Expected Return and the Bid-Ask Spread

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ABSTRACT

This paper empirically examines the relation between the expected stock return and the bid-ask spread. Using the same portfolio formation method as in Amihud and Mendelson (1986) but different test methodologies, we do not find any clear reliable relation between the CAPM risk-adjusted return and the relative bid-ask spread. Our empirical results are more consistent with the conclusions of Constantinides (1986) that transaction costs play a minor role in the determination of expected returns in security markets.

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Most asset pricing models, such as Sharpe (1964), Lintner (1965), Merton (1973), Ross (1976), Rubinstein (1976), Breeden (1979) and Cox, Ingersoll and Ross (1985), are derived with the assumption of a perfect capital market. In these cases, the expected return of a financial asset is a function of risks alone. In actual trading, however, the transaction costs can be very different for different classes of financial assets and the equilibrium expected gross return can be a function of the transaction costs also.

Since investors can only consume the investment income net of transaction costs, the expected gross return should be positively correlated with transaction costs. Unfortunately, there is little agreement among the researchers of how important this positive relation between the expected return and the transaction costs is. On one hand, differences in transaction costs can induce a clientele effect. For example, high transaction cost securities tend to be held for long-term investment purposes and such cost-minimization strategies on the part of investors would attenuate any positive relation between return and transaction costs. Furthermore, Constantinides (1986) argues theoretically that transaction costs can only have a second-order effect on the liquidity premium implied by the equilibrium asset returns in an intertemporal portfolio selection model. On the other hand, Amihud and Mendelson (1986) (hereafter AM) empirically find that there is an economically and statistically significant positive relation between the expected return and the relative bid-ask spread.

The relative bid-ask spread is measured as the dealer's bid-ask spread divided by the average of the bid-price and the ask-price. Since many recent studies have found a reliable relation between expected return and (functions of) price, it is obvious that we cannot unambiguously interpret the results in AM as a relation between the expected return and the relative bid-ask spread. Miller and Scholes (1982) observe a positive relation between return and the inverse of price and point out that the inverse of price may be proxying for the errors in the estimation of risk. Since firm size and relative bid-ask spread are both functions of price and both are possibly related to risk mismeasured or unmeasured, the explanation for the size effect proposed in the Chan and Chen (1988) study may also apply to the relative bid-ask spread variable. Furthermore, given the recent empirical evidence in Fama and French (1988) and Chen, Grundy and Stambaugh (1990), the slope coefficient of the relative bid-ask spread in a joint time-series cross-sectional test may also be picking up time-series variations in the market risk and risk premium.

In this study, we shall investigate the relation between the expected return and the relative bid-ask spread taking into account the recent empirical regularities mentioned above. We examine the robustness of this relation estimated with (i) the Fama-MacBeth (1973) approach, (ii) the SUR framework as in Zellner (1962), Gibbons (1982), and Stambaugh (1982), (iii) the unconditional approach in Chan and Chen (1988), and (iv) an empirical design proposed in Chen, Grundy and Stambaugh (1990) where the changing risk measure is modeled directly. In most cases, we find that the relation between the CAPM risk-adjusted return and the relative bid-ask spread is economically and statistically indistinguishable from zero.

I. The Data

The relative bid-ask spread variable is the same as that constructed in Stoll and Whaley (1983) and used in the AM study.[1] The primary data source is Fitch's Stock Quotations on the NYSE. The relative spread is the year-end bid-ask spread divided by the average of the bid-price and the ask-price. For each year in the sample period 1961–1980, the spread for stock i, S_{it}, is the average of the beginning and end-of-year relative spread in year $t-1$.

To test the relation between the expected return and the relative bid-ask spread conditional on the market beta, we form 49 equally-weighted portfolios based on the same criteria used in AM. For each "test" year from 1961 to 1980, the stocks traded through the eleven-year period from $t-10$ to t on the NYSE are first ranked by their relative spreads in $t-1$ and divided into seven spread groups. Within each spread group, the stocks are ranked by their equally-weighted NYSE market beta estimated with monthly excess returns from year $t-10$ to $t-6$, and divided into seven subgroups. Thus, we form 49 equally-weighted portfolios and the portfolio composition is updated once a year. The monthly portfolio excess returns are equally-weighted monthly rebalanced returns in excess of the 30-day Treasury-bill rates.

