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# The Postwar Stability of the Fisher Effect

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## ABSTRACT

Most research concerning the Fisher relationship has concentrated on the magnitude and significance of the response of nominal interest rates to anticipated inflation. Recently, attention has shifted to the stability of that response. According to previous estimates, the impact of anticipated inflation on interest rates varies substantially over time. By extending a standard model to include tax and aggregate supply shock effects, we are able to reduce such instability considerably. Our results also reveal that increased foreign demand for bonds lowers the interest rate.

FOR MORE THAN A DECADE, empirical studies appearing in economic journals have investigated the Fisher neutrality hypothesis that nominal interest rates adjust to changes in the anticipated inflation rate so as to leave real interest rates unaffected. This research has intensified with each upward ratchet in interest and inflation rates. Initially, studies focused on the magnitude of the nominal interest rate response to the anticipated inflation rate. Estimates based on data for the 1950s provided low and insignificant values for the response (see, for example, Cargill and Meyer [6]). As the sample period was extended, however, the puzzle of low coefficient estimates disappeared. Larger point estimates of the response were obtained until finally they began to approach unity. Darby [11] and Feldstein [13] then introduced income taxes into interest rate models, and the search began for a response equal to $1/(1 - t)$ (or at least greater than unity), rather than unity, where $t$ was the relevant marginal tax rate on interest income (for example, Carr et al. [10], Cargill [5]). Recent theoretical work, however, suggests that such a value is not necessarily implied by a properly specified model. Levi and Makin [18] show that a wide range of coefficient values is plausible. Specifically, they demonstrate that even in the presence of income taxation a response in the vicinity of unity is reasonable.

While this partially reconciled the conflicting theoretical and empirical evidence concerning the magnitude of the nominal interest rate response to anticipated inflation, important issues remained. One unsettling aspect of these studies was the volatility of the estimated response over various sample periods (see, for example, Cargill [4], Wachtel [24], Cargill and Meyer [7], [8], Levi and Makin [19], and Carlson [9]). In particular, Carlson [9], Cargill and Meyer [8], and Levi and Makin [19] each report estimates that drop, often dramatically, when the sample period is extended to include the first half of the 1970s. These results conflict with an explanation of the rise in that response during the 1960s based on a period of "learning" (for example, Cargill and Meyer [7]). Substantial

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residual autocorrelation is also a common characteristic of interest rate regressions. The combination of coefficient instability and the presence of autocorrelated residuals suggests that these equations may have been misspecified, perhaps by having omitted important explanatory variables.

This paper presents and estimates for the postwar period a model of interest rates that produces a statistically stable interest rate response to anticipated inflation. The model includes proxies for two important but usually omitted effects: aggregate supply forces and income taxes. Wilcox [25] provides evidence that supply shocks represent an important factor in interest rate determination. Peek [21] demonstrates that income taxes have significant effects on interest rates. Here we include both effects in our model and estimated equations. The empirical evidence suggests that income tax rates and aggregate supply forces each exert a significant impact on nominal interest rates over the 1952–79 period. Our results also imply that much of the previously reported coefficient instability may be spurious, having been generated by inappropriately specified models. The evidence then indicates that estimates of the anticipated inflation effect have been biased due to the failure to incorporate these additional factors, and furthermore that the apparent instability reflects the changing relationships between the omitted and included explanatory variables over this sample period.

## I. The Model

The analysis is based on a simple macro model composed of IS, LM, wage, and aggregate supply equations. These four relationships can be expressed in linearized form as

$$Y - Y^N = a_0 - a_1 r^* + a_2 Z + a_3(X - Y^N) + a_4(M - P - Y^N) - a_5\mathrm{SS} - a_6\mathrm{FB} \quad (1)$$

$$M - P - Y^N = b_0 + b_1(Y - Y^N) - b_2 i^* - b_3\mathrm{FB} \quad (2)$$

$$W = c_0 + P^e - c_1\mathrm{SS} \quad (3)$$

$$P = W + d_1(Y - Y^N) + d_2\mathrm{SS}, \quad (4)$$

where the coefficients of all the variables are positive and

$Y$ = the logarithm of actual real output
$Y^N$ = the logarithm of natural (i.e., potential) real output
$Z$ = the percentage change in real actual output lagged one period
$X$ = the logarithm of the sum of real exports and real government expenditures
$M$ = the logarithm of the nominal 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
FB = the domestic bonds held by foreigners
$r^*$ = 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 (5) and (6)

