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The M2 Demand in Japan: Shifted and Unstable?

TOMOO YOSHIDA and ROBERT H. RASCHE

This paper investigates the stability of M2 demand in Japan using the statistical technique recently developed by Johansen (1988) and Johansen and Juselius (1989). Long-run equilibrium money demand function is identified and estimated along with the four-variable vector error correction model (VECM). The results strongly suggests that the function shifted upwards — while leaving the income elasticity unchanged — in mid-1985, when the interest rates on large time deposits were deregulated.

I. Introduction

This study investigates the structure and stability of the long-run demand for M2+CD in Japan. The analysis is conducted within a four variable vector error correction model involving M2+CD, real GNP, the GNP deflator and the interest rate spread between the average own rate of return on M2+CD components and the call rate. The analysis covers the period from 1956 through 1989 which includes the post-war "high growth" period, the two "oil shock" experiences, and the experience under financial deregulation during the 1980s.

In Section II a brief review of existing studies on the demand for M2+CD in Japan is presented. This section also summarizes the principal unresolved issues in this literature. In Section III we outline the major events in the history of deposit interest rate regulation in Japan. In Section IV we sketch the econometric technique that we employ to construct maximum likelihood estimates of multivariate error correction models. Section V contains the empirical results of the analysis.

Our principal conclusions are that the long-run demand for M2+CD in Japan

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remained stable throughout the period 1956:1–85:2, and that the equilibrium income elasticity of the demand for real M2+CD is approximately 1.2. The equilibrium elasticity with respect to the interest rate spread is not significantly different from zero. We conclude that this equilibrium was affected by deregulation which introduced large denomination time deposits in 1985. However, this action only appears to have produced a shift in the level of equilibrium real M2+CD, and left the equilibrium elasticities and the short-run reduced form dynamics unchanged.

II. Empirical Studies on Japan's Money Demand Function

There are numerous empirical studies on the money demand function in Japan, although only a fraction of these is available in English. As in studies on many other countries, their major concern is to check the stability of the money demand function, or to find out whether and when shift(s) occurred in the course of financial innovation and deregulation.

Early studies on Japan's money demand function draw intensively on the Goldfeld-type (Goldfeld, 1973, 1976) partial adjustment model. In this context, a shift is said to be found if the estimated model fails to predict the subsequent developments in money demand. Tsutsui and Hatanaka (1982) are probably the first to point out the possibility of shifts in Japan's money demand function. In out-of-sample simulation, they found an upward shift in the M2 demand after the end of 1970s, what they call "excavated money." However, when Komura (1986) tested this upward shift with a dummy variable (set to unity from 1980 to 1982), it turned out to be insignificant.

Other research found a shift in the demand for M2 even before the deregulation process started. Hamada and Hayashi (1983) estimated the money demand function for M2 for both 1962-73 and 1974-82 and found the F-statistic for a Chow test was "highly significant." Suzuki (1985) argued that "the money-demand function shifted around 1973-74 under the influence of structural changes in the financial system and financial innovation, but there is no evidence that further shifts have occurred in the post-1974 period" (p.27).

More recently, the Bank of Japan (1988) reactivated the shift/nonshift controversy by saying that "there is a high probability of the money demand function having shifted recently" (i.e. in 1987) and that it could possibly be attributable either to "a once-and-for-all shift in the money demand function... which cannot be measured as a change in income or interest elasticity" or to a sudden increase in the interest elasticity. Ueda (1988) immediately disagreed. He dismissed the idea of recent shift in money demand, arguing that the increase in money demand reflects a wealth effect from the increase in equity and land prices in Japan.

Independently, Ito (1989) argues that the introduction of large denomination time deposits in October 1985 and subsequent reductions in the minimum denomination of such deposits must have attracted funds from bond and mutual funds markets. However, he is uncertain about the magnitude of the shift, since he obtains completely different prediction errors when the specification of his model is changed from a real to a nominal partial adjustment mechanism.

Recently several studies have been carried out outside the context of the partial adjustment model. First, Ito (1989) investigates the demand for M2 with VAR models in both log-levels and in log first differences. He finds that, while former exhibits a large shift in the demand for M2, the latter shows little sign of instability. Second, Baba (1989) and Yoshida (1990) both estimate a single equation error correction model (ECM) for the period of 1980-85 and 1968-89 respectively. They demonstrate that ECM has a far better fit than conventional partial adjustment models and, unlike partial adjustment models, no evidence of structural breaks or shifts in these models are found during the sample periods.

Methodologically, it is fair to say that the ECM is superior both to the partial adjustment and the ordinary VAR models. The excess simplicity in the lag structure of the partial adjustment model is discussed in Yoshida (1990). Engle and Granger (1987) make it clear that the (vector) ECM is a better estimation method than a VAR model in first differences when the economic variables involved are individually nonstationary but cointegrated. True, Sims, Stock and Watson (1989) show that a VAR model in levels can yield consistent estimators of cointegrated systems. In practice, though, such a VAR is useful only when the major concern is forecasting, for its estimators do not have asymptotically standard distribution, and thus any inference based on estimated standard errors is unreliable.

Although the ECM seems to be a promising way to proceed, the existing empirical studies by Baba (1989) and Yoshida (1990) are less than satisfactory at least in four respects. First, they estimate the money demand function in a single equation framework, making it likely that the results suffer from simultaneous equation bias. As Engle, Hendry and Richard (1985) correctly point out, unbiased estimation in a single equation framework is possible as long as a weak exogeneity condition holds. However, the problem is that tests for this condition are not yet fully developed. To be on the safe side, therefore, it seems better to proceed with a vector error correction model (VECM).

Second, the estimation of their cointegrating vector is problematic. Baba (1989) merely assumes that the cointegrating vector between real M2 and real GNP has unitary elasticity without estimating it. Yoshida (1990) uses the Engle and Granger's two-step estimation method and finds the cointegration vector (or scalar, to be precise) to be 1.4. Although this procedure is based on Stock's (1987) finding of superconsistency, Monte Carlo experiments by Banerjee, *et al.* (1986) show that small sample bias can be still large.

Third, the sample periods in these studies seem to be too short to make a decisive judgement on whether there is a shift(s) in the Japanese money demand function. Among the two studies, Yoshida's (1990) sample period is three times as long as Baba's, yet it spans only 20 years and has around 80 observations. Moreover, the only pre-deregulation data it includes are those from the 1970s when the Japanese economy was in turmoil after flotation of yen and the first oil crisis. Thus it may be possible that a different picture of money demand is observed once the sample period is extended to include the 1960s.

