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ABSTRACT

This study investigates the equilibrium demand for narrowly defined monetary aggregate during the Great Depression. We find evidence in support of a stable demand for real balance, but no evidence in support of stable demand functions for real currency and real monetary base. This is consistent with the Friedman-Schwartz interpretation of this period.

We do not reject the hypothesis that the equilibrium demand for real M1 is stable between the pre and post WWII sample periods. We find that the "shift in the drift" of M1 velocity after 1945 and at the end of 1981 as well as the "shift in the drift" of currency and base velocities in 1981 is the image of corresponding "shift in the drift" of short-term interest rates. We interpret this as consistent with the hypothesis that the dramatic change in velocity patterns after WWII and in 1981 result from changes in inflationary expectations.

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Analyses by Lucas [1988], McCallum and Goodfriend [1987] and McCallum [1989] establish a microeconomic equilibrium relationship between real balances, short-term interest rates and a measure of the volume of transactions for a utility maximizing consumer unit. Lucas [1988] argues that it is appropriate to interpret the demand in real balances as proportional to "real permanent income" (p. 154) and to regard this as a relation that "will be stable over time provided only that preferences are, and that the trading technology...is stable" (p. 153).

In a recent paper (Hoffman and Rasche [1989]), we have investigated the nature and stability of the long-run or equilibrium aggregate demand functions for both M1 and the adjusted monetary base in the United States during the post-Accord period. We utilize actual real (personal) income as a measure of transactions volume, however, since we are concerned with equilibrium states, it is appropriate to interpret our estimate of the income elasticity as equal to that of real "permanent income." This study extends that analysis to the period of the Great Depression (January, 1929 through February, 1942).[1] In section I we review briefly the statistical methodology we apply to the estimation of income and interest elasticities in equilibrium money demand functions. In section II the sources of the various data series used in this study are discussed. In Section III we present the results of our estimations for the period of the Great Depression. In section IV we present reestimates of our previous results using data series that are comparable to those available for the 1930s, and compare the results of the reestimation to the estimates for the Great Depression. In section V we discuss the observed changes in velocity drift that occurred at the end of World War II and in 1981 in terms of our estimated equilibrium money demand functions. In section VI, some implications for monetary policy of our "shift in velocity drift" hypothesis are discussed. Finally, in section VII a summary of our major conclusions is presented.

I. Estimation and Testing Methodology

The notion of cointegration is formalized for the general case by Engle and Granger [1987]. For the purpose at hand, a set of variables which are integrated of order one are said to be cointegrated if there exists one or more linear combinations (cointegrating vectors) of these variables that are integrated of order zero.

Recent papers by Johansen [1988, 1989a], and Johansen and Juselius [1989] develop tests for both the number of cointegrating vectors and tests of hypotheses regarding elements of the

[1]January, 1929 is the earliest data for which the income data described in section II are available. February, 1942 is the last month before the Federal Reserve implemented the policy of pegging the yields on Treasury securities for the duration of World War II.

cointegrating vectors. The basic idea (using Johansen's notation) is to rewrite a
p-dimensional vector autoregression:

$$(1) \quad X_t = \sum_{i=1}^{k} \Pi_i X_{t-i} + \epsilon_i$$

as:

$$(2) \quad \Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} - \Pi X_{t-k} + \epsilon_i$$

where:

$$(3) \quad \Gamma_i = -I + \Pi_1 + \dots + \Pi_i \qquad i = 1, 2, \dots, k-1$$

and:

$$(4) \quad \Pi = I - \Pi_1 - \Pi_2 - \dots - \Pi_k.$$

so that the matrix Π conveys the long-run information in the data. When $0 < \text{rank}\ (\Pi) = r < p$ express $\Pi = \alpha\beta'$ where β may be interpreted as a $p \times r$ matrix of cointegrating vectors α a $p \times r$ matrix of vector "error correction" parameters.

Johansen [1988] shows that estimates of β can be obtained from the eigenvectors associated with the r largest eigenvalues obtained by solving the eigenvalue problem:

$$(5) \quad |\lambda S_{kk} - S_{ko} S_{oo}^{-1} S_{ok}| = 0$$

where S_{ij}; i, j = o,k represent residual moment matrices formed from least squares regressions of ΔX_t and X_{t-k} on ΔX_{t-1}, i = 1 ..., k-1. Hence the eigenvalues in this problem are the squared canonical correlations of the "levels" regression residuals with respect to those in the "differenced" regressions. The concentrated likelihood is also formed from these eigenvalues ($\hat{\lambda}_i$; i = 1,..., p) so that a test statistic for the hypothesis that there are <u>at most</u> r cointegrating vectors is:

$$(6) \quad -2 \log(Q) = - T \sum_{i=r+1}^{p} \log(1-\hat{\lambda}_i)$$

where $\hat{\lambda}_{r+1} > \ldots > \hat{\lambda}_p$ are the p-r smallest eigenvalues. This statistic has a nonstandard distribution and Johansen [1988] develops appropriate critical values.

Johansen [1989] develops tests of hypotheses regarding individual elements of α and β. The likelihood test statistic suggested for H_0: $\beta = H\Theta$ where H is an arbitrary p x s matrix, is:

$$(7) \quad -2 \log(Q) = \sum_{i=1}^{r} \log \{(1 - \lambda_i^*)/(1 - \hat{\lambda}_i)\}$$

where λ^*_i are the r largest eigenvalues obtained by solving the eigenvalue problem under the s linear restrictions conveyed by H_0. Similarly $\hat{\lambda}$ are the r largest eigenvalues obtained without the restriction. Johansen proves that this statistic is distributed as χ^2 with r(p-s) degrees of freedom. Johansen develops a comparable Wald test that is essentially based on estimates of the asymptotic covariance matrix of $\hat{\beta}$.

II. Data

The data for the pre World War II period used in this study are monthly observations on M1, the currency component of the money stock, the Adjusted Monetary Base, Personal Income, the Consumer Price Index, the commercial paper rate, and the long-term Aaa bond rate. The first six of these series are seasonally adjusted. M1 and currency are taken from Friedman and Schwartz [1970], Table I. Data on the adjusted monetary base were supplied by the Federal Reserve Bank of St. Louis. The personal income data are obtained from National Income Supplement, Survey of Current Business, 1954, pp 238-42.[2] There is no monthly deflator for personal consumption expenditures for this period, so we have used the consumer price index to deflate all of the monetary aggregates and to construct a measure of real personal income. Treasury bill rates and long-term government rates comparable to those observed in the post-war period are not available, so we have

[2]It would be extremely interesting to extend the sample back through the 1920s. Unfortunately, the available personal income data begin in January, 1929. Prior to this only fragments of the data base used to construct the monthly personal income estimates for the 1930s are available. We are investigating the possibility of constructing a monthly personal income series for the 1920s at the present time using the fragmentary data that is available. For the details on the construction of monthly personal income estimates from 1929 see Nathan and Cone [1938].

utilized the commercial paper and Aaa bond rates instead.[3] In section III we show that the interpretation of our earlier estimates for the post-war period is not sensitive to the use of the short-term private rates rather than the Treasury bill rate during that period though there are significant differences between the Aaa rate and the 10 year Treasury rate.