Using their methodology, we find that our results are largely consistent with the findings in AM. For example, AM (1986, p. 238) report the following result from a joint time-series cross-sectional empirical design:

$$r_{pt} = \underset{}{0.0036} + \underset{(6.18)}{0.00672}\beta_{pt} + \underset{(6.83)}{0.211}\, S_{pt} + \sum_{t=1}^{19} d_t DY_t + e_{pt},$$

and our estimation of the same equation, using the same methodology, is

$$r_{pt} = 0.0039 + \underset{(6.45)}{0.00712}\beta_{pt} + \underset{(6.79)}{0.212}\, S_{pt} + \sum_{t=1}^{19} d_t DY_t + e_{pt},$$

where r_{pt} is the average monthly excess return for portfolio p in year t, β_{pt} is the equally-weighted NYSE market beta estimated with monthly excess returns from year $t-5$ to $t-1$, S_{pt} is the average relative spread of the stocks in portfolio p as of year $t-1$, and DY_t is the yearly dummy variable with $DY_t = 1$ in year t $(t = 1, 2, \ldots, 19)$.[2]

As AM recognize in their discussion, the t-statistics in these equations are overstated because of the correlations among residuals across portfolio groups. Using the OLS standard errors that are corrected for the residual correlations,[3] the above equation becomes

$$r_{pt} = 0.0039 + \underset{(2.23)}{0.00712}\beta_{pt} + \underset{(2.46)}{0.212}\, S_{pt} + \sum_{t=1}^{19} d_t DY_t + e_{pt}. \quad (1)$$

II. Empirical Results

A. Empirical Design Consideration

Since the market risk (beta) of a security is not directly observable, the precision of the estimated betas can be a very important issue in an empirical design. Miller and Scholes (1982) point out that the inverse of price, $1/P$, may be better than the estimated beta as a proxy for risk. This is especially true if a recent drop in the price of a stock reflects an increase in the risk. Thus, a spurious cross-sectional relation between return and a variable which is a function of price (e.g., size, P/E, dividend-yield, relative bid-ask spread) may arise because of the imprecision of the estimated betas. This is the first of the two design issues that we want to address in this study.

The second issue is related to the stochastic nature of the parameters of the asset pricing model. Given the evidence in Fama and French (1988) and Chen, Grundy, and Stambaugh (1990), a variable which is a function of the share price can be related to return in a joint time-series cross-sectional test because it proxies for the changes in the conditional parameters of the asset pricing model. The following experiments provide some direct evidence that the relative bid-ask spread is related to the stochastic expected market premium and the stochastic beta over time.

Let $\bar{S}_t$ be the cross-sectional average of the 49 relative spreads measured the previous year and $\bar{r}_{mt}$ be the average monthly realized excess return of the equally-weighted NYSE index in year t. The following equation is estimated with annual data over the sample period 1961–1980.

$$\bar{r}_{mt} = \underset{(-1.78)}{-0.038} + \underset{(2.19)}{3.194}\,\bar{S}_t + \epsilon_{mt}.$$

Even though the average spread variable is only updated once a year, the result indicates that the relative bid-ask spread, like the dividend-yield, can forecast changes in expected return.

Next, we document the relation between the relative bid-ask spread and the conditional beta. We assume the following simple linear relation:

$$\beta_{pt} = \beta_{0p} + \beta_{1p} S_{pt}, \tag{2}$$

where β_{0p} and β_{1p} are portfolio specific constants and S_{pt} is the relative spread measured in the previous year. Thus the market model of portfolio p becomes

$$r_{pt} = \alpha_{0p} + \beta_{0p} r_{mt} + \beta_{1p} S_{pt} r_{mt} + e_{pt}. \tag{3}$$

In this specification, the beta for portfolio p is stochastic as a linear function of the relative spread if and only if β_{1p} is non-zero. We estimate the above equation with monthly excess returns for all 49 portfolios in the sample period 1961–1980. The test statistic rejects the null hypothesis that all $\beta_{1p} = 0$ at a p-level less than 10^{-7}, indicating that indeed portfolio betas are stochastic as a function of the relative spread.[4]

In the estimation of equation (1) above, observations across portfolios (49 portfolios) and across time (20 years) are stacked together for the regression. A spurious positive relation between the risk-adjusted return and the relative bid-ask spread can arise because (i) the spread variables proxy for the errors in the conditional betas estimated with only 60 monthly observations (from year $t-1$ to $t-5$) — both across portfolios each year and across time for each portfolio, and (ii) the spread variables proxy for changes in the market risk premium since the estimated risk premium in (1) is constrained to be constant over time.

