Monetary Policy and Short-Term Real Rates of Interest

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TEXTBOOK descriptions of the channels of monetary policy's impact on the economy usually outline a two-step procedure: "The first is that an increase in real balances generates a portfolio disequilibrium — at the prevailing interest rate and level of income, people are holding more money than they want. This causes portfolio holders to attempt to reduce their money holdings by buying other assets, thereby changing asset yields. In other words, the change in the [real] money supply changes [real] interest rates. The second stage of the transmission process occurs when the change in interest rates affects aggregate demand."1

The rational expectations literature, however, has raised serious questions about this description, especially the first stage wherein an increase in real money balances lowers expected real interest rates. Shiller, for example, drawing from previous work in rational expectations, hypothesizes that the expected real interest rate is unaffected by changes in monetary policy.2 While Shiller found little support for this hypothesis, other recent empirical work supports it. Fama, for instance, is unable to reject the hypothesis that the expected real rate on short-term financial assets was constant over much of the post-Accord period in the United States.3 This hypothesis is even stronger than Shiller's. It holds that monetary actions, as well as everything else, have had no systematic effect on expected real interest rates.

This article re-evaluates the evidence suggesting that the expected (ex ante) real interest rate on short-term financial assets is constant. Evidence is provided that allows us to reject this hypothesis for the 1955-79 period. Following this, data are examined to determine whether evidence supports the typical textbook description in which changes in expected real interest rates are associated with changes in real money growth.

THE FRAMEWORK OF ANALYSIS

Consider first the relationship between nominal interest rates and inflation expectations embodied within the so-called Fisher relationship,4

\[ \text{Fisher equation} \]

\[ (1) \quad i_t = r_f + \pi_t^e, \]

where \( i_t \) is a nominal (or market) rate of interest (the rate measuring how many dollars must be repaid in the future for a given dollar loaned today), \( r_f \) is the expected real interest rate (the rate measuring how

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2Robert J. Shiller, "Can the Fed Control Real Interest Rates?" in Stanley Fischer, ed., *Rational Expectations and Economic Policy* (The University of Chicago Press, 1989), pp. 117-56. Shiller also outlined two other (non-exclusive) hypotheses: (1) the Fed can affect real rates only through unexpected policy moves and (2) Fed policies known far enough ahead of time have no effect on real rates. These hypotheses are not as stringent as the hypothesis considered in this paper.

This equilibrium relationship also should include the cross-product term $\tilde{i} P_t$. Like most empirical analyses of this relationship, we ignore this term, assuming that the magnitude of the variable is sufficiently small.

**IS THE EXPECTED REAL RATE OF INTEREST CONSTANT?**

To test the relationship specified by equation 1, one can make two assumptions: First, assume that the expected real interest rate is a constant, such that

$$(2) \tilde{r}^e = \tilde{r}.$$  

Second, to circumvent the problem of measuring inflation expectations, assume that next period's actual inflation ($\tilde{P}_{t+1}$) is equal to what is currently expected (at time $t$), plus a random disturbance $\mu_{t+1}$, where $\mu_{t+1}$ is independent and distributed N(0, $\sigma^2$):

$$(3) \tilde{P}_{t+1} = \tilde{P}_t + \mu_{t+1}.$$  

This relationship specifies that one-period-ahead inflation forecasts are unbiased, on average the actual inflation rate over the next time period will be the expected rate.

Substituting equations 2 and 3 into 1 yields

$$(4) \tilde{i} = \tilde{r} + \tilde{P}_{t+1} - \mu_{t+1}.$$  

This equation can be arranged to test empirically the hypothesis that today's interest rate accurately predicts tomorrow's inflation as follows:

$$(5) \tilde{P}_{t+1} = \tilde{r} + \beta_0 \tilde{i} + \mu_{t+1}.$$  

Assuming that financial markets are efficient, we would expect to find $\beta_0$ not to be statistically different from unity and the estimated constant term to be negative. If the estimated coefficient $\beta_0$ is not statistically different from unity, the proposition that current interest rates fully reflect the market's anticipations of the future inflation rate cannot be rejected. Similarly, if the estimated constant term is negative, the expected real rate of return is then positive as suggested by the underlying economic theory. More-
Table 1
Empirical Estimates of Equation 5

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<td>0.439</td>
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<td>1.63</td>
<td>1.92</td>
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<td>(\rho)</td>
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\(R^2\) represents the coefficient of determination adjusted for degrees of freedom. SE is the regression standard error. DW is the Durbin-Watson test statistic and \(\rho\) is the estimate of the autocorrelation coefficient. Absolute value of t-statistics appear in parentheses.

