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PITFALLS IN FISHER MODEL BUILDING:
INTEREST RATES AND INFLATION IN THE INTERWAR PERIOD

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ABSTRACT

Contrary to the postwar evidence, no response of nominal interest rates to expected inflation for the pre-World War II era has been reported heretofore. Our interest rate model illustrates why an operative Fisher effect during this period may have escaped detection. Estimates based on that model reveal a statistically significant reaction of nominal interest rates to changes in expected inflation before World War II. The reduced-form estimates for the interwar period, in fact, are very similar to what structural parameter estimates based on postwar data predict. The estimates based on interwar data are also similar to those obtained when the same specification is applied to postwar data.
I. Introduction: Review and Preview

Assessment of the reduced-form effect of expected inflation on nominal interest rates has passed through several stages. The effect was apparently little supported by the data at the time it was enunciated by Irving Fisher (1896) or for half a century afterward. Fisher (1930) himself concluded that, unless extraordinarily long expectations lags were in operation, even the response of short-term nominal interest rates to expected inflation was feeble.¹ The 1960s' and 1970s' historically high nominal interest rates convinced many that these rates reflected a significant response to those decades' higher level of expected inflation. The 1980s, however, brought renewed suggestions that nominal interest rates had rarely, if ever, adjusted reliably or by large amounts in response to expected inflation.²

We propose to demonstrate that, in spite of the stability of their underlying parameters, the empirically-relevant functional forms of the structural relations are likely to imply a substantially nonlinear reduced form for the nominal interest rate. When the economy operates over highly nonlinear parts of the system, econometric estimates of linear approximations to a reduced form may be econometrically unstable and economically misleading.

¹Fisher did not compute the size of the response; he calculated correlation coefficients. Summers (1983) has replicated Fisher's technique for generating expected inflation. He finds an economically small and statistically insignificant response of short-term nominal interest rates to that proxy for expected inflation.

²Summers (1983) and Carmichael and Stebbing (1983) argue that nominal interest rates have responded very little to changes in inflation.
Specifically, we document that the economy moved across dramatically differently sloped portions of the (parametrically stable) portfolio balance schedule (the "LM" curve) in the 1920s and 1930s. The traversing of these nonlinearities over this period generated observations which are best not characterized with linear reduced forms.

Section II shows estimates of simple and expanded linear interest rate reduced forms based on the full interwar sample period. Though the resulting coefficient estimates otherwise coincide with conventional notions about the determinants of interest rates, they confirm the widespread impression that this period exhibits virtually no effect of expected inflation on the nominal interest rate.

Section III presents econometric evidence that the demand for money was better approximated by a functional form that is linear in the logs of each of its arguments. This suggests that the reduced forms were highly and intrinsically nonlinear. The slope of the LM curve in the (i,Q) plane dictated by this functional form for the demand for money was an increasing function of the level of the interest rate. Since the reduced-form response of nominal interest rates to expected inflation is a function of that slope, that reduced-form response fell as interest rates fell. Estimates of constant-coefficient, linear approximations are therefore likely to be misleading and unstable when they cover the full interwar period, over which substantial interest rate movement took place.

Section IV presents the results of estimating an interest rate
reduced form over interwar sub-samples distinguished by differing mean interest rate levels to show how sensitive estimates can be to such nonlinearities. It also documents the similarity of the estimates for the 1920s to those for the postwar period. Confirming evidence from estimates of the reduced form for output is also introduced. Section V concludes.

II. Functional Forms and Reduced Forms

The presence of nonlinear transformations of variables in structural relations may preclude consistent estimation of reduced forms. To illustrate this point, we initially specify a simple, conventional model of the macroeconomy in which each variable enters in levels. Real expenditure (Q) is a function of the expected, real, after-tax interest rate (r*) and the expected condition of the economy. We proxy the latter with ILI, an "index of leading economic indicators" for the interwar period. This index has been constructed from data series that were available during this period by mimicking as closely as possible the choices of data series and procedures the government has used to construct the index of leading economic indicators for the postwar period.¹

The expenditure function also contains a "liquidity effect". On the presumption that financial markets adjust more quickly than the real side of the economy, a shift in monetary policy will initially generate a liquidity effect on interest rates. One way

³The model is similar to those used in Peek and Wilcox (1983, 1984).

¹Details of the construction of ILI are given in Appendix B. The values for the index are available from the authors on request.
this effect can be viewed as a transitory shift of the IS function in conjunction with an LM shift of the same vertical distance.\textsuperscript{5} To allow for such an effect, we include in the expenditure function the first-difference of the real money supply ($d(M/P)$). Real expenditure is taken to be equal to output. Thus, the expenditure function ("IS curve"), with disturbance term, $u$, can be written:\textsuperscript{6}

\begin{equation}
Q/QN = E(r^*, \text{ILI}, d(M/P), u).
\end{equation}

The portfolio balance relation ("LM curve") equates the supply of real balances ($M/P$) to real money demand, which depends on real income, the after-tax, nominal interest rate ($i$), and a disturbance term, $e$:

\begin{equation}
M/P = L(Q, i, e).
\end{equation}

The expected, real, after-tax interest rate is defined as the after-tax, nominal rate minus the expected inflation rate over the same horizon ($\pi^*$):

\begin{equation}
\hat{r}^* = i - \pi^*.
\end{equation}

The current nominal interest rate is affected by the current expectation of future inflation. The inflation rate that is currently expected to occur in the future is taken to be a function of current and past conditions. On the other hand, the current-quarter price level is assumed to be pre-determined for the current

\textsuperscript{5}This argument was laid out in Peek and Wilcox (1986). An earlier example of this effect is in Carlson (1979).