Fourth, while statistical tests by Baba (1989) and Yoshida (1990) fail to detect a shift in money demand, anecdotal evidence points in the other direction. This is:

a) Since large companies were allowed to issue commercial paper (CP) in November 1987, CP rates have stayed lower than that of large denomination time deposits offered to those companies. No wonder some of those companies, especially trading companies, are heavily engaged in "financial engineering"; earning handsome profits justs by issuing CP and redirecting the funds to banks. The outstanding value of CP is almost 17 trillion yen, which amounts to 3% of total M2, and most is said to be issued solely for such purposes.

b) Although interest rate on deposits of 10 million yen and higher are already deregulated, banks have not fully succeeded in persuading borrowers to accept their new formula for short-term prime lending rates introduced in January 1989. As a result, it is claimed that there is large-scale merry-go-rounding: companies are cashing in on the negative spread between prime lending rates and the rates paid on large denomination time deposits.

c) Influx of money into postal savings, which is not counted as M2, is reported to have dwindled, reflecting the fact that the postal system cannot offer large denomination time deposits. At the moment the maximum amount of deposit it can take from an individual is set at 7 million yen, 3 million less than the threshold of large denomination time deposits. The "*Chukoku* fund" (medium-term government bond fund), once regarded as a formidable rival for regulated-rate time deposits, is dwindling since the introduction of large denomination time deposits. The outstanding balance of this fund hit its peak at 6.8 trillion yen in August 1987 and is 4.3 trillion yen at the end of April 1990.

To summarize, there is clearly a need for further research on the demand for M2 in Japan, which a) is methodologically better than single-equation ECM, b) has a longer sample period, and c) can shed some light on the recent shift/nonshift controversy.

III. A Brief History of Interest Rate Deregulation in Japan

Before the war, interest rates on deposits were set by gentlemen's agreements among private banks. However, when the anti-trust law in 1947 made this practice illegal, the Temporary Money Rates Adjustment Law, (apparently intended to be in effect for a short period, but still in effect 35 years later) was legislated for the continued regulation of interest rates. It authorizes the Bank of Japan to impose interest rates ceilings for all types of deposits. These ceilings are known as "the BOJ guidelines."

The deregulation process began in May 1979 when banks were allowed to issue CDs up to 10% of their individual capital: the minimum amount was set to 500 million yen and maturity was restricted to between 3 and 6 months. Subsequently, the volume restriction gradually was relaxed and finally abolished in October 1987. Similarly, the minimum amount has been lowered to 50 million yen and the maturity range is now 2 weeks to 2 years. The key feature of Japan's interest rate deregulation in the first half of 1980s is that changes occurred only gradually. At the end of September 1985, the share of CDs in M2 was only 3.3%.

The process shifted into high gear in October 1985 when interest rates on large denomination time deposits, 1 billion yen or higher, were made exempt from the guideline. Since then, the threshold on large time deposits has been lowered to 10 million yen and recently the share of deregulated deposits has reached nearly 40% of M2.

The Small Saver's MMC was introduced in June 1989; its minimum amount was originally 3 million yen, subsequently lowered to 1 million yen in April 1990, and maturities are 3 months to 3 years. Interest rates are not fully deregulated, but are set at 0.7–1.75% lower than market rates, depending on maturity, and at least 0.15% higher than corresponding regulated deposits.

IV. A Description of Maximum Likelihood Procdures for Estimation of Vector ECM

Although Engle and Granger's (1987) two-step method is a pioneering work in the estimation of cointegration vectors, it has several shortcomings. The small-sample bias problem mentioned earlier is one of these. However, a major difficulty is that the method assumes there is just one cointegrating vector among the variables concerned. It is quite possible in terms of economic theory that there exist two or more cointegrating vectors (or long-run equilibrium relationships) among three or more variables. Appling the two-step method to such a system may well yield erroneous results.

An alternative estimation method for cointegrating vectors and subsequent error correction models is that of Johansen (1988, 1989a, 1989b). This procedure is useful in that a) it can estimate all the cointegrating vectors in the system in question, b) it is based on the maximum likelihood principle, c) it enables us to do tests on the number of cointegrating vectors and an estimation of related VECM in one exercise, and d) the tests have well-defined and invariant limiting distributions.

In what follows, we explain briefly the essence of Johansen's procedure. Suppose the vector p variables $X_t=(X_{1t},\dots,X_{pt})'$ is generated from the following vector autoregressive process:

$$X_t= \mu + \sum_{i=1}^{k} \Pi_i X_{t-1} + \epsilon_t \tag{1}$$

where μ is a vector of constants, Π_i are (pxp) coefficient matrices and ϵ_t is i.i.d. N(0, Λ).

Note at this point that individual components of X_t can be either I(0) or I(1).

The process (1) can be rewritten as (2) without any loss of generality:

$$\Delta X_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} - \Pi X_{t-k} + \epsilon_t \tag{2}$$

where

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

$$\Pi = I - \Pi_1 - \dots - \Pi_k.$$

Without the ΠX_{t-k} term, (2) becomes an ordinary VAR in first differences. Hence, it is the Π matrix that contains information on the long-run property of this process. Following three cases are possible in this setting:[1]

a) If rank(Π)=p, i.e. Π has a full rank, all the variables in X_t are individually I(0).
b) If rank(Π)=r and $0<r<p$, there exist $p-r$ cointegrating vectors.
c) If rank(Π)=0, i.e. Π is a null matrix, all the variables in X_t are individually I(1) and there is no cointegration.

Moreover, if there is cointegration in the system, Π can be expressed as $\Pi=\alpha\beta'$, where both α and β are $(p\times r)$ matrices, and (2) becomes

$$\Delta X_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} - \alpha\beta' X_{t-k} + \epsilon_t. \tag{3}$$

In (3) the rows of β' create linear combinations of the elements in X_{t-k}, thus can be viewed as cointegration vectors. On the other hand, rows of α are called loading vectors since they load past errors $\beta' X_{t-k}$ into the system for error correction.

Johansen (1988) and Johansen and Juselius (1989) develop a maximum likelihood estimation procedure for μ, Γ_i, α, β and Λ, which also provides tests for the number of cointegrating vectors. Note that α and β are overparameterized in (3) so that point estimation of α and β is impossible. Instead, spaces spanned by α and β are estimated and, for the ease of interpretation, columns of β are usually normalized to have one element equal to unity.

The procedure begins with two auxiliary VARs in which ΔX_t and X_{t-k} are regressed on ΔX_{t-i} (i=1,..., $k-1$) and two residual matrices R_0 and R_k are produced. Next, three moment matrices of residuals are calculated as

$$S_{ij} = T^{-1} R_i R_j \qquad i, j = 0, k \tag{4}$$

where T is the number of observations.

