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INFLATION AND REAL INTEREST RATES ON ASSETS WITH DIFFERENT RISK CHARACTERISTICS

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ABSTRACT

Several recent studies find that ex ante real returns for short-term U.S. Treasury securities are negatively correlated both with inflation and with nominal interest rates. This paper examines whether these findings extend to the short-term holding return on publicly and privately issued securities of longer maturity, are robust with respect to the choice of price index, and are stable over time. Our results show that before 1979 a negative relationship of ex ante real returns with inflation and nominal interest rates does appear for the longer maturity assets. In fact, the relationship grows stronger with increases in maturity length. This suggests that although short-term U.S. Treasury bills were, of all the assets we study, the best hedge against expected inflation, none of the assets were a perfect hedge. We find a statistically significant change in the stochastic process of bond returns in 1979, with nominal interest rates and ex ante real holding returns being positively correlated in this latter period. This is not true for stocks, however. While the above results are robust to the choice of price index, we show that estimating the level of ex ante real returns depends crucially on the price index chosen.

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Given the importance of ex ante real interest rates to economic decision making, it is no surprise that much empirical research focuses on them.[1] Several recent studies find that the ex ante real returns for short-term U.S. Treasury securities are not constant and are negatively correlated both with inflation and with nominal interest rates. This paper examines the extent to which these findings hold for assets of longer maturity, are robust with respect to the choice of price index, and are stable over time.

We expand the set of assets studied to include those of longer maturity, both publicly and privately issued, for two reasons. First, their ex ante real returns may differ substantially from that of a short-term nominally riskless security. Second, because longer-maturity assets have risk characteristics closer to capital, their ex ante real returns are potentially more relevant for investment and savings decisions. We explore the use of various price indices because the index used in most studies, the Consumer Price Index, severely overstates inflation during the 1970s as a result of its inappropriate treatment of residential housing costs.[2] This overstatement casts doubt on an important finding of previous studies -- that there are long stretches of time during the 1970s when ex ante real rates are significantly negative. Given the existence of holding costs for the goods relevant to most price indices and the absence of a riskless real return, no arbitrage conditions are violated by a negative ex ante real rate. However, its persistence is problematical because asset holders are typically assumed to have a positive rate of time preference, which ensures a positive ex ante real rate in the long run.

We find that before 1979 a negative relationship of ex ante real rates both with inflation and with nominal interest rates appears for the longer maturity assets as well as the short-term nominally riskless asset; indeed the relationship grows monotonically with the maturity of the asset. This suggests that although short-term U.S. Treasury bills are, of all the assets we study, the best hedge against expected inflation, none of the assets are perfect hedges.[3] Our results also indicate that nominal interest rates are not a reliable indicator of the tightness of credit markets, since they are negatively correlated with ex ante real returns on all assets examined.

The use of a more accurate price index than the Consumer Price Index results in noticeably different estimates of ex ante real rates -- rates that are generally higher and statistically significantly negative for far shorter periods of time. Our findings on the relationship of ex ante real interest rates with inflation and nominal interest rates are, however, robust to the choice of price index.

When we examine the behavior of financial markets since October 1979, we find a statistically significant change in the stochastic process of ex ante real returns on bonds. In contrast to the earlier period, after October 1979 nominal interest rates are positively correlated with ex ante real returns. However, this change in the stochastic process appears only for bonds and not for common stocks. Given the small amount of data since 1979, we clearly consider these results tentative; nonetheless they suggest topics for future research.

II. METHODOLOGY

Following Fisher (1930), at each time t the ex ante, or expected, real return earned from holding an asset for j periods can be decomposed into an expected nominal return and an expected inflation rate:

$$rr_{t,j} = i^e_{t,j} - \pi^e_{t,j} , \qquad (1)$$

where $rr_{t,j}$ is the expectation at time t of the real return earned by holding an asset from time t to time $t+j$, $i^e_{t,j}$ is the expectation at time t of the nominal return earned by holding an asset from time t to time $t+j$, and $\pi^e_{t,j}$ is the expectation at time t of the inflation rate from time t to time $t+j$. Similarly, the ex post, or realized, real return can be decomposed into an ex post nominal return and an ex post inflation rate:

$$eprr_{t,j} = i_{t,j} - \pi_{t,j} , \qquad (2)$$

where $eprr_{t,j}$ is the realized real return earned by holding an asset from time t to time t+j, $i_{t,j}$ is the realized nominal return earned by holding an asset from time t to time t+j, and $\pi_{t,j}$ is the realized inflation rate from time t to time t+j. Given equations (1) and (2), the ex post real return can be expressed as

$$eprr_{t,j} = rr_{t,j} + \varepsilon 1_{t,j} - \varepsilon 2_{t,j}, \qquad (3)$$

where $\varepsilon 1_{t,j} = i_{t,j} - i^e_{t,j}$ and $\varepsilon 2_{t,j} = \pi_{t,j} - \pi^e_{t,j}$. From equation (3) it is clear that ex ante and ex post real returns will diverge when either the nominal return on the asset or the inflation rate is uncertain. It is also clear that if the nominal interest rate and inflation surprises are observable, the ex ante real interest rate is observable. Much measurement of ex ante real returns is based on equation (3), with survey data used to construct $\varepsilon 2_{t,j}$ and a j-period default free bond that ensures $\varepsilon 1_{t,j} = 0$.

