Japanese Demand for M1 and Demand Deposits: Cross-Sectional and Time-Series Evidence from Japan

Hiroshi Fujiki and Kiyoshi Watanabe

We investigate the relationship between money, short-term interest rates, and scale variables. We use three monetary aggregates: M1, demand deposits, and cash currency in circulation. Regional cross-sectional data yield stable estimates of the income elasticity of demand deposits that are positive and close to unity. We impose the estimated income elasticity obtained from cross-sectional data and estimate double-log interest rate elasticities of demand for M1 velocities and demand-deposit velocities using time-series data.

Keywords: Zero interest rate policy; Demand for money
JEL Classification: E41, E52

Hiroshi Fujiki: Director and Senior Economist, Institute for Monetary and Economic Studies, Bank of Japan (E-mail: hiroshi.fujiki@boj.or.jp)
Kiyoshi Watanabe: Economist, Institute for Monetary and Economic Studies, Bank of Japan (E-mail: kiyoshi.watanabe@boj.or.jp)

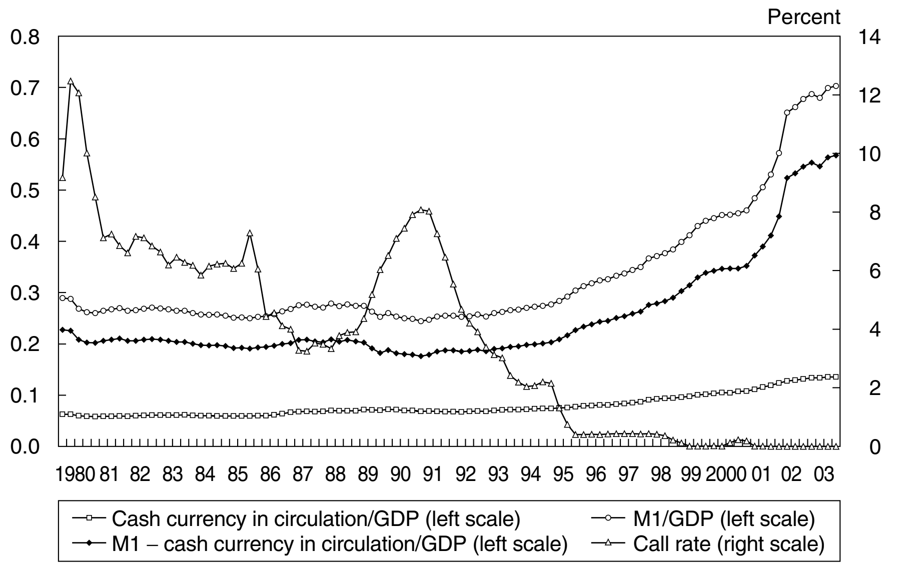
The authors thank Yuzo Honda, Kiyotaka Nakashima, Makoto Saito, Toshitaka Sekine, and the staff of the Bank of Japan for their helpful comments. We thank especially Ryuzo Miyao for his very useful comments and discussions that substantially improved the earlier version of this paper. The views expressed in this paper are solely those of the authors and do not necessarily represent the views of the Bank of Japan or the Institute for Monetary and Economic Studies.

I. Introduction

This paper examines the relationship between macroeconomic variables, such as nominal GDP, and M1, demand deposits, and cash currency in circulation. We focus on M1 for theoretical and empirical reasons. Theoretically, there is a testable implication of the so-called "liquidity trap." In this paper, the existence of a "liquidity trap" is implied by a highly nonlinear M1 demand function with respect to the short-term nominal interest rate when the nominal interest rate is close to zero, holding real income constant. Empirically, Figure 1 shows that the ratio of M1 to nominal GDP increased rapidly after 1995, and a substantial part of that increase was due to the increase in demand deposits, rather than cash currency in circulation. If the income elasticities of M1 and demand deposits are close to unity, the increase in the ratio of M1 to nominal GDP, especially after 1995, could be explained by changes in the nominal interest rate. Is demand for M1, demand deposits, and cash currency in circulation interest rate elastic according to Japanese data? To answer this question, this paper first presents a theoretical model developed by Fujiki and Mulligan (1996a) and then updates the empirical results presented in Fujiki (2002).

Two empirical studies have motivated this paper. First, Nakashima and Saito (2002) used Japanese money market data from January 1985 to March 2001 to examine whether nominal prices move with inertia when nominal interest rates are extremely low. They found that the demand for money had been extremely responsive to interest rates since the Bank of Japan had started to guide overnight

Figure 1 Money Stock Relative to Nominal GDP



Note: Seasonally adjusted, average outstanding amount.

call rates below 0.5 percent in 1995. They also found that nominal prices did not respond to changes in the nominal money supply under the low interest rate policy. Second, Miyao (2003) analyzed the presence and stability of a cointegrating relationship between the ratio of M1 to GDP and the call rate by using quarterly data from 1985/I to 2002/IV. He found evidence of a log-linear cointegrating relationship. This cointegrating relationship, between the log ratio of M1 to GDP and the log of the call rate, was stable (i.e., has no structural break in the interest rate elasticity) even after 1995, when nominal rates were virtually zero. Miyao (2003) argued that the stable double-log interest rate elasticity could be reconciled with the unstable interest rate semi-elasticity obtained by Nakashima and Saito (2002).

In the remainder of this paper, we update the results of Fujiki (2002) and check the robustness of the empirical evidence presented by Miyao (2003) and Nakashima and Saito (2002) in three ways. First, we use three monetary aggregates: M1, demand deposits, and cash in circulation. Since Fujiki (2002) used income elasticities obtained from cross-sectional data on demand deposits to estimate the interest rate elasticity using time-series analysis, our use of demand deposits is a natural extension of Miyao (2003). Second, we pay attention to the subsample properties of our estimates. Third, we use longer sample periods for estimation. While Miyao (2003) used data from 1985/I to 2002/IV, and Nakashima and Saito (2002) used data from January 1985 to March 2001, we use data from 1980/I to 2003/II.

This paper is organized as follows. In Section II, we briefly summarize the theoretical model developed by Fujiki and Mulligan (1996a) and specify empirical models. In Section III, we describe the data used for analysis. In Section IV, we report the main results of the empirical analysis based on cross-sectional and time-series data. In Section V, we report the results of robustness checks. Section VI summarizes the paper.

II. Theoretical and Empirical Models

This section explains our theoretical model and statistical models.

A. Theoretical Model

Fujiki and Mulligan (1996a) showed that a parametric model of production by households and firms leads to a conventional log-linear money demand function, in which money demand depends on real income, nominal interest rates, and the prices of productive inputs. They showed that both the income elasticity of money demand and the own opportunity cost elasticity of money demand are equal to the structural parameters of the household production functions and firm production functions. Moreover, both the income and own opportunity cost elasticities of money demand obtained from household and firm production functions are invariant to aggregation if the cost of transaction services relative to total household income is negligible. Therefore, the use of our cross-sectional estimates for prior information on time-series estimates is reasonable.

B. Cross-Sectional Statistical Model

We obtain a *cross-sectional estimator*, hereafter $\hat{\beta}k_{cs}(t)$, which is an ordinary least squares (OLS) estimator for equation (1) using annual data from 1990 to 2000:

$$\begin{aligned}\ln(\text{prefectural demand deposits})_{it} &= \alpha_{cs}(t) + \beta 1_{cs}(t)\ln(\text{prefectural GDP})_{it} \\ &\quad + \beta 2_{cs}(t)(\text{population density})_{it} + u_{it}, \\ &\quad i = 1, \ldots, 47,\ t \text{ given}, \end{aligned} \tag{1}$$

where the subscript i denotes the prefecture and the subscript t denotes the fiscal year t (t = 1990, . . . , 2000), and k = 1 and 2. Following Fujiki and Mulligan (1996b), we include population density (PD) to control for the level of financial technology in each prefecture. Standard errors are computed by using the method of White (1980). Data on prefectural demand deposits and prefectural GDP are per capita constant (fiscal 1990) values.

C. Time-Series Statistical Model

We obtain a *time-series estimator*, hereafter $\hat{\beta}k_{ts}$, which is an estimator for equation (2) by using quarterly time-series data from 1980/I to 2003/II:

$$\ln(Money)_t = \alpha_{ts} + \beta 1_{ts}\ln(\text{GDP})_t + \beta 2_{ts}\ln(\text{call rate})_t + u_t, \tag{2}$$

where the subscript t denotes the period (t = 1980/I, . . . , 2003/II), and k = 1 and 2. The dependent variable, money, is either M1, demand deposits, or cash currency in circulation deflated by the GDP deflator. We estimate equation (2) by using standard time-series techniques, OLS, fully modified OLS (FMOLS), and dynamic OLS (DOLS), imposing the income elasticity of M1 minus cash currency in circulation, $\hat{\beta}1_{cs}(t)$, obtained from the cross-sectional data described above. Miyao (2003) estimated equation (2) by imposing the restriction that $\beta 1_{ts}$ is unity. We checked the robustness of his assumption.

III. Data

This section describes the cross-sectional data and the time-series data.

A. Cross-Sectional Data

We use three types of annual data: data on prefectural GDP, data on prefectural demand deposits, and data to account for differences in financial technology between regions from 1990 to 2000. We only use data from 1990 onward to avoid discrepancies in statistics.

First, statistics on gross prefectural expenditure from the Economic and Social Research Institute, Cabinet Office of Japan, 2003, provide a suitable counterpart to annual data on national GDP. We have consistent estimates of gross prefectural expenditure and the gross prefectural expenditure deflator from fiscal 1990 to fiscal 2000. Prefectural GDP data are based on the 1993 System of National Accounts

(SNA), which are available only after 1990. The estimates of prefectural GDP differ significantly from those based on the 1968 SNA, which provides a consistent data series from 1975 to 1999. We use only data based on the 1993 SNA to avoid discrepancies in the statistics. Our data sample period is from 1990 to 2000.

Second, data on prefectural demand deposits held by individuals and firms at domestically licensed banks by prefecture (end of month outstanding) are available from "Financial and Economic Statistics Monthly" of the Bank of Japan (hereafter, MF1 data). Since national M1 statistics are defined as the sum of cash currency in circulation and total demand deposits net of deposits held by financial institutions, MF1 represents the prefectural counterpart of national M1 minus cash currency in circulation.[1] MF1 data do not include demand deposits at community banks, Norinchukin Bank, and Shoko Chukin Bank, which are included in the M1 statistics. However, as Table 1 shows, MF1 data explain about 70 percent of M1 from 1992 to 2001, and about 80 percent from 1990 to 1991, at least from 1992 to 2000. Therefore, given a careful choice of sample periods, MF1 accounts for a reasonably constant proportion of M1. MF1 data refer to the end of the month outstanding, and we use fiscal-year averages (e.g., 1991 data represent the average from April 1991 to March 1992).

