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MONETARY POLICY REGIME SHIFTS AND THE UNUSUAL BEHAVIOR OF REAL INTEREST RATES

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ABSTRACT

A striking phenomenon of the early 1980s is the climb in real interest rates to levels unprecedented in the post-World War II period. In order to understand this phenomenon, this paper investigates the nature and timing of shifts in the real rate process to determine if the recent unusual behavior of real rates is associated with monetary policy regime changes. We find that not only are there significant shifts in the stochastic process of real interest rates in October 1979 and October 1982 when the Federal Reserve alters its behavior, but these dates are also found to be the most likely breakpoints in the real rate process. When we analyze another monetary policy regime change with many similarities to that of October 1979, the sharp rises in the discount rate in 1920, we also reach a similar conclusion: there is a striking correspondence between the monetary policy regime change and the shift in the real rate process.

Other studies have examined competing explanations for the recent unusual behavior of real interest rates -- e.g., budget deficits or favorable changes in business taxation. Although these competing explanations have met with mixed success, our evidence lends substantial support to the view that monetary policy regime changes have been and continue to be an important source of shifts in the behavior of real interest rates.

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I. Introduction

A striking phenomenon of the early 1980s is the climb in real interest rates to levels unprecedented in the post-World War II period. In order to understand this phenomenon, it is not enough to focus on the recent high level of real rates, as in several recent papers;[1] research has also revealed a significant shift in the stochastic process of real interest rates.[2] Modern monetary theory suggests that regime changes have an important impact on the stochastic process of many economic variables. Thus, a change in a policy regime (by which we mean a change in the direction of policy or the way in which policy is conducted) may explain the unusual real rate experience in recent years.

In this paper, we investigate the nature and timing of shifts in the real rate process to determine if the unusual behavior of real rates is associated with monetary policy regime changes. We find that when the Federal Reserve alters its behavior in October 1979 and October 1982, there are statistically significant shifts in the stochastic process of real rates. Statistical analysis designed to determine the timing of the shifts in the real rate process indicates that these dates are also the most likely choices for breakpoints.

The above evidence suggests that the Fed's changes in the monetary policy regime are a likely candidate for explaining the recent unusual behavior of real interest rates. Our search for an understanding of how these changes might have altered real rate behavior proceeds along two

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[1] See for example, Blanchard and Summers (1984) and Holland (1984).

[2] See Huizinga and Mishkin (1984).

dimensions. First, since the Fed's 1979 changes in operating procedures were associated with increased variability of interest rates, money supply growth and inflation, we explore whether increased uncertainty about these variables plays a significant role in real rate movements.[3] We find little evidence supporting an important role for uncertainty in the recent unusual behavior of real rates.

Second, we look for other periods in which there is a clearly defined change in the monetary policy regime that bears similarities to the recent regime changes.[4] In a sense, we are looking for other "controlled experiments" that provide information on how monetary policy regime changes affect real interest rates. One such "experiment" is in 1920, when the Federal Reserve sharply raised its discount rate twice. This episode is a natural one to study because the economy was suffering from a high and persistent inflation before the regime shift, while afterwards a sharp disinflation occurred. This is similar to what we have seen in recent years and thus we might expect to find parallels between the two periods.

Our analysis of the period surrounding 1920 reveals a significant shift in the stochastic process of real interest rates which has many similarities to the recent experience. For example, the 1920 monetary regime change and the subsequent disinflation is associated with a

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[3] Mascaro and Meltzer (1983) have argued that increased monetary uncertainty is an important factor in the recent rise in real rates. Fama (1976) has found a link between real interest rates, inflation uncertainty, and nominal interest rate uncertainty. Hartman and Makin (1982) also find a link between inflation uncertainty and interest rates.

[4] Shiller (1980) has also examined historical episodes to see if a change in the monetary policy regime affects real interest rates. For the end of Fed-Treasury Accord he finds a noticeable change in the behavior of real interest rates.

weakening of the Fisher effect and a shift to a sustained higher level of real interest rates, which are also characteristic of the recent period. Although our analysis does not reveal the precise mechanism through which changes in monetary policy regimes affect real interest rates, it does suggest that disinflation is a more important factor in the recent policy regime shift than is the increased variability of money growth, inflation and interest rates.

Other studies have examined competing explanations for the recent unusual behavior of real interest rates -- e.g., budget deficits, investment booms, favorable changes in business taxation, and the declining relative price of energy. Although these competing explanations have met with mixed success, our evidence lends substantial support to the view that monetary policy regime changes have been and continue to be an important source of shifts in the behavior of real interest rates. Future research should focus on how changes in monetary policy regimes produce shifts in real interest rate movements.

II. The Methodology

The real interest rate of concern to economists is more precisely referred to as the ex ante real interest rate. At time t, this ex ante rate is defined as the expectation at time t of the ex post real return on an asset when it is held from time t to t+1. In the case of a one-period, default-free, nominal bond, whose expected nominal return is equal to its nominal interest rate at time t, the ex ante real interest rate is:

(1) $$rr_t \equiv i_t - \pi_t^e$$

where,

rr_t = the ex ante real interest rate on the one-period bond at time t: i.e., the ex ante real return from time t to t+1.

i_t = the nominal interest rate on the one-period bond at time t: i.e., the nominal return from

time

time t to t+1.

π_t^e = the inflation rate from time t to t+1 expected at time t.

The ex post real rate is defined as,

(2) $$eprr_t = i_t - \pi_t = rr_t - \varepsilon_t$$

where,

$eprr_t$ = the ex post real interest rate on the one-period bond at time t: i.e., the realized real return from time t to t+1.

π_t = the actual inflation rate from time t to t+1.

ε_t = the inflation forecast error, $\pi_t - \pi_t^e$.

Although the ex post real rate is observable, the ex ante rate is not. Thus, in order to measure and make inferences about the ex ante real rate, some identifying assumptions are needed. In this paper we use the identifying assumptions that expectations of inflation are rational and that the ex ante real rate can be reasonably approximated by linear projection onto an observable information set. The assumption of rational expectations implies that inflation forecast errors are unforecastable given any information available at time t: i.e.,

(3) $$E(\epsilon_t \mid \phi_t) = 0$$

where,

ϕ_t = represents all the information available at time t.

$E(\ldots \mid \phi_t)$ = the mathematical expectations operator conditional on the information set ϕ_t.

Describing the ex ante real interest rate as a linear projection onto an observable information set X_t gives,

(4) $$rr_t = X_t\beta + u_t$$

where,

X_t = a subset of ϕ_t.

u_t = the error from projecting rr_t on X_t, which by construction is uncorrelated with X_t.

Combining equations (1) through (4) generates the following ex post real rate regression equation,

$$\text{(5)} \qquad e\dot{prr}_t = X_t\beta + u_t - \varepsilon_t$$

As shown in Mishkin (1981), ordinary least squares (OLS) estimation of this regression equation yields consistent estimates of the parameters in β. This conclusion stems from the orthogonality of the composite error term $u_t - \varepsilon_t$ with X_t.[5]

Within the above framework, a logical way to test for a shift in the stochastic process of the ex ante real rate is to test for the constancy of β when equation (5) is estimated over different samples. We do this with a conventional F-test (Chow test). We date when the real rate process shifts using a procedure proposed by Quandt (1958, 1960). This procedure involves finding the breakpoints that maximize the likelihood of the ex post real rate regression. Specifically, for each set of potential breakpoints, we calculate minus twice the log of the Quandt likelihood ratio -- the maximized likelihood assuming there are no breaks, divided by the maximized likelihood conditional on that set of breakpoints. When this statistic reaches its maximum, we have found the set of breakpoints that maximizes the likelihood function.

Are the procedures testing for and dating of breakpoints meaningful? How are they to be interpreted if there is information Z_t in ϕ_t that is relevant for predicting the ex ante real rate but is unknown to the econometrician and has thus been excluded from X_t? To examine these questions, consider the hypothetical case in which the ex ante real rate can be expressed as

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[5] Note that there has been a change in the timing of the notation here from Mishkin (1981, 1984a,b). What was called rr_t, u_t and ε_t in Mishkin (1981,1984a,b) corresponds to rr_{t-1}, u_{t-1} and ε_{t-1} in the notation here.

$$(6) \qquad rr_t = X_t\beta^* + Z_t\delta$$

This implies that the coefficient β in equation (4) and (5) can be expressed as,

$$(7) \qquad \beta = \beta^* + \lambda\delta$$

where,

λ = the regression coefficient of X_t in a regression of Z_t on X_t.

A change in β represents either (1) a change in the way information is used to predict ex ante real rates (a change in β^* or δ), or (2) a change in the relation between the variables in the information set (a change in λ). Either of these changes can cause a change in the stochastic process of real rates.

We are interested not only in verifying a break in the stochastic process of real interest rates, but also in examining measures of the ex ante real rate to see how the stochastic process changes. Given the consistency of $\hat{\beta}$, it is logical to estimate the ex ante real rate as the fitted values obtained from OLS estimation of equation (5): i.e.,

$$(8) \qquad \hat{rr}_t = X_t\hat{\beta}$$

where,

$\hat{\beta}$ = the OLS estimate of β.

How good is this estimate? In a large sample the estimated ex ante real rate defined in equation (8) will differ from the true ex ante real rate by u_t, so that when u_t has small variance, $\hat{rr}_t$ is an accurate measure.

