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THE REACTION OF REDUCED-FORM COEFFICIENTS TO REGIME CHANGES: THE CASE OF INTEREST RATES

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ABSTRACT

This study investigates whether the apparent intertemporal instability of a particular reduced-form equation (that for interest rates) can be explained by changing government policy parameters, or regimes, and otherwise stable structural parameters. We hypothesize that major fiscal, monetary, and regulatory policy parameter shifts have been important sources of that instability. Direct tests imply that reduced-form coefficients move by statistically significant and economically meaningful amounts in response to policy parameter change. Allowing for this systematic parameter variation produces greater stability in the remaining parameters. Furthermore, in-sample and out-of-sample forecasts from the proposed model outperform those from the non-responsive parameter specification.

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A. Introduction

For more than a decade, empirical studies have sought to determine whether nominal interest rates adjust such that real rates are unaffected by changes in the anticipated inflation rate (the Fisher neutrality hypothesis). An unsettling aspect of these studies is the volatility of the estimated interest rate response to anticipated inflation over various sample periods.[1] Estimates based on data for the 1950s provided low and insignificant values for the response.[2] As the sample period was extended, larger point estimates of the response were obtained, finally approaching unity. Carlson (1979), Cargill and Meyer (1980) and Levi and Makin (1979) produce estimated interest rate responses that decline, often dramatically, when the sample period is extended to include the first half of the 1970s. Peek and Wilcox (1983) find that the inclusion of income tax and aggregate supply shock effects reduces, but does not eliminate, the observed coefficient instability. This suggests other relevant factors remain.

The Lucas (1976) critique suggests that conventional reduced-form coefficients may vary over time due to the dependence of private sector expectational parameters on government policy parameters. Sims (1982) has recently countered that this objection should be regarded as no more than a "cautionary footnote" (p. 108) since policy rules "have not changed frequently or by large amounts" (p. 138). He argues that in fact there has been little drift in (final form) parameter estimates through time. Here we test directly whether changes in policy parameters have produced changes in reduced-form parameters. We hypothesize that significant, quantifiable changes have occurred in the policy parameters that are especially relevant to the reduced form for interest rates. We incorporate these para-meters in our model and investigate whether the apparent intertemporal instabil-ity of the reduced-form equation for interest rates can be explained by the changing values of government policy parameters through time and otherwise stable structural parameters.

We address three major sources of change in the government's policy parameters:[3]

1) changes in fiscal policy parameters,

2) changes in monetary policy parameters, and

3) changes in financial regulatory policy parameters.

The first category is exemplified by changes in personal tax rates. Peek (1982) presents evidence that changing tax rates significantly affect interest rates and that incorporating their movements substantially reduces the instability of interest rate coefficients. Furthermore, Peek and Wilcox (1984) find that such tax effects are complete; that is, there is no evidence of the "fiscal illusion" suggested by Tanzi (1980). The second category of policy parameter change is typified by the October 1979 change in monetary policy (as well as by the 1951 Treasury Accord). The creation of negotiable certificates of deposit (CDs) in the early 1960s and of six-month money-market certificates in the late 1970s and the elimination of interest rate ceilings on large CDs in the early 1970s exemplify the regulatory changes most directly relevant to financial markets. By reducing the degree of disintermediation when rates rise, these financial innovations may have reduced the impact of monetary restraint through credit availability and thus may have decreased the interest elasticity of private expenditures. To the extent each of these policies influence structural parameters, the reduced-form response of nominal interest rates to anticipated inflation (and to other factors) will vary with regime changes.

Below we present a simple macro model which highlights the link between policy parameters and reduced-form coefficients. Sections C and D describe the estimation and measurement methodology and present our empirical results. The final section concludes.

B. A Model of Interest Rates

This section presents a model of interest rates that embodies both fiscal and monetary policy parameters. Solution of the model produces reduced forms that highlight the link between policy parameter variation and movements in the reduced-form coefficients. The model consists of IS, LM, wage, aggregate supply, and monetary policy rule equations. These five relationships can be expressed in linearized form as

$$Y - Y^N = a_0 - a_1 r^* + a_2 \Delta Y + a_3(X - Y^N) - a_4 LIQ - a_5 SS \qquad (1)$$

$$M - P - Y^N = b_0 + b_1(Y - Y^N) - b_2 i^* + b_3 \sigma \qquad (2)$$

$$W = c_0 + P^e - c_1 SS \qquad (3)$$

$$P = d_o + W + d_1(Y - Y^N) + d_2 SS, \qquad (4)$$

$$M = M_x + P^e + Y^N + e_1 r \qquad (5)$$

where the coefficients of all the variables are assumed to be positive and:

Y = the logarithm of actual real output,

Y^N = the logarithm of natural real output,

ΔY = the percentage change in real output lagged one period,

X = the logarithm of the sum of real exports and real government expenditures,

LIQ = the current growth rate of the nominal money supply relative to its recent growth trend,

M = the logarithm of the nominal money supply,

M_x = the logarithm of the non-interest-rate-reactive component of the real money supply,

P = the logarithm of the actual price level,

P^e = the logarithm of the expected price level,

W = the logarithm of the nominal wage,

SS = the supply shock variable,

σ = the standard deviation of the after-tax nominal interest rate,

r = the real interest rate,

r^* = the after-tax real interest rate,

i^* = the after-tax nominal interest rate.

