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by

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R. W. HAFER

I. INTRODUCTION

A substantial amount of research has been devoted to examining the relationship between money supply changes and economic activity during the Great Depression. Schwartz (1981) argues that the economic decline was initiated solely by overly restrictive monetary policy. In contrast, Temin's (1976, 1981) position is that monetary policy had little impact, if any, on the economy. Although they differ in emphasis, the analyses of Friedman and Schwartz (1963) and Gordon and Wilcox (1981) present a picture of monetary and nonmonetary forces combining to explain the Depression.

Little has been done, however, to specifically investigate the issue of money demand stability during this period of time. Previous time series studies are hampered by the use of annual data which severely restrict the available degrees of freedom and, consequently, do not afford rigorous tests of stability.[1/] In fact, the only published analyses specifically aimed at examining the stability of money demand during the Great Depression are those of Gandolfi (1974) and an extension by Gandolfi and Lothian (1976). In each of these studies, annual cross-sectional data is utilized in place of time series data. Based on their innovative cross-sectional estimates, they conclude that the demand for bank deposits (i.e., demand and time deposits) was stably related to income and interest rates across the 1929-1933 period.

In this paper we propose to investigate the stability of the money demand function during the Great Depression in a more conventional manner by utilizing time series data on money, prices, real income and interest rates. This study is unique in that a quarterly money demand function is estimated for the period I/1920-IV/1939. The data set recently developed by Gordon (1982) enables us to estimate a specification commonly used in studies of post-WWII money demand. This approach permits a comparison of empirical estimates of relevant income and interest rate elasticities from the 1920-39 period to those estimated using data from the past twenty years. More importantly, because we have greatly increased the available degrees of freedom, we can confidently apply several stability tests to investigate the stability of money demand during the Great Depression.

The format of the paper is as follows. Section II presents the model, data and our estimation results. Evidence is presented for money demand functions using both the M1 and M2 definitions of money. The stability of these estimated relationships is investigated in Section III. The evidence presented there suggests that each function was subject to some sort of statistical instability during 1933. This finding is explored further in Section IV where we examine the nature of the instability. Anticipating the outcome, it appears that the underlying economic relationship embodied in each demand function is in fact stable across the 1920-39 period. Section V closes the paper with concluding remarks.

II. MODEL, DATA AND EMPIRICAL RESULTS

The conventional specification used in analyses of recent money demand behavior takes the general form of

$$(1) \quad \ln (M/P)_t = \alpha_0 + \beta_1 \ln y_t + \beta_2 \ln i_t + \beta_3 \ln(M/P)_{t-1} + \varepsilon_t$$

where M is some measure of nominal money, P is the price level, y represents current real income and i is a representative market rate of interest. The use of a lagged dependent variable reflects the hypothesis that economic agents partially adjust their actual holdings of real balances to some desired level.

We use the above specification based on evidence from previous studies. Lieberman (1980) addressed the issue of choosing between the transactions model or the utility model of money demand. Using Chow's (1966) data, Lieberman found that, properly estimated, the transactions model--i.e, the model using current real income as the scale measure--was supported by the data.[2/] Moreover, Goldfeld's (1976) analysis of various quarterly models lead to the conclusion that little was gained by incorporating a wealth variable.[3/] Finally, incorporating permanent income into the quarterly model does not significantly alter or improve the estimation results from recent periods.[4/] Thus, the use of current real income as the scale measure seems appropriate.

Choosing the best opportunity cost measure is problematic. This is because until 1933, explicit interest was paid on both demand and time deposits. From then on, only time deposits earned interest, and the rate was subject to newly enacted Regulation Q ceilings. Unfortunately for empirical analyses, reliable time series for these interest data do not exist. The data available on rates paid on demand deposits are annual primarily and only for the period 1927-32. With regard to a reliable series on savings deposit rates, the early years of the sample period only have annual figures and, as mentioned, rates were restricted by Regulation Q ceilings after 1933. It should be noted, however, that incorporating a constructed savings deposit rate into an equation similar to (1) and estimated for the period 1920-31 produced no change in the estimated interest elasticity when the commercial paper rate was used as the opportunity cost measure.[5/] Thus, although one may quibble with our choice, we shall use the prime commercial paper rate as the opportunity cost variable.[6/]

The data used to estimate equation (1) are quarterly series for the period I/1920 to IV/1939. Both M1 and M2 money demand functions are estimated.[7/] The series for real income and prices are measured by real GNP and the GNP deflator, respectively. These data are taken from Gordon (1982) wherein a generalized least squares technique based on Chow and Lin (1971) is used to interpolate existing annual data by using closely related monthly series.[8/]

Initial estimates of equation (1) using ordinary least squares (OLS) indicated the presence of significant serial correlation among the residuals. Consequently, a maximum likelihood estimation procedure is employed which yields parameter estimates that are consistent and asymptotically efficient.[9/] The maximum likelihood estimates for equation (1) using M1 and M2 are presented in Table 1. The overall explanatory power of the equations is quite high and the Durbin h-statistics indicate that the problem of first-order serial correlation in the residuals has been removed.[10/] Moreover, each variable's estimated coefficient is signed correctly and achieves statistical significance at the 95 percent level (one-tailed test).

The coefficient estimates are surprising in their overall comparability to estimates obtained from recent sample periods. For instance, the estimated speed of adjustment $(1 - \hat{\beta}_3)$ is about 18 percent per quarter for M1 and about 15 percent per quarter for M2. These estimates are well in range of more recent findings, which place the speed of adjustment for M1 to be around 20 percent per quarter and, for M2, about 12 percent per quarter.[11/] The estimated income elasticities also conform with theory and recent empirical findings.[12/] The 1920-39 sample estimates indicate that the long-run income elasticity $(\hat{\beta}_1/1-\hat{\beta}_3)$ for M1 is 0.57 and, for M2, the figure is almost 0.75.[13/] Thus, as expected from the transactions model, the long-run income elasticities fall in the range of 0.5 to 1.0.