While it would be desirable to know that the variance of u_t is small, unfortunately its variance is not identifiable. However, there may be information in the data concerning the variance of u_t. Our information is limited to the variances and autocovariances of the regression residual from the ex post real rate regression, $u_t - \varepsilon_t$. Under rational expectations, ε_t is serially uncorrelated, yet u_t may be autocorrelated or correlated with past values of ε_t. If this is the case and the variance of u_t is large, then the regression residual, $u_t - \varepsilon_t$, is serially correlated. However, if the variance of u_t is small, then the regression residual $u_t - \varepsilon_t$ is dominated by ε_t and is serially uncorrelated. Thus a diagnostic check of the X_t specification involves examining the residuals from the ex post real rate regressions to see if they are white noise.[6]

The last remaining issue is how we choose the information set X_t. As described above, the choice of the variables in X_t comprise part of the identifying assumptions for measuring real interest rates. A variety of previous research uses the identifying assumption that ex ante real rates follow a random walk:[7] i.e., $X_t = rr_{t-1}$ and the coefficient $\beta = 1$. Evidence in Litterman and Weiss (1985) is inconsistent with this assumption: they find the ex ante real rate follows an AR(1) process with the

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[6] Although the diagnostic check described here is worth doing, it is not powerful against certain alternatives. If u_t is not serially correlated or correlated with past ε_t, then the regression residual $u_t - \varepsilon_t$ will not be serially correlated even if the variance of u_t is large. For example, if the X_t specification includes many lags of rr_t then u_t will not display any serial correlation and the diagnostic check will have little power.

[7] See, for example, Garbade and Wachtel (1978), Fama and Gibbons (1982), and Antoncic (1983).

AR coefficient significantly less than one.[8]

In light of this finding and the lack of a compelling theoretical justification for the random walk specification, we feel it appropriate to use a less restrictive set of identifying assumptions about the time series process for the ex ante real rate. This can be accomplished by including a distributed lag of ex post real rates and inflation rates in X_t.[9] This methodology also allows us to relax another assumption used by proponents of the random walk specification -- that unexpected inflation between time t-1 and t has no explanatory power for the ex ante real rate at t. Again, there is no compelling theoretical justification for this assumption and it is inconsistent with evidence in Litterman and Weiss (1985).[10]

By including other economic variables known at time t in X_t, we allow additional factors to help predict ex ante real interest rate movements, making our approach even more general. Specifically, we can include a measure of supply shocks which may affect ex ante real interest rates by affecting current and/or expected future investment

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[8] Evidence in Huizinga and Mishkin (1985) also suggests that the random walk assumption for the ex ante real rate is inappropriate.

[9] See Huizinga and Mishkin (1985).

[10] This assumption is not an assumption about monetary neutrality. Real factors can lead to unexpected inflation and have a lasting effect on the ex ante real interest rate. Litterman and Weiss (1985) estimate the correlation coefficient between unexpected inflation and the innovation in the ex ante real rate to be -.64.

[11] It is not clear a priori what effect a supply shock that raises the relative price of energy should have on ex ante real rates. The most common story, outlined by Wilcox (1983), has a rise in the relative price of energy decreasing investment opportunities, which lowers the demand for loanable funds and hence ex ante real rates. However, it is also possible that a rise in the relative price of energy might increase the returns to new capital even though it reduces the value of old capital. Then we might expect ex ante real rates to

opportunities.[11] We can also include the nominal interest rate at time t which may be able to capture other, hard to measure, influences on the ex ante real rate. The nominal rate is likely to reflect these other influences because it is composed of the ex ante real rate itself, as well as expected inflation, which a variety of evidence has shown to be negatively related to ex ante real rates.[12]

III. The Data

Two monthly data sets are used in this paper: one for the time period from January 1953 to December 1984 and the other for the time period from January 1916 to December 1927. We begin the postwar period in January 1953 because the quality of the CPI was substantially upgraded at this time. We stop in December 1984 because of data availability. We begin the earlier period in January 1916 because this is the beginning of the sustained inflation that lasts until the middle of 1920. We end in December 1927 because data availability requires us to use assets with risk premiums which begin to undergo substantial fluctuations after this date.

For the 1953-84 period, the nominal interest rates are computed from one-month U.S. Treasury bill prices (end of the month) obtained from the bond file of the Center for Research in Security Prices (CRSP) at the University of Chicago. The rates are continuously compounded

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rise rather than fall, with an increase in the relative price of energy. Both Wilcox (1983) and Huizinga and Mishkin (1985) find that a rise in the relative price of energy is associated with a fall in ex ante real interest rates.

12 Fama and Gibbons (1982), Huizinga and Mishkin (1985), Litterman and Weiss (1985), Nelson and Schwert (1977), Mishkin (1981), and Summers (1983).

rates, expressed at a monthly rate.

The inflation rate is computed as the one-month change in the log of the consumer price index (CPI). To obtain a price index that appropriately treats housing costs on a rental equivalence basis, we use four series to construct our CPI measure. From January 1983 to December 1984, we use the Bureau of Labor Statistics' (BLS) widely reported CPI-U series, which is on a rental equivalence basis. Before January 1983, however, the CPI-U series is not on a rental equivalence basis which leads to serious biases in measures of real rates in the 1970s.[13] Thus, from January 1954 to December 1982, we use the series CPI-UX, a series which is consistent with the currently produced CPI-U series. This is actually two series because the data from January 1967 to December 1982 is obtained from the BLS, while the data from January 1954 to December 1966 is obtained from the Congressional Budget Office. Prior to January 1954, we use the CPI-U series.

Our measure of supply shocks is calculated as the log of the relative price of fuel and related products in the producer price index. The data were obtained from the BLS. Money supply growth is calculated as the one-month change in the log of monthly averages of seasonally adjusted M1. The money supply data was obtained from the Board of Governors of the Federal Reserve System. It includes the February 1984 benchmark revisions, but not those of February 1985.

For the 1916-27 period, we use three nominal interest rate series because a series for a one-month government bond is unavailable. The three series are monthly averages of the rate on 4-6 month prime commer-

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[13] See Huizinga and Mishkin (1984). Note that the BLS does not revise earlier CPI-U data when it changes the procedures for its calculations. This requires us to construct a consistent CPI series using the procedures outlined above.

cial paper, the rate on 90 day stock exchange time loans, and the average rate on stock exchange call loans (renewal rate). These data are obtained from the Board of Governors of the Federal Reserve System, Banking and Monetary Statistics 1914-1941.[14] Monthly averaged data is used because last-week-of-the-month data is available only after 1919. We have used the last-week-of-the-month data in the 1919-1927 period and have obtained similar results to those found for monthly averaged data. All rates are treated as though they were one-month rates and have been converted to continuously compounded rates, expressed at a monthly rate.

Inflation in the 1916-27 period is measured as the one-month change in the log of the CPI-U series provided by the BLS.

The timing of variables is as follows. The real interest rates we examine are for a one-month holding period.[15] The ex post real interest rate ($eprr_t$) for 1953:1 is the realized real return for a one-month U.S. Treasury bill over the month of January 1953. It equals the nominal interest rate on the one-month bill (i_t) on December 31, 1952 minus the inflation rate (π_t) calculated using the December 1952 and January 1953 CPI. The supply variable for 1953:1 ($SUPPLY_t$) is calculated using the January 1953 fuel and related products component of the PPI and the

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[14] Macaulay (1938), Appendix E, describes background information on these rates including a description of why they are the most representative of market-clearing interest rates in this period.

[15] In previous work [Huizinga and Mishkin (1984, 1985)] we have argued that one-month holding periods have timing problems that arise because the appropriate dating for the CPI is unclear. Although three-month holding periods have the advantage of reducing the relative magnitude of this timing problem, they have the disadvantage of reducing the number of non-overlapping observations. In our empirical analysis here, which centers on dating breakpoints and thus requires estimation over short sample periods, we feel that the loss of degrees of freedom resulting from the use of three month holding periods outweighs the benefits from reducing the timing problem.

January 1953 overall PPI. For 1916:1, the ex post real interest rate for each of the three assets ($eprr_t$) equals its average nominal interest rate (i_t) for December 1915 minus the inflation rate (π_t) calculated using the December 1915 and January 1916 CPI.

IV. The Nature and Timing of Shifts in the Real Rate Process: 1953-84

Our discussion of methodology in Section II suggests that X_t, the explanatory variables in the ex post real rate regression, should include lags of ex post real rates and inflation, as well as other economic variables known at time t, such as the nominal interest rate and a supply shock variable. In practice, the number of variables that should be included in X_t must be limited in order to avoid overfitting the data. This is especially necessary in the present analysis because we are examining time periods as short as 26 months.

This need for parsimony has led us to terminate lag lengths and omit variables when doing so does not significantly diminish the explanatory power of the ex post real rate regression. To arrive at our specification, we examined separate regressions for the 1953:1 to 1979:10, 1979:11 to 1982:10, and 1982:11 to 1984:12 sample periods. We chose this division of sample periods to correspond with our prior belief that the changes in the Federal Reserve operating procedures in October 1979 and October 1982 constitute regime shifts. Variables which entered significantly at the 5% level in any of the three periods were included in our final specification for X_t. The resulting specification of X_t includes a constant term, the one-month nominal interest rate (i_t), the one-month inflation rate lagged one and two months (π_{t-1} and

π_{t-2}), and the supply shock variable lagged one month ($SUPPLY_{t-1}$).[16]

Recall that a diagnostic check of the X_t specification involves examining the residuals from the ex post real rate regression to see if they are white noise. Table 1 reports the autocorrelations of the regression residuals for the three sample periods mentioned above. In all three sample periods there is little evidence of serial correlation of the residuals.[17] The first twelve autocorrelations are small and the Ljung-Box (1978) Q(12) statistic, which is distributed asymptotically as $\chi_2(12)$ under the null hypothesis that the residuals are serially uncorrelated, is not significant at the 5% level for any of the three periods. Given that we have found no evidence which contradicts our specification, we proceed to an examination of the nature and timing of shifts in the real rate process.