The sample period of this study ends in February, 1942. This date was chosen because the Federal Reserve implemented a policy of pegging the Treasury Bill rate as part of the wartime finance effort in March, 1943 (Goldfeld and Chandler [1986], p. 541).

A critical assumption that underlies our tests for cointegration and estimation of the equilibrium income and interest elasticities of the demand for money is that the time series under consideration are not stationary in levels (or log-levels), but achieve stationarity when first differenced. (The time series must be "difference stationary" (Nelson and Plosser [1982]) or integrated of order 1 (Engle and Granger [1987])). Thus it is critical to check the statistical properties of the time series used in the following analysis. In Table 1 the results of a battery of unit root tests on these data series and their log differences are presented. The tests uniformly <u>fail</u> to reject the hypothesis of a unit root in the log-levels of all the data series used in this analysis. The relevant

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[3]Recently, Cecchetti [1988] has constructed estimates of the term structure of government rates from January, 1929 through December, 1949. Through 1940, interest on U.S. government securities was exempt from Federal Income taxation, so these data do not provide a consistent series for our purposes.

subset of the unit root tests (i.e. excluding those that assume a deterministic trend in the data) applied to the log first differences of the various data series strongly reject the hypothesis of a unit root in any of these differences data. Therefore, we conclude that there is no evidence to contradict the assumption that these data series are integrated of order 1 during the period under consideration.

III. Tests for Cointegration 1929-42.

The test for a log-linear cointegrating vector among the three nonstationary variables, real M1, real personal income, and the commercial paper rate is presented in the top part of Table 2. We present results of the test using the residuals from augmented vector autoregressions of lengths 4 and 7. The computed values of the Jorgensen test statistics fail to reject the hypothesis of one or fewer cointegrating vectors against the alternatives of stationarity ($p > .5$ under the trace test) and two cointegrating vectors ($p > .5$ for the computed values of $6.48 - .83 = 5.65$ {k=4} and $8.14 - .50 = 7.64$ {k=7} under the maximum eigenvalue test). The computed values of the maximum eigenvalue test statistic for the maintained hypothesis of zero cointegrating vectors against the alternative hypothesis of one cointegrating vector are $28.83 - 6.48 = 22.35$ {k=4} and $28.54 - 8.14 = 20.46$ {k=7}. In both cases $.10 > p > .05$ which we interpret as consistent with the conclusion that there is a single cointegrating vector among the three variables. Furthermore, the estimated elements of the unique cointegrating vector are consistent with an interpretation as an equilibrium demand function for real balances.

The coefficient on real income has the opposite sign to that on real M1, and is of roughly the same absolute value. The coefficient on the log of the commercial paper rate has the same sign as the coefficient of real M1 but is smaller in magnitude. A further test that the real M1 and real income coefficients are equal in absolute value does not reject this restriction. Thus, as in the post-war data used in Hoffman and Rasche [1989], the conclusion that the equilibrium demand for real M1 is a velocity function cannot be rejected. Under the velocity constraint, the implied interest elasticity of the demand for real M1 is .45 as indicated in Table 2.

The estimate of the interest elasticity of the equilibrium velocity relationship here is remarkably close to that reported in Hoffman and Rasche [1989] for the post-war period. This is particularly remarkable in light of the different ranges of variation of short-term interest rates during the two sample periods. In the 29-42 sample period the commercial paper rate varied from .5 to 6.25 percent; in the 53-87 sample period the Treasury bill rate varied from .64 to 16.3 percent. The similarity of the estimated interest elasticities is not the result of the use of different interest rates during the two samples. The estimated interest elasticity of the equilibrium velocity vector for the 53-87 sample period using the commercial paper rate reported in Table 2 is only marginally larger than that reported in Hoffman and Rasche [1989] for the Treasury bill rate. We interpret this as strong evidence in support of the approximate log-linearity of the equilibrium of the aggregate demand for real M1 in the U.S.[4]

The results of a test for a cointegrating vector among the logs of the three variables real M1, real personal income and the Aaa corporate interest rate are given in the bottom part of Table 2, and contrast sharply with the results for the commercial paper rate. The trace and maximum eigenvalue test here fail to reject the hypothesis of one or fewer cointegrating vectors against the alternatives of stationarity ($p > .5$) and two cointegrating vectors ($p > .5$). However, the maximum eigenvalue test also fails to reject the hypothesis of zero cointegrating vectors (nonstationarity) against the alternative of one cointegrating vector. The computed values of this test statistic are $22.38 - 6.73 = 15.65$ and $21.54 - 8.37 = 13.17$ for k=4 and 7 respectively ($.5 < p < .8$). Furthermore, where the lag length is 4, the estimated coefficient on real M1 is extremely small relative to the coefficients on real income and the Aaa rate, and is

At first glance, the result of the estimation with the Aaa rate appears disturbing, particularly in light of the result reported in Hoffman and Rasche [1989] that in the post-war period there are no

[4]The post-war data on the commercial paper rate strongly support the hypothesis of a single cointegrating vector. They fail to reject the hypothesis of one of fewer cointegrating vectors against the alternatives of stationarity ($p > .5$ under the trace test for both k=4 and 7) and two cointegrating vectors ($p > .5$ under the maximum eigenvalue test for both k=4 and 7). They strongly reject the hypothesis of zero cointegrating vectors in favor of the alternative of one cointegrating vector. For k=4 the computed value of the maximum eigenvalue statistic is $33.49 - 2.19 = 31.30$ ($p < .01$) while for k=7 it is $24.89 - 3.35 = 21.54$ ($p = .05$).

differences of economic significance in the estimated cointegrating vector between real M1, real income and Treasury bill rate compared with the estimated cointegrating vector between real M1, real income and the 10-year government rate. More detailed consideration suggests that the differences between the estimates using short-term interest rates and long-term interest rates found here may be attributable to measurement errors. First, as discussed in Temin [1976], long-term corporate interest rates for a particular class of bond ratings are likely to inadequately reflect changes in risk that occurred at the depths of the Great Depression. Second, the post-war estimation reported at the end of Table 2 for the same interest rate series also fails to support the hypothesis of a cointegrating vector among real M1, real income and the Aaa rate. The only change here from the results reported in Hoffman and Rasche [1989] is the substitution of the Aaa rate for the 10 year government rate. In the latter case, there is substantial evidence in support of a cointegrating vector involving long-term interest rates. These results suggest that the specification of equilibrium money demand equations for the U.S. with short-term interest rates are more robust than with long-term interest rates. This is consistent with the conclusion of an early investigation by Laidler [1966].