\textsuperscript{6}ILI is defined as the index of leading indicators normalized by natural real GNP ($QN$). In the expenditure function, both the index and real GNP are normalized by $QN$ to render each of the variables scale invariant.
period, depending only on past conditions.

The reduced forms for the nominal interest rate and for output implied by (1)-(3) are given by:  

\[ i = f(\pi*, \text{ILI}, d(M/P), M/P, QN, u, e) \]  

\[ Q = g(\pi*, \text{ILI}, d(M/P), M/P, QN, u, e). \]

The reduced form for the level of the nominal interest rate can be expressed as a linear (in the parameters) function of the pre-determined variables when the real and nominal interest rates appear only in level form (as opposed to in logs, reciprocals, and so on) and each endogenous variable enters subject to only one transformation in the system (for example, always as logs). For example, suppose that in the IS function the log of output (relative to its natural rate) depends on the difference between the level of the after-tax, nominal interest rate and the level of expected inflation, that is, on the level of the expected, real, after-tax interest rate (1'):

\[ \log(Q/QN) = a_0 - a_1*\pi - a_2*d(M/P)' + a_3*\text{ILI} + e. \]

Suppose the portfolio balance equation reflects the equating of the pre-determined real money supply and a real money demand function that is linear (in the parameters) when the money and

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7Cecchetti (1987) and Childs (1947) present details concerning the tax-status of these Treasury issues during this period. On the basis of their evidence and arguments, we take the relevant marginal income tax rate for these issues during this period to be zero. Henceforth, we will refer only to the interest rate.

'This was the case in Peek and Wilcox (1983, 1984). The functional form of the exogenous variables is irrelevant for our argument. Specifying their form is, of course, relevant to obtaining satisfactory estimates.
income variables are expressed in logs and the interest rate is expressed in levels (2'):

\[ (2') \log(M/P) = b_0 + b_1 \log(Q) - b_2 i + u \]

or, equivalently:

\[ (2'') \log(M) - \log(P) - b_1 \log(QN) = b_0 + b_1 \log(Q/QN) - b_2 i + u. \]

This system has a constant-coefficient, linear (in the coefficients, not the variables) reduced form for the nominal, after-tax interest rate (4'):

\[ (4') i = (b_0 + a_0 b_1 + a_1 b_1 \pi^* - (M/P)^* - a_2 b_1 d(M/P)^* + a_3 b_1 IIL + u + b_1 e)/(a_1 b_1 + b_2) \]

where \((M/P)^*\) is defined to be \(\log(M) - \log(P) - b_1 \log(QN)\) and \(d(M/P)^*\) is its first difference.

Estimation of the reduced form for the nominal interest rate, (4'), requires a proxy for the rate of inflation expected to occur from the issuance until maturity of the Treasury issue, \(\pi^*\). As the proxy for expected inflation, we take the vector of fitted values from the following forecasting equation:

\[ (6) \quad \pi_{t+1} = \pi(IIL, IIL_{t-1}, (M/P)^*, d(M/P)^*, i_{t-1}). \]

This forecasting equation is specified such that the current expectation of the future inflation rate is a function of current and lagged variables. \(\pi_{t+1}\) is the annualized, future (i.e., over the ensuing three months) consumer price inflation rate.* IIL and IIL_{t-1} are the current and three-month-lagged values of the

*We use the National Industrial Conference Board's measure of consumer prices, as reported in Sayre (1948), which unlike the GNP deflator, is available monthly. The results were little changed when the BLS measure of the monthly CPI was substituted. See Cecchetti (1989) for a discussion of the two series.
(normalized) index of leading economic indicators that we have constructed. \( (M/P)^* \) is \( (\log(M) - \log(P) - 0.75 \log(QN)) \).\(^{10}\) \( d(M/P)^* \) is the first difference of \( (M/P)^* \). i.e., is the three-month-lagged value of the nominal interest rate.

Table 1 shows the results of estimating (6) over the full sample and over the first and second halves of the sample. This sample split corresponds to the periods of above and below average interest rates. The most noticeable feature of table 1 may be the extent to which the estimated coefficients shift. In fact, a standard F-test allows us to reject the hypothesis of coefficient stability over the mid-sample split. What can be seen is that the statistically optimal forecasting equation can differ considerably over time. The prediction of this apparent instability and its causes will be addressed in what follows.