Third, following Fujiki and Mulligan (1996b), to account for differences in financial technology between regions, we use PD. These data are based on the population of each prefecture for the beginning of October of each year.

MF1 and GDP figures constructed in these ways are deflated by the gross prefectural expenditure deflator and divided by the population in each prefecture to obtain per capita real money balances and real gross prefectural expenditure. Figure 2 shows that there is a stable positive correlation between the logs of real MF1 per capita and the log of regional GDP per capita during the sample period.

Table 1 Ratio of M1 and MF1

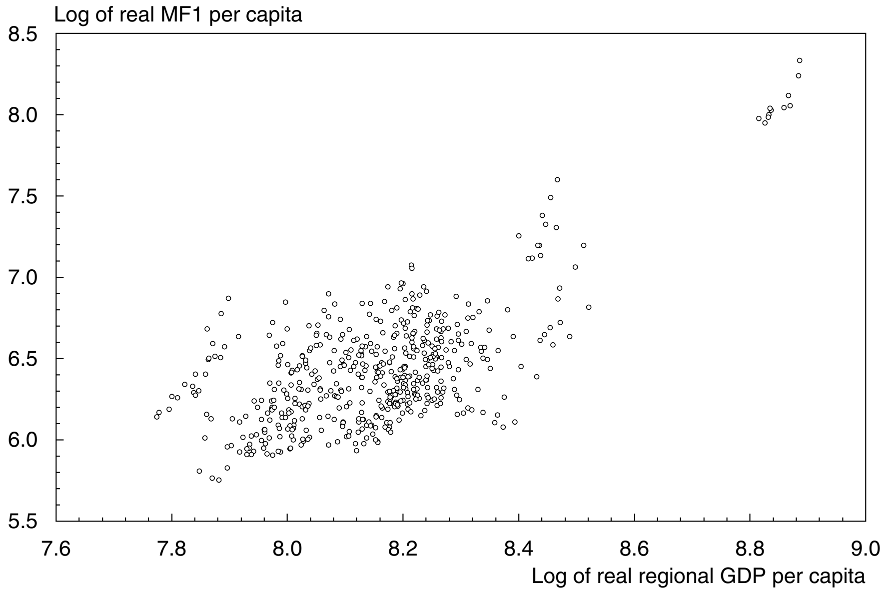
Fiscal year	MF1	M1AVG	MF1/M1
1990	908,493	1,119,869	0.8112
1991	921,532	1,192,225	0.7729
1992	900,270	1,229,769	0.7320
1993	913,550	1,275,002	0.7165
1994	944,268	1,344,552	0.7022
1995	1,045,545	1,489,961	0.7017
1996	1,169,644	1,672,461	0.6993
1997	1,269,304	1,818,555	0.6979
1998	1,333,857	1,959,787	0.6806
1999	1,491,488	2,191,495	0.6805
2000	1,587,963	2,332,027	0.6809
2001	1,813,519	2,618,135	0.6926

Note: Units are ¥100 millions. MF1 stands for average MF1 for the fiscal year, and M1AVG shows average outstanding amount of M1.

Source: Bank of Japan.

1. MF1 data do not include cash currency in circulation, because regional data on the amount of currency held by individuals are not available.

Figure 2 Regional Money Stock and Regional GDP



B. Time-Series Data

We use GDP (seasonally adjusted, 1993 SNA base),[2] the GDP deflator (seasonally adjusted, 1993 SNA base), Indices of Industrial Production (hereafter IIP),[3] M1, demand deposits, and cash currency in circulation (average outstanding, seasonally adjusted).[4]

The use of M1 as a dependent variable follows Nakashima and Saito (2002) and Miyao (2003). Nakashima and Saito (2002) used cash currency in circulation in addition to M1, but did not use demand deposits.

The use of GDP as a scale variable follows Miyao (2003). We also tried IIP as a scale variable following Nakashima and Saito (2002), but were unable to replicate their results, mainly because the base year for the IIP has been revised from 1995 to 2000. Thus, we mainly report results based on GDP.

We use the uncollateralized overnight call rate (hereafter, the call rate) to represent the own opportunity cost of M1, demand deposits, and cash currency in circulation. The use of the call rate follows Nakashima and Saito (2002) and Miyao (2003).[5]

2. Seasonally adjusted SNA statistics are available from the Cabinet Office of Japan's website at http://www.esri.cao.go.jp/en/sna/menu.html.
3. We used seasonally adjusted IIP data, which are released on the website of the Ministry of Economy, Trade and Industry (METI) at http://www.meti.go.jp/english/statistics/index.html.
4. The Bank of Japan publishes these monetary data on its website at http://www.boj.or.jp/stat/stat_f.htm.
5. U.S. studies usually use three-month treasury bills or three-month commercial paper for the opportunity costs of M1 (see Serletis [2001, p. 97], or Hayashi [2000, p. 660], for example). Data on Japanese treasury bills (six-month) are available only after 1992, and the Japanese financing bill (three-month) is sold at market price only after 1999.

IV. Main Results

In this section, first we report the estimates of the income elasticity of money demand from cross-sectional data. Having obtained a plausible estimate of the income elasticity of money demand from cross-sectional regressions, we use these estimates to estimate the interest rate elasticity of money demand from time-series data.

A. Results of Cross-Sectional Estimation

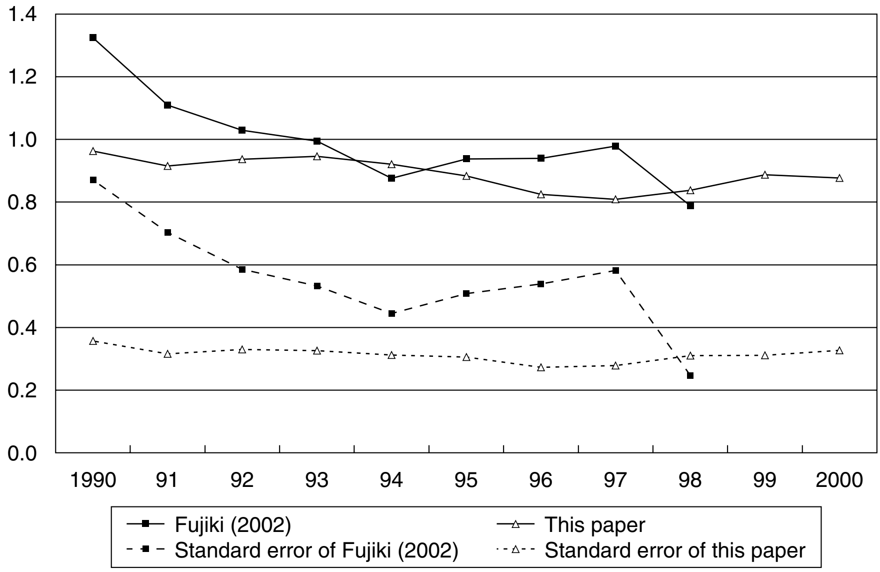
The second, fourth, and sixth columns of Table 2 show the estimates of the constant, α_{cs}, the income elasticity of demand deposits, $\beta 1_{cs}$, and the PD elasticity of demand deposits, $\beta 2_{cs}$, from equation (2). All estimates have the expected signs, and their standard errors are sufficiently small for the parameters to be significantly different from zero. The cross-sectional estimates of the income elasticity of demand deposits are positive and take reasonably stable values close to unity. Figure 3 reports estimates

Table 2 Cross-Sectional and Panel Estimates

Sample	α_{cs}	Standard error	$\beta 1_{cs}$	Standard error	$\beta 2_{cs}$	Standard error
1990	–3.681	2.417	0.963	0.303	1.157	0.250
1991	–3.245	2.453	0.915	0.300	1.116	0.239
1992	–3.195	2.472	0.937	0.304	0.993	0.205
1993	–3.128	2.552	0.946	0.310	0.925	0.191
1994	–2.812	2.536	0.921	0.305	0.889	0.176
1995	–2.465	2.423	0.883	0.289	0.925	0.171
1996	–1.939	2.345	0.824	0.276	0.955	0.165
1997	–1.789	2.251	0.809	0.265	0.991	0.146
1998	–1.956	2.236	0.837	0.264	0.986	0.133
1999	–2.480	2.447	0.887	0.288	1.092	0.142
2000	–2.405	2.349	0.877	0.275	1.128	0.141
Pooling model with time dummies			0.838	0.104	0.182	0.016
Pooling model with time dummies and random effects	–2.049	0.314	0.914	0.041	0.176	0.008
Pooling model with time dummies and region dummies			0.555	0.104	–0.375	0.270
Pooling model with time dummies			1.327	0.104		
Pooling model with time dummies and random effects	–4.445	0.465	1.333	0.057		
Pooling model with time dummies and region dummies			0.631	0.088		

Note: The estimation method is OLS. Dependent variable is deflated per capita ln(demand deposits). Standard errors are computed by using the method of White (1980). The estimations include a constant term as the set of explanatory variables. The Hausman test statistic for comparing the pooling model with time dummies and the random-effects model with PD and the pooling model with time dummies and region dummies with PD is 14.20, with two degrees of freedom, which supports the pooling model with time dummies and region dummies with a p-value of 0.0008. The Hausman test statistic for comparing the pooling model with time dummies and random effects and the pooling model with time dummies and region dummies without PD is 82.86, with one degree of freedom, which supports the pooling model with time dummies and region dummies with a p-value of 0.000.

Figure 3 Cross-Sectional Income Elasticity of Demand Deposits



Note: Dashed lines show the lower bound of the estimates, defined as the parameter estimates minus twice their standard errors.

for the income elasticity for demand deposits that are consistent with those obtained by Fujiki (2002) from historical data on regional GDP (using the level of PD, rather than its log).

The bottom six rows of Table 2 show the estimates based on a static panel model. The results are consistent with those of Fujiki and Mulligan (1996b). The pooling model with time dummies and region dummies yields a significantly lower income elasticity of demand deposits, and the parameter estimates of PD are incorrectly signed. However, the estimates from the pooling model with time dummies and pooling model with time dummies and random effects are again around 0.8 to 0.9. Without PD, the estimates from the pooling model with time dummies and pooling model with time dummies and random effects are around 1.3. This suggests that the omission of PD introduces an upward bias to the income elasticity of demand deposits. The pooling model with time dummies and region dummies yields a lower income elasticity of demand deposits of 0.63. However, using PD, time dummies and region dummies for instrumental variables for regional GDP, two-stage least squares estimation of the income elasticity of demand deposits, and conditional on time dummies and region dummies, yields an income elasticity of demand deposits of 0.83 (standard error = 0.17). Thus, we may conclude that the pooling estimates should be around 0.8 to 0.9 on average.