We use an alternative procedure to estimate ex ante real interest rates. The procedure does not require survey measures of inflation,[4] allows for estimation of the ex ante real return on all assets - not just j-period default free bonds - and simultaneously allows us to measure the correlation of ex ante real rates with variables that are observable at time t. The critical assumption underlying the methodology is rationality of expectations; this assumption ensures that forecast errors are uncorrelated with past information. Specifically,

$$E(\varepsilon 1_{t,j}|\psi_t) = E(\varepsilon 2_{t,j}|\psi_t) = 0, \quad (4)$$

where ψ_t represents all the information available to agents in the economy at time t and $E(\ |\psi_t)$ is the mathematical expectations operator, conditional on ψ_t. While ψ_t is the information set used by economic agents to form their ex ante real returns, we assume that as econometricians we can observe only a subset of ψ_t, X_t. Since any estimate of ex ante real returns must be based solely on X_t, a logical choice is the best linear predictor of $rr_{t,j}$ given X_t, which we represent as

$$P(rr_{t,j}|X_t) = X_t\beta. \qquad (5)$$

Note that when the ex ante real interest rate is constant, $\beta = 0$ for all regressors except a constant term.

Obviously $P(rr_{t,j}|X_t)$ cannot be used directly, because β is unknown, and so we use

$$\hat{rr}_{t,j} = X_t\hat{\beta}, \qquad (6)$$

where $\hat{\beta}$ is an estimate of the unknown parameter β. To obtain $\hat{\beta}$, we estimate by ordinary least squares the equation

$$eprr_{t,j} = X_t\beta + \eta_{t,j}, \qquad (7)$$

where $u_{t,j} = rr_{t,j} - P(rr_{t,j}|X_{t,j})$ and $\eta_{t,j} = u_{t,j} + \varepsilon1_{t,j} - \varepsilon2_{t,j}$. A separate equation is estimated for each of the seven assets we study.

There are several important econometric issues concerning this estimation strategy that need to discussed. Formal demonstration of the following points can be found in Mishkin (1981b, 1982, 1984) and Cumby, Huizinga and Obstfeld (1983).

1. Given the rationality of expectations, $\hat{\beta}$ will be a consistent estimate of β. That is, using ex post real rates will asymptotically yield the same estimate of β as a regression using ex ante real rates. Although we cannot observe the ex ante real rate, we can estimate it and can infer information about its relationship with other variables known at time t via ex post real rate regressions.
2. We do lose information by using ex post real interest rates in the regression rather than ex ante rates. The presence of the forecast errors $\varepsilon1_{t,j}$ and $\varepsilon2_{t,j}$ in equation (7) means that β will be estimated less precisely; that is, the standard errors of the estimated paramters will be larger. In addition, as the maturity of the asset lengthens we expect more volatility in the asset price, so the variance of the holding period return

should increase, increasing the variance of $\varepsilon 1_{t,j}$. We thus expect less precise estimates of β for common stocks and long-term bonds than for short-term bills.

3. The coefficient β, and hence its estimate $\hat{\beta}$, does not imply that X_t causes the ex ante real interest rate, only that X_t helps to predict it. Without further information, β should be interpreted only as measuring the correlation between X_t and ex ante real returns.

4. For the case $j > 1$ that we consider here, the data is overlapping; that is, the holding period is longer than the observation interval. In this case, estimated standard errors of the parameters reported by typical regression packages will generally be inconsistent. The problem stems from the fact that when $j > 1$, $\varepsilon 1_{t-1,j}$ and $\varepsilon 2_{t-1,j}$ are not in the information set ψ_t because they are not known until time $t+j-1 > t$. Thus, in contrast to the nonoverlapping case of $j=1$ where the rational expectations conditions of equation (4) imply that $\varepsilon 1_{t,j}$ and $\varepsilon 2_{t,j}$ are serially uncorrelated, $\varepsilon 1_{t,j}$, $\varepsilon 2_{t,j}$ and therefore $\eta 2_{t,j}$ are likely to be serially correlated. In this case, consistent estimates of the standard errors for $\hat{\beta}$ can be computed with methods outlined by Hansen (1982) or Cumby, Huizinga and Obstfeld (1983). The estimates used here are similar to those used by Hansen and Hodrick (1980) but are more general, because they allow for conditional heteroscedasticity of the regression residuals. Tests of the null hypothesis that the residuals were conditionally homoscedastic indicated rejection in our sample.[5]

5. With respect to testing hypotheses such as the constancy of ex ante real returns, tests that are joint for all assets together will have increased power even though the equation for each asset is efficiently estimated in isolation. In doing a joint test, however, the correlation of the $\hat{\beta}$ estimates obtained from regression equations for different assets must not be ignored. One reason this correlation can be expected to be nonnegligible is that all regression errors contain the common element $\varepsilon 2_{t,j}$, unexpected inflation.