Table 2 contains estimates of truncated and complete versions of the linear interest rate reduced form (4'). In each case, we estimated (4') over the entire interwar sample period.\(^{11}\) The data frequency is quarterly. The estimation technique for each row was

\(^{10}\)M is the nominal money supply. \( P \) is the implicit GNP deflator. We use 0.75 as the estimate of \( b1 \). This is the estimated long-run income elasticity of money demand over the 1922:1-1941:2 period. (See tables 3a and 3b.) This adjusts for the long-run, upward drift in velocity produced by the less-than-unity income elasticity of money demand (and perhaps by technological advance). Estimates of the money demand function in the form given by Goldfeld (1973), with real balances on the left-hand-side and their lag on the right-hand-side along with real income and nominal interest rates produced an implied long-run income elasticity of 0.76. Complete definitions of each of the series are given in the Data Appendixes.

\(^{11}\)The endpoints were chosen to avoid the effects of the two world wars.
either ordinary least squares (OLS) or maximum likelihood in the presence of first-order autoregressive error terms (AR1). The dependent variable in table 2 is the nominal yield on short-term Treasury issues (i).

The interest rate reduced-form estimates contain a regression-based proxy for the inflation expected to take place over the upcoming period. We take the fitted values from row 1 of table 1 as the expected inflation proxy, termed PIFIT. As Pagan (1984) has shown, when the regressors in the focus (interest rate) equation are a subset of those used to generate fitted values and when the estimates are to be evaluated under the null that the coefficient on this generated regressor is zero, OLS produces estimates of the focus equation coefficients that are consistent, are fully efficient, and permit valid inference. These conditions are satisfied herein.

The first three rows of table 2 show that the simple relation between expected inflation and interest rates during the interwar period is economically and statistically insignificant, regardless of whether we introduce an autoregressive disturbance term correction or a lagged dependent variable (for which we report the Durbin h-statistic to test for an autocorrelated disturbance term). In each case, only the lagged error or the lagged dependent variable is significant.

The bottom three rows add the remaining presumed determinants of real interest rates. In each row, both ILI and \((M/P)^*\) carry the expected signs and are significant. (These two variables have had
their respective sample means subtracted to enable us to estimate the "steady-state" zero-expected-inflation real rate of interest using the estimates of the constant term and of the lagged dependent variable.) Each variable has nearly the same estimated long-run effect across these specifications. These specifications provide only weak evidence that liquidity effects proxied by \( d(M/P) \) were significant. The expected inflation proxy, PIFIT, is now significant until we add the lagged dependent variable, but in both cases it is negative and quite small. These estimates suggest that real interest rates were systematically driven by expenditure function and monetary shifts. There is no evidence, however, that nominal rates rose with expected inflation.

The justification for estimating (4') is provided by the assumed functional forms in (1') and (2'). The LM curve implied by this functional form is plotted as LM₁ in figure 1.¹² Note that it has a curvature opposite to that often presumed. Suppose, on the other hand, the demand for money has constant parameters when each of its variables is expressed in logs.¹³ Then, in (2'), (4'), and (5'), b₂ is replaced by \( b₂' \times \log(i)/i \), where \( b₂' \) is the (constant) interest elasticity of money demand and \( \log(i) \) is the log of the nominal interest rate. The LM curve implied by this specification is labelled LM₂ in figure 1. The reduced-form equations for the

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¹²To facilitate comparison of the 1920s to the 1930s, real GNP has been standardized by natural real GNP.

¹³We assume that, in absolute value, the income elasticity exceeds the interest elasticity of money demand, as is almost universally thought and estimated.
nominal interest rate and for output are now intrinsically nonlinear; neither endogenous variable can be expressed as a function solely of exogenous variables and a separable disturbance term. Equation (4'), for example, becomes

$$8 \quad \alpha_1 b_1 i + b_2' \log(i) = (b_0 + a_0 b_1) + a_1 b_1 \pi^* - a_2 b_1 d(M/P)^* - (M/P)^* + a_3 b_1 \text{ILI} + b_1 e + u.$$  

The reduced-form coefficients, or total derivatives with respect to the exogenous variables, are now each functions of the system's endogenous variables, specifically the nominal interest rate. The reduced-form response of the interest rate to expected inflation, \( \frac{\text{di}}{\text{d}\pi^*} \), for example, is

$$9 \quad \frac{\text{di}}{\text{d}\pi^*} = \frac{\alpha_1 b_1}{\alpha_1 b_1 + (b_2'/i)}.$$  

As interest rates fall, the "reduced-form" response of interest rates to each of the exogenous factors in (4') falls. The corresponding "reduced-form" response of output to money falls and the remaining responses rise. Given this model specification, the "reduced-form" responses will be constant only if the nominal interest rate is constant, which would render the interest rate function, the focus of our concern, non-estimable (and uninteresting). Indeed, the linear-in-the-logs specification of the money demand function precludes the possibility of constant "reduced-form" coefficients.

We can ameliorate, but not completely eliminate, the problems associated with the intrinsic nonlinearity implied by (8) by splitting the sample into halves, which are distinguished by being either high or low interest rate eras. Such sample halves
correspond approximately to the 1920s and the 1930s, respectively. The irony is that only if the interest rate varies can we estimate (4'), but it is precisely that variation that renders linear estimation inappropriate.