B. Application to Time-Series Data

We use data from 1980 to 2003 and new cross-sectional estimates for the income elasticity of money demand. More specifically, we define the velocity of M1 with income elasticities of unity, 0.838 and 0.915, as M1V1, M1V2, and M1V3, respectively, based on the pooling estimates in Table 2. We define velocity for cash in circulation with an income elasticity of unity as CAV1. We define the velocity of demand deposits with income elasticities of unity, 0.838 and 0.915, as DDV1, DDV2, and DDV3, respectively. Since our cross-sectional income elasticities are measured with respect to demand deposits, the use of DDV1, DDV2, and DDV3 is a natural extension of Miyao (2003). Miyao (2003) considered only counterparts of M1V1, M1V2, and M1V3 based on data from 1985/I to 2002/IV.

We begin by applying the augmented Dickey-Fuller test (ADF test) (Dickey and Fuller [1979]) and the Phillips-Perron test (PP test) (Phillips and Perron [1988]) to the logs of these velocities and the log of the call rate, because the structural model of Fujiki and Mulligan (1996a) suggests double-log specification. We used three specifications for each test: first, neither a constant term nor trend was included; second, only a constant term was included; and third, both a constant term and trend were included. We apply those tests for the level of the call rate following Miyao (2003).[6] The results of the ADF tests and PP tests are summarized in Table 3. All except the level of the call rate have unit roots in level, and are integrated of order 1 because their differenced series are stationary.

We apply the Engle-Granger cointegration test (the ADF *t*-test based on the OLS residuals). We begin by specifying the maximum length of the augmented autoregression term used for the ADF test, and then choose the optimal length of the augmented autoregression term based on the Akaike Information Criterion (AIC). We use two criteria to determine the maximum length of the augmented autoregression ($p(\max)$ hereafter). The first criterion, $p(\max)_1$, is due to Schwert (1989), as suggested by Hayashi (2000):

$$p(\max)_1 = \left[12\left(\frac{T}{100}\right)^{1/4}\right]\left(\text{integer part of } 12\left(\frac{T}{100}\right)^{1/4}\right),$$

where T is the size of sample.
The second criterion, $p(\max)_2$, is used following Said and Dickey (1984):

$$p(\max)_2 = [(T)^{1/3}].$$

6. It might not be appropriate to use standard unit root tests for the level of the call rate if one takes seriously the problem of a zero bound for nominal interest rates. A zero bound for nominal interest rates might lead to a nonuniform variance for the level of the interest rate.

Table 3 Unit Root Tests

[1] ADF Test

	No constant				Constant added				Constant and trend			
	AR parameter	Test statistics	Lag order		AR parameter	Test statistics	Lag order		AR parameter	Test statistics	Lag order	
			AIC	BIC			AIC	BIC			AIC	BIC
M1V1	0.997	–1.636	(3)	(1)	1.023	2.295	(1)	(1)	0.999	–0.053	(1)	(1)
M1V2	1.005	2.586	(1)	(1)	1.022	2.415	(1)	(1)	0.998	–0.128	(1)	(1)
M1V3	1.025	2.714	(1)	(1)	1.022	2.366	(1)	(1)	0.999	–0.092	(1)	(1)
CAV1	0.998	–2.462	(3)	(1)	1.018	2.624	(1)	(1)	0.989	–0.684	(1)	(1)
DDV1	0.997	–1.658	(2)	(1)	1.022	1.992	(1)	(1)	0.999	–0.070	(1)	(1)
DDV2	1.007	2.434	(1)	(1)	1.021	2.144	(1)	(1)	0.998	–0.133	(1)	(1)
DDV3	1.003	0.314	(1)	(1)	1.022	2.078	(1)	(1)	0.999	–0.103	(1)	(1)
Call rate	0.963	–4.010**	(1)	(1)	0.950	–3.483**	(1)	(1)	0.917	–2.527	(1)	(4)
ln(call rate)	1.029	1.380	(5)	(6)	1.076	3.914	(5)	(5)	1.050	1.482	(5)	(5)
ΔM1V1	0.692	–3.092**	(2)	(1)	0.554	–4.527**	(1)	(1)	0.316	–5.444**	(1)	(1)
ΔM1V2	0.708	–3.039**	(2)	(1)	0.557	–4.573**	(1)	(1)	0.330	–5.445**	(1)	(1)
ΔM1V3	0.700	–3.064**	(2)	(1)	0.555	–4.550**	(1)	(1)	0.324	–5.443**	(1)	(1)
ΔCAV1	0.745	–2.660**	(2)	(2)	0.567	–3.753**	(2)	(1)	0.350	–5.180**	(1)	(1)
ΔDDV1	0.618	–4.064**	(1)	(1)	0.547	–4.558**	(1)	(1)	0.319	–5.447**	(1)	(1)
ΔDDV2	0.633	–4.003**	(1)	(1)	0.548	–4.608**	(1)	(1)	0.328	–5.465**	(1)	(1)
ΔDDV3	0.626	–4.032**	(1)	(1)	0.548	–4.583**	(1)	(1)	0.324	–5.455**	(1)	(1)
Δ (call rate)	0.507	–4.642**	(1)	(2)	0.467	–4.758**	(1)	(2)	0.460	–4.707**	(1)	(2)
Δln(call rate)	0.413	–2.751**	(5)	(6)	0.028	–4.529**	(4)	(5)	–0.477	–8.344**	(4)	(4)

(Continued on next page)

Table 3 (continued)

[2] PP Test

	No constant			Constant added			Constant and trend		
	AR parameter	Z_α	Z_t	AR parameter	Z_α	Z_t	AR parameter	Z_α	Z_t
M1V1	0.995	–0.561	–1.386	1.043	3.858	4.030	1.009	0.815	0.669
M1V2	1.010	0.850	2.844	1.041	3.731	4.663	1.009	0.774	0.609
M1V3	1.048	4.271	4.373	1.042	3.806	4.381	1.009	0.796	0.640
CAV1	0.997	–0.279	–2.388	1.029	2.537	3.325	0.996	–1.375	–0.683
DDV1	0.995	–0.498	–1.247	1.044	3.873	3.490	1.010	0.918	0.778
DDV2	1.013	1.152	2.786	1.043	3.870	4.164	1.010	0.901	0.742
DDV3	1.018	0.282	0.151	1.044	3.890	3.852	1.010	0.910	0.761
Call rate	0.974	–3.459	–1.863	0.974	–5.029	–1.723	0.918	–22.411*	–3.357
ln(call rate)	1.021	1.511	0.767	1.025	2.945	2.358	0.978	–0.873	–0.334
ΔM1V1	0.578	–54.490**	–5.666**	0.515	–54.568**	–5.841**	0.306	–56.733**	–6.630**
ΔM1V2	0.606	–50.379**	–5.422**	0.530	–51.258**	–5.665**	0.331	–54.270**	–6.425**
ΔM1V3	0.593	–52.291**	–5.537**	0.524	–52.770**	–5.746**	0.320	–55.406**	–6.520**
ΔCAV1	0.581	–61.234**	–5.938**	0.448	–65.804**	–6.488**	0.278	–73.377**	–7.325**
ΔDDV1	0.565	–54.242**	–5.685**	0.515	–53.446**	–5.794**	0.317	–53.126**	–6.474**
ΔDDV2	0.587	–51.210**	–5.502**	0.526	–50.793**	–5.655**	0.336	–50.982**	–6.306**
ΔDDV3	0.577	–52.617**	–5.587**	0.521	–52.007**	–5.719**	0.328	–51.969**	–6.384**
Δ (call rate)	0.324	–79.678**	–8.480**	0.295	–78.975**	–8.813**	0.288	–82.219**	–8.986**
Δln(call rate)	0.340	–52.631**	–6.511**	0.305	–45.366**	–6.710**	0.276	–37.112**	–7.323**

Note: The length of lags selected for the ADF tests are chosen according to AIC, from up to 11 lags. * shows that the null hypotheses are rejected at the 5 percent level, and ** shows that the null hypotheses are rejected at the 1 percent level.

The upper panel of Table 4 shows the results of the Engle-Granger cointegration test for the double-log specification. Based on the Schwert criterion, $p(\max)_1$, we set $p(\max)_1 = 11$. Given $p(\max)_1$, we report the optimal lag length suggested by the AIC in the column headed P1. We cannot reject the null of a unit root in the residuals for all cases, as the ADF statistics show. The evidence is consistent with Miyao (2003), who could not reject the null of a unit root in the residuals based on M1V1 data from 1975/I to 2002/IV, with five lags based on double-log specification. Based on the Said and Dickey criterion, we first set $p(\max)_2 = 4$. We report the optimal lag length suggested by the AIC in the column headed P2. We find evidence for cointegration for M1V2 and M1V3 at the 1 percent level, for M1V1 and DDV2 at the 5 percent level, and for DDV3 at the 10 percent level.

The lower panel of Table 4 shows the results of the Engle-Granger cointegration test for the semi-log specification. For the semi-log specification, we cannot reject the

Table 4 Cointegration Test

Dependent variable	Constant	ln(call rate)	ADF statistics	P1	ADF statistics	P2
M1V1	–1.161 (0.007)	–0.108 (0.003)	–2.600	(5)	–3.635**	(3)
M1V2	1.317 (0.007)	–0.116 (0.003)	–2.880	(5)	–6.116***	(2)
M1V3	0.139 (0.007)	–0.112 (0.003)	–2.806	(5)	–5.698***	(2)
CAV1	–2.579 (0.007)	–0.090 (0.003)	–1.705	(5)	–2.897	(4)
DDV1	–1.441 (0.009)	–0.113 (0.003)	–2.293	(5)	–2.898	(3)
DDV2	1.038 (0.008)	–0.121 (0.003)	–2.673	(5)	–3.806**	(3)
DDV3	–0.140 (0.008)	–0.117 (0.003)	–2.469	(5)	–3.339*	(3)

Dependent variable	Constant	Call rate	ADF statistics	P1	ADF statistics	P2
M1V1	–0.941 (0.032)	–0.061 (0.006)	–0.150	(3)	–0.150	(3)
M1V2	1.566 (0.032)	–0.069 (0.006)	–0.272	(3)	–1.541	(1)
M1V3	0.374 (0.032)	–0.065 (0.006)	–0.211	(3)	–1.469	(1)
CAV1	–2.365 (0.023)	–0.059 (0.005)	–1.105	(5)	–2.115	(1)
DDV1	–1.219 (0.035)	–0.062 (0.007)	–0.082	(3)	–1.258	(1)
DDV2	1.289 (0.035)	–0.070 (0.007)	–0.163	(3)	–1.398	(1)
DDV3	0.097 (0.035)	–0.066 (0.007)	–0.121	(3)	–1.332	(1)

Note: The estimation method is OLS. The numbers in the column headed P1 report the optimal lag length chosen by the AIC under the Schwert criterion. The numbers in the column headed P2 report the optimal lag length chosen by the AIC under the Said and Dickey criterion. * shows that the null hypotheses are rejected at the 10 percent level, ** shows that the null hypotheses are rejected at the 5 percent level, and *** shows that the null hypotheses are rejected at the 1 percent level. Sample period is from 1980/I to 2003/II.

null hypothesis of a unit root (no cointegration) for all cases. This evidence is again consistent with that of Miyao (2003), who could not reject the null of a unit root in the residuals based on M1V1 data from 1975/I to 2002/IV, with one lag based on semi-log specification.