$$i^* \equiv i(1 - t) \tag{5}$$

$$r^* \equiv i^* - p^e \tag{6}$$

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, real exogenous export and government demand, a real balance effect, and an investment accelerator term. The opportunity cost of holding money is the after-tax nominal interest rate. The wage and price equations embody the natural rate hypothesis. While the basic structure of this model is similar to that of Peek [21], it differs by incorporating the impact of supply shocks (SS) and bond demand by foreigners (FB). Supply shocks exert their effects through three channels. First, a shock to the supply of one factor input, such as energy, alters the demand for capital. The resulting change in investment demand shifts the IS curve (see Wilcox [25]). Berndt and Wood [3], Berndt and Khaled [2], and Hudson and Jorgenson [17] each provide empirical evidence that negative supply shocks reduce capital and investment demand. The supply shock variable also enters the aggregate supply equation (4). A shock that increases the relative price of a factor input, ceteris paribus, raises the cost of production. The wage equation (3) exhibits the third path of the supply shock effect. An adverse supply shock would be expected to reduce the equilibrium real wage as well. Gordon [14] and Gramlich [15] document empirically the importance of this factor in U. S. wage determination.

During the 1970s these negative energy shocks transferred real income to OPEC countries. If these countries save a higher proportion of this transferred income, world saving increases. Part of the associated decline in world spending represents a concomitant decline in expenditure on U. S. goods. Although such income redistribution effects are usually ignored, we incorporate this additional IS curve shift factor for two reasons. The magnitude of the redistribution during this period is very large relative to that historically encountered. Second, allowing for this effect permits us to distinguish between IS shifts brought about by declines in investment and those due to increases in world saving. We assume that part of this increase in world saving generates an increase in the foreign demand for U. S. bonds. Thus, we take foreign holdings of bonds (FB) as a proxy for this effect. As a result of this transfer, OPEC countries also accumulate wealth. If these OPEC wealth-holders desire a portfolio composition with a much higher proportion of U. S. government securities than do domestic wealth-holders, the demand for money will shift downward, causing a corresponding downward (rightward) LM shift (see Adams [1] and Hartman [16] for a discussion of the importance of foreign holdings of domestic securities). Either of these effects would reduce the interest rate. Since the effects are reinforcing, if either effect operates, the bond variable appears in the reduced-form interest rate equation with a negative coefficient.

Equations (1) to (4) and (6) can be combined to yield the reduced-form equation for the after-tax nominal interest rate:

$$\begin{matrix} i^* = A_0 + A_1 p^e + A_2 M' + A_3 X' + A_4 Z + A_5 \mathrm{SS} + A_6 \mathrm{FB} \\ \quad\quad (+) \quad\quad (?) \quad\quad (+) \quad\quad (+) \quad\quad (?) \quad\quad (-) \end{matrix} \tag{7}$$

where $M'$ and $X'$ are $(M - P^e - Y^N)$ and $(X - Y^N)$, respectively. Liquidity and real balance effects have conflicting impacts on the interest rate, resulting in an ambiguous sign for $A_2$. If liquidity forces dominate, a negative value for $A_2$ is predicted. Likewise, the sign of $A_5$ 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 $A_5$. The results presented in Wilcox [25] can be so interpreted.

To convert (7) into the standard form of the Fisher equation which explains before-tax nominal interest rate movements, substitute (5) into (7) and divide both sides of the equation by $(1 - t)$ to obtain

$$i = \frac{A_0}{(1-t)} + A_1\tilde{p}^e + A_2\tilde{M}' + A_3\tilde{X} + A_4\tilde{Z} + Z_5\widetilde{\mathrm{SS}} + A_6\widetilde{\mathrm{FB}}, \tag{8}$$

where the $\sim$ over a variable indicates that it has been divided by $(1 - t)$. From (8) it can be seen that

$$\frac{di}{dp^e} = \frac{A_1}{(1-t)} = \frac{1}{(1-t)\left[1 + \dfrac{b_2(1 + a_4\, d_1)}{a_1(b_1 + d_1)}\right]} \tag{9}$$

In general, $di/dp^e$ will be less than the value of $1/(1 - t)$ suggested by Darby [11]. It will be equal to $1/(1 - t)$ if, for example, the LM curve is vertical or the IS curve is horizontal. These conditions are consistent with those required for a unitary coefficient on the anticipated inflation rate when taxes are ignored (see Sargent [22]).