The two after-tax interest rates are related to the nominal interest rate (i) by (6) and (7):

$$i^* \equiv i(1 - t) \tag{6}$$

$$r^* \equiv i^* - p^e \tag{7}$$

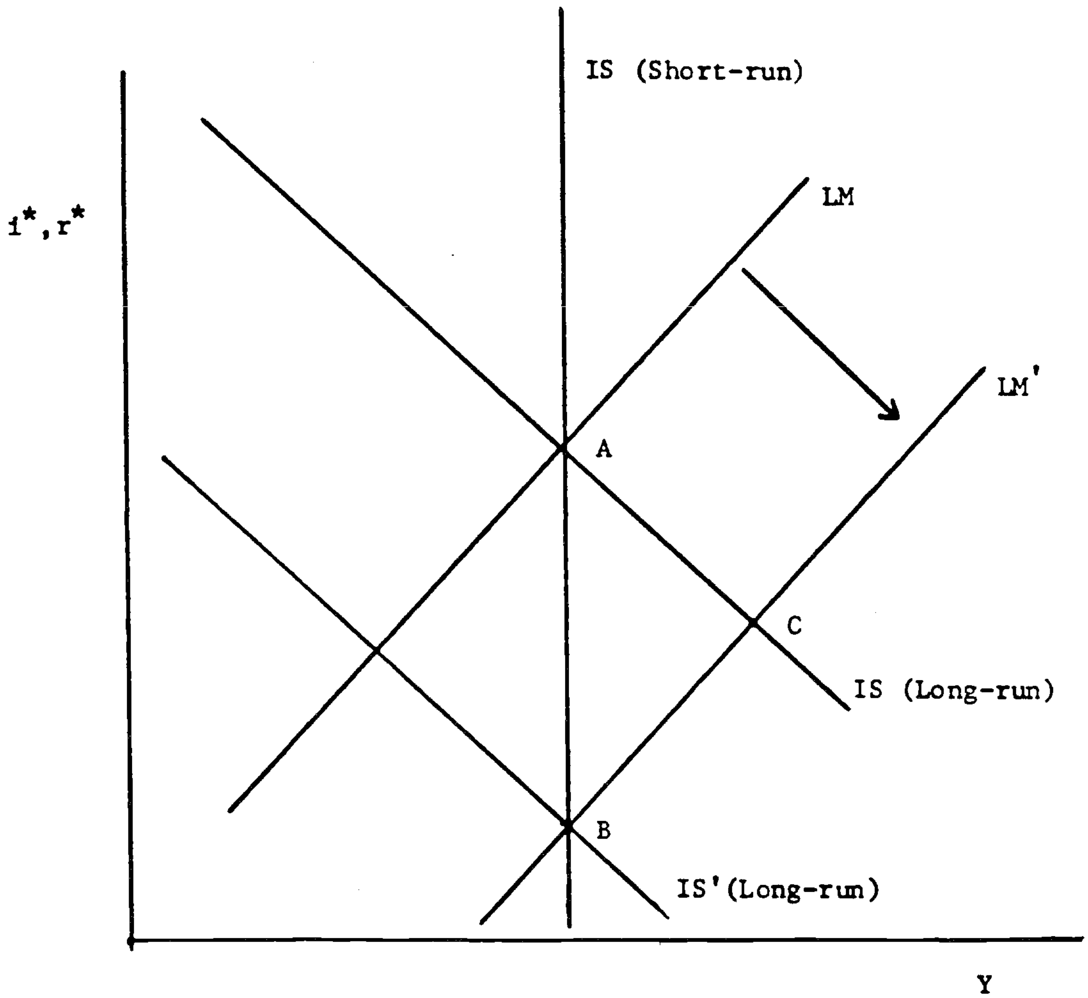
where t is the marginal tax rate on interest income and p^e is the anticipated inflation rate.

Real expenditures depend on the real after-tax interest rate, an investment accelerator term, exogenous real export and real government demand, and real shocks emanating from the supply side. In addition, the presence of the liquidity variable (LIQ) allows us to capture the difference between short-run and long-run IS curves. This difference follows from the assumed differential adjustment speeds in real and financial markets. Since real variables (such as output) adjust more slowly than financial variables (such as interest rates), we hypothesize a steeper (e.g., vertical) short-run IS curve. The liquidity term represents accelerations (or decelerations) in nominal money growth relative to normal. Given an acceleration in nominal money growth, the LM curve would shift to the right and the economy would move from point A to point B along the vertical short-run IS curve in Figure 1. Movement down the short-run IS curve is captured by a temporary downward shift of the flatter long-run IS curve to IS'. Thus, this term allows us to capture the well known liquidity effect. If this higher growth rate persists, LIQ returns to its long-run value of zero and the IS curve returns to its original position.

Money demand is hypothesized to depend on output and on the after-tax nominal interest rate, which represents the opportunity cost of holding money when interest income is taxed. The third argument in the money demand function (σ) represents a measure of the capital-value risk associated with holding bonds as alternatives to money in wealth portfolios (see Slovin and Sushka (1983) for discussion and empirical evidence in favor of this hypothesis). The wage and price equations embody the natural rate hypothesis.

Equation (5) posits a monetary authority (Fed) policy whereby the money supply rises and falls with the expected price level, natural real output, and the real interest rate. The Fed almost certainly has reacted to other factors as well, e.g., cyclical unemployment, inflation, international forces, and the preferences of individual policymakers. Shifts in the slope parameter e_1 in (5), however, can be viewed as capturing some of the major policy shifts of the postwar period. That parameter measures the extent to which the Fed stabilizes interest rates in practice, i.e., accommodates. This parameter can be thought of as a measure of the degree to which the nominal money supply (given P^e and

FIGURE 1



Y^N) moves in response to fluctuations in the real interest rate. The reduced pegging of interest rates after the 1951 Treasury Accord, the increased emphasis on monetary aggregates in the 1970s, and the October 1979 shift to reserves targeting can each be represented by changes in the policy parameter e_1. Since each of these changes involved moving toward a less procyclical monetary policy (and a steeper effective LM curve), each can be characterized as a reduction in e_1.