A particular interesting test is to examine the data from the Great Depression period for the monetary base and the currency component of M1. In our previous study of post-war data, Hoffman and Rasche [1989], we found that a single cointegrating vector can be meaningfully interpreted in terms of a demand function for the real base. The same result is replicated with the commercial paper rate in Table 3. In addition, the evidence strongly supports the hypothesis of a unique cointegrating vector among real currency, real income, and the commercial paper rate in the post-Accord period (Table 3).

Friedman and Schwartz [1963] argue that the demand function for money (in their definition M2) was stable during the Great Depression, but that the banking panics during 1931-33 provoked massive portfolio shifts out of bank deposits into currency. Under this hypothesis, we expect that our statistical techniques should fail to find a cointegrating vector during the interwar sample period between either the real monetary base or real currency holdings and real income and interest rates, whose estimated coefficients are consistent with the parameters of demand functions for these aggregates.

The results of these tests are reported in Table 3. The results using the corporate bond rate and either the real monetary base or real currency holdings are similar to those for real M1 in Table 2: the data reject even a single cointegrating vector among the three variables. The results of the estimations using the commercial paper rate fail to reject conclusively the hypothesis of a single cointegrating vector among the three variables.[5] However, this evidence does not support the hypothesis of a stable demand function for these two aggregates during this period. In all four cases considered in Table 3, the estimated coefficient on real income in the

[5]The p values for both the trace and maximum eigenvalue tests are in the range .10 to .05 for the monetary base. The p values of the tests for currency vary considerably, but for the maximum eigenvalue test with $k=4$, $p < .05$.

cointegrating vector is very small, both in absolute value and relative to the estimated coefficient on the real monetary aggregate. In all four cases the estimated real income coefficients are not significantly different from zero. Furthermore, in three of the four cases using the commercial paper rate, the estimated coefficient on real income has the same sign as the estimated coefficient on the real monetary aggregate, contrary to the hypothesized demand function with a positive real income elasticity. Thus we conclude that the results of our statistical procedures support the Friedman-Schwartz hypothesis that the demand function for money (in this case M1) was stable during the Great Depression, but at the same time the demand functions for real currency balances and the real adjusted monetary base were not stable during the banking panics.

IV. Tests of Equality of Elements of Cointegrating Vectors Across Sample

We can test for such stability within the framework proposed by Johansen [1988], by testing whether the Π matrices which convey the long-run information are equal in the inter-war and post-war sample. To illustrate write:

$$(8)\quad \Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-k+1} - \Pi X_{t-k} + \mu + \epsilon_t$$

or in matrix terms:

$$Z_o = \Gamma Z_1 - \Pi Z_k + \epsilon$$

Where Z_o, Z_1 and Z_k are Txp, Tx(k-1)p and Txp respectively, and VEC(ϵ) $\sim N(0, \Lambda \otimes I_T)$ the likelihood function associated with this system is:

$$(9) \quad \mathcal{L} = |\Lambda|^{-T/2} \exp\{-\tfrac{1}{2}\mathrm{trace}\Lambda^{-1}(Z_o - Z_1\Gamma+Z_k\ \Pi)'(Z_o-Z_1\Gamma+Z_k\Pi)$$

The likelihood over two independent samples is:

$$(10) \quad \mathcal{L} = |\Lambda_1|^{-T1/2}\exp\{-\tfrac{1}{2}\mathrm{trace}\Lambda_1^{-1}(Z_{01}-Z_{11}\Gamma_1+Z_{k1}\Pi_1)'(Z_{01}-Z_{11}\Gamma_1+Z_{k1}\Pi_1)\}$$

$$* \ |\Lambda_2|^{-T2/2}\exp\{-\tfrac{1}{2}\mathrm{trace}\Lambda_2^{-1}(Z_{02}-Z_{12}\Gamma_2+Z_{k2}\Pi_2)'(Z_{02}-Z_{12}\Gamma_2+Z_{k2}\Pi_2)\}$$

where Z_{oi}, Z_{1i} and Z_{ki}, $i=1, 2$ have T_i rows respectively.

Now assume that the hypothesis $\Pi_1=\Pi_2$ is to be tested _without_ imposing restrictions on Γ_i or Λ_i, $i=1,2$. To accommodate the heteroskedacity, adopt a transformation Q_i for the errors of each subsample such that $Q_iQ_i'=\Lambda_i^{-1}$ $i=1,2$ using a Cholesky decomposition.

Then establish the transformation:

$$\left.\begin{aligned} W_{j1} &= Z_{j1}Q_1Q_2^{-1} \\ W_{j2} &= Z_{j2} \end{aligned}\right\} \quad j = 0,\ 1,\ 2$$

This provides homoskedacity across both regimes since:

$$\mathrm{VAR}(\mathrm{VEC}\{\epsilon_1Q_1Q_2^{-1}\}) = ([Q_2^{-1}]'Q_1^{1}\Lambda_1Q_1Q_2^{-1}) \otimes I_{T1} = \Lambda_2\otimes I_{T1}$$

Where ϵ_1 denotes the first T_1 rows of ϵ.

The likelihood based on the transformed data over the two regimes is then:

$$(11) \quad \mathcal{L} = |\Lambda_2|^{-(T_1+T_2)/2} \exp\{-\tfrac{1}{2}\mathrm{trace}\Lambda_2\ (\bar{W}^{1}_{01}-W_{11}\Gamma\ +W_{1k}\Pi_1)'(W_{01}-W_{11}\Gamma_1+W_{1k}\Pi_1)\}$$

$$* \ \exp\{-\tfrac{1}{2}\mathrm{trace}\Lambda_2^{-1}(W_{02}-W_{12}\Gamma_2+W_{k2}\Pi_2)'(W_{02}-W_{12}\Gamma_2+W_{k2}\Pi_2)\}$$

The likelihood can be concentrated with respect to Γ_i and Π_i, i=1,2, by noting that both

$$\hat{\Lambda}_{21}=(W_{01}-W_{11}\hat{\Gamma}_1+W_{1k}\hat{\Pi}_1)'(W_{01}-W_{11}\hat{\Gamma}_1+W_{1k}\hat{\Pi}_1)/T_1 \text{ and}$$

$$\hat{\Lambda}_{22}=(W_{02}-W_{12}\hat{\Gamma}_2+W_{2k}\hat{\Pi}_2)'(W_{02}-W_{12}\hat{\Gamma}_2+W_{2k}\hat{\Pi}_2)/T_2$$

where $\hat{\Gamma}_i$ and $\hat{\Pi}$ are the maximum likelihood estimates of Γ_i and Π_i provide consistent estimates of Λ_2. Hence the concentrated likelihood is:

$$(12) \quad \mathcal{L}_{w,c}= |\hat{\Lambda}_{21}|^{-(T_1)/2}|\hat{\Lambda}_{22}|^{-(T_2)/2}\exp\{\frac{-(T_1+T_2)p}{2}\}$$