## II. Empirical Results

The reduced-form estimates are based on semiannual observations corresponding to the Livingston survey data collected each June and December. The sample extends from June 1952 through December 1979. This sample period avoids both the pre-1952 pegging of interest rates by the Federal Reserve and any structural changes arising from the imposition of credit controls in 1980. Monthly averages of the 1-year Treasury bill yield during June and December are used as the nominal interest rate measure to match the maturity of the Livingston 1-year anticipated inflation rate data.[1]

The anticipated inflation rate series, PE12, is the percentage of change in the CPI expected over the next 12 months derived from the Livingston survey. This series was provided by the Federal Reserve Bank of Philadelphia. This measure of anticipated inflation has two advantages: it is a truly ex ante expectation and it reflects whatever sophistication agents use to process information.

Second- and fourth-quarter observations are used for the remaining explana-

[1] Before December 1959, when 1-year Treasury bills were introduced, the interest rate measure is the yield on bills with maturities of 9 to 12 months.

tory variables. The logarithm of the sum of real exports and real government expenditures on goods and services divided by the level of natural real output (EXG) and the percentage change in real GNP lagged one period ($DGNP_{-1}$) are constructed from the National Income and Product Accounts data. We use the potential real GNP series constructed by the Council of Economic Advisors as our measure of natural real output. The logarithm of the nominal money supply deflated by the expected price level and natural real output (M1B) is constructed using the M1B definition of the money supply and the Livingston survey measure of the expected price level.[2]

The tax rate on interest income is an average marginal tax rate constructed from data contained in annual editions of *Statistics of Income, Individual Income Tax Returns.* Following Wright [26], 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. This provides one value for each year. The same value is assigned to each semiannual observation within a given year. The series used here has been smoothed with a centered three-period moving average. The calculated series starts with a value of approximately 0.32 in 1952, declines to 0.24 by 1965, and then generally increases until it reaches 0.35 in 1979. The major tax cuts in 1954 and 1964–65 as well as the income tax surcharge in 1968–70 account for the major sharp fluctuations in the tax rate proxy. Bracket creep is also apparent as nominal income growth pushes individuals into higher marginal tax brackets.

The supply shock variable, SUPPLY, is measured by the ratio of the implicit price deflator for imports to the GNP deflator.[3] It is intended to reflect changes in the world supply of materials, broadly defined to include both materials for further processing, such as minerals, as well as inputs consumed in production, such as oil. The import deflator responds to both supply shifts and changes in the exchange rate. However, exchange rate changes and real materials supply shocks may have different effects on real interest rates since the former implies an offsetting shock to foreign markets while the latter may be a positive or negative shock to all countries. Consequently, in constructing SUPPLY, exchange rate changes were eliminated from the import deflator by multiplying it by the effective exchange rate.

The proxy for foreign holdings of bonds (FB) is the ratio of foreign holdings to the sum of private domestic and foreign holdings of U. S. government short-term marketable securities taken from the Government Securities Table in the Federal Reserve Flow of Funds Accounts. Movements in this ratio reflect the changing importance of foreign investors relative to the size of the outstanding stock of securities.

[2] The interest rate data are from [27]. The post-1958 money supply data are also from [27]. The earlier money supply data were provided by the Board of Governors of the Federal Reserve System. The potential real output series is taken from [28]. The national income account data are from [29] and supplemented by the July issues of [30].

[3] Wilcox [25] finds that a proxy which is limited to the energy sector, the producer price index for fuel and power relative to the GNP deflator, is highly correlated with SUPPLY and delivers similar regression results.

**Table I**

Reduced Form Estimates for Interest Rates 1952:06–1979:12, Semiannually

| Estimation Method | Coefficient on | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | Constant | PE12 | M1B | EXG | $DGNP_{-1}$ | SUPPLY | FB | $\hat{\rho}$ | $R^2$ | D.W. | S.E.E. |
| 1. OLSQ[a] | 2.92 | 0.895 | −0.742 | 0.90 | 20.70 | — | −9.06 | — | 0.8698 | 1.25 | 0.839 |
| | (0.81)[b] | (5.37) | (−0.46) | (0.31) | (2.98) | | (−3.40) | | | | |
| 2. CORC | 9.40 | 0.769 | −0.797 | 6.34 | 15.08 | — | −4.34 | 0.47 | 0.8911 | 2.10 | 0.763 |
| | (1.58) | (3.82) | (−0.35) | (1.46) | (2.33) | | (−1.22) | (3.98) | | | |
| 3. OLSQ | 16.88 | 1.103 | 1.988 | 5.20 | 17.80 | −4.24 | −6.71 | — | 0.9126 | 1.68 | 0.694 |
| | (4.10) | (7.65) | (1.37) | (2.06) | (3.08) | (−4.90) | (−2.97) | | | | |
| 4. Tax-adjusted | 12.79 | 0.727 | 1.184 | 4.00 | 12.12 | −3.60 | −4.69 | — | 0.9190 | 1.76 | 0.668 |
| OLSQ | (4.57) | (7.42) | (1.19) | (2.34) | (3.06) | (−6.12) | (−3.05) | | | | |