Then the eigenvalue problem

$$|\lambda S_{kk} - S_{k0} S_{00}^{-1} S_{0k}| = 0$$

[1]This follows from the Granger Representation Theorem (Engle and Granger, 1978).

is solved and this yields p eigenvalues $\hat{\lambda}_i$ (i=1,..., p and $\hat{\lambda}_1>\hat{\lambda}_2>\ldots>\hat{\lambda}_p$).

With these eigenvalues, two types of tests for the number of cointegrating vectors are performed. First, a test statistic for a "trace test" which tests the null hypothesis that there are at most r cointegrating vectors against the alternative of no cointegration is given as:

$$-T\sum_{i=r+1}^{p} \ln(1-\hat{\lambda}_i). \tag{5}$$

Second, a test statistic for "maximum eigenvalue test" which tests the null hypothesis of r cointegrating vectors against the alternative of $r-1$ cointegrating vectors is given as:

$$-T\ln(1-\hat{\lambda}_{r+1}).$$

The critical values for these tests can be found in Tables D.2 and D.3 in Johansen and Juselius (1989).

Suppose there are r cointegration vectors and the estimate of β is obtained as the eigenvectors associated with the r largest eigenvalues. Once we have $\hat{\beta}$, the estimation of other parameters is straightforward. $\hat{\alpha}$, $\hat{\Lambda}$ and $\hat{\Pi}$ are given as:

$$\hat{\alpha} = -S_{0k}\beta\ (\beta' S_{kk}\beta)^{-1}$$

$$\hat{\Lambda} = S_{00}-S_{0k}\beta\ (\beta' S_{kk}\beta)^{-1}\beta' S_{k0}$$

$$\hat{\Pi} = \hat{\alpha}\hat{\beta}'$$

The coefficients on constant and lagged differences $\Gamma=(\mu, \Gamma_1, \Gamma_2,\ldots, \Gamma_{k-1})$ can be calculated from the coefficient matrices of two auxiliary VAR, V_0 and V_k. Namely: [2]

$$\hat{\Gamma} = V_0 + \hat{\Pi}V_k.$$

It is also possible to place linear restrictions on both α and β in the course of estimation and test their validity. For details, see Johansen and Juselius (1989).

V. Empirical Results

A. The general model

The starting point for this analysis is a 4-dimensional reduced-form vector error correction model in the logs of nominal M2+CD, real GNP, the GNP deflator and the level of the spread between the own rate of return on M2+CD and the call rate.[3] Since our focus

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[2] V_0 is equal to $M_{0i}M_n^{-1}$ and V_k is equal to $M_{ki}M_n^{-1}$ in equation (3.5) of Johansen and Juselius (1989).

[3] Equivalently this can be viewed as a model involving nominal M2+CD, nominal GNP and the GNP deflator, or real M2+CD, real GNP and the GNP deflator. Tests for unit roots in nominal M2+CD and the GNP deflator can be found in Yoshida (1990), Table 2. Unit root tests for real GNP are constructed in Rasche (1990). In addition we have tested for, and rejected, a unit root in the inflation rate (a second unit root in the log of the GNP deflator).

is the structure of the long-run demand function for M2+CD, it is necessary to impose some identifying restrictions on the reduced-form structure in order to identify the equilibrium real income and spread elasticities of the demand for M2 +CD. In particular, we assume that in equilibrium the demand for M2+CD is the form:

$$1\text{n}(\text{M2}) - 1\text{n}(P) = \beta_0 + \beta_1 1\text{n}y + \beta_2 S \tag{6}$$

where

M2 = nominal M2+CD
P = GNP deflator
y = real GNP
S = own rate of return on M2+CD − call rate.[4]

The general reduced-form model, subject to this identifying restriction is:

$$\Delta X_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-1} + \Pi H' Z_{t-k} + \epsilon_t \tag{7}$$

where

$$\Delta X_{t-i}\ (4\times1) = \begin{bmatrix} \Delta 1\text{n}(\text{M2})_{t-i} \\ \Delta 1\text{n}y_{t-i} \\ \Delta 1\text{n}P_{t-i} \\ \Delta S_{t-i} \end{bmatrix} \quad i = 0, \ldots, k-1$$

$$Z_{t-k}\ (4\times1) = \begin{bmatrix} 1\text{n}(\text{M2})_{t-k} \\ 1\text{n}y_{t-k} \\ 1\text{n}P_{t-k} \\ S_{t-i} \end{bmatrix}$$

and

$$H\ (4\times3) = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ -1 & 0 & 0 \\ 0 & 0 & 1 \end{bmatrix}$$

$\mu\,(4\times1)$; $\Gamma_i\,(4\times4)$, $i=1,\ldots, k-1$; and $\Pi\,(4\times3)$ are the parameters of the error correction model and $\epsilon_t\,(4\times1)$ is the reduced-form error vector.

Granger (1986) and Engle and Granger (1987) show that when the elements of the X_t

[4]The own rate on M2+CD is constructed on a quarterly basis from the own rates on the individual components. The weights are the share of the component in M2+CD during the previous quarter.

vector are nonstationary, but there also exist linear combinations of those elements such as (6) that are stationary (cointegration vectors), then the rank of the Π matrix in (7) is less than three. When rank $(\Pi)=0$ then no stationary linear combinations of the elements of X_t exist, and thus the hypothesis of a long-run demand for M2+CD in (log) level terms is rejected.

The particular econometric procedures that are used in this analysis are these developed by Johansen (1988, 1989a, 1989b) that are discussed above. Equation (7) was estimated recursively for samples beginning in 1956:1 (the earliest date for which we have data on the own rate on M2+CD) and ending at various points from 1969:4 through 1989:3. In those estimations k was set at 3, which is large enough to eliminate significant autocorrelation from the estimated residuals of all four equations. Summary information from three sample periods (1956:1-73:2, 1956:1-79:4 and 1956:1-85:2) is presented in Table 1. These sample illustrate the estimates for a) the period of "high growth" prior to the 1973-74 "Oil Shock"; b) the period before the beginning of rapid changes in Japanese financial markets in 1980; and c) the period prior to the authorization of large time deposits (1985:3).

The results presented in Table 1 provide strong support for the hypothesis that cointegration exists between real M2+CD, real income and the interest rate spread (rank $\Pi > 0$). The hypothesis of stationary of all three variables is rejected, since the hypothesis that rank $\Pi \leq 2$ is *never rejected*. However, it appears that there are two cointegration vectors, since the hypothesis rank $\Pi \leq 1$ is rejected in the sample that ends in 1985:2 and is on the margin of rejection in the sample ending in 1979:4. This hypothesis is not rejected in the shortest sample period, but the sample size here is so small that the asymptotic tests may not discriminate well.

