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. Previous estimates of the impact of anticipated inflation on interest rates vary substantially over time. Our findings indicate that such instability is reduced considerably in an extended model. In addition to significant income tax and aggregate supply shock effects, the results also point to the separate effect on interest rates of increased foreign demand for bonds.
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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. Previous estimates of the impact of anticipated inflation on interest rates vary substantially over time. Our findings indicate that such instability is reduced considerably in an extended model. In addition to significant income tax and aggregate supply shock effects, the results also point to the separate effect on interest rates of increased foreign demand for bonds.

I. Background

Empirical investigations of the Fisher neutrality hypothesis that nominal interest rates adjust to changes in the anticipated inflation rate so as to leave real interest rates unaffected have appeared frequently in economic journals for more than a decade. In fact, the research intensified with each upward ratchet in interest and inflation rates. Initially, studies focused on the magnitude of the point estimates of the response of nominal interest rates 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 (1974)). 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 (1975) and Feldstein (1976) introduced income taxes into interest rate models and the search began for a response equal to \( \frac{1}{(1-t)} \) (or at least greater than unity), rather than unity, where \( t \) was the relevant
marginal tax rate on interest income (for example Carr, Pesando, and Smith (1976), Cargill (1977)). Recent theoretical work, however, suggests that such a value is not necessarily implied by a properly specified model. Levi and Makin (1978) 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 reduced the controversy over the exact magnitude of the point estimate of the impact of anticipated inflation on nominal interest rates, important issues remained. One unsettling aspect of these studies was the volatility of the estimated response over various sample periods (see for example Cargill (1976), Wachtel (1977), Cargill and Meyer (1977, 1980), Levi and Makin (1979), and Carlson (1979)). In particular, Carlson (1979), Cargill and Meyer (1980) and Levi and Makin (1979) 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 in the 1960s based on a period of "learning" (for example Cargill and Meyer (1977)). Another common characteristic of interest rate regressions is substantial residual autocorrelation. 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 a model of interest rates that leads to an empirically stable response to anticipated inflation for the postwar period. The model includes proxies for two important, but
usually omitted, effects: aggregate supply forces and income taxes. Wilcox (1983) provides evidence that supply shocks represent an important factor in interest rate determination. Peek (1982) 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, having been generated by inappropriately specified models, may be spurious. 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.

II. 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

\begin{align*}
(1) \quad Y - Y^N &= a_0 - a_1 r^* + a_2 Z + a_3 (X - Y^N) + a_4 (M - P - Y^N) - a_5 SS - a_6 FB \\
(2) \quad M - P - Y^N &= b_0 + b_1 (Y - Y^N) - b_2 i^* - b_3 FB \\
(3) \quad W &= c_0 + p^e - c_1 SS \\
(4) \quad P &= W + d_1 (Y - Y^N) + d_2 SS,
\end{align*}

where the coefficients of all the variables are positive and:

\( Y \) = the logarithm of actual real output,
\( Y^N \) = the logarithm of natural real output,
\( Z \) = the percentage change in real 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):

\[
(5) \quad i^* \equiv i(1 - t) \\
(5) \quad r^* \equiv i^* - p^e
\]

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

The model is similar to that of Peek (1982). The primary difference here is that we also allow for 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 (1983)). Berndt and Wood (1979), Berndt and Khaled (1979), and Hudson and Jorgenson (1978) each provide empirical evidence that capital and investment
demand are reduced by such negative supply shocks. 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 (1980) and Gramlich (1979) document empirically the importance of this factor in U.S. wage determination.

FB enters the equations to allow for the effect of the foreign demand for bonds. This variable is related to, but not identical to, the supply shock variable. For example, one result of the OPEC energy shock was a transfer of real income to OPEC countries. To the extent that these countries save a higher proportion of this transferred income, saving is increased. This shifts the IS curve downward (leftward). As a result of this transfer, OPEC countries also accumulate wealth. To the extent that the OPEC holders of this wealth have a desired portfolio composition with a much higher proportion of U.S. government securities than domestic wealth holders, there is a downward shift in the demand for money and a corresponding downward (rightward) LM shift (see Adams (1973) and Hartman (1980) 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, the bond variable appears in the reduced-form interest rate equation if either effect operates.

