ft}\right),$$

where $\frac{dF_{it+1}}{F_{it}} \equiv R^i_{Ft} \equiv \frac{F_{it+1} - F_{it}}{F_{it}}$ is asset return, $r_t = \frac{1}{E[m_t]}$ is the risk-free rate which we assume is deterministic, and the left-hand side is the expected asset premium.

Following Merton (1974), we assume that all firms are levered with predetermined debt. Firm i defaults if its value $F_{i\tau}$ in the future hits some lower boundary as a fraction of its initial value. With this setup equity and bond are contingent claims on the asset value. It follows that equity return is

$$\begin{aligned} R^i_{St} = \frac{dS_{it}}{S_{it}} &= \left(\frac{\partial S_{it}}{\partial F_{it}}\right)\left(\frac{F_{it}}{S_{it}}\right)\left(\frac{dF_{it}}{F_{it}}\right) \\ &= \left(\frac{\partial S_{it}}{\partial F_{it}}\right)\left(\frac{F_{it}}{S_{it}}\right) R^i_{Ft}, \end{aligned} \tag{A1}$$

and

$$\begin{aligned} R^i_{Bt} = \frac{dB_{it}}{B_{it}} &= \left(\frac{\partial B_{it}}{\partial F_{it}}\right)\left(\frac{F_{it}}{B_{it}}\right)\left(\frac{dF_{it}}{F_{it}}\right) \\ &= \left(\frac{\partial B_{it}}{\partial F_{it}}\right)\left(\frac{F_{it}}{B_{it}}\right) R^i_{Ft}. \end{aligned} \tag{A2}$$

Expected equity return and bond return must satisfy

$$E\left[R^i_{St}\right] - r = -\mathrm{cov}\left(m_t, R^i_{St}\right) \tag{A3}$$

$$E\left[R^i_{Bt}\right] - r = -\mathrm{cov}\left(m_t, R^i_{Bt}\right). \tag{A4}$$

Proposition 1 follows by substituting (A1) into (A3), substituting (A2) into (A4), and taking the ratio of the two equations.

Because bankruptcy happens as a function of asset value, the multiple-factor vector X_t affects expected risk premium only indirectly through its impact on asset value F_{it}. Therefore, even though expected equity premium and bond premium are driven by asset value, and thus by multiple risk factors, they are only direct functions of asset value. As a result, the relation between expected equity premium and bond premium is conditionally linear. Also note that for equity and bond as functions of asset value alone, we have assumed that the interest rate is deterministic and that asset volatility is a function of asset value alone.[12] ■

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[12] We have presented Proposition 1 in discrete time for simplicity. The intuition of the proposition can also be seen alternatively through a continuous-time example. Assume the risk vector X_t follows geometric Brown motion:

$$\frac{dx_{it}}{x_{it}} = \mu_i dt + \sqrt{\Omega_{ii}} dw_{it}$$

for $i = 1, \ldots, N$ and Ω_{ii} is the ith diagonal element of the constant matrix Ω. Because asset value F_{it} is a function of multiple risk factors X_t, according to Ito's lemma,

$$\frac{dF_{it}}{F_{it}} = \sum_{j=1}^{N} \frac{\partial F_{it}}{\partial x_{jt}} \frac{x_{jt}}{F_{it}} \frac{dx_{jt}}{x_{jt}} + \sum_{g=1}^{N}\sum_{h=1}^{N} \frac{\partial^2 F_{it}}{\partial x_{gt}\partial x_{ht}} \frac{1}{F_{it}} \Omega_{gh} dt.$$

Here we assume that the asset volatility is either a constant or a function of asset value alone. The expected asset premium satisfies

$$E_t\left[\frac{dF_{it}}{F_{it}}\right] - r = -\sum_{j=1}^{N} \frac{\partial F_{it}}{\partial x_{jt}} \frac{x_{jt}}{F_{it}} \mathrm{cov}\left(\frac{d\Lambda_t}{\Lambda_t}, \frac{dx_{jt}}{x_{jt}}\right)$$

and $\frac{d\Lambda_t}{\Lambda_t}$ is the stochastic discount factor. That is, the expected asset premium is driven by multiple risk factors. Because bankruptcy happens as a function of asset value, with risk-neutral valuation, the values of equity and

A.2 Proof of Proposition 2

Similar to Jarrow (1978), we start with the bond yield equation:

$$B_{it} = \sum_{j=1}^{n} C_i e^{-Y_{it}T_j} + K_i e^{-Y_{it}T_n},$$

where C_i is the coupon payment of the bond, n is the number of remaining coupons, Y_{it} is the bond's yield to maturity, T_j, $j = 1, \ldots, n$ are length of time period for each coupon payment, and K_i is the face value of debt.

