Financial Liberalization, the Wealth Effect, and the Demand for Broad Money in Japan

Toshitaka Sekine

This paper examines the demand for broad money in Japan from 1975 to 1994. In spite of the large shocks resulting from financial liberalization and the subsequent "boom and bust" of the "bubble" economy, the paper confirms that a stable money demand function can still be set up by taking proper account of financial liberalization and the wealth effect, and by adopting an adequate econometric strategy. In addition, a super exogeneity test is conducted, and its implication is considered in the context of the monetary transmission mechanism.

Key words: Demand for money; Monetary policy; Wealth effect; Cointegration; Error correction; Exogeneity; Japan

Policy Development and Review Department, International Monetary Fund (tsekine@imf.org)

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I. Introduction

A stable money demand function is thought to be essential for monetary policy, because it is supposed to play a key role in enabling the monetary authorities to set an intermediate target, money supply, consistent with final targets such as the inflation rate and/or real income. Because of its importance in macroeconomic policy, a large amount of research has been conducted on this topic (see Judd and Scadding [1982], Goldfeld and Sichel [1990], and Laidler [1993] for extensive surveys). This is true in Japan as well. Recent work mainly concerns the question of whether financial liberalization and the subsequent "boom and bust" of the "bubble" economy make the money demand function in Japan unstable. This paper shows that a stable money demand function can still be set up by taking proper account of financial liberalization and the wealth effect, and by adopting an adequate econometric modeling strategy.

In particular, the paper confirms the role of wealth in a Japanese broad money demand function. In fact, in the literature on the demand for money, the wealth effect itself is an old issue and we can find it in such classical papers as Goldfeld (1976) and Friedman (1978). Here, the effect is carefully examined by using a definition of wealth that includes not only financial assets but also nonfinancial assets, and adding it to income as a second scaling variable.

On the other hand, for its methodology, the paper exploits the "general-to-simple" approach. That is, it derives a single equation of the demand for broad money from a sequential reduction of a general model of system equations. The technique of cointegration analysis is employed and the validity of the model is finally checked by a super exogeneity test.

The outline of the paper is as follows. Section II briefly summarizes empirical research into Japanese money demand functions, and refers to financial liberalization and the "bubble" economy as the background of these works. Section III explains the data set. Section IV describes the sequence of investigations and the result of seeking a stable money demand function. A super exogeneity test is conducted. Section V concludes the paper.

II. Financial Liberalization and the "Bubble" Economy

There are a number of studies concerning money demand in Japan, but most of them—and particularly the early works, such as that of Tsutsui and Hatanaka (1982), and the excellent survey of Tsutsui (1986)—are only available in Japanese. Apart from the studies cited below, Hamada and Hayashi (1985) and Ishida (1984) are available in English.

First, concerning the demand for narrow money (M1),[1] Tsutsui and Hatanaka (1982) suggested the possibility of "missing money" from 1980 by applying a partial adjustment model on the lines that Goldfeld (1973, 1976) used for U.S. money demand. Afterward, this was confirmed by Rasche (1990), who used a vector

1. The definitions of money used in Japan are summarized in the next section.

autoregression (VAR) including an error correction term and a dummy variable that was set to unity from 1982.

On the other hand, concerning the demand for broad money (M2+CDs), it was generally believed that a partial adjustment model could trace the actual development of data satisfactorily up to the mid-1980s. In 1985, money market certificates (MMCs) and large time deposits were introduced,[2] and the financial market in Japan was liberalized rapidly. Moreover, the "bubble" boom followed. It reached its peak around the end of 1989 and burst after that. Considering this large shock to Japan's economy, some researchers suggested that there might be a shift in the money demand function (Bank of Japan [1988, 1992], Suzuki [1989], and Yoshida and Rasche [1990]). However, other researchers such as Ueda (1988), Corker (1989), Shiba (1989), Yoshida (1990), Arize and Shwiff (1993), Baba (1995), and Ishida and Shirakawa (1996) (the latter two studies are available only in Japanese) showed that the money demand function remained stable.

As seen above, financial liberalization and the "bubble" economy are the background to recent discussions of broad money demand. First, let us consider financial liberalization in Japan.[3] Apart from McKenzie (1992), who examined the impact of the amendment of the Foreign Exchange Law in 1980, the above authors are mainly concerned with the effect of domestic financial liberalization. Even though market-determined deposit interest rates were first introduced for CDs in 1979, it is widely believed that the introduction of MMCs and large time deposits in 1985 had a greater impact on the Japanese financial market. This is because initially the issue of CDs was highly regulated in the form of a minimum deposit requirement (¥500 million) and a maximum limit for a bank (10 percent of its capital). More or less similar regulations were imposed on MMCs and large time deposits at the outset, but they were relaxed—the deregulation of time deposit interest rates was completed in 1993. The deregulation of demand deposit interest rates was completed in 1994.

On the other hand, partly because of the extremely low level of interest rates after the Plaza Accord of September 1985, the prices of existing assets such as land and stocks increased significantly. It is thought that banks in Japan were able to lend more money to companies and investors by using such assets as highly appreciated collateral. This money was invested not only in the newly liberalized deposits, but also again in land and stocks, increasing their prices. The value of collateral then increased and the financial sector lent more money to the nonfinancial sector, and so on. This circle—the "bubble" economy—became particularly noticeable after 1987. The bubble reached its peak around the end of 1989 and "burst" after the Bank of Japan tightened monetary policy and the Ministry of Finance restrained bank lending for land purchasing. Since the bubble burst, the circle has gone in the opposite direction. That is, banks have been able to lend less and less money as the value of collateral has decreased following a drop in asset prices.

These episodes suggest that we should be extremely careful about the impact of financial liberalization and the wealth effect when we model the demand for money

2. MMCs had interest rates that followed those of CDs, and their minimum deposits (initially, ¥50 million) were smaller than those of large time deposits (initially, ¥1 billion). Eventually, MMCs were absorbed by large time deposits because of the subsequent reductions of the latter's minimum deposits (e.g., ¥3 million from November 1991).

3. The details of financial liberalization in Japan can be found in Nakao and Horii (1991).

in Japan. Most researchers have incorporated the impact of financial liberalization by taking into account the liberalized interest rates. This paper follows this line. On the other hand, only Ueda (1988), Corker (1989), Bank of Japan (1992), Arize and Shwiff (1993), Baba (1995), and Ishida and Shirakawa (1996) have examined the wealth effect. Baba claimed that his money demand function violated the exogeneity condition with respect to wealth. As we will see below, the definition of wealth employed by these studies did not have enough coverage, and hence this paper will use more comprehensive wealth data.