The potential existance of two cointegration vectors among the three variables creates a problem of identification of the equilibrium real income and interest rate elasticities from the reduced form parameters. It is well known that cointegration vectors are not unique: If $\beta\,(p\times r)$ is a matrix of r cointegration vectors among p variables, then any nonsingular transformation, matrix, $R(r\times r)$ applied to the matrix β produces another matrix of cointegration vectors: $\beta^*=\beta R$. Identification requires the restriction of the admissible transformations (Hoffman and Rasche, 1990).

A likely source of a second cointegration vector in this analysis is stationarity of the spread between the own rate on M2 and the call rate. Even for a sample period of 1956:1 through 1979:4, during which rates on M2 components were heavily regulated, unit root tests on the spread variable uniformily reject the unit root hypothesis. The same conclusion is supported when the sample period is extended to include the 80s.

In order to allow for the stationary behavior of the interest rate spread, the reduced form error correction model is modified by restricting Π to

$$\Pi = \left[\begin{array}{c|c} \pi_{11} & \pi_{12} \\ \hline 0 & \pi_{22} \end{array}\right] \tag{8}$$

where π_{11} is $(p-1)\times(p-2)$; π_{12} is $(p-1)\times 1$; π_{22} is (1×1) and the zero vector is dimensioned conformably. Under this restriction the interest rate spread is modeled as a k^{th} order stationary autoregressive process that can be affected (Granger caused) by lagged inflation and lagged changes in nominal M2+CD and real GNP.

Under the hypothesis of cointegration among real M2+CD, real GNP and the spread, Π can be written as

Table 1. Summary of the Estimation of Unrestricted Vector Error Correction Model

Sample Period 1956:1–73:2 (T=67)			
Estimated Eigenvalues of π:	.2722	.0909	.0460
Johansen Test Statistics H_0	$r=0$	$r\leq 1$	$r\leq 2$
Trace Test	30.83*	9.54	3.15
Maximum Eigenvalue Test	21.29*	6.39	3.15
Estimated Cointegration Vectors:	ln (M2/P)	ln$y$	Spread
	–7.77	8.05	.92
	20.92	–26.36	.07
Sample Period 1956:1–79:4 (T=93)			
Estimated Eigenvalues of π:	.2126	.1163	.0427
Johansen Test Statistics H_0	$r=0$	$r\leq 1$	$r\leq 2$
Trace Test	37.78***	15.56*	4.06
Maximum Eigenvalue Test	22.22**	11.50	4.06
Estimated Cointegration Vectors:	ln (M2/P)	ln$y$	Spread
	–3.96	3.36	.75
	19.90	–23.72	–.09
Sample Period 1956:1–85:2 (T=115)			
Estimated Eigenvalues of π:	.2158	.1211	.0455
Johansen Test Statistics H_0	$r=0$	$r\leq 1$	$r\leq 2$
Trace Test	48.16***	20.20***	5.35
Maximum Eigenvalue Test	27.96***	14.85**	5.35
Estimated Cointegration Vectors:	ln (M2/P)	ln$y$	Spread
	–.37	– .87	.75
	16.83	–19.27	–.26

* significant at .10 level
** significant at .05 level
*** significant at .01 level
Note: Critical values for the trace test and maximum eigenvalue test are found in Johansen and Juselius (1989) Table D2.

$$\Pi = \left[\begin{array}{c|c} \theta\,\beta'_1 & \theta\,\beta'_2 \\ \hline 0 & \pi_{22} \end{array}\right] \tag{9}$$

where θ is $(p-1)\times r$, β_1 is $(p-2)\times r$, and $\beta_2=1\times r$; $0<r<p-2$. When $r=1$ then there exists a unique cointegration vector among the three variables of the form:

$$\beta_{11}\ln(M2/P) + \beta_{12}\ln y + \beta_2 S. \tag{10}$$

B. The restricted model

We are unable to discover any simple linear restrictions on the reduced-form error correction model (7) that correspond to (9). Rather than proceeding with the nonlinear estimation subject to the restrictions in (9) we have used an indirect approach. First we estimate (7) subject to the following restriction on Π:

$$\Pi = \left[\begin{array}{c|c} \theta\,\beta'_1 & \pi_{12} \\ \hline 0 & \pi_{22} \end{array}\right] \tag{11}$$

where π_{12} is $(p-1)\times 1$. This is less restrictive than (9), and is straightforward to estimate.[5] The full error correction model estimated under these restrictions is:

$$\Delta X_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-1} + \left[\begin{array}{c} \pi_{12} \\ \hline \pi_{22} \end{array}\right] S_{t-k} + \left[\begin{array}{c|c} \theta\,\beta'_1 & 0 \\ \hline 0 & 0 \end{array}\right] X_{t-k} + \epsilon_t$$

which is equivalent to (7) subect to the restrictions in (11). After estimating this model we tested the linear restrictions $\pi_{12}=0$, using likelihood ratio test. These latter restrictions imply a block diagonal π matrix:

[5]Estimation of (7) subject to the restrictions in (11) is done by using the restriction matrix H for β; imposing linear restrictions on α (Johansen and Juselius, 1989) of the form $\alpha=A\theta$ where

$$A = \begin{bmatrix} I_{p-1} \\ 0 \end{bmatrix}$$

and then adding the variable S_{t-k} as an additional regressor, just as Johansen and Juselius add seasonal dummy variables in their analysis. The addition of S_{t-k} as a separate regressor causes no statistical problems since it is stationary. The restriction matricies A and H result in a Π matrix of the form

$$\Pi = \left[\begin{array}{c|c} \theta\,\beta'_1 & 0 \\ \hline 0 & 0 \end{array}\right]$$

$$\Pi = \left[\begin{array}{c|c} \theta\,\beta'_1 & 0 \\ \hline 0 & \pi_{22} \end{array}\right] \tag{12}$$

The economic implication of the full set of restrictions is that the *equilibrium* elasticity of the demand for real M2+CD with respect to the interest rate spread is zero, but that changes in the rate spread can affect the short-run movements in real and nominal M2+CD. A failure to reject the restrictions in (12) makes the procedurally more tedious tests of the restrictions in (9) redundant. A summary of the estimation of (7) subject to restrictions (11) are shown in Table 2 for two sample periods: 1956:1-79:4 and 1956:1-85:4 and the corresponding estimations subject to restrictions (12) are shown in Table 3.[6]