Equations (1-7) can be combined to yield the reduced-form equation for the nominal interest rate:

$$i = \beta_0 + \underset{(+)}{\beta_1 p^e} + \underset{(+)}{\beta_2 \Delta Y} + \underset{(+)}{\beta_3 X'} + \underset{(-)}{\beta_4 LIQ} + \underset{(-)}{\beta_5 M_x} + \underset{(?)}{\beta_6 SS} + \underset{(+)}{\beta_7 \sigma} \tag{8}$$

where X' represents $(X - Y^N)$ and:

$$\beta_0 = \frac{a_0(b_1+d_1) + b_0 + c_0 + d_0}{D} \tag{9}$$

$$\beta_1 = \frac{a_1(b_1+d_1) + e_1}{D}\ , \tag{10}$$

$$\beta_2 = \frac{a_2(b_1+d_1)}{D}, \tag{11}$$

$$\beta_3 = \frac{a_3(b_1+d_1)}{D}, \tag{12}$$

$$\beta_4 = \frac{-a_4(b_1 + d_1)}{D}\ , \tag{13}$$

$$\beta_5 = \frac{-1}{D}\ , \tag{14}$$

$$\beta_6 = \frac{(d_2-c_1) - a_5(b_1+d_1)}{D}, \tag{15}$$

$$\beta_7 = \frac{b_3}{D}, \quad \text{and} \tag{16}$$

$$D = (1-t)[a_1(b_1+d_1) + b_2] + e_1. \tag{17}$$

The sign of β_6 is indeterminate a priori. An adverse supply shock reduces investment and real wages and thus the interest rate, while at the same time increasing input costs operating through the aggregate supply equation which raises the interest rate. The investment-real wage effect might be expected to dominate, suggesting a negative value for β_6. The results presented in Peek and Wilcox (1983) and Wilcox (1983a, 1983b) support this interpretation.

The three policy parameters t, e_1, and a_1 are of particular interest. To the extent that any of these parameters (or for that matter, any of the structural parameters) vary over time, the reduced-form coefficients will also change. Insofar as the structural parameter in question enters more than one reduced-form coefficient, the β's will covary deterministically. For example, the marginal tax rate (t) enters all of the reduced-form coefficients in the same way. An increase in the tax rate will raise not only the interest rate response to expected inflation, but also all of the other reduced-form coefficients. A decrease in the response of private expenditures to the real after-tax interest rate (a_1) will reduce the denominators of all eight β's by the same amount. However, since a_1 also appears in the numerator of β_1, the interest rate response to the expected inflation rate will be differentially affected; the decrease in a_1 will raise all of the other reduced-form coefficients, while the response to expected inflation will be reduced. Similarly, a decrease in e_1 will raise all of the β's except β_1 through its effect on D. But, as with

a_1, e_1 also enters the numerator of β_1. In this instance, however, the direction of the net effect of the change in e_1 on β_1 is ambiguous a priori.

C. Methodology

Equations (9)-(17) illustrate the relationship between policy parameters and the reduced-form coefficients. Our hypothesis is that failure to allow for movements over time in these policy parameters has contributed to observed reduced-form estimate instability. We rectify this shortcoming by including values of the time series of the proxies for fiscal, monetary, and regulatory policy parameters. This allows us to test directly for the significance of policy changes in explaining reduced-form coefficient variability and to evaluate whether the remaining, deeper parameters are stable. Incorporating fiscal policy parameter movements requires a measure of the marginal tax rate of the marginal investor, t. If a tax-exempt institution is the marginal investor, the marginal tax rate is zero. If individuals are the marginal investors, the appropriate tax rate is the marginal personal income tax rate. The progressivity of the personal income tax rates makes measuring that rate problematic. As our measure of t, we use the average marginal tax rate on interest income constructed from data contained in annual editions of Statistics of Income, Individual Income Tax Returns (see Peek (1982)). The tax rate is calculated as a weighted average of the marginal personal income tax rate for each adjusted gross income class. The weight for each class is equal to its share of the total interest received by all income classes.[4]

A downward drift in the measured interest rate response to expected inflation as the sample period was extended into the 1970s has been noted by Carlson (1979), Cargill and Meyer (1980), and Levi and Makin (1979). This may be due to the continuing financial institution deregulation and consequent reduction in disintermediation. A number of regulatory changes reduced the potential for disintermediation throughout that period (e.g., the creation of negotiable certificates of deposit, increases in Regulation Q ceilings, the introduction of six-month money market certificates). Such changes might reduce the interest response of private expenditures (a_1), thereby lowering the reduced-form coefficient on expected inflation (see equation (10)).[5]

We allow for changing regulatory policy (and innovation) with a measure of the effect of such changes rather than attempting to quantify the changes themselves. To do so, we take as our regulatory policy indicator, s, the share of commercial banks' and thrift institutions' liabilities that pay market-related interest rates. We assume that the interest response of expenditures is a function of s and DCC:

$$a_1 = f_o + (f_1 + f_2/s)DCC, \tag{18}$$

where DCC is a dummy variable that takes a value of unity when the three-month Treasury bill yield exceeds the regulation Q ceiling interest rate on savings deposits and is zero otherwise. Thus, DCC switches on when disintermediation is likely. The extent to which a_1 changes then depends upon the share of liabilities subject to disintermediation. We also allow the IS curve intercept (a_o) to move during these periods so as not to constrain the IS function to pivot about its horizontal intercept:

$$a_o = f_3 + f_4DCC. \tag{19}$$

In this specification, $f_o(>0)$ represents the (absolute value of the) interest response of expenditures in the non-disintermediation periods. During the disintermediation periods, the value of a_1 will increase. Thus, we expect

$(f_1 + f_2/s)$ to be positive. Furthermore, we expect the increase in a_1 to be larger the smaller is the value of s ($f_2 > 0$). We also expect f_4 to be positive in (19).