Table 1 presents estimates of equation 5 for various periods. The inflation data used to estimate equation 5 are based on quarterly observations of the GNP deflator, expressed as annual rates of change. Since the GNP deflator provides an average measure of prices over the quarter, the quarterly average three-month Treasury bill rate is used as the nominal interest rate measure.

Consider first the results obtained by estimating equation 5 over the full sample period, I/1955-IV/1979. The constant term is negative (although not significantly different from zero), and the coefficient on the interest rate variable is not statistically different from unity as suggested by the theory. Unfortunately, the low Durbin-Watson statistic provides evidence of first-order serial correlation.

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The GNP deflator is used to avoid recent problems with the consumer price index. For a discussion of problems with this index, see Alan S. Blinder, "The Consumer Price Index and the Measurement of Recent Inflation," Brookings Papers on Economic Activity (2,1980), pp. 539-85.

Eugene F. Fama and Michael R. Gibbons, "Inflation, Real Returns and Capital Investment," Working Paper No. 41 (Graduate School of Business, University of Chicago, 1980), also find evidence of serially correlated disturbance terms when quarterly data are employed. In addition, in that study as well as in his "Stock Returns, Real Activity, Inflation, and Money," American Economic Review (September 1981), pp. 545-85, Fama drops the assumption that the expected real rate of interest is constant. Both studies estimate the inflation/interest rate relationship assuming that the expected real rate is a random walk.
This result, by itself, is enough to reject the framework in equation 5. Focusing solely on the constancy of the expected real rate, however, the accompanying estimation problem can be corrected by using generalized least-squares estimation. These results appear in the lower half of Table 1.

The full sample results reported there again indicate that next period’s rate of inflation does mirror, one-for-one, a rise in today’s interest rates. Moreover, the constant term remains insignificantly different from zero.

Table 1 further reports estimation results for subperiods arbitrarily truncated at the end of each decade. If the expected real rate of interest is temporally invariant, the constant terms in these subperiods should not differ statistically. Yet, as the table immediately shows, they do differ significantly across the various subperiods shown. In fact, the estimated constant term is positive and significant in the first subperiod (late 1950s), while not different from zero in the last decade (1970s). It has the anticipated negative sign only in the decade of the 1960s. Moreover, the coefficient on the interest rate variable is not statistically different from zero in the late 1950s, even though theory suggests that it should equal unity. Thus, the coefficient estimates, as well as summary statistics such as the R² and the standard errors of the equation, vary substantially across subperiods, irrespective of the estimation technique used.

The statistical significance of the variation in the constant term (the estimate of the ex ante real interest rate) can be investigated by including dummy variables for possible shifts in the intercept. Thus, equation 5 was re-estimated with two dummy variables: D1 equal to 1 for I/1955-IV/1959 and D2 equal to 1 for I/1960-IV/1969. Estimating such an equation with ordinary least squares again yielded residuals that were significantly autocorrelated. To improve hypothesis testing, the equation was estimated using a generalized least-squares routine to correct for assumed first-order autocorrelation. The I/1955-IV/1979 estimation results are (absolute value of t-statistics in parentheses):

\[
\hat{\beta}_{t+1} = 1.40 - 0.52 D1 - 1.88 D2 + 0.83 i_t
\]

\[
R^2 = 0.55 \quad SE = 1.37 \quad DW = 2.07 \quad \hat{\rho} = 0.35
\]

These results support the previous subperiod findings: the estimated real interest rate is significantly positive only in the 1960s. The point estimates of the expected real interest rate for the 1950s, 1960s and 1970s, respectively, are \(-0.52, +0.48\) and \(-1.40\). While the point estimates for the 1950s and the 1970s are negative, they are not significantly different from zero. On the other hand, the positive point estimate for the 1960s is significantly different from zero. Thus, the hypothesis that the expected real interest rate has been constant over the past 25 years must be rejected.

