Is There a Stable Money Demand Function under the Low Interest Rate Policy? A Panel Data Analysis

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We use annual Japanese prefecture data on income, population, demand deposits, and saving deposits from 1992 to 1997 to investigate the issue of whether there exists a stable money demand function under the low interest rate policy. The evidence appears to support the contention that there does exist a stable money demand function with long-run income elasticity greater than one for M2 and less than one for M1. Furthermore, we find that Japan's money demand is sensitive to interest rate changes. However, there is no evidence of the presence of a liquidity trap.

Key words: Money demand; Interest rate; Panel data; Prefecture data

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I. Introduction

In response to the deterioration of the Japanese economy in the 1990s, expansionary fiscal and monetary policies have been implemented. However, according to "Highlights of the Budget for Fiscal 2001 (April 2001)" published by the Ministry of Finance of Japan, the dependence ratio of the general account of the national budget (*ippan kaikei*) on the issuance of the bonds on an ongoing annual basis has dramatically increased and reached 38.5 percent in the fiscal 2000 budget, up from 10.6 percent in 1990. On a stock basis, the government's gross debt relative to GDP was approximately 135.3 percent in fiscal 2000, the worst level among industrialized countries (also see Fujiki, Okina, and Shiratsuka [2001]). Has Japan's fiscal position deteriorated to an unsustainable level? Bohn (1998) suggests checking this issue in terms of (1) whether the GDP ratio of the primary balance increases as the GDP ratio of public debt rises, and (2) whether the GDP ratio of public debt does not exceed some fixed level. According to his method, both conditions need to be satisfied. Doi (2000) has used this method for the Japanese general account from fiscal 1956 to fiscal 1998 and found that the conditions for the sustainability of debt were not met. Given that the sustainability of fiscal debt is uncertain, it is natural that one might wonder if monetary policy could play a more important role in stimulating the Japanese economy. However, the effectiveness of monetary policy could be affected by many factors. Economists probably would agree that the stability of the following two relationships is critical. First, is there a stable money demand function? Second, to what extent is the money supply responsive to the operational target of the central bank? This paper focuses on the first relationship.

Nakashima and Saito (2000) use monthly aggregate time-series data to analyze whether nominal prices move inertially when nominal interest rates are extremely low in Japan. They find that the real money balance was highly elastic with respect to the nominal interest rate and real output had no clear impact on real money demand in the period between 1995 and 1999. The almost horizontal money demand function makes the nominal price level irresponsive to changes in the money supply, thus rendering the use of low interest rates to stimulate aggregate demand ineffective.

In this paper, we use data on 47 prefectures in Japan from 1992 to 1997 to study whether there exists a stable money demand function under the policy of low interest rates. There are many advantages to using panel data as opposed to using time-series or cross-sectional data. First, it allows more accurate estimates of parameters because it contains many more degrees of freedom and reduces the problem of multicollinearity that is often present in time-series data by appealing to inter-individual differences. Second, it allows a more accurate modeling of dynamic adjustment behavior with a short time-series. Third, it provides the possibility of controlling the impact of omitted variables. Fourth, it provides a possibility of controlling the impact of structural changes without relying on the conventional tests of structural breaks that are based on large sample theory with dubious finite sample properties. Fifth, it allows the possibility of controlling the problem of measurement errors (e.g., Hsiao [2001]).

We present our model in Section II. In Section III, we discuss statistical issues in connection with estimating our models. Section IV describes the data. Section V presents the empirical analysis and compares our results with other studies. Conclusions and policy implication are in Section VI.

II. The Model

The basic model for our analysis is a combination of the stock adjustment principle with a money demand equation by households and firms proposed by Fujiki and Mulligan (1996). Assuming that an agent chooses the real money balance to minimize the rental cost subject to a CES-type production function for output and transaction service, Fujiki and Mulligan (1996) derive a log-linear (desired) money demand equation of the form

$$m_{it}^{*} = \alpha_i^{*} + b^{*}y_{it} + c^{*}r_t + \epsilon_{it}, \quad \begin{array}{l} i = 1, \ldots, N \\ t = 1, \ldots, T, \end{array} \tag{1}$$

where m_{it}^{*} denotes the logarithm of the desired real money balance for agent i at time t, y denotes the logarithm of real income, and r denotes the interest rate. The intercept α_i^{*} is an approximation of the effects of rental costs of inputs to the production function of output and transaction service, which may vary across i.