Similarly, we seek a measure of the time series for the money supply policy parameter, e_1. We obtain an estimate of this series from the following regression estimated with semi-annual data over our full 1952:06-1983:06 sample (t-statistics in parentheses):

$$\begin{aligned} M-P^e-Y^N = {} & \underset{(14.7)}{-1.082WMM} - \underset{(13.5)}{1.086AFB} - \underset{(11.8)}{1.121GWM} - \underset{(12.1)}{1.149PV} + \underset{(0.99)}{0.35WMMRE} \\ & + \underset{(2.66)}{1.38AFBRE} + \underset{(1.33)}{1.12GWMRE} - \underset{(0.04)}{0.012PVRE} - \underset{(2.08)}{0.044D8006} - \underset{(10.8)}{0.016T} + \underset{(16.1)}{0.90u_{-1}} \end{aligned} \tag{20}$$

$$S.E.E. = 0.017 \qquad D.W. = 1.17$$

WMM, AFB, GWM, and PV are dummy variables that take a value of one during the terms of the sample's Fed Chairmen Martin, Burns, Miller, and Volcker, respectively. The same variables with the suffix RE are those dummies multiplied by the expected real interest rate, $i - p^e$. D8006 is a dummy variable that takes the value one during the credit control period in 1980. A linear time trend is included to pick up the long-term increase in velocity. The last term in (20) indicates a first-order autocorrelation correction coefficient estimate of 0.90.

This specification allows e_1, the reaction of the money supply to expected real interest rates, to vary across the regimes of the different Fed chairmen, but restricts it to be constant within regimes. The four intercept dummies are included to lessen the likelihood that variations in the overall stringency of monetary policy across regimes be mistakenly attributed to variations in the systematic-response coefficient, i.e., to avoid empirically confusing intercept and slope shifts in the money supply function. The estimated time series for the money supply reaction coefficient can be read directly from (20). The values for e_1 are 0.35, 1.38, 1.12, and -0.012 for 1952:06-1970:06, 1970:12-1978:06, 1978:12-1979:06, and 1979:12-1983:06, respectively. This series suggests that the Volcker regime has accommodated real rate shocks least while the Burns and Miller years saw the most accommodation. The William McChesney Martin regime appears to have been in between. F-tests allow us to reject the hypothesis that the money supply reaction coefficient, e_1, was the same across Fed chairmen.[6] M_x is also based on (20). M_x is $M-P^e-Y^N$ minus the estimated trend, credit control and real-rate-reaction elements. Movements in M_x, the intercept dummies plus the error term, u, imply that monetary policy loosened from the latter 1950s until the early 1970s and has been tighter than average since the mid-1970s, especially during the Volcker years.

D. Empirical Results

1. Estimates Based on Constant Policy Parameters

This section presents the results of estimating (8) subject to (9)-(19). When e_1, f_1, f_2, and f_4 are taken to be zero through time, constant-coefficient, ordinary least squares (OLS) suffices. These restrictions are equivalent to setting M_x to equal $(M-P^e-Y^N)$, e_1 equal to zero, and a_1 equal to a constant. As a result, (8) can be expressed as:

$$i = \beta_0 + \beta_1 P^e + \beta_2 \Delta Y + \beta_3 X' + \beta_4 LIQ + \beta_5 M' + \beta_6 SS + \beta_7 \sigma \qquad (21)$$

where M' is $(M-P^e-Y^N)$,

$$\beta_1 = \frac{a_1(b_1 + d_1)}{D}, \text{ and} \qquad (22)$$

$$D = (1-t)(a_1(b_1 + d_1) + b_2). \qquad (23)$$

Tanzi (1980) suggests that individuals have suffered from "fiscal illusion" by failing to take complete account of tax rates. Peek (1982), using a 1960-1979 sample, can be interpreted as testing the hypothesis that individuals rather than tax-exempt institutions are the marginal investors in the six-month and one-year Treasury bill market. Alternatively, the null hypothesis could be that individuals are the marginal investors but ignore income tax considerations in making their financial decisions. Using the Davidson-McKinnon (1981) non-nested model specification test, the non-tax adjusted model (equivalent to our equations (21)-(23) with $t = 0$) was rejected in favor of the tax-adjusted formulation (equations (21)-(23)) using an index of marginal personal income tax rates. Employing a more detailed model incorporating supply shock and foreign-held bond effects, Peek and Wilcox (1983) reconfirmed these results for the entire 1952-79 period for the one-year Treasury bill rate. We also showed how changes over time in the correlations between the anticipated inflation rate and the tax rate and supply shock variable contributed to previously measured intertemporal instability in the estimated expected inflation coefficients. Further, Peek and Wilcox (1984) estimate a specification similar to (21)-(23) with $(1-t)$ replaced by $(1-\theta t)$, where θ reflects the degree of (lack of) fiscal illusion. Using nonlinear least squares, the estimate of θ closely approximates one, indicating no fiscal illusion. Therefore, we here restrict θ to unity, implying complete adjustment to changes in tax rate policies. From (23), it can be seen that $(1-t)$ can be factored out of the coefficient of each explanatory variable. Using our tax rate series, we can express (21) with constant reduced-form coefficients when we divide each of the right-hand-side variables by $(1-t)$. The implied reduced-form coefficients at any time are then the estimated constant coefficients times the value of $(1-t)$ for that period.

Table 1 presents the results of estimating (21). June and December averages of the secondary market yield (on a bond equivalent basis) on one-year U.S. Treasury bills are used as the dependent variable. $\dot{p}^e$ is the Livingston one-year expected inflation rate, recorded in June and December. This measure of expected inflation has the advantages of being truly ex ante and of embodying whatever sophistication agents actually use to form their expectations.[7] The remaining variables are measured with second and fourth quarter data (except σ). M1 is the nominal money supply. P^e is the price level expected six months ahead from the Livingston survey data. Y^N, natural real output, is from the Council of Economic Advisors. $(M-P^e-Y^N)$ has been detrended by regressing its log on a linear time trend and using the residual as M'. LIQ is the difference between the annualized growth rate of M1 measured from the current to the previous quarter and its four-quarter growth rate up to that previous quarter. X' is the logarithm of the ratio of real government expenditures plus real exports to real natural output. ΔY is the four-quarter growth rate of real GNP up to the preceding quarter. SS is the ratio of the import deflator to the GNP deflator, adjusted for exchange rate changes. σ is the 18-month moving standard deviation of the after-tax nominal interest rate, lagged one month. D7983 is a dummy variable that takes the value one starting with the December 1979 observation. The June 1980 observation has been omitted due to the presence of credit controls; otherwise, the full sample is 1952:06-1983:06.