III. DATA

This paper analyzes monthly data from 1959 to 1981 on real returns over a three month holding period for the following seven securities: (1) three month Treasury bills (TBILL3), (2) six month Treasury bills (TBILL6), (3) twelve month Treasury bills (TBILL12), (4) intermediate-term (5-10 years maturity) Treasury bonds (INTBOND) (5) long-term Treasury bonds (over 10 years maturity) (LONGBOND), (6) long-term corporate bonds (CORPBOND), (7) common stocks (STOCK). The nominal return data is obtained from the Center for Research in Security Prices (CRSP) at the University of Chicago. Because there are no Treasury bills with more than three months of maturity until 1959, the data for the six- and twelve-month bills cannot be collected before this date. In fact, twelve month bond data are not available until late 1963. The bonds chosen for the intermediate-term bond returns are an update (supplied to us by Lawrence Fisher) of the same securities used in Fisher and Lorie (1977). The bonds chosen for the long-term bond returns are an update (supplied by Roger Ibbotson) of those used in the Ibbotson and Sinquefeld (1976) long-term Treasury bond return index. The corporate bond data are derived from the Ibbotson-Sinquefeld corporate bond returns index and the CRSP NYSE value-weighted index, is used to calculate the stock returns.

The three month holding period has been chosen because a shorter holding period would increase timing problems created by the sampling methods used to construct the price indices. The problem is that the price components in the indices are sampled at different times over the course of the month. Thus it is not clear what is the appropriate dating for the price index. With monthly holding periods, the dating can easily be off by as much as half of the holding period. With three month holding periods, the dating can be off by only one-sixth.

Four price indices are used here to construct inflation and real interest rates. The first is the CPI, the price index most frequently used in empirical research on real rates. In January of 1983 the Bureau of Labor Statistics began calculating the official CPI using a new "rental equivalence" measure for the residential housing component. Our second price index, denoted CPIX, reconstructs the CPI from 1959 to 1981 using this new procedure.[6] This is likely to be the most accurate of our four price indices and its use should lead to the most reliable measures of inflation and real rates. The third price index is the Personal Consumption Expenditure Deflator, denoted PCED, which comes from the National Income and Product Accounts but is available monthly. The index, like the CPIX, does not suffer from the inappropriate treatment of housing found in the CPI, but it differs from the CPIX in being a variable weight index. The PCED is not without problems for our study; unlike other indices, it is available only in seasonally adjusted form and, more important, it uses more interpolation of prices than the other indices do. Finally, we use the Producer Price Index, denoted PPI. It does not include residential housing in any manner but has the drawback that it is constructed with list prices rather than transactions prices.

IV. EMPIRICAL RESULTS

A. Results for the pre-October 1979 period

The relationship of real rates with inflation and nominal interest rates for the 1959:5 to 1979:10 sample period is described in Table 1. The post 1979:10 period is dealt with later in the paper because, as we shall see, the stochastic process of real rates undergoes a major shift near the end of 1979. The inflation coefficients are generated from regressions where the X_t variables include a constant term and the inflation rate for the past three- months. The nominal three month bill coefficients are obtained from regressions where the X_t variables include a constant term and the nominal rate on three- month U.S. Treasury bills known at time t. The timing of both the inflation rate and the nominal bill rate clearly reflects that X_t must be a subset of ψ_t, the information set agents have available to form their ex ante real rates for period t.[7]

The inflation coefficients tell us that ex ante real rates on all seven assets are negatively correlated with inflation. Regardless of the price index used, all

TABLE 1

Relationship of Ex Ante Real Rates with Inflation and Nominal Interest Rates
1959:5 to 1979:10

Estimation of β in eprr = $\alpha + \beta$ X

X	INFLATION				NOMINAL THREE MONTH BILL RATE			
PRICE INDEX[a]	CPI	CPIX	PCED	PPI	CPI	CPIX	PCED	PPI
ASSET								
TBILL3	-.349** (.047)	-.280** (.050)	-.242** (.060)	-.309** (.096)	-.562** (.101)	-.319** (.103)	-.206* (.097)	-1.015** (.310)
TBILL6	-.354** (.062)	-.282** (.062)	-.240** (.082)	-.312** (.105)	-.556** (.118)	-.313** (.112)	-.200 (.113)	-1.009** (.322)
TBILL12	-.433** (.145)	-.379** (.144)	-.338 (.190)	-.312* (.149)	-.627* (.273)	-.343 (.264)	-.216 (.274)	-.977* (.492)
INTBOND	-.746* (.376)	-.701 (.390)	-.809 (.440)	-.478 (.264)	-1.432** (.606)	-1.189* (.597)	-1.077 (.613)	-1.886** (.732)
LONGBOND	-.777 (.518)	-.648 (.538)	-.774 (.616)	-.374 (.350)	-1.711* (.795)	-1.468 (.787)	-1.355 (.809)	-2.164** (.912)
CORPBOND	-.971 (.638)	-.878 (.677)	-1.060 (.783)	-.540 (.430)	-2.153* (.971)	-1.910* (.962)	-1.797 (.983)	-2.606** (1.102)
STOCK	-1.793 (1.126)	-1.924 (1.240)	-2.095 (1.444)	-1.390** (.552)	-3.645 (1.878)	-3.402 (1.883)	-3.289 (1.900)	-4.098* (2.008)
χ^2 VALUE[b]	71.85	46.75	35.93	25.37	48.16	23.56	16.86	23.90
MARGINAL SIGNIFICANCE LEVEL	$<10^{-10}$	6×10^{-8}	8×10^{-6}	7×10^{-4}	3×10^{-8}	1×10^{-3}	2×10^{-2}	1×10^{-3}