An implication of the changing "reduced-form" coefficients is that both (4') and (5') are characterized by heteroskedastic errors. The disturbance term associated with the interest rate relation given in (4') when b2 is replaced by b2'log(i)/i is:

\[ z = (u + b1*e)/(a1b1 + (b2'/i)). \]

Even if we assume that a1 and b1 (and b2') are constant, the variance of this composite error term varies with the interest rate when specified as discussed above. To test for the presence of heteroskedastic errors in (4'), we estimate a modified version of the specification used for row 6 of table 2. The modification entails adding to that specification six variables (denoted by primes), each of which is the product of one of the original variables and a dummy variable which takes on the value one for the observations after 1932:1 and is zero otherwise. The resulting specification allows for different coefficients during the high and low interest rate portions of the sample, thereby reducing the effect that the reduced-form coefficient shifts implied by (7) have on the estimation of the disturbance terms.

The modification also entails moving to an IV technique whereby the generated expected inflation proxy, PIFIT, is replaced by the actual future values of inflation, PI, and PIFIT is used as an instrument. Pagan (1984) concludes that this procedure will
produce the appropriate estimates of the error terms when the null hypothesis is no longer that the coefficient on expected inflation is zero, as it was for the estimates in table 2. The results are (t-statistics in parentheses):

\[(11)i = 1.43 + 0.492i_{x} + 0.106\pi + 0.070\text{ILI} - 0.046d(M/P) - 8.82(M/P)^{2}
\]
\[
(1.44) (1.30) (1.22) (0.59) (1.42) (1.00)
\]
\[-1.78D + 0.951i_{x} + 0.091\pi - 0.229\text{ILI} - 0.042d(M/P)^{2}
\]
\[
(1.74) (1.16) (1.04) (1.39) (1.24)
\]
\[+ 10.3(M/P)^{2}
\]
\[
(1.16)
\]

where D is the product of the constant term and the dummy variable and \(x'\) is variable x times the dummy variable.

The sample residual variances of this regression for the first and second halves of the sample period were 0.29 and 0.08, respectively. An F-test of the hypothesis that the variances in the first and second halves of the sample were equal allows us to reject the null hypothesis of homoskedastic errors. Equation (10) further suggests the form that heteroskedasticity might take. To assess whether the variance of the reduced-form disturbance term systematically moves according to (10), the absolute values of the residuals from (11) are regressed on the interest rate and its square. This specification allows the data to choose whether the most appropriate way to characterize the disturbance variance is as a positive and concave function of the interest rate. The results of that OLS regression over the full sample period are (t-statistics in parentheses):

\[(12) \quad |z| = 0.11 + 0.277i_{x} - 0.052i_{x}^{2}.
\]
\[
(2.15) (3.63) (-2.76)
\]
Thus we have evidence not only that the error variances differ across high and low interest rate eras, but also that that heteroskedasticity follows the pattern suggested by the existence of a linear-in-the-logs money demand function.

III. Money Demand Function Estimates

In this section, we make the case that the empirically-relevant functional form for the interest rate in the demand for money is closer to being logarithmic, and thus nonlinear (in the level of the interest rate), than to being linear. The case is made on the basis of the observed instability of the coefficients on linear interest rate terms and on the basis of estimates of nonlinear specifications. Tables 3a and 3b present instrumental variables estimates of money demand functions.\textsuperscript{14} In each case, the nominal interest rate was regressed on the real money stock \((M/P)\) and on real GNP \((Q)\), each measured as a quarterly average.\textsuperscript{15}

The top three rows specify both interest rates and their determinants (the real money stock and real GNP) in levels; the

\textsuperscript{14}The real money stock is taken to be econometrically pre-determined. Real GNP is endogenous. For each set of estimates the instruments were a constant term and the levels and logs of current and one-period lagged real money stock, the log and level of the lagged nominal interest rate, the level and log of lagged real GNP, plus the pre-determined variables used in the inflation forecasting equation. This same list of instruments was used for each of the instrumental variables estimates in this article.

\textsuperscript{15}None of these variables were standardized by the real natural rate of output. Nor were the data in equation (13) below. The alternative Romer (1988) real GNP series very closely tracked the series we used. For example, the standard deviations of the annual growth rates for our series and the Romer series are 4.50\% and 4.53\% during the decade of the 1920s. The correlations between the levels of these real GNP series and between their growth rates were each above 0.99.
dependent variable in the middle three rows is the level of the interest rate and logs of the determinants are included; the bottom three rows specify each variable in logs. For each specification, we present estimates for the full sample period, as well as for the first and second halves of that full sample period. Table 3b differs from table 3a in that it shows autocorrelated-residual-corrected results.\textsuperscript{16}

A priori we anticipated the real GNP and real money supply coefficients to be positive and negative, respectively, consistent with the elasticities of money demand with respect to real GNP and the nominal interest rate being positive and negative, respectively. In table 3a, the functional forms that include the level of the interest rate (rows 1 - 6) produce estimates of money coefficients that vary considerably in size and significance across sample periods. The pattern of lower interest rate coefficients (for each of the interest rate level specifications) during the lower interest rate period of the 1930s is suggestive of a piecewise-linear approximation to the log-of-the-interest-rate specification. The real income coefficient estimates vary

\textsuperscript{16} Table 3b was estimated by instrumental variables using transformed variables. The money demand variables were each transformed by \((1-\varphi B)\), where \(\varphi\) is the estimated first-order-autoregressive-residual coefficient estimated from the full sample and B is the backshift operator. For the log-log specification, the autoregressive coefficient was 0.44; for the level-log specification, it was 0.76; and for the level-level specification, it was 0.89.