The estimates in row 1 imply that rises in expected inflation, exogenous spending, faster real growth, and more volatile interest rates raise rates while accelerations in money growth, higher real balances, and positive supply shocks each lower them. After 1979 interest rates were both surprisingly high and volatile. Row 2 shows that when the post-1979 period is added to the sample,

TABLE 1

OLS ESTIMATES OF EQUATION (21)
dependent variable: nominal yield on 1 year Treasury bill
semi-annual observations
(t-statistics in parentheses)

Sample Period	Constant	p^e	ΔY	X'	LIQ	M'	SS	σ	D7983	R^2	DW	SEE	SSR
1.1952:06-1979:06	15.4 (6.20)	0.562 (14.73)	4.10 (1.31)	7.32 (4.27)	-0.101 (3.15)	-2.23 (0.84)	-3.39 (4.73)	0.25 (0.62)	---	.9141	1.41	0.715	24.06
2.1952:06-1983:06	13.1 (3.73)	0.577 (11.53)	2.93 (0.63)	6.58 (2.60)	-0.081 (1.88)	-7.29 (2.28)	-2.68 (2.60)	1.63 (3.80)	---	.8887	1.34	1.211	79.22
3.1952:06-1983:06	12.4 (4.48)	0.513 (12.45)	2.25 (0.61)	5.58 (2.77)	-0.082 (2.39)	-0.48 (0.17)	-2.76 (3.37)	0.47 (1.20)	4.21 (5.74)	.9313	1.58	0.960	48.86

the standard error of the estimate rises by 69 percent and the measure of interest rate variability gains in significance. This appears to confirm the popular attribution of the high post-1979 interest rates to sharply increased interest rate volatility. Including D7983 in row 3, however, reduces the effect of volatility to insignificance. The coefficient of 4.21 on D7983 in row 3 indicates that the surprises were large and primarily on the upside. However, even allowing for this nonexplained upward shift leaves the standard error of the estimate much larger than before 1979.

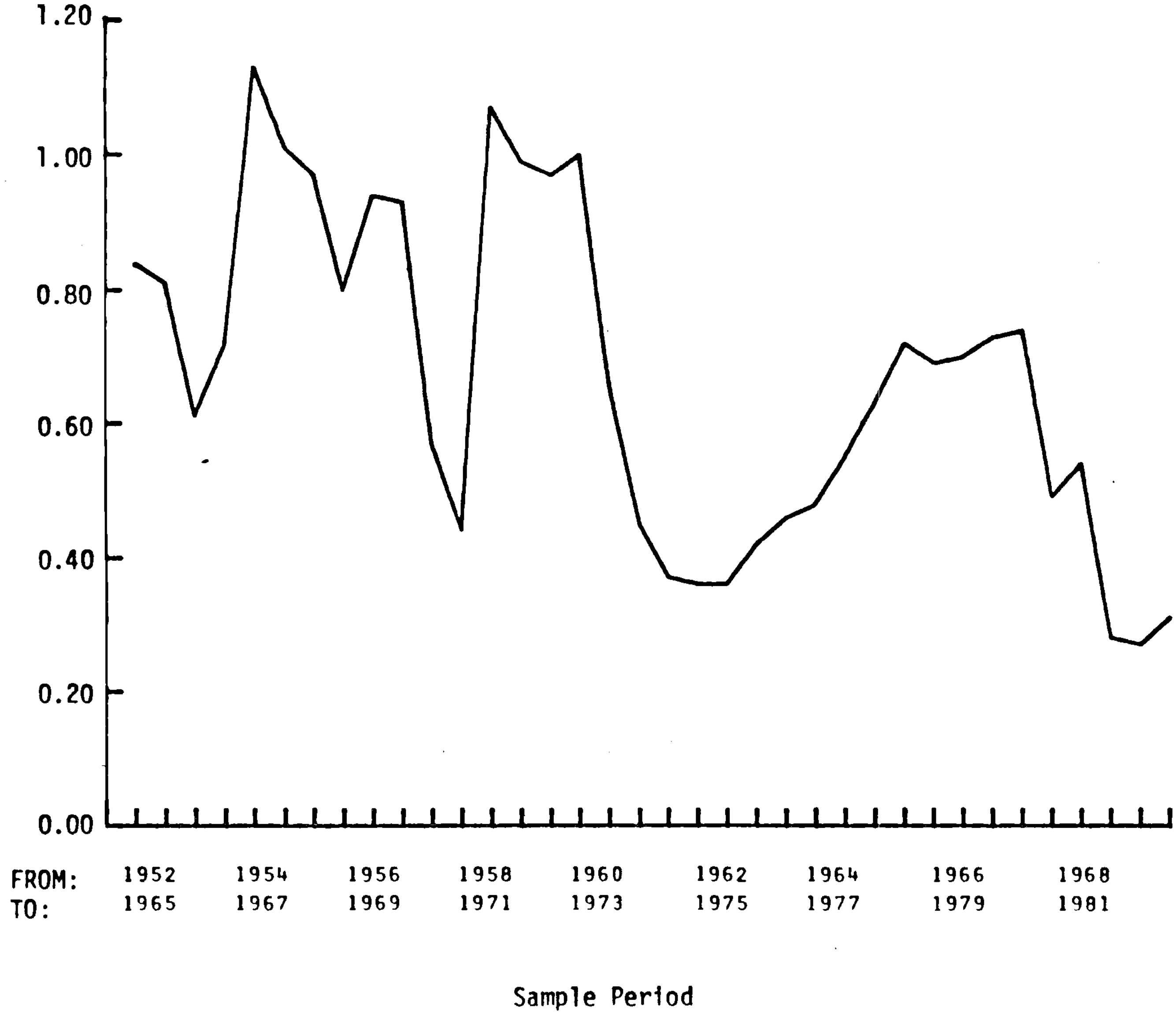
Adding the post-1979 period seems to change the estimated coefficients relatively little. When a time series of the expected inflation coefficient estimate is generated by rolling over a fourteen-year (n=28) sample using the specification of rows 1 and 2, however, a different picture emerges. Figure 2 plots this series. Although these coefficients abstract from changes due to changes in tax rates, they still exhibit considerable movement. In particular, the sharp jump in the series after the 1952 and 1953 observations are eliminated and the early 1970s observations are added to the rolling sample, the downward drift as the sample leaves the 1950s and enters the 1970s, the rise as the sample moves into the second half of the 1970s, and the resumption of the decline as the 1980s observations are included, suggests that major movements are left to be explained.