The estimated short-run interest elasticities for M1 are noticeably larger for the 1920-39 period than those derived from recent sample period, which have tended to be around -0.02. This temporal decline is even more pronounced when considering the long-run interest elasticities: the figure for M1 is about -0.20 for the period 1920-39 and about -0.15 for M2. The M1 result contrasts sharply with Goldfeld's (1976) estimates of around -0.06 using the sample period 1952-73. Compared to Goldfeld's M2 finding, the estimate for the long-run interest elasticity (-0.18) is similar, however. This drop in the short- and long-run interest sensitivity of real M1 money balances suggests that the number of financial innovations occurring since the 1920-39 period have not reduced the responsiveness of narrowly defined real balances to changes in market rates, as posited by Gurley and Shaw (1960) and their more recent advocates.[14/]

The in-sample estimates do not appear appreciably different from those of more recent vintage. The estimated speeds of adjustment are well-within historical bounds, and the estimated income elasticities are theoretically acceptable. The major difference between the 1920-39 results and recent estimates is the higher interest response during the earlier period. Next we turn to the question of stability.

III. STABILITY TEST RESULTS

In-sample estimates of the M1 and M2 money demand equations suggest that the conventional money demand function estimated was "stable" across the 1920-39 sample period. Following Friedman (1956), we use the term stability to denote the statistical finding that estimated coefficients of the explanatory variables remain constant across differing economic environments.[15/] To address this issue, several tests are conducted.

The backward and forward recursive tests formulated by Brown, Durbin and Evans (1975, hereafter BDE) are used first.[16/] Unlike the Chow (1960) test, the BDE technique does not rely upon prior specification of the break point. Rather, it is based on comparing the ratio of the one-period-ahead prediction errors based on a moving sample period. These tests yield a statistic, called the cusum-squares statistic, that is compared to a predetermined significance level. Values of the test statistics exceeding the critical value allow one to reject the hypothesis of stability.[17/]

The outcome of the BDE forward and backward cusum-squares test are reported in Table 2. Looking first at the M1 results, the test statistics indicate that the function is not stable at the 5 percent level. At the 1 percent level, however, we cannot reject the hypothesis of stability using either test. The M2 results offer conflicting evidence: the backward cusum-squares statistic allows us to reject the stability hypothesis at the 5 percent level, but the forward cusum-squares test does not. As with M1, however, stability cannot be rejected at the 1 percent level of significance for either test.

The most common stability test used is the Chow test. One unfortunate feature of this test is the necessity to specify a priori the break in the relationship being examined. While certain historical events may lead one to prefer one period over another, Farley, Hinich and McGuire (1975) have demonstrated that the power of the test diminishes as the break point moves away from the mid-point of the sample. In order that we may apply the Chow test more objectively than usual, we employ Quandt's (1960) log-likelihood ratio statistic to determine the "most likely" break point in the estimated relationship.[18/] At the point where Quandt's likelihood ratio statistic (λ_t) reaches a global minimum, that point is selected as the most likely date of the structural change. Once the break point is determined, the Chow test then is used to assess the statistical significance of the change.

The plots of Quandt's likelihood ratio λ_t for M1 and M2 are presented in Figures 1 and 2, respectively. Looking at Figure 1, the λ_t for M1 reaches a minimum in 1933/I. Although two local minima are indicated at 1935/II and 1937/III, suggesting that some instability may be present at that time, the global minimum at 1933/I is assumed to be the most likely time of the shift in keeping with the assumptions underlying the Quandt test. The λ_t plot for M2, shown in Figure 2, indicates 1932/IV to be the likely break point. Interestingly, both M1 and M2 appear to break around early 1933, the time of the final banking crisis and the attendant financial changes.

Using the Quandt test results, a Chow test was performed using 1932/IV as the break point.[19/] The outcome is reported in Table 2. There we see that the calculated F-statistic for both M1 and M2 exceeds the 5 percent critical value allowing us to reject the hypothesis of structural stability. Note again, however, that the F-value for M1 is less than the 1 percent critical value.

The evidence presented in this section suggests that the M1 and M2 money demand functions experienced some sort of statistical instability beginning in 1933.[20/] The test statistics from the BDE and Chow test procedures reject stability at standard significance levels. We now intend to examine the nature of these instabilities.

IV. THE NATURE OF THE INSTABILITY

The above evidence suggests the existence of an unstable relationship between real balances, real income and interest rates. To further investigate why the instabilities occur, we use a test procedure recently developed by Dufour (1980, 1982b). The Dufour procedure amounts to estimating the regressions for the full sample period and entering (0, 1) dummy variables for each quarter beyond the suspected point of structural change. Since the parameter estimates for the right-hand-side variables (i.e., real income, interest rate and the lagged term) are based on pre-break data points, the estimated dummy variable coefficients are analogous to out-of-sample forecast errors. Consequently, the Dufour procedure provides valuable diagnostic information in our attempt to examine the stability of the money demand function across the 1920-39 period.

The Dufour test was implemented by re-estimating both money demand equations with individual dummy variables included for the period I/1933-IV/1939. Each equation was estimated again using the maximum-likelihood grid-search method. The results are presented in Table 3. The Dufour test results indicate that, for M1, the dummy variables for II/1933, III/1933, I/1937 and IV/1937 are significant at the 5 percent level. Thus, the M1 results indicate that the instability is localized. The results for M2 are quite different. There we find that ten out of twenty-eight of the dummy variables' t-scores exceed the 5 percent level. The M2 findings suggest a more general deterioration in the function during the post-1932 period. Because these findings are so different in nature, each is examined in more detail.