Table 5 shows the results of testing the null of no cointegration against a structural break with regime shift following Gregory and Hansen (1996) using three test statistics. The null hypothesis of no cointegration between seven measures of velocity and the log of the call rate is rejected, with a break point around the late 1990s. However, we cannot reject the null hypotheses of no cointegration between seven measures of velocity and the level of the call rate. The results from the two tests summarized in Tables 4 and 5 suggest that only M1V1, M1V2, M1V3, DDV2, and DDV3 from the double-log specification have a stable relationship, with or without a break.

Table 6 reports the estimates of the parameters of equation (2) given the income elasticity based on FMOLS and DOLS assuming stable cointegration. Surprisingly, the double-log interest rate elasticities are similar at around –0.1 to –0.15. The results are consistent with those of Miyao (2003), who reports a double-log interest rate elasticity of –0.131, based on a DOLS regression of M1V1 on the log of the call rate using a sample from 1985/I to 2002/IV.[7]

Table 5 Cointegration Tests with Regime Shift (Gregory and Hansen [1996])

ln(call rate)									
	M1V1	M1V2	M1V3	CAV1	DDV1	DDV2	DDV3	Critical value (5 percent)	Critical value (10 percent)
Inf-ADF	–6.100** (1995/III)	–6.091** (1998/IV)	–6.436** (1995/III)	–5.497** (1987/II)	–4.695* (1997/IV)	–5.469** (1997/IV)	–5.113** (1997/IV)	–4.95	–4.68
Inf-Z_t	–4.877* (1996/I)	–4.806* (1996/IV)	–4.947* (1996/IV)	–4.376 (1987/I)	–4.795* (1989/IV)	–4.833* (1996/I)	–4.812* (1995/III)	–4.95	–4.68
Inf-Z_α	–39.825* (1996/I)	–39.130* (1996/I)	–41.143** (1996/IV)	–34.102 (1987/I)	–38.998* (1989/IV)	–39.093* (1995/III)	–39.170* (1995/III)	–40.48	–36.19
Call rate									
	M1V1	M1V2	M1V3	CAV1	DDV1	DDV2	DDV3	Critical value (5 percent)	Critical value (10 percent)
Inf-ADF	–4.058 (1999/III)	–4.395 (1999/IV)	–4.361 (1999/IV)	–3.610 (1998/III)	–3.877 (1999/III)	–4.069 (1999/III)	–3.988 (1999/III)	–4.95	–4.68
Inf-Z_t	–3.540 (1999/IV)	–3.545 (1999/IV)	–3.553 (1999/IV)	–3.027 (1999/IV)	–3.445 (1999/IV)	–3.520 (1999/IV)	–3.491 (1999/IV)	–4.95	–4.68
Inf-Z_α	–22.527 (1999/IV)	–22.942 (1999/IV)	–22.860 (1999/IV)	–17.421 (1999/IV)	–21.237 (1999/IV)	–22.417 (1999/IV)	–21.932 (1999/IV)	–40.48	–36.19

Note: The three test statistics are the smallest value (the largest negative value) of ADF test statistics and Z_t and Z_α statistics in Phillips (1987) across all possible break points in the data sample. Those statistics test the null of no cointegration against the alternative of cointegration in the presence of possible regime shift. See the definition and details of those statistics in Gregory and Hansen (1996), especially p. 106. The points in the parentheses designate when the structural break occurs. Sample period is from 1980/I to 2003/II. The estimation method is based on the procedure proposed by Gregory and Hansen (1996). We used a GAUSS code programmed by Professor Bruce Hansen. * shows that the null hypotheses are rejected at the 5 percent level, and ** shows that the null hypotheses are rejected at the 1 percent level.

7. The estimation in this section could benefit from a consideration of the effects of generated regressors (see McKenzie and McAleer [1997]). For example, standard OLS estimates from a regression of M1V1 on a constant and the call rate (which is stationary and exogenous) yield biased standard errors.

Table 6 Estimates of Time-Series Models: Full Sample

Methods	Dependent variable	Constant	ln(call rate)
FMOLS	M1V1	–1.163 (0.011)	–0.112 (0.005)
	M1V2	1.317 (0.009)	–0.122 (0.004)
	M1V3	0.138 (0.010)	–0.117 (0.004)
	DDV1	–1.442 (0.018)	–0.117 (0.007)
	DDV2	1.037 (0.012)	–0.127 (0.005)
	DDV3	–0.141 (0.014)	–0.122 (0.006)
DOLS (2)	M1V1	–1.157 (0.077)	–0.117 (0.034)
	M1V2	1.324 (0.036)	–0.126 (0.016)
	M1V3	0.145 (0.049)	–0.121 (0.022)
	DDV1	–1.438 (0.137)	–0.122 (0.061)
	DDV2	1.044 (0.064)	–0.131 (0.029)
	DDV3	–0.136 (0.095)	–0.127 (0.043)
DOLS (4)	M1V1	–1.155 (0.198)	–0.119 (0.094)
	M1V2	1.330 (0.132)	–0.129 (0.063)
	M1V3	0.149 (0.151)	–0.124 (0.072)
	DDV1	–1.440 (0.287)	–0.122 (0.137)
	DDV2	1.045 (0.176)	–0.133 (0.084)
	DDV3	–0.136 (0.227)	–0.128 (0.108)
DOLS (8)	M1V1	–1.116 (0.182)	–0.138 (0.086)
	M1V2	1.368 (0.079)	–0.148 (0.038)
	M1V3	0.187 (0.075)	–0.143 (0.035)
	DDV1	–1.404 (0.284)	–0.140 (0.135)
	DDV2	1.080 (0.189)	–0.151 (0.089)
	DDV3	–0.101 (0.234)	–0.146 (0.111)

Note: This table shows the results of estimation based on FMOLS proposed by Phillips and Hansen (1990) and DOLS proposed by Stock and Watson (1993). In employing FMOLS, we used the FM procedure in the GAUSS Coint package version 2. We adopted the pre-whitened spectral quadratic kernel as in Nakashima and Saito (2002). "DOLS (n)" refers to results based on DOLS with n leads and n lags (n = 2, 4, and 8). The numbers in parentheses under estimated parameters denote standard errors. Sample period is from 1980/I to 2003/II.

C. Subsample Properties

To examine the assumption of cointegration throughout the sample period, we apply the test proposed by Hansen (1992) to check the stability of the parameters obtained from FMOLS in Table 7.[8] Table 7 shows that both the Sup-F statistic (which tests the null of constancy against parameter change of unknown timing) and the Mean-F statistic (which tests the null of stability against the alternative of parameter instability following a random walk) reject the null hypothesis of parameter stability at the 5 percent significance level, except in the case of demand deposits. However, the LC statistics (which test the null of cointegration against the alternative of no cointegration) support cointegration in relation to M1V1, M1V2, M1V3, DDV1, DDV2, and DDV3, but not CAV1. Thus, the results for parameter stability in relation to M1V1, M1V2, M1V3, and DDV3 are unclear.

Given the results of the tests based on Gregory and Hansen (1996) shown in Table 5, we further divide the sample at 1995/II for M1V1, M1V2, DDV2, and DDV3 following the results reported by Miyao (2003). The upper panel of Table 8 shows the results based on the sample from 1980/I to 1995/II, and the lower panel shows the results based on the sample from 1995/III to 2003/II. Table 8 shows that the double-log interest rate elasticities are larger in the latter sample. Table 8 also shows that there is cointegration between velocities and the log of the call rate. This result is inconsistent with the finding of Miyao (2003), who finds no stable relationship with a break.

To examine the source of inconsistency with the results of Miyao (2003), we start the initial sample period from 1985/I, instead of 1980. The upper panel of Table 9 reproduces the results of Miyao (2003). The results show that there is no cointegration between velocities and the log of the call rate. The lower panel of Table 9 shows the robustness of the results reported by Miyao (2003). The double-log interest rate elasticity for M1V1 changes by no more than 0.03 in absolute value between the two subsamples. We conclude that the source of inconsistency between our results in Table 8 and those of Miyao (2003) is our inclusion of data from 1980 to 1984.

Table 7 Stability of FMOLS Parameters

	LC	Mean-F	Sup-F
M1V1	0.434	7.757***	15.331**
M1V2	0.181	15.051***	28.120***
M1V3	0.263	11.137***	22.031***
CAV1	0.582**	18.003***	26.769***
DDV1	0.534*	3.790*	7.200
DDV2	0.442	8.980***	18.270***
DDV3	0.528*	5.993**	12.329**

Note: The estimation method is based on the procedure proposed by Hansen (1992). We used a GAUSS code programmed by Professor Bruce Hansen. For each test, we adopted the pre-whitened spectral quadratic kernel. * shows that the null hypotheses are rejected at the 10 percent level, ** shows that the null hypotheses are rejected at the 5 percent level, and *** shows that the null hypotheses are rejected at the 1 percent level. Sample period is from 1980/I to 2003/II.

8. Note that the values of parameter estimates are not sensitive to the choice of estimation methods, and we restrict our attention to FMOLS.