2. Estimates based on Changing Policy Parameters

To estimate (8), allowing for variations in monetary and regulatory policy as well as tax policy changes, we substitute (18) and (19) for a_1 and a_o. Later we wish to calculate the marginal significance levels of each explanatory variable. We require an estimate of the numerator of β_5 to obtain a marginal significance level on M_x. To obtain a unique set of parameter estimates when

FIGURE 2

Anticipated Inflation Coefficient From Rolling Regressions



that numerator is estimated, we choose to normalize the parameter estimates by setting the numerator of β_o to unity. We can then rewrite (9)-(17) as:

$$\beta_o = 1/D \tag{24}$$

$$\beta_1 = N/D \tag{25}$$

$$\beta_2 = g_2/D \tag{26}$$

$$\beta_3 = g_3/D \tag{27}$$

$$\beta_4 = g_4/D \tag{28}$$

$$\beta_5 = g_5/D \tag{29}$$

$$\beta_6 = g_6/D \tag{30}$$

$$\beta_7 = g_7/D \tag{31}$$

$$\beta_8 = g_8/D \tag{32}$$

$$N = h_o + h_1 DCC + h_2 DCC/s + h_3 e_1 \tag{33}$$

$$D = (1-t)(h_4 + h_1 DCC + h_2 DCC/s) + h_3 e_1 \tag{34}$$

where the h's and g's are constants.

Allowing a_o in (19) to change over time requires that we add DCC as an explanatory variable; its reduced-form coefficient is β_8. The coefficient g_9 applies to D7983 (not divided by D). Our model predicts that h_o, h_2, h_3, and h_4 are all positive. The sign of h_1 may be positive or negative. We anticipate positive values for g_2, g_3, g_7 and g_8, and negative values for g_4, g_5, and g_6.

The estimates for the 1952:06-1983:06 sample, omitting D7983, are:

$$i = (1 + Np^e + \underset{(0.26)}{0.367}\Delta Y + \underset{(1.85)}{1.36}X' - \underset{(1.85)}{0.024}LIQ - \underset{(2.08)}{2.78}M_x - \underset{(3.76)}{1.22}SS + \underset{(1.29)}{0.175}\sigma + \underset{(3.57)}{0.680}DCC)/D \tag{35}$$

$$N = \underset{(1.34)}{0.114} - \underset{(0.48)}{0.050}DCC + \underset{(2.99)}{0.00221}DCC/s + \underset{(4.99)}{0.245}e_1 \tag{36}$$

$$D = (1-t)(\underset{(3.11)}{0.310} - \underset{(0.48)}{0.050}DCC + \underset{(2.99)}{0.00221}DCC/s) + \underset{(4.99)}{0.245}e_1 \quad (37)$$

where the numbers in parentheses below the coefficient estimates can be interpreted approximately as t-statistics. They are calculated as the square root of the chi-square test statistic used to perform the likelihood ratio tests of the restriction that each of the relevant coefficients was zero.[8] The general pattern of signs and significance of the coefficients in (35) mirrors the OLS results of Table 1. The coefficients of particular concern here, however, are those associated with the regulatory and monetary policy parameters, s and e_1.

The estimates of their coefficients are both positive and decisively significant, indicating that the reduced form coefficients do in fact respond to regime changes. As e_1 falls, i.e., as monetary policy becomes less accommodative of real rate shocks, D falls. The response of interest rates to changes in the explanatory variables (except possibly for p^e) then rises, as the economy is effectively operating with a steeper LM curve. As s rises, i.e., as the share of liabilities which are unregulated rises, D falls, increasing all of the reduced-form coefficients except that on p^e. Due to the effect of s on the numerator of β_1, the increase in s (reduction in a_1) will reduce the reduced-form expected inflation coefficient. As fewer liabilities are regulated, market interest rate increases induce less disintermediation and less credit rationing. Less expenditure is deterred by given interest rate increases, generating an effectively steeper IS curve, as hypothesized.

Adding D7983 to this specification (not shown) permits us to test whether the reduced-form coefficient movement that we ascribe to policy shifts remains when a dummy variable for the later part of the sample is included. The coefficient estimate of 2.87 (t-statistic = 2.67) on D7983 does indicate that our model still leaves an important part of the recent interest rate story to be told. Our expanded specification has reduced the post-1979 dummy coefficient by one-third, however, and the coefficients associated with e_1 and s retain their significance.

The summary statistics listed in Tables 1 and 2 suggest that allowing for changing policy parameters improves the interest rate model. The standard error of the estimate falls 24 and 30 percent for the short and long samples, respectively, when we incorporate policy-induced parameter movement. Likelihood ratio tests more formally confirm the improvement. The specification that jointly restricts the a_o and a_1 parameters to be constant and e_1 to be zero in favor of the expanded model that posits their movement according to (18)-(20) is rejected at the 5 percent level. This is true for the shorter sample (1952:06-1979:06) as well as for the sample that includes the post-1979 period, either with or without the D7983 dummy. For the longer period, the data also decisively reject the hypotheses that the coefficients associated with monetary policy and regulatory policy individually were zero. Further, the expanded model is statistically superior in-sample to the model that includes D7983 but ignores regulatory and monetary policy changes. Thus, in addition to producing reduced-form coefficients that move as our theory hypothesizes, allowing for structural change in response to policy change produces a significantly improved specification.

