Treasury Bill Rates as Proxies for Expected Inflation

Clifton T. Jones*
Murray State University

Douglas R. Kahl
University of Akron

Jerry L. Stevens
University of Richmond

This paper first replicates tests found in prior studies of the hypothesis that Treasury bill rates are appropriate proxies for expected inflation over a given holding period. The replication is extended using a time-variant coefficient model to test for stability of prior test results. Finally, the test models are examined over more recent periods and over different inflation patterns. Time variant coefficient models, rather than traditional constant coefficient models, are shown to be necessary to test for the use of the Treasury bill rate as a proxy for expected inflation. Results from this study suggest that the Treasury bill rate serves as an appropriate proxy for expected inflation only over periods of low and stable inflation. For other periods, the common practice of using the Treasury bill rate as the expected inflation rate is found to be inappropriate.

INTRODUCTION

The Treasury bill rate commonly is used as a proxy for expected inflation in empirical studies. Yet the conditions required for a high correlation between expected inflation and Treasury bill rates may not hold for all periods. For example, Ratner (1989) presents test results suggesting that the Treasury bill rate is an appropriate proxy for monthly expected inflation over the period studied by Fama and Schwert (1977) but not over a later period studied by Ferris and Makhija (1987).

Ratner's concern for the appropriate use of the Treasury bill rate as a proxy for expected inflation is well-founded. Tests of whether, over a given period, the Treasury bill rate serves as a reliable proxy for expected inflation should be the precondition for such use. Ratner's test method, however, implicitly assumes that the correlation between expected inflation and the Treasury bill rate is constant. Ratner's finding that the Treasury bill rate is a good proxy for expected inflation in the period studied by Fama and Schwert but not in the period studied by Ferris and Makhija simply may reflect time variation in the coefficients of the test equation. Ratner's findings also suggest that time variation in the real interest rate and in the expected inflation-Treasury bill correlation exists in some periods but not in others.

This paper has several purposes. The first aim is to determine whether a replication of Ratner's study over the same periods produces findings consistent with constant coeffi-

* This research was supported in part by research grants from the Stoller Fund, College of Business Administration, the University of Akron, and from the E. C. Robins School of Business, the University of Richmond.

0747-5535/95/1600-0080 $01.00

cients for the test model. A time-variant coefficient model, of which the constant coefficient model is a special case, is used to estimate the coefficients in Ratner's model and test for coefficient stability. A second objective of the study is to analyze the expected inflation–Treasury bill rate relationship in more recent periods and in periods with different inflation patterns. A 40 year interval from 1951 through 1990 is divided into four different subperiods with substantially different inflation patterns to discover whether the relationships between expected inflation and Treasury bill rates vary.

The underlying premise of the Ratner and Fama and Schwert test model is developed, and empirical results from the replication of the tests are provided. Empirical results from a time-variant coefficient extension of the Ratner and Fama and Schwert test model also are provided for the periods of the replication. Finally, the analysis is extended over more recent subperiods and different inflationary environments.

THE FISHER EQUATION AND CONSTANT COEFFICIENT MODELS

Research in finance and economics gives close attention to the time series relationship between nominal rates of return on assets, real interest rates, and inflation.[1] Fisher (1930) was the first to present the notion that nominal interest rates on risk free securities (n) reflect a required real rate of return (r) plus a return covering expected inflation (i). The Fisher equation specification is:[2]

$$(1)\ (1+n_{t-1}) = (1+r_{t-1})(1+i_t) = 1 + r_{t-1} + i_t + (r_{t-1})(i_t).$$

Within this framework, studies focus on topics ranging from money illusion to the role of stocks as a hedge against inflation. In most studies, market data are available for the nominal rate of return, but proxies are needed for the real rate of return and expected inflation. For expected inflation proxies, one approach uses exogenous estimates from either macroeconomic models of interest rate determination (Wilcox, 1983) or asset pricing models (Evans and Wachtel, 1989).

An alternative approach, and the focus of Ratner's paper, was developed by Fama (1975) and is employed by Fama and Schwert (1977) and Ferris and Makhija (1987), among others. Fama notes that interest rates in an efficient market provide a proxy for expected inflation over the holding period of the security if the real rate is constant. Fama's proposition is illustrated by rearranging the Fisher equation

$$(1+i_t) = (1+n_{t-1})/(1+r_{t-1})$$

and taking a log transformation as follows:[3]

[1] For example, see studies by Bodie (1976), Jaffe and Mandelker (1976), Nelson (1976), Fama and Schwert (1977), Fama and Gibbons (1982), Gultekin (1983), Wilcox (1989), Ferris and Makhija (1987), and Evans and Wachtel (1989), to mention only a few.