Evidence on the Shift in M1

The second and third quarters of 1933 are the most significant departures from the pre-1933 model of M1 money demand, both statistically and historically. The most likely reason for this localized deviation clearly is the advent of the national banking holiday on March 6, 1933. Prior to I/1933, the public's desire to convert bank deposits into currency is evident: from IV/1932 to I/1933, currency in circulation increased at a 50 percent annual rate (from $4,845 million to $5,359 million) and demand deposits fell at a 24 percent rate ($15, 539 million to $14, 528 million). Once the holiday was declared, late in I/1933, the currency and demand deposit figures begin to move in opposite directions: demand deposits begin to rise in III/1933 while currency holdings fall off throughout the last three quarters of 1933. The consequence of these counteracting forces is revealed in real M1 balances. After remaining relatively constant through 1932, real M1 fell sharply between II/1933 and III/1933: from $838 million to $765 million, or a 31 percent rate of decrease. During the next two quarters, however, real M1 holdings advanced at annual rates of 3 and 10 percent. Thus, the period of apparent instability in the M1 function closely is associated with the sharp decline in the money stock between the first and third quarters of 1933.[21/]

To consider the impact of these two quarters on the stability findings, the M1 demand function was re-estimated for the full period <u>excluding</u> the II/1933 and III/1933 data points. The regression results are (absolute value of t-statistics in parentheses)[22/]

$$(2) \quad \ln(M1/P)_t = \underset{(2.27)}{-0.557} + \underset{(2.34)}{0.101} \ln y_t - \underset{(3.94)}{0.038} \ln RCP_t + \underset{(13.94)}{0.842} \ln(M1/P)_{t-1}$$

$\bar{R}^2 = 0.992$ SE = 0.016 h = 0.41 $\hat{\rho} = 0.44$ (3.49)

The results obtained from omitting II/1933 and III/1933 are quite similar to those reported in Table 1. The estimated speed of adjustment is 16 percent per quarter compared to 18 percent in Table 1. The income elasticity--both short- and long-run--evidences a slight increase in the new regression, and the interest elasticities show little change.

The stability tests reported in Table 2 again were applied to equation (2) and the test results are reported in Table 4. Both the forward cusum-squares and the Chow test statistics do not permit rejection of the null hypothesis of structural stability at the 5 percent level. The backward cusum-squares statistic indicates instability at the 5 percent level, although at the 1 percent level, the stability hypothesis cannot be rejected.

The new evidence suggests a statistically stable M1 money demand function across the 1920-39 period, once the instabilities emanating from II/1933 and III/1933 are excluded. This finding that the instability found earlier for M1 may be the result of large swings in the nominal money stock and not a break in the underlying economic relationship being estimated is consistent with recent theories attempting to explain the problems of money demand estimation.[23/]

Investigating the Shift in M2

Recall from Table 3 that the Dufour results for M2 indicates a more general breakdown in the function than M1. One explanation for the shift is the change in accounting procedures after I/1933 used to calculate the money supply after I/1933. As discussed in Friedman and Schwartz (1963, pp. 428-34), prior to the banking holiday in March 1933, the money stock data includes the deposits held at licensed and unlicensed banks.[24/] After the holiday, however, the deposits of unlicensed banks are excluded from the data. The consequence of this shift in treatment is the sharp decline in M2 figures beginning in March 1933. For example, the recorded M2 figure in March 1933 is $29,970 million; the figure including the unlicensed bank deposits is $33,270 million. This discrepancy in measuring the money stock, although the magnitude of the difference declined markedly during late 1933, lasted through June 1935.

To see if this accounting artifact caused the apparent statistical shift in the M2 function, deposits at unlicensed banks were added to the recorded M2 data. The resulting "adjusted" M2 series (M2*) then was used to re-estimate equation (1). The outcome is (absolute value of t-statistics in parentheses):

$$(3) \quad \ln(M2^*/P)_t = \underset{(1.66)}{-0.377} + \underset{(1.82)}{0.077} \ln y_t - \underset{(2.67)}{0.018} \ln RCP_t + \underset{(19.43)}{0.882} \ln(M2^*/P)_{t-1}$$

$\bar{R}^2 = 0.990$ SE = 0.017 h = 0.78 $\hat{\rho} = 0.44$ (3.49)

Compared to the M2 results in Table 1, there are a few minor changes in the estimates. For example, the estimated speed of adjustment from equation (3) is 12 percent per quarter, compared to 15 percent when unlicensed bank deposits are excluded. The estimated interest elasticities also show little difference. The estimated income elasticities do, however, decline by 31 percent for the short-term elasticity and, for the long-term figure, by 12 percent.

Does this definitional change affect the earlier stability test results? The calculated stability test statistics based on the estimates of equation (3) are reported in Table 5. The cusum-squares test results offer conflicting evidence: the forward recursive test statistic exceeds the 5 percent critical value, allowing us to reject stability. The backward recursive statistic, however, is less than the 5 percent value. The Chow test result does reject stability: based on a Quandt test break of IV/1932, the calculated F-statistic easily exceeds the 1 percent critical value. Thus, the overall evidence from Table 5 suggests that the previously found instability in the M2 function cannot be explained by the change in statistical procedures used to measure bank deposits.[25]

Another hypothesis consistent with the Dufour results is that the M2 function experienced a significant, once-and-for-all downward shift in 1933. The direction of the shift is suggested empirically not only by the consistently negative signs on the Dufour dummies in Table 3, but with the movements in time deposit holdings relative to demand deposits and currency during this period. Friedman and Schwartz (1963, pp. 303-4) suggest this line of reasoning, noting that the bank failures of 1929-33 "greatly reduced the attractiveness of [bank] deposits as a form of holding wealth . . ." This "unattractiveness" of the relatively illiquid time deposit holdings, it could be argued, would be greater during the final banking crisis. This would be especially true if the larger proportion of restricted deposits (see footnote 24) was in the form of time deposits, a very reasonable assumption.

To illustrate this change in the public's preference for time versus demand deposits and currency, Figure 3 plots the ratio of the latter two items to time deposits for the period 1927-39. The sharp increase in 1933 indicates that the public's preference for relative liquidity in the form of currency and demand deposits. Although currency, demand and time deposits all declined in early 1933, the drop in time deposits was much greater. For example, from January through May of 1933, the period before which deposits began to increase, currency and demand deposit holdings declined $1,178 million. In contrast, time deposits alone fell by $2,876 million, or 144 percent of the decline in M1. As these figures suggest, although M1 declined during early 1933 (possibly explaining the II/1933, III/1933 points of instability) the decline in time deposits was much greater. Thus, a downward level shift in the M2 function caused by the unprecented flight from time deposits could explain the previous statistical results.