The results in Table 2 are consistent with the conclusion that there is no more than one stationary linear combination of real M2+CD and real GNP, once the model is modified to allow for the stationary behavior of the interest rate spread. A common feature of the results from both sample periods shown is the small point estimate of the elements of the vector π_{12}. This suggests that the restrictions of (12) may not be rejected. The estimations reported in Table 3 impose restrictions that are sufficient to restrict Π to the block diagonal

Table 2. Summary of the Estimation of Restricted Vector Error Correction Model [Restrictions from Equation (6)]

Sample Period 1956:1–79:4		
Estimated Eigenvalues of π:	.1087	.0497
Johansen Test Statistics H_0	$r = 0$	$r \leq 1$
Trace Test	15.45*	4.74
Maximum Eigenvalue Test	10.71	4.74
Estimated Cointegration Vector	−21.06	24.75
$\pi_{12}' = [-.00014 \quad .00035 \quad -.00039]$	$\pi_{22} = -.1716$	
Sample Period 1956:1–85:4		
Estimated Eigenvalues of π_{11}:	.1130	.0525
Johansen Test Statistics H_0	$r = 0$	$r \leq 1$
Trace Test	19.99**	6.20
Maximum Eigenvalue Test	13.79*	6.20
Estimated Cointegration Vector	16.98	-19.19
$\pi_{12}' = [-.00008 \quad .00078 \quad -.00051]$	$\pi_{22} = -.1945$	

* significant at .10 level
** significant at .05 level

[6]We discuss the stability of our estimates over different sample periods in considerable detail in the following section.

Table 3. Summary of the Estimation of Restricted Vector Error Correction Model [Restrictions from Equation (7)]

Sample Period 1956:1–79:4		
Estimated Eigenvalues of π_{11}:	.1452	.0355
Johansen Test Statistics H_0	$r = 0$	$r \leq 1$
Trace Test	17.96**	3.46
Maximum Eigenvalue Test	14.60**	
	ln (M2/P)	ln$y$
Estimated Cointegration Vector	−18.179	21.588
Test Statistic for $\pi_{12} = 0 : \chi^2_{(3)} =$ .39		

Sample Period 1956:1–85:2		
Estimated Eigenvalues of π_{11}:	.1523	.0403
Johansen Test Statistics H_0	$r = 0$	$r \leq 1$
Trace Test	23.74***	4.73
Maximum Eigenvalue Test	19.01***	
	ln (M2/P)	ln$y$
Estimated Cointegration Vector	−15.39	17.97
Test Statistic for $\pi_{12} = 0 : \chi^2_{(3)} =$ .95		

* significant at .10 level
** significant at .05 level
*** significant at .01 level

Parameter Estimates of Error Correction Model 1956:1–85:2

$$\hat{\alpha} = \begin{bmatrix} -.0224 \\ -.0089 \\ .0211 \\ .0000 \end{bmatrix} \qquad 10^3 * \hat{\Lambda} = \begin{bmatrix} .0271 & -.0035 & .0102 & .8323 \\ -.0035 & .1486 & -.0108 & -.3334 \\ .0102 & -.0108 & .0852 & -.5712 \\ .8323 & -.3334 & -.5712 & 498.15 \end{bmatrix}$$

$$\begin{array}{cc} & \begin{array}{ccc} \ln M2_{t-3} & \ln y_{t-3} & \ln P_{t-3} \end{array} \\ \hat{\pi}_{11} = & \begin{bmatrix} .0224 & -.0261 & -.0224 \\ .0137 & -.0160 & -.0137 \\ -.0324 & .0378 & .0324 \end{bmatrix} \end{array}$$

$$\hat{\pi}_{22} = -.1945$$

$$\begin{array}{cc} & \begin{array}{ccccccccc} \text{const} & \Delta S_t & \Delta S_t & \Delta y_t & \Delta y_t & \Delta \ln M2_{t-1} & \Delta \ln M_{t-2} & \Delta \ln P_{t-1} & \Delta \ln P_{t-2} \end{array} \\ \hat{\Gamma} = & \begin{bmatrix} .0061 & .0018 & .0009 & .0172 & .0207 & .5865 & .2132 & -.1247 & .0790 \\ .0037 & -.0008 & .0014 & -.0546 & .2071 & .1471 & .2162 & -.0253 & -.1815 \\ -.0046 & -.0009 & -.0001 & -.1978 & -.0167 & .0960 & .3602 & .3527 & .0802 \\ -.1453 & .0453 & .1311 & -.3418 & .1039 & 7.9639 & -19.4363 & -17.0682 & 10.0323 \end{bmatrix} \end{array}$$

form of (12).[7] The restrictions of (12) relative to (11) can be tested by a likelihood ratio test which is distributed as χ^2 with 3 degress of freedom. In neither case shown in the table are the restrictions sufficient to produce (12) rejected relative to (11). Thus we conclude that the data are consistent with the hypothesis that real M2+CD and real GNP are cointegrated, and that the euilibrium elasticity of the demand for real M2+CD with respect to the interest rate spread is not significantly different from zero.

From the information reported in Table 3 it is also possible to construct point estimates of the equilibrium real income elasticity of the demand for M2+CD from the estimates of the elements of the cointegration vector. The estimated income elasticities are (1.19 (21.588/18.179) and 1.17 (17.971/15.393)) for 1956:1-79:4 and 1956:1-85:2, respectively. These estimates are broadly consistent with that of Yoshida (1990) though they are derived using a completely different estimation techniques and using one sample (1956: 1-79:4) that is not affected by the rapid changes in Japanese financial markets during the 80s.

C. Stability of the restricted model

In this section we examine the stability of the error correction model (7) subject to the restrictions (12) to the choice of sample period. The model is estimated recursively for sample periods beginning with 1956:1 and ending with each quarter from 1974:1 through 1989:3. The minimum effective sample size is 70 observations; the maximum 132 observations.

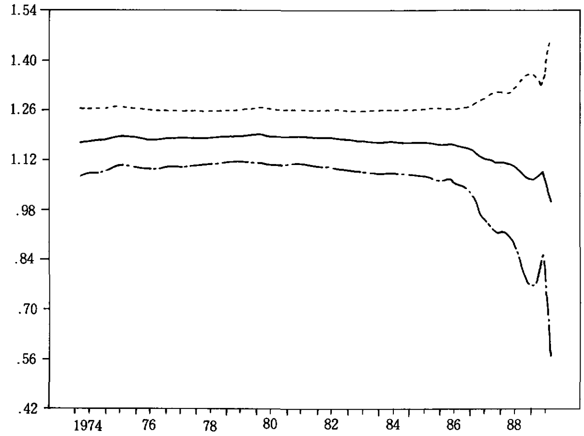
It is necessary to summarize all the available information, since the full error correction model contains a large number of estimated parameters. Fortunately a clear and accurate picture of the effects of increasing the sample size is obtained by concentrating on the estimated equilibrium real income elasticity and the precision with which that parameter is estimated. In Figure 1 the estimate of the equilibrium real interest elasticity is plotted against the data of the final observation of the estimation period. An estimate of the confidence interval for this coefficient is also plotted as ± 1.96 standard errors.