The actual logarithm of real money demand, m_{it}, is assumed to follow a stock adjustment principle,[1]

$$(m_{it} - m_{i,t-1}) = \gamma^{*}(m_{it}^{*} - m_{i,t-1}) + u_{it}, \tag{2}$$

where γ^{*} denotes the speed of adjustment, which is assumed to be between zero and one, and u_{it} is the error term that is assumed to be independently, identically distributed across i and over t with mean zero and variance σ_u^2. Substituting equation (1) into equation (2) yields

$$m_{it} = (1 - \gamma^{*})m_{i,t-1} + by_{it} + cr_t + \alpha_i + v_{it}, \quad \begin{array}{l} i = 1, \ldots, N \\ t = 1, \ldots, T, \end{array} \tag{3}$$

where $b = \gamma^{*}b^{*}$, $c = \gamma^{*}c^{*}$, $\alpha_i = \gamma^{*}\alpha_i^{*}$, and $v_{it} = \gamma^{*}\epsilon_{it} + u_{it}$.

1. As pointed out by a referee, equation (2) is known as a "real adjustment mechanism" (Goldfeld [1973]). An alternative adjustment mechanism in time-series literature is a "nominal adjustment mechanism." We have performed the analysis using the nominal form as well, and the results are similar. Therefore, we only report the results in the real term.

III. Statistical Issues

A model of the form equation (3) is commonly referred to as a dynamic panel data model,

$$m_{it} = \gamma m_{i,t-1} + \underset{\sim}{\beta}'\underset{\sim}{x}_{it} + \alpha_i + v_{it}, \quad \begin{array}{l} i = 1, \ldots, N \\ t = 1, \ldots, T, \end{array} \tag{4}$$

where $\gamma = (1 - \gamma^*)$, $\underset{\sim}{x}'_{it} = (y_{it}, r_t)$, $\underset{\sim}{\beta}' = (b, c)$. When the regional-specific effect, α_i, is treated as a fixed constant, it is commonly referred to as a fixed effects (FE) model. When α_i is treated as randomly distributed across i with mean μ and variance σ^2_α, it is commonly referred to as a random effects (RE) model.

The advantage of FE specification is that it allows the presence of regional differences that can fundamentally differ across regions, and these regional-specific effects are allowed to be correlated with the included explanatory variables $(m_{i,t-1}, \underset{\sim}{x}'_{it})$. The disadvantage of FE specification is that it introduces the classical incidental parameter problem if the time-series dimension, T, is short (e.g., Neyman and Scott [1948]). The RE specification assumes that the regional differences are random draws from a common distribution and the observed differences are attributable to chance outcomes. The advantage of RE specification is that there is no incidental parameter problem. The disadvantage is that it typically does not allow the correlation between the regional-specific effect, α_i, and $\underset{\sim}{x}_{it}$. However, it does allow α_i to be correlated with $m_{i,t-1}$.

Applying the covariance transformation eliminates the regional specific effect, α_i, from the specification. However, in a dynamic model the usual covariance or within estimator is biased if T is finite (Anderson and Hsiao [1981, 1982]). To obtain a consistent estimator of γ and $\underset{\sim}{\beta}$ when N is large, we can first take the difference of equation (4) to get rid of α_i for $t = 2, \ldots, T$,

$$\Delta m_{it} = \gamma \Delta m_{i,t-1} + \underset{\sim}{\beta}'\Delta \underset{\sim}{x}_{it} + \Delta v_{it}, \quad \begin{array}{l} t = 2, \ldots, T, \\ i = 1, \ldots, N. \end{array} \tag{5}$$

where $\Delta = (1 - L)$, L denotes the lag operator that shifts the observation back by one period, $Lm_{it} = m_{i,t-1}$. Although the least squares estimator of equation (5) is inconsistent because $\Delta m_{i,t-1}$ is correlated with Δv_{it}. However, lagged $m_{i,t-j}$, $j = 2, \ldots, t-1$ are uncorrelated with Δv_{it}. Therefore, one may apply the instrumental variable (IV) or generalized method of moments estimator (GMM) to equation (5) (e.g., Ahn and Schmidt [1995] and Arellano and Bover [1995]).