The October 1979 change in operating procedures away from interest rate smoothing is a natural candidate for a monetary policy regime

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[16] Variables which we have examined but have dropped from our final specification include lags of the ex post real rate, additional lags of inflation and supply shocks, and lags of industrial production and detrended industrial production. Note that adding lags of nominal rates is equivalent to adding lags of the ex post real rate, given that we have included lags of inflation. Other potentially relevant variables -- for example, the unemployment rate, investment, budget deficits, and, most notably, money supply growth -- were not included in our specification search either because monthly data were unavailable or because previous research fails to find that these variables have any significant explanatory power for ex ante real rates. For example, Litterman and Weiss (1985) and Mishkin (1981) find that money supply growth does not help to predict ex ante real rates. However, it is important to realize this finding does not imply that monetary policy is unimportant for the determination of ex ante real interest rates.

[17] There is some evidence of seasonality in the residuals because the autocorrelation at lag 12 in the 1953:1 to 1979:10 sample period is significantly different from zero at the 5% level, although not at the 1% level. However, as pointed out in Mishkin(1981), we do not devote much attention to this seasonality because it may well be spurious because of the fixed-weight nature of the CPI.

Table 1

Autocorrelations of Ex Post Real Rate Regression Residuals, 1953-1984

Sample period	Autocorrelation at Lag 1	2	3	4	5	6	7	8	9	10	11	12	SE	Q(12)	MSL
1953:1 - 1979:10	.01	-.01	-.03	-.02	-.04	-.01	-.04	-.02	.07	.11	.03	.14	.06	14.72	.26
1979:11- 1982:10	-.06	-.04	-.07	.08	.15	.07	-.15	.05	-.15	-.12	.10	.10	.17	6.07	.91
1982:11- 1984:12	.14	.06	.03	-.14	.07	-.14	.04	-.10	.17	-.03	-.06	.18	.20	5.61	.93

SE = asymptotic standard error of the estimated autocorrelations under the assumption that the true residuals are serially uncorrelated.

Q(12) = adjusted Q-statistic suggested by Ljung and Box (1978) which is distributed asymptotically as $\chi^2(12)$ under the null hypothesis that the residuals are serially uncorrelated.

MSL = marginal significance level of Q-statistic, i.e., the probability of obtaining that value of the χ^2 statistic or higher under the null hypothesis that the residuals are serially uncorrelated.

Residuals from a regression of the ex post real rate, $eprr_t$, on a constant, π_{t-1}, π_{t-2}, i_t and $SUPPLY_{t-1}$.

shift. The deemphasis of monetary aggregates by the Fed in October 1982 is another natural candidate. Our next set of results examine whether there were shifts in the real rate process at these dates. The stability test reported in the first column of Table 2 indicates that there was a highly significant change in the real rate process in October 1979. The $F(5,348) = 12.14$ statistic indicates that the null hypothesis that the coefficients of the ex post real rate regression remain unchanged from the 1953:1-1979:10 period to the 1979:11-1982:10 period is rejected at a marginal significance level of 1.7×10^{-10}.[18]

The F-test presented in the second column of Table 2 shows that there is a second significant shift in the real rate process in October 1982. The finding of a second significant shift is surprising. Although the finding that there is a shift in the real rate process in October 1979 could have been guessed by a casual examination of dramatic rise in ex post real rates in 1980, such an examination would not have revealed a further shift in October 1982 because ex post real rates remained high through the end of 1984.

An additional interesting finding, reported in the third column of Table 1, is that the stochastic process of real rates after October 1982 does not represent a return to the pre-October 1979 process. The F-statistic which tests for coefficient equality in the pre-October 1979

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[18] We did not find any evidence of significant heteroscedasticity across the sample periods examined in Table 2 and thus did not correct for it when computing the F-tests. (The test for heteroscedasticity involved an F-test of the ratio of the estimated variances over the differing sample periods.) We also tested for equality of coefficients in the 1953:1-1979:10 and 1979:11-1984:12 sample periods, obtaining $F(5,374) = 11.44$ which is significant at a marginal significance level of 3.7×10^{-10}. Our results are consistent with the finding of Clarida and Friedman (1984) that the relationship of nominal interest rates with other economic variables shifted after October 1979.

Table 2

Stability Tests for Ex Post Real Rate Regressions, 1953-1984

F-test for Equality of Coefficients in			
1953: 1 - 1979:10 and 1979:11 - 1982:10	1979:11 - 1982:10 and 1982:11 - 1984:12	1953: 1 - 1979:10 and 1982:11 - 1984:12	1953: 1 - 1966: 5 and 1966: 6 - 1979:10
$F(5,348) = 12.14$**	$F(5,52) = 3.99$**	$F(5,338) = 8.90$**	$F(5,312) = 1.12$
(1.7×10^{-10})	(.0039)	(5.9×10^{-8})	(.34)

Regressions of the ex post real rate, $eprr_t$, on π_{t-1}, π_{t-2}, i_t and $SUPPLY_{t-1}$.

Marginal significance levels in parentheses; i.e., the probability of obtaining that value of the F-statistic or higher under the null hypothesis the coefficients in the ex post real rate regression are equal in the two periods.

* = significant at the 5% level
** = significant at the 1% level

and post-October 1982 periods is highly significant.

Since we have found significant shifts in the stochastic process of real rates after October 1979, it is interesting to ask whether similar shifts occurred in our sample prior to this date. The final column addresses this question by testing for coefficient stability in the two halves of the pre-October 1979 period. The F-statistic does not reveal a shift in the real rate process. We have also examined coefficient stability over various three year samples in the 1953:1-1979:10 period. Although we find one three year sample whose coefficients are significantly different from those in the remainder of the period, the change is not nearly as significant as the change in October 1979.[19] Further, dropping this three year sample from the 1953:1-1979:10 period does not alter the significant rejection of coefficient stability pre- and post-October 1979.

While we have found significant shifts in the coefficients of the ex post real rate regressions, we have not yet determined the dates at which the real rate process shifts: i.e, determined the breakpoints. Table 3 reports minus twice the log of the Quandt likelihood ratio for pairs of breakpoints surrounding the October 1979 and October 1982 dates. Remarkably, the Quandt procedure indicates that October 1979 and October 1982 are the breakpoints that maximize the likelihood of the ex

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[19] Of the nine 3-year periods from 1953:1-1979:10, we find significantly different coefficients only in 1973:11-1976:10 [F(5,312) = 2.99, marginal significance level = .012.] We also tested for coefficient equality in each 3-year sample and the subsequent 26-month sample. Of the eight tests we conducted, none is significant at the 5% level. This latter finding gives us some confidence that our rejection of equal coefficients in 1979:11-1982:10 and 1982:11-1984:12 is not due to small sample problems.

[20] We have also examined minus twice the log of the Quandt likelihood ratio for breakpoint dates outside the range reported in Table 3.

Table 3

Dating of Breakpoints in 1953-84 Period

Minus Twice the Log of the Quandt Likelihood Ratio

Second Breakpoint	First Breakpoint 1979:4	1979:5	1979:6	1979:7	1979:8	1979:9	1979:10	1979:11	1979:12	1980:1	1980:2	1980:3	1980:4
1982: 4	72.16	71.04	71.52	71.76	71.72	71.68	71.68	69.90	69.58	67.60	66.96	67.52	59.52
5	70.08	68.98	69.46	69.70	69.68	69.66	69.94	68.48	68.30	66.08	65.36	65.96	57.54
6	75.64	74.54	75.00	74.98	75.22	74.92	75.46	73.82	73.58	71.26	70.44	71.00	62.34
7	74.98	73.88	74.36	74.60	74.58	74.56	75.10	73.56	73.42	71.26	70.36	71.08	62.20
8	71.94	70.62	71.38	71.64	71.90	71.64	72.40	71.22	71.40	69.00	68.14	68.60	60.00
9	76.86	75.80	76.32	76.56	76.82	76.58	77.08	75.82	75.72	73.34	72.78	73.52	64.08
10	76.74	75.68	76.46	76.70	76.96	76.98	77.48	76.52	76.72	74.30	73.46	74.20	65.06
11	73.12	71.86	72.88	73.16	73.42	73.48	74.20	73.26	73.74	71.38	70.58	71.06	62.40
12	71.64	70.40	71.42	71.68	72.20	72.24	72.74	72.06	72.28	69.92	69.10	69.60	61.48
1983: 1	71.02	69.76	70.78	71.04	71.56	71.60	72.34	71.52	71.72	69.34	68.78	69.00	60.92
2	72.06	70.84	71.88	72.16	72.68	72.76	73.22	72.28	72.26	70.22	69.46	69.76	62.52
3	74.00	72.78	73.82	74.10	74.62	74.46	74.92	73.46	73.44	71.42	70.92	71.22	63.96
4	73.96	72.78	73.82	74.34	74.84	74.96	75.40	74.52	74.54	72.52	71.82	72.12	64.76

post real rate regressions.[20] The evidence in Table 3 reinforces that of Table 2 by suggesting that the shifts in the monetary policy regime may account for the unusual behavior of real interest rates in recent years.

One reason why we have reported minus twice the log of the Quandt likelihood ratio for so many pairs of breakpoints is to indicate that the likelihood function around these breakpoints is quite flat.[21] For example, the pair of breakpoints at (1979:4, 1982:10) has a value of 76.74 which is very close to the maximum of 77.48 found at (1979:10, 1982:10). Similarly, the pair breakpoints at (1979:10, 1983:4) has a value of 75.40, again close to the maximum of 77.48. Therefore, although we have found that the most likely breakpoints in the real rate process coincide with the recent changes in the Fed operating procedures, the evidence does not appear to be sufficiently strong to convince someone with prior beliefs that the breakpoints occur at 1979:10 and 1983:4 that they are incorrect.