Table 8 Structural Break and FMOLS

1980/I–1995/II	Dependent variable	Constant	ln(call rate)	ADF statistics	P1	ADF statistics	P2
Standard OLS	M1V1	–1.289	–0.032	–3.911**	(1)	–3.911**	(1)
		(0.017)	(0.010)				
	M1V2	1.233	–0.068	–4.355***	(1)	–4.355***	(1)
		(0.015)	(0.009)				
	DDV2	0.901	–0.041	–4.132***	(1)	–4.132***	(1)
		(0.018)	(0.011)				
	DDV3	–0.298	–0.023	–3.439**	(1)	–3.439**	(1)
		(0.021)	(0.012)				
FMOLS	M1V1	–1.286	–0.035	Automatic bandwidth selected			
		(0.033)	(0.019)	9.000			
	M2V2	1.255	–0.082				
		(0.026)	(0.015)	7.179			
	DDV2	0.901	–0.042				
		(0.034)	(0.020)	8.288			
	DDV3	–0.320	–0.011				
		(0.046)	(0.027)	11.682			

1995/III–2003/II	Dependent variable	Constant	ln(call rate)	ADF statistics	P1	ADF statistics	P2
Standard OLS	M1V1	–1.156	–0.108	–3.581**	(4)	–4.087***	(2)
		(0.026)	(0.007)				
	M1V2	1.349	–0.109	–3.582**	(4)	–4.116***	(2)
		(0.027)	(0.008)				
	DDV2	1.054	–0.119	–3.664**	(4)	–4.230***	(2)
		(0.029)	(0.008)				
	DDV3	–0.137	–0.119	–3.667**	(4)	–4.218***	(2)
		(0.029)	(0.008)				
FMOLS	M1V1	–1.182	–0.118	Automatic bandwidth selected			
		(0.035)	(0.010)	3.722			
	M2V2	1.321	–0.120				
		(0.036)	(0.010)	3.825			
	DDV2	1.023	–0.131				
		(0.038)	(0.011)	3.834			
	DDV3	–0.167	–0.130				
		(0.038)	(0.011)	3.788			

Note: Based on the results of the structural break test reported in Table 5, we divided the sample into two subsamples before and after the break point. The sample periods of the former are from 1980/I to 1995/II. Those of the latter are from 1995/III to 2003/II. In each panel, the upper part denotes the results based on standard OLS and the lower part corresponds to the results of estimation based on FMOLS. The numbers in the column headed P1 report the optimal lag length chosen by the AIC under the Schwert criterion. The numbers in the column headed P2 report the optimal lag length chosen by the AIC under the Said and Dickey criterion. As for FMOLS, we adopted the non-pre-whitened Bartlett quadratic kernel, and the selection of automatic bandwidth is based on Andrews (1991). The numbers in the parentheses denote standard errors. * shows that the null hypotheses are rejected at the 10 percent level, ** shows that the null hypotheses are rejected at the 5 percent level, and *** shows that the null hypotheses are rejected at the 1 percent level. The points in the parentheses designate when the F statistics are highest.

Table 9 Structural Break and FMOLS (2)

1985/I–1995/II	Dependent variable	Constant	ln(call rate)	ADF statistics	P1	ADF statistics	P2
Standard OLS	M1V1	–1.239	–0.073	–2.054	(1)	–2.054	(1)
		(0.014)	(0.009)				
	M1V2	1.257	–0.085	–2.459	(1)	–2.459	(1)
		(0.015)	(0.010)				
	DDV2	0.955	–0.084	–2.374	(1)	–2.374	(1)
		(0.014)	(0.009)				
	DDV3	–0.231	–0.079	–1.979	(1)	–1.979	(1)
		(0.016)	(0.011)				
FMOLS	M1V1	–1.223	–0.084	Automatic bandwidth selected			
		(0.020)	(0.013)	9.729			
	M1V2	1.280	–0.100				
		(0.024)	(0.015)	8.725			
	DDV2	0.968	–0.094				
		(0.023)	(0.015)	8.102			
	DDV3	–0.082	–0.227				
		(0.019)	(0.030)	11.005			

1995/III–2003/II	Dependent variable	Constant	ln(call rate)	ADF statistics	P1	ADF statistics	P2
Standard OLS	M1V1	–1.156	–0.108	–3.581**	(4)	–4.087***	(2)
		(0.026)	(0.007)				
	M1V2	1.349	–0.109	–3.582**	(4)	–4.116***	(2)
		(0.027)	(0.008)				
	DDV2	1.054	–0.119	–3.664**	(4)	–4.230***	(2)
		(0.029)	(0.008)				
	DDV3	–0.137	–0.119	–3.667**	(4)	–4.218***	(2)
		(0.029)	(0.008)				
FMOLS	M1V1	–1.182	–0.118	Automatic bandwidth selected			
		(0.035)	(0.010)	3.722			
	M1V2	1.321	–0.120				
		(0.036)	(0.010)	3.825			
	DDV2	1.023	–0.131				
		(0.038)	(0.011)	3.834			
	DDV3	–0.167	–0.130				
		(0.038)	(0.011)	3.788			

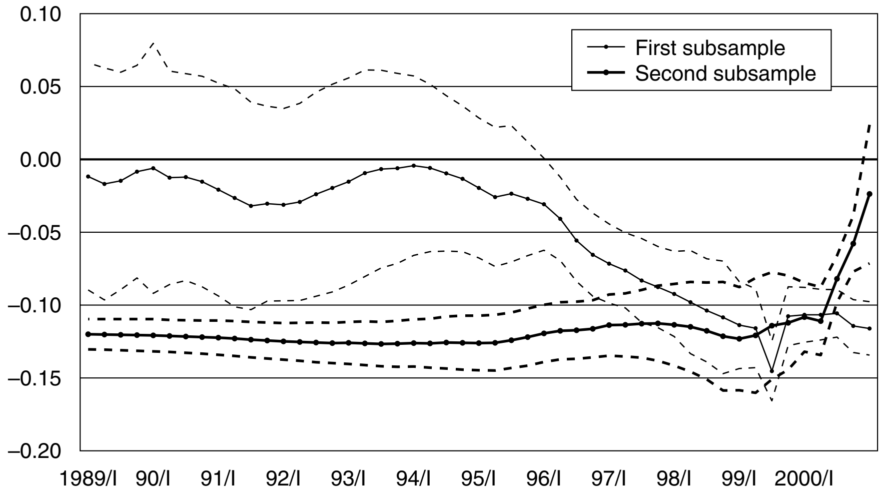
Note: We truncated the sample and made a subsample from 1985/I to 2003/II. We set 1995/II as the break point, following Miyao (2003). Estimation is based on standard OLS and FMOLS. The numbers in the column headed P1 report the optimal lag length chosen by the AIC under the Schwert criterion. The numbers in the column headed P2 report the optimal lag length chosen by the AIC under the Said and Dickey criterion. For FMOLS, we adopted the non-pre-whitened Bartlett quadratic kernel, automatic bandwidth selection as before.

To check the robustness of our estimated income and interest rate elasticities of money demand, we divide our sample period into two and compute the interest rate elasticities of M1V1, M1V2, and DDV2 for each sub-period.

The thin solid line in Figure 4 shows our estimates of the interest rate elasticity of M1V1 based on the data before the break point (the first sub-period), and the thick line shows the double-log interest rate elasticity of M1V1 based on the data after the break point (the second sub-period). The dotted lines show the upper and lower bounds for the interest rate elasticities, which are the estimated coefficients plus and minus two standard errors, respectively. The estimates are based on FMOLS. The horizontal axis in Figure 4 corresponds to the break points of the sample. The break points are from 1986/I to 2000/IV. (We only report the results from 1989/I to 2000/I in the figure.) Figure 4 shows the double-log interest rate elasticities estimated from the two subsamples. The second period provides reasonable interest rate elasticities of around –0.1. The first period yields statistically insignificant interest rate elasticities if we restrict the end of the sample period to before 1995. The results suggest that the interest rate elasticities from the double-log specification are stable, especially if the sample includes only recent data. This finding is consistent with the results of Miyao (2003).

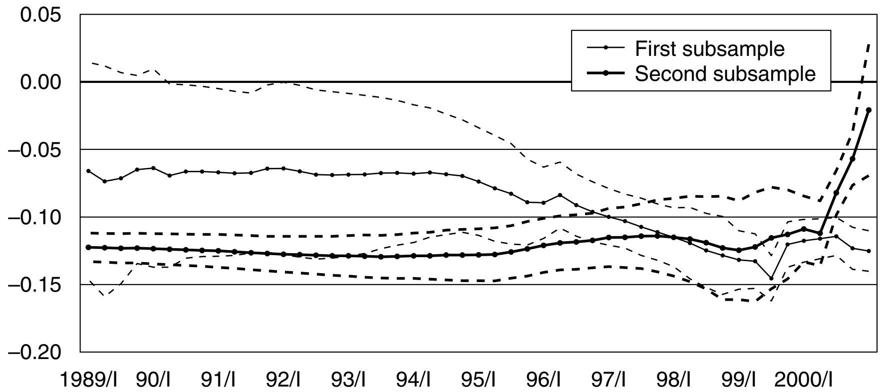
The double-log interest rate elasticities of M1V2 estimated from the two subsamples are shown in Figure 5, which is constructed in the same way as Figure 4. The second period provides reasonable interest rate elasticities of around –0.1, as in

Figure 4 Double-Log Interest Rate Elasticity of M1V1



Note: The horizontal axis corresponds to the break points of the sample. The thin solid line (first subsample) shows estimates obtained from the data before the break point, and the thick solid line (second subsample) shows those obtained from data after the break point. The dotted lines denote the upper and lower bounds of the estimated values, which are constructed by adding and subtracting two standard errors to and from the estimated coefficients. The vertical axis measures the estimated values. Estimation is by FMOLS.

Figure 5 Double-Log Interest Rate Elasticity of M1V2



Note: The horizontal axis corresponds to the break points of the sample. The thin solid line (first subsample) shows estimates obtained from the data before the break point, and the thick solid line (second subsample) shows those obtained from data after the break point. The dotted lines denote the upper and lower bounds of the estimated values, which are constructed by adding and subtracting two standard errors to and from the estimated coefficients. The vertical axis measures the estimated values. Estimation is by FMOLS.

the case of M1V1. The first period yields statistically significant negative values if we include observations after 1992. The figure suggests that small changes in the absolute value of the income elasticities do not affect the estimates of the interest rate elasticity in the second sub-period, which is again consistent with the findings of Miyao (2003).