Equations (1) - (4) and (6) can be combined to yield the reduced-form equation for the after-tax nominal interest rate:
\( i^* = A_0 + A_1 p^e + A_2 M' + A_3 X' + A_4 Z + A_5 S + A_6 F B \)

\( (+) \quad (?) \quad (+) \quad (?) \quad (-) \)

where \( M' \) and \( X' \) are \( (M - p^e - \gamma N) \) and \( (X - \gamma N) \), respectively. The ambiguous sign of \( A_2 \) is due to the conflicting impact of money on the interest rate due to liquidity and real balance effects. If liquidity forces dominate, a negative value for \( A_2 \) is predicted. The sign of \( A_5 \) is likewise indeterminate \textit{a priori}. An adverse supply shock tends to reduce investment and real wages and thus the interest rate, while the resulting increased input costs operating through the aggregate supply equation have the opposite effect on 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 (1983) 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

\( i = \frac{A_0}{(1-t)} + A_1 \tilde{p}^e + A_2 \tilde{M}' + A_3 \tilde{X}' + A_4 \tilde{Z} + A_5 \tilde{S} + A_6 \tilde{F} B, \)

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

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

In general, \( \frac{di}{dp^e} \) will be less than the value of \( \frac{1}{(1-t)} \) suggested by Darby (1975). It will be equal to \( \frac{1}{(1-t)} \) 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 1976)).
III. 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 one-year Treasury bill yield during June and December are used as the nominal interest rate measure for consistency with the Livingston one-year anticipated inflation rate data.¹

The anticipated inflation rate series, PE12, is the percentage change in the CPI expected over the next twelve months derived from the Livingston survey. This series was provided by the Federal Reserve Bank of Philadelphia. This measure of anticipated inflation has the advantages of being a truly \textit{ex ante} expectation and of being able to reflect sophisticated information processing.

Second and fourth quarter observations are used for the remaining explanatory 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) are constructed from the National Income and Product Accounts data and from the potential real GNP series constructed by the Council of Economic Advisors. 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.

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 (1969), 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 supply shock variable, SUPPLY, is measured by the ratio of the implicit price deflator for imports to the GNP deflator. It is intended to reflect changes in the world supply of materials, broadly defined to include both materials for further processing, like minerals, as well as inputs consumed in production, like 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. Movements in
this ratio reflect the changing importance of foreign investors relative to the size of the outstanding stock of securities.

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

\[ i = \frac{A_0}{(1-t)} + A_1(1-t) + A_{PE12} \frac{M1B}{2(1-t)} + A_3(1-t) + A_{EXG}\\frac{DGNP_{-1}}{4(1-t)} + A_5\frac{SUPPLY}{5(1-t)} + A_6\frac{FB}{6(1-t)}. \]

The results of estimating complete and restricted versions of (10) are presented in Table 1. The restricted equation, exhibited in row 1, excludes SS and omits tax adjustment by setting \( t = 0 \). Only PE12 and DGNP_{-1} are statistically significant in this specification. Furthermore, a sizable (0.47) and significant autocorrelation correction is required. This specification, which ignores supply shocks and taxes but includes liquidity and demand proxies, resembles recent empirical formulations (e.g., Carlson (1979), Cargill and Meyer (1980)). We present these results to facilitate comparison with our expanded model.

Row 2 presents the results obtained 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 five. The introduction of SUPPLY causes the anticipated inflation rate coefficient to rise from 0.769 to 1.103, a rise of over one-and-a-half times the value of that coefficient's row 1 standard error. Furthermore, the significant estimated coefficient associated with FB implies that the increase in foreign demand for bonds reduced the interest rate. No significant autocorrelation remains. These changes indicate that a relevant, but previously omitted, variable has been introduced.
### TABLE 1

**REDUCED FORM ESTIMATES FOR INTEREST RATES**

1952:06-1979:12, SEMIANNUALLY (t-STATISTICS IN PARENTHESES)