By definition, the bond risk premium is a function of bond yield and time. Taking a second-order Taylor expansion, and taking the expectation on both sides, we can write bond risk premium as a function of the bond yield and other observable bond characteristics:

$$\frac{E_t\left[dB_{it}\right]}{B_{it}} = E_t\left[\frac{\partial B_{it}}{\partial t}\frac{dt}{B_{it}} + \frac{\partial B_{it}}{\partial Y_{it}}\frac{dY_{it}}{B_{it}} + \frac{1}{2}\frac{\partial^2 B_{it}}{\partial Y_{it}^2}\frac{1}{B_{it}}\left(dY_{it}\right)^2\right], \tag{A5}$$

where

$$\frac{\partial B_{it}}{\partial t} = Y_{it}B_{it} \tag{A6}$$

$$\frac{\partial B_{it}}{\partial Y_{it}} = -H_{it}B_{it} \quad \text{with} \quad H_{it} = \sum_{j=1}^{n}\frac{T_j C_i e^{-Y_{it}T_j}}{B_{it}} + \frac{T_n K_i e^{-Y_{it}T_n}}{B_{it}} \tag{A7}$$

$$\frac{\partial^2 B_{it}}{\partial Y_{it}^2} = G_{it}D_{it} \quad \text{with} \quad G_{it} \equiv \sum_{j=1}^{n}\frac{T_j^2 C_i e^{-Y_{it}T_j}}{B_{it}} + \frac{T_n^2 K_i e^{-Y_{it}T_n}}{B_{it}} \tag{A8}$$

and H_{it} and G_{it} are modified duration and convexity, respectively. In the above derivation we have assumed that $(dt)^2$ is too small and is omitted. (A5) thus

bond, as contingent claims, are completely determined by the asset value, the risk-free rate, and the volatility of asset value. Therefore, with the assumption of constant risk-free rate and that asset volatility is only a function of asset value, equity and bond are only driven stochastically by asset value. Applying Ito's lemma,

$$\frac{dS_{it}}{S_{it}} = \frac{\partial S_{it}}{\partial F_{it}}\frac{F_{it}}{S_{it}}\frac{dF_{it}}{F_{it}} + \frac{\partial^2 S_{it}}{\partial F_{it}^2}\frac{1}{S_{it}}\left(dF_{it}\right)^2$$

where the second item is locally deterministic and does not affect expected equity premium. The expected equity premium is then

$$E_t\left[\frac{dS_{it}}{S_{it}}\right] - r = -\frac{\partial S_{it}}{\partial F_{it}}\frac{F_{it}}{S_{it}}\operatorname{cov}\left(\frac{d\Lambda_t}{\Lambda_t}, \frac{dF_{it}}{F_{it}}\right),$$

and similarly, the expected bond risk premium is

$$E_t\left[\frac{dB_{it}}{B_{it}}\right] - r = -\frac{\partial B_{it}}{\partial F_{jt}}\frac{F_{jt}}{B_{it}}\operatorname{cov}\left(\frac{d\Lambda_t}{\Lambda_t}, \frac{dF_{it}}{F_{it}}\right).$$

Again the expected equity premium is linearly related to the expected bond premium.

becomes

$$E_t\left[R^i_{Bt}\right] - r_t = E_t\left[\frac{dB_{it}}{B_{it}}\right]/dt - r_t = (Y_{it} - r_t) - H_{it}\frac{E_t\left[dY_{it}\right]}{dt} + \frac{1}{2}G_{it}\frac{E_t[(dY_{it})^2]}{dt}. \tag{A9}$$

■

A.3 Proof of Proposition 3

The proposition follows by combining (2) with

$$E_t[dY_{it}] = \pi_{it}E_t[dY^-_{it} \mid \text{default}] + (1 - \pi_{it})E_t[dY^+_{it} \mid \text{no default}]$$
$$E_t[(dY_{it})^2] = \pi_{it}E_t[(dY^-_{it})^2 \mid \text{default}] + (1 - \pi_{it})E_t[(dY^+_{it})^2 \mid \text{no default}].$$

■

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