III. Data Description

A. Money

The definitions of money in Japan are as follows:

[1] M1 = cash currency + demand deposits;

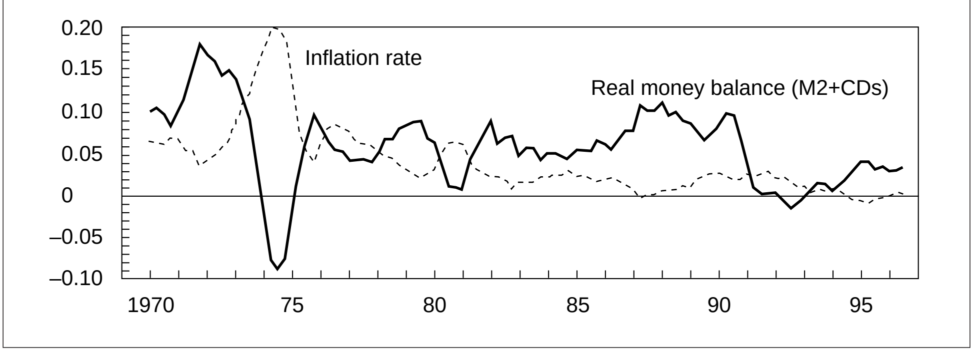
[2] M2+CDs = M1 + time deposits + certificates of deposit; and

[3] broad liquidity = M2+CDs + deposits (including CDs) of post offices, agricultural cooperatives, etc., + money trusts and loan trusts of domestically licensed banks (excluding foreign trust banks) + bonds with repurchase agreement + bank debentures + government bonds + investment trusts + money deposited other than money in trust + foreign bonds.

Most researchers have estimated M2+CDs demand as broad money demand partly because broad liquidity data are only available from 1980. M1 and M2+CDs data are available from 1955 on end-of-month bases, and from 1963 (M1) and 1967 (M2+CDs) on daily average bases. In this paper, the three-month average of end-of-month M2+CDs is used because the wealth data that will be discussed below are end-of-quarter based and the end-of-month M2+CDs data are more consistent with this scale variable.

Figure 1 shows the annual change in real balance of M2+CDs deflated by the GDP deflator. There seems to be a regime shift in the series around 1973 that reflects the transition from a high-growth to a stable-growth era in Japan's economy. After the "bubble" economy, it plunged into negative growth in 1992, then recovered

Figure 1 Annual Growth of Real Balance and Annual Inflation



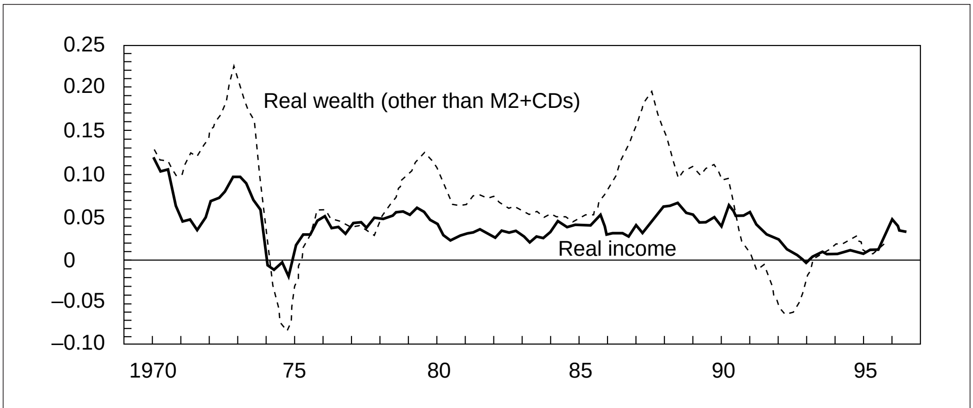
somewhat. Subsequently, it continued to show moderate growth until 1996. As seen above, some researchers have argued that this swing in the series came from a shift in the money demand.

B. Scale Variables

In the literature on money demand functions, income or wealth is often chosen as the scale variable (see Judd and Scadding [1982] and Goldfeld and Sichel [1990], for example). The portfolio approach developed by Tobin (1969) says that both income and wealth play an important role in the demand for money. Since we are modeling the demand for broad money, the portfolio theory is likely to hold. Moreover, money is likely to be used for transactions of asset dealing such as purchasing land or equities. Considering the bubble episode, we should take into account not only the conventional measure of transactions—income—but also wealth all the more.

However, the wealth data seem problematic. Conceptually, there is no reason why we should restrict ourselves to financial assets. The importance of including physical wealth in consumption functions is pointed out by Muellbauer (1991). This is the case with money demand functions as well. In fact, Corker (1989), Bank of Japan (1992), Arize and Shwiff (1993), Baba (1995), and Ishida and Shirakawa (1996) used financial assets held by personal and nonfinancial corporate sectors in Flow of Funds (FOF) data when they investigated the wealth effect in their money demand functions. On the other hand, the System of National Accounts (SNA)-based definition of wealth, which covers not only financial assets but also nonfinancial assets such as land, housing, inventories, and so on, is more promising. The major drawback is that it is only available on an annual basis. This paper tries to solve the problem by interpolating it into a quarterly series. The procedure of the interpolation is summarized in the Appendix. Since wealth contains money as its component, the M2+CDs balance is subtracted from the wealth data.[4] The resulting data, which are deflated by the GDP deflator, are shown in Figure 2, together with real GDP, and the

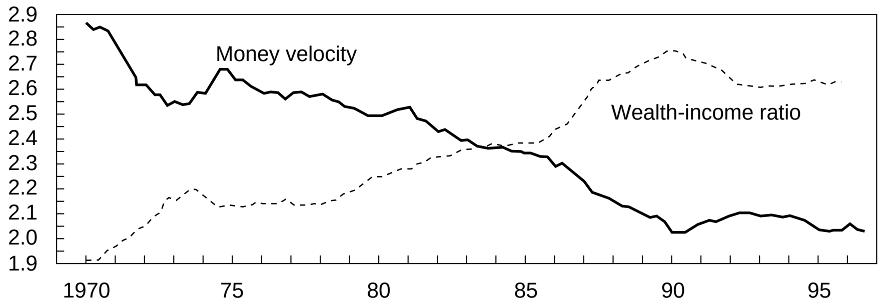
Figure 2 Annual Growth of Real Income and Real Wealth



4. At the end of 1995, M2+CDs amounts to nearly 40 percent of FOF financial assets held by personal and nonfinancial corporate sectors. This large proportion could be the reason why the exogeneity condition did not hold in Baba (1995).

wealth-income ratio can be found in Figure 3.

Figure 3 Money Velocity and the Wealth-Income Ratio



Note: 1. In log scale. Match means and ranges.

C. Opportunity Costs

In order to take into account the opportunity costs, two interest rates, that is, the interest rate of money itself (hereafter, the own interest rate) and the return on rival assets, are considered.