<table>
<thead>
<tr>
<th>Estimation Method</th>
<th>Constant</th>
<th>PE12</th>
<th>M1B</th>
<th>EXG</th>
<th>DGNP₋₁</th>
<th>SUPPLY</th>
<th>FB</th>
<th>FIT</th>
<th>β</th>
<th>$R^2$</th>
<th>D.W.</th>
<th>S.E.E.</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. CORC</td>
<td>9.40</td>
<td>0.769</td>
<td>-0.797</td>
<td>6.34</td>
<td>15.08</td>
<td>---</td>
<td>-4.34</td>
<td>---</td>
<td>0.47</td>
<td>.8911</td>
<td>2.10</td>
<td>0.763</td>
</tr>
<tr>
<td></td>
<td>(1.58)</td>
<td>(3.82)</td>
<td>(−0.35)</td>
<td>(1.46)</td>
<td>(2.33)</td>
<td>(-1.22)</td>
<td>(3.98)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>2. OLSQ</td>
<td>16.88</td>
<td>1.103</td>
<td>1.988</td>
<td>5.20</td>
<td>17.80</td>
<td>-4.24</td>
<td>-6.71</td>
<td>---</td>
<td>---</td>
<td>.9126</td>
<td>1.68</td>
<td>0.694</td>
</tr>
<tr>
<td></td>
<td>(4.10)</td>
<td>(7.65)</td>
<td>(1.37)</td>
<td>(2.06)</td>
<td>(3.08)</td>
<td>(-4.90)</td>
<td>(-2.97)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>3. Tax-Adjusted OLSQ</td>
<td>12.79</td>
<td>0.727</td>
<td>1.184</td>
<td>4.00</td>
<td>12.12</td>
<td>-3.60</td>
<td>-4.69</td>
<td>---</td>
<td>---</td>
<td>.9190</td>
<td>1.76</td>
<td>0.668</td>
</tr>
<tr>
<td></td>
<td>(4.57)</td>
<td>(7.42)</td>
<td>(1.19)</td>
<td>(2.34)</td>
<td>(3.06)</td>
<td>(-6.12)</td>
<td>(-3.05)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>4. Tax-Adjusted OLSQ</td>
<td>-11.26</td>
<td>-0.711</td>
<td>-1.344</td>
<td>-3.57</td>
<td>-11.30</td>
<td>2.77</td>
<td>4.09</td>
<td>1.64</td>
<td>---</td>
<td>.9201</td>
<td>1.78</td>
<td>0.670</td>
</tr>
<tr>
<td></td>
<td>(-0.81)</td>
<td>(-0.82)</td>
<td>(−0.64)</td>
<td>(−0.74)</td>
<td>(−0.76)</td>
<td>(0.81)</td>
<td>(0.74)</td>
<td>(2.12)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>5. Tax-Adjusted OLSQ</td>
<td>20.82</td>
<td>1.190</td>
<td>1.922</td>
<td>6.47</td>
<td>19.85</td>
<td>-5.89</td>
<td>-7.73</td>
<td>-0.64</td>
<td>---</td>
<td>.9201</td>
<td>1.77</td>
<td>0.670</td>
</tr>
<tr>
<td></td>
<td>(2.05)</td>
<td>(2.04)</td>
<td>(2.04)</td>
<td>(1.87)</td>
<td>(1.95)</td>
<td>(-2.08)</td>
<td>(-1.93)</td>
<td>(-0.82)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: OLSQ is ordinary least squares. CORC is least squares with the Cochrane-Orcutt autocorrelation correction.
Row 3 is based on the income-tax adjusted specification (10). 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 six associated with the estimated coefficient of SUPPLY substantiates, in a tax-adjusted specification, the claim by Wilcox (1983) that SUPPLY is an important factor in the determination of interest rates.

The evidence presented in row 3 suggests that the tax-adjusted formulation provides a superior empirical explanation of nominal interest rate movements. The results of formal statistical tests of the hypothesis that the relevant specification includes income tax effects are presented in rows 4 and 5. These model specification tests for two nonnested alternative models explaining the same dependent variable were proposed by Davidson and MacKinnon (1981). If the null hypothesis ($H_0$) is that (10) is the true model and the alternative hypothesis ($H_1$) is that

\begin{equation}
(11) \quad i = B_0 + B_1 \text{PEI2} + B_2 \text{M1B} + B_3 \text{EXG} + B_4 \text{DGNP}_{-1} + B_5 \text{SUPPLY} + B_6 \text{FB}
\end{equation}

is the true model (i.e., equation (10) omitting taxes), the test equation is

\begin{equation}
(12) \quad i = \frac{C_0}{(1-t)} + C_1 \frac{\text{PEI2}}{(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{FIT11}
\end{equation}

where \text{FIT11} is the predicted series for $i$ based on equation (11). 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 (11) is the true model, we reestimate (12), reversing the roles of (10) and (11) (i.e. estimating (11) with the addition of FIT10 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.