[2] The last term often is ignored because it is the product of small fractions. If so, the Fisher equation is expressed as $n_{t-1} = r_{t-1} + i_t$.

[3] The log transformation is needed when the full Fisher equation is used. Ratner and others ignore the interaction term of the Fisher equation and simply work with the $i_t = -r_{t-1} + n_{t-1}$ version of the Fisher equation.

$\ln(1+i_t) = -\ln(1+r_{t-1}) + \ln(1+n_{t-1})$.

If the expected real rate of return is a constant and the bond market is rational, an empirical model based on *ex post* inflation and *ex ante* interest rates is obtained:

(2) $I_t = b_0 + b_1 B_t + e_t$

where:

I_t = The log of 1 plus the current observed inflation rate in period t; and
B_t = The log of 1 plus the current Treasury bill rate,[4] which is set at time t-1.

The intercept term (b_0) represents the negative of the log of 1 plus the constant expected real rate of return. If the empirical estimates of b_0 and b_1 do not differ significantly from 0 and 1, respectively, then the Treasury bill rate set at time t-1 (B_t) is an appropriate proxy for expected inflation.

Ratner and Fama and Schwert both test for the appropriate use of the Treasury bill rate as a proxy for expected inflation by using ordinary least squares (OLS) estimation of the empirical model in equation (2). Using a traditional F-test, Ratner supports the maintained hypothesis of H_O: $b_0 = 0$, $b_1 = 1$ for the February 1953 through July 1971 period of the Fama and Schwert study, but rejects it for the January 1977 through December 1983 period of the Ferris and Makhija study. Ratner concludes that because the Treasury bill proxy is shown to be appropriate in some time periods but not in others, researchers should conduct similar tests before using it. The assumption by Ratner and Fama and Schwert of a constant coefficient model as well as the sensitivity of the findings to the model specification, however, remain to be tested. Moreover, Ratner does not explore why the Treasury bill rate is a good proxy in some periods but not others.

REPLICATION OF CONSTANT COEFFICIENT TESTS

The constant coefficient model of equation (2) is estimated using OLS over the two periods examined by Ratner. These results serve mainly to validate data sources and variable measurement across separate studies. Monthly data on Treasury bill rates and the *ex post* rate of inflation as measured by the Consumer Price Index are from the Ibbotson and Associates (1990) data files. For purposes of comparison, the first four rows of Table 1 present both the results from Ratner's paper and the estimated results from the present replication.

The comparative findings shown in the first four rows of Table 1 support the same conclusions reached by Ratner. The Treasury bill rate appears to serve as an appropriate proxy for expected inflation in the February 1953 through July 1971 period, as the inter-

Empirical results, however, do not appear to be sensitive to these alternative versions of the Fisher equation.

[4] Because Treasury bills are discount securities, the rate measure at time t is set at time t – 1.

Table 1—Traditional Constant Coefficient Tests of the Treasury Bill Rate as a Proxy for Expected Inflation (Model: $I_t = b_0 + b_1 B_t + e_t$)

Source	Period	b_0	b_1	R^2	DW	F-test	Obs.
Ratner#	2/1953-7/1971	-.0006** (-1.97)	.91* (9.18)	.27	1.81	.653	222
Replication	2/1953-7/1971	-.0007** (-2.19)	.94* (9.46)	.29	1.77	.421	222
Ratner#	1/1977-12/1983	.0004* (2.95)	.37** (2.33)	.05	.61	15.88*	84
Replication	1/1977-12/1983	.0033* (2.95)	.42* (2.59)	.07	.63	13.34*	84
Low Inflation	1/1951-12/1964	-.0003 (-0.58)	.7127* (2.91)	.05	1.80	1.38	156
Rising Inflation	1/1965-12/1973	-.0009 (-1.05)	1.09* (5.36)	.21	2.00	0.22	108
High Inflation	1/1974-12/1982	.0041* (4.55)	0.40* (3.24)	.08	.70	23.5*	108
Moderate Inflation	1/1983-12/1990	.0030** (2.54)	.0529 (0.28)	.00	1.16	25.3*	96

#Taken from Table 1 on p. 79 of Ratner (1989)

The t-statistics are in parentheses

* and ** indicate significance at the 1 percent and 5 percent levels, respectively

The R^2 is adjusted for degrees of freedom

cept term is close to zero and the slope coefficient is about 1, with an acceptable value for the Durbin-Watson test statistic of 1.77. The F-statistic of 0.421 for the test of the joint restrictions H_O: $b_0 = 0$, $b_1 = 1$ is too low to reject the null hypothesis.