A test of the hypothesis $\Pi_1 = \Pi_2$ is accomplished without imposing $\Gamma_1 = \Gamma_2$ on the transformed model by forming the restricted likelihood which concentrates to:

$$(13) \quad \mathcal{L}_{w,c} =|\hat{\Lambda}_{2c}|^{-(T_1+T_2)/2} \exp\{\frac{(T_1+T_2)p}{2}\}$$

$$\hat{\Lambda}_{2c}= \frac{1}{T_1+T_2} \left\{ \left(W_o-\begin{bmatrix}W_{11}\tilde{\Gamma}_1\\ W_{12}\tilde{\Gamma}_2\end{bmatrix}+\begin{bmatrix}W_{k1}\\ W_{k2}\end{bmatrix}\tilde{\Pi}\right)'\left(W_o-\begin{bmatrix}W_{11}\tilde{\Gamma}_1\\ W_{12}\tilde{\Gamma}_2\end{bmatrix}+\begin{bmatrix}W_{k1}\\ W_{k2}\end{bmatrix}\tilde{\Pi}\right)\right\}$$

and $\tilde{\Gamma}_i$, $\tilde{\Pi}$ are the maximum likelihood estimates for Γ_i and Π assuming $\Pi_1=\Pi_2=\Pi$. Then

$$\lambda = \mathcal{L}^*_{w,c}/\mathcal{L}_{w,c} \text{ and}$$

$$(14)\quad -2\ln\lambda = (T_1+T_2)\ell n\ |\hat{\Lambda}_{2c}|\ -T_1\ln|\hat{\Lambda}_{21}|\ -T_2\ln|\hat{\Lambda}_{22}|$$

is the likelihood ratio statistic.

The test can be simplified further by noting that

$$(15)\quad |\hat{\Lambda}_i| = |\ S^{(i)}_{oo} - S^{(i)}_{ok}\beta_i(\beta_i' S^{(i)}_{kk}\beta_i)^{-1}\ \beta_i' S^{(i)}_{ko}|$$

$$= |\ S^{(i)}_{oo}||\ \beta_i' S^{(i)}_{kk}\beta_i - \beta_i' S^{(i)}_{ko} S^{(i)-1}_{oo} S^{(i)}_{ok}\beta_i|\ |\ \beta_i' S^{(i)}_{kk}\beta_i|$$

for i=1,2,c where the $S^{(i)}_{ij}$ are the moments matrices defined in Johansen (1988) for the respective samples.[6] The β_i are estimated by solving the eigenvalue problems:

$$(16)\quad |\lambda^{(i)} S^{(i)}_{kk} - S^{(i)}_{ko} S^{(i)-1}_{oo} S^{(i)}_{ok}| = 0$$

Let $D^{(i)}$ be the diagonal matrices of eigenvalue obtained as the solutions to those problems, while V_i are the matrices of the corresponding eigenvectors such that $V_i'\ S^{(i)}_{kk} V_i = I$. Then (Johansen, [1988]):

$$(17)\quad S^{(i)}_{kk} V_i D^{(i)} = S^{(i)}_{ko} S^{(i)-1}_{oo} S^{(i)}_{ok} V_i$$

[6] Note that $T_1 S^{(1)}{}_{x,y} + T_2 S^{(2)}{}_{x,y} = (T_1+T_2) S^{(c)}{}_{x,y}$ for x,y=o,k but this does <u>not</u> establish any particular relationship among $|S^{(1)}{}_{x,y}|, |S^{(2)}{}_{x,y}|$ and $|S^{(c)}{}_{x,y}|$.

Let $\hat{\beta}_i$ be the (pxr) submatrix of V_i corresponding to the r largest eigenvalue from (17), then

$$(18) \qquad |\hat{\beta}_i S_{kk}^{(i)} \hat{\beta}_i - \hat{\beta}_i S_{ko}^{(i)} S_{oo}^{(i)-1} S_{ok}^{(i)} \hat{\beta}_i | / |\hat{\beta}_i S_{kk}^{(i)} \hat{\beta}_i| =$$

$$|\hat{\beta}_i S_{kk}^{(i)} \hat{\beta}_i - \hat{\beta}_i S_{kk}^{(i)} \hat{\beta}_i D_r^{(i)} \; | \, |\hat{\beta}_i' S_{kk}^{(i)} \hat{\beta}_i| = |I_r - D_r^{(i)}|$$

where $D_r^{(i)}$ is a diagonal matrix of the r largest eigenvalue from (17). From (16):

$$(19) \qquad |\hat{\Lambda}_{2i}| = |S_{oo}^{(i)}| \, |I_r - D_r^{(i)}| \quad \text{so:}$$

$$(20) \qquad \mathcal{L}_{w,c} = |S_{oo}^{(1)}| \, |S_{oo}^{(2)}| \, |I_r - D_r^{(1)}| \, |I_r - D_r^{(2)}|$$

$$= |S_{oo}^{(1)}| \, |S_{oo}^{(2)}| \prod_{j=1}^{r} \ln \, (1-\lambda_j^{(1)}) \prod_{j=1}^{r} \ln \, (1-\lambda_j^{(2)})$$

and

$$(21) \qquad \mathcal{L}^*_{w,c} = |S_{oo}^{(c)}| \prod_{j=1}^{r} \ln \, (1-\lambda_j^{(c)})$$

Under the maintained hypothesis that the rank $(\Pi_i)=r$, there are $r*p+(p-r)*r$ independent elements in each Π_i matrix, so the degree of freedom (the number of parameter restrictions) for the likelihood ratio test is $(2p-r)r$.

Unfortunately, we do not know the true value of the Λ_i, but have only the estimates of these matrices shown in Table 4. The approach used here is a single iteration procedure: first construct the estimated transformation matrices, Q_i, from the estimated covariance matrices, $\hat{\Lambda}_i$, and then apply these transformation matrices to the original data. Then the maximized values of the concentrated likelihood functions, $\ell^*_{W,C}$ and $\ell_{W,C}$ are used to compute the value of the likelihood ratio statistic. The results of the estimations are presented in Table 4. The computed value of the test statistic is 3.38 which is distributed as χ^2 with 5 degrees of freedom. This fails the reject equality of the Π_i matrices across the two samples, hence we conclude that we cannot reject the hypothesis of a stable equilibrium demand function for real M1 over the entire 1929-87 period.