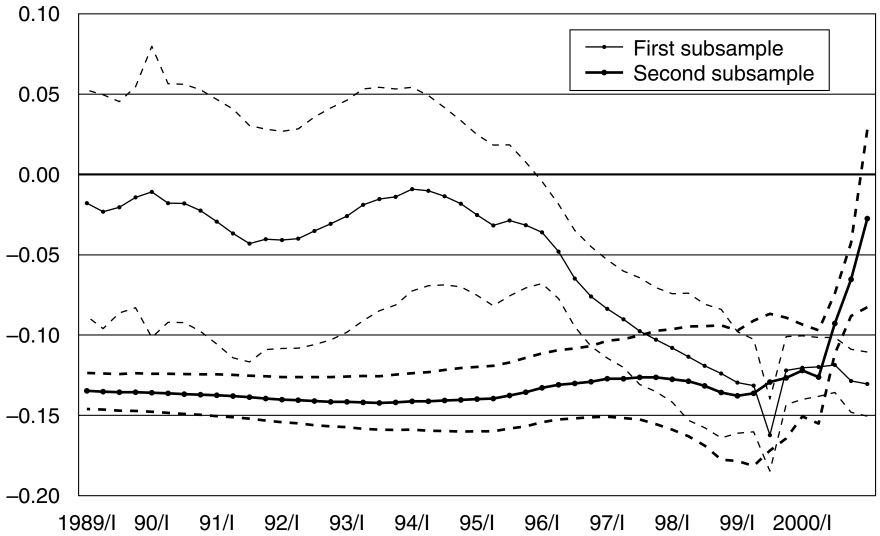
Figure 6 shows the interest rate elasticities of DDV2 estimated from the two subsamples. The second period provides reasonable interest rate elasticities of around –0.1. The first period yields statistically insignificant interest rate elasticities if we restrict the end of the sample period to before 1995. The results are quite similar to the results based on M1V1.

D. Reservations

The double-log interest rate elasticity seems to have a structural break in 1995 or 1998 based on the tests reported in Table 5. Might this be related to changes in the statistical properties of the nominal interest rate, due, for instance, to a shift to a low interest rate period? In particular, it might be better to treat the log of the call rate as a unit root variable with a shift in the slope of the trend. Hence, tests proposed by Perron (1997) were applied to the log of the call rate.

Following Soejima (1995), we set the maximum lag length for the test at 12. Table 10 shows that we cannot reject the null hypothesis that the log of the call rate has a unit root with a shift in the trend, and the structural break might occur in 1999 or 1990. The break in 1999 might be related to the introduction of the zero interest rate policy. However, the results are not consistent with the analysis in this section,

Figure 6 Double-Log Interest Rate Elasticity of DDV2



Note: The horizontal axis corresponds to the break points of the sample. The thin solid line (first subsample) shows estimates obtained from the data before the break point, and the thick solid line (second subsample) shows those obtained from data after the break point. The dotted lines denote the upper and lower bounds of the estimated values, which are constructed by adding and subtracting two standard errors to and from the estimated coefficients. The vertical axis measures the estimated values. Estimation is by FMOLS.

Table 10 Results of Test Based on Perron (1997)

Sample period: 1983/III–2002/III Sample size = 77	Break point	k(max) = 12 (k = the optimal number of lags)							Critical value (5 percent)	
		k	μ (t-value)	θ (t-value)	β (t-value)	δ (t-value)	γ (t-value)	α (t-value)	$t^*(\alpha)$ k(t-sig)	$t^*(\alpha, \gamma)$ k(t-sig)
ln(interest rate)	1990/I	12	0.013	0.485	0.001	–0.263	–0.045	0.919		
			(0.061)	(–2.117)	(0.147)	(–0.772)	(–3.345)	(–0.944)	–5.59	
	1999/IV	12	–0.113	0.852	–0.005	–1.377	–0.115	1.090		
			(–0.576)	(3.253)	(–1.451)	(–3.249)	(–3.563)	(1.208)		–4.98

Note: We tested the null hypothesis that $\alpha = 1$ in the following model:

$$y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t,$$

where y is the data we are interested in testing (in our case, log of the call rate), μ is a constant term, DU_t is a dummy variable that takes one if $t < T_b$ and zero otherwise, t is a linear time trend, T_b denotes the point when the structural break occurs, DT_t is t if $t > T_b$ and 0 if $t \leq T_b$, $D(T_b)_t$ is a dummy variable 1 if $t = T_b + 1$ and zero otherwise, and e_t is an error term. The parameters μ, θ, β, δ, γ, α, and c_i are estimated by the regression. The panel shows our results. In each column, the upper numbers denote the estimated values and the numbers in parentheses are t-values of each parameter except for c_i. In this paper, we employed two test statistics. One is the minimum t-value of α, which corresponds to the value shaded in the upper panel. The other is the minimum t-value of γ, which corresponds to the value shaded in the lower panel. For details, see Perron (1997).

which suggest a break in 1995 or 1998. Given this evidence of a unit root with a break, the results reported in this section should be treated with caution.

V. Robustness Checks

A. Results Based on Quarterly GDP, Double-Log Functional Form

Why should we have to impose the income elasticity of M1 minus cash currency in circulation, $\hat{\beta}1_{cs}(t)$, obtained from the cross-sectional data to estimate equation (2)? This section reports the results of estimating a standard double-log demand for money equation (2) without imposing cross-sectional estimates.

For that purpose, we first apply the ADF test and PP test to the logs of the time series for real M1, real cash currency in circulation (hereafter real cash), real demand deposits, real GDP, the log of the call rate, and the level of the call rate. For the sake of subsequent analysis, IIP is also tested. Table 11 shows that we can only reject the null hypothesis of a unit root for the level of the call rate. For the first differences of the variables, we reject the null hypothesis of a unit root.

We tested for cointegration between the following: (1) log real M1, log real GDP, and the log call rate; (2) log real cash, log real GDP, and the log call rate; and (3) log real demand deposits, log real GDP, and the log call rate from 1980/I to 2003/II. We used two statistical methods to test for cointegration between these variables.

The first panel of Table 12 shows the results of the Engle-Granger residual-based tests for cointegration using the ADF *t*-test. Based on the Schwert criterion, we cannot reject the null of a unit root in the residuals for all cases. Based on the Said and Dickey criterion, we reject the null hypothesis of the existence of a unit root in the residuals. There is cointegration for M1, cash, and demand deposits. Second, the lower panel of Table 12 shows the results of three tests of the null hypothesis of no cointegration against the alternative hypothesis of a structural break with a regime shift following Gregory and Hansen (1996). The results show that the null hypotheses of no cointegration are rejected for M1 and demand deposits. We cannot reject the null of no cointegration for cash.

The two tests in Table 12 suggest that the cointegration results are not necessarily stable for cash. On balance, M1 and demand deposits seem to have a stable cointegrating relationship with GDP and the call rate, at most before 1996. This evidence is supported by the fact that the Bank of Japan did not set the call rate below 1 percent before 1995, and is consistent with the interpretation of Nakashima and Saito (2002). The instability of the cointegrating relationship after 1995 suggests that the use of cross-sectional income elasticities as prior information for the analysis of money demand makes sense, especially for the period of the low interest rate policy.

B. Results Based on Quarterly GDP and the Semi-Log Form

The evidence in Section V.A is not consistent with Nakashima and Saito (2002), who found that, after the break point of June 1995, the income elasticity is small and statistically insignificant and the semi-interest rate elasticity is large. To check this, we examine the results based on the level of the call rate, rather than the log of the call rate.

Table 11 Unit Root Tests

[1] ADF Test

	No constant				Constant added				Constant and trend			
	AR parameter	Test statistics	Lag order		AR parameter	Test statistics	Lag order		AR parameter	Test statistics	Lag order	
			AIC	BIC			AIC	BIC			AIC	BIC
Real M1	1.001	3.121	(1)	(1)	1.014	2.431	(1)	(1)	0.991	–0.557	(1)	(1)
Real cash	1.001	3.805	(2)	(2)	1.003	0.851	(2)	(1)	0.970	–1.540	(3)	(2)
Real demand deposits	1.001	2.848	(1)	(1)	1.016	2.363	(1)	(1)	0.992	–0.483	(1)	(1)
Real GDP	1.000	2.757	(3)	(1)	0.988	–2.115	(3)	(1)	0.984	–0.841	(3)	(1)
IIP	1.000	1.228	(3)	(3)	0.976	–2.201	(3)	(3)	0.975	–1.328	(3)	(3)
Call rate	0.963	–4.010**	(1)	(1)	0.950	–3.483**	(1)	(1)	0.917	–2.527	(4)	(1)
ln(call rate)	1.029	1.380	(6)	(5)	1.076	3.914	(5)	(5)	1.050	1.482	(5)	(5)
Δ (real M1)	0.737	–3.385**	(1)	(1)	0.533	–4.948**	(1)	(1)	0.363	–5.583**	(1)	(1)
Δ (real cash)	0.904	–1.546	(2)	(2)	0.618	–3.708**	(2)	(1)	0.576	–3.680*	(2)	(1)
Δ (real demand deposits)	0.697	–3.673**	(1)	(1)	0.521	–4.973**	(1)	(1)	0.345	–5.657**	(1)	(1)
Δ (real GDP)	0.709	–2.425*	(2)	(2)	0.354	–3.761**	(2)	(2)	0.178	–4.286**	(2)	(1)
Δ IIP	0.513	–4.847**	(2)	(2)	0.481	–5.020**	(2)	(2)	0.431	–5.358**	(2)	(2)
Δ (call rate)	0.507	–4.642**	(2)	(1)	0.467	–4.758**	(2)	(1)	0.460	–4.707**	(2)	(1)
Δln(call rate)	0.413	–2.751**	(6)	(5)	0.028	–4.529**	(5)	(4)	–0.477	–8.344**	(4)	(4)

(Continued on next page)

Table 11 (continued)

[2] PP Test

	No constant			Constant added			Constant and trend		
	AR parameter	Z_α	Z_t	AR parameter	Z_α	Z_t	AR parameter	Z_α	Z_t
Real M1	1.001	0.104	3.750	1.026	2.436	4.865	1.006	0.094	0.053
Real cash	1.001	0.108	5.550	1.012	0.971	1.773	0.977	–7.488	–1.973
Real demand deposits	1.001	0.108	3.317	1.030	2.800	5.388	1.008	0.540	0.346
Real GDP	1.000	0.038	4.359	0.987	–1.246	–2.293	0.993	–1.018	–0.531
IIP	1.000	0.024	1.208	0.974	–3.099	–1.836	0.987	–2.742	–1.031
Call rate	0.974	–3.459	–1.863	0.974	–5.029	–1.723	0.918	–22.411*	–3.357
ln(call rate)	1.021	1.511	0.767	1.025	2.945	2.358	0.978	–0.873	–0.334
Δ (real M1)	0.704	–37.100**	–4.567**	0.549	–43.910**	–5.317**	0.402	–48.464**	–5.919**
Δ (real cash)	0.802	–22.734**	–3.490**	0.529	–51.741**	–5.738**	0.457	–64.109**	–6.377**
Δ (real demand deposits)	0.668	–41.130**	–4.847**	0.540	–44.108**	–5.346**	0.392	–45.322**	–5.858**
Δ (real GDP)	0.320	–135.654**	–8.955**	0.014	–140.380**	–10.152**	–0.064	–113.432**	–10.245**
Δ IIP	0.561	–36.019**	–4.879**	0.544	–33.791**	–4.833**	0.523	–27.561**	–4.722**
Δ (call rate)	0.324	–79.678**	–8.480**	0.295	–78.975**	–8.813**	0.288	–82.219**	–8.986**
Δln(call rate)	0.340	–52.631**	–6.511**	0.305	–45.366**	–6.710**	0.276	–37.112**	–7.323**

Note: The length of lags used for ADF tests are chosen according to the AIC, from up to 11 lags. * shows that the null hypotheses are rejected at the 5 percent level, and ** shows that the null hypotheses are rejected at the 1 percent level. Sample period is from 1980/I to 2003/II.