For the January 1977 through December 1983 period the estimated coefficients again are close to Ratner's, failing to support the joint restrictions of the null hypothesis, as evidenced by a strongly significant F-statistic of 13.34. As in Ratner's study, the Durbin-Watson test statistic is significant, indicating the presence of serial correlation in the residuals from the constant coefficient model over this later period. In light of this evidence of possible model misspecification, Ratner uses a Cochrane-Orcutt iterative procedure to correct for first-order autocorrelation, although his conclusions remain unchanged. Overall, the findings in the first four rows of Table 1 demonstrate consistency between Ratner's study and the variable measures and data of this study.

Given the difference in coefficient stationarity for Ratner's two sample periods (February 1953 through July 1971 and January 1977 through December 1983) the appropriateness of the Treasury bill rate proxy over other sample periods and inflation patterns is questioned. To investigate this issue, the constant coefficient model is estimated over four separate subperiods from January 1951 through December 1990, during which inflation follows markedly different patterns.[5] The first subperiod, from January 1950 to December 1964, is characterized by low inflation rates, averaging only 1.8 percent per

[5] We are grateful to an anonymous reviewer for the suggestion to extend the analysis to more recent periods and define subperiods by differences in inflation patterns.

year. The second subperiod, from January 1965 to December 1973, is characterized by rising inflation, going from 1.6 percent in 1965 to 6.2 percent in 1973 and averaging 4.1 percent per year. The third subperiod, from January 1974 to December 1982, is characterized by relatively high inflation, averaging 9.0 percent per year. The fourth and final subperiod, from January 1983 to December 1990, is characterized by moderately low and stable inflation, averaging 3.9 percent per year.

The last four rows of Table 1 present regression results from the constant coefficient test model for the smaller subperiods representing different inflation environments. The insignificant F-statistics for the low and rising inflation periods do not allow rejection of the joint restriction of $b_0 = 0$ and $b_1 = 1$. The coefficient t-statistics also offer support for the use of the Treasury bill as a proxy for expected inflation, because the intercepts are insignificantly different from zero and the slope coefficients are about one. For more recent periods, characterized by high (January 1974 through December 1982) and moderate (January 1983 through December 1990) inflation, the F-statistics and coefficient estimates do not support the use of the Treasury bill rate to estimate expected inflation. These findings are consistent with the periods studied by Ratner, since his test period of February 1953 through July 1971 spans the low and rising inflation subperiods and the test period of January 1977 through December 1983 matches the high inflation period.

All of the regression results in Table 1 are based on an implicit assumption that the slope and intercept coefficients are constant over the test period. As the subperiod regression results show, the coefficients change dramatically from period to period, suggesting instability. It is also possible that the coefficients of the test equation exhibit random variation within a test period, violating a necessary restriction required for using the Treasury bill rate as a proxy for expected inflation.

AN ALTERNATIVE TIME-VARYING COEFFICIENT MODEL

The constant coefficient model of equation (2) is a special case of a more general time-variant coefficient model. When written in state space form, the time-variant coefficient model consists of one measurement equation,

$$(2') \quad I_t = b_{0t} + b_{1t} B_t + e_t,$$

and two state equations,

$$(2'') \qquad b_{0t} = b_{0,t-1} + u_{1t}$$

and

$$(2''') \qquad b_{1t} = b_{1,t-1} + u_{2t},$$

where the two disturbance terms (u_{1t}, u_{2t}) in the state equations are assumed to be uncorrelated with each other and with the disturbance term in the measurement equation (e_t).

The time series processes followed by the coefficients of the measurement equation (b_{0t}, b_{1t}) are indicated by the standard errors of the disturbance terms in the state equations (σ_1, σ_2). Should one (or both) of these standard errors be nonzero, the associated coefficient(s) follow a simple random walk.[6] There is no tendency for the coefficient to return to any particular value or underlying mean in a random walk process, so the best estimate for the time-variant coefficient in this case is its most recent value, b_{it-1}. On the other hand, whenever one of these standard errors is zero, then $b_{it} = b_{i,t-1}$ for all t, and the associated coefficient is a constant. If both standard errors are zero, then both coefficients are constants and the measurement equation (2') is identical to the empirical model of equation (2).