An alternative test of the equality of cointegrating vectors in the inter and postwar periods may be obtained by forming a Wald test that relies on the distributional properties of the cointegrating vector estimates. Johansen (1989b) provides an expression for the asymptotic distribution of a normalized vector of cointegrating parameter estimates. Specifically, select a normalization; $\hat{\beta}_C=\hat{\beta}(C'\hat{\beta})^{-1}$ where $C=(I_r,0)$. Then $T(\hat{\beta}_C-\beta_C)$ is asymptotically mixed Gaussian with means zero and variance-covariance matrix that may be estimated by

$$(22) \qquad \Omega=(I-\hat{\beta}_C C')\ S_{KK}^{-1}\ (I-C\hat{\beta}_C')\otimes(C'\hat{\beta}(D^{-1}-I)^{-1}\hat{\beta}'C)^{-1}$$

where D is a diagonal matrix of the r largest eignevalues. We use these distributional properties to calculate a Wald statistic of the form

$$T(K'\mathrm{Vec}\{\hat{\beta}_C-\beta_C\})'(K'\Omega K)^{-1}(K'\mathrm{Vec}\{\hat{\beta}_C-\beta_C\}) \tag{23}$$

that is distributed as $\chi^2_{\dim K}$. The restrictions must be imposed on the normalized cointegrating vector to ensure that they are effectively constraints on π.

Our interest is to test the equality of the parameters of the cointegrating vectors that prevail in the inter and postwar periods. We form the Wald test by first augmenting the interwar series with 420 "zeros" and adding 158 "zeros" at the outset of each postwar series. The six variables formed in this manner may then be subjected to the Johansen procedure outlined above.[7] We know from Table 2 that two cointegrating vectors prevail in this system and joint estimation allows us to test hypotheses among elements of the two vectors. The estimated cointegrating vectors and associated standard errors appear in Table 5. The estimates obtained in the joint estimation using M1-money and commercial paper rates are very similar to those obtained in Table 2 with the two distinct "money demand" vectors appearing in the normalized vectors.

The Wald tests are designed to test the equality of the two free parameters across the two vectors. For the k=4 specification the $\chi^2_{(2)}$ statistic is .435 and for k=7, the value is 1.427. We conclude

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[7]Observations at beginning of the sample and at the "splice" are truncated to accommodate lags.

that we cannot reject the hypothesis of a stable equilibrium demand function for real M1 over the entire 1929-1987 period.

V. Stable Equilibrium Demand and "Shifts in Velocity Drift"

It is well documented that during the post-Accord period in the United States, the velocity of both M1 (as presently measured) and the monetary base behave like a random walks with drift. (Haraf [1986], Rasche [1987], [1988]). This statistical model is stable through late 1981. Beginning in 1982, a substantial change occurs in the drift parameter. Subsequently, both measures of velocity behave like random walks without significant drift.

This is not the first time that this phenomenon has occurred. Friedman and Schwartz [1963] refer to a change in the trend of velocity at the end of World War II. While their focus is on M2 (as they measured it; not as currently measured by the Board of Governors), they also show that the velocity of M1 displays the same change in trend. In light of the work of Gould and Nelson [1974] it is more appropriate to characterize these post World War II phenomena as changes in the drift of an I(1) process. This post-war change in the behavior of velocity has gone largely unexplained.

The 1981 "shift in the drift" is the source of the widespread conclusion that there no longer exists a stable relationship between the nominal money supply and nominal measures of economic activity. The alleged breakdown of a stable relationship between such narrowly defined monetary aggregates and measures of economic activity is the official rationale given by the FOMC for the downgrading of M1 to the status of a "monitored" aggregate and the readoption of a borrowed reserves (or free reserves) operating procedure for the conduct of monetary policy in the fall of 1982. (Volcker [1983], Wallich [1984], Heller, [1988])

Numerous hypotheses have been proposed as explanations for the abrupt change in the behavior of velocity measures in the early 1980s (Rasche [1987], Stone and Thornton [1987]). However, substantive explanations that are consistent with the observed statistical properties of the data have eluded analysts.

The problem here is to reconcile the two well documented shifts in velocity drift with the single cointegrating vector between real M1, real personal income, and the commercial paper rate that is presented in this study. Since we do not reject the hypothesis that the equilibrium real income elasticity of the demand for real M1 is unity, the existence of a single cointegrating vector implies that there is only one independent trend between the velocity of M1 and the commercial paper rate over the entire period 29-42 and 53-87. Similarly, there is only one independent trend between base and currency velocities and the Treasury bill rate (or the commercial paper rate) in the post-war period). This implies that the observed drift in any of the velocities is proportional to the drift in the corresponding nominal interest rate during the respective sample periods (Engle and Yoo [1987]. Thus any "shift in the drift" of these velocity measures is the image of a shift in the drift of nominal interest rates.

These conclusions are verified by the tests reported in Table 5. There regressions are presented for log changes of M1 velocity and the commercial paper rate against a constant and two dummy variables (D45, D82). The first of these dummies is zero through February, 1942 and 1.0 thereafter. The second is zero through December, 1981 and 1.0 thereafter. The results for velocity clearly indicate a strong upward shift in the drift after WWII, and the decline after 1981. The results indicate that the commercial paper rate also shows shifts in the drift in the same direction as the velocity shifts, but these shifts are not measured with as much precision because of the high variance in short-term interest rates.

Linear restrictions on the estimated coefficients across the two equations are tested using seemingly unrelated regression estimation (SUR). These linear restrictions are determined by the estimated equilibrium interest elasticity of M1 from Table 4.

The drift restrictions implied by the cointegrating vector for real balances are <u>not rejected</u>. Thus we conclude that the shifts in velocity drift that are observed for M1 after WWII and after 1981 are the images of corresponding shifts in the drift of nominal interest rates at those times and are consistent with a stable equilibrium demand function for real balances.

An additional hypothesis is that the drift in M1 velocity after 1981 is zero (Rasche [1988]). If this is true, then the equilibrium demand function for real balances implies that the drift in nominal interest rates after 1981 is also zero. This adds an additional (fourth) restriction across the estimated regression coefficients, and leaves only one independent parameter in the regressions. Tests of the four joint restrictions are presented in Table 6. Again the data do not reject the linear restrictions implied by the equilibrium demand for real balances, nor do they reject the hypothesis that the drifts in the two variables are zero subsequent to 1981.[8]

An even stronger test can be constructed in the same fashion for the post-war data. After 1981 there is a significant shift in the drift of M1 velocity, base velocity, and currency velocity (Rasche [1987]). The evidence in Table 3 strongly supports a single cointegrating vector for both base velocity and currency velocity during the post-war period. If these results are consistent, then the same linear restrictions among the drift of M1 velocity, base velocity, currency velocity and the commercial paper rate must hold jointly both before and after 1981. A test of the six cross equation restrictions is presented in Table 7. The data do not reject the restrictions implied by the cointegration vectors [χ^2 = 9.81; p = .13].