Notes: Numbers in parentheses are estimated standard errors.
- * denotes significantly different from zero at the 5% level.
- ** denotes significantly different from zero at the 1% level.
- a price index used to construct eprr and, for the first four columns, the regressor X.
- b distributed according to the χ^2 distribution with seven degrees of freedom under the null hypothesis that all coefficients in the column are zero.

the inflation cefficients are negative, many of them significantly so.[8] An extremely interesting finding is that as maturity length increases, and the asset gets closer in its risk characteristics to capital, an increase in inflation is associated with an even larger decrease in the ex ante real return. For example, the CPIX results indicate that a one-percentage-point increase in the inflation rate is associated with a 28-basis-point drop (a basis point is 1/100 of a percentage point) in the ex ante real rate on three month U.S. Treasury bills, a 65-basis-point decline in the ex ante real rate on long-term U.S. Treasury bonds, and a 192-basis-point decline in the ex ante real rate on common stocks.[9]

The χ^2 statistic at the bottom of each column tests the null hypothesis that all seven coefficients in the column equal zero.[10] This is a joint test of the constancy of ex ante real rates on all seven assets. The low values of the marginal significance levels indicate that the null hypothesis can always be rejected at the 1% level. As the asterisks indicate, constancy of ex ante real rates is rejected less frequently for the longer maturity assets. Despite their more negative coefficients, there is less statistical significance because the standard errors increase. As described earlier, this is exactly what we expect since the forecast error of the nominal return, $\varepsilon 1_{t,j}$, should become more variable with lengthening maturity. Both because of this and because of the consistent pattern of the coefficients, our view is that an appropriate interpretation of the Table 1 results is that ex ante real returns on intermediate term bonds, long-term bonds, and common stocks are negatively associated with inflation. The absence of statistical significance is a reflection of the low power of the tests.[11]

The nominal three month bill rate coefficents follow a pattern similar to that of the inflation coefficients. They are all negative, many are significantly so, and they increase as we go to the longer maturity assets. Again, constancy of the ex ante real return for all seven assets is soundly rejected by the χ^2 values. Rejection is widely spread across the seven assets. The negative coefficients indicate that a nominal interest rate, such as that on three month U.S. Treasury bills, is a poor indicator of the tightness of credit markets. When we use the nominal yield to maturity for each asset as the explanatory variable in the regression for that asset, the findings are similar. During times of high nominal rates, credit market conditions are not tight; rather the contrary is true, with low ex ante real rates on all assets.[12]

For the sample period studied here, previous evidence has found that movements in the nominal three month bill rate primarily reflects movements in expected inflation.[13] Since inflation is a highly positively autocorrelated series, the inflation rate over the past three months should also be highly positively correlated with the expected inflation rate. Thus the inflation and interest rate coefficients of Table 1 convey information about how ex ante real rates are correlated with expected inflation. They suggest that, while of all the assets we study three month U.S. Treasury bills are the best hedge against expected inflation, all assets are an imperfect hedge. Additional evidence supporting this proposition is presented in Huizinga and Mishkin (1983). Our results are in general agreement with those of Fama and Schwert (1977), who also find that as assets lengthen in maturity and become more like equity, they become worse inflation hedges. However Fama and Schwert's conclusion that U.S. Treasury bills and bonds of five-year maturity or less are reasonably good hedges against expected inflation is not supported here. The main reason for the difference is that our sample includes the high inflation sample period after 1971 while Fama and Schwert's does not. As has been documented in Nelson and Schwert (1977), Mishkin (1981b), and Fama and Gibbons (1982), the period before 1971 is one with little variation in ex ante real rates and inflation. This makes it difficult to detect a negative relationship between ex ante real rates and expected inflation. However, with the additon of the highly variable data after 1971, the negative asociation clearly emerges.

The conclusions described above are robust to the use of different price indices -- a comforting finding. We should still be concerned, however, that the overstatement of inflation by the CPI in the 1970s may lead to misleadingly low estimates of ex ante real interest rates. To see how severe the bias is from using the CPI, we plot in Figure 1 the estimated ex ante real return for three month U.S. Treasury bills using both the CPI, the broken line, and CPIX, the solid line. We use the CPIX because we consider it to be the most reliable of our price indices. Although it is not reported here, the effect of choosing the CPIX instead of the CPI is similar for the other six assets. As described in Section II, the estimated ex ante real rates are the fitted values from regressions of the ex post real rate on information available at the beginning of the period. Our regressors were a constant term, the inflation rate over the past three months, the current nominal return on three-month U.S. Treasury bills and a fourth order polynomial in time that proxies for economic variables left out of the specification.