Given that the evidence is consistent with breakpoints at October 1979 and October 1982, we construct estimates of the ex ante real rate using fitted values from three separate ex post real rate regressions, estimated over the periods 1953:1-1979:10, 1979:11-1982:10, and

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The pair of breakpoints at October 1979 and October 1982 is the global maximum.

[21] Although we report minus twice the log of the Quandt likelihood ratio, Quandt (1960) has pointed out that the maximum value across all possible breakpoints of this statistic is not distributed as χ^2. His Monte Carlo simulations indicate that the distribution of the statistic has fatter tails than the χ^2 distribution.

[22] The regressions for the three sample periods are as follows:

1953:1 to 1979:10

$$eprr_t = \underset{(.00025)}{.000697} + \underset{(.1035)}{.2467}\, i_t - \underset{(.0544)}{.1204}\, \pi_{t-1} - \underset{(.0542)}{.2000}\, \pi_{t-2}$$

1982:11-1984:12.[22] The heavy solid line in Figure 1 displays the estimated ex ante real rates, while the dotted lines surrounding the estimated real rates provide the 95% confidence interval for these estimates under the assumption that the variance of inflation forecast errors are large relative to the variance of u_t.[23]

The estimated ex ante real rates from 1953:1 to 1979:10 are consistent with earlier findings that real rates were generally positive prior to late 1973 when the first oil shock occurs, ranging from -1% to 2.5% at an annual rate. After 1973, they are generally negative, falling to as low as -2.5%. Following the break in October 1979, estimated real rates remain low until late 1980 when they begin to rise to over 7%. They fluctuate around this level until June 1982 when they fall sharply with the decline in nominal interest rates. Following the break in

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$$- \underset{(.00083)}{.002390}\ SUPPLY_{t-1}$$

Standard Error = .001982 R^2 = .1598 Durbin-Watson = 1.98

1979:11 to 1982:10

$$eprr_t = - \underset{(.00520)}{.017305} + \underset{(.1750)}{.9368}\ i_t - \underset{(.1709)}{.6088}\ \pi_{t-1} + \underset{(.1563)}{.2920}\ \pi_{t-2}$$

$$+ \underset{(.00623)}{.01652}\ SUPPLY_{t-1}$$

Standard Error = .002133 R^2 = .6809 Durbin-Watson = 1.98

1982;10 to 1984:12

$$eprr_t = \underset{(.01246)}{-.00093} - \underset{(.6864)}{.0528}\ i_t - \underset{(.2047)}{.1940}\ \pi_{t-1} + \underset{(.2005)}{.3195}\ \pi_{t-2}$$

$$+ \underset{(.012032)}{.006256}\ SUPPLY_{t-1}$$

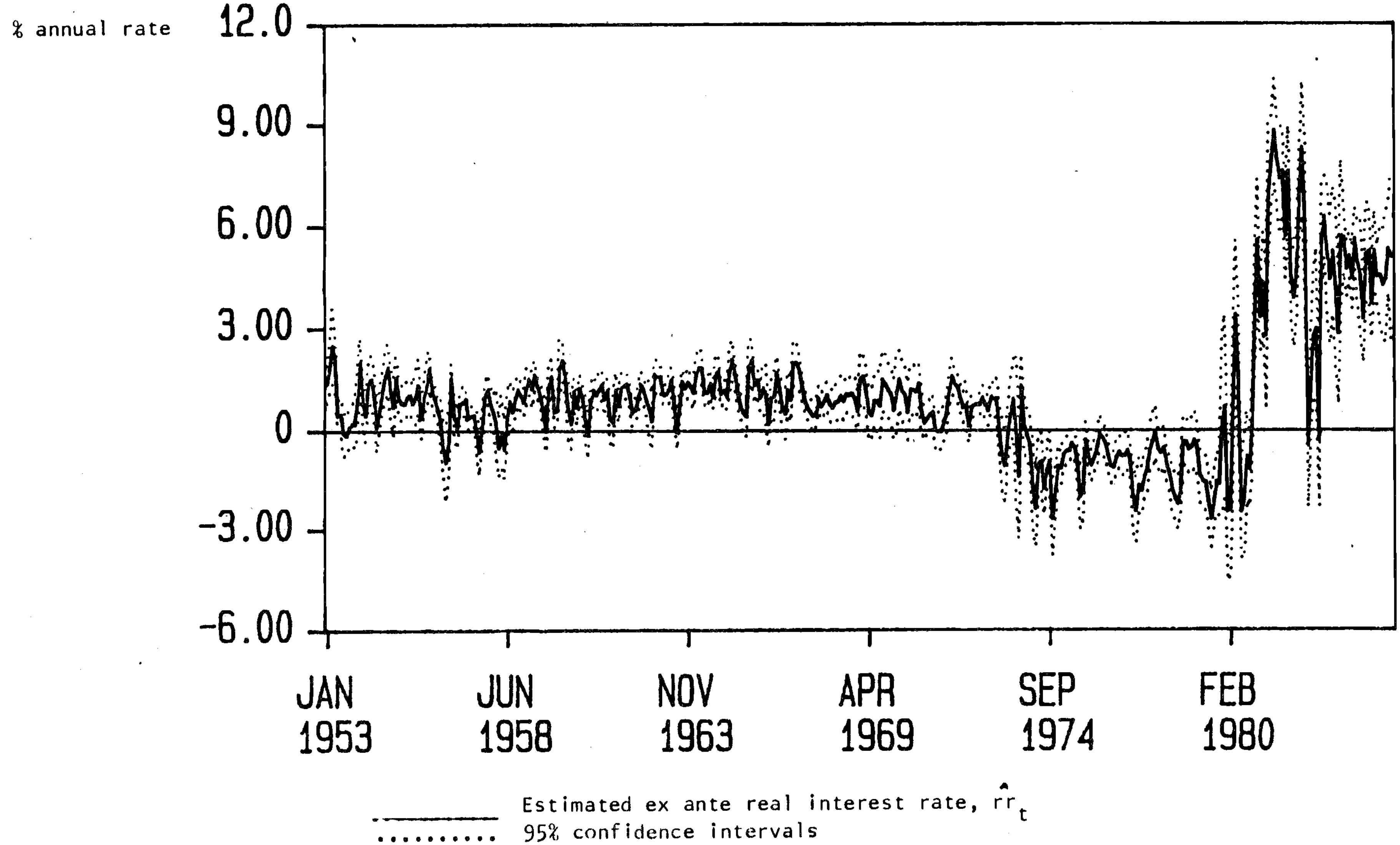
Standard Error = .001679 R^2 = .1521 Durbin-Watson = 1.70

Note that the coefficients of the constant term and $SUPPLY_{t-1}$ above have not been multiplied by 100 as in Table 4.

[23] The formula for the confidence interval is given in Mishkin (1981).

Figure 1

Estimated Ex Ante Real Interest Rate, 1953-84



October 1982, estimated real rates fluctuate around the 5% level.[24]

One striking feature of Figure 1 is that the Fed's switch to their new operating procedures in October 1979 is associated with a notable rise in the variability of ex ante real interest rates. Following the Fed's abandonment of these operating procedures in October 1982, the variability of ex ante real rates falls sharply to a level comparable to that before October 1979. The standard deviation of estimated ex ante real rates (annual rate) is 1.0% from 1953:1 to 1979:10, 3.5% from 1979:11 to 1982:10, and .8% from 1982:11 to 1984:12.[25]

Another interesting feature of Figure 1 is that our dating of the first break in the real rate process at October 1979 does not coincide with the substantial rise in the estimated real rates which occurs in late 1980. This finding indicates why focusing solely on the level of real interest rates does not provide an accurate characterization of the unusual behavior of real interest rates.

Figure 1 provides one way of characterizing the nature of the shifts in the real rate process. The regression results presented in Table 4 provide another way. Table 4 describes the relationship of ex ante real rates with expected inflation, nominal interest rates and the supply shock variable in the three sample periods defined by the

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[24] The sharp rise in the estimated real rates from 1982:10 to 1982:11 is probably an artifact of splitting the sample for the ex post real rate regressions at 1982:10.

[25] The standard deviation of the estimated ex ante real rate for the 1953:1 to 1979:10 period may give a somewhat misleading impression of the variability of ex ante real interest rates because there is an apparent shift in the mean at the end of 1973 (see Figure 1). The standard deviation of the estimated ex ante real rate (annual rate) for the 1953:1 to 1973:12 period is .6%, while for the 1974:1 to 1979:10 period it is .7%. These standard deviations suggest that the variability of ex ante real interest rates before October 1979 may be slightly lower than the standard deviation for the 1953:1 to 1979:10 period suggests.