[a] OLSQ, ordinary least squares; CORC, least squares with the Cochrane-Orcutt autocorrelation correction.

[b] Numbers in parentheses, $t$-statistics.

Equation (8) can be rewritten using these proxy variables as

$$i = \frac{A_0}{(1-t)} + A_1 \frac{\text{PE12}}{(1-t)} + A_2 \frac{\text{M1B}}{(1-t)} + A_3 \frac{\text{EXG}}{(1-t)} + A_4 \frac{\text{DGNP}_{-1}}{(1-t)} + A_5 \frac{\text{SUPPLY}}{(1-t)} + A_6 \frac{\text{FB}}{(1-t)} \tag{10}$$

Table I displays the results of estimating complete and restricted versions of (10). The restricted equation, exhibited in Row 1, differs from the unrestricted specification (10) by excluding SUPPLY and omitting the income tax adjustment by setting $t = 0$. The low Durbin-Watson statistic and the sizable (0.47) and significant autocorrelation coefficient obtained in Row 2 argue for Cochrane-Orcutt estimation. Only PE12 and $\text{DGNP}_{-1}$ remain statistically significant after the autocorrelation correction. This specification, which ignores supply shocks and taxes but includes liquidity and demand proxies, resembles recent empirical formulations (e.g., Carlson [9], Cargill and Meyer [8]). We present these results to facilitate comparison with our expanded model.

Row 3 summarizes our findings when the supply shock proxy, SUPPLY, is included. The standard error of the regression is reduced by 9% and all of the estimated coefficients except that for M1B are now significant. In particular, the $t$-statistic associated with the SUPPLY coefficient has a value of nearly 5. The introduction of SUPPLY causes the anticipated inflation rate coefficient to rise from 0.769 to 1.103, a rise of over 1.5 times the value of that coefficient's Row 2 standard error. Furthermore, the significant estimated coefficient associated with FB implies that an increase in foreign demand for bonds reduces the interest rate. No significant autocorrelation remains. These changes indicate that a relevant, but previously omitted, variable has been introduced.

Row 4 is based on the income tax-adjusted specification (10).[4] This modification further improves the overall fit of the equation, slightly increases the Durbin-Watson statistic, and generally raises the significance level of the explanatory variables. The $t$-statistic of over 6 associated with the estimated coefficient of SUPPLY substantiates, in a tax-adjusted specification, the claim by Wilcox [25] that SUPPLY is an important factor in the determination of interest rates. Furthermore, in the presence of the previously omitted SUPPLY variable, these results confirm the importance of income taxes in the determination of the nominal interest rate found by Peek [21]. The evidence presented in Row 4 suggests that the tax-adjusted formulation provides a superior empirical explanation of nominal interest rate movements. In fact, Davidson-McKinnon [12] non-nested specification tests reject the non-tax-adjusted model in favor of the tax-adjusted formulation.[5] With the income tax adjustment, the estimated coef-

[4] Tanzi [23] incorporated tax effects in a similar way.

[5] When the null hypothesis ($H_0$) is that

$$i = \frac{A_0}{(1-t)} + A_1 \frac{\text{PE12}}{(1-t)} + A_2 \frac{\text{M1B}}{(1-t)} + A_3 \frac{\text{EXG}}{(1-t)} + A_4 \frac{\text{DGNP}_{-1}}{(1-t)} + A_5 \frac{\text{SUPPLY}}{(1-t)} + A_6 \frac{\text{FB}}{(1-t)} \tag{1}$$

ficient on PE12 is reduced from 1.103 to 0.727 and is significantly below unity. This finding can be related to our model. According to Equation (9), an $A_1$ estimate significantly below unity suggests that $b_2(1 + a_4 d_1)/a_1(b_1 + d_1) > 0$. For example, our results support a non-zero interest elasticity of money demand, a finite income elasticity of money demand, and a finite interest elasticity of expenditures, none of which is surprising.