The comparison of the two measures is quite illuminating. Both measures show that ex ante real interest rates were higher in the 1960s than in the 1970s. However, from 1964 onward, estimated real rates using CPIX are almost always above those using CPI, especially so during 1968-71, 1973-75, and 1977-79. These are all periods of rising inflation and nominal interest rates, periods when the CPI is most likely to overstate inflation.

Although both estimated ex ante real returns turn negative in mid 1972, those based on CPIX are significantly negative only during the short period from late 1975 to 1977, while those using the CPI are continuously significantly negative from 1974 onward. This is easily seen because, as an approximation, only rates that lie outside the range ±1% are significantly different from zero at the 5% level of significance.[14] Previous findings of persistently negative ex ante real interest rates have apparently been spurious, a result of the CPI's mismeasurement of inflation.

B. Stability of the Stochastic Process of Real Rates

Casual observation of nominal interest rates and inflation makes it obvious

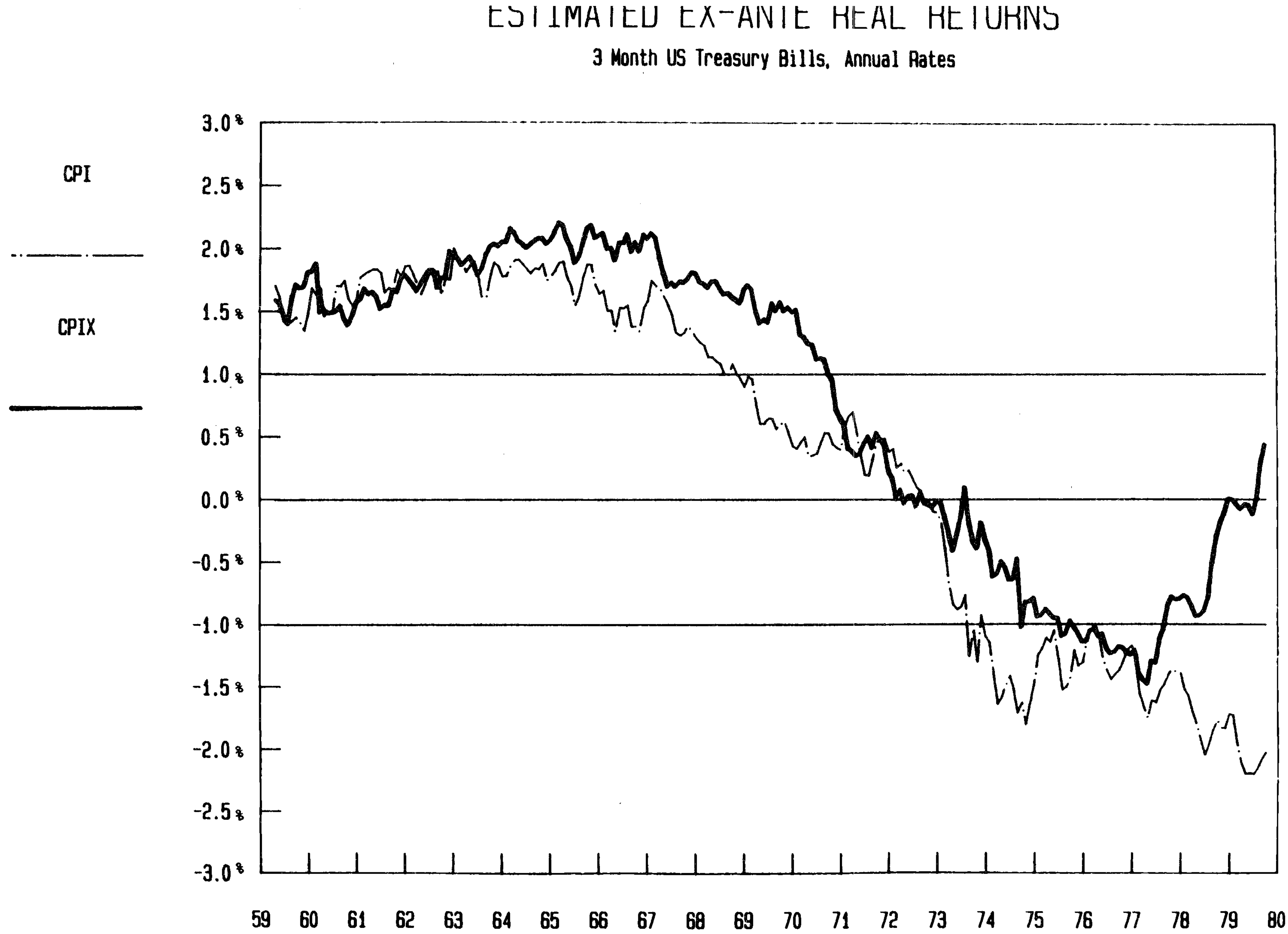
ESTIMATED EX-ANTE REAL RETURNS
3 Month US Treasury Bills, Annual Rates
CPI
CPIX
3.0%
2.5%
2.0%
1.5%
1.0%
0.5%
0.0%
-0.5%
-1.0%
-1.5%
-2.0%
-2.5%
-3.0%
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FIGURE 1