Although the IV or GMM is consistent, Monte Carlo studies conducted by Hsiao, Pesaran, and Tahmiscioglu (2001) show that it is subject to serious bias and size distortion in a finite sample, in particular, if γ is close to one. On the other hand, the likelihood approach performs remarkably well in a finite sample. However, Δm_{i1} is a random variable and cannot be treated as a fixed constant when T is finite.

To complete the system, we need to add a specification for the initial value,

$$\begin{aligned}\Delta m_{i1} &= E(\Delta m_{i1} \mid \Delta \underset{\sim}{x}_{i2}, \ldots, \Delta \underset{\sim}{x}_{iT}) + v_{i1} \\ &= g + \sum_{t=2}^{T} \underset{\sim}{\pi}_t' \Delta \underset{\sim}{x}_{it} + v_{i1}, \quad i = 1, \ldots, N.\end{aligned} \tag{6}$$

We can apply a minimum distance or maximum likelihood type estimator to the combined system of equations (5) and (6). The resulting estimator is consistent and asymptotically normally distributed as $N \to \infty$ and has very good finite sample properties (Hsiao, Pesaran, and Tahmiscioglu [2001]).

When α_i is treated as a random variable, there is no incidental parameter problem. Therefore, there is no need to take the first difference of equation (4) to eliminate the individual effect, α_i. However, there is still an initial value problem because m_{i1} is a random variable and cannot be treated as a fixed constant (e.g., Hsiao [1986]). To complete the system of equation (4), Bhargava and Sargan (1983) suggest the following specification,

$$\begin{aligned}m_{i1} &= E(m_{i1} \mid \underset{\sim}{x}_{i1}, \ldots, \underset{\sim}{x}_{iT}) + v_{i1}^{*} \\ &= g^{*} + \sum_{t=1}^{T} \underset{\sim}{\pi}_t^{*\prime} \underset{\sim}{x}_{it} + v_{i1}^{*}.\end{aligned} \tag{7}$$

Applying the generalized least squares (GLS) or maximum likelihood estimator to equations (4) and (7) is consistent and asymptotically efficient (Hsiao [1986]).

IV. Data

This section explains the definition of prefectural income statistics, population, and prefectural money aggregates.

A. Prefectural Income Statistics

Prefectural income statistics compiled by the Economic and Social Research Institute (the former Economic Planning Agency of Japan) for each fiscal year provide a good counterpart to national GDP. We downloaded the data from 1987–97 from the homepage of the Economic and Social Research Institute, and supplemented them with data for 1985–86 from Fujiki and Mulligan (1996).

The prefectural income data are deflated by the gross prefectural expenditure deflator during the period from fiscal 1985 to fiscal 1997.

B. Population

We use population to convert prefectural data to per capita data. The population of each prefecture is as of the beginning of October of each year.

C. Prefectural Money Aggregates

1. MF1

First, data on demand deposits[2] held by individuals and firms at domestically licensed bank by prefecture (end of month outstanding) are available from *Financial and Economic Statistics Monthly* from the BOJ (hereafter, MF1 data).[3] Due to the extension of the number of banks covered that are included in these statistics in April 1989 and occasional consolidation of banks, MF1 data sometimes show an unusual increase, particularly in April 1989.[4]

Since the national M1 statistics are defined as the sum of cash currency in circulation and total demand deposits, net of the deposits held by the financial institutions, MF1 is the prefectural counterpart of national M1 minus cash. However, the following caveats are in order. First, MF1 data do not include cash, because regional data on the amount of currency held by individuals are not available. Second, they do not contain a breakdown by individuals or firms. Third, they do not include demand deposits at Shinkin banks, Norinchukin Bank, or Shoko Chukin Bank, which are included in the computation of M1 statistics. Therefore, the aggregate MF1 is not M1. However, MF1 data explain about 70 percent of M1 during the period from 1985 to 1988, about 80 percent from 1989 to 1991, and about 70 percent from 1992 to 1997. Therefore, if we are careful about the sample periods, MF1 predicts an almost constant proportion of M1, because M1 minus cash is almost proportional to M1 as shown in Figure 1.