Previous investigators have obtained a wide range of estimated coefficients for the anticipated inflation rate by altering the sample period (e.g., Carlson [9], Cargill and Meyer [6, 7]). Figure 1 plots the values of the estimated PE12 coefficient obtained using a rolling-regression technique. These estimates are based on 14-year subsamples (half the sample), the beginning and ending dates of which are rolled forward one period at a time. This produces an estimate for the first half, last half, and all contiguous halves of the total sample in between. The estimated PE12 coefficients based on the unrestricted and retricted specifications are labeled $\hat{A}_1^U$ and $\hat{A}_1^R$, respectively. The unrestricted model is (10); the restricted model differs by omitting SUPPLY and setting $t = 0$. Both series of estimates vary over time and $\hat{A}_1^U$ is not obviously more stable than $\hat{A}_1^R$. Both series rise very sharply, although temporarily, in the early subsamples. Both decline appreciably as the mid-1970s are added to the rolling samples and both rise again in later subperiods.

Table II presents the results of estimating the non-SUPPLY, non-tax-adjusted specification (Table I, Row 1) and the complete model (Table I, Row 4) when the complete sample is halved in two different ways. Rows 1, 2, 5, and 6 present the evidence obtained by splitting the sample period at the end of 1965. This

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is the true model and the alternative hypothesis ($H_1$) is that

$$i = B_0 + B_1\text{PE12} + B_2\text{M1B} + B_3\text{EXG} + B_4\text{DGNP}_{-1} + B_5\text{SUPPLY} + B_6\text{FB} \tag{2}$$

is the true model [i.e., Equation (1) omitting taxes], the test equation is

$$i = \frac{C_0}{(1-t)} + C_1\frac{\text{PE12}}{(1-t)} + C_2\frac{\text{M1B}}{(1-t)} + C_3\frac{\text{EXG}}{(1-t)} + C_4\frac{\text{DGNP}_{-1}}{(1-t)} + C_5\frac{\text{SUPPLY}}{(1-t)} + C_6\frac{\text{FB}}{(1-t)} + C_7\text{FIT2} \tag{3}$$

where FIT2 is the predicted series for $i$ based on Equation (2). The true value of $C_7$ is zero when $H_0$ is true. Consequently, $H_0$ cannot be rejected if the estimated value of $C_7$ does not differ significantly from zero. If $H_1$ is true, the estimated value of $C_7$ should be near unity. To test the null hypothesis that the non-tax-adjusted specification (2) is the true model, we reestimate (3), reversing the roles of (1) and (2) (i.e., estimating (2) with the edition of FIT1 as an explanatory variable). The results may reject neither, either, or both of the proposed specifications. Thus, the test does not force acceptance of one specification and rejection of the other.

In the alternative version of Equation (3), the estimated value of the FIT1 coefficient was 1.64 with a $t$-statistic of 2.12 enabling us to reject the non-tax-adjusted specification. Furthermore, the estimated coefficients associated with the other explanatory variables were statistically insignificant. Reversing the procedure, the estimated coefficient of FIT2 was −0.64 with a $t$-statistic of only −0.82. Thus, we were unable to reject the tax-adjusted model. In addition, all of the other explanatory variables (except M1B) retained their statistical significance at the 10% level or better. The reduction in their $t$-statistics is related to the increased multicollinearity associated with the addition of FIT2 which is highly correlated with the other explanatory variables.