that real rates have undergone a major shift in the 1980s. We look at this phenomenon with a standard test for the stability of the coefficients in the CPIX regressions of Table 1, testing for a stable relationship of ex ante real rates with inflation and with nominal rates. We use 1959:5 to 1979:10 as one time period and 1979:11 to 1981:10 as the other. The breakpoint was chosen to coincide with the shift in the Federal Reserve policy regime towards a monetary aggregates target and away from interest rate targets. The choice of this breakpoint is somewhat arbitrary and we are by no means sure that the shift in the stochastic process for real rates occurred at this time. We intend to pursue a statistically based approach to the dating of the breakpoint in future research. Nevertheless, the advent of the new Federal Reserve policy has seemed to coincide with major changes in the behavior of U.S. financial markets.

Table 2 contains the χ^2 values for testing coefficient stability and the coefficients estimated for the 1979:11 to 1981:10 period. Because there are only twenty four highly autocorrelated observations for this latter period, whether using asymptotic distribution theory to test coefficient stability is appropriate is a serious question. A set of Monte Carlo simulations[15] indicates that the small sample distribution of the test statistics is quite different from the asymptotic. For example, in the case of using CPIX in the nominal bill rate regression for TBILL3, the real probability of rejecting the null hypothesis of coefficient stability when it is true, using a 5% significance level and the asymptotically valid χ^2 distribution, is 26.1%. Using the asymptotic distribution can, therefore, lead to rejecting the null hypothesis far too often. For this reason, the significance levels calculated by the Monte Carlo simulation are used for both the tests of coefficient stability in Panel A and tests of coefficients equal to zero in Panel B.

The results of Panel A indicate that for the nominal bill rate regressions, the equality of coefficients in the pre- and post-October 1979 periods is strongly rejected for the short-term assets. However, there are no rejections for any assets with the inflation coefficients. Although we do not report them here, we also performed stability tests where the 1959:5 to 1979:10 period was split in half. These tests rarely reveal a rejection of coefficient stability; the shift in the stochastic process of real rates after 1979:10 is more noticeable than any shift we have detected before this date.

A deeper understanding of why stability of the nominal bill rate coefficients is rejected is provided by Panel B of Table 2. The coefficients for the latter period tell a very different story than than those for the earlier one. After 1979:10 an increase in the nominal interest rate on three month U.S. Treasury bills is associated with a rise in the ex ante real return on all assets except common stock. In contrast to the pre 1979:10 period, when the nominal and ex ante real rates on three month U.S. Treasury bills were negatively correlated, in the latter period an increase in the nominal rate relects a one-for-one increase in the ex ante real rate. That is, the nominal bill rate coefficient in the TBILL3 regression is not statistically different from one using CPIX. Furthermore, with the exception of stocks, the magnification effect found in the earlier period is still present; but it now indicates that high short-term nominal rates are associated with higher real rates on longer maturity assets relative to those on short-maturity assets. The inflation coefficients in Panel B also undergo a large shift from their pre-October 1979 levels. However, this shift is not statistically significant because, despite some large coefficients, the relationship between real rates and inflation is not statistically significant in the latter period.

Taken together, the combined results of Tables 1 and 2 can be characterized by the following relationship among ex ante real rates, expected inflation, and the nominal interest rate on three-month U.S. Treasury bills. In the pre-October 1979 period, the ex ante real return on all seven assets studied here is negatively related to expected inflation, and given the positive autocorrelatation of inflation, negatively related to actual inflation. Further, although the ex ante

TABLE 2

Stability Tests and Coefficient Estimates for eprr = $\alpha + \beta X$
Using the CPIX Price Index

	PANEL A 1959:5 to 1979:10 vs. 1979:11 to 1981:10 Stability Test for β		PANEL B Estimates of β for 1979:11 to 1981:10	
X	INFLATION	NOMINAL BILL RATE	INFLATION	NOMINAL BILL RATE
ASSET				
TBILL3	0.67	40.91**	.014	1.160**
TBILL6	1.69	33.19*	.721	1.898**
TBILL12	3.43	23.09*	2.452	3.181**
INTBOND	5.97	13.13	7.649	5.926
LONGBOND	10.99	11.91	12.711	7.984
CORPBOND	10.72	12.37	12.217	7.166
STOCK	11.37	2.11	6.990	-6.810*

Notes: The stability tests of Panel A are standard Wald tests. Under the null hypothesis of coefficient stability, the statistics are asymptotically distributed as a χ^2 random variable with one degree of freedom. The α's were not constrained to be equal in the two periods.