The picture presents a dramatic contrast. From the beginning of 1974 through the middle of 1985, the recursive regressions generate almost constant estimates of both the equilibrium real income elasticity and its standard error. The average point estimate of the elasticity is approximately 1.17, the average estimate of its standard error is approximately 0.045. In every case it is significantly greater than 1.0. Furthermore there is no evidence of any instability of the estimate when the sample periods extend beyond the end of 1981. Thus the error correction model for M2+CD is strikingly different from that for Japanese M1 investigated by Rasche (1990). This difference and the economic rationale for such a difference is discussed in the following section.

The contrast with this stability occurs when the terminal date of the sample period is

[7]The restrictions imposed here are stronger than the necessary conditions to assure $\pi_{12}=0$.

Figure 1. Real Income Elasticity of M2+CD



after the middle of 1985. As the sample period is lengthened, and point estimate of the equilibrium real income elasticity decreases rapidly, and the precision with which it is estimated deteriorates even more rapidly. This occurs even though the additional observations represent only a small portion of the total observations in the sample. Clearly the data generation process is different in the late 1980s that it was previously.

A potential source for this change is the major financial deregulation that occurred at that time. There is considerable disagreement over the impact of this deregulation on the demand for M2+CD. One view is that the availability of market determined rates on large time deposits caused a shift in asset portfolios out of assets not included in M2+CD into large time deposits. Alternatively it is proposed that a portion of the holdings of large time deposits represents a shift in portfolios from other components of M2+CD. A third view is the portfolio shifts into M2+CD have occurred as a result of large capital gains on real estate and corporate equities. The first and third hypotheses are consistent with an increase in the equilibrium level of real M2+CD. The third hypothesis also suggests a change in the deterministic rate of growth or "drift" in M2+CD in the face of continuing appreciation of equities and land during the latter part of the 80s.

In this analysis we add to the error correction model a simple dummy variable, D85, that takes a value of zero prior to 1985:3, and a value of 1.0 subsequent to 1985:2. The expanded model is estimated recursively for sample periods with terminal dates from 1985:3 through 1989:3. The results of these estimations are shown in Figures 2 and 3 in terms of the point estimate of the equilibrium real income elasticity of M2+CD and its standard error. The addition of the dummy variable eliminates the sharp downward trend in the recursive estimates of the equilibrium real income elasticity, and stabilizes the recursive estimates of the standard error of the parameter. All other parameters of the error correction model also remain stable after the addition of the dummy variable to the specification. This can be seen from a comparison of the individual parameter estimates in

Figure 2. Real Income Elasticity of M2+CD

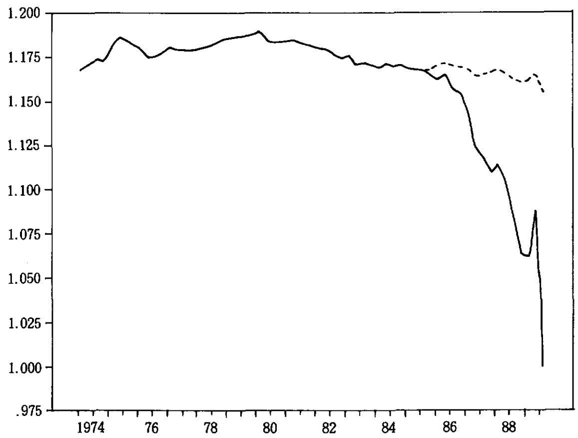
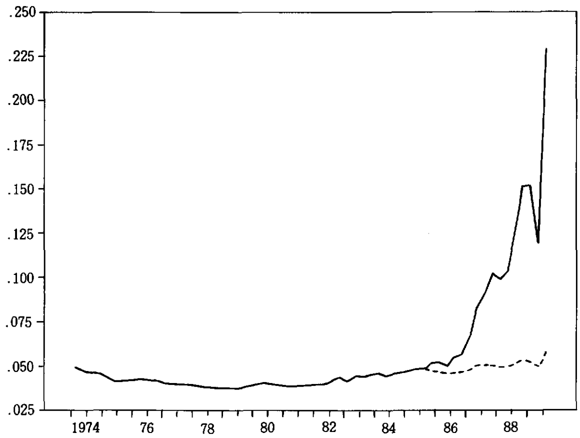


Figure 3. S.E. of Income Elasticity



part C of Table 3 for the sample period ending in 1985:2 with the corresponding parameter estimates in Table 4 for the sample period ending in 1989:2.[8]

D. An interpretation of the economic significance of the augmented error correction model

There are two significant features of the stability of the estimated error correction model during the 1980s. The first is the stability of the parameter estimates after including the D85 dummy variable. The second is a change in the structure that does *not* occur.

The reduced-form error correction model, expressed in deviations from means is:

$$(\Delta X_t - \mu_0) = \sum_{i=1}^{k-1} \Gamma_i(\Delta X_{t-i} - \mu_0) - \alpha[\beta' X_{t-k} - \beta'_0] + \epsilon_t. \tag{13}$$

[8]We have also considered an additional dummy variable that is 0.0 through 74.4 and 1.0 thereafter to represent any differences between the "high growth" period and the period since the oil shock. The estimated parameters of the vector error correction model reported in Tables 3 and 4 are robust to the addition of this dummy variable.