VI. Implications for Disinflationary Monetary Policy

These results reported in the previous section help sort out the numerous hypotheses that prevail about the change in velocity behavior in the 1980s. Nine such hypotheses are outlined in Rasche [1987]. Many of those hypotheses are inconsistent with the evidence presented in that analysis. The remaining hypotheses not conclusively ruled out

[8]The autocorrelation function of monthly log changes in the commercial paper rate suggests that this series is IMA(1). This is indicated in Table 6 by the low Durbin-Watson statistic. The dummy variables in Table 6 have been incorporated in a VAR of length 4 (corresponding to k=4) and the tests of the cross equation restrictions have been replicated in these models. The test results are the same as in the simple regressions.

by the earlier analysis are not consistent with stable equilibrium demand functions for real M1, real currency, and the real base before and after 1982.

One hypothesis which is investigated indirectly in Rasche [1987] and received little support, is that the observed change in velocity behavior is the result of a break in inflation expectations. In the absence of direct effects of measures of expected inflation rates in the demand for real balances, there is no intuitive explanation of how such a break generates a shift in the drift of the velocity measures.

The realization that the shift velocity drift is just the image of a shift in the drift of nominal interest rates provides the missing intuition for the expected inflation hypothesis. If the post-Accord period through 1980 is characterized by a steady upward drift in inflation expectations, then it is reasonable to conjecture that this drift is reflected in a positive drift in nominal interest rates. If the inflation expectations stabilized during the 1981-2 recession, and subsequently remain stable, then a reasonable conjecture is that there is no drift in nominal interest rates in the 1980s. This inflation expectations hypothesis is also a plausible rationale for the increase in M1 velocity drift after WWII compared to the earlier experience.

One source of evidence consistent with the hypothesis of a break in the drift in inflation expectations around the end of 1981 is the Livingston survey data on inflation expectations. The survey dates from the late 1940s and is plotted in Figure 1 beginning in 1954. The data are on one year ahead inflation expectations formed at the end of the previous year.[9] The series shows a general upward trend through 1980 and then breaks sharply downward. Since 1982 the series has fluctuated without trend in the 3-5 percent range.[10]

If published inflation forecasts are taken as representative of inflation expectations there is a second source of evidence in support of a break in drift of inflation expectations in the 1981-82 period. The annual CEA forecasts of the GNP deflator, as tabulated by McNees [1988], trend steadily upward from 1962 through 1981. Then the forecast rates drop precipitously in 1981-2 and stabilize in the 3-4 percent range through 1987. The forecasts for 1988 and 1989 in the respective Annual Reports of the Council of Economic Advisors are 3.9 and 3.7 percent respectively. Belongia [1988] analyzes GNP deflator forecasts for the 1976-87 period from five sources: the CEA, the CBO, the ASA/NBER survey panel, and from two major economic consulting firms. He finds that the forecasts of the latter four sources closely parallel those of the CEA. Thus the historical ex-ante inflation forecasts are consistent with the hypothesis that inflation expectations stabilized in the early 1980s and have not drifted since.

This interpretation suggests that there are "Lucas effects" associated with the implementation of a credible disinflationary monetary policy. Under this hypothesis, when agents come to believe

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[9]The data from the Livingston survey are provided by the research department of the Federal Reserve Bank of Philadelphia.

[10]It would be interesting to know if these inflation expectations series are "trend stationary" or "difference stationary". With only about 30 observations, it is unlikely that any test of the unit root hypothesis provides a reliable discrimination between the two alternatives.

that the monetary policy regime has switched from one which permits accelerating inflation, to one of stable or decelerating inflation, then the monetary authorities should expect that there will be a change in the average growth rate of velocity of narrowly defined monetary aggregates. If this average growth rate declines, then it will not be necessary to slow the growth of the monetary aggregates as much as appears from an examination of the historical data generated by the accelerating inflation policy regime in order to accomplish the objective of a constant or declining rate of inflation. For example, using the data from the 60s and 70s it was generally believed that to stabilize the inflation rate at four percent per annum, the Fed would have to achieve a long-run growth rate of the monetary base of the order of four percent per annum to allow for the historical drift of base velocity of around two percent. If agents come to believe that inflation is stabilized, and this in turn eliminates the drift of base velocity, then reducing the long-run growth rate of the monetary base to only six percent per annum will accomplish the objective of a stable inflation at a four percent rate.

VII. Conclusions

The evidence examined here supports the hypothesis that there has been a stable equilibrium demand for real M1 balance over the entire period 1929-87. The evidence also supports the Friedman-Schwartz hypothesis that the demand for real balances (here M1) remained stable throughout the Great Contraction although the banking panics during that period provoked portfolio shift out of demand deposits and into currency such that there is no meaningful equilibrium demand function for either the monetary base or currency during that period. Stable demand functions for those latter aggregates are present in the post-war period.

Our analysis also suggests an explanation for the observed changes in _reduced form_ equations relating various monetary aggregates to nominal measures of economic activity that occurred after World War II and again in 1982. The data and analysis are consistent with the hypothesis that such changes in _reduced form_ relationships occur when there are distinct changes in inflation expectations. While further research is needed to validate this conclusion, the results of this study conclusively retect that the hypothesis that financial innovation and financial deregulation are significant distabilizing factors for the demand for narrowly defined real balances in the 1980s.

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Table 1
Unit Root Test Statistics

	τ_μ		τ_τ		$T(\rho_\mu-1)$			
	AR-4	AR-12	AR-4	AR-12	AR-4	AR-4^c	AR-12	AR-12^c
α=.05,θ=-.5	-3.02	-2.82	-3.61	-3.36	-29.2	-19.9	-31.1	-36.4
α=.05,θ=0	-2.87	-2.82	-3.41	-3.36	-14.4	-16.6	-16.1	-35.8
α=.05,θ=.5	-2.93	-2.82	-3.49	-3.36	-9.8	-17.4	-10.8	-37.6
ln(Y/P)	.35	-.55	-1.50	-2.01	.64	.75	-1.12	-2.71
ln(RCP)	-.20	-2.24	-1.99	-1.87	-3.30	-3.34	-4.21	-3.30
ln(RAaa)	-.53	-.76	-2.17	-2.44	-.70	-.71	-1.05	-1.26
ln(MB/P)	-.94	-1.23	-2.41	-2.60	-.79	-1.20	-1.10	-1.63
ln(M1/P)	.75	.16	-1.81	-1.81	.75	.90	.15	.25
Δln(Y/P)	-4.82	-2.08			-117.8	-94.3	-69.5	-15.1
Δln(RCP)	-5.88	-4.08			-152.3	-271.3	-195.1	84.1
Δln(RAaa)	-5.69	-3.65			-146.7	-201.9	-155.1	208.95
Δln(MB/P)	-3.92	-3.01			-90.6	-47.6	-99.8	-105.1
Δln(M1/P)	-4.20	-2.85			-112.3	-57.1	-83.1	-53.8