A * in Panel A denotes rejection of the null hypothesis of coefficient stability at the 5% level of significance, ** represents rejection at the 1% level. In Panel B * denotes a coefficient which is significantly different from zero at the 5% level of significance and ** significantly different at the 1% level. Significance values were determined using the Monte Carlo simulation described in footnote 15.

real rate on three-month U.S. Treasury bills is not constant, their movement is sufficiently small, as is demonstrated by Fama and Gibbons (1982) and Huizinga and Mishkin (1983), so that nominal three month bill rate movements mainly reflect changes in expected inflation. This, when combined with the negative correlation between real rates and expected inflation, accounts for the negative correlation between the nominal bill rate and the seven ex ante real rates displayed in Table 1.

In the post-October 1979 period, movements in the ex ante real rates appear to have been highly variable and uncorrelated with expected inflation, thereby breaking the link between ex ante real rates and actual inflation. Because of the high variablilty in the ex ante real rate on three month U.S. Treasury bills, movements of their nominal rate no longer primarily reflect changes in expected inflation but rather movement of the real rate. The result for CPIX and TBILL3 presented in Panel B of Table 2 emphasizes this result, because it implies that in the post-October 1979 period a regression of inflation on a constant and the nominal three-month bill rate yields a coefficient on the bill rate of -.16. This is significantly different from one, the value we should get if the nominal rate is primarily reflecting expected inflation. The positive correlation of the nominal bill rate and the ex ante real rates on all assets except stock displayed in Panel B of Table 2 then means that the real rates on bills and bonds have been positively correlated in this latter period. The negative relationship between the nominal rate on three-month bills and the ex ante real return on common stock indicates a divergence of the real rate on stock from that on other assets. This result is certainly one that merits further investigation.

V. CONCLUSIONS

This paper discusses a methodology for measuring ex ante real interest rates on assets with different risk characteristics and for examining their correlation with inflation and nominal interest rates. It is applied to the 1959-1981 time period for the following seven assets: (1) three-month U.S. Treasury bills, (2) six-month U.S.

Treasury bills, (3) twelve-month U.S. Treasury bills, (4) intermediate-term (5-10 years maturity) U.S. Treasury bonds, (5) long-term (10-20 years maturity) U.S. Treasury bonds, (6) long-term corporate bonds, and (7) common stock. The importance of choosing a price index for constructing inflation and real rates is also examined. The major findings are as follows:

1. Pre-October 1979, the negative association between ex ante real rates and both inflation and nominal interest rates, which was previously found for short-term U.S. Treasury bills, appears for all the assets in this study. Indeed, the negative association is even larger for the longer-maturity assets. This suggests that all seven assets have been poor hedges against expected inflation, and that the longer-maturity assets have been the worst hedges.
2. The results described above are robust with respect to the use of price indices other than the Consumer Price Index (CPI). The puzzling finding of previous research that ex ante real returns on short-term U.S. Treasury securities were persistently negative in the 1970s, however, appeals spurious and due to the CPI's mismeasurment of inflation. A more reliable estimate of ex ante real returns, one that uses a more accurate price index, finds statistically significant ex ante real rates for far shorter periods of time.
3. A statistically significant shift in the stochastic process of real rates does occur sometime around the October 1979 change in the Federal Reserve's policy regime. In contrast to the pre-October 1979 period where ex ante real rates and nominal rates are negatively correlated, the post-October 1979 period has real and nominal rates moving together. In particular, movements in the nominal interest rate on three-month U.S. Treasury bills no longer primarily indicate changes in expected inflation, but rather the movement of the ex ante real return. There also seems to be a change in the relationship between the ex ante real returns on stock and the ex ante real return on bills and bonds near the end of 1979.

Footnotes

1. See Litterman and Weiss (1982), Barro (1982), Fama and Gibbons (1982), Summers (1982), Mishkin (1981b,1982,1984) and Huizinga and Mishkin (1983). More detailed references on earlier work can be found in Mishkin (1981b).

2. See Blinder (1980)

3. The notion of a hedge against expected inflation was suggested in Fama and Schwert (1977). It is not a standard use of the term because the holder of the asset is not protected against an unanticipated event. Nevertheless, the meaning of hedge as used here should be clear.

4. As is indicated in Mishkin (1981a), survey data on expected inflation may not be a reliable guide to the bond market's rate of expected inflation.

5. Conditional heteroscedasticity in the error term is heteroscedasticity that can be explained by movements in the regressors. We test for this by regressing the squared residuals from our estimated equations on the squared regressors. Significant coefficients on any variable except the constant term is evidence of conditional heteroscedasticity. Space does not permit us to report the results of all these tests. However, for the regressions where the CPIX price index is used, the t-statistics on the squared values of inflation (in a regression of the squared residual on a constant and squared inflation) are TBILL3: 1.51, TBILL6: 1.71, TBILL12: 2.01, INTBOND: 2.56, LONGBOND: 2.23, CORPBOND: 3.36, and STOCK: 1.79. It should be noted that the method of correcting the estimated parameter standard errors used here does not require a parameterization of how the variance of the error is related to the regressors.