Figure 3 plots the estimated time series values for β_1, the reduced-form coefficient for expected inflation implied by (36) and (37). The implied coefficient exhibits a slight downward drift until the mid-1960s, due to a small decline in the effective tax rate series. Tax schedule reductions in 1954 and 1964-65 and a slight cyclical response to economic slack in the late 1950s and

TABLE 2

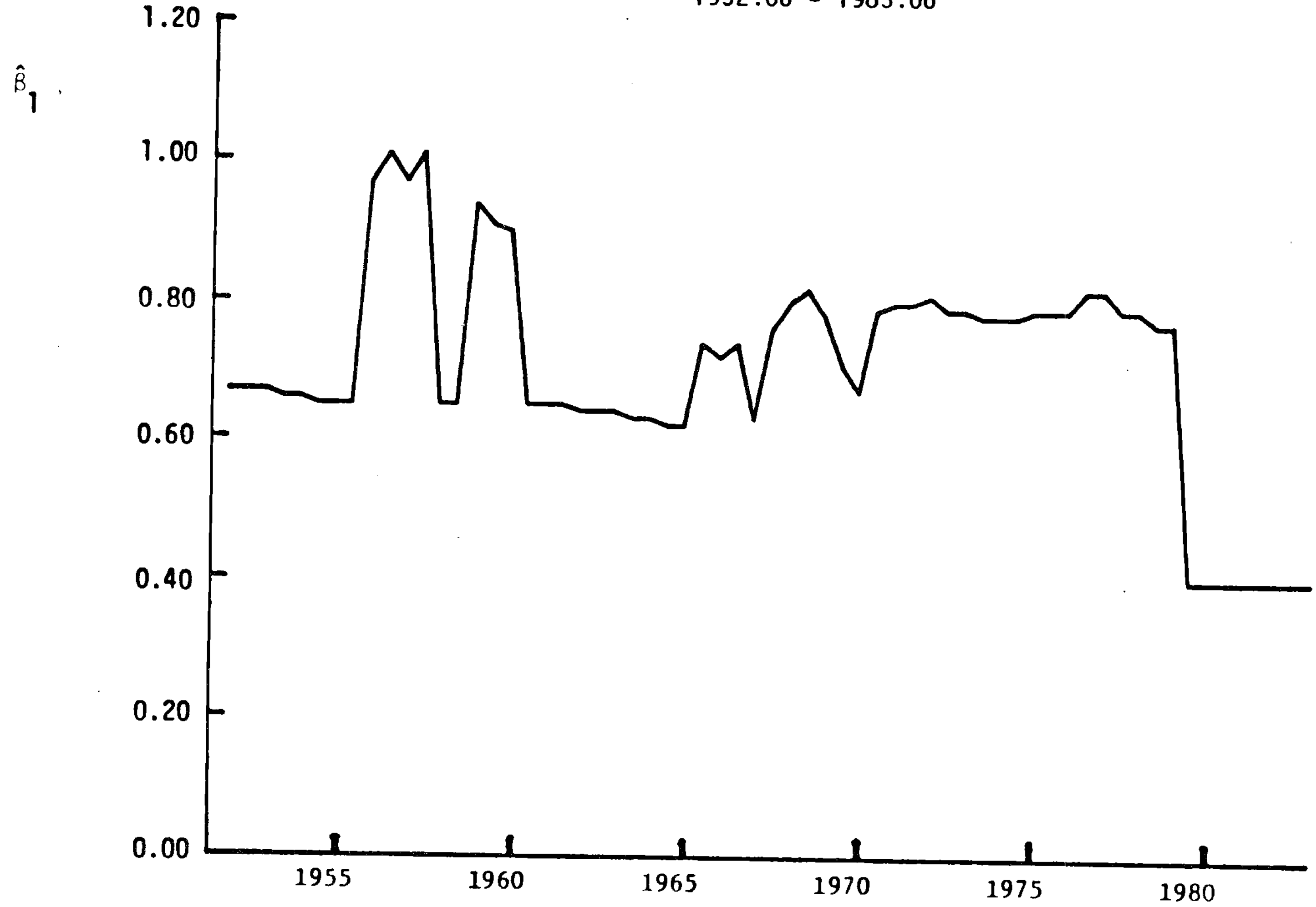
Summary Statistics for
Nonlinear Least Squares
Estimates of Equation (8)

	Sample Period	R^2	DW	SEE	SSR
1.	1952:06-1979:06	.9543	1.97	0.545	12.78
2.	1952:06-1983:06 (without D7983)	.9502	2.05	0.842	35.43
3.	1952:06-1983:06 (with D7983)	.9533	2.11	0.823	33.19

FIGURE 3

Implied Reduced-Form Coefficient for the Expected Inflation Rate (β_1)

1952:06 - 1983:06



early 1960s combined to reduce effective tax rates. After 1965, strong nominal income growth lifted the effective tax rate and, hence, β_1. β_1 rose sharply in 1970 with the higher e_1 of the Burns regime. The large fall-off in β_1 in 1979 is associated with the dramatic decline in e_1 attributed to the Volcker regime. The sharp upward spikes in β_1 reflect the disintermediation effects. During potential disintermediation periods (when DCC takes on a nonzero value), the interest sensitivity of expenditures (a_1) and β_1 rise. The magnitude of these increases is related to the share of financial institutions' liabilities that pay market-related interest rates (s). As s increased over time, the size of the spikes diminished. After the removal of interest ceilings on large CDs in the early 1970s, the spikes become almost indistinguishable.