Kahl and Ledolter (1983) show how a recursive Kalman filtering approach obtains simultaneous maximum likelihood estimates of the two parameters of the general model[7] of equations (2'), (2"), and (2'"): σ_1 and σ_2. A likelihood ratio statistic proposed by Ledolter (1979) and Kahl and Ledolter (1983) (distributed as χ^2_p, where the degrees of freedom, p, is given by the number of time-variant coefficients in the general model) provides a test of the null hypothesis that both parameters (σ_1, σ_2) are jointly zero (i.e., a constant coefficient process), against the alternative that they are non-zero (i.e., a time-variant coefficient process).

Empirical results from estimation of the time-variant coefficient model are provided in Table 2. For the period February 1953 through July 1971 the maximum likelihood estimate of σ_1 is zero, while σ_2 is estimated to be 0.0305. The likelihood ratio test statistic's value of 3.40 is not statistically significant, even at the 10 percent level; the null hypothesis of a constant coefficient model ($\sigma_1 = \sigma_2 = 0$) cannot be rejected. This result

Table 2—Maximum Likelihood Estimates of the Time-Variant Coefficient Model Under Different Inflation Patterns

Time Period	Standard Errors for State Equations σ_1	σ_2	Likelihood Ratio Test Statistic χ^2_2
2/1953-7/1971 Ratner Period	0.0	0.0305	3.40
1/1977-12/1983 Ratner Period	0.0025	0.0097	47.84*
1/1951-12/1964 Low Inflation	0.0	0.0	0.0
1/1965-12/1973 Rising Inflation	0.0	0.0577	8.68**
1/1974-12/1982 High Inflation	0.0012	0.0044	57.38*
1/1983-12/1990 Moderately Low Inflation	0.0003	0.0	5.22

* and ** indicate significance at the 1 percent and 5 percent levels, respectively

[6] While there are other, more complex time series processes that could be specified for the model coefficients, we follow Garbade (1977), who uses a random walk model to examine the stability of regression coefficients in the demand for money in the U.S. In his study, Garbade shows that the well-known Cooley-Prescott (1976) model is a reparameterization of this simple random walk model.

[7] The recursive Kalman filtering algorithm allows coefficient estimates to vary with each monthly observation in the sample. For more detailed applications of the procedure, see Gombola and Kahl (1990).

confirms Ratner's (and Fama and Schwert's) untested assumption of constant coefficients over the February 1953 through July 1971 period.

In contrast, for the period January 1977 through December 1983 the maximum likelihood estimates of σ_1 and σ_2 are both nonzero (0.0025 and 0.0097), and the likelihood ratio test statistic's value of 47.84 is highly significant, resulting in rejection of the null hypothesis of a constant coefficient model ($\sigma_1 = \sigma_2 = 0$). Both the intercept term (b_{0t}) and the slope term (b_{1t}) follow a random walk over time. These results, with time-variant properties for both b_0 and b_1, suggest constantly changing monthly patterns for expected inflation during this later period. Thus, the inappropriateness of the Treasury bill rate proxy over this later period is attributable in part to the presence of significant time variation in the model's coefficients, a variation not shown for the earlier period studied by Fama and Schwert.

Table 2 also presents maximum likelihood estimates of σ_1 and σ_2 and the likelihood ratio test statistics for each of the four subperiods used in Table 1. During the 1950s and early 1960s when inflation was low, both standard errors are estimated to be zero, leaving a constant coefficient model as in equation (2). For the next subperiod of rising inflation from the mid-1960s to the first oil price shock of 1973, the likelihood ratio test statistic indicates that the constant coefficient model is rejected at the 5 percent level, because the estimate of σ_1 is still zero but the estimate of σ_2 is nonzero (0.0577). This finding illustrates the importance of testing for time variation in subperiods, as the January 1965 through December 1973 period is part of the longer February 1953 through July 1971 period tested by Ratner.