To test this hypothesis, the M2 equation (using <u>un</u>adjusted M2 data) was re-estimated for the 1920-39 period including an intercept shift term. This term (D1) was constructed as a (0, 1) dummy variable with a value of unity for the period I/1933-IV/1939 and zero elsewhere. Again using the maximum likelihood estimation procedure, the full-period regression results are (absolute value of t-statistics in parentheses):

$$\begin{aligned}(4)\quad \ln (M2/P)_t = \underset{(2.45)}{-0.454} - \underset{(4.02)}{0.051}\, D1 + \underset{(2.93)}{0.100} \ln Y_t \\ \underset{(5.38)}{-0.046} \ln RCP_t + \underset{(24.88)}{0.850} \ln (M2/P)_{t-1}\end{aligned}$$

$\bar{R}^2 = 0.992$ SE = 0.016 h = 0.70 $\hat{\rho} = 0.44$ (4.01)

The intercept shift term (D1) is statistically significant at the 1 percent level and correctly signed to support the hypothesis of a downward shift in the M2 function in I/1933. Indeed, the results indicate that the constant term declined about 11 percent, from -0.454 during the I/1920-IV/1932 period to -0.505 for the I/1933-IV/1939 sample. Moreover, this specification indicates that accounting for the one-time shift leads to a substantial increase in the short- and long-run interest elasticities relative to the outcome reported in Table 1.

To see if accounting for the one-time level shift restores stability to the M2 function, the Chow test and the Dufour procedure are used. (Because of the one-time shift term, the BDE tests are not applicable.) The break point assumed for the Chow test again is IV/1932. The calculated F-statistic is F(5, 70) = 2.12 which is less than the 5 percent critical value of 2.35. Thus, the Chow test result indicates a stable relationship once the downshift is captured.

The Dufour procedure is used to determine if accounting for the intercept shift in I/1933 affects the number of significant dummy variables. To do this, equation (4) was estimated with (0, 1) dummy variables entered for each quarter from I/1934 to IV/1939.[26] The results, reported in Table 6, are striking: capturing the intercept shift in M2 reduces all the post-1933 dummy terms, except II/1936, to statistical insignificance. In addition, the dummy variables in Table 6 are, with only 2 exceptions, positively signed. This is another indication that the intercept shift term has captured a significant downward level shift in M2.

To summarize, subsequent examination of the M1 function revealed that the statistical instability reported earlier probably is localized in the middle two quarters of 1933: the apparent instability in the M1 function is related to the sharp decline in the nominal stock of M1 between the first and third quarters of 1933. The basis for the previous instability results for M2 comes from the fact that the function experienced a significant, once-and-for-all down-shift in I/1933. This hypothesis is supported by the regression results given in equation (4) and the Dufour results presented in table 6. Thus, accounting for the aberration in M1 and the intercept shift in M2 results in money demand functions that are stable in their underlying economic relationships between real money balances and real income and interest rates during the Great Depression.

V. CONCLUSION

The use of a recently developed quarterly data series has permitted us to estimate a conventional, short-run money demand function for the period 1920-39. Using both the M1 and M2 definitions, our empirical results support the position that real money balances were stably related to real income and interest rates across the Great Depression. Although localized instability was discovered for the M1 function, and M2 was found to have experienced a one-time intercept shift in I/1933, these results do not detract from the argument that the functions were stable in their economic arguments. Moreover, the finding of such minor deviations in the relationships across such an economic dislocation as the Great Depression attests to the reliability of the money demand function.

Our money demand results compliment those of Gandolfi (1974) and Gandolfi and Lothian (1976). Taken together, they cast serious doubt on the hypothesis that expansionary monetary policy during the Great Depression would have been fruitless. Given the presence of a stable demand function, a sufficient condition for effective, expansionary monetary policy existed. This finding, along with the analysis of Trescott (1982), which indicates that monetary policy was highly restrictive before and during the 1929-33 decline relative to the pre-1929 policy regime, favors the hypothesis that the sharp contraction in the money supply was a primary factor in causing the Great Depression.

FOOTNOTES

[1]Examples of such time series studies are Laidler (1966) and Meltzer (1963). In these studies, stability analysis is conducted by simply comparing parameter estimates taken from different sample periods.

[2]Thornton (1982) has questioned Lieberman's empirical results due to his treatment of the autocorrelation correction procedure. Using the same data, Thornton presents conclusions that are exactly the opposite of Lieberman's.

[3]For other studies analyzing the role of wealth or income in money demand functions, see inter alia, B. Friedman (1978), Meltzer (1963), Laidler (1966), Goldfeld (1973, 1976), Hafer and Hein (1979) and Laumas and Ram (1980).

[4]Hafer and Hein (1982a) present evidence against incorporating permanent income in the quarterly model.

[5]See Cagan and Schwartz (1975). A similar specification also is used in Anderson and Butkiewicz (1980).

[6]The interest rate data are taken from Banking and Monetary Statistics.

[7]The M1 and M2 series are quarterly averages constructed from the data presented in Friedman and Schwartz (1963), Table A-1.

[8]In other words, if annual measures of both real GNP and industrial production are highly correlated, industrial production (which is available monthly) can be used to "guide" the intra-year interpolation of real GNP. For the period 1920-39, quarterly real GNP is interpolated from the monthly series on the Index of Industrial Production, along with retail sales deflated by the consumer price index. The GNP deflator is generated using the Consumer and Wholesale Price Indexes. See Gordon (1982) for a discussion of the comparability in actual and interpolated economic time series. See also, Gordon and Wilcox (1981).

[9]To obtain the coefficient estimates, the residual sum of squares is minimized conditional on the autocorrelation coefficient, which is found using a grid search across the region -1.0 to 1.0 to an accuracy of 0.01. This approach circumvents the possible iteration to a non-global minimum, a possibility using the popular Cochrane-Orcutt iterative technique. On these and related issues, see Bentacourt and Kelejian (1981) and Offenbacher (1981).

[10]The Durbin h-statistic is reported instead of the Durbin-Watson (DW) statistic because the latter is inappropriate in the presence of a lagged dependent variable. See Durbin (1970).

[11]See, for example, Offenbacher (1981) and Goldfeld (1976). Recent evidence presented in Hafer and Hein (1982b) indicates that, based on comparable 20-year sample periods across 1915-79, the estimated adjustment term for M1 ranges from 9 percent for the sample I/1950-IV/1969 to 19 percent for the I/1945-IV/1964 period. Thus, the estimates for I/1920-IV/1939 are well within historical experience.

[12]See Barro (1976) for a generalization of the theoretical constraints under the transactions model approach.