In the empirical tests that follow, each design will address some or all of these problems. In the Fama-MacBeth (1973) cross-sectional regressions, the risk premium is allowed to change every month and the relative bid-ask spread cannot proxy for changes in the conditional beta over time. However, the precision of the estimated betas may not be high because they are only estimated with 60 monthly observations. In the Gibbons (1982)–Stambaugh (1982) SUR framework, assuming that the unconditional return is linearly related to the unconditional beta (see, e.g., Chan and Chen (1988) for sufficient conditions), the precision of the unconditional beta is improved when it is estimated with 240 monthly observations, but the relative bid-ask spread may still proxy for changes in the conditional beta over time. With the Chan and Chen (1988) approach and the Chen, Grundy and Stambaugh (1990) approach, if the assumptions governing the stochastic beta are good working approximations, these two approaches will address both the estimation error problem and the stochastic parameter problem at the same time. In the empirical tests that follow, we find that when the problems discussed above are taken into account, we can no longer observe any reliable relation between the CAPM risk-adjusted return and the relative bid-ask spread.

B. The Fama-MacBeth Approach

All of the parameters necessary to implement the Fama-MacBeth approach have been estimated already and have been used as inputs in estimating (1). The only difference is how the overall estimates for the slope coefficients are obtained. In (1), we stack all the observations across portfolios and across time in one regression. In Fama-MacBeth, we run monthly cross-sectional regressions with the 49 portfolios, thus obtaining a monthly estimate of each slope coefficient, and we use the mean and the standard deviation of the mean of the monthly estimates as the overall estimate of the slope coefficient and its standard error. Unlike (1), the Fama-MacBeth approach allows the estimated market risk premium to change every month. Consequently, the relative bid-ask spread cannot be proxying for the changes in the expected return that arise from the changes in the expected market premium. Furthermore, since the monthly slope coefficients for the spread variable are estimated in cross-sectional regressions, they cannot be contaminated by any positive relation between the relative spread and the estimation error of the conditional beta over time.[5] The drawback of this approach is that only 60 observations are used to estimate the betas and the relative bid-ask spread may be proxying for risk mismeasured in the cross-sectional regressions.

In implementing the Fama-MacBeth approach, the null hypothesis is nested in the alternative hypothesis that there is a positive relation between the risk-adjusted return and the relative bid-ask spread:

$$r_{pt} = \gamma_{0t} + \gamma_{1t}\beta_{pt} + \gamma_{2t}S_{pt} + e_{pt}, \tag{4}$$

where β_{pt} is the portfolio beta estimated with monthly excess return from the previous five years and S_{pt} is the relative bid-ask spread. If the estimated γ_2 is reliably positive, we reject the null hypothesis in favor of the alternative hypothesis. We also examine another alternative hypothesis suggested by AM

$$r_{pt} = \sum_{i=1}^{7} \gamma_{0it} + \gamma_{1t}\beta_{pt} + \sum_{i=1}^{7} \gamma_{2it}\hat{S}^{i}_{pt} + \epsilon_{pt}, \tag{5}$$

where the intercept, γ_{0i}, and the slope coefficient for the spread variable, γ_{2i}, are allowed to be different for each of the seven spread groups. Here, $\hat{S}^{i}_{pt}$ is the mean-adjusted relative spread variable defined as $S_{pt} - \bar{S}_i$ where $\bar{S}_i$ is the average of the relative spreads in group i if portfolio p is in spread group i, and zero if portfolio p is not in spread group i. The results are reported in columns 1 and 2 in Table 1.