First, the own interest rate. One of the keys to success in modeling the demand for narrow money in the United Kingdom and the United States by Hendry and Ericsson (1991) and Baba, Hendry, and Starr (1992) is modeling the effect of financial liberalization by introducing the "learning-adjusted" interest rate. This paper will imitate them by weighting the corresponding interest rates. That is, the own interest rate is calculated as

$$\begin{aligned}\text{(The own interest rate)}_t = {} & w_{m1t} \cdot \text{(regulated time deposit interest rate)}_t \\ & + w_{m2t} \cdot \text{(free time deposit interest rate)}_t \\ & + w_{m3t} \cdot \text{(CD rate)}_t,\end{aligned}$$

where $w_{m1t}, \ldots, w_{m3t}$ are the calculated shares of each deposit.

Next, the interest rate on rival assets. Possible candidates are interest rates on assets that are included in broad liquidity but not in M2+CDs. That is, the rival rate is calculated as the maximum return on the following assets:

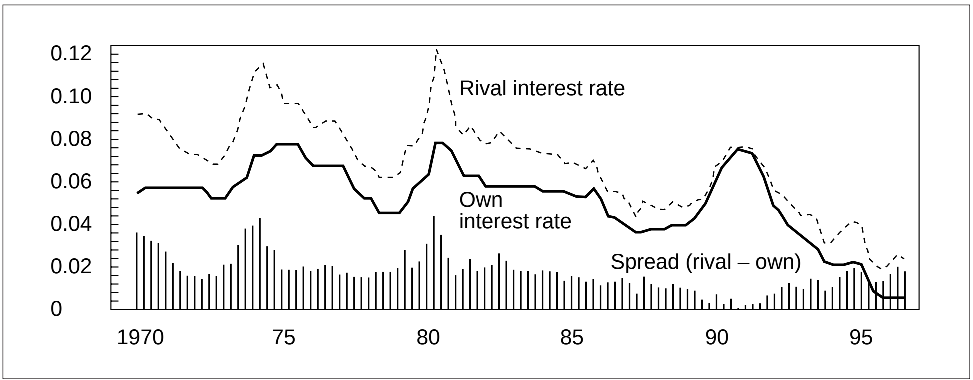
[1] three-month *gensaki* (bonds with repurchase agreement);
[2] five-year money trusts;
[3] five-year loan trusts;
[4] five-year bank debentures (subscription market rates, yields to subscribers);
[5] five-year bank debentures (secondary market rates, over-the-counter standard bond quotation); and
[6] five-year postal savings[5] (*teigaku chokin*, saving certificates).

5. The five-year rate was calculated as the compound interest rate from the corresponding uncompound (single) interest rate.

Other than these assets, which are characterized by broad liquidity, the return on stock could also be included. However, a preliminary investigation showed that the coefficients on the stock return were positive since the return is dominated by the wealth effect or the transaction effect. For this reason, the paper does not include the stock return as a rival interest rate.

In Figure 4, the own interest rate and the rival interest rate are plotted together with the spread between them.

Figure 4 Interest Rates



D. Time Properties of the Series

Before going into modeling the demand for money, the time property of each series is checked by an augmented Dickey-Fuller test (ADF) (Table 1).[6] The procedure to determine the inclusion of deterministic terms and lag length is as follows (see Banerjee et al. [1993] for a rationale of the procedure). First, check the significance of a constant and a trend by the conventional student-t distribution. If the trend is significant, then retain these variables and check the significance of lagged terms by the conventional student-t distribution, where the lag length is originally set at five. Then determine the lag length as the highest lag that is significant within five lags.

Table 1 ADF Statistics for Testing a Unit Root

Variables	t-adf	Lag	Additional regressor
m	–1.92	4	Constant
p	–2.89	2	Constant
y	–1.66	4	Constant, trend
w	–2.71	4	Constant, trend
Rm	–3.86*	3	Constant, trend
Rr	–3.52*	4	Constant, trend
m – p	–2.08	3	Constant, trend
Δp	–4.41**	1	Constant, trend

Notes: 1. The estimation periods are 1976/III–1994/IV.
2. t-adf denotes the t-value of the augmented Dickey-Fuller (ADF) statistics. The critical values are based on response surface in MacKinnon (1991). Lag denotes its lag order.
3. Here and elsewhere in the paper, * and ** denote rejection at the 5 percent and 1 percent critical values.
4. Lowercase letters denote logarithms of the corresponding variables, and Δ denotes a difference operator.

6. All calculations in this paper, except otherwise noted, are conducted by PcGive 9.0 (Doornik and Hendry [1996]) or PcFiml 9.0 (Doornik and Hendry [1997]).

In the case where the trend is not significant, drop the trend and test the significance of the constant term. If it is significant, then retain it in the ADF regression and determine the lag length just mentioned. Otherwise, drop the constant term from the regression and determine the lag length.

Sample periods for estimation are after 1975, that is, after a regime shift from a high-growth to a stable-growth era in Japan. These sample periods are chosen because, as is well known, the ADF test has low power in the presence of regime shifts and we should pay more attention to this problem in the case of Japan, where a regime shift is apparent—see Soejima (1995), for example.

Test results show that all the variables except two interest rates and the inflation rate do not reject the null hypotheses that they have unit roots. However, it seems safe to treat interest rates, whose unit roots are rejected at only the 5 percent significance level, as $I(1)$ variables since these t-adf become insignificant if the constant and the trend are dropped from the regression, whereas that of the inflation rate is significant even in such a case.

IV. Modeling the Demand for Money

Using the data discussed above, we proceed to model a Japanese money demand function. The modeling strategy followed in this paper is the "general-to-simple" approach (Hendry and Ericsson [1991], Baba, Hendry, and Starr [1992], Ericsson and Sharma [1996], etc.). Summarizing the procedure, first a system analysis of cointegration is conducted and exogeneity status with respect to long-run parameters is checked. Then, we proceed to a single equation analysis. In this framework, the long-run solution is examined again by an autoregressive distributed lag (ADL) model and is compared to the cointegrating vector in the above system analysis. After confirming an error correction term, the model is reduced so that it becomes parsimonious. Various tests, particularly those concerning stability, are conducted. Finally, a test for super exogeneity follows.

A. System Analysis of Cointegration

First, an error correction term is found as a cointegration relationship. Following the approach cited above, we will try a Johansen test (Johansen [1988], Johansen and Juselius [1990]), since its VAR approach is the most general model. We start with a six-variate VAR that involves M2+CDs, m_t; price, p_t; income, y_t; wealth other than M2+CDs, w_t; the own interest rate, Rm_t; and the rival interest rate, Rr_t.[7] Four lags are taken for each variable, and a trend,[8] a constant, and a second oil crisis dummy—

7. Lowercase letters denote logarithms of the corresponding variables.
8. A deterministic time trend is included for the following statistical reason. When variables in a cointegrating vector have deterministic time trends and these trends are linearly independent with coefficients that bring a linear combination of stochastic components of the variables (i.e., the case of stochastic cointegration; see Ogaki [1993] for further explanation, and also see Soejima [1996] for critics on the existing money demand functions from the viewpoint of stochastic cointegration), a cointegrating vector should also contain a deterministic time trend in order to absorb the linear independence among the deterministic time trends. For this reason, we had better check the significance of a deterministic time trend in our model.

which takes a value of unity at 1980/II, nought otherwise—are included. The sample period is from 1975/I to 1994/IV.