Table 4

The Relationship of Ex Ante Real Interest Rates with Nominal Interest Rates, Expected Inflation and Supply Shocks, 1953-84

Dependent Variable: Ex Post Real Interest Rate on a 1-month U.S. Treasury Bill, $eprr_t$

Estimation Method: Ordinary Least Squares

Sample Period	Constant term[a]	Coefficient of π_t^e	i_t	$SUPPLY^a_{t-1}$	SER	DW
1953: 1 - 1979:10	.12** (.01)	-.29** (.04)			.0021	1.71
	.12** (.03)		-.26** (.07)		.0021	1.59
	.07** (.01)			-.41** (.07)	.0020	1.68
1979:11 - 1982:10	.89** (.13)	-.91** (.19)			.0031	1.55
	-.61** (.21)		.93** (.22)		.0029	.81
	-2.21** (.60)			3.08** (.74)	.0029	1.22
1982:11 - 1984:12	.56** (.05)	-.53** (.16)			.0016	1.42
	.69 (.40)		-.40 (.57)		.0017	1.29
	-.28 (.79)			.88 (1.01)	.0017	1.35

standard errors in parentheses
* = significant at the 5% level
** = significant at the 1% level
SER = standard error of the regression
DW = Durbin-Watson Statistic

[a] Coefficients and their standard errors have been multiplied by 100.

breakpoints.[26] The regression results of Table 4 indicate that in all three sample periods, there is a significant negative correlation between ex ante real rates and expected inflation.[27] This finding is a surprisingly robust one, having been found in previous research for many different sample periods in the U.S. and for other countries.[28]

In contrast to the robustness of the relationship between expected inflation and ex ante real rates, the relationship between nominal interest rates and ex ante real rates changes dramatically with the policy regime shifts in October 1979 and October 1982. In the pre-October 1979 period, nominal rates are a misleading indicator of ex ante real interest rates because the two rates are significantly negatively correlated. With the Fed's deemphasis of interest rate smoothing following October 1979, the correlation of nominal and ex ante real interest rates becomes significantly positive. Following the Fed's abandonment of

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[26] The regressions of Table 4 use the ex post real rate as the dependent variable. The coefficient estimates would be numerically the same if the dependent variable were the estimated ex ante real rate because, by construction, the regressors are orthogonal to the difference between the ex post and the estimated ex ante real rates. Further, the coefficient estimates would be asymptotically the same as those obtained in regressions with the true ex ante real rate as the dependent variable. See Mishkin (1981).

[27] The expected inflation variable, π_t^e, which is an explanatory variable in the regressions is calculated as the nominal interest rate minus our estimated values of the ex ante real rate. Since this involves the two step procedure of first estimating expected inflation and then regressing the ex post real rate on this measure, the standard errors of the coefficients typically reported will be incorrect. We have obtained consistent estimates of the standard errors by using the instrumental variables procedure outlined by McCallum (1976) and Pagan (1984). Since some of the Durbin-Watson statistics are low, we have also estimated the standard errors correcting for possible serial correlation of the error terms using the procedure outlined in Cumby, Huizinga and Obstfeld (1983). Although the standard errors rose slightly, there was no appreciable difference in the significance of the coefficients.

[28] Mishkin (1981,1984) and Summers (1983).

the new operating procedures, the correlation of nominal and estimated ex ante real rates returns to being negative.

The shifts in the real rate process are also reflected in the changing strength of the Fisher effect (i.e., the positive correlation between expected inflation and the nominal interest rate). The Fisher effect is strong in the pre-October 1979 period, is nonexistent from November 1979 to October 1982, and is strong again after October 1982: the correlation coefficient between the nominal interest rate and our measure of expected inflation is .95 for 1953:1-1979:10, .08 for 1979:11-1982:10, .81 for 1982:11-1984:12. The lack of a Fisher effect from 1979:11 to 1982:10 is not surprising considering the dramatic increase in the variability of the estimated ex ante real rate during this period. Only if there was a similar rise in the variability of expected inflation or a change in the correlation of expected inflation with ex ante real rates could this fail to be the case.

The relationship of ex ante real rates and the supply shock variable also shifts following October 1979. Whereas before October 1979 a rise in the relative price of energy is associated with a significant decrease in ex ante real rates, after this date it is associated with a rise in ex ante real rates. Given the theoretical ambiguity associated with the effects of supply shocks on ex ante real rates and this change in sign of its regression coefficient, we have some concern that the

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[29] Because of our concern about the interpretation of the supply shock variable, we have conducted the stability tests described in Table 2 excluding the supply shock variable from the list of regressors. We continue to find significant shifts in the real rate process in October 1979 and October 1982. The test for equality of coefficients in 1953:1-1979:10 and 1979:11-1982:10 yields $F(4,350) = 10.73$ (marginal significance level = 3.3×10^{-8}). The test for equality of coefficients in 1979:11-1982:10 and 1982:11-1984:12 yields $F(4,54)= 3.71$ (marginal significance level = .0096).

supply shock variable may be proxying for other economic factors.[29]

Our characterization of the real rate process suggests that there are strong similarities between the 1953:1-1979:10 and 1982:11-1984:12 periods. Both periods display a negative correlation between nominal interest rates and ex ante real rates, a strong Fisher effect, and similar variability in ex ante real rates.[30] In contrast, the 1979:11-1982:10 period, when the Fed deemphasizes interest rate smoothing,

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[30] Blanchard and Summers (1984) suggest that an upward shift in investment profitability may have been an important source of high ex ante real interest rates after 1982. Following Blanchard and Summers, we have also examined forecast errors from time series models of real gross private fixed investment to see if investment has been strong since October 1979.

Because residuals from a regression of investment on current and past values of real GNP appear to be nonstationary, our forecasting models regress the change in investment on the current and lagged changes in real GNP over the the 1953:I-1979:III period. We then conducted dynamic simulations for the period beginning with 1979:IV to obtain forecasts of investment to compare with actual investment. As long as four lags of real GNP changes are included as regressors, further lags of real GNP and alternative ARIMA specifications for the error term produce very similar findings.

For real gross business fixed investment (gross private fixed investment excluding residential investment), the dynamic forecast errors are positive after 1980 and grow in size in 1983 and 1984, averaging $20 billion (1972 dollars) for all of 1984. This evidence provides some support for Blanchard and Summers contention that investment spending has been strong recently. However, for real gross private fixed investment (including residential investment), there is much less evidence for buoyant investment spending before 1983. Dynamic forecast errors are negative until 1983:IV, and they average only $5 billion (1972 dollars) in 1984.

An upward shift in profitability may help to explain why the stochastic process of real rates after October 1982, although similar in many ways to that before October 1979, is significantly different. In particular, the overall level of ex ante real rates is substantially higher after October 1982 than it is before October 1979 and this may be related to the strong investment performance in this period. A test for whether coefficients remain stable across these two periods if the constant terms are allowed to differ [$F(4,338) = 1.41$ (marginal significance level = .23)] indicates that the source of the significant difference between the two periods is an upward shift in the level of ex ante real rates. (Note, however, that this test may not have very much power because we also found that, when comparing the 1982:11-1984:12 and 1979:11-1982:10 periods, we could not reject stability of the other coefficients besides the constant term.)

displays a positive correlation between nominal and ex ante real rates, a weak Fisher effect, and variability in ex ante real rates that is substantially higher than in the periods before or after.

The characterization of the ex ante real rate process from 1979:11-1982:10 is not totally without precedent, however. As found in Mishkin (1984), West Germany, the Netherlands and Switzerland also have a positive correlation between nominal and ex ante real rates, a weak Fisher effect and greater variability of ex ante real rates than the U.S. in the period 1967-79. Could it be that the change in the Fed operating procedures in October 1979 produced a monetary policy regime similar to that found in these countries before 1979?

V. Uncertainty and Ex Ante Real Rates

Casual inspection of the data for the October 1979 to October 1982 period reveals that the adoption of the Fed's new operating procedures led to increased uncertainty about interest rate movements, money supply growth, and possibly inflation. This, combined with the results of the previous section which suggest that the Fed's change in operating procedures may account for the shifts in the real rate process, makes it worth examining whether this increased uncertainty is the critical feature of the monetary policy regime shift that alters real rate movements in recent years. Such a view has been espoused by Mascaro and Meltzer (1983) who argue that increased money growth uncertainty caused the rise in ex ante real interest rates after October 1979. Research by Fama (1976) and Hartman and Makin (1982) has also pointed to uncertainty about inflation and interest rates as possible factors that affect ex ante real interest rates.

Because the Mascaro and Meltzer (1983) results concerning the effects of money growth uncertainty on interest rates are so striking, we extend our analysis to include a measure of money growth uncertainty which was constructed using their procedure. Specifically, we have fit univariate time series models to the M1 money growth series in order to extract a series of forecast errors.[31] The money growth uncertainty measure (σ^{mm}_{t-1}) equals the square root of the average sum of squared forecast errors over the previous twelve months.

Table 5 presents estimates of ex post real rate regressions that include σ^{mm}_{t-1} in the information set X_t. We look at the 1953:1 to 1979:10 period to see whether money growth uncertainty has any explanatory power before the Fed adopts its new operating procedure. We look at the 1953:1 to 1984:12 period to see whether uncertainty can help explain the recent unusual behavior of ex ante real interest rates. The results do not support the proposition that money growth uncertainty is an important determinant of ex ante real interest rates. In contrast to Mascaro and Meltzer (1983), we find that the σ^{mm}_{t-1} coefficients enter with a negative sign in three out of four cases so that they cannot explain the rise in ex ante real rates after October 1979. In addition they are never sig-

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[31] After performing Box-Jenkins (1970) identification procedures, the following ARIMA models were fit to data for the three periods: for 1952:1-1979:10, an autoregressive model with AR parameters at lags 2, 3, 4, and 6; for 1979:11-1982:10 and 1982:10-1984:12, autoregressive models with AR parameters at lag 1. The σ^{mm} measure constructed from the one-step-ahead forecast errors from these models is the one used in Table 5. In order to check for robustness, we also constructed another σ^{mm} measure from the forecast errors of an autoregressive model with AR parameters at lags 1, 3 and 6, fit to the entire 1952:1-1984:12 sample period. Results of using this measure were roughly similar: the signs of the estimated coefficients on σ^{mm}_{t-1} were identical to those reported in Table 5 and were insignificant in all but the last regression where it entered with a coefficient of .12 and a t-statistic of 2.28.