**EX POST REAL RATES: FURTHER CONSTANCY TESTS**

Equation 4 can be rewritten as

\[
h_t - \hat{r}_{t+1} = \hat{r} - \mu_{t+1}
\]

This equation states that the ex post real rate should equal a constant (the ex ante real rate), minus a white noise random error term. A feel for the statistical variation in the real rate can be obtained by plotting its behavior for our sample period. Chart 1 shows the quarterly ex post real rate for the I/1955-IV/1979 period and its mean values for the I/1955-IV/1959 (−0.03), I/1960-IV/1969 (1.21) and I/1970-IV/1979 (−0.39) subperiods. If equation 6 holds for the whole period, the means across subperiods should be equal, since the expected value of the disturbance term in each subperiod is zero.

Tests for equality of the ex post real interest rate means across the subperiods provide another investigation of the constancy hypothesis. Such tests again lead to a rejection of this hypothesis. The t-statistic,

\[
(6) \quad h_t - \hat{r}_{t+1} = \hat{r} - \mu_{t+1}
\]

These results are somewhat different from that used by others. Many take the difference between today’s interest rates and today’s inflation rate as an ex post real rate measure. Theory suggests, however, that the preferable measure is the difference between today’s interest rates and tomorrow’s inflation.

In the test subsequently developed and others which follow, interest rates are assumed to adjust one-for-one with inflation expectations, a hypothesis that can be rejected in equation 5. The reader should be cautioned that there are counter-theoretical arguments and some empirical evidence to suggest that the nature of the U.S. tax system has invalidated this relationship, with interest rates rising more than one-for-one with an increase in inflation expectations. For theoretical discussions, see Michael R. Darby, “The Financial and Tax Effects of Monetary Policy on Interest Rates,” *Economic Inquiry* (June 1975), pp. 296-76; and Martin Feldstein, “Inflation, Income Taxes, and the Rate of Interest: A Theoretical Analysis,” *American Economic Review* (December 1976), pp. 800-20. For empirical evidence on the matter, see John A. Carlson, “Expected Inflation and Interest Rates,” *Economic Inquiry* (October 1979), pp. 597-608.
used to test whether the mean ex post real rate for the latter half of the 1950s is equal to that of the 1960s, is 3.67, sufficiently large to reject the null hypothesis at the 5 percent significance level. Further, the t-statistic used to test the equality of mean ex post real rates in the 1960s relative to the 1970s is 4.86, again allowing rejection of the null hypothesis of constant real interest rates at the 5 percent level. Thus, if one accepts the propositions that interest rates move in direct proportion with expected inflation and that inflation expectations are unbiased, one must reject the constancy of the ex post real interest rate over the subperiods investigated.

**MONETARY POLICY AND THE EXPECTED REAL RATE**

These findings suggest that the real interest rate has not been constant over the past 25 years. In this light, is there any evidence that links the real rate of interest to monetary policy? After all, the textbook description of monetary policy's transmission mechanism relates changes in the real rate to changes in real money balances. In particular, it maintains that an increase in real money balances lowers expected real rates, at least temporarily.

The previous framework, linking ex post and ex ante real rates, can be used to address this issue. If inflation expectations are unbiased and financial markets are efficient, then the ex post real rate \((i_t - \hat{P}_{t+1})\) is equal to the ex ante real rate \((r^e_t)\), minus a random disturbance term \(\mu_{t+1}\) capturing unexpected inflation:

\[
(7) \quad i_t - \hat{P}_{t+1} = r^e_t - \mu_{t+1}.
\]