Specifications that allow for stock adjustment of money demand often include the lagged money stock. In no case was the lag of the money stock significant here.
appreciably not only in size, but also in significance, even changing signs across samples. The estimated interest elasticities in the log specifications (rows 7 - 9), on the other hand, change relatively little and remain statistically significant across sample periods.\textsuperscript{17} Similar conclusions emerge from table 3b.

We also employed the Box-Cox technique to allow the data to select the most appropriate functional form (common to all variables) for the demand for money. To do so, we estimated (13) with a nonlinear, instrumental variables technique (t-statistics in parentheses):

\begin{align*}
(13) \quad i' &= 18.5 + 5.73Q' - 6.77M'. \\
(1.75) \quad (1.95) \quad (1.50)
\end{align*}

where each variable, \(x\), has been transformed by \(x' = (x^\alpha - 1)/\alpha\). The point estimate for \(\alpha\) is 0.068 (standard error = 0.090). Here, \(\alpha\) and the remaining coefficients are estimated simultaneously.\textsuperscript{18} The

\textsuperscript{17}The elasticities from the full-sample, log-log specification are strikingly similar to those found by Goldfeld(1973). Our implied long-run income and interest elasticities, for example, are 0.75 and -0.09. His are 0.68 and -0.11. He includes two interest rates and thus his average interest elasticity should perhaps be doubled and then compared to ours. However, since his sample period consisted almost entirely of a time when time deposit interest rates were at regulatory ceilings, it is not straightforward to assess how much each of his interest rates changed when the open market rates changed. Nonetheless, our estimates parallel his. Given the fate of empirical estimates of money demand over the last fifteen years, however, that may not be very comforting.

Goldfeld apparently considered the issue of the functional form for interest rates sufficiently settled that, in spite of the exhaustive range of his study, he did not discuss alternative transformations to the log-log specification.

\textsuperscript{18}Estimation of this equation produced a Durbin-Watson statistic of 1.10. The already very considerable extent of nonlinearity, coupled with the use of an instrumental variables
remaining coefficients are not estimated very precisely, but \( \alpha \) is. Its two-standard deviation range is fairly narrow: -0.10 to 0.25. Chi-square tests decisively reject \( \alpha = 1 \), but do not reject \( \alpha = 0 \). \( x' \) approaches the log of \( x \) as \( \alpha \) approaches zero and \( x' \) approaches \( x \) itself as \( \alpha \) approaches one. Thus, the linear functional form is ruled out, while the data freely choose a form that is almost exactly linear in the logs.

We also tested for the appropriate functional form of the interest rate alone by estimating the money demand function with all variables, except the interest rate, in logs. For the interest rate variable, the unconstrained estimate of \( \alpha \) in that case was 0.11 (standard error = 0.11). Chi-square tests again decisively rejected \( \alpha = 1 \), but could not reject \( \alpha = 0 \). Thus we again reject the linear, but not the log, form of the interest rate in the money demand function.

Finally, we obtained estimates for the money demand function where the nominal interest rate was transformed with the Box-Cox formulation, while the real money stock and real GNP entered in level form. The estimated value of \( \alpha \) there (for the interest rate alone), 0.022 (standard error = 0.12), was similar to those reported above.

The implication of the estimates of these various specifications is that the data point to the log of interest rates as entering the portfolio balance schedule, suggesting that the LM technique, precluded simultaneously estimating the residual autocorrelation coefficient.
curve looks like LM, in figure 1. Rather convincingly, the data rule out the otherwise convenient specification of the level of interest rates entering the demand for money and thereby delivering a linear (in the parameters) reduced form for interest rates.

IV. Split-sample Estimates

Based on the results of tables 3a and 3b and on the results of estimating Box-Cox specifications, we re-estimated the interest rate reduced forms given by (4') using the same sample splits. PIFIT for each sub-sample is based on the appropriate row of table 1. Table 4 also reports estimates for the postwar samples of 1952:1-1979:2 and 1960:1-1969:4, periods in which previous studies have found sizable Fisher effects.¹⁻¹

The results are strikingly different from the full-sample results in table 2: The impact of expected inflation during the first half of the sample ("the 1920s") is now generally significant; the lagged effects are much less pronounced; and the liquidity effects are estimated to be negative. The estimates for the second half of the sample ("the 1930s") look much like those for the entire sample period. Insignificant coefficients are pervasive and expected inflation, the variable with the coefficient closest to being statistically significant, carries a negative sign.

But this is precisely what we should have anticipated based on the money demand estimates shown above. The higher interest rates

¹⁻¹The postwar estimates are based on after-tax values for interest rates.
of the 1920s compared to those in the 1930s should have produced larger interest rate reduced-form coefficients for each of the exogenous variables in the 1920s compared to the 1930s. Table 4 exhibits exactly that feature: across the board, the short-run and long-run coefficient estimates are larger for the 1920s than for the 1930s.