[13]The M2 income elasticity reported here is smaller than that found using recent data. For instance, Goldfeld (1976) reports the M2 long-run income elasticity to be 1.75 using the II/1952-IV/1973 sample period. Our estimate, however, is within the theoretical range specified by the transactions model. It may be argued, of course, that this model is not appropriate for the M2 definition. This discrepancy is clearly an avenue for further research.

[14]Recent arguments in this vein are reported in Enzler, Johnson and Paulus (1976) and Simpson and Porter (1980). The argument is that the growth of final intermediaries increases the interest elasticity of money demand, because a wide array of money substitutes would be available. As with any good, increasing the opportunity cost of holding it is reflected in an increased responsiveness to a relatively small change in the price of a close substitute. Thus, as money substitutes proliferate, changes in interest rates should be reflected by relatively large changes in real money balances. The data, however, do not support this view. Indeed, our results, when compared to more recent estimates suggest that real money balances have become less interest sensitive over time. This finding supports the conclusion of Cagan and Schwartz (1975) who also found that "the coefficient (in absolute value) of a short-term interest rate is not higher in this period [1954-1971] than in the 1920s and by most estimates is lower" (p. 153, emphasis added). See also, Hafer and Hein (1982b).

[15]This concept of stability is also used in Khan (1974), Heller and Khan (1979) and Boughton (1981).

[16]The BDE tests are used to analyze the stability of money demand in Khan (1974), Heller and Khan (1979), Hafer and Hein (1979) and Boughton (1981).

[17]The BDE cusum-squares statistic is written as:

$$S_r = \frac{\sum_{t=k+1}^{r} W_t^2}{\sum_{t=k+1}^{T} W_t^2} \qquad r = k+1, \ldots, T$$

where W_t^2 represents the squared one-period-ahead prediction errors generated by the estimated equation. Consequently, S_r is essentially the ratio of the squared one-period-ahead prediction errors based on the sample period k+1 (k being the number of regressors plus the constant) to r, to the errors based on a regression estimated over a sample period k+1 to T. If the estimated relationship is stable, S_r will be less than a predetermined critical value. The critical values for the backward and forward cusum-squares tests are found in Durbin (1969). For a critical appraisal of these tests, see Garbade (1977). See also, Dufour (1982a).

[18]Quandt's log-likelihood rato is premised on the hypothesis that only shift occurs during the sample period. Quandt's likelihood ratio is found by computing the time series:

$$\lambda_t = 1/2\ [r \ . \log \hat{\sigma}_1^2 + (T-r) \ . \log \hat{\sigma}_2^2 - T \ . \log \hat{\sigma}^2]$$

where $r = k+1, \ldots, T-k-1$, $\hat{\sigma}_1^2$ and $\hat{\sigma}_2^2$ are the ratios of the sum of squares residuals to the number of observations r.

[19]The point IV/1932 was used as the break, even though the Quandt test indicated I/1933 as the likely candidate for M1. This was done for convenience of comparability of results, and because historical evidence suggests that the period after IV/1932 comes from a different population. This does not seem too great an assumption given the advent of the national banking holiday in I/1933, the beginning of deposit insurance, prohibition of explicit interest payments on demand deposits along with introduction of Regulation Q ceilings on time deposits and other institutional/economic changes. Thus, we plead guilty to allowing a bit of historically founded subjectivity guide our research. For a discussion of the changes occurring during this period, see Friedman and Schwartz (1963) or Chandler (1971).

[20]Khan (1974), using annual data, found that M1 and M2 were stably related to real income and a long-term interest rate across the 1901-65 period.

[21]These figures may overstate the impact due to the exclusion of unlicensed bank deposits from the recorded money stock data. To see if including these deposits alters the textual discussion, we calculated a crude "adjusted" M1 series which includes a proportion of total deposits held at unlicensed banks. Since the data is listed only for total deposits, it was assumed that 52.5 percent of these deposits were demand deposits. This is the percent of total deposits at commercial banks held as demand deposits during 1932. Adding the demand deposits at unlicensed banks to the recorded series gives us the adjusted series. Surprisingly, the calculated growth rates for the adjusted M1 series shows a greater decrease in II/1933 than does the recorded M1 series: -20.5 percent for the adjusted series and -9.4 percent for the recorded M1 series. This suggests that the change in money stock coverage, an aspect that will be covered more in our examination of M2, is not responsible for the instability result. See Friedman and Schwartz (1963), pp. 420-34, for the relevant discussion and data.

[22]Again the coefficients are based on the maximum likelihood grid search estimation procedure. Because of the missing data points, the estimation was carried out following the procedure described in Savin and White (1978).

[23]Readers familiar with previous studies will have noticed that the specification estimated here does not incorporate a proxy for a "supply side" shock suggested by some. This measure--bank failures -- was not used for several reasons: First, the data on suspensions is

available on an annual basis. Second, the description of the data suggests that it may not be very reliable. For example, if a bank closed but later re-opened or was absorbed by another institution, it still counts as a suspended bank. Consequently, the data do not dinstiguish between temporary and permanent closings. Third, and most important, the data for 1933 are "not wholly comparable with those of other years." See Banking and Monetary Statistics, 1914-1941, p. 282 for amplification of this latter observation. Based on these conditions we have opted to exclude the variable. Moreover, the study by Anderson and Butkiewicz (1980) revealed that bank failures play a minor role in explaining the decline in M1 and M2 money demand during the Great Depression. A different view is taken in Boughton and Wicker (1979).

For studies attempting to explain recent money demand instabilities by focusing on money supply "shocks," see Carr and Darby (1981), Coats (1982) and Khan and Knight (1982).

[24]The recorded money data also does not distinguish between restricted and unrestricted bank deposits. "Restricted" deposits are those that are not available at the originally contracted date: "the payment of which has been deferred beyond the time originally contemplated." Comptroller of the Currency, Annual Report, 1933; cited in Friedman and Schwartz (1963), pp. 427, footnote 3.

[25]Equation (3) also was estimated with the Dufour dummy variables entered from I/1933-IV/1939. The result using the adjusted M2 series supports the findings in the text: a large number of dummy variables remained statistically significant.

[26]The dummy variables were entered in this fashion to give the intercept shift term an opportunity to capture the decline.