Table 5

Correlation of the Ex Ante Real Interest Rate with Mascaro-Meltzer Measure of Money Growth Uncertainty

Dependent Variable: Ex Post Real Interest Rate, $eprr_t$
Estimation Method: Ordinary Least Squares

	Constant term[a]	Coefficients of i_t	π_{t-1}	π_{t-2}	$SUPPLY^a_{t-1}$	σ^{mm}_{t-1}	SER	DW
1953: 1 - 1979:10	.11** (.03)	.23* (.10)	-.11* (.05)	-.20** (.05)	-.24** (.08)	-.09 (.05)	.0020	1.97
	.07** (.02)					-.10 (.06)	.0021	1.52
1953: 1 - 1984:12	.01 (.03)	.61** (.08)	-.31** (.05)	-.26** (.05)	.09 (.06)	-.02 (.05)	.0022	2.02
	.05 (.03)					.10 (.06)	.0025	1.13

standard errors in parentheses
* = significant at the 5% level
** = significant at the 1% level
SER = Standard Error of the Regression
DW = Durbin-Watson Statistic

[a]coefficients and their standard errors have been multiplied by 100.

nificantly different (at the 5% level) from zero.[32]

We have also tested whether the inclusion of this measure of money growth uncertainty affects the conclusion that significant shifts in the real rate process occur in October 1979 and October 1982. Not surprisingly, the rejections of stability continue to be strong. For example, the test for equality of coefficients in the 1953:1-1979:10 and 1979:11-1982:10 periods yields an $F(6,346) = 10.80$ (marginal significance level $= 5.1\times10^{-11}$). The test for equality of coefficients in the 1979:11-1982:10 and 1982:11-1984:12 periods yields an $F(6,50) = 2.69$ (marginal significance level $= 2.4\times10^{-4}$).

The failure of the results in Table 5 to provide evidence that money growth uncertainty is an important determinant of ex ante real interest rates may be attributable to econometric difficulties. Measuring uncertainty with a weighted average of past squared forecast errors and then including it in an OLS regression can be thought of as a classic errors-in-variables problem.[33] A way around this problem has been suggested by Pagan (1984). His procedure involves including the contemporaneous measure of uncertainty as a regressor and estimating the equation by instrumental variables using lagged measures of uncertainty as instruments.

Table 6 reports results obtained when Pagan's procedure is used to

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[32] While negative coefficients on σ_{t-1}^{mm} are the "wrong" sign from the point of view of explaining the rise in ex ante real rates after October 1979, they are not inconsistent with some theoretical models of how uncertainty affects ex ante real interest rates. For example, see Bohn (1985).

[33] The uncertainty variable that presumably belongs in the regression is the true standard deviation of the forecast error which is measured imperfectly as the square root of an equally weighted average of past squared forecast errors. This mismeasurement never disappears, even if the sample size grows to infinity, because the number of forecast errors used in the estimation is independent of sample size.

Table 6

Correlation of the Ex Ante Real Interest Rate with Money Growth, Inflation and Nominal Interest Rate Uncertainty

Dependent Variable: Ex Post Real Interest Rate, $eprr_t$
Estimation Method: Instrumental Variables

Sample Period	Constant[a]	Coefficients of i_t	π_{t-1}	π_{t-2}	$SUPPLY^{a}_{t-1}$	σ^m_t	σ^π_t	σ^i_t	SER	DW
1953: 1 - 1979:10	.13**	.25*	-.12*	-.22**	-.22**	-.19			.0021	1.97
	(.04)	(.11)	(.06)	(.06)	(.09)	(.10)				
	.09**					-.18			.0022	1.55
	(.03)					(.11)				
	-.01	.26*	-.16*	-.21**	-.19		.57		.0023	1.98
	(.08)	(.12)	(.07)	(.06)	(.10)		(.52)			
	.01						.21		.0022	1.52
	(.07)						(.45)			
	.10**	.32**	-.11*	-.16	-.27**			-2.65	.0021	1.95
	(.04)	(.12)	(.06)	(.06)	(.09)			(2.04)		
	.15**							-4.33**	.0023	1.64
	(.04)							(1.61)		
1953: 1 - 1984:12	.02	.61**	-.31**	-.27**	.09	-.06			.0022	2.02
	(.03)	(.08)	(.05)	(.05)	(.06)	(.10)				
	.07*					.03			.0025	1.12
	(.04)					(.11)				
	-.14	.57**	-.36**	-.26**	.14		1.10		.0028	2.02
	(.09)	(.11)	(.07)	(.06)	(.08)		(.61)			
	.02						.40		.0027	1.17
	(.08)						(.54)			
	.00	.62**	-.31**	-.26**	.09			-.11	.0022	2.02
	(.02)	(.09)	(.05)	(.05)	(.06)			(.71)		
	.04							1.20**	.0025	1.18
	(.02)							(.45)		

standard errors in parentheses
* = significant at the 5% level
** = significant at the 1% level
SER = Standard Error of the Regression
DW = Durbin-Watson Statistic

Instruments are a constant, i_t, π_{t-1}, π_{t-2}, $SUPPLY_{t-1}$, and four lags of the σ variable in the equation. [a] indicates that the coefficient, and its standard error, have been multiplied by 100.

examine the association of ex ante real rates with uncertainty not only about money growth, but also about inflation and nominal interest rates. The uncertainty variable for money growth (σ_t^m), inflation (σ_t^π) and interest rates (σ_t^i) are contemporaneous absolute values of forecast errors obtained from time series models.[34] The results for the money growth uncertainty variable are similar to those reported in Table 5. The signs of the σ_t^m coefficients are identical to those found for σ_{t-1}^{mm} and none of these coefficients is significant at the 5% level.

In all the regressions, the results on inflation uncertainty are consistent with the premise that increases in inflation uncertainty raise ex ante real interest rates. However, none of the σ_t^π coefficients is significantly different from zero. Results on interest rate uncertainty are somewhat more mixed. The σ_t^i coefficients are significant when

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[34] For money growth, we use the same forecast errors that were used in constructing the Mascaro-Meltzer uncertainty measure. Forecast errors for nominal interest rates were constructed from univariate time series models developed using Box-Jenkins identification procedures. The following ARIMA models for the change in the nominal interest rate were fit to data for the three periods: for 1952:1-1979:10, an autoregressive model with AR parameters at lags 1, 4, and 6; for 1979:11-1982:10 and 1982:10-1984:12, a white noise specification. Forecast errors for inflation were generated from the ex post real rate regressions estimated separately over the three periods. The results in Table 6 use four lags of the absolute values of the forecast errors as instruments. In order to check for robustness, we also estimated these regressions with only one lag and also twelve lags of the absolute of the forecast errors as instruments. We also obtained forecast errors from time series models estimated over the entire 1952:1-1984:12 sample period. The ARIMA model for money growth and the ex post real rate regression used to generate inflation forecast errors have already been described for this period, while the ARIMA model for the change in the nominal interest rate is an autoregressive model with AR parameters at lag 7 and 8. Results with all of these alternative specifications are similar to those found in Table 6. The signs of the σ coefficients are typically the same, and they are only significant in the same cases that they are significant in Table 6. These results are available from the authors by request.

[35] The finding of a significant negative coefficient in the 1953:1-

no other information is included in the regressions.[35] However, when other information is added to the regressions, nominal interest rate uncertainty has no significant additional explanatory power for ex ante real interest rates. Furthermore, in the specification in which the coefficients of σ_t^i are significantly different from zero, the coefficients change sign when the 1953:1-1979:10 sample is extended through 1984:12.[36]

In summary, uncertainty variables do not have much explanatory power for ex ante real rates.[37] Thus to understand how the recent changes in monetary policy regime might have affected ex ante real rates, it is necessary to look elsewhere.

VI. Ex Ante Real Rates and the 1920 Monetary Regime Shift

We have presented evidence suggesting that recent shifts in a

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1979:10 period is in agreement with the findings of Fama (1976).

[36] Bodie, Kane and McDonald (1983) also fail to find evidence that the riskiness of nominal bonds can explain the recent rise of ex ante real rates on short maturity assets.

[37] Although a variable may not be statistically significant, it is always worth asking whether the size of its coefficient might make it economically significant. Given that uncertainty for money growth, inflation and interest rates rose in the 1979:11-1982:10, we can see what would be the largest positive effect on ex ante real interest rates arising from increased uncertainty by applying the most positive coefficients in Table 6 to the change in the standard deviations of these variables from 1953:1-1979:10 to 1979:11-1982:10. The result of this exercise is that the largest positive effect on ex ante real rates from uncertainty is 10 basis points for monetary growth, 20 basis points from inflation and 189 basis points from nominal interest rates. The first two are clearly economically insignificant. Although the third can be considered economically important, recall that we must assume that no other information other than nominal interest rate uncertainty is helpful for predicting ex ante real rates (something we can statistically reject) in order to arrive at this figure.

monetary policy regime have been an important determinant of the behavior of real interest rates. If this is true, there should be breaks in the stochastic process of real rates in other periods which contain similar changes in the direction of monetary policy. One such period is 1916-27. In January and June of 1920, the Federal Reserve carried out the steepest rises in the discount rate in the first 50 years of its Fed's history -- raising the discount rate by 1 1/4% in January and 1% in June. Two aspects of this period make it especially attractive to study. First, since discount rate changes were the primary tool for conducting monetary policy in this period, the sharp rises in the discount rate represented a definite change in the direction of monetary policy. Changes in the discount rate were particularly potent at this time because the total amount of member bank borrowing from the Fed exceeded the amount of nonborrowed reserves. Second, this regime shift has several similarities to the more recent regime shift. In both periods, the U.S. had been experiencing a sustained high level of inflation and the dollar was facing foreign exchange pressure prior to the shift. After the regime change there was a rapid disinflation in both periods.

Our analysis of ex ante real interest rate behavior surrounding the time of the regime shift is conducted using the same techniques we applied previously. The major difficulty we encounter when examining this period is that reliable monthly data on one-month, default-free, government bonds are not available. The only reliable interest rate data are for 4-6 month commercial paper, 90 day stock exchange time loans, and stock exchange call loans. Since maturities of these assets span the one-month maturity we desire, and also because they have different risk characteristics, our strategy is to examine results based on interest rates from all three assets. Findings that are consistent for all three assets, would presumably extend to the ex ante real rate on a one-month U.S. Treasury bill, had such an asset existed. Another difficulty is that inflation data for this period are of lower quality than current inflation data. Nonetheless, the results for this period are so striking that we doubt that they would be substantially altered if higher quality data were available.