An alternative specification of the IS curve might separately include the log of the nominal interest rate and the level of the expected inflation rate. (Separately entering the log of the expected inflation rate is ruled out at least for this era by the fact that expected inflation is sometimes negative.) Such a specification implies that nominal interest rates or inflation rates (or both) have effects apart from those captured by expected real interest rates. This alternative does lead to a constant coefficient reduced form for the log of the nominal interest rate. The level of expected inflation still enters linearly.

When we estimated that specification using the full interwar sample, the estimated expected inflation coefficient was significantly negative. Under this specification, the reduced-form effect of expected inflation on the level of the nominal interest rate, \( \frac{di}{d\pi^*} \), equals the level of the interest rate times the expected inflation coefficient. The significantly positive coefficient based on the 1920s sample period implies an average \( \frac{di}{d\pi^*} \) of 0.23. The coefficient estimated from the 1930s data, on the other hand, is insignificantly positive. The estimates of this specification led to rejection of the null hypothesis of
coefficient stability over the mid-sample split. Thus, this specification of the IS curve also delivers the general pattern of a declining response and coefficient instability over the interwar period as well.

The estimates of (4') based on postwar data are given in rows 9 and 10. These estimates are similar to those for the 1920s sample. The estimated, long-run, expected inflation coefficients from the 1920s, the postwar period as a whole, and the 1960s are 0.18, 0.30, and 0.38, respectively. Thus, when we direct our attention to periods when the economy was operating along similar portions of its LM curves, we obtain similar, though hardly identical, reduced-form estimates.20

We can also use postwar estimates to generate predictions of the expected inflation effect during the 1920s and 1930s. If we take the estimate of the regression-based-proxy expected inflation effect on after-tax interest rates based on line 10 of table 4 (0.38), the average Goldfeld (1973) estimate of the long-run interest rate elasticity of money demand (0.113), and a 1960s'

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20 These postwar Fisher-effect estimates may seem surprisingly small. The postwar estimates shown here use the same method of constructing an expected inflation proxy that we use for the interwar period. The difference between using a regression-based expected inflation data series and what we have regarded to be the superior Livingston Survey data series can be dramatic. In an earlier paper (1984), we showed that the estimated expected inflation coefficient declined by half (0.654 to 0.330) when we substituted a similar regression-based expected inflation proxy for the Livingston Survey data series and kept all other variables the same. That estimate is similar to the postwar estimate shown here. Since the remaining interwar exogenous variables are not precisely the same in definition or scaling as those for the postwar period, their coefficients are not directly comparable to those based on postwar data.
average Treasury bill nominal interest rate of 4 percent, the value of \( a_{bl} \) implied by (9) is 0.0173. Thus,

\[
(14) \quad \frac{di}{d\pi^*} = \frac{0.0173}{(0.0173 + (0.113/i))}.
\]

Inserting the 1920s and 1930s mean values of the nominal interest rate of 3.10 and 0.22 percent, respectively, into (14) we get estimates of \( \frac{di}{d\pi^*} \) for those sub-periods of 0.32 and 0.033, predictions which are not too dissimilar to those presented in table 4. Thus, based on postwar parameter estimates, we would expect to find estimated expected inflation coefficients for the 1920s in the range of one-third, and for the 1930s in the range of zero, even when the underlying parameters were the same in the 1920s and 1930s as they were in the 1960s. It seems likely that there have been changes over the past half-century in some, if not all, of the structural parameters that determine reduced-form coefficients, but the proximity of the estimated coefficients across sample periods is noteworthy.

Figure 2 shows the location and shape of the LM curves implied by the money demand function estimate presented in row 7 of table 3a. The average location of the LM curve for the 1920s (LM20s) and similarly for the 1930s (LM30s) is plotted based on the average level of the real money supply during that period.\(^{21}\) On each LM curve, the point consistent with the average level of the nominal interest rate is marked with an asterisk. Figure 2 shows that not only was the real money supply larger on average during the 1930s

\(^{21}\)To compare the LM curves of the different decades, we have standardized real GNP by the average level of natural real GNP in each decade.
than it was during the 1920s, leading to a rightward (or downward) shift of the LM curve, but that the economy operated at considerably lower real GNP and nominal interest rate levels. It therefore operated primarily along much flatter-sloped portions of the LM curve. The difference in the relevant LM curve slopes across decades is dramatic: the average slope of the points observed on the LM curve in the 1920s was 14 times as large as it was in the 1930s.

Figure 3 plots the course of $\frac{d_i}{d\pi^*}$ implied by the movement of the nominal interest rate over the interwar period. The value of albl is calculated using (9), the interest elasticity of money demand from row 7 of table 3a, the expected inflation coefficient from row 3 of table 4, and the 1920s average interest rate of 3.10 percent.\(^{22}\) That value is then used in (9) to calculate the time series of $\frac{d_i}{d\pi^*}$. The interwar interest rate range of 0.01 to 4.80 percent implies $\frac{d_i}{d\pi^*}$ values ranged from a low of 0.001 to a high of 0.257. Therefore it is not surprising the full interwar sample tends to produce statistically insignificant and economically puzzling point estimates for the expected inflation coefficient. The surprise perhaps is that, although each of the reduced-form coefficients is subject to this kind of variability, the full interwar sample produces estimates (for example, those in row 6 of table 2) that otherwise appear to be so reasonable.