Row 4 presents the test results for the null hypothesis that the non-tax-adjusted hypothesis (11) is the true model. The estimated value of $C_7$, reported in the column titled FIT, differs significantly from zero (but not from unity) enabling us to reject the non-tax-adjusted specification. Further, the estimated coefficients associated with the other explanatory variables are now statistically insignificant. Reversing the procedure generates row 5. The estimated value of $C_7$ does not differ significantly from zero. Thus the tax-adjusted model cannot be rejected. All of the other explanatory variables (except M1B) retain their statistical significance at the 10 percent level or better. The reduction in their t-statistics is related to the increased multicollinearity associated with the addition of FIT11 which is highly correlated with the other explanatory variables. These results confirm, in the presence of the previously omitted SUPPLY variable, the importance of income taxes in the determination of the nominal interest rate found by Peek (1982).

Previous investigators have obtained a wide range of estimated coefficients for the anticipated inflation rate as they altered the sample period (e.g. Carlson (1979), Cargill and Meyer (1977, 1980)). Figure 1 plots the values of the estimated coefficient of PEL2 over time for
FIGURE 1

Anticipated Inflation Coefficients From 14-Year Rolling Regressions

Center of Sample Period
the restricted (B) and for the unrestricted (BTS) specifications of
the interest rate equation. These estimates were obtained
by using a sample length of 14 years (half the sample) whose beginning
and ending dates are rolled forward one period at a time. This produces
an estimate for each model for the first half, last half, and all
contiguous halves of the total sample in between. Both series of
estimates vary over time and BTS is not obviously more stable than B.
Both rise very sharply, though 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 2 presents the results of estimating the non-SUPPLY non-tax-
adjusted specification (Table 1, row 1) and the complete model (Table 1,
row 3) 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 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 2 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, 60s and 70s was different. In particular,
those findings suggest that the interest rate reacted more strongly
to anticipated inflation in the 1960s.

In the restricted specification the estimated coefficient of PE12
jumps from .762 to 1.399 between the first and second halves of the full
### TABLE 2
INTEREST RATE ESTIMATES OVER VARIOUS SAMPLE PERIODS
ORDINARY LEAST SQUARES
(t-STATISTICS IN PARENTHESES)