Though the two series are not directly comparable, the pattern in Figure 3 broadly mirrors the rolling sample estimates in Figure 2 (with the exception of Figure 2's mid-sample spike). The implied time series of the other reduced-form coefficients are proportional to (1/D). Given the form of the model and the particular set of coefficient estimates, changes in a_1 and e_1 produce values for these remaining reduced-form coefficients that move inversely to β_1. Figure 3 then shows movements in β_1 and the <u>mirror image</u> of the movements of the other reduced-form coefficients. Changes in t move all coefficients proportionately.

Ignoring regime changes may have been responsible for spuriously unstable estimates over this period. Our model explains not only statistically significant movement of structural, and thus reduced-form, coefficients, but economically meaningful changes in those parameters. Stability tests conducted over a mid-sample split using the constant policy parameter specification indicate that specification is unstable. By contrast, when we allow reduced-form parameters to respond to regime changes, the resulting specification is stable.[8]

Table 3 lists the actual values of interest rates and the values predicted out-of-sample by row 1 and in-sample by rows 2 and 3 from Table 1 (constant policy parameters) and Table 2 (variable policy parameters). The 1980:06 credit control observation is omitted. The summary measures in Table 3 show that, out-of-sample or in, the Table 2 equations outperform those from Table 1 over the most recent period.[9] The relative improvement appears to be approximately the same for either in- or out-of-sample forecasts. In the out-of-sample case, the mean error is reduced by (4.33-3.01=) 132 basis points (or 30 percent) and the mean square error by 45 percent.

Fama and Gibbons (forthcoming) suggest that interest rate-based forecasts may be better predictors of inflation than survey-based forecasts. To investigate the robustness of our findings, we have substituted a measure of expected inflation based on prior interest rates for the Livingston survey measure. We generate these forecasts using only information publicly available when the interest rate was determined each period. The forecast during June depends on the six, monthly-average Treasury bill yields from December to May and on the forecast equation coefficients. The coefficients are obtained by regressing inflation on a constant and six lags of one-month Treasury yields over the forty-eight months ending two months before the forecast is made. Using coefficient estimates from a sample that edged closer to the forecast dates (June and December) would require more information than agents actually had. Using these expected inflation proxies, likelihood ratio tests of the hypotheses that the coefficients on the regulatory regime measures and the coefficient on the monetary regime proxy are zero are rejected for both the pre-1979 and the longer sample period. Thus our finding that the reduced-form coefficients respond systematically to government policy parameter shifts is further substantiated.

TABLE 3

A Comparison of Nominal Interest Rate Predictions from the Constant and the Variable Policy Parameter Models

		Constant Policy Parameters			Variable Policy Parameters		
Date	Actual	out-of sample	in-sample	in-sample with D7983	out-of sample	in-sample	in-sample with D7983
1979:12	12.26	12.02	13.50	15.56	12.72	14.21	14.70
1980:12	15.25	11.20	14.90	15.26	12.14	14.59	14.71
1981:06	15.23	9.34	13.54	13.63	10.52	13.21	13.29
1981:12	13.08	9.87	14.04	13.89	11.35	14.74	14.42
1982:06	14.38	7.27	9.43	11.46	8.42	12.26	12.19
1982:12	8.97	4.22	8.19	9.30	5.93	9.78	9.84
1983:06	9.73	4.66	8.10	9.79	6.76	9.61	9.76
MEAN ERROR		4.33	1.03	--	3.01	0.07	--
MEAN ABSOLUTE ERROR		4.33	1.66	1.29	3.14	1.34	1.34
MEAN SQUARE ERROR		22.89	4.74	3.24	12.67	2.33	2.49

E. Concluding Remarks

There is considerable intertemporal instability in previous interest rate equation estimates. We hypothesize that major fiscal, monetary, and regulatory policy parameter shifts have been important sources of that instability. We embed estimates of the time series values of these policy parameters in our model and estimate the deeper, more stable, underlying parameters. The estimates generate reduced-form coefficients that move by sizeable amounts in response to policy parameter change. Statistical tests imply that allowing for varying policy parameters provides a significantly better explanation of interest rates. These explainable movements in the reduced-form coefficients correspond rather closely to the heretofore unexplained movements in the rolling sample estimates obtained from the constant policy parameter model. The variable policy parameter model passes the test of stability that the model which ignores regime switches cannot. Furthermore, in-sample and out-of-sample forecasts from the proposed model outperform the more traditional specification.

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FOOTNOTES

1. See, for example, Cargill (1976), Wachtel (1977), Carlson (1979)).

2. See, for example, Cargill and Meyer (1974)).

3. This agenda ignores technological changes such as improvements in information processing and data transmission. Though the "deep" parameters of taste and technology may vary over time, their shifts are less readily quantified and are outside the range of this study.

4. This tax series serves as an index of the marginal tax rate of the marginal individual, moving with that rate but perhaps not measuring its level exactly.

5. Gordon (1984) discusses this effect.

6. Test statistics were calculated from the residuals obtained by ordinary least squares from the levels and from the first differences of (20). An alternative we have not pursued is that regime switches occur with changes in presidents rather than with Fed chairmen. The most discussed monetary policy regime switch clearly seems to be associated with the installation of Paul Volcker as Fed chairman in the middle of the Carter presidency.

7. In Peek and Wilcox (1984), we found that substituting an expected inflation measure based on prior interest rates did not affect our qualitative findings.

8. The chi-square test statistics and critical values for the constant policy parameter model are 23.9 and 15.5. For the variable parameter model they are 17.6 and 21.0.

9. The sample size is 62. The square root of the critical values for the chi-square distribution and (the absolute value of) the critical values for the t distribution converge as the sample size grows. These likelihood ratio tests reject the insignificance of the individual coefficients in (35)-(37) when the calculated chi-square test statistics exceed 3.84 or, equivalently, when the statistics in parentheses in (35)-(37) exceed 1.96.