To the extent that the LM curve was flatter due to the lower interest rates of the 1930s compared to the 1920s, the impact of

\(^{22}\)In this case, $\frac{d_i}{d\pi^*}=0.0063/(0.0063+(0.0893/i))$. 
given (horizontal) IS curve shifts should have translated into larger changes in output and given (horizontal) LM curve shifts should have translated into smaller changes in output. Table 5 presents estimates of the reduced form for output for the full-and half-sample periods. When we compare across sub-periods, the implied long-run point estimates do coincide with the interest rate results to some extent. The ILI effects are estimated to be considerably larger and more significant during the flatter LM period of the 1930s. The liquidity variable was introduced to capture the effectively vertical short-run IS curve, and its effect is uniformly insignificant, as we would expect. On the other hand, contrary to what we would have expected, the coefficients on the money supply variable are larger and the expected inflation coefficient is smaller in the latter period.

V. Conclusion

Long ago, Brainard and Tobin (1968) argued "for the importance of explicit recognition of the essential interdependences of markets in theoretical and empirical specifications of financial models. Failure to respect some elementary interrelationships . . . can result in inadvertent but serious errors of econometric inference. . . ." From them, we have taken not only the title of this paper, but inspiration. Whereas Brainard and Tobin focussed on the issues raised by lags, we focussed on the issues raised by logs. We have attempted to show that serious errors can result from inattention to the implications of nonlinearities. Specifically, we directed our attention to nonlinearities generated
by the interaction of relations with different functional forms.

Our handling of these issues has been imperfect. Though splitting the estimation period into high and low interest rate segments ameliorates the problems produced by nonlinearities, it does not completely remove them. It does reduce the degree of reduced-form coefficient variability within the estimation period, but the reduced-form coefficients implied by this model are not consistently estimable.

The empirical results suggest that the underlying structure of the aggregate demand side of the interwar macroeconomy may have been little different from that of the postwar economy. Predictions of the interwar coefficients based on postwar estimates turned out to be very close to those we estimated from interwar data. Indeed, interwar and postwar estimates of the same relations looked very similar.

The empirical results also serve to remind us that, no matter how closely estimates conform to our priors, they may (both) be badly flawed. It is natural enough to not probe as relentlessly when regression estimates coincide with our desired or presumed results. It can also lead to serious errors. With regard to the effect of expected inflation on interest rates in particular, we saw how the evidence on the Fisher effect points to its being as operative before, as after, World War II.
REFERENCES


DATA APPENDIX A

Unless otherwise noted, data are seasonally adjusted, quarterly averages for 1921-1941.

Interest Rate The not-seasonally-adjusted, last-month-of-the-quarter average nominal interest rate is a bond-equivalent annual yield. For 1921-1930, the yield on three-to-six-month U.S. Treasury notes is used. The new-issue, three-month Treasury bill yield is used for 1931-1933. For 1934-1941, the dealers' quotation on three-month Treasury bills is used. The data come from the Board of Governors of the Federal Reserve System's Banking and Monetary Statistics, 1914-1941, page 460.

Inflation The not-seasonally-adjusted measure of consumer prices from Sayre (1949) is used to calculate the inflation rate from end-of-quarter-month to end-of-quarter-month. This rate is annualized by multiplying that growth rate in percentage points by 4. The NRA-induced inflation effects estimated in Gordon (1981) are then subtracted.


Real GNP Real GNP is taken from Gordon and Wilcox (1981).

Natural Real GNP Natural real GNP is taken from Gordon (1984), Appendix B. The second quarter observations are set equal to the annual observation; the remaining quarters' values are the result of linear interpolation between the second quarter values.

GNP Deflator The GNP deflator is taken from Gordon and Wilcox (1981).
DATA APPENDIX B

Index of Leading Economic Indicators for the Interwar Period

We constructed an index of leading economic indicators (ILI) for the interwar period basically following the method currently used by the Commerce Department to construct the ILI. The monthly growth rate of each series was standardized by its mean absolute monthly growth rate. The monthly growth rate of the composite index is the sum of those standardized growth rates. Monthly levels of the ILI were next calculated. Quarterly averages of the monthly levels were then obtained.

All data came from Moore (1961). We use seven component series, each of which was classified as a leading indicator in Moore (1961) and was available for the entire interwar period. Some of the series required (ratio) splicing. The seven series, which correspond closely to components currently included in the ILI are: average manufacturing workweek, the wholesale price index, square feet of residential building contracts, the Dow-Jones Industrial average, liabilities of failed industrial and commercial businesses, and real new orders for durable goods. Each was seasonally adjusted, except for the DJIA and the WPI.
FIGURE 1