Table 12 Cointegration Test

[1] Engle-Granger Test

Dependent variable	Constant	ln(GDP)	ln(call rate)	ADF statistics		ADF statistics	
					P1		P2
M1	1.060 (0.734)	0.855 (0.048)	–0.115 (0.004)	–2.879	(5)	–6.050***	(2)
Cash	–8.258 (0.567)	1.371 (0.037)	–0.071 (0.003)	–2.580	(5)	–4.694***	(1)
Demand deposits	3.313 (0.806)	0.689 (0.053)	–0.129 (0.004)	–2.999	(5)	–6.395***	(2)

Note: The estimation method is OLS. The numbers in the column headed P1 report the optimal lag length chosen by the AIC under the Schwert criterion. The numbers in the column headed P2 report the optimal lag length chosen by the AIC under the Said and Dickey criterion. *** shows that the null hypotheses are rejected at the 1 percent level. Sample period is 1980/I to 2003/II.

[2] Cointegration Tests with Regime Shift (Gregory and Hansen [1996])

	M1	Cash	Demand deposits	Critical value (5 percent)	Critical value (10 percent)
Inf-ADF	–6.638** (1996/III)	–4.879 (1997/I)	–6.947** (1996/III)	–5.50	–5.23
Inf-Z_t	–6.349** (1996/IV)	–4.839 (1996/IV)	–6.512** (1996/IV)	–5.50	–5.23
Inf-Z_α	–54.191** (1996/IV)	–35.435 (1996/IV)	–56.371** (1996/IV)	–58.33	–52.85

Note: The three test statistics are the smallest value (the largest negative value) of ADF test statistics and Z_t and Z_α statistics in Phillips (1987) across all possible break points in the data sample. Those statistics test the null of no cointegration against the alternative of cointegration in the presence of possible regime shift. See the definition and details of those statistics in Gregory and Hansen (1996), especially p. 106. The points in the parentheses designate when the structural break occurs. Sample period is from 1980/I to 2003/II. The estimation method is based on the procedure proposed by Gregory and Hansen (1996). We used a GAUSS code programmed by Professor Bruce Hansen. ** shows that the null hypotheses are rejected at the 1 percent level.

We begin by testing for cointegration between the following: (1) log real M1, log real GDP, and the call rate; (2) log real cash, log real GDP, and the call rate; and (3) log real demand deposits, log real GDP, and the call rate. We used two statistical methods to test for cointegration between these variables. Panel 1 of Table 13 shows that the ADF *t*-tests do not support cointegration for any case. Panel 2 of Table 13 shows that we cannot reject the null hypothesis of no cointegration against a structural break with regime shift for all cases. These results are consistent with the results reported in Table 4, which show that the choice between the semi-log specification and the double-log specification matters for the evaluation of a stable cointegrating relationship.

Table 13 Cointegration Test

[1] Engle-Granger Test

Dependent variable	Constant	ln(GDP)	Call rate	ADF statistics		ADF statistics	
					P1		P2
M1	–2.375 (2.965)	1.093 (0.192)	–0.057 (0.011)	–0.096	(3)	–0.096	(3)
Cash	–10.937 (1.922)	1.554 (0.124)	–0.033 (0.007)	–1.393	(1)	–1.393	(1)
Demand deposits	–0.226 (3.292)	0.936 (0.213)	–0.065 (0.012)	–0.110	(3)	–1.314	(1)

Note: The estimation method is OLS. The numbers in the column headed P1 report the optimal lag length chosen by the AIC under the Schwert criterion. The numbers in the column headed P2 report the optimal lag length chosen by the AIC under the Said and Dickey criterion. In this table, there is no evidence for cointegration relation.

[2] Cointegration Tests with Regime Shift (Gregory and Hansen [1996])

	M1	Cash	Demand deposits	Critical value (5 percent)	Critical value (10 percent)
Inf-ADF	–4.360 (1997/II)	–4.322 (1996/IV)	–4.374 (1997/II)	–5.50	–5.23
Inf-Z_t	–4.164 (1996/IV)	–4.285 (1996/IV)	–4.119 (1996/III)	–5.50	–5.23
Inf-Z_α	–29.057 (1996/III)	–29.820 (1996/II)	–28.678 (1996/III)	–58.33	–52.85

Note: The three test statistics are the smallest value (the largest negative value) of ADF test statistics and Z_t and Z_α statistics in Phillips (1987) across all possible break points in the data sample. Those statistics test the null of no cointegration against the alternative of cointegration in the presence of possible regime shift. See the definition and details of those statistics in Gregory and Hansen (1996), especially p. 106. The points in the parentheses designate when the structural break occurs. Sample period is from 1980/I to 2003/II. The estimation method is based on the procedure proposed by Gregory and Hansen (1996). We used a GAUSS code programmed by Professor Bruce Hansen.

C. Results Based on Quarterly IIP and the Semi-Log Form

Our analysis in Section V.B used GDP and the level of the call rate, while Nakashima and Saito (2002) used IIP and the level of the call rate. Are the results sensitive to whether GDP or IIP is chosen as the scale variable?[9]

To answer this question, we begin by testing for cointegration between the following variables: (1) log real M1, log IIP, and the call rate; (2) log real cash, log IIP, and the call rate; and (3) log real demand deposits, log IIP, and the call rate. Note that we use 2000 as the base year for IIP, while Nakashima and Saito (2002) used 1995. The results summarized in this section are best interpreted as robustness checks, because we do not use the same data sets as did Nakashima and Saito (2002). Given these limitations, the results reported in Table 14 differ from those reported in Nakashima and Saito (2002). The ADF *t*-tests in the upper panel of Table 14 indicate an absence of cointegration. Panel 2 shows that we cannot reject the null hypothesis of no cointegration against a structural break with regime shift.

9. Our analysis in this section uses quarterly data on IIP and the level of the call rate, while Nakashima and Saito (2002) used monthly data on IIP and the level of the call rate from January 1985 to March 2001. The results reported in this section are qualitatively similar even if we use monthly data.

Table 14 Cointegration Test

[1] Engle-Granger Test

Dependent variable	Constant	ln(IIP)	Call rate	ADF statistics	P1	ADF statistics	P2
M1	7.219 (2.937)	0.556 (0.223)	–0.095 (0.009)	–0.497	(3)	–1.794	(1)
Cash	–2.200 (2.277)	1.164 (0.173)	–0.077 (0.007)	–2.189	(1)	–2.189	(1)
Demand deposits	9.572 (3.147)	0.356 (0.239)	–0.101 (0.010)	–0.510	(3)	–1.770	(1)

Note: The estimation method is OLS. The numbers in the column headed P1 report the optimal lag length chosen by the AIC under the Schwert criterion. The numbers in the column headed P2 report the optimal lag length chosen by the AIC under the Said and Dickey criterion. In this table, there is no evidence for cointegration relation.

[2] Cointegration Tests with Regime Shift (Gregory and Hansen [1996])

	M1	Cash	Demand deposits	Critical value (5 percent)	Critical value (10 percent)
Inf-ADF	–4.393 (1998/IV)	–4.376 (1993/III)	–4.424 (1993/IV)	–5.50	–5.23
Inf-Z_t	–3.236 (1999/I)	–3.361 (1998/IV)	–3.232 (1999/I)	–5.50	–5.23
Inf-Z_α	–21.126 (1999/I)	–22.413 (1998/IV)	–21.118 (1999/I)	–58.33	–52.85

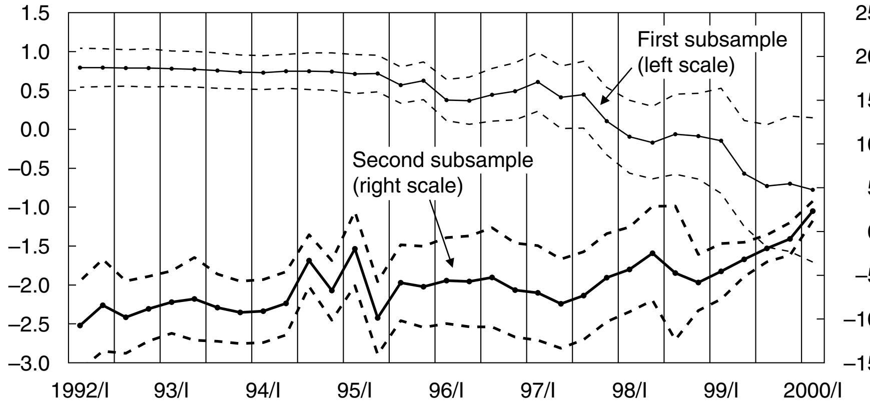
Note: The three test statistics are the smallest value (the largest negative value) of ADF test statistics and Z_t and Z_α statistics in Phillips (1987) across all possible break points in the data sample. Those statistics test the null of no cointegration against the alternative of cointegration in the presence of possible regime shift. See the definition and details of those statistics in Gregory and Hansen (1996), especially p. 106. The points in the parentheses designate when the structural break occurs. Sample period is from 1980/I to 2003/II. The estimation method is based on the procedure proposed by Gregory and Hansen (1996). We used a GAUSS code programmed by Professor Bruce Hansen.