The typical textbook relationship can be represented as

\[
(8) \quad r^e_t = \beta_0 + \beta_1 (M_t/P_t) + \beta_2 (M_{t-1}/P_{t-1}) + \ldots + \varepsilon_t,
\]

where \(M\) is the nominal money stock, \(P\) is the price.
level and $\epsilon$ is a random error term. This relationship represents the hypothesis that the expected real rate is related to real money balances. Since nothing in macroeconomic theory indicates how long it takes for changes in monetary policy to have an effect, lagged real balances are included in an effort to capture empirically the dynamics of the process. Theory does suggest, however, that some of the coefficients should be significantly negative. While it is impossible to estimate equation 8 because of a lack of observations on $r_{n}$, equation 7 indicates that it is impossible to estimate equation 8 because of a lagged real balances. Trying 10 lags on real money balances in the relationship implies a more stable pattern of real interest rates. The coefficient estimate for real money balances lagged one period is significantly positive and is not statistically different from the absolute value of the coefficient on contemporaneous real money balances. This finding indicates that a current increase in real money balances will be associated with a current decline in real rates, but followed by a rise in real rates of equal size at time $t+1$. This suggests that monetary authorities, to the extent that they can change real balances, cannot permanently affect real rates of interest.

While earlier evidence showed that the ex post real rate ($i_t - \bar{P}_{t+1}$) behaved differently across subperiods, there is little evidence to suggest that its relationship to real money balances has changed over the period. For example, a conventional Chow test evaluating a hypothesized break in the relationship at IV/1969 yields a calculated $F$-statistic of $F(3.94) = 0.39$, well below the 5 percent critical value of 2.70. Thus, the regression coefficients are not statistically different before or after IV/1969.\(^{14}\) Changes in real balances have the same statistical effect on real interest rates across the sample period.

Finally, it is appropriate to note that the estimated relationship implies a positive relationship between the volatility in real money balances and the volatility in real interest rates. If the frequency of change in real money balances increases, the estimated relationship implies an increase in the frequency of change in real interest rates. The evidence presented here suggests that more stable real money growth, even over periods as short as a quarter, will produce a more stable pattern of real interest rate movements.\(^{15}\)

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\(^{12}\)We do not mean to suggest that monetary authorities can control real money balances over long periods of time. On this point, see Denis S. Karnowsky, "Real Money Balances: A Misleading Indicator of Monetary Actions," this Review (February 1974), pp. 2-10.

\(^{13}\)Money ($M/P$) is measured (in billions of 1972 dollars) by the adjusted monetary base for all results reported here. Thus, the empirical results indicate that a $1 billion increase in real balances will reduce the real interest rate by 89 basis points in the current period. This decline is offset, however, by an 83 basis-point increase in the real rate in the subsequent period. We also tried the M1 measure and obtained similar results.

\(^{14}\)In addition, we tested the hypothesis that the variance of the error term was larger in the 1970s than in the earlier period. The calculated $F$-statistic (with 37 and 57 degrees of freedom, respectively) was 1.44, less that the 5 percent critical value of 1.59. Thus, the hypothesis of equal variance across these two periods cannot be rejected.

\(^{15}\)An interesting investigation into the effects of monetary policy on both short- and long-term real interest rates is provided in Dean W. Hughes and Diane Weimer, "The Impact on Business Investment of the Federal Reserve System’s Operating Procedures," Federal Reserve Bank of Kansas City Economic Review (February 1982), pp. 14-25.
CONCLUSION

This article has provided evidence counter to the hypothesis that the expected real rate of return on short-term financial assets was constant over the period 1955-79. If such a hypothesis were valid, monetary policy would be powerless in affecting real economic activity through the conventional transmission mechanism. While rejecting the constancy hypothesis, this article also provides evidence consistent with conventional macroeconomic theory whereby increases in real money balances temporarily lower expected real rates. This effect is contemporaneous on a quarterly basis. While such an effect is significant, it is relatively small and is offset in the following quarter by an identical rise in expected real rates. Thus, there is no evidence of a long-run effect running from changes in real money balances to changes in real interest rates. Finally, the evidence presented here suggests that more volatile short-run real money growth is likely to produce more volatile real interest rate fluctuations. Thus, contrary to recent claims, stable money growth and stable interest rates are hardly inconsistent policy objectives.\textsuperscript{16}

\textsuperscript{16}For another view, see Bryon Higgins, “Should the Federal Reserve Fine Tune Monetary Growth?” Federal Reserve Bank of Kansas City Economic Review (January 1982), pp. 3-16.