LM CURVES IMPLIED BY DIFFERENT
FUNCTIONAL FORMS FOR MONEY DEMAND
FIGURE 2

THE AVERAGE POSITION OF LM CURVES IN THE 1920S AND 1930S
FIGURE 3

THE MOVEMENT OF THE ESTIMATED EXPECTED INFLATION COEFFICIENT

1922 - 1941
<table>
<thead>
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<th>Sample Period</th>
<th>Coefficient on</th>
<th>( R^2 )</th>
<th>S.E.E.</th>
<th>D.W.</th>
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<td>0.2536</td>
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<td>1.76</td>
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<tr>
<td></td>
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<td></td>
<td>ILI_1: -4.41, (3.02)</td>
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<tr>
<td></td>
<td>((M/P)^*): -1.4, (0.13)</td>
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</tr>
<tr>
<td></td>
<td>(d(M/P)^*): 0.163, (1.78)</td>
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<tr>
<td></td>
<td>(i_{-1}): -0.81, (0.85)</td>
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<td>1922:1-32:1</td>
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<td>0.5261</td>
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<td>ILI: 3.30, (1.96)</td>
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<td>ILI_1: -3.78, (2.31)</td>
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<tr>
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<td>((M/P)^*): 77.3, (1.84)</td>
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<td>(d(M/P)^*): 0.151, (1.05)</td>
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<td>(i_{-1}): 3.05, (2.29)</td>
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<td>ILI_1: -1.30, (0.57)</td>
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<td></td>
<td>((M/P)^*): -27.6, (1.85)</td>
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<td>(d(M/P)^*): 0.279, (2.22)</td>
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<td>(i_{-1}): 4.13, (1.56)</td>
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<td>Constant</td>
<td>i₁</td>
<td>PIFIT</td>
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TABLE 2

INTEREST RATE EQUATION ESTIMATES

(t-statistics in parentheses)

Quarterly, 1922:1 - 1941:2
### TABLE 3a

**MOONEY DEMAND FUNCTION ESTIMATES**

Instrumental Variables Estimation  
dependent variable: nominal interest rate  
(t-statistics in parentheses)

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<thead>
<tr>
<th>Functional Form</th>
<th>Sample Period</th>
<th>Coefficient on</th>
<th>R²</th>
<th>S.E.E.</th>
<th>D.W.</th>
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<td>M/P</td>
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<td>-0.0114</td>
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<td></td>
<td>(0.21)</td>
<td>(10.94)</td>
<td>(16.02)</td>
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<td>(4.63)</td>
<td>(6.75)</td>
<td>(7.01)</td>
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<td>0.0002</td>
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<td>4. level-log</td>
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<td>(0.37)</td>
<td>(2.56)</td>
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# TABLE 3b

MONEY DEMAND FUNCTION ESTIMATES

Instrumental Variables Estimation
dependent variable: nominal interest rate
(t-statistics in parentheses)

<table>
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<th>Functional Form</th>
<th>Sample Period</th>
<th>Coefficient on</th>
<th>R²</th>
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<td>M/P</td>
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<td>1. level-level</td>
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<td>1.58 (0.92)</td>
<td>0.020 (3.01)</td>
<td>-0.0064 (3.47)</td>
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<td>-0.03</td>
<td>0.009</td>
<td>0.001</td>
<td>0.251</td>
</tr>
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<td></td>
<td>(0.41)</td>
<td>(0.00)</td>
<td>(0.27)</td>
<td>(0.02)</td>
<td>(5.48)</td>
</tr>
<tr>
<td>9. 1952:1-79:2</td>
<td>0.69</td>
<td>0.632</td>
<td>0.11</td>
<td>0.554</td>
<td>0.251</td>
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<tr>
<td></td>
<td>(5.03)</td>
<td>(10.25)</td>
<td>(4.38)</td>
<td>(6.69)</td>
<td>(5.48)</td>
</tr>
<tr>
<td>10. 1960:1-69:4</td>
<td>1.21</td>
<td>0.353</td>
<td>0.24</td>
<td>0.317</td>
<td>0.251</td>
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<tr>
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<td>(4.97)</td>
<td>(3.46)</td>
<td>(4.61)</td>
<td>(2.19)</td>
<td>(5.48)</td>
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</table>
### TABLE 5

**OUTPUT EQUATIONS**

Ordinary Least Squares

(t-statistics in parentheses)

<table>
<thead>
<tr>
<th>Sample Period</th>
<th>Coefficient on</th>
<th>R²</th>
<th>S.E.E</th>
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</thead>
<tbody>
<tr>
<td></td>
<td>Constant</td>
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<td></td>
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<tr>
<td></td>
<td>QQN₀.₁</td>
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<tr>
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<td>PIFIT</td>
<td></td>
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</tr>
<tr>
<td></td>
<td>ILI</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>d(M/P)*</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(M/P)*</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1. 1922:1-41:2</td>
<td>-0.035</td>
<td>0.9616</td>
<td>0.033</td>
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<tr>
<td></td>
<td>(2.87)</td>
<td>(10.59)</td>
<td>(2.29)</td>
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<td>2. 1922:1-32:1</td>
<td>-0.005</td>
<td>0.9289</td>
<td>0.024</td>
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<td>(0.40)</td>
<td>(8.71)</td>
<td>(2.53)</td>
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<td>3. 1932:2-41:2</td>
<td>-0.181</td>
<td>0.9342</td>
<td>0.029</td>
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<td>(6.01)</td>
<td>(2.06)</td>
<td>(1.63)</td>
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<table>
<thead>
<tr>
<th>R²</th>
<th>S.E.E</th>
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<tr>
<td>0.51</td>
<td>1.44</td>
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<tr>
<td>0.51</td>
<td>0.04</td>
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