<table>
<thead>
<tr>
<th>Sample Period</th>
<th>Constant</th>
<th>PE12</th>
<th>MLB</th>
<th>EXG</th>
<th>DGNP&lt;sub&gt;-1&lt;/sub&gt;</th>
<th>SUPPLY</th>
<th>FB</th>
<th>R&lt;sup&gt;2&lt;/sup&gt;</th>
<th>D.W.</th>
<th>S.E.E</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. 1952:06-1965:12</td>
<td>5.62 (0.77)</td>
<td>0.762 (1.96)</td>
<td>-1.53 (-0.82)</td>
<td>5.45 (0.97)</td>
<td>19.1 (2.47)</td>
<td>---</td>
<td>9.33 (0.63)</td>
<td>0.6503 (1.56)</td>
<td>0.686</td>
<td></td>
</tr>
<tr>
<td>2. 1966:06-1979:12</td>
<td>12.83 (1.64)</td>
<td>1.399 (4.42)</td>
<td>7.43 (2.08)</td>
<td>-1.81 (-0.27)</td>
<td>19.0 (1.64)</td>
<td>---</td>
<td>-7.70 (-2.11)</td>
<td>0.7537 (1.71)</td>
<td>0.886</td>
<td></td>
</tr>
<tr>
<td>3. 1959:06-1972:12</td>
<td>-1.97 (-0.20)</td>
<td>1.308 (4.40)</td>
<td>1.11 (0.28)</td>
<td>-5.37 (-1.13)</td>
<td>20.5 (2.30)</td>
<td>---</td>
<td>-12.40 (-3.52)</td>
<td>0.8432 (1.52)</td>
<td>0.589</td>
<td></td>
</tr>
<tr>
<td>4. 1952:06-1958:12, 1973:06-1979:12</td>
<td>13.57 (2.39)</td>
<td>1.048 (4.39)</td>
<td>3.18 (1.25)</td>
<td>6.18 (1.58)</td>
<td>18.2 (1.86)</td>
<td>---</td>
<td>3.06 (0.49)</td>
<td>0.9168 (1.47)</td>
<td>0.914</td>
<td></td>
</tr>
<tr>
<td>5. 1952:06-1965:12</td>
<td>19.02 (1.23)</td>
<td>0.701 (2.29)</td>
<td>2.23 (0.60)</td>
<td>7.31 (1.44)</td>
<td>12.6 (2.22)</td>
<td>-5.05 (-1.06)</td>
<td>7.78 (0.72)</td>
<td>0.6686 (1.66)</td>
<td>0.683</td>
<td></td>
</tr>
<tr>
<td>6. 1966:06-1979:12</td>
<td>17.54 (3.73)</td>
<td>0.663 (3.63)</td>
<td>-0.72 (-0.28)</td>
<td>9.41 (2.03)</td>
<td>8.7 (1.37)</td>
<td>-5.14 (-3.64)</td>
<td>-3.52 (-1.76)</td>
<td>0.8616 (2.12)</td>
<td>0.679</td>
<td></td>
</tr>
<tr>
<td>7. 1959:06-1972:12</td>
<td>-0.73 (-0.06)</td>
<td>0.853 (2.71)</td>
<td>0.11 (0.04)</td>
<td>-2.84 (-0.77)</td>
<td>15.4 (2.28)</td>
<td>-0.37 (-0.06)</td>
<td>-8.86 (-3.14)</td>
<td>0.8416 (1.48)</td>
<td>0.606</td>
<td></td>
</tr>
<tr>
<td>8. 1952:06-1958:12, Adjusted</td>
<td>17.89 (4.83)</td>
<td>0.735 (5.90)</td>
<td>2.12 (1.57)</td>
<td>6.73 (3.09)</td>
<td>10.1 (1.95)</td>
<td>-4.34 (-3.91)</td>
<td>0.10 (0.03)</td>
<td>0.9549 (2.11)</td>
<td>0.689</td>
<td></td>
</tr>
</tbody>
</table>
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, the difference in the estimated coefficient is relatively minor, .701 compared to .663. The change between the middle half and the first and final quarters is slightly larger, .853 compared to .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 3 displays the results of formal coefficient stability tests based on the evidence presented in Table 2. The test statistics in row 1 pertain to the mid-sample split. Those in row 2 refer to the halves of the full 1952-79 sample 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. The last two columns are derived from the complete specification. The reported F-statistic is for a standard Chow test.\(^3\) To obtain the t-statistic, a dummy variable was added to the regression and the equation was estimated over the entire sample period. For row 1, the dummy variable is zero for the first half of the sample and for the second half of the sample takes on the values of PE12 and PE12/(1-t) in columns 2 and 4, respectively. Similarly, in row 2 the dummy variable takes on the nonzero values in the 1959-72.
TABLE 3
TESTS OF COEFFICIENT STABILITY

<table>
<thead>
<tr>
<th>Subsample Dates</th>
<th>Specification</th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>(First Subsample/Second Subsample)</td>
<td>Row 1, Table 1</td>
<td>Row 3, Table 1</td>
<td></td>
</tr>
<tr>
<td></td>
<td>$\hat{F}$</td>
<td>$\hat{t}$</td>
<td>$\hat{F}$</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td>1. 1952:06-1965:12/1966:06-1979:12</td>
<td>2.00</td>
<td>-0.20</td>
<td>0.74</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>-1.17</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>-1.25</td>
</tr>
</tbody>
</table>

*Significant at the 0.05 level.
period. The reported t-statistic is associated with the estimated coefficient of the dummy variable.