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Table 1
Money Demand Regression Results (I/1920-IV/1939)

Dependent Variable	Estimated Coefficients[1/]				Summary Statistics[2/]			
	Constant	y	RCP	$(M/P)_{t-1}$	$\bar{R}^2$	SE	h	$\hat{\rho}$
M1/P	-0.536	0.100	-0.036	0.825	0.990	0.018	0.66	0.45
	(1.91)	(1.96)	(3.29)	(12.00)				(3.40)
M2/P	-0.560	0.112	-0.022	0.848	0.990	0.017	1.15	0.56
	(2.39)	(2.56)	(3.22)	(18.06)				(5.04)

1/ All variables enter logarithmically. The equations are estimated using a maximum-likelihood grid search procedure. The numbers in parentheses are absolute values of asymptotic t-values.

2/ $\bar{R}^2$ is the coefficient of determination adjusted for degrees of freedom; SE is the standard error of the equation; Dh represents Durbin's h-statistic; ρ is the estimate of the serial correlation coefficient.

Table 2
Stability Test Results: M1 and M2

Test	Calculated Test Statistic		Critical Values	
	M1	M2	5%	1%
Cusum-Squares 1/				
Forward	0.198	0.196	0.180	0.227
Backward	0.181	0.176	0.180	0.227
Chow Test 2/	2.84	6.54	2.50	3.60

1/ Cusum-squares test based on Brown-Durbin-Evans (1975).

2/ The break point tested is IV/1932.

Table 3
Dufour Test Results[1/]

Equation Tested: $\ln(M/P_t) = \alpha_0 + \beta_1 \ln y_t + \beta_2 \ln RCP_t + \beta_3 \ln (M/P)_{t-1}$

$$+ \sum_{s = I/1933}^{IV/1939} \beta_s D_{ts} + \epsilon_t$$

Coefficient	M1	M2
α_0	-0.483 (1.79)	-0.456 (2.00)
β_1	0.089 (1.87)	0.097 (2.26)
β_2	-0.046 (3.19)	-0.035 (4.05)
β_3	0.807(10.96)	0.859(20.32)
β_s		
1933/I	0.001 (0.05)	-0.030 (1.94)
II	-0.040 (2.43)*	-0.084 (5.46)*
III	-0.091 (5.57)*	-0.083 (5.12)*
IV	-0.036 (1.96)	-0.042 (2.45)*
1934/I	-0.036 (1.96)	-0.039 (2.26)*
II	-0.036 (1.64)	-0.030 (1.56)
III	-0.023 (1.09)	-0.027 (1.42)
IV	-0.012 (0.59)	-0.020 (1.08)
1935/I	-0.022 (0.97)	-0.034 (1.75)
II	-0.023 (1.09)	-0.028 (1.45)
III	0.015 (0.66)	-0.007 (0.36)
IV	-0.016 (0.79)	-0.032 (1.66)
1936/I	-0.024 (1.20)	-0.036 (1.98)
II	0.020 (1.01)	-0.004 (0.19)
III	-0.008 (0.42)	-0.032 (1.72)
IV	-0.016 (0.81)	-0.039 (2.08)*
1937/I	-0.041 (2.06)*	-0.058 (0.00)
II	-0.035 (1.86)	-0.054 (3.03)*
III	-0.037 (1.99)	-0.048 (2.74)*
IV	-0.039 (2.09)*	-0.046 (2.73)*
1938/I	-0.002 (0.08)	-0.018 (1.06)
II	-0.026 (1.35)	-0.040 (2.32)*
III	0.001 (0.03)	-0.024 (1.34)
IV	0.013 (0.62)	-0.017 (0.93)
1939/I	-0.023 (1.06)	-0.040 (2.38)*
II	-0.003 (0.16)	-0.022 (1.57)
III	0.027 (1.29)	-0.002 (0.00)
IV	0.016 (0.76)	-0.020 (1.45)
$\bar{R}^2$	0.993	0.992
SE	0.012	0.012
h	1.013	0.99
$\hat{\rho}$	0.46 (3.40)	0.54 (4.93)

* Indicates significance at the 5 percent level.
1/ Absolute value of t-statistics appear in parentheses.

Table 4
Stability Test Results: M1 (II/1933, III/1933 omitted)

Test	Calculated Test Statistic	Critical Values 5%	1%
Cusum-Squares1/			
Forward	0.144	0.178	0.224
Backward	0.207	0.178	0.224
Chow Test2/	1.69	2.50	3.60

1/ Cusum-squares test based on Brown-Durbin-Evans (1975).

2/ The break point tested is IV/1932.

Table 5
Stability Test Results: Adjusted M2

Test	Calculated Test Statistic	Critical Values 5%	Critical Values 1%
Cusum-Squares 1/			
Forward	0.189	0.180	0.227
Backward	0.169	0.180	0.227
Chow Test 2/	3.76	2.50	3.60

1/ Cusum-squares test based on Brown-Durbin-Evans (1975).

2/ The break point tested is IV/1932.

Table 6
Dufour Test Results: M2 With I/1933 Intercept Shift[1/]

Equation Tested: $\ln (M2/P_t) = \alpha_0 + D_1 + \beta_1 \ln y_t + \beta_2 \ln RCP_t + \beta_3 \ln (M2/P)_{t-1}$

$$+ \sum_{s = I/1934}^{IV/1939} \beta_s D_{ts} + \epsilon_t$$

Coefficient	Estimate
α_0	-0.319 (1.51)
D_1	-0.056 (4.39)
β_1	0.072 (1.81)
β_2	-0.037 (3.52)
β_3	0.878(24.06)
β_s	
1934/I	0.009 (0.56)
II	0.019 (0.96)
III	0.021 (1.04)
IV	0.029 (1.36)
1935/I	0.016 (0.70)
II	0.022 (0.95)
III	0.044 (1.87)
IV	0.019 (0.81)
1936/I	0.013 (0.57)
II	0.047 (2.02)*
III	0.018 (0.80)
IV	0.012 (0.50)
1937/I	-0.010 (0.41)
II	-0.018 (0.08)
III	0.003 (0.14)
IV	0.003 (0.16)
1938/I	0.032 (1.47)
II	0.009 (0.42)
III	0.025 (1.09)
IV	0.032 (1.32)
1939/I	0.005 (0.22)
II	0.021 (0.85)
III	0.037 (1.53)
IV	0.024 (0.95)
R^2	0.991
SE	0.014
h	0.94
$\hat{\rho}$	0.46 (4.16)

* Indicates significance at the 5 percent level.