For the third subperiod of high inflation, from 1973 through 1982, estimates of both σ_1 and σ_2 are nonzero (0.0012, 0.0044) and the likelihood ratio test statistic is highly significant. The constant coefficient model is easily rejected. Finally, for the last subperiod, after 1982 when inflation is stable and moderately low, the estimate of σ_1 is nonzero (0.0003) but the estimate of σ_2 is zero. In this final case the likelihood ratio test statistic is again too low to be significant, so the constant coefficient model cannot be rejected.

Overall, the results in Table 2 demonstrate a relationship between different inflation patterns and time variation of the coefficients in equation (2). In general, constant coefficient models are more appropriate in periods of low or moderately low and stable inflation. For periods of high and rising inflation, there is significant time variation in the coefficients of equation (2). This finding is intuitive, as the assumptions of the constant coefficient model are most likely to be violated in periods of high and rising inflation. Differences between *ex post* inflation and expected inflation should be greatest when inflation is high and difficult to predict. Note also that the *ex post* real rate of interest varies as *ex post* inflation deviates from the expected inflation rate. In the context of the Fisher equation (equation (1)), the *ex post* real rate is lower (higher) when inflation is higher (lower) than expected.[8]

[8] The findings are also consistent with the notion that inflation tends to be overestimated in periods following high and rising inflation. Once lenders experience the consequences of unexpected inflation, it takes a long time to lower expectations. In this case, the *ex post* inflation rate is lower than the *ex ante* inflation rate, and the *ex post* real rate in equation (2) is higher. The net result is a positive correlation between inflation and the *ex post* real rate estimate (b_{0t}).

Table 3—Correlation Between Time-Variant Coefficients and Inflation

Time Period	Correlation Between I_t and b_{0t}	Correlation Between I_t and b_{1t}
1/1951-12/1964 Low Inflation	N/A	N/A
1/1965-12/1973 Rising Inflation	N/A	.3772*
1/1974-12/1982 High Inflation	.8610*	-.1115
1/1983-12/1990 Moderately Low Inflation	.2486*	N/A

* and ** indicate significance at the 1 percent and 5 percent levels, respectively

N/A means that correlations between coefficients and inflation are not applicable because a constant coefficient model is supported by the likelihood ratio test statistics

The results in Table 2 introduce the possibility that the coefficients of the test model vary systematically with the inflation rate. To test this hypothesis, correlations between monthly coefficient estimates and monthly inflation rates are calculated for the 1974 through 1982 and 1983 to 1990 subperiods, where coefficients are shown to have time variation. Results of the correlation analysis appear in Table 3; three of the four possible correlation coefficients are positive and significant at the 5 percent level. The estimate of the intercept term (b_{0t}), representing the real rate of interest, is significantly positively correlated with inflation in both subperiods where it shows time variation. The estimate of the expected inflation–Treasury bill rate relationship (b_{1t}) has a significantly positive correlation with inflation in the 1965 to 1973 subperiod, where inflation is rising. Inflation does not have a significant correlation with the variation of coefficient b_{1t} during the 1974 through 1982 subperiod, however, suggesting that more than just the level of inflation may be a factor in coefficient instability.

Collectively, the findings of Table 2 and Table 3 offer an intuitive explanation for constant coefficients in equation (2) over some periods but not others. The constant coefficient model appears to be appropriate in periods where the inflation rate is low and stable, making it easier to predict. In a period of low inflation the differences between *ex post* inflation and expected inflation are likely to be small, meeting an important assumption of the constant coefficient model of equation (2). The finding of a positive correlation between *ex post* inflation and the real rate estimate (b_{0t}) in subperiods when the intercept term is time variant also is consistent with macroeconomic phenomena. The marginal productivity of capital, represented by the real rate of interest, and inflation both tend to rise with economic expansion and to be more stable during periods of slow growth. Thus, a positive correlation exists between the real rate of interest (b_{0t}) and *ex post* inflation in periods when the real rate varies significantly.