When we compare the result in column 1 with the result reported in (1), we notice a striking difference. Recall that we are using numerically exactly the same estimated betas. The main difference arises from the Fama-MacBeth design which does not stack the time series observations in a single regression. This difference in the methodology is sufficient to lower the estimated slope coefficient for the spread variable to less than half of what is reported in (1), and the point estimate of 0.0893 has a t-statistic of only 1.07.

The result reported in column 2 also indicates that there is no reliable relation between the risk-adjusted return and the relative bid-ask spread. A Hotelling T^2 test [$F(7, 233) = 1.133$, p-value$= 0.34$] indicates that we cannot reject the null hypothesis that jointly $\gamma_{21} = \gamma_{22} = \gamma_{23} = \gamma_{24} = \gamma_{25} = \gamma_{26} = \gamma_{27} = 0$.[6] The slope coefficients across the seven spread groups are highly variable and do not conform to any discernible pattern. This is not too surprising because there is little cross-sectional dispersion in the relative spread variable within the same spread group. Overall, even though the market betas may be estimated with errors, we do not find any positive relation between the CAPM risk-adjusted return and the relative bid-ask spread using the classical Fama-MacBeth approach.[7]

C. Results Using the SUR Approach

In this section, the methodology that we use is the nonlinear SUR approach of Gibbons (1982) and Stambaugh (1982). Only the non-overlapping monthly excess returns in the "test" years are used in the tests. Due to the fairly large number of equations, we use all the time-series observations in estimating the system so that the asymptotic distribution is a reasonable approximation. The null hypothesis is that there is a linear relation between the unconditional beta and the unconditional return. With this approach, the precision of the estimated unconditional betas is likely to be high because the unconditional betas are essentially estimated with 240 monthly observations. The only drawback of this approach is that the relative bid-ask spreads may be proxying for changes in expected returns that arise from the changes in the conditional betas.

The model that we estimate is:

$$r_{pt} \;=\; \gamma_0 + \gamma_{1t}\beta_p + \gamma_2 S_{pt} + e_{pt}, \tag{6}$$

where S_{pt} is the relative spread for portfolio p, $\gamma_{1t} = \gamma_1' + r_{mt}$, γ_1' is a constant and r_{mt} is the excess equally-weighted NYSE market return for month t. We also examine the model:

$$r_{pt} \;=\; \sum_{i=1}^{7}\gamma_{0i} + \gamma_{1t}\beta_p + \sum_{i=1}^{7}\gamma_{2i}\hat{S}_{pt}^{i} + e_{pt}, \tag{7}$$

where the intercept, γ_{0i}, and the slope coefficient for the spread variable, γ_{2i}, are allowed to be different for each of the seven spread groups. The results are reported in columns 3 and 4 of Table 1.

With only a single spread variable (column 3), we find that the estimated slope coefficient for the relative spread variable is 0.0998 with an asymptotic t-statistic of 2.00, indicating that it is somewhat related to the return. When we allow each spread group to have its own slope coefficient for the relative spread, most of them are positive. However, they do not resemble the monotonic pattern (AM, Table 4) that may reflect the clientele effect induced by the differential transaction cost. We also test the null hypothesis that jointly $\gamma_{21} = \gamma_{22} = \ldots = \gamma_{27} = 0$ and the test statistic, which is distributed asymptotically as chi-square with 7 degrees of freedom, is 14.42 with a p-value of 0.044. The overall picture again indicates that there may be some relation between return and the relative bid-ask spread, but it is not clear if this relation is arisen from the higher compensations for the larger spreads.

D. The Unconditional Two-Step Approach

In this section, we use the two-step methodology in Chan and Chen (1988) to investigate the relation between spread and return. In the first step, we estimate the unconditional market betas for the 49 portfolios by regressing excess monthly portfolio returns (in the "test" years) on the excess equally-weighted NYSE market return. In the second step, we regress cross-sectionally the monthly returns of the 49 portfolios on the estimated unconditional beta and the portfolio relative spread. Thus we have a series of 240 estimates, one for each month in the sample period 1961 to 1980, for each of the slope coefficients. The overall estimate and the standard error of the slope coefficient are given by the mean and the standard deviation of the mean of the 240 estimates. If the slope coefficient for the spread variable is reliably different from zero, it must reflect a cross-sectional relation between return and spread.