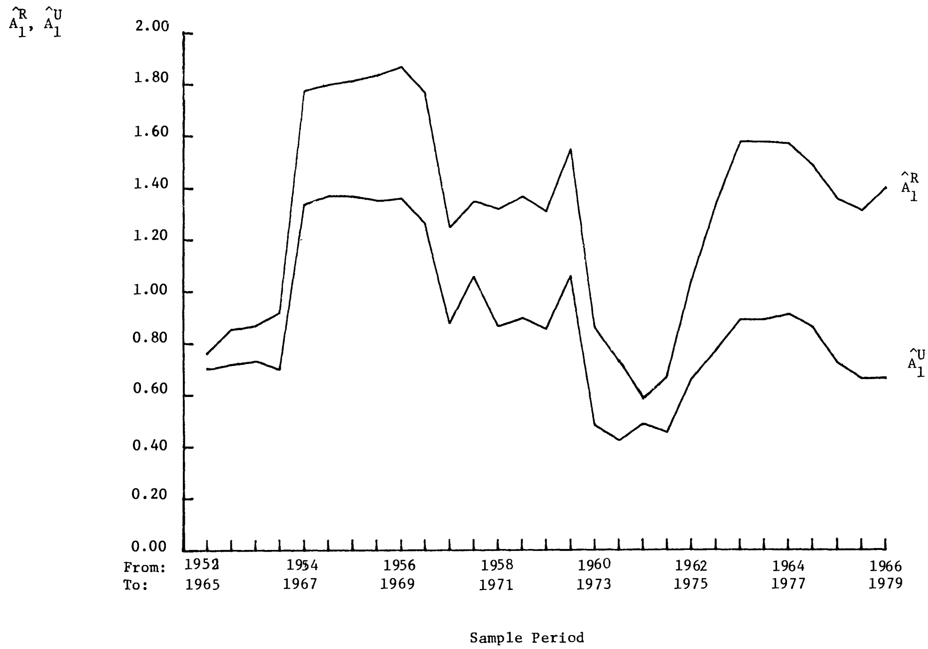


**Figure 1.** Anticipated Inflation Coefficients from Unrestricted and Restricted Regressions. $A_1^U$ and $\hat{A}_1^R$ are the estimated PE12 coefficients from the unrestricted and the restricted specifications, respectively. Their time series are generated using a rolling sample of length 14 years.

conveniently separates the earlier, low-inflation period from the high-inflation period that starts in the mid-1960s. This provides an opportunity to assess whether markets reacted differently to the higher average rates of inflation in the latter period. The second split presented in the remaining four rows of Table II isolates the middle half of the full sample period. Much of the previous evidence suggests that the reaction to the expected inflation rate in the 1950s, 1960s, and 1970s was different. In particular, those findings suggest that the interest rate reacted more strongly to anticipated inflation in the 1960s.

In the restricted specification (Rows 1 to 4) the estimated coefficient of PE12 jumps from 0.762 to 1.399 between the first and second halves of the full sample period. In comparing the middle-half subsample to the subsample consisting of the first and final quarters of the full sample period, the coefficient changes from 1.308 to 1.048. While this is not as great a difference as between the first and second halves, it is still a substantial change. In the complete model (Rows 5 to 8), the difference in the estimated coefficient is relatively minor, 0.701 compared to 0.663. The change between the middle half and the first and final quarters is slightly larger, 0.853 compared to 0.735, but is still substantially less than the difference obtained in the alternative restricted specification. The overall standard errors of the complete specification are also generally smaller than those for the restricted version, especially for subsamples in the 1970s.

Table III displays the results of formal coefficient stability tests based on the subperiods presented in Table II. The test statistics in Row 1 pertain to a mid-sample split. Those in Row 2 refer to the halves of the full 1952–79 sample

**Table II**

Interest Rate Estimates over Various Sample Periods (Ordinary Least Squares)

| Sample Period | | Coefficient on | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | | Constant | PE12 | M1B | EXG | $DGNP_{-1}$ | SUPPLY | FB | $R^2$ | D.W. | S.E.E. |
| 1. 1952:06–1965:12 | | 5.62 (0.77) | 0.762 (1.96) | −1.53 (−0.82) | 5.45 (0.97) | 19.1 (2.47) | — | 9.33 (0.63) | 0.6503 | 1.56 | 0.686 |
| 2. 1966:06–1979:12 | | 12.83 (1.64) | 1.399 (4.42) | 7.43 (2.08) | −1.81 (−0.27) | 19.0 (1.64) | — | −7.70 (−2.11) | 0.7537 | 1.71 | 0.886 |
| 3. 1959:06–1972:12 | | −1.97 (−0.20) | 1.308 (4.40) | 1.11 (0.28) | −5.37 (−1.13) | 20.5 (2.30) | — | −12.40 (−3.52) | 0.8432 | 1.52 | 0.589 |
| 4. 1952:06–1958:12, 1973:06–1979:12 | | 13.57 (2.39) | 1.048 (4.39) | 3.18 (1.25) | 6.18 (1.58) | 18.2 (1.86) | — | 3.06 (0.49) | 0.9168 | 1.47 | 0.914 |
| 5. 1952:06–1965:12 | Tax-adjusted | 19.02 (1.23) | 0.701 (2.29) | 2.23 (0.60) | 7.31 (1.44) | 12.6 (2.22) | −5.05 (−1.06) | 7.78 (0.72) | 0.6686 | 1.66 | 0.683 |
| 6. 1966:06–1979:12 | Tax-adjusted | 17.54 (3.73) | 0.663 (3.63) | −0.72 (−0.28) | 9.41 (2.03) | 8.7 (1.37) | −5.14 (−3.64) | −3.52 (−1.76) | 0.8616 | 2.12 | 0.679 |
| 7. 1959:06–1972:12 | Tax-adjusted | −0.73 (−0.06) | 0.853 (2.71) | 0.11 (0.04) | −2.84 (−0.77) | 15.4 (2.28) | −0.37) (−0.06) | −8.86 (−3.14) | 0.8416 | 1.48 | 0.606 |
| 8. 1952:06–1958:12, 1973:06–1979:12 | Tax-adjusted | 17.89 (4.83) | 0.735 (5.90) | 2.12 (1.57) | 6.73 (3.09) | 10.1 (1.95) | −4.34 (−3.91) | 0.10 (0.03) | 0.9549 | 2.11 | 0.689 |