The performance of the VAR is shown in Table 2 [1]. *AR* represents Lagrange-multiplier tests for the fifth-order autocorrelated residuals of each variable [~ $F(5, 48)$] and the system as a whole [~ $F(180, 114)$]. *Normality* represents the Doornik-Hansen normality tests for each variable [~ $\chi^2(2)$] and the system [~ $\chi^2(12)$]. *ARCH* is a test for the fourth-order autoregressive conditional heteroscedasticity [~ $F(4, 45)$]. X_i^2 represents the White heteroscedasticity tests for

Table 2 System Analysis of Cointegration

[1] Properties of VAR residuals							
	m	*p*	*w*	*y*	*Rm*	*Rr*	Vector
AR	2.22	2.33	0.27	0.76	1.83	0.77	1.70**
Normality	1.90	1.44	0.97	2.13	2.16	1.52	10.18
ARCH	0.69	0.51	0.57	0.96	0.97	0.42	—
X_i^2	0.09	0.07	0.14	0.06	0.06	0.06	1,063
[2] Tests for the number of cointegrating vectors							
Eigenvalues	0.423	0.392	0.302	0.193	0.150	0.061	
Hypotheses	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$	$r \leq 4$	$r \leq 5$	
λ_{max}	44.0*	39.8*	28.8	17.2	13.0	5.1	
λ_{trace}	147.9**	103.9**	64.1*	35.3	18.1	15.1	
[3] Standardized adjustment coefficients α							
m	–0.30	0.14	0.03	0.007	–0.12	0.02	
p	–0.06	–0.30	0.003	0.004	–0.08	0.05	
y	–0.13	0.08	–0.05	0.012	0.25	–0.03	
w	–0.03	0.08	0.005	0.009	–0.54	–0.22	
Rm	0.02	0.11	–0.01	0.002	–0.15	0.03	
Rr	–0.13	0.06	–0.05	–0.006	–0.16	0.04	
[4] Standardized eigenvectors β'							
	m	*p*	*y*	*w*	*Rm*	*Rr*	Trend
	1.00	–0.91	–0.62	–0.35	–3.33	3.84	–0.0017
	–0.92	1.00	0.48	0.40	0.84	–1.86	0.0011
	–3.41	2.85	1.00	1.31	3.41	–0.70	0.0212
	6.96	–5.47	–12.52	1.00	–14.68	8.84	–0.0228
	–0.03	0.06	–0.12	0.04	1.00	–0.76	0.0005
	0.16	–0.36	0.18	–0.02	–0.97	1.00	–0.0022
[5] Weak exogeneity tests for long-run parameter							
	m	*p*	*w*	*y*	*Rm*	*Rr*	
$\chi^2(1)$	9.89**	1.01	1.02	1.09	1.01	1.25	

Notes: 1. The vector autoregression model includes four lags on each variable (*m*, *p*, *y*, *w*, *Rm*, *Rr*), a trend, a constant, and the second oil crisis dummy ID80/II. The estimation period is 1975/I–1994/IV. The trend is restricted so that it lies in the cointegration space. The VAR model can be reexpressed as a vector error correction model:

$$\Delta X_t = \alpha\beta' X^*_{t-1} + \sum_i \Gamma_i \Delta X_{t-i} + \Phi d_t + e_t,$$

where X_t is (m_t, p_t, y_t, w_t, Rm_t, Rr_t), X_t^* is (X_t, trend) and d_t is deterministic components other than the trend.

2. The statistics λ_{max} and λ_{trace} are Johansen's maximal eigenvalue and trace eigenvalue statistics for testing cointegration.

3. The weak exogeneity test statistics are evaluated under the assumption that $r = 2$, *m* is homogeneous with respect to *p* in the long run, and the second cointegrating vector has no feedback effect on *m*. Its asymptotic distribution is $\chi^2(1)$, except that for *m*, which is $\chi^2(2)$.

each variable [~ $F(50, 2)$] and the system [~ $\chi^2(1050)$]. See Doornik and Hendry (1996, 1997) for details of these tests. They reveal that each equation is reasonably congruent. However, the system as a whole shows a sign of residual autocorrelation. This could undermine the validity of the following system analysis, but with this caveat, we proceed to the next step.

Table 2 [2] is the test for the number of cointegrating vectors. From this, we may conclude that there are two, possibly three, cointegration relationships in this VAR model. This is because, based on the maximum eigenvalue test, the hypothesis of one cointegrating vector is rejected at the 5 percent significance level, but that of two cointegrating vectors cannot be rejected. Meanwhile, the trace test leads us to accept three cointegrating vectors. Below, some tests are based on the assumption of two cointegrating vectors, but the final outcome does not change even if we assume that there are three cointegrating vectors.

From the adjustment coefficients in Table 2 [3], it seems quite likely that the second and possibly the third cointegrating vector are not relevant for the money equation. That is, the first cointegrating vector goes into the money equation with a feedback parameter of –0.30, while the second and the third cointegrating vector have feedback coefficients of 0.14 and 0.03, which are less significantly different from nought.[9] Then, if we concentrate on the first eigenvector in Table 2 [4], the coefficient on p is –0.91 and the long-run homogeneity with respect to p seems to hold. These intuitions are formally checked by the restriction test on α and β' and accepted [1.01 ~ $\chi^2(1)$]. On the other hand, the test of equality in absolute values between Rm and Rr in the first eigenvector is rejected [6.6 ~ $\chi^2(1)$].[10]

Furthermore, assuming that m is homogeneous with respect to p in the long run and the second cointegrating vector has no feedback effect on m, we test whether the first cointegrating vector has a feedback effect on variables other than money. As Johansen (1992) shows, these tests correspond to weak exogeneity tests for the long-run parameter and such a condition is indispensable for a single equation analysis of an error correction model.[11] As Table 2 [5] shows, this condition seems to hold since zero coefficients on α of the first cointegrating vector into the other variables cannot be rejected. Moreover, the test statistics of jointly setting zero restrictions on these coefficients is accepted [10.93 ~ $\chi^2(5)$].

Finally, maintaining these restrictions, the first cointegrating vector becomes

$$m - p = 0.42y + 0.42w + 2.37Rm - 2.99Rr + 0.002\text{trend}, \qquad (1)$$

and the adjustment coefficient for money is –0.50 with a standard error of 0.09.

B. Single Equation Analysis

In order to investigate the cointegration relationship further, an ADL model is estimated as a general model in the single equation analysis. From the above system analysis, we can

9. Of course this is a subjective judgment, but it is formally tested as we will see.
10. This might result from different treatment of the two interest rates; the own interest rate is calculated as the *average* of the deposit rates, while the rival interest rate is calculated as the *maximum* of the rival assets' rates.
11. The concept of weak exogeneity is summarized, together with that of super exogeneity, in Section IV.C below.

safely assume the long-run homogeneity with respect to p, and hence regress $(m-p)_t$ on the other five variables (y_t, w_t, Δp_t, Rm_t, Rr_t) plus a constant and a trend.[12] Other than Δp_t, which includes only three lags, four lags are taken for each variable. Table 3 is the outcome of the estimation. Other than those referred to in the Johansen test, *RESET* is a test for functional form. There is no sign of a violation of these diagnostic checks. The solved static long-run solution is

$$\underset{}{(m-p)_t} = \underset{(0.12)}{0.42y_t} + \underset{(0.04)}{0.42w_t} + \underset{(0.48)}{2.37Rm_t} - \underset{(0.47)}{3.03Rr_t} + \underset{(0.001)}{0.002\text{trend}}, \tag{2}$$

which surprisingly resembles the first cointegrating vector (equation [1]) in the Johansen test. These equations suggest that the long-run elasticity with respect to income is around 0.4. In Japanese broad money demand functions, Yoshida and Rasche (1990) find that the long-run income elasticity is 1.2 according to a Johansen test. Further, Fujiki and Mulligan (1996) estimate the income elasticity at 1.2 to 1.4 by cross-sectional regressions using data for Japanese prefectures (M2 minus cash). Looking at the above regression, one may infer that such large elasticity beyond unity might result from the omission of an important scale variable, i.e., wealth.

In addition, the stability of the estimated cointegration relationship is tested by the procedure developed by Hansen (1992). Although his fully modified (FM)

Table 3 Autoregressive Distributed Lag Model

Variables	Lag *i*					
	0	1	2	3	4	Σ
m – p	–1.000	0.565	–0.025	0.060	–0.099	–0.499
	—	(0.142)	(0.150)	(0.153)	(0.112)	(0.122)
y	–0.015	0.159	0.037	–0.101	0.129	0.209
	(0.112)	(0.127)	(0.125)	(0.120)	(0.104)	(0.095)
w	0.066	0.153	–0.022	–0.037	0.049	0.210
	(0.064)	(0.105)	(0.102)	(0.099)	(0.068)	(0.046)
Rm	0.493	0.217	0.422	0.510	–0.456	1.190
	(0.316)	(0.408)	(0.412)	(0.412)	(0.281)	(0.336)
Rr	–0.275	–0.342	–0.443	–0.631	0.180	–1.510
	(0.188)	(0.209)	(0.217)	(0.219)	(0.206)	(0.373)
Δ*p*	–0.790	0.181	0.124	–0.054	—	–0.539
	(0.152)	(0.173)	(0.164)	(0.159)	—	0.276
Constant	0.594	—	—	—	—	0.594
	(0.678)	—	—	—	—	(0.678)
Trend	0.0010	—	—	—	—	0.0010
	(0.0005)	—	—	—	—	(0.0005)
T = 1975/I–1994/IV, R^2 = 0.99, $\hat{\sigma}$ = 0.45 percent, *DW* =1.99, *AR*: $F(5, 45)$ = 2.31, *ARCH*: $F(4, 42)$ = 0.25, *Normality*: $\chi^2(2)$ = 3.84, *RESET* $F(1, 49)$ = 2.98.						

Note: 1. The figures in the parentheses are standard errors for coefficients.

12. Δ denotes a difference operator.

estimators give slightly different coefficients (they are rather similar to those of the unrestricted cointegrating vector in Table 2) such as

$$\underset{}{m - p_t} = \underset{(0.14)}{0.67y_t} + \underset{(0.03)}{0.31w_t} + \underset{(0.42)}{1.19Rm_t} - \underset{(0.37)}{1.46Rr_t} + \underset{(0.001)}{0.003\text{trend}} - \underset{(1.70)}{0.81}, \quad (3)$$

the test statistics confirm the stability of the relation. That is, Lc = 0.29, $MeanF$ = 5.02, $SupF$ = 21.70, where Lc and $MeanF$ capture a gradual shift that follows some Martingale process and $SupF$ is against a one-time break. None of them rejects the null hypothesis of no shift at the 5 percent significance level of his simulated critical values.[13]

Since we have a reliable long-run solution, the next step is to reduce the model to a more parsimonious one. Imposing the restriction on the coefficients of y_t and w_t as 0.4, we have the following result.

$$\begin{aligned}\Delta(m-p)_t = {} & \underset{(0.05)}{-0.30}(m-p-0.4y-0.4w)_{t-1} + \underset{(0.13)}{0.65Rm_{t-1}} - \underset{(0.12)}{0.75Rr_{t-1}} \\ & + \underset{(0.0002)}{0.0008\text{trend}} + \underset{(0.14)}{0.40\mathrm{M}\Delta y_{t-1}} + \underset{(0.05)}{0.18\Delta w_t} - \underset{(0.08)}{0.44E\Delta p_t} \qquad (4) \\ & + \underset{(0.16)}{0.55\Delta Rm_t} + \underset{(0.08)}{0.51},\end{aligned}$$

T = 1975/I–1994/IV, R^2 = 0.79, $\hat{\sigma}$ = 0.49 percent, DW = 2.05,
AR: $F(5, 66)$ = 0.96, $ARCH$: $F(4, 63)$ = 0.06, $Normality$: $\chi^2(2)$ = 1.05,
X_i^2: $F(16, 54)$ = 0.60, X_iX_j: $F(44, 26)$ = 0.74, $RESET$ $F(1, 70)$ = 2.44.

Here $M\Delta y_t$ is a two-quarter moving average ($\sum_{i=0}^{1}\Delta y_{t-i}$), and $E\Delta p_t$ is the data-based expectation of inflation ($\Delta p_t + \Delta^2 p_t$) à la Hendry (1995). X_iX_j is another type of White heteroscedasticity test. The signs of the coefficients allow for the ordinary interpretation. The estimated model retains the features of the original model. For example, the solved static solution of the model (putting zero to the Δ terms and ignoring a constant) is

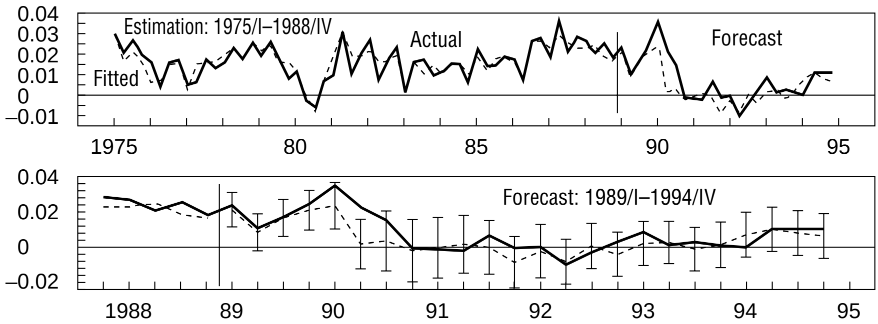
$$m - p = 0.4y + 0.4w + 2.19Rm - 2.54Rr + 0.003\text{trend}, \qquad (5)$$

and is almost identical to equations (1) and (2). In fact, a test to see whether the parsimonious model encompasses the ADL model gives the result 1.71, which follows $F(21, 50)$ and is accepted.

Next, we test the stability of this model. In Figure 5, a forecast test of 1989/I–1994/IV, which covers the peak and the subsequent bursting of the bubble, is reported. Although the model fails to forecast the actual outcome on 1990/II, it reasonably succeeds in forecasting the rest of the period. The "forecast" Chow test, which tests the equality between coefficients obtained from the pre-forecast interval and those from the forecast interval, is 1.25 ~ $F(22, 49)$ and is not rejected. Then, Figure 6 and Figure 7 show the stability of the model in the form of recursive tests of

13. The FM regression and the Hansen test are calculated by COINT 2.0 on GAUSS. Automatic bandwidth is used.

Figure 5 Forecast Test



Note: 1. Bars are approximate 95 percent confidence intervals of one-step forecasts.

Figure 6 Recursively Estimated Chow Tests

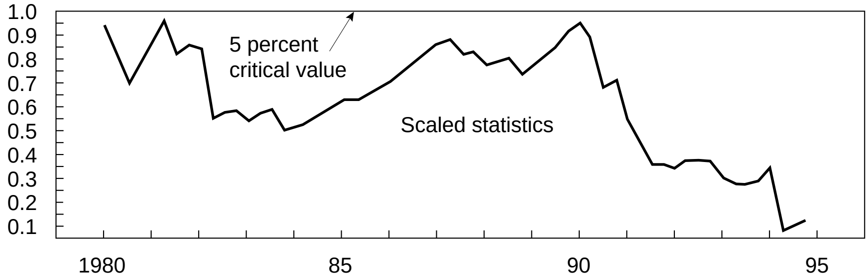
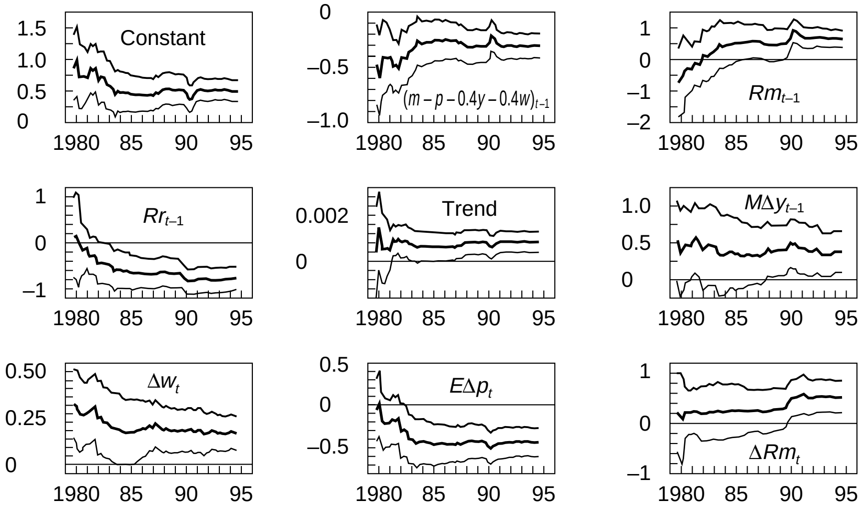


Figure 7 Recursively Estimated Coefficients



Note: 1. The upper and lower lines denote approximately two standard error bands.

the model. First, in Figure 6, the recursively estimated Chow tests are reasonably below their 5 percent critical values. Furthermore, in Figure 7, recursively estimated coefficients are plotted with their standard errors. In fact, almost all parameters show some fluctuations around 1990. However, these fluctuations are within bands of standard errors through the 1980s. For this reason, we may conclude that these parameters are reasonably stable.[14]

In sum, all of the evidence suggests that an appropriately modeled demand function is still stable even after financial liberalization and the bubble, which were believed to make it unstable.

C. Exogeneity Test

In order to confirm the validity of the above model, we further test the exogeneity status of two contemporaneous variables of interest, the inflation rate and the own interest rate.[15] The test applied below is a test for super exogeneity proposed by Engle and Hendry (1993).

According to Engle, Hendry, and Richard (1983), super exogeneity is defined as follows.[16] If the data generation process (DGP) can be expressed as a joint density, it can be factorized as

$$D_X(m_t, z_t|X_{t-1}, \theta) = D_{M|Z}(m_t|z_t, X_{t-1}, \phi_1)\cdot D_Z(z_t|X_{t-1}, \phi_2), \qquad (6)$$

where $X_t' = (m_t, z_t)$. In this form, z_t is super exogenous for ψ (the parameters of interest), if

[1] it is weakly exogenous for ψ; and

[2] ϕ_1 is invariant to a change in ϕ_2, i.e., $\partial\phi_1/\partial\phi_2 = 0$.

The former is the condition that enables us to obtain all the necessary information for efficient estimates of the parameters of interest from the conditional model $D_{M|Z}(m_t|z_t, X_{t-1}, \phi_1)$ alone. For this reason, this can be said to be one of the minimum requirements for the regression model, because we derive the linear regression model from the conditional model by approximating it to the normal distribution through adequate data transformation. That is,

$$D_{M|Z}(m_t|z_t, X_{t-1}, \phi_1) \approx N_{M|Z}(m_t^*|z_t^*, X_{t-1}^*, \phi_1^*), \qquad (7)$$

where $m_t^* = h_m(m_t)$, $z_t^* = h_z(z_t)$. More formally, the weak exogeneity condition is written as

[1] $\psi = f(\phi_1)$ alone, i.e., ψ does not depend on ϕ_2; and

[2] ϕ_1 and ϕ_2 are variation free, i.e., $\Phi_1 \neq g(\phi_2)$, where Φ_1 is the parameter space of ϕ_1.

14. Parameter stability of $E\Delta p$ and ΔRm_t will be further tested by the super exogeneity test in the next section.
15. In an error correction model, there are two types of parameters whose exogeneity conditions should be checked. One is long-run parameters, i.e., coefficients of error correction terms, and the other is short-run parameters, i.e., coefficients of contemporaneous variables in Δ form. The former can be tested by a restriction test on the adjustment coefficient, as we did in a system analysis of cointegration. On the other hand, the super exogeneity test discussed below is an exogeneity test for short-run parameters. As Urbain (1992) summarized, required exogeneity conditions differ depending on which parameters are of interest to researchers.
16. Other than cited here, see chapter 5 of Hendry (1995) and Ericsson (1992) for various concepts of exogeneity.

On the other hand, the latter condition claims that the parameters of interest remain constant when the process of generating z_t, the marginal model $D_Z(z_t|X_{t-1}, \phi_2)$ in equation (6), changes. For example, if z_t is generated by a policy rule, variance of the marginal process is brought about by a policy intervention. Although no one seriously believes that a model describing an economic phenomenon would be absolutely robust for any change in an input, there may be robustness for a certain class of input changes including some policy interventions.

Putting the above two conditions together, super exogeneity enables us to conduct policy analysis or simulation. This is because, regarding z_t as policy instruments, from the first condition we can derive the parameters of interest ψ by regression, and from the second condition we can simulate the change in policy rule—say, $z_{T+1} \rightarrow z'_{T+1}$ as $\psi z'_{T+1}$—since ψ does not change as a result of the policy intervention.

In addition, there is an important derivation here. Engle and Hendry (1993) prove that if z_t is super exogenous for ψ in the conditional model, the reverse cannot happen, that is, m_t cannot be super exogenous in a simple reverse model (z_t is regressed on m_t). Intuitively their proof implies that when z_t is super exogenous, the reverse model cannot be constant. For example, in our framework, if the inflation rate is super exogenous in a money demand function, we cannot have a constant inflation function by means of simply reversing that money demand function.

Meanwhile, as for a test for super exogeneity, Engle and Hendry check both weak exogeneity and parameter invariance as follows. If $(m_t, z_t)'$ follow jointly a linear normal distribution given past observations,

$$\begin{pmatrix} m_t \\ z_t \end{pmatrix} | X_{t-1} \sim N \left[\begin{pmatrix} \mu_t^M \\ \mu_t^Z \end{pmatrix}, \begin{pmatrix} \sigma_t^{MM} & \sigma_t^{MZ} \\ \sigma_t^{MZ} & \sigma_t^{ZZ} \end{pmatrix} \right]. \quad (8)$$

In this case, it can be shown that a conditional distribution (equation [7]) is expressed as

$$D_{M|Z}(m_t|z_t, X_{t-1}, \phi_1) \approx N[\beta_t z_t + X_{t-1}\gamma + (\delta_t - \beta_t)(z_t - \mu_t^Z), \omega_t], \quad (9)$$

where $\delta_t = \sigma_t^{MZ}/\sigma_t^{ZZ}$. Weak exogeneity with respect to β_t implies that $\delta_t = \beta_t$, since in this case, we can see that $D_{M|Z}(.)$ does not include any parameter of D_Z. Moreover, parameter invariance can be tested as follows. Using a Taylor Series approximation, $\beta_t z_t$ can be expressed as

$$\beta_t z_t \approx \beta_0 \mu_t^Z + \beta_1 (\mu_t^Z)^2 + \beta_2 \sigma_t^{ZZ} + \beta_3 \mu_t^Z \sigma_t^{ZZ}. \quad (10)$$

Parameter invariance means that $\beta_1 = \beta_2 = \beta_3 = 0$. In other words, the test corresponds to the test of the null coefficient of $z_t - \mu_t^Z$ and σ_t^{ZZ}, which are added to the original regression model of $m_t = \beta z_t + X_{t-1}\gamma$.[17]

17. Frequently, terms of higher order, $(\mu_t^Z)^2$ and $\mu_t^Z \sigma_t^{ZZ}$, are not added because they take quite small values and are insignificantly different from zero.

In practice, as proxies for the unobservable values $z_t - \mu_t^Z$ and σ_t^{ZZ}, Engle and Hendry used ordinary least square (OLS) estimated errors, $\hat{u}_t$, and a moving average of $\hat{u}_t^2$, $\hat{\sigma}_t$, which are derived from the marginal model of z_t,

$$z_t = \sum_{i=0}^{k} b_i z_{t-i} + u_t.$$

Following a procedure similar to Engle and Hendry, first, we try to model the univariate marginal processes of the inflation rate and the own interest rate.

$$\underset{}{\Delta p_t} = \underset{(0.10)}{0.27\Delta p_{t-2}} - \underset{(0.09)}{0.16\Delta p_{t-3}} - \underset{(0.08)}{0.21\Delta p_{t-4}} - \underset{(0.005)}{0.04 p_{t-1}} + \underset{(0.00078)}{0.003}, \tag{11}$$

T = 1975/I–1994/IV, $R^2 = 0.54$, $\hat{\sigma} = 0.45$ percent, $DW = 1.48$,
AR: $F(5, 70) = 2.28$, $ARCH$: $F(4, 67) = 0.55$, *Normality*: $\chi^2(2) = 2.42$,
X_i^2: $F(8, 66) = 0.91$, X_iX_j: $F(14, 60) = 0.75$, $RESET$ $F(1, 74) = 3.70$.

$$\begin{aligned}\Delta Rm_t = {} & \underset{(0.08)}{0.51\Delta Rm_{t-1}} + \underset{(0.10)}{0.22\Delta Rm_{t-3}} - \underset{(0.03)}{0.12 Rm_{t-1}} + \underset{(0.003)}{0.012\text{ID80/II}} \\ & \underset{(1.8\cdot 10^{-5})}{-3.9\cdot 10^{-5}\text{trend}} - \underset{(0.003)}{0.011},\end{aligned} \tag{12}$$

T = 1975/I–1994/IV, $R^2 = 0.54$, $\hat{\sigma} = 0.28$ percent, $DW = 1.88$,
AR: $F(5, 69) = 0.78$, $ARCH$: $F(4, 66) = 1.45$, *Normality*: $\chi^2(2) = 4.73$,
X_i^2: $F(9, 64) = 0.89$, X_iX_j: $F(15, 58) = 0.75$, $RESET$ $F(1, 73) = 0.008$.

From these equations, $\hat{u}(.)_t$ and $\hat{\sigma}(\cdot)_t$, which is calculated as the four-period moving average of $\hat{u}^2(\cdot)_t$, are obtained. A super exogeneity test is conducted by including these terms together. The obtained result is

$$\begin{aligned}\Delta(m-p)_t = {} & \underset{(0.06)}{-0.29(m-p-0.4y-0.4w)_{t-1}} + \underset{(0.16)}{0.65 Rm_{t-1}} - \underset{(0.16)}{0.71 Rr_{t-1}} \\ & + \underset{(0.0002)}{0.0009\text{trend}} + \underset{(0.17)}{0.39 M\Delta y_{t-1}} + \underset{(0.06)}{0.18\Delta w_t} - \underset{(0.19)}{0.21 E\Delta p_t} \\ & + \underset{(0.26)}{0.62\Delta Rm_t} + \underset{(0.09)}{0.49} - \underset{(0.003)}{0.005\hat{u}(\Delta p)_t} - \underset{(0.003)}{0.003\hat{\sigma}(\Delta p)_t} \\ & - \underset{(0.005)}{0.002\hat{u}(\Delta Rm)_t} + \underset{(0.011)}{0.001\hat{\sigma}(\Delta Rm)_t},\end{aligned} \tag{13}$$

T = 1975/IV–1994/IV, $R^2 = 0.79$, $\hat{\sigma} = 0.50$ percent, $DW = 2.07$,
AR: $F(5, 59) = 1.32$, $ARCH$: $F(4, 56) = 0.13$, *Normality*: $\chi^2(2) = 1.73$,
X_i^2: $F(24, 39) = 0.54$, $RESET$ $F(1, 63) = 6.03$*.

None of the additional terms is significant. In fact, the linear restriction tests for zero coefficients of the additional terms is $0.65 \sim F(6, 64)$, which is well below its 5 percent critical value. Furthermore, if each pair of $\hat{u}(.)_t$ and $\hat{\sigma}(\cdot)_t$ is added individually, the results are the same, i.e., none of the added terms is significant.

Additional terms	$\hat{u}(\Delta p)$ $\hat{\sigma}(\Delta p)$	$\hat{u}(\Delta Rm)$ $\hat{\sigma}(\Delta Rm)$
Test statistics	1.12	0.09

Test statistics correspond to the linear restriction tests for setting zero coefficients to the additional terms and distributes as $F(2, 66)$. From them, we may conclude that these terms are super exogenous in the money demand function.

V. Conclusion

Despite the "boom and bust" of the "bubble" economy, the paper shows that the underlying demand for money has been stable, if we properly take into account financial liberalization and the wealth effect. At this moment, Japan is experiencing another big wave of financial liberalization, the financial "Big Bang" reform, which is scheduled to be completed by 2001. The reform might require researchers to reconsider the above specification of money demand in the future. For example, the amendment of the Foreign Exchange Law, effective from April 1998, is likely to make the money demand more sensitive to movements in the exchange rate.[18] How to incorporate the impacts of the ongoing financial reform is an item on the agenda of future research.

Furthermore, the paper reveals that the function satisfies the super exogeneity condition with respect to the inflation rate and the money's own interest rate. The latter implies that money demand can be controlled by the central bank, if it can deliberately alter the own interest rate. This is because super exogeneity claims that such policy intervention has been effective in the sense that the parameter used for policy simulation is invariant (see Favero and Hendry [1992] for the issue related to the Lucas critique). On the other hand, the former implies that we cannot obtain an inflation function by means of a simple inversion of the money demand. Because of super exogeneity, this function loses parameter constancy. While money may have some information content that permits predicting future inflation rates (Sekine [1996]), its transmission mechanism is far more complex than a simple inversion of the money demand. In order to clarify the transmission mechanism of monetary policy, we have to look at an inflation function together with the demand for money and other structural equations. An extension of Hendry and Mizon (1993) of the "general-to-simple" approach in the context of a simultaneous equations model seems to be one of the more promising candidates for this.

18. A preliminary investigation has not confirmed any significant role of the exchange rate in the above money demand function. As Ishida and Shirakawa (1996) emphasized, this is probably because of the past regulation that prohibited Japanese residents from maintaining overseas deposits.

APPENDIX: INTERPOLATION OF WEALTH DATA

The annual SNA-based wealth data are interpolated into quarterly data. This appendix describes the procedure. All data used here are evaluated at current prices.

SNA wealth data can be decomposed into five components. These are

[1] inventory;

[2] net fixed assets; buildings for residence and nonresidence, machinery, etc.;

[3] non-reproducible tangible assets; lands, forest, etc.;

[4] financial assets other than stocks; currency, deposits, bonds, etc.; and

[5] stocks.

The interpolation is conducted individually into each component.

A. Inventory

Quarterly based changes in inventories (private and public sectors) are cumulated to the stock of inventory at the end of 1955. However, because of estimation error, the cumulated series does not coincide with the stock data, which are available at the end of each year. Then, the quarterly based stock of inventory is obtained by putting the cumulated series into x^* in the following equation:

$$x^{**}_{t+i} = \bar{x}_{t+i} + \frac{T_x}{T_{x^*}}(\bar{x}^*_{t+i} - x^*_{t+i})\,i, \tag{A.1}$$

where i = 1, 2, 3 and denotes the i-th quarter of each year. Over-bar denotes the linear interpolation and $T_x = (x_t - x_{t-4})/4$, $T_{x^*} = (x^*_t - x^*_{t-4})/4$. The equation comes from interpolation by related series in the textbook by Maddala (1977, pp. 205–207).

B. Net Fixed Assets

The related series—x^* in equation (A.1)—is obtained by cumulating quarterly net investment (private housing investment + private business investment + public investment – depreciation) to the stock of net fixed assets at the end of 1955. Then, similar to inventory, the quarterly based net fixed assets are derived from equation (A.1).

C. Non-Reproducible Tangible Assets

First, the quarterly land price index is calculated by linearly interpolating the ratio of (land price index)/(nominal GDP), where the land price index (urban district, all purpose, six major cities) comprises biannual data. Next, the annual SNA data are divided by the corresponding land price index and the result is linearly interpolated into the quarterly data. Then, the quarterly series of non-reproducible tangible assets is obtained by multiplying these two quarterly sets of data.

D. Financial Assets Other than Stocks

The FOF statistics provide information on quarterly basis financial assets other than stocks. However, because of some discrepancies in bond valuation, the series does not coincide with the corresponding SNA data. Here, the quarterly SNA-based financial assets other than stocks are calculated by linearly interpolating the ratio of (FOF-based series)/(SNA-based series).

E. Stocks

There are differences in treatment between SNA and FOF. These are

	SNA	FOF
Nonfinancial sector	Book values up to 1968, market values after 1969.	Book values up to 1982/IV, market values after 1983/I.
Financial sector	Book values up to 1968, market values after 1969.	Book values up to 1994/IV, market values after 1995/I.

Our estimation is conducted for sample periods after 1975, hence SNA-based stock assets are valued at market prices. First, quarterly stocks held by the nonfinancial sector are calculated by linearly interpolating the ratio of (FOF nonfinancial stock)/(SNA nonfinancial stock). Although there is a change in valuation of FOF-based stock, I cannot find an apparent shift in the ratio. For this reason, such a shift in valuation is neglected in the interpolation.

Next, for stocks held by the financial sector before 1995, linear interpolation is conducted to the ratio of (FOF financial stock) to (SNA financial stock), both on a book value basis. Then, the ratio of (market values in SNA)/(book values in SNA) is linearly interpolated. The market value of quarterly stocks held by the financial sector is obtained by multiplying the interpolated book value of stock and the interpolated ratio of (market values in SNA)/(book values in SNA).[19]

19. Since both statistics value the stock at market prices after 1995, quarterly stocks held by the financial sector can be calculated simply by linear interpolation of the ratio (FOF financial stock)/(SNA financial stock), both on a market price basis. However, because of substantial revisions of the preliminary data in SNA—only preliminary data were available when I started this study—I decided not to use the data after 1995/I and conducted all the estimation within periods prior to that quarter.

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