If the assumptions underlying the Chan and Chen model are satisfied,[8] this approach has the advantage that the estimated slope coefficient for the spread variable would not be contaminated by any time-series correlations between the spread and the stochastic market risk and risk premium. We can also use the full sample period observations to estimate the unconditional betas so that the precision is maximized. The models that we estimate are:

$$r_{pt} = \gamma_{0t} + \gamma_{1t}\beta_p + \gamma_{2t}S_{pt} + e_{pt}, \qquad (8)$$

and

$$r_{pt} = \sum_{i=1}^{7}\gamma_{0it} + \gamma_{1t}\beta_p + \sum_{i=1}^{7}\gamma_{2it}\hat{S}_{pt}^{i} + \epsilon_{pt}. \qquad (9)$$

The means of the estimated γ's are reported in columns 5 and 6 in Table 1.

The results are qualitatively similar to those in columns 1 through 4. In column 5, the slope coefficient for the spread variable is 0.026 (t-ratio = 0.29). Both the point estimate and the t-statistic are smaller than those estimated with the Fama-MacBeth methodology (where the precision of the estimated betas is lower) and those estimated with SUR in the previous section (where the relative spread may proxy for the changes in the conditional betas). In column 6, the seven slope coefficients are all within two standard errors from zero with no obvious pattern. A Hotelling T^2 test [$F(7, 233) = 1.024$, p-value= 0.41] indicates that we cannot reject the joint hypothesis that $\gamma_{21} = \gamma_{22} = \ldots = \gamma_{27} = 0$.[9] Therefore, we conclude that there is no reliable evidence relating the CAPM risk-adjusted return to the relative bid-ask spread.

E. An Approach that Models the Stochastic Beta

In the above two sections, we make an implicit assumption about the stochastic behavior of the conditional beta that enables us to test the model with the unconditional moments. Alternatively, we can handle the stochastic conditional beta by modeling it directly. Since one possibility of why the relative bid-ask spread variable is nonzero in a joint time-series cross-sectional test is that the conditional beta may be a function of the relative spread, we model the conditional beta here the same way as in section A:

$$\beta_{pt} = \beta_{p0} + \beta_{p1}S_{pt}.$$

In the corresponding tests for the relation between return and spread, the models that we estimate, using a nonlinear SUR methodology,[10] are, respectively:

$$r_{pt} \quad = \quad \gamma_0 + \gamma_{1t}\beta_{pt} + \gamma_2 \hat{S}^i_{pt} + \epsilon_{pt}, \tag{10}$$

and

$$r_{pt} \quad = \quad \sum_{i=1}^{7} \gamma_{0i} + \gamma_{1t}\beta_{pt} + \sum_{i=1}^{7} \gamma_{2i} \hat{S}^i_{pt} + \epsilon_{pt}. \tag{11}$$

where β_{pt} is defined above and γ_{1t} is defined as in equations (6) and (7). The estimates of the γ's are reported in columns 7 and 8 in Table 1.

Not surprisingly, the results are qualitatively the same as those in columns 1 through 6. In column 7, the estimated slope coefficient for the spread variable is 0.0670 (t-ratio = 1.33). In column 8, the coefficients do not have any discernible pattern across the spread groups. The chi-square statistic for the hypothesis that γ_{21} to γ_{27} are jointly zero is 14.66 (p-value equals 0.041). Overall, there is no strong indication that the data reject the null hypothesis that there is no relation between the CAPM risk-adjusted expected return and the relative bid-ask spread.

III. Conclusion

We investigate the relation between the expected return and the relative bid-ask spread using the same portfolio formation method as in Amihud and Mendelson (1986) but different test methodologies. We do not find any clear reliable relation between the CAPM risk-adjusted return and the relative bid-ask spread. Our empirical results are more consistent with the conclusions of Constantinides (1986) that transaction costs play a minor role in the determination of expected returns in security markets.