Our specification of the X_t variables in the ex post real rate regressions for the 1916-1927 period is the same as for the 1953-84 period, with the exception that data were not readily available to construct the supply shock variable so it was not included. This means that for each asset, the information set X_t contains a constant term, the nominal interest rate on that asset (i_t), and two lags of inflation (π_{t-1} and π_{t-2}). The results from the diagnostic check of the X_t-specification for the 1916-27 period appear in Table 7. There is no evidence of any serial correlation of the residuals from the ex post real rate regressions for any asset in either time period. Thus, we have found no evidence that we have omitted information relevant for predicting ex ante real rates from our regressions.

Table 8 contains the tests of coefficient stability in the ex post real rate regressions in the 1916-27 period. Given the two sharp rises in the discount rate in January and June of 1920, choosing a date for the monetary policy regime shift is somewhat ambiguous. In practice it does not make a difference: choosing either date, we find a highly significant shift of the real rate process. Since timing evidence presented in Table 9 indicates that the breakpoint occurs in June of 1920, the results of Table 8 use this dating.

The test for equality of the coefficients in the 1916:1-1920:6 and

Table 7

Autocorrelations of Ex Post Real Rate Regression Residuals, 1916-1927

Sample period and asset	Autocorrelation at Lag 1	2	3	4	5	6	7	8	9	10	11	12	SE	Q(12)	MSL
1916:1 - 1920:6 4-6 month commercial paper	-.00	.02	-.10	-.11	-.10	-.12	-.14	.05	-.15	-.11	-.06	.00	.14	6.81	.87
Stock exchange time loans, 90 days	-.01	.02	-.11	-.10	-.11	-.12	-.14	.05	-.15	-.10	-.06	.01	.14	6.76	.87
Stock exchange call loans	-.02	.02	-.11	-.09	-.10	-.09	-.12	.05	-.15	-.07	-.05	.01	.14	5.62	.93
1920:7 - 1927:12 4-6 month commerical paper	-.04	.06	-.12	.06	.06	.02	-.02	-.04	-.02	-.06	.07	-.01	.11	3.59	.99
Stock exchange time loans, 90 days	-.02	.02	-.11	.05	.06	.03	-.01	-.03	-.01	-.05	.04	-.04	.11	2.71	.99
Stock exchange call loans	-.02	.04	-.08	.09	.10	.02	-.02	-.02	-.02	-.08	-.01	-.07	.11	3.94	.99

SE = asymptotic standard error of the estimated autocorrelations under the assumption that the true residuals are serially uncorrelated.

Q(12) = adjusted Q-statistic suggested by Ljung and Box (1978) which is distributed asymptotically as $\chi^2(12)$ under the null hypothesis that the residuals are serially uncorrelated.

MSL = marginal significance level of Q-statistic, i.e., the probability of obtaining that value of the χ^2 statistic or higher under the null hypothesis that the residuals are serially uncorrelated.

Table 8

Stability Tests for Ex Post Real Rate Regressions, 1916-1927

Ex Post Real Rate Regression For	F-Tests for Equality of Coefficients in		
	1916: 1 - 1920: 6 and 1920: 7 - 1927:12	1916: 1 - 1918: 3 and 1918: 4 - 1920: 6	1920: 7 - 1924: 3 and 1924: 4 - 1927:12
4-6 month commerical paper	$F(4,136) = 13.60^{**}$ (2.4×10^{-9})	$F(4,46) = .50$ (.74)	$F(4,82) = .96$ (.43)
Stock Exchange Time Loans, 90 days	$F(4,136) = 12.77^{**}$ (7.5×10^{-9})	$F(4,46) = .47$ (.76)	$F(4,82) = 1.16$ (.33)
Stock Exchange Call Loans	$F(4,136) = 12.89^{**}$ (6.3×10^{-9})	$F(4,46) = 1.02$ (.41)	$F(4,82) = 1.79$ (.14)

Marginal significance levels in parentheses: i.e., the probability of obtaining that value of the F-statistic or higher under the null hypothesis that the coefficients in the ex post real rate regressions are equal in the two periods.

* = significant at the 5% level
** = significant at the 1% level.

1920:7-1927:12 is very similar for all three assets.[38] The F(4,136) statistics all are larger than 12 and are significant at the 10^{-8} level.[39] The additional tests presented in Table 8 indicate that there is only one significant shift in the real rate process during the 1916-27 period.[40]

Table 9 presents the results of using the Quandt procedure for dating the timing of the shift in the real rate process. In contrast to the likelihood surface examined for the 1953-84 period, the likelihood function is very well defined for all three assets in the 1916-27 period. Evidence for all three assets is in agreement that the most likely date for the break in the real rate process is June 1920, the date of the second large increase in the discount rate.

The timing evidence for the 1916-27 period indicates that, just as in October 1979, a monetary regime change is associated with a significant shift in the stochastic process of real interest rates. This evidence adds further support to the view that the recent changes in the

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[38] The reason the results are so similar for all three assets is that there is very little variation in the nominal interest rates for each asset, with the result that most of the variation in the ex post real interest rates is attributable to changes in inflation.

[39] An F-test using the ratio of the estimated variances of the two sample periods reveals that that the variance of the residuals is significantly higher in 1916:1-1920:6 than in 1920:7-1927:12. Thus, we also conducted the stability tests correcting for heteroscedasticity. For 4-6 month commercial paper, F(4,136) = 12.54; for 90 day stock exchange time loans, F(4,136) = 11.58, for stock exchange call loans, F(4,136) = 11.73. These results lead to the same conclusion as those presented in Table 8.

[40] To check for the possibility of a second breakpoint, we used the Quandt procedure to locate the most likely pair of breakpoints in the 1916-27 sample period. The result was June 1920 and March 1922. When conducting the standard F-test for coefficient equality before and after the second breakpoint, we failed to find a significant shift despite using this search technique which maximizes the probability of finding a significant shift in the real rate process.

Table 9

Dating of Breakpoint in 1916-27 Period

Real Rate Regression for	Minus Twice the Log of Quandt Likelihood Ratio												
	Break Point												
	1919:12	1920:1	1920:2	1920:3	1920:4	1920:5	1920:6	1920:7	1920:8	1920:9	1920:10	1920:11	1920:12
4-6 month commercial paper	32.20	33.24	33.18	33.56	47.04	50.02	58.46	52.86	44.88	43.54	43.52	42.82	40.26
90-day Stock Exchange Time Loans	29.90	30.82	30.64	30.90	43.90	46.76	54.44	48.84	40.40	39.02	39.12	37.84	34.92
Stock Exchange Call Loans	29.70	30.90	30.76	31.58	45.78	48.58	54.80	49.50	40.84	38.48	37.30	35.64	32.94

monetary regime have been an important source of the unusual behavior of real rates.

We now proceed to examine the way in which the stochastic process of real rates changes at the time of the 1920 regime shift. Is the nature of the changes in the real rate process at the time of the regime shift similar to that found recently? We first examine this question by generating estimates of ex ante real interest rates using fitted values from the separate ex post real rate regressions for the 1916:1-1920:6 and 1920:7-1927:12 periods.[41] Because the estimates are so similar for all the assets, we only report those for stock exchange call loans in Figure 2.

We see that estimated ex ante real rates during the inflationary period prior to June 1920 are persistently negative, just as they were prior to October 1979. After the June 1920 monetary regime shift, they climb to extremely high levels, sometimes exceeding 25%, and then settle down to around the 4 1/2% level after mid 1922 when the deflation is

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[41] The stock exchange call loan regressions for the two sample periods are as follows

1916:01 to 1920:06

$$eprr_t = -\underset{(.00365)}{.00876} - \underset{(.8401)}{.1507}\, i_t - \underset{(.1400)}{.1190}\, \pi_{t-1} + \underset{(.1412)}{.1388}\, \pi_{t-2}$$

Standard Error = .0101 R^2 = .0301 Durbin-Watson = 2.04

1920:07 to 1927:12

$$eprr_t = -\underset{(.00300)}{.01151} + \underset{(.7718)}{4.4196}\, i_t - \underset{(.1014)}{.2524}\, \pi_{t-1} + \underset{(.0950)}{.0597}\, \pi_{t-2}$$