[a] Numbers in parentheses, $t$-statistics.

**Table III**

Tests of Coefficient Stability

| Subsample Dates (First Subsample/Second Subsample) | Specification: Restricted — Expected Inflation $\hat{t}$ (1) | Restricted — All $\hat{F}$ (2) | Unrestricted — Expected Inflation $\hat{t}$ (3) | Unrestricted — All $\hat{F}$ (4) |
|---|---|---|---|---|
| 1. 1952:06–1965:12/1966:06–1979:12 | −0.20 | 2.00 | −1.17 | 0.74 |
| 2. 1952:06–1958:12, 1973:06–1979:12/1959:06–1972:12 | 2.70[a] | 2.58[a] | −1.25 | 1.42 |

[a] Significant at the 0.05 level.

composed of the first and last quarters of the sample (1952–58 and 1973–79) and the middle half of the sample (1959–72). Columns 1 and 2 contain the test results based on the non-tax, non-SUPPLY specification.[6] The last two columns are derived from the unrestricted specification. Rows 1 and 2 of Table II imply a rather large shift in the PE12 point estimate as the sample moves from the first to the second half of the period. Table III, however, shows that neither the hypothesis that the expected inflation coefficient alone is stable nor that the entire coefficient vector is stable can be rejected when the sample is halved at its midpoint,[7] nor can stability be rejected for the tax-adjusted, SUPPLY specification as indicated by the small values of the test statistics in Columns 3 and 4.

When the sample is split in half by grouping the first and last quarter of the full sample as one subperiod and the middle half as the other, a different picture emerges. Although the point estimates across Rows 3 and 4 in Table II vary less than those across Rows 1 and 2, the $t$- and $F$-tests, respectively, allow us to reject the stability of the expected inflation rate coefficient alone and the coefficient vector as a whole in the unrestricted specification. The test statistic for each hypothesis is significant at the 0.05 level. When the sample split is used for the tax-adjusted SUPPLY specification, we cannot reject stability, either of the coefficient vector as a whole or of the expected inflation rate coefficient alone. Thus, incorporation of supply and tax effects jointly reduces the instability reported elsewhere (e.g., Carlson [9], Cargill and Meyer [8]) and documented here for our restricted specification.

Omission of tax and supply effects from Equation (8) can produce biased estimates, as shown in Table II. Even when the true reduced-form parameters are constant, coefficients may appear to vary if the estimation bias varies over

[6] Each of these tests is based on ordinary least squares estimates. The reported $F$-statistics result from Chow-type stability tests based on the equations presented in Table II. The equations in Table II were then reestimated with an additional explanatory variable. This variable had a value of zero in the first subsample and took on the values of PE12 in the second subsample. The $t$-statistics reported in Columns 1 and 3 are associated with the estimated coefficients on this additional variable. The hypothesis of intertemporal stability of the PE12 coefficient is rejected when the $t$ value exceeds the relevant critical level.

[7] Makin [20] also attains stable coefficients when his 1959–81 sample is split near its midpoint at the end of 1969. His empirical model includes the effects of unexpected money and inflation uncertainty, but excludes tax and supply effects.