Our test methods differ from those by previous authors mainly in our attempt to accommodate the stochastic risk and risk premium and to improve the precision of the estimated betas in the empirical design. If we consider collectively the evidence presented in Chan and Chen (1988), Chen, Grundy and Stambaugh (1990) and in this study, an interesting pattern emerges. When returns are not "properly" adjusted for risk, variables that are functions of the most recently observed price of a stock, such as size, dividend yield and the relative bid-ask spread, are often found to possess explanatory power on the cross-sectional differences in the risk-adjusted return. As Miller and Scholes (1982) point out, the most recently observed price of a security may well be a very good proxy for the security risk, and the explanatory power of a price-related variable can derive from its ability to proxy for risks mismeasured or unmeasured. The evidence presented here as well as the above-mentioned studies are consistent with this view. In the present case, after we estimate the models taking into account the recent empirical regularities related to the stochastic beta, the stochastic risk premium and the importance of the precision of the estimated betas, we do not find any reliable relation between the CAPM risk-adjusted return and the relative bid-ask spread.

Table I
The Relation Between Return and the Relative Bid-Ask Spread (1961–1980)

The table reports estimates of the relation between return and the relative bid-ask spread for eight different models:

$$r_{pt} = \gamma_{0t} + \gamma_{1t}\beta_{pt} + \gamma_{2t}S_{pt} + e_{pt} \quad (1)$$

$$r_{pt} = \sum_{i=1}^{7}\gamma_{0it} + \gamma_{1t}\beta_{pt} + \sum_{i=1}^{7}\gamma_{2it}\hat{S}^i_{pt} + \epsilon_{pt} \quad (2)$$

$$r_{pt} = \gamma_0 + \gamma_{1t}\beta_p + \gamma_2 S_{pt} + e_{pt} \quad (3)$$

$$r_{pt} = \sum_{i=1}^{7}\gamma_{0i} + \gamma_{1t}\beta_p + \sum_{i=1}^{7}\gamma_{2i}\hat{S}^i_{pt} + \epsilon_{pt} \quad (4)$$

$$r_{pt} = \gamma_{0t} + \gamma_{1t}\beta_p + \gamma_{2t}S_{pt} + e_{pt} \quad (5)$$

$$r_{pt} = \sum_{i=1}^{7}\gamma_{0it} + \gamma_{1t}\beta_p + \sum_{i=1}^{7}\gamma_{2it}\hat{S}^i_{pt} + \epsilon_{pt} \quad (6)$$

$$r_{pt} = \gamma_0 + \gamma_{1t}\beta_{pt} + \gamma_2 S_{pt} + e_{pt} \quad (7)$$

$$r_{pt} = \sum_{i=1}^{7}\gamma_{0i} + \gamma_{1t}\beta_{pt} + \sum_{i=1}^{7}\gamma_{2i}\hat{S}^i_{pt} + \epsilon_{pt} \quad (8)$$

Models (1) and (2) are estimated with the Fama-Macbeth (1973) method. Models (3), (4), (7) and (8) are estimated with nonlinear SUR. Models (5) and (6) are estimated with the Chan and Chen (1988) unconditional two-step approach. r_{pt} is the monthly portfolio return for the 49 spread-beta ranked portfolios in excess of 30-day Treasury-bill rate. In models (3), (4), (7) and (8), $\gamma_{1t} = \gamma_1' + r_{mt}$ and r_{mt} is the monthly equally-weighted NYSE market return in excess of 30-day Treasury-bill rate. $\gamma_1 = \gamma_1' + \bar{r}_{mt}$ where $\bar{r}_{mt}$ is the average of r_{mt} over 1961 to 1980. In models (1), (2), (5) and (6), the γ's reported in the table are time-series averages of the monthly estimated γ's. Define S_{pt} as the average relative spread [$\equiv$ (ask$-$bid)/(average of ask and bid)] for portfolio p measured in the previous year and $\bar{S}_i$ is the average of the relative spreads in group i, then $\hat{S}^i_{pt} = S_{pt} - \bar{S}_i$ if portfolio p is in spread group i, and zero otherwise. In models (7) and (8), $\beta_{pt} = \beta_{p0} + \beta_{p1}S_{pt}$.