Table 4. Estimation of Restricted Vector Error Correction Model
[Restrictions from Equation (7)]
with D85 Dummy Variable

Sample Period 1956:1–89:2		
Estimated Eigenvalues of π_{11}:	.1448	.0384
Johansen Test Statistics H_0	$r=0$	$r \leq 1$
Trace Test	25.625	5.126
Maximum Eigenvalue Test	20.499	
	ln (M2/P)	ln$y$
Estimated Cointegration Vector	15.453	−18.013
Estimated Equilibrium Real Income Elasticity		1.166
Estimated Standard Error (Wald)		.049

$$\hat{\alpha} = \begin{bmatrix} -.0217 \\ -.0121 \\ .0302 \\ .0000 \end{bmatrix} \qquad 10^3 * \hat{\Lambda} = \begin{bmatrix} .0249 & -.0018 & .0085 & .6763 \\ -.0018 & .1394 & -.0113 & -.4835 \\ .0085 & -.0113 & .0765 & -.4827 \\ .6763 & -.4835 & -.4827 & 467.71 \end{bmatrix}$$

$$\begin{array}{cc} & \begin{array}{ccc} \ln M2_{t-3} & \ln y_{t-3} & \ln P_{t-3} \end{array} \\ \hat{\pi}_{11} = & \begin{bmatrix} .0217 & -.0252 & -.0217 \\ .0121 & -.0141 & -.0121 \\ -.0302 & .0352 & .0302 \end{bmatrix} \end{array}$$

$$\hat{\pi}_{22} = -.1935$$

$$\begin{array}{cc} & \begin{array}{cccccccccc} \text{const} & D85 & \Delta S_{t-1} & \Delta S_{t-2} & \Delta y_{t-1} & \Delta y_{t-2} & \Delta \ln M2_{t-1} & \Delta \ln M2_{t-2} & \Delta \ln P_{t-1} & \Delta \ln P_{t-2} \end{array} \\ \hat{\Gamma} = & \begin{bmatrix} .0067 & .0016 & .0016 & .0010 & .0179 & .0158 & .5954 & .2083 & -.1305 & .0772 \\ .0037 & -.0016 & -.0005 & .0014 & -.0733 & .2021 & .1386 & .2494 & -.0460 & -.1713 \\ -.0050 & -.0074 & -.0010 & -.0001 & -.1964 & -.0118 & .0947 & .3472 & .3582 & .0802 \\ -.1564 & .1749 & .0369 & .1080 & -.5182 & .1499 & 11.3240 & -22.1554 & -17.1574 & 9.6823 \end{bmatrix} \end{array}$$

This can be rewritten with a constant term as:

$$\Delta X_t = \{(I - \sum_{i=1}^{k-1} \Gamma_i)\mu_0 + \alpha\beta'_0\} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} - \alpha\beta' X_{t-k} + \epsilon_t. \tag{14}$$

From (14) it is clear that the vector of constant terms in the reduced-form error correction model, $\mu = \{(I - \sum_{i=1}^{k-1} \Gamma_i)\mu_0 + \alpha\beta'_0\}$ has two separate components one of which depends on the deterministic trends in the variables and a second that depends on the constant terms of the cointegration vectors. Johansen (1989b) shows that

$$\mu_0 = \beta_\perp [\alpha'_\perp \Pi_1(1) \beta_\perp]^{-1} \alpha'_\perp \mu \tag{15}$$

where $\beta'_\perp$ and $\alpha_\perp$ are $p \times (p-r)$ matrices such that:

$$\beta'_{\perp}\beta = 0$$

$$\alpha'_{\perp}\alpha = 0$$

$$\alpha'_{\perp}\Lambda\alpha_{\perp} = I$$

$$\alpha_{\perp}\alpha_{\perp} = \Lambda^{-1}(I - \alpha(\alpha'\Lambda^{-1}\alpha)^{-1}\alpha'\Lambda^{-1})$$

$$\Pi_1(1) = I - \sum_{i=1}^{k-1} \Gamma_i - (k-2)\,\Pi$$

then

$$\beta_0' = (\alpha'\alpha)^{-1}\alpha'\{I + [\sum_{i=1}^{k-1}\Gamma_i - I]\beta_{\perp}[\alpha'_{\perp}\Pi_1(1)\beta_{\perp}]^{-1}\alpha'_{\perp}\}\mu. \tag{16}$$

The presence of a dummy variable such as D85 in the error correction model with coefficient vector δ, indicates a shift in the constant vector to $\mu+\delta$. This reflects either a change in the deterministic trends of the variable, a change in the means of the cointegration vectors, or both. When β and α remain unchanged in the presence of a dummy variable, as they do in the models estimated above, estimates of β'_0 and μ_0 for the subsamples through 1985:2 and since 1985:3 can be computed from (15) and (16) as

$$\mu_0 = \beta_{\perp}[\alpha'_{\perp}\Pi_1(1)\beta_1]^{-1}\alpha_{\perp}(\mu + \delta \mathrm{D85}) \tag{17}$$

and

$$\beta_0' = (\alpha'\alpha)^{-1}\alpha'\{I + [\sum_{i=1}^{k-1}\Gamma_i - I]\beta_{\perp}[\alpha'_{\perp}\Pi_1(1)\beta_{\perp}]^{-1}\alpha'_{\perp}\}(\mu + \delta \mathrm{D85}). \tag{18}$$

If $\alpha_{\perp}\delta=0$ then there is no change in the deterministic trends of the variable, and the dummy variable reflects only a change in the means of the cointegration vectors. Conversely when

$$\{I+[\sum_{i=1}^{k-1}\Gamma_i - I]\ \beta_{\perp}\ [\alpha'_{\perp}\Pi_1(1)\beta_{\perp}]^{-1}\alpha_1\}\,\delta=0$$

then there is no change in the means of the cointegration vectors and the dummy variable reflects only a change in the deterministic trends, or drift, of the individual variables.

From the estimates in Table 4, the computed shift in the mean of the cointegration vector after 1985:2 is 0.418 (after normalization of the cointegration vector). This point estimate suggests a substantial *increase* in equilibrium real M2+CD after the middle of 1985.

Johansen and Juselius (1989) develop a test for the absence of deterministic trends in cointegrated variables that is easily modified to test the hypothesis that the *shifts* in the deterministic trends are zero.[9] The computed value of the χ^2 statistic for this test is 3.59,

[9]The relevant hypothesis test is H_2^* in H_2 in the notation of Johansen and Juselius (1989). The only difference between this situation and that constructed by Johansen and Juselius is that here the D85 dummy variable is substituted for the constant vector in the construction of the test statistic.

with three $(p-r)$ degrees of freedom. Thus we do not reject the hypothesis that the D85 variable represents just an increase in the equilibrium level of real M2+CD balances subsequent to the introduction of large denomination time deposits. This result strongly supports the hypothesis that there is a portfolio shift into large denomination time deposits from other non M2+CD assets after the deregulation in 1985. The stability of the estimated coefficients of the error correction model after we have allowed for a one-time change in the equilibrium level of real M2+CD and the absence of any significant changes in the drift of real M2+CD appears to be inconsistent with the hypothesis that continuing capital gains from land and equity revaluations are shifting the real M2 – real GNP relationship.