	$T(\rho_\tau-1)$				t_{α^*}		$t_{\bar{\alpha}}$	
	AR-4	AR-4^c	AR-12	AR-12^c	AR-4	AR-12	AR-4	AR-12
α=.05,θ=-.5	-44.2	-33.8	-52.8	-108.3	-5.30	-6.16	-6.59	-7.36
α=.05,θ=0	-22.4	-28.3	-27.8	-114.9	-2.93	-2.95	-3.53	-3.47
α=.05,θ=.5	-15.6	-30.4	-19.0	-110.8	-2.73	-2.67	-3.15	-2.96
ln(Y/P)	-3.24	-3.24	-5.03	-7.14	.35	-.001	-1.34	-1.35
ln(RCP)	-8.70	-9.59	-9.55	-9.24	-1.72	-1.74	-1.93	-1.49
ln(RAaa)	-9.10	-10.22	-11.70	-26.04	-.55	-.52	-2.38	-2.30
ln(MB/P)	-8.27	-15.00	-10.46	-38.57	-.72	-.77	-1.87	-2.20
ln(M1/P)	-4.80	-5.81	-4.66	-9.39	1.11	.78	-1.57	-1.68
Δln(Y/P)					-8.46	-9.86		
Δln(RCP)					-8.70	-9.78		
Δln(RAaa)					-8.56	-9.57		
Δln(M1/P)					-10.32	-12.48		
Δln(MB/P)					-16.45	-21.58		

Table 1 Continued
Unit Root Test Statistics

	Z_{α^*}		$Z_{\hat{\alpha}}$	
	AR-4	AR-12	AR-4	AR-12
α=.05,θ=-.5	-44.9	-65.9	-65.8	-90.5
α=.05,θ=0	-14.3	-14.2	-21.8	-21.0
α=.05,θ=.5	-11.9	-11.1	-17.6	-14.9
ln(Y/P)	.69	-.004	-2.84	-2.87
ln(RCP)	-3.01	-2.62	-8.47	-5.41
ln(RAaa)	-.78	-.71	-10.11	-9.40
ln(MB/P)	-.71	-.88	-7.42	-10.13
ln(M1/P)	1.08	.89	-3.93	-4.69
Δln(Y/P)	-28.66	-11.17		
Δln(RCP)	-32.31	-13.59		
Δln(RAaa)	-30.61	-13.75		
Δln(MB/P)	-27.62	-12.02		
Δln(M1/P)	-28.60	-12.63		

Note:

Specifically τ_μ and τ_τ refer to tests of Ho: $\rho_\mu = 1$ or $\rho_\tau = 1$ in $y_t = \alpha + \rho_\mu y_{t-1} + \sum_{i=1}^{q}\phi_i \Delta y_{t-i} + \epsilon_t$ or $y_t = \alpha + \beta t + \rho_t y_{t-1} + \sum_{i=1}^{q}\phi_i \Delta y_{t-i} + \epsilon_t$. The normalized bias tests $T(\rho_\mu - 1)$ and $T(\rho_\tau - 1)$ use the estimated values for ρ_μ or ρ_τ in the structures defined above. Again these are tests of H_o: $\rho_m = 1$ or $\rho_\tau = 1$. The Phillips corrected normalized bias tests are formed by weighting these statistics by $c = 1/(1-\sum_{i=1}^{q}\phi_i)$. The t_{α^*} and $t_{\hat{\alpha}}$ defined as Z_{τ_μ} and Z_{τ_τ} by Schwert are adjusted Dickey-Fuller tests suggested by Phillips. The adjustments are designed to cope with potential ARMA errors in the augmented Dickey-Fuller equations. Similarly Z_{α^*} and $Z_{\hat{\alpha}}$ values are Phillips corrected normalized bias statistics. Again the corrections allow for potential ARMA errors in the Dickey-Fuller equation. We set the Phillips "lag truncation" value equal to the number of lags in the Dickey-Fuller specification (either 4 or 12). Schwert provides the details of the Phillips correction. Alternatively, a complete development of Phillips' argument can be found in Phillips and Perron [1986].

Table 2
Cointegration Tests for Real M1, Real Personal Income and Interest Rates
Log Specifications

Sample	k	Johansen Test Statistic[a]			Unconstrained Cointegrating Vector			Test for Velocity Restriction	Implied Interest Elasticity of Velocity[b]
		r=0	r<=1	r<=2	M/P	Y/P	R		
		(28.4)	(15.6)	(6.7)				$\chi^2_{(1)}$	
					Commercial Paper Rate (RCP)				
29,1-	4	28.83	6.48	.83	-1.00	.765	-.457	.53	.464
						(.12)	(.19)		(.21)
42,2	7	28.54	8.14	.50	-1.00	.940	-.448	.05	.447
						(.26)	(.17)		(.17)
53,1-	4	33.49	2.19	.29	-1.00	.928	-.596	.04	.647
						(.32)	(.23)		(.06)
87,12	7	24.89	3.35	.29	-1.00	.939	-.614	.02	.657
						(.38)	(.28)		(.07)
					Aaa Bond Rate (RAaa)				
29,1-	4	22.38	6.73	1.13	-1.00	36.91	44.36		
						(793.8)	(982.1)		
42,2	7	21.54	8.37	1.76	-1.00	18.09	23.02		
						(212.9)	(270.7)		
53,1-	4	19.21	5.64	.14	-1.00	1.32	-.936	.65	.693
						(.54)	(.49)		(.04)
87,12	7	19.90	6.84	.23	-1.00	1.30	-.977	.45	.742
						(.59)	(.49)		(.052)

[a] numbers in parentheses are the 95 percent critical values for the trace version of Johansen's test

[b] numbers in parentheses are estimated asymptotic standard errors

Table 3
Cointegration Tests for Real Monetary Base and Real Currency
Real Personal Income and Interest Rates
Log Specifications

Sample	k	Johansen Test Statistic[a]			Unconstrained Cointegrating Vector			Test for Velocity Restriction	Implied Interest Elasticity of Velocity[b]
		r=0 (28.4)	r<=1 (15.6)	r<=2 (6.7)	B/P	Y/P	R	$\chi^2_{(1)}$	
Real Monetary Base									
Commercial Paper Rate (RCP)									
29,1-	4	29.57	8.57	1.71	-1.00	-.30	-1.34		
						(1.78)	(.96)		
42,2	7	30.03	10.0	1.20	-1.00	.08	-1.13		
						(.80)	(.58)		
53,1-	4	43.84	3.22	1.37	-1.00	.809	-.337	1.23	.464
						(.12)	(.08)		(.03)
87,12	7	32.90	5.35	2.16	-1.00	.802	-.335	1.02	.467
						(.13)	(.09)		(.04)
Aaa Corporate Rate (RAaa)									
29,1-	4	24.39	7.27	1.84	-1.00	5.80	4.56		
						(9.24)	(9.17)		
42,2	7	21.72	6.73	1.67	-1.00	4.37	3.70		
						(6.13)	(6.65)		
53,1-	4	27.09	4.32	1.52	-1.00	1.93	-1.89	11.84	-.81
						(.17)	(.49)		(.47)
87,12	7	24.65	5.03	2.14	-1.00	1.99	-1.94	10.68	-1.02
						(.18)	(.51)		(.56)