1/ Absolute value of t-statistics appear in parentheses. The shift term D1 = 1 for I/1933-IV/1939; and, 0 otherwise.

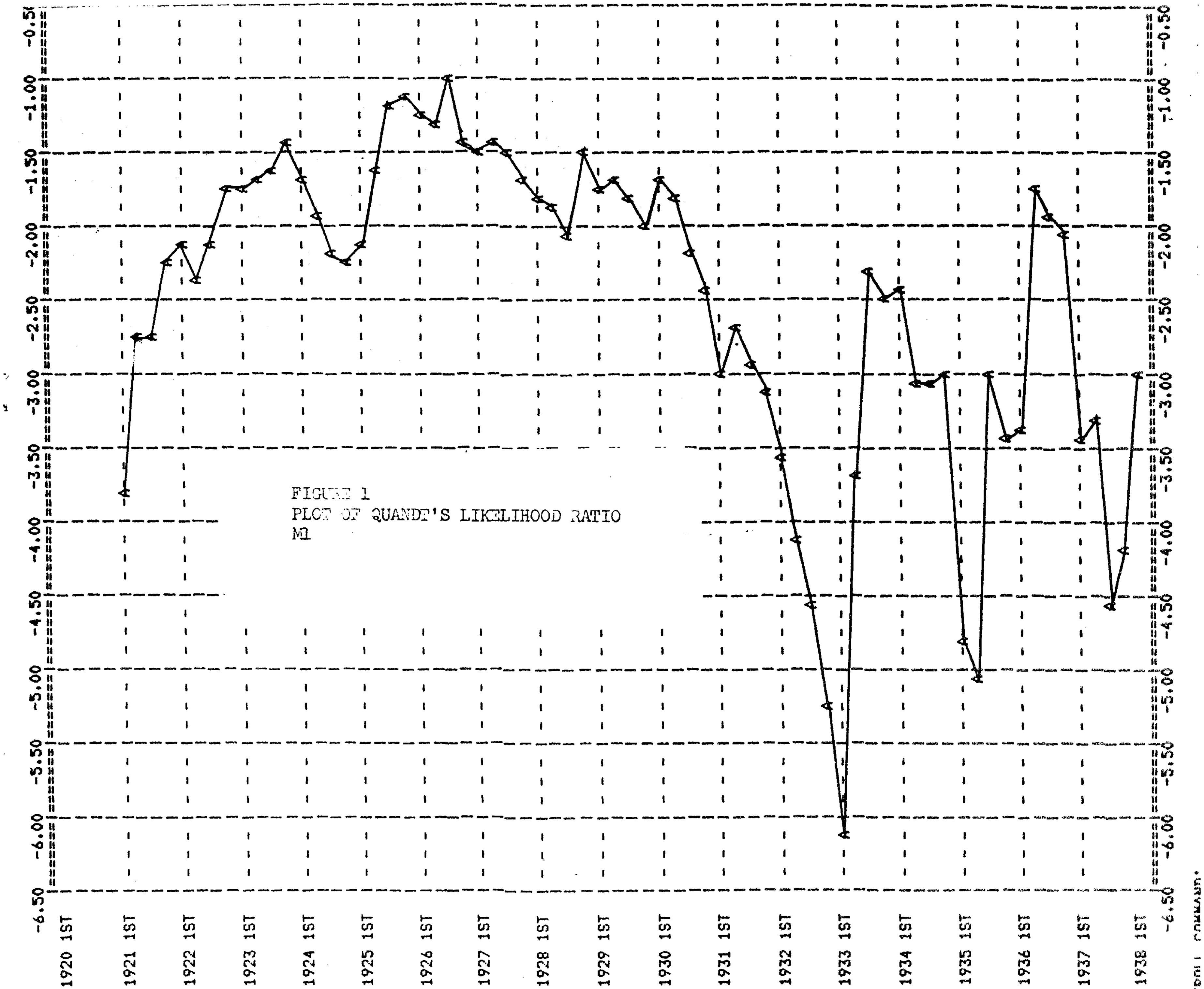

FIGURE 1
PLOT OF QUANDT'S LIKELIHOOD RATIO