**Table IV**
Correlations of PE12 with SUPPLY and $1/(1-t)$ over Various Subsamples

| | Simple Correlation Between PE12 and | |
|---|---|---|
| Sample Period | SUPPLY (1) | $1/(1-t)$ (2) |
| 1. 1952:06–1979:12 | −0.01 | 0.46 |
| 2. 1952:06–1965:12 | −0.79 | −0.80 |
| 3. 1959:06–1972:12 | −0.94 | 0.22 |
| 4. 1966:06–1979:12 | 0.78 | 0.94 |
| 5. 1952:06–1958:12, 1973:06–1979:12 | −0.15 | 0.55 |

time. For the expected inflation coefficient, the size of that bias will vary with the extent of the correlation of expected inflation with the omitted variables. Changes over time in the correlations between PE12 and $1/(1-t)$ and SUPPLY then lead to intertemporal instability in the estimated expected inflation coefficient (and in the remaining coefficient estimates as well). Equation (8) implies that the level of $1/(1-t)$ also affects the coefficient estimates unless tax adjustments are made. Table IV presents values of the simple correlations between PE12 and both SUPPLY and $1/(1-t)$ for the subsamples considered in Tables II and III. The relatively low correlations in the first row mask the fact that these correlations have been high (in absolute value) within subsamples and changeable across subsamples. The change in the pattern of correlation between PE12 and excluded SUPPLY and $1/(1-t)$ is dramatic. For both variables the strong, negative correlations with PE12 in the 1952–65 period give way to equally strong, positive correlations in the 1966–79 period.

The fact that these changing correlations between PE12 and SUPPLY and $1/(1-t)$ are consistent with the instability of the estimated coefficient of PE12 in the restricted relative to the unrestricted specification is demonstrated in (11) below. There the ratio of $\hat{A}_1^R$ to $\hat{A}_1^U$ from the rolling regressions is regressed on the average level of $1/(1-t)$, the correlation between PE12 and $1/(1-t)$, and the correlation between PE12 and SUPPLY. These correlations are labeled CORRPET and CORRPESS, respectively. Each explanatory variable is calculated over the same rolling sample periods used to derive the $\hat{A}_1^R$ and $\hat{A}_1^U$ series. We include the mean of $1/(1-t)$ since this term enters the reduced-form coefficient of PE12 multiplicatively.[8] The results of this regression are ($t$-statistics in parentheses)[9]

$$\hat{A}_1^R/\hat{A}_1^U = \underset{(-3.25)}{-18.9} + \underset{(3.51)}{14.7}\,\overline{\frac{1}{(1-t)}} + \underset{(4.38)}{0.39}\text{CORRPET} - \underset{(-1.93)}{0.22}\text{CORRPESS} + \underset{(1.34)}{0.24}\hat{e}_{-1}$$

$$R^2 = 0.8055 \quad \text{D.W.} = 2.06 \quad \text{S.E.E.}/(\hat{A}_1^R/\hat{A}_1^U) = 0.076 \tag{11}$$

[8] The mean value of SUPPLY is impounded in the constant term and does not bias PE12 since SUPPLY, unlike the income tax rate, does not enter the reduced form coefficient for PE12.

[9] The coefficient on $\hat{e}_{-1}$ is the autocorrelation-correction coefficient. The regression was also performed using a centered, three-term moving average of $(B/BTS)$ as the independent variable. Similar but more precise point estimates were obtained.

Each of the three effects is statistically significant and has the hypothesized sign. Higher average tax rates increase the spread between the non-tax-adjusted coefficient and the tax-adjusted coefficient. A larger correlation of PE12 with $1/(1 - t)$ produces a more upwardly biased coefficient on PE12. Similarly, a larger correlation between PE12 and SUPPLY results in a more downwardly biased PE12 coefficient.

## III. Conclusion

In the last decade as inflation rose to historically high levels, the magnitude, the precision, and the stability of estimates of the response of nominal interest rates to anticipated inflation have each come into question. The estimates presented here demonstrate separate, significant reactions of interest rates to income tax rates, to aggregate supply shocks, and to wealth redistributions among groups with different saving and wealth-holding propensities. The results reveal that the substantial residual autocorrelation extant in models that omit these factors is removed in our more inclusive specification. Formal tests imply that the inclusion of tax and supply effects reduces coefficient instability. A regression test supports the hypothesis that the changing relationship between the included and excluded variables is responsible for much of this instability. Formal tests, however, do not reject the stability of the expanded formulation, although our rolling-sample results suggest that there are still more factors relevant to the determination of interest rates that remain to be identified.

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