	Models							
	1	2	3	4	5	6	7	8
γ_0	−0.0048	•	−0.0071	•	−0.0065	•	−0.0065	•
	(−1.95)		(−3.20)		(−2.42)		(−3.28)	
γ_1	0.0108	0.0107	0.0175	0.0044	0.0136	0.0126	0.0175	0.0196
	(2.87)	(2.74)	(4.02)	(4.38)	(2.83)	(2.52)	(4.15)	(4.57)
γ_2	0.0893	•	0.0998	•	0.0260	•	0.0670	•
	(1.07)		(2.00)		(0.29)		(1.33)	
γ_{21}		−4.058		−0.056		−3.713		0.145
		(−1.87)		(−0.06)		(−1.78)		(0.15)
γ_{22}		4.840		0.477		4.688		0.683
		(1.31)		(0.92)		(1.27)		(1.26)
γ_{23}		1.587		0.785		1.776		1.231
		(0.41)		(2.12)		(0.47)		(3.13)
γ_{24}		1.606		0.283		2.134		0.297
		(0.38)		(1.04)		(0.51)		(1.01)
γ_{25}		1.249		0.515		2.495		0.485
		(0.47)		(2.41)		(0.97)		(2.12)
γ_{26}		2.551		0.195		1.711		0.190
		(1.43)		(1.38)		(0.96)		(1.27)
γ_{27}		−0.079		0.171		−0.116		0.068
		(−0.33)		(2.04)		(−0.45)		(0.75)

FOOTNOTES

1. We thank Stoll and Whaley for providing us with the relative bid-ask spread data.

2. They also report the results from another empirical design where they have a separate slope coefficient for each spread group. Their results (p.349, Table 4, column 1) for the seven slope coefficients (with t-statistics in parentheses) are 3.641 (2.76), 3.242 (3.50), 2.854 (3.93), 1.657 (3.06), 2.224 (5.69), 1.365 (5.28) and 0.605 (5.28), respectively. Using their methodology but our data, our estimates for the slope coefficients are 3.649 (2.73), 2.949 (3.24), 2.545 (3.60), 1.773 (3.37), 1.730 (4.51), 0.809 (3.46) and 0.594 (6.59), respectively for the seven spread groups from the smallest to the largest spread group.

3. The standard errors are estimated with $(X'X)^{-1}X'\Sigma X(X'X)^{-1}$, where X is the matrix of observations of the independent variables and Σ is a block diagonal matrix with the block covariance matrix of the residuals for the 49 portfolios on the diagonal. The residuals are estimated with their method (AM, p.237) adapted for the average excess monthly returns. We have also estimated (1) with GLS with the same estimated covariance matrix and the slope of the spread variable is 0.067 (t-ratio = 1.15).

4. The test statistic, which is asymptotically distributed as chi-square with 49 degrees of freedom, is 175.46.

5. For example, if, at every point of time, all the estimated betas are biased towards 1 because of estimation error, but they are still cross-sectionally perfectly correlated with the true betas, then there is still a cross-sectional linear relation between expected return and the estimated betas. However, a variable that is a function of price may still be spuriously related to return in a joint time-series cross-sectional design due to the misestimated conditional betas over time.

6. The t-test and the F-test are based on the null hypothesis that all the slope coefficients are zero. If we test the hypothesis with the assumption that $\gamma_1 \neq 0$, Shanken (1985) shows that under certain assumptions, both the t and the F-statistics should

be adjusted downward. Since the null hypothesis cannot be rejected even without the adjustment, we conclude that there is no reliable relation between risk-adjusted return and the relative bid-ask spread.

7. We have also estimated the models in columns 1 and 2 using GLS in the cross-sectional regressions (with the residual covariance matrix estimated from the residuals in equation (3)), and the results are qualitatively the same. The estimated slope coefficient is 0.085 (t-ratio = 0.90) for the spread variable in equation (4). In equation (5), all of the estimated slope coefficients are negative and the F-statistic for the null hypothesis that all seven spread slope coefficients are jointly zero is 1.26 (p-value = 0.27).