Strictly interpreted, the D85 variable indicates a one-time shift in equilibrium real M2+CD balances after the deregulation of large time deposits. In fact, this deregulation has involved several steps, each of which has lowered the minimum balance requirement for large denomination time deposits. It would be interesting to know whether the reductions in the minimum balance requirement had any impact on equilibrium real M2+CD balances, beyond the impact of the initial interest rate deregulation. Unfortunately, such precise discriminations are not possible with the available data since there are only 17 quarterly observations from 1985:3 through 1989:3.

In an analysis of the equilibrium demand for Japanese real M1 balances, Rasche (1990) finds that a stable error correction model existed throughout the 80s after the introduction or a dummy variable which is 1.0 after 1981 (D82). That analysis shows the role of the dummy variable is to incorporate shifts in the deterministic trends in the nonstationary variables (M1 velocity and the call rate) which appear in the cointegration vector. There is no evidence of a change in the equilibrium income or interest elasticities of the demand for real M1, nor is there any supprt for the hypothesis that the equilibrium level of real M1 balances changed during the 80s. The dummy variable only reflects the maintained interdependence of the trends in M1 velocity and the call rate, imposed by the unchanged equilibrium demand function (cointegration vector) for real M1.

There is no evidence for this effect in the results shown here for M2+CD. The explanation for this is the difference in the structure of the equilibrium demand function for real M1 compared to that for real M2. Rasche (1990) finds a significant negative equilibrium interest elasticity of the demand for real M1. Here we conclude that the equilibrium elasticity of the demand for real M2 with respect to interest rate spreads is not significantly different from zero, and the cointegration vector involves only real M2 and real GNP. Alternatively, we conclude that in equilibrium M2 velocity depends only on real GNP (since the estimated income elasticity is significantly greater than one) and not on nominal interest rates. Out of equilibrium, the behavior of M2 velocity depends only on real GNP, inflation, and *interest rate spreads*. Thus, even if there is a change in the deterministic trend in Japanese interest rates in the early 80s, as postulated in Rasche (1990), there should not be a change in the trend in M2 velocity. Thus there is no inconsistency between the appearance of the D82 dummy variable in the error correction model for M1 and its absence from this analysis.

A final caution that should be noted is the effect of the most recently available data point (1989:3). It can be seen in Figures 2 and 3 that the addition of the final data point to the recursive regression analysis, causes a reduction in the estimated equilibrium real income elasticity and a decrease in the precision of the estimate in the model with the D85 variable. Both of these changes are large relative to the fluctuations observed in the shorter samples. It is possible that this is just a random movement. However, as noted earlier, at the end of 1989:2 money market certificates first became available to individuals. While maximum interest rates on MMCs are not totally deregulated, they are closely tied to market rates of interest and provide a significant liberalization of the options available to individuals with smaller wealth holdings. It is possible that this could generate another upward shift in the level of equilibrium real M2+CD balances such as we believe occurred with the deregulation of rates on large time deposits. This hypothesis can only be confirmed or refuted as the future unfolds.

E. The error correction structure as a "real" or "nominal" adjustment mechanism

Since considerable attention has been given to the issue of "real" or "nominal" adjustment mechanisms in the context of single equation partial adjustment models, some mention of the implied impact of inflation on real balances within the error correction model is appropriate. The reader is cautioned that the equation developed here is *not* a short-run demand function for real M2+CD. That demand function is *not identified* in this analysis (see Rasche (1990) for a detailed discussion of this issue). Instead the equation discussed in this section is only one of a set of reduced form equations.

The error correction model in Table 4 is written in terms of nominal M2+CD, real GNP, the price level, and the interest rate spread. There are a number of algebraically equivalent alternative representations. In particular a representation in terms of *real* M2+CD, real GNP, the price level and the interest rate spread can be obtained by subtracting the third equation of the system in Table 4 from the first equation. The resulting *reduced form* equation for real M2+CD is:

$$\begin{aligned}\Delta \ln(\mathrm{M2}/P) &= .0117 + .0090\ \mathrm{D85} + .0026\Delta S_{t-1} + .0011\Delta S_{t-2} \\ &+ .2143\Delta \ln y_{t-1} + .0276\Delta \ln y_{t-2} + .5007\Delta \ln(\mathrm{M2}/P)_{t-1} \\ &- .1389\Delta \ln(\mathrm{M2}/P)_{t-2} + .0120\Delta \ln P_{t-1} - .1419\Delta \ln P_{t-2} \\ &- .0519[-\ln(\mathrm{M2}/P)_{t-3} + 1.166\ln y_{t-3}] + (\epsilon_{1t} - \epsilon_{3t}). \qquad (19)\end{aligned}$$

Note that given the history of real balances, $\Delta \ln P_{t-1}$ has almost no effect on current real balances. This does not appear to be true for $\Delta \ln P_{t-2}$ whose coefficient is substantially different from zero. Therefore the estimated error correction model suggest some independent role of inflation in the adjustment of real balances and thus is more closely related to a "nominal" adjustment mechanism than to a "real" adjustment mechanism.

VI. Conclusions

The results of the analysis suggest that the equilibrium demand for real M2+CD in Japan over the past 25 years has constant elasticities with respect to real income and the spread between the own rate of interest and the call rate. The equilibrium real income elasticity is around 1.2 and is significantly greater than 1.0. The equilibrium elasticity with respect to the interest rate spread is not significantly different from zero. Consequently there exists a stationary linear combination of real M2+CD and real GNP over this period, but the velocity of M2+CD is nonstationary.

Real GNP, inflation *and* changes in interest rate spreads all have impacts on short-run movements in real M2+CD through a reduced-form vector autoregressive process, even though in the long-run the path of real M2+CD is driven only by real income.

The evidence suggests that the behavior of M2+CD *has* been affected by financial deregulation in the 1980s in two distinct ways. First the deregulation of interest rates on large time deposits in 1985 appears to produce a one-time upward shift in the equilibrium demand for real M2+CD. It appears that this change took the form of a discrete jump, though the available data do not permit discrimination of a discrete change at the time of deregulation from gradual changes over a short period following deregulation. In addition, the data are too limited to identify any subsequent effects of deregulation associated with the reduction of minimum balance requirements on large time deposits.

A second source of effects of deregulation is through the behavior of the interest rate spread. With interest rate deregulation deposit rates can in principle adjust more quickly to changes in market rates of interest. This in turn can affect the short-run relationship between real M2+CD and real GNP, though not the equilibrium relationship between the two. Our analysis suggests that this latter impact of deregulation has been relatively small since the time series characteristics of the interest rate spread have remained quite stable as deposit rates have been deregulated.

Tomoo Yoshida: Research Division I, Institute for Monetary and Economic Studies, The Bank of Japan

Robert H. Rasche: Professor of Economics, Michigan State University, East Lansing, Michigan, U.S.A.

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