2. MF2

The definition of MF2 is the sum of the deposits at domestically licensed banks, Shinkin banks, and Shoko Chukin Bank. MF2 consists of both demand deposits and savings deposits. MF2 is our counterpart of national M2+CDs minus cash, with the existence of the following statistical discrepancies.[5]

First, the prefectural breakdown of certificates of deposit (CDs) outstanding does not exist, hence we ignore it. Second, we only eliminate the deposits held by the financial institutions for domestically licensed banks, since the breakdown of deposits held by financial institutions by prefecture is available for domestically licensed banks only. Third, we exclude the data for Norinchukin Bank from the regional deposit statistics to avoid possible double-counting of the same deposits. Again, the aggregate MF2 is not M2. However, MF2 data explain about 98 percent of M2 during the period from 1985 to 1992, about 95 percent from 1993 to 1995, and about 90 percent from 1996 to 1997. Therefore, if we are careful about the sample periods, MF2 predicts an almost constant proportion of M2, because M2 minus cash is almost proportional to M2 as shown in Figure 2.

2. Substantial parts of demand deposits are either current deposits or ordinary deposits. Current deposits are deposits that the depositor may demand as freely as his or her needs require. Corporations use this account for the sake of settlement, but this account does not pay interest. The individuals and corporations with temporary excess funds mostly hold ordinary deposits.
3. Domestically licensed banks include city banks, regional banks, regional banks II, trust banks, and long-term credit banks. Note that the location of branches of each financial institution determines the prefecture to which the deposit belongs.
4. The data before March 1989 do not cover deposits at regional banks II.
5. National M2+CDs adds saving deposits and certificates of deposit to M1. The financial institutions that are authorized to accept deposits have been allowed to issue CDs since 1979. The interest rate for CDs is not regulated, and CDs may be sold to third parties.

Figure 1 Natural Logarithm of Real M1 with and without Currency

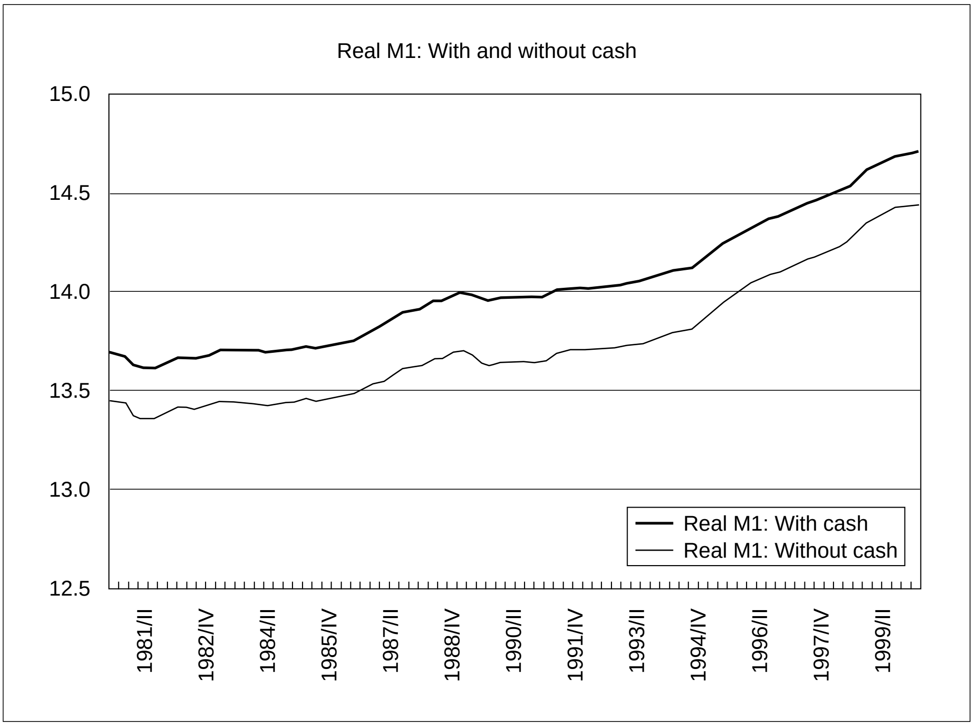