8. The main assumption is that there is a linear function that relates the conditional portfolio betas to the means of the distributions of the conditional portfolio betas cross-sectionally (plus noise). It implies that the unconditional portfolio betas estimated in non-overlapping periods are highly cross-sectionally correlated. The cross-sectional correlation between the 49 unconditional betas estimated over the first 10 years and the 49 unconditional betas estimated over the second 10 years is 0.83. Although the correlation is not as high as that reported in Chan and Chen (1988) for the size portfolios, we regard the assumptions underlying the Chan and Chen (1988) approach here as approximations in order to investigate how robust the results are to different methodologies.

9. We have also estimated the models in columns 5 and 6 using GLS in the cross-sectional regressions and the results are qualitatively the same. The estimated slope coefficient is 0.058 (t-ratio = 0.60) for the spread variable in equation (8). In equation (9), the F-statistic for the null hypothesis that all seven spread slope coefficients are jointly zero is 0.716 (p-value = 0.66). We have also estimated the models described in this footnote and footnote 7 with GLS using the residual covariance matrix estimated from (i) a simple market model for each of the 49 portfolios and (ii) the procedure described in AM (p.237), and the results are qualitatively the same.

10. This is essentially the same as the generalized method of moments (see Hansen (1982)) methodology used in Chen, Grundy and Stambaugh (1990).

REFERENCES

Amihud, Yakov, and Mendelson, Haim, 1986, Asset pricing and the bid-ask spread, *Journal of Financial Economics* 17, 223–249.

Breeden, Douglas T., 1979, An intertemporal asset pricing model with stochastic consumption and investment opportunities, *Journal of Financial Economics* 17, 265–296.

Chan, K. C., and Chen, Nai-fu, 1988, An unconditional asset pricing test and the role of firm size as an instrumental variable for risk, *Journal of Finance* 43, 309–325.

Chen, Nai-fu, Grundy, Bruce, and Stambaugh, Robert F., 1990, Changing risk, changing risk premiums, and dividend yield effects, *Journal of Business* 63, S51–S70.

Constantinides, George, 1986, Capital market equilibrium with transaction costs, *Journal of Political Economy* 81, 842–862.

Cox, John C., Ingersoll, Jonathan E., Jr., and Ross, Stephen A., 1985, An intertemporal general equilibrium model of asset pricing, *Econometrica* 53, 363–384.

Fama, Eugene F., and French, Kenneth R., 1988, Dividend yields and expected stock returns, *Journal of Financial Economics* 22, 3–26.

Fama, Eugene F., and MacBeth, James D., 1973, Risk, return, and equilibrium: empirical tests, *Journal of Political Economy* 81, 607–636.

Gibbons, Michael R., 1982, Multivariate tests of financial models: a new approach, *Journal of Financial Economics* 10, 3–28.

Hansen, Lars P., 1982, Large sample properties of generalized method of moments estimators, *Econometrica* 50, 1029–1054.

Lintner, John, 1965, The valuation of risky assets and the selection of risky investments in stock portfolios and capital budgets, *Review of Economics and Statistics* 47, 13–37.

Merton, Robert C., 1973, An intertemporal capital asset pricing model, *Econometrica* 41, 867–887.

Miller, Merton H., and Scholes, Myron F., 1982, Dividends and taxes: empirical evidence, *Journal of Political Economy* 90, 1118–1141.

Ross, Stephen A., 1976, The arbitrage theory of capital asset pricing, *Journal of Economic Theory* 113, 341–360.

Rubinstein, Mark, 1976, The valuation of uncertain income streams and the pricing of options, *Bell Journal of Economics* 7, 407–425.

Shanken, Jay, 1985, Multivariate tests of the zero-beta CAPM, *Journal of Financial Economics* 14, 327–348.

Sharpe, William, 1964, Capital asset prices: a theory of market equilibrium under conditions of risk, *Journal of Finance* 19, 425–442.

Stambaugh, Robert F., 1982, On the exclusion of assets from tests of the two-parameter model, *Journal of Financial Economics* 10, 237–268.

Stoll, Hans R., and Whaley, Robert E., 1983, Transaction costs and the small firm effect, *Journal of Financial Economics* 12, 57–59.

Zellner, Arnold, 1962, An efficient method of estimating seemingly unrelated regressions and tests of aggregation bias, *Journal of the American Statistical Association* 57, 348–368.