In spite of the rather large shift in the point estimates of the PE12 coefficient from the first to the second half of the sample reported in rows 1 and 2 of Table 2, neither the hypothesis of vector nor of individual coefficient stability can be rejected. 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. Though the point estimates in rows 3 and 4 in Table 2 are much closer together than those in row 1 and 2, the F- and t-tests allow us to reject the stability of both the coefficients as a whole and the anticipated inflation rate coefficient in particular. The test statistic for each hypothesis is significant at the .05 level. When the same sample split is used for the tax-adjusted SUPPLY specification, we cannot reject stability, either of the vector of coefficients or of the anticipated inflation rate coefficient alone. Thus, incorporation of supply and tax effects jointly reduces the instability reported elsewhere (e.g., Carlson (1979), Cargill and Meyer (1980)) and documented here for our less-inclusive specification.

Table 4 presents values of the simple correlations between PE12 and both SUPPLY and 1/(1-t) for the subsamples considered in Tables 2 and 3. The change in the pattern of correlation between PE12 and excluded SUPPLY and 1/(1-t) is dramatic. For both variables the
TABLE 4
CORRELATION OF PE12 WITH SUPPLY AND \(1/(1-t)\)
OVER VARIOUS SUBSAMPLES

<table>
<thead>
<tr>
<th>SAMPLE PERIOD</th>
<th>SIMPLE CORRELATION BETWEEN PE12 AND</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>SUPPLY (1)</td>
</tr>
<tr>
<td>1. 1952:06-1979:12</td>
<td>-0.01</td>
</tr>
<tr>
<td>2. 1952:06-1965:12</td>
<td>-0.79</td>
</tr>
<tr>
<td>3. 1959:06-1972:12</td>
<td>-0.94</td>
</tr>
<tr>
<td>4. 1966:06-1979:12</td>
<td>0.78</td>
</tr>
<tr>
<td>5. 1952:06-1958:12,</td>
<td>-0.15</td>
</tr>
<tr>
<td>1973:06-1979:12</td>
<td>0.55</td>
</tr>
</tbody>
</table>
strong negative correlations with PE12 in the 1952-65 period give way to equally strong, positive correlations in the 1966-79 period. It is this change in the correlations over time that contributes to the temporal instability of the estimated anticipated inflation coefficient (and of the remaining coefficients as well). The large negative correlation between PE12 and SUPPLY through the early 1970s, coupled with the negative effect of SUPPLY on interest rates, adds an upward bias to the estimated coefficient of PE12. As a consequence, when we isolate the middle half of the full sample period we are able to reject the stability of the non-supply specification.

To test the hypothesis that the changing correlations between PE12 and SUPPLY and 1/(1-t) contribute to the instability of the estimated coefficient of PE12 in the restricted relative to the unrestricted specification, we regressed the ratio of B to BTS from the rolling regressions on the average level of 1/(1-t), the correlation between PE12 and 1/(1-t) (CORRPET), and the correlation between PE12 and SUPPLY (CORRPRESS), where each is calculated over the same rolling sample periods used to derive the B and BTS series. We include the mean of 1/(1-t) since this term enters the reduced-form coefficient of PE12 multiplicatively. The results of this regression were (t-statistics in parentheses):

\[
\frac{B}{BTS} = -20.1 + 15.6 \left( \frac{1}{(1-t)} \right) + 0.39 \text{ CORRPET} - 0.24 \text{ CORRPRESS} + 0.20 \hat{e}_{-1} \\
(-3.24) \quad (3.48) \quad (4.17) \quad (-1.97) \quad (1.09)
\]

\[
R^2 = .7673 \quad \text{D.W.} = 2.10 \quad \text{S.E.E.}/(B/BTS) = 0.085
\]
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.

IV. Conclusion

The magnitude, precision, and stability of estimates of the response of nominal interest rates to anticipated inflation have each come into question in the last decade as inflation rose to historically high levels. 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 coefficient instability is reduced by the inclusion of tax and supply effects. 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, though our rolling-sample results suggest there are still more factors relevant to the determination of interest rates which have not been considered here.
FOOTNOTES

1Before December 1959, when one-year Treasury bills were introduced, the interest rate measure is the yield on bills with maturities of nine to twelve months.

2Wilcox (1983) 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.

3Each of these tests is based on ordinary least squares estimates.

4The 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.

5The 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 dependent variable. Similar but more precise point estimates were obtained.
REFERENCES


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