CONCLUSIONS

This paper explores the common practice in time series studies of using the Treasury bill rate as a proxy for expected inflation. Ratner's tests of the validity of the Treasury

Table 4—Summary of Test Results for Time Varying Coefficients and Joint Restrictions on Constant Coefficients

Time Period	Appropriate Model	Treasury Bill Rate as a Proxy for Expected Inflation
Low Inflation 1/1951-12/1964	Constant Coefficient Equation (2)	Yes
Rising Inflation 1/1965-12/1973	Time Varying Coefficient Equation (2′)	No
High Inflation 1/1974-12/1982	Time Varying Coefficient Equation (2′)	No
Moderate Inflation 1/1983-12/1990	Constant Coefficient Equation (2)	No

bill rate as a proxy for expected inflation are replicated over periods for which other authors (Fama and Schwert and Ferris and Makhija) use it. Ratner's constant coefficient test model is extended to allow for a more flexible time-variant coefficient model. Results from the time-variant model indicate that the constant coefficient test model and the use of the Treasury bill rate as a proxy for expected inflation are appropriate only for a period of low inflation (January 1951 to December 1964). In all later periods, however, the coefficients of the test model are either time-variant or do not conform to the joint hypothesis that $b_0 = 0$ and $b_1 = 1$, invalidating use of the Treasury bill rate as a proxy for expected inflation. Table 4 summarizes these findings.

Results from the replication of Ratner's study demonstrate that the coefficients of the test model are constant in some subperiods and not in others. In general, the coefficients of the test model are constant in subperiods characterized by low and stable rates of inflation. Tests for stability of the coefficients must be conducted first before results of constant coefficient test equations can be interpreted.

For subperiods where time variation in coefficients of the test model is found, findings of significant positive monthly correlations between *ex post* inflation and estimated coefficients reinforce the conclusion that variation in the test model's coefficients is related to inflation patterns. Extensions of this analysis to consider systematic overestimation (underestimation) of expected inflation in different inflation environments are straightforward, but beyond the scope of this paper.

REFERENCES

1. Bodie, Z., "Common Stocks as a Hedge Against Inflation," *Journal of Finance*, 31 (1976), pp. 459-470.

2. Cooley, T.F., and E.C. Prescott, "Estimation in the Presence of Stochastic Parameter Variation," *Econometrica*, 44 (1976), pp. 167-184.

3. Evans, M., and P. Wachtel, "A Modern Approach to the Pricing of Riskless Bonds," working paper, New York University, 1989.

4. Fama, E.F., "Short Term Interest Rates as Predictors of Inflation," *American Economic Review*, 65 (1975), pp. 269-282.

5. Fama, E.F., and M. Gibbons, "A Comparison of Inflation Forecasts," *Journal of Monetary Economics*, 13 (1984), pp. 327-351.

6. Fama, E.F., and G.W. Schwert, "Asset Returns and Inflation," *Journal of Financial Economics*, 5 (1977), pp. 115-146.

7. Ferris, S.P., and A.K. Makhija, "A Search for Common Stock Inflation Hedges," *Review of Business and Economic Research* (Spring 1987), pp. 27-36.

8. Fisher, I., *The Theory of Interest: As Determined by Impatience to Spend Income and Opportunity to Invest It* (New York: Augustus M. Kelly Publishers, 1930, republished in 1965).

9. Garbade, K., "Two Methods for Examining the Stability of Regression Coefficients," *Journal of the American Statistical Association* (March 1977), pp. 54-63.

10. Gombola, M., and D.R. Kahl, "Time-Series Processes of Utility Betas: Implications for Forecasting Systematic Risk," *Financial Management* (Autumn 1990), pp. 84-93.

11. Gultekin, N.B., "Stock Market Returns and Inflation: Evidence from other Countries," *Journal of Finance*, 38 (1983), pp. 49-65.

12. Ibbotson, R.G., *Stocks, Bonds, Bills and Inflation* (Chicago, Illinois: Ibbotson Associates Incorporated, 1990).

13. Jaffee, J., and G. Mandelker, "The 'Fisher Effect' for Risky Assets: An Empirical Investigation," *Journal of Finance*, 31 (1976), pp. 447-458.

14. Kahl, D.R., and J. Ledolter, "A Recursive Kalman Filter Forecasting Approach," *Management Science*, (Autumn 1983), pp. 1325-1333.

15. Ledolter, J., "A Recursive Approach to Parameter Estimation in Regression and Time Series Models," *Communications in Statistics*, A8 (1979), pp. 1227-1245.

16. Nelson, C.R., "Inflation and Rates of Return on Common Stock," *Journal of Finance*, 31 (1976), pp. 471-483.

17. Ratner, M., "Testing a Proxy for Expected Inflation: A Note," *Review of Business and Economic Research*, 25 (Fall 1989), pp. 78-81.

18. Wilcox, J.A., "Why Interest Rates Were So Low in the 1970s," *American Economic Review*, 73 (March 1983), pp. 44-53.