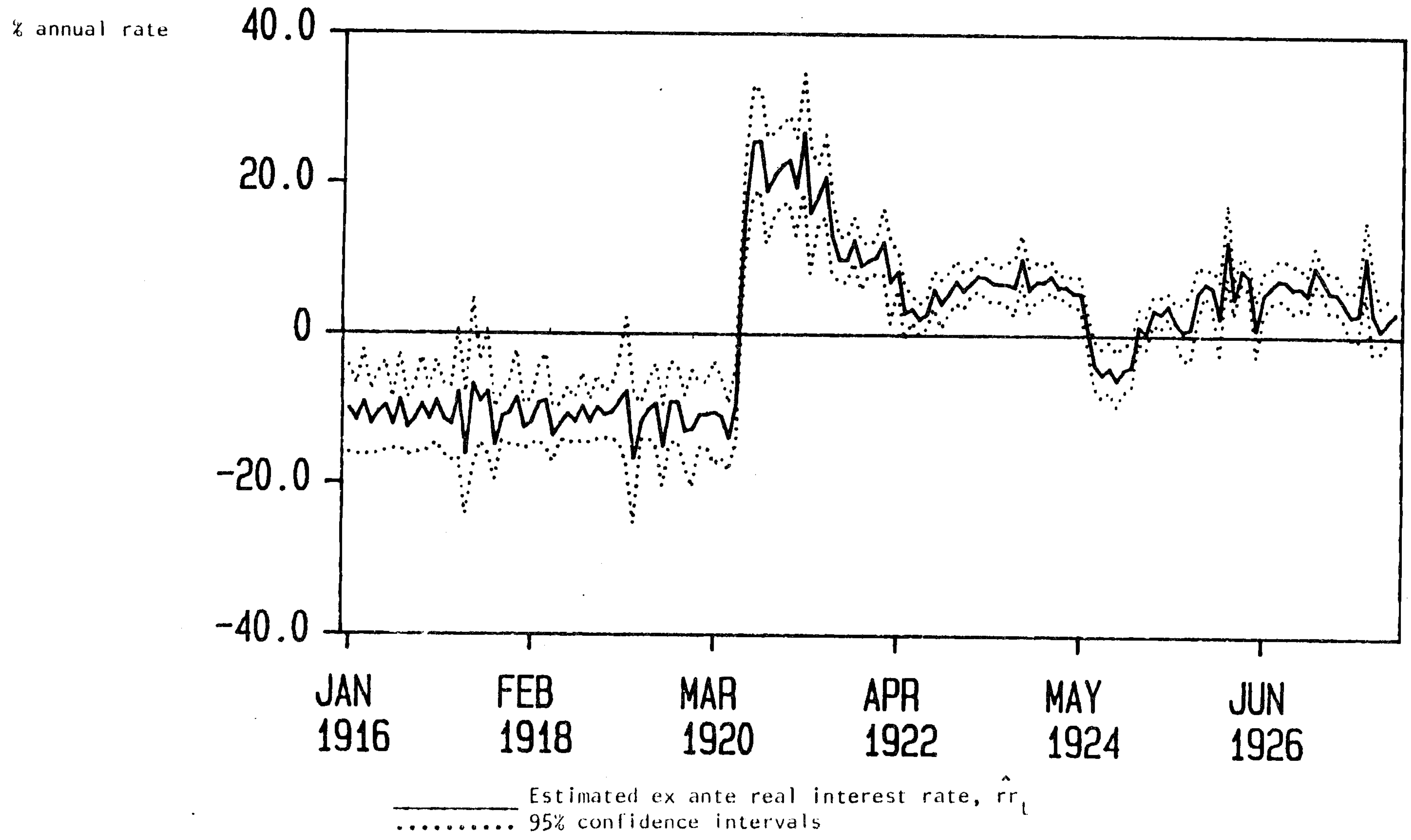
Standard Error = .007019 R^2 = .4020 Durbin-Watson = 2.04

Note that the coefficient of the constant term has not been multiplied by 100 as in Table 10.

[42] As was the case in Figure 1, the sharp rise in the estimated real

Figure 2

Estimated Ex Ante Real Interest Rates
for Stock Exchange Call Loan, 1916-27



over and price stability is maintained.[42] The variability of the ex ante real rates also appears to be higher after June 1920 than before: the standard deviation of the estimated ex ante real rate is 2.1% in 1916:1-1920:6 and is 6.8% in 1920:7-1927:12. Even after price stability has been achieved, the variability of ex ante real rates continues to be higher than before June 1920: the standard deviation of the estimated ex ante real rate is 4.0%.

Although the rise in ex ante real rates is far greater after the 1920 regime shift than after October 1979, the general pattern we saw in the post-October 1979 period seems to be repeated.[43] After the regime shift in both periods, ex ante real rates climb to a temporarily high level and then settle down to a level that is permanently higher than in the pre-regime shift period. Although the securities used in Figures 1 and 2 are not strictly comparable, adjusting for these differences would probably not change the conclusion that the ex ante real interest rates we see currently are at a similar level to those found after 1922.

Further similarities between the behavior of real rates after the monetary regime shifts of June 1920 and October 1979 are evident in the regression results of Table 10. As we found for the entire post World War II period, ex ante real rates are negatively correlated with expected inflation, usually significantly so. We also again find that during the inflationary period before the regime shift, the correlation between nominal interest rates and ex ante real rates is negative, while afterwards the correlation becomes positive. Furthermore, during the

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rate at a breakpoint is likely to be an artifact of splitting the sample for the ex post real rate regressions at 1920:6.

[43] One possible reason why real rates climb to much higher levels after June 1920 than after October 1979 is that the disinflation after June 1920 was far more rapid than the recent disinflation.

Table 10

The Relationship of Ex Ante Real Interest Rates with Nominal Interest Rates and Expected Inflation: 1916-1927

Ex Post Real Rate Regression for	1916:1 to 1920:6 Coefficients of Constant term[a]	π_t^e	i_t	SER	DW	1920:7 to 1927:12 Coefficients of Constant term[a]	π_t^e	i_t	SER	DW
4-6 month commerical paper	.12 (.17)	-.76** (.13)		.0099	1.94	.37** (.02)	-1.21** (.03)		.0068	2.17
	-.34 (.75)		-1.24 (1.69)	.0101	1.78	-1.54** (.28)		5.27** (.67)	.0069	1.68
Stock Exchange Time Loans	-.03 (.25)	-.65** (.19)		.099	1.93	.37** (.02)	-1.21** (.03)		.0070	2.14
	-.66 (.50)		-.50 (1.12)	.0101	1.79	-1.48** (.31)		5.11** (.72)	.0072	1.58
Stock Exchange Call Loans	-.27 (.36)	-.49 (.27)		.0100	1.92	.34** (.02)	-1.21** (.03)		.0070	2.14
	-.88 (.34)		-.09 (.79)	.0101	1.81	-1.36** (.29)		5.07** (.72)	.0072	1.55

standard errors in parentheses
* = significant at the 5% level
** = significant at the 1% level
SER = standard error of the regression
DW = Durbin-Watson Statistic

[a] Coefficients and their standard errors have been multiplied by 100.

inflationary period we find a strong Fisher effect -- a positive correlation between nominal interest rates and our measure of expected inflation equal to .74; while after June 1920, the correlation between nominal interest rates and estimated expected inflation even becomes strongly negative, equaling -.92.

Our examination of the 1916-27 period reveals that the surprising behavior of ex ante real rates that we have been experiencing in the last ten years has historical precedent. In both periods, a central feature of the changes in monetary policy regime is that they are followed by a substantial disinflation. Our evidence that both regime shifts are closely associated with changes in the stochastic process of real interest rates suggests that disinflation may well be an important characteristic of the regime changes affecting ex ante real interest rates.

VII. Conclusions

This paper has examined evidence that changes in monetary policy regimes are an important factor in explaining the recent unusual be-

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[44] Another possible factor often cited as a source of the current high real interest rates are the current and prospective large budget deficits. Although evidence analyzing the link between budget deficits and real interest rates does not support a strong connection [Blanchard and Summers (1984), and Evans (1985)], the perspective in this paper suggests that the possible link between deficits and real rates might be better understood by concentrating on a regime shift in the budgetary process. Evidence in Eisner and Pieper (1984) indicates that when the high employment budget surplus as a percentage of GNP is appropriately corrected for price and interest effects, it is rarely negative in the postwar period. Beginning in 1982, it swings sharply into the negative range and can be expected to remain there for some time in the future. This may indicate that there has been a shift in the budgetary regime which might be linked to recent behavior of real interest rates. However, Barro (1984)

havior of ex ante real interest rates.[44] We find that not only are there significant shifts in the stochastic process of real interest rates in October 1979 and October 1982 when the Federal Reserve alters its behavior, but these dates are also found to be the most likely breakpoints in the real rate process.

Although these results point the finger at monetary regime changes as a factor explaining the unusual behavior of ex ante real rates, we run the danger of fitting one historical episode with one tailor-made theory. Truly convincing evidence that monetary policy is important must rely on examining similar "controlled experiments" in other time periods. If we examine other time periods with similar monetary policy regime changes and we fail to find a correspondence between shifts in the stochastic process of ex ante real rates and the regime changes, then the results from the recent experience become less convincing. In contrast, if we do find such a correspondence, we have more confidence that monetary policy regime shifts have played an important role in the recent real rate experience.

Our examination of another time period, that surrounding the 1920 monetary regime shift when the discount rate was raised sharply, produces a striking correspondence between the regime change and the shift in the real rate process. The shift away from an inflationary monetary policy regime is associated with a change from a negative correlation between nominal and real interest rates to a positive correlation. In addition, the Fisher effect significantly diminishes with the change in monetary regime. Finally, the shift away from the inflationary monetary regime is associated with a sharp rise in ex ante real

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presents evidence that no regime shift for fiscal policy has occurred.

rates, while after the disinflation is complete, ex ante real rates remain at a high level.

In this paper, we have examined several pieces of evidence to provide an explanation of recent real rate behavior. Although each piece of evidence by itself may not be compelling, we find that taken together, there is strong support for the view that the recent shift in real rate behavior is a monetary phenomenon. Although our evidence does not rule out affects from such factors as the high budget deficits or financial deregulation in the 1980s, it does suggest that monetary factors are more important since high budget deficits or financial deregulation were not a feature of the 1920s, a period which displays real rate behavior similar to that in recent years.[45]

Our findings also provide a different perspective on the recent behavior of real interest rates. When many economists discuss the recent behavior of real interest rates, they describe it as unusual. An important conclusion that comes out of our analysis is that what appears to be unusual behavior of real interest rates is only unusual from the perspective of the post World War II U.S experience. In a wider historical context the recent behavior of real interest rates is not unusual at all.

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[45] Although the federal government ran substantial budget deficits in the years 1917-1919 as a result of World War I, there were budget surpluses in every year from 1920 to 1927. (See the federal budget series in the Historical Statistics of the United States, page 1104.)

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