M1

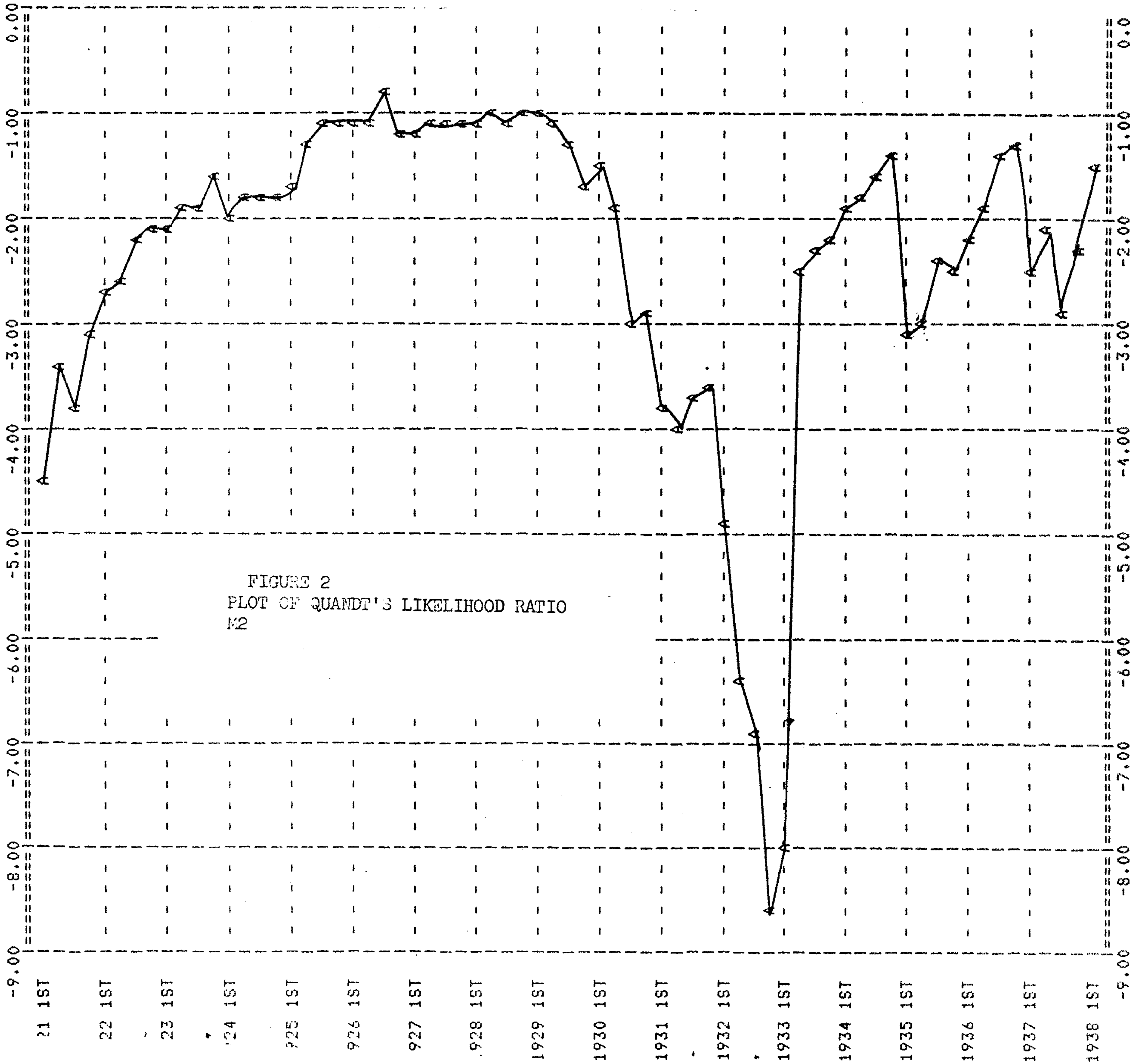

FIGURE 2
PLOT OF QUANDT'S LIKELIHOOD RATIO
M2

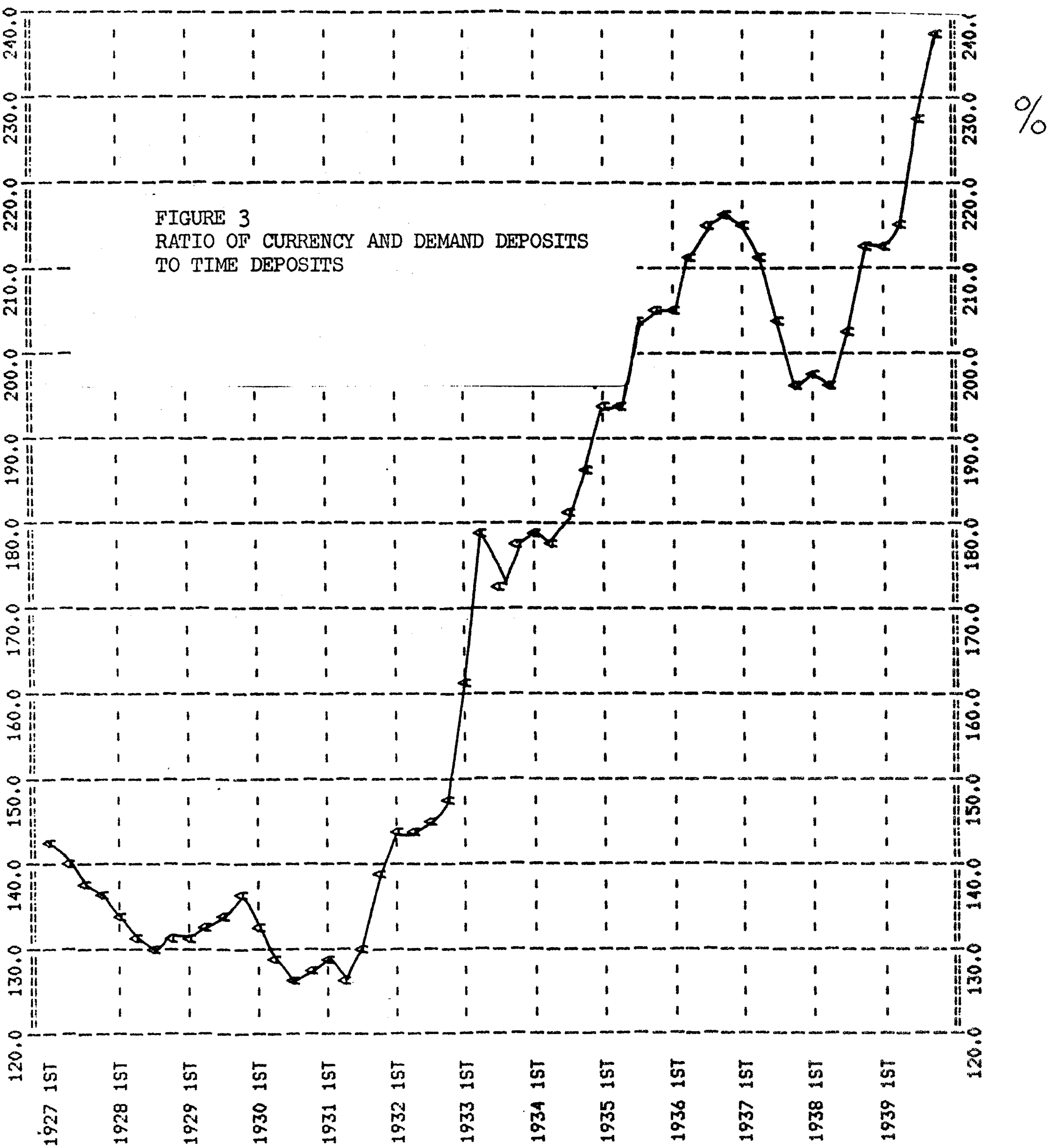

FIGURE 3
RATIO OF CURRENCY AND DEMAND DEPOSITS
TO TIME DEPOSITS