6. The rental equivalence measure, first announced in October 1981, became the official residential housing component for the CPI for All Urban Consumers in January 1983. The CPI for Urban Wage Earners and Clerical Workers will continue to use the old home-ownership method until January 1985. A detailed description of the new rental equivalence measure can be found in the CPI Detailed Report, January 1983 issue. The CPIX should be a better index than the CPI less home mortgage interest costs, a series often used to minimize the mismeasurment of prices by the CPI, because this latter index still includes a large interest rate component via its inclusion of the purchase price for new homes. We thank Robert Dennis of the Congressional Budget Office for providing us with the CPIX data back to 1959.

7. Our sample period starts in 1959:5 because the PCED price index is not available until 1959. The dating of the real returns follows the convention described in Section II, so that a time period 1959:5 to 1979:10 has the return from May 1, 1959 to July 31, 1959 as its first observation and the return from October 1, 1979 to December 31, 1979 as its last. The first observation for the inflation rate regressor is constructed using the January and April 1959 price indices. By using the April index to explain the real return from May to August, we are assuming that agents know the April price index on April 30. In fact, the April index will not be announced until sometime in May. However, replacing the January-April inflation rate with the December-March rate does not change our results appreciably. See Huberman and Schwert (1983) for evidence on this issue

in the Israeli bond market. The nominal three-month bill rate used as a regressor for the first observation is the rate on April 30, 1959. There is no seasonal adjustment of any data, except for PCED which we could not get in seasonally unadjusted form. We have constructed tests where all data except PCED are seasonally adjusted with Box-Jenkins (1976) seasonal models and the results are essentially the same.

8. In computing the standard errors reported in Table 1, the error terms of the regressions are assumed to have nonzero autocorrelations at lags one and two. This is based on examination of the sample residuals. For example, when using CPIX, the first twelve autocorrelations of the residuals from a regression of the ex post real rate for TBILL3 on a constant and the inflation rate are .68, .35, -.03, -.07, -.08, -.02, -.01, .04, .12, .22, .25 and .25. The non-zero autocorrelations at lags one and two are exactly what one should expect given the quarterly real rates and monthly data. Since lagged real rates do not appear as regressors the u_t component of the error term, and hence the entire error term, can display autocorrelation at any lag without violating our assumptions of rational expectations. The rise in the autocorrelations near lag twelve can be interpreted as a mild seasonal fluctuation in the ex ante real rate that is not captured by the inflation rate. Sample residuals from other equations display essentially the same pattern of autocorrelations, though for the bonds and stock equations there is no evidence of a seasonal. This is not surprising since the larger variance of the forecast errors $\varepsilon 1_{t,j}$ and $\varepsilon 2_{t,j}$ in the equations for the longer maturity assets should make a seasonal harder to find.

9. When the coefficients in Table 1 differ significantly from equation to equation, it is evidence of a risk premium in the returns. This issue is formally addressed in Huizinga and Mishkin (1983).

10. The test statistic at the bottom of each column is distributed as a χ^2 random variable with seven degrees of freedom under the null hypothesis that all seven coefficients in the column are zero. Its marginal significance level is the probability, when the null hypothsis is true, of getting a test statistic as high as, or higher than, the one observed. Thus a marginal significance level smaller than .01 indicates rejection of the null hypothesis at the 1% level.

11. When we test the null hypothesis that there is no relationship between inflation and the ex ante real return on the four long-term assets -- i.e. jointly test that the coefficient of inflation is zero for the INTBOND, LONGBOND, CORPBOND, and STOCK equations -- we get χ^2 values of 13.90 for CPI, 14.68 for CPIX, 11.68 for PCED, and 18.74 for PPI. This indicates rejection of the null hypothesis at the 1% level for CPI, CPIX, and PPI and rejection at the 5% level for PCED. Thus our tests do have some power to reject the hypothesis that inflation is uncorrelated with the ex ante real return for long-term assets.

12. When we included both the yields to maturity and the three month bill rates as regressors, the coefficients for the yield variables became insignificantly different from zero. Thus the term structure of nominal interest rates apparently provides no additional information about short-term ex ante real returns above and beyond what is captured by the nominal three-month bill rate.

13. See Fama(1975), Mishkin (1981b), Fama and Gibbons (1982), and Huizinga and Mishkin (1983).

14. The standard errors of the estimated real rates are calculated under the assumption that unexpected inflation is the major component of the error term. These standard errors do show some variation over time but do not differ greatly from .5%. The calculations used are described in Mishkin (1981a).

15. To do the Monte Carlo simulations, the inflation rate, nominal interest rate on three month U.S. Treasury bills and residuals from each of the fourteen CPIX regressions reported in Table 1 were fit to a univariate time series process. Using the estimated time series representation, including the estimated correlation of the innovations in the series, two thousand and replications of each of the fourteen regressions were run. A replication consisted of generating a new regressor, generating a new residual, generating a new dependent variable from the regressor, the residual and the coefficients reported in Table 1 and finally, estimating the equation.

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