Table 3 Continued
Cointegration Tests for Real Monetary Base and Real Currency
Real Personal Income and Interest Rates
Log Specifications

Sample	k	Johansen Test Statistic[a]			Unconstrained Cointegrating Vector			Test for Velocity Restriction	Implied Interest Elasticity of Velocity[b]
		r=0	r<=1	r<=2	C/P	Y/P	R		
		(28.4)	(15.6)	(6.7)				$\chi^2_{(1)}$	
					Real Currency Balances				
					Commercial Paper Rate (RCP)				
29,1-	4	27.36	4.55	.15	-1.00	-.226 (.65)	-.856 (.38)		
42,2	7	24.04	7.98	.24	-1.00	-.42 (1.77)	-1.82 (2.47)		
51,1	4	58.38	15.97	2.81	-1.00	.899 (.10)	-.282 (.06)	.67	.339 (.02)
87,12	7	49.20	19.20	3.45	-1.00	.887 (.10)	-.282 (.06)	.72	.35 (.01)
					Aaa Corporate Rate (RAaa)				
29,1-	4	22.01	5.35	.02	-1.00	-8.57 (16.67)	-11.06 (20.82)		
42,2	7	19.33	6.94	.04	-1.00	5.47 (10.68)	5.67 (11.01)		
53,1-	4	40.22	9.44	3.52	-1.00	1.09 (.18)	-.40 (.11)	.30	.35 (.03)
87,12	7	36.65	10.6	3.22	-1.00	1.02 (.16)	-.34 (.10)	.26	.33 (.04)

[a]numbers in parentheses are the 95 percent critical values for the trace version of Johansen's test

[b]numbers in parentheses are estimated asymptotic standard errors

Table 4

Estimated Residual Covariance Matricies (Λ_i) for Real M1

	29,1-42,2			53-87 omitting 80,2-80,6;81,1-81,4		
	ln RCP	ln Y/P	ln M/P	ln RCP	ln Y/P	ln M/P
ln RCP	.8989 -2			.3050 -2		
ln Y/P	-.1510 -2	.3437 -2		.3657 -5	.2132 -4	
ln M/P	-.3737 -3	.9179 -4	.2496 -3	-.2300 -5	.5371 -5	.2153 -4

Estimated Residual Correlation Matricies

	29,1-42,2			53-87 omitting 80,2-80,6;81,1-81,4		
	ln RCP	ln Y/P	ln M/P	ln RCP	ln Y/P	ln M/P
ln RCP	1.00			1.00		
ln Y/P	-.09	1.00		.01	1.00	
ln M/P	-,25	.31	1.00	-.35	.25	1.00

Estimated Maximum Eigenvalue (after Heteroskedacity Transformation)

Sample period	T	λ_i
29-42	154	.1351
53-87	404	.0784
29-42;53-87	558	.0884

Estimated Cointegrating Vector
29,1-42,2; 53-87 omitting 80,2-80,6;81,1-81.4

	ln M/P	ln Y/P	ln RCP	interest elasticity
Unconstrained	6.82	-5.72	3.41	
Constrained	5.32	-5.32	3.19	.60

Table 5

Stability Test for M1 Demand based on the Joint Estimation of Interwar and Postwar Cointegrating Vectors

	29,1-42,2			53,1-87,12			$\chi^2_{(2)}$
k	ln M/P	ln Y/P	ln RCP	ln M/P	ln Y/P	ln RCP	
	-1.00	.762	-.456	0.00	.004	-.003	
		(.16)	(.10)		(.18)	(.13)	.435
4	.00	.012	.001	-1.00	.927	-.597	
		(.23)	(.15)		(.27)	(.19)	
	-1.00	.945	-.446	0.00	-.006	.004	
		(.14)	(.09)		(.17)	(.12)	1.427
7	.00	-.100	-.168	-1.00	.965	-.630	
		(.09)	(.13)		(.35)	(.26)	

Table 6

Velocity and Commercial Paper Rate SUR Regressions
29,1-42,2; 53,1-87,12, omitting 80,2-80,6 and 81,1-81,4

Unrestricted

	Constant	D45	D82	R^2	se	d-w
$1200*\Delta \ln V_{M1}$	-3.208	6.591	-5.651	.04	15.26	2.03
	(1.22)	(1.47)	(1.98)			
$1200*\Delta \ln RCP$	-16.392	24.410	-16.066	.01	91.39	1.32
	(7.31)	(8.86)	(11.90)			

Cross Equation Restriction (.6)

	Constant	D45	D82	R^2	se	d-w
$1200*\Delta \ln V_{M1}$	-3.522	6.973	-5.840	.04	15.26	2.03
	(1.20)	(1.45)	(1.95)			
$1200*\Delta \ln RCP$	-5.869	11.622	-9.733	.01	92.09	1.31
	(1.99)	(2.42)	(3.25)			

$\chi^2_{(3)} = 2.61$ (p=.46)

Cross Equation Restriction (.6) Plus drift = 0 after 1981

	Constant	D45	D82	R^2	se	d-w
$1200*\Delta \ln V_{M1}$	-3.522	6.973	-3.451	.03	15.28	2.03
	(1.20)	(1.45)	(.82)			
$1200*\Delta \ln RCP$	-5.869	11.622	-5.752	.01	92.12	1.31
	(1.99)	(2.42)	(1.36)			

$\chi^2_{(4)} = 4.44$ (p=.35)

Table 7

Velocity and Commercial Paper Rate SUR Regressions
53,1-87,12, omitting 80,2-80,6 and 81,1-81,4

Unrestricted

	Constant	D82	R^2	se	d-w
$1200*\Delta lnV_{M1}$	3.406	-5.408	.07	7.24	1.83
	(.39)	(.94)			
$1200*\Delta lnV_B$	2.821	-3.823	.05	6.46	2.00
	(.35)	(.84)			
$1200*\Delta lnV_C$	2.436	-3.344	.04	6.08	2.17
	(.33)	(.79)			
$1200*\Delta lnRCP$	8.022	-16.071	.00	78.29	1.05
	(4.24)	(10.13)			

Cross Equation Restrictions

	Constant	D82	R^2	se	d-w
$1200*\Delta lnV_{M1}$ (.61)	3.305	-5.216	.07	7.25	1.83
	(.38)	(.92)			
$1200*\Delta lnV_B$ (.43)	2.330	-3.677	.04	6.48	1.99
	(.27)	(.65)			
$1200*\Delta lnV_C$ (.31)	1.680	-2.651	.03	6.12	2.14
	(.20)	(.47)			
$1200*\Delta lnRCP$	5.418	-8.551	.00	78.35	1.04
	(.63)	(1.50)			

$\chi^2_{(6)} = 9.81$ (p=.13)

Figure 1

Six–Month Price Expectations