The evidence against cointegration presented above prevents us from proceeding with the cointegrating regressions. However, for demand deposits, the value of Johansen's maximal eigenvalue test of no cointegration against the alternative of one cointegrating relationship is 18.97 (with two lags), which is statistically significant at the 10 percent level.[10]

Since Nakashima and Saito (2002) insist that a structural break occurs around 1995, we examine the robustness of our stability test by dividing our sample into two sub-periods and compute the income and interest rate elasticities of demand deposit for each for the sake of comparison. Figures 7 and 8 show the income and semi-interest rate elasticities based on the two subsamples. Figure 7 shows that the income elasticities from the second period are negative. Figure 8 shows two large decreases

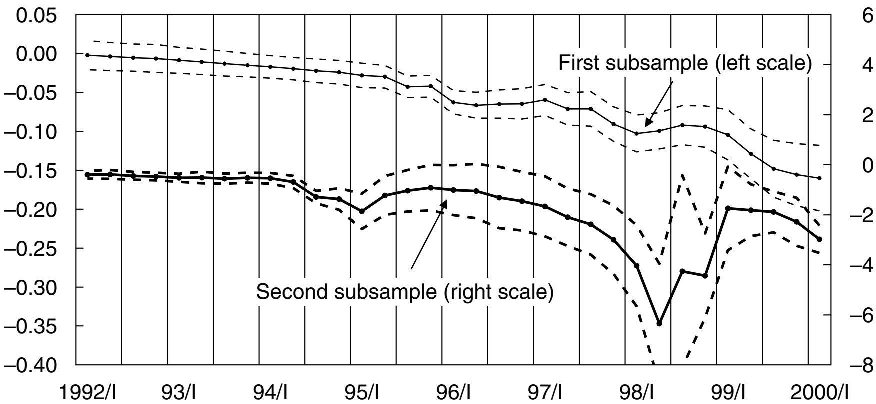
10. The estimation is done by E-Views, version 4.0. We choose optimum lag length 2 based on the Schwartz information criterion applied for unrestricted vector autoregression (VAR). We set maximum lag length 12. Critical value for the test is 18.60, based on Osterwald-Lenum (1992). The evidence for cointegration might not be robust, since we did not adjust the critical value for finite sample. Thus, the following discussion is for the sake of comparison.

Figure 7 Income Elasticity of Demand Deposits



Note: The horizontal axis corresponds to the break points of the sample. The thin solid line (first subsample) shows estimates obtained from the data before the break point, and the thick solid line (second subsample) shows those obtained from data after the break point. The dotted lines denote the upper and lower bounds of the estimated values, which are constructed by adding and subtracting two standard errors to and from the estimated coefficients. The left vertical axis measures the estimated values of the first sample. The second subsample's estimated values are on the right vertical axis. Estimation is by FMOLS.

Figure 8 Semi-Interest Rate Elasticity of Demand Deposits



Note: The horizontal axis corresponds to the break points of the sample. The thin solid line (first subsample) shows estimates obtained from the data before the break point, and the thick solid line (second subsample) shows those obtained from data after the break point. The dotted lines denote the upper and lower bounds of the estimated values, which are constructed by adding and subtracting two standard errors to and from the estimated coefficients. The left vertical axis measures the estimated values of the first sample. The second subsample's estimated values are on the right vertical axis. Estimation is by FMOLS.

in the semi-interest rate elasticity in 1995 and 1998 based on the estimates from the second period. This evidence is partly consistent with the findings of Nakashima and Saito (2002), which suggest that a structural break occurred around 1995. However, the structural break in the interest rate elasticities seems to occur in 1998. Overall, our results are similar to those obtained by Nakashima and Saito (2002) when we use IIP and the semi-log specification. However, the choice of sample period matters for identifying the break point.

VI. Summary

Imposing the relatively stable cross-sectional estimate of the income elasticity yields stable double-log interest rate elasticities for M1 and demand deposits, but not for cash. If the level of financial technology is adequately represented by population density and the nominal interest rate is constant in a cross-section of regions, stable cross-sectional estimates are to be expected, even from a short estimation period. Thus, one might benefit from prior information on the size of the income elasticity obtained from cross-sectional estimation.

Regarding the instability of the semi-log interest rate elasticity and the stability of the double-log interest rate elasticity, Miyao (2003) argues that two interest rate elasticities are not inconsistent, if one compares them using the identity (double-log interest rate elasticity)/(nominal interest rates) = (semi-log interest rate elasticity). We agree with his assertion. Note that he drew this conclusion having estimated the demand for money using both functional forms. Our results suggest that if we use the semi-log form, we might conclude that there is no stable relationship between the logs of the velocity measures and the call rate. Such a conclusion suggests that we should not use information about the demand for money for policy analysis, especially when a low interest rate policy applies. For example, in the zero interest rate environment, monetary assets included in M1 and other short-term assets might be perfect substitutes, in which case money demand might be indeterminate, so there is no equilibrium or long-run money demand relationship. Given the evidence of no cointegration from the semi-log functional form, information about the demand for money when interest rates are close to zero may not be useful. However, the stable double-log functional form suggests that one may well forecast velocity using the double-log form. Thus, analysis of the demand for money is useful.

The semi-log functional form is standard in the empirical literature on the demand for money. To test some classes of structural model, it is better to use the semi-log form. For example, Nakashima and Saito's (2002) objective was to highlight nonlinear changes in the shape of the M1 demand function. Although their statistical tests implied structural breaks, their objective was to identify jumps in the interest rate elasticities and income elasticities, whereas our focus is on cointegrating relationships. For their purpose, the choice of the semi-log form makes sense. This paper does not suggest that use of the semi-log functional form is inappropriate. Rather, it suggests that the choice of functional form is important for policy analysis.[11] Note also that the results of using the double-log form might change drastically when one allows for the possibility of a structural break.

If the inconsistency between our results and those reported by Miyao (2003) is simply due to our inclusion of observations from 1980 to 1984, we return to the issue raised by Ball (2001). He points out that even when using the same method, expanding the sample size to 1996, rather than to 1987, yields significantly smaller estimates in absolute value of the income and interest rate elasticities from U.S. money demand functions for M1. This point is particularly relevant for Japanese time-series estimates. We need to update the estimates of the money demand function, and pay attention to the sample period to be analyzed. The results show that policy recommendations based on the demand for M1 must be made cautiously.

11. Note that the so-called liquidity trap equilibrium could emerge as an equilibrium phenomenon before the adoption of a zero interest rate policy. Benhabib, Schmitt-Grohe, and Uribe (2002) demonstrate that a stable liquidity-trap equilibrium may emerge before the economy reaches the zero interest rate bound. Their liquidity traps may emerge even when an equilibrium money demand relation is well defined. Thus, our evidence of stable double-log interest rate elasticity does not necessarily mean that Japan is not in liquidity-trap equilibrium.

References

Andrews, Donald W. K., "Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimation," *Econometrica*, 59, 1991, pp. 817–858.

Ball, Laurence, "Another Look at Long-Run Money Demand," *Journal of Monetary Economics*, 47, 2001, pp. 31–44.

Benhabib, Jess, Stephanie Schmitt-Grohe, and Martin Uribe, "Avoiding Liquidity Traps," *Journal of Political Economy*, 110, 2002, pp. 535–563.

Dickey, David A., and Wayne A. Fuller, "Distribution of the Estimators for Autoregressive Time Series with a Unit Root," *Journal of the American Statistical Association*, 74, 1979, pp. 427–431.

Fujiki, Hiroshi, "Money Demand near Zero Interest Rate: Evidence from Regional Data," *Monetary and Economic Studies*, 20 (2), Institute for Monetary and Economic Studies, Bank of Japan, 2002, pp. 25–42.

———, and Casey B. Mulligan, "Production, Financial Sophistication, and the Demand for Money by Households and Firms," *Monetary and Economic Studies*, 14 (1), Institute for Monetary and Economic Studies, Bank of Japan, 1996a, pp. 65–103.

———, and ———, "A Structural Analysis of Money Demand: Cross-Sectional Evidence from Japan," *Monetary and Economic Studies*, 14 (2), Institute for Monetary and Economic Studies, Bank of Japan, 1996b, pp. 53–78.

Gregory, W. Allan, and Bruce E. Hansen, "Residual-Based Tests for Cointegration in Models with Regime Shifts," *Journal of Econometrics*, 70, 1996, pp. 99–126.

Hansen, Bruce E., "Test for Parameter Instability in Regressions with I(1) Processes," *Journal of Business and Economic Statistics*, 10, 1992, pp. 321–325.

Hayashi, Fumio, *Econometrics*, Princeton: Princeton University Press, 2000.

McKenzie, Colin R., and Michael McAleer, "On Efficient Estimation and Correct Inference in Models with Generated Regressors: A General Approach," *Japanese Economic Review*, 48 (4), 1997, pp. 368–389.

Miyao, Ryuzo, "Liquidity Traps and the Stability of Money Demand: Is Japan Really Trapped at the Zero Bound?" mimeo, 2003.

Nakashima, Kiyotaka, and Makoto Saito, "Strong Money Demand and Nominal Rigidity: Evidence from the Japanese Money Market under the Low Interest-Rate Policy," Discussion Paper No. 2001-11, Graduate School of Economics, Hitotsubashi University, http://wakame.econ.hit-u.ac.jp/~koho/4links/DP/saito-rigidity.pdf, 2002.

Osterwald-Lenum, Michael, "A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics," *Oxford Bulletin of Economics and Statistics*, 54 (3), 1992, pp. 461–471.

Perron, Pierre, "Further Evidence on Breaking Trend Functions in Macroeconomic Variables," *Journal of Econometrics*, 80 (2), 1997, pp. 355–385.

Phillips, Peter C. B., "Time Series Regression with a Unit Root," *Econometrica*, 55, 1987, pp. 277–301.

———, and Bruce E. Hansen, "Statistical Inference in Instrumental Variables Regression with I(1) Process," *Review of Economic Studies*, 57, 1990, pp. 99–125.

———, and Pierre Perron, "Tests for a Unit Root in Time Series Regression," *Biometrika*, 75, 1988, pp. 335–346.

Said, Said E., and David A. Dickey, "Testing for Unit Roots in Autoregressive-Moving Average Models of Unknown Order," *Biometrika*, 71, 1984, pp. 599–607.

Schwert, G. William, "Tests for Unit Roots: A Monte Carlo Investigation," *Journal of Business and Economic Statistics*, 7, 1989, pp. 147–159.

Serletis, Apostolos, *The Demand for Money*, Norwell: Kluwer Academic Publishers, 2001.

Soejima, Yutaka, "A Unit Root Test with Structural Change for Japanese Macroeconomic Variables," *Monetary and Economic Studies*, 13 (1), Institute for Monetary and Economic Studies, Bank of Japan, 1995, pp. 53–68.

Stock, James H., and Mark W. Watson, "A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems," *Econometrica*, 61, 1993, pp. 787–820.

White, Halbert, "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity," *Econometrica*, 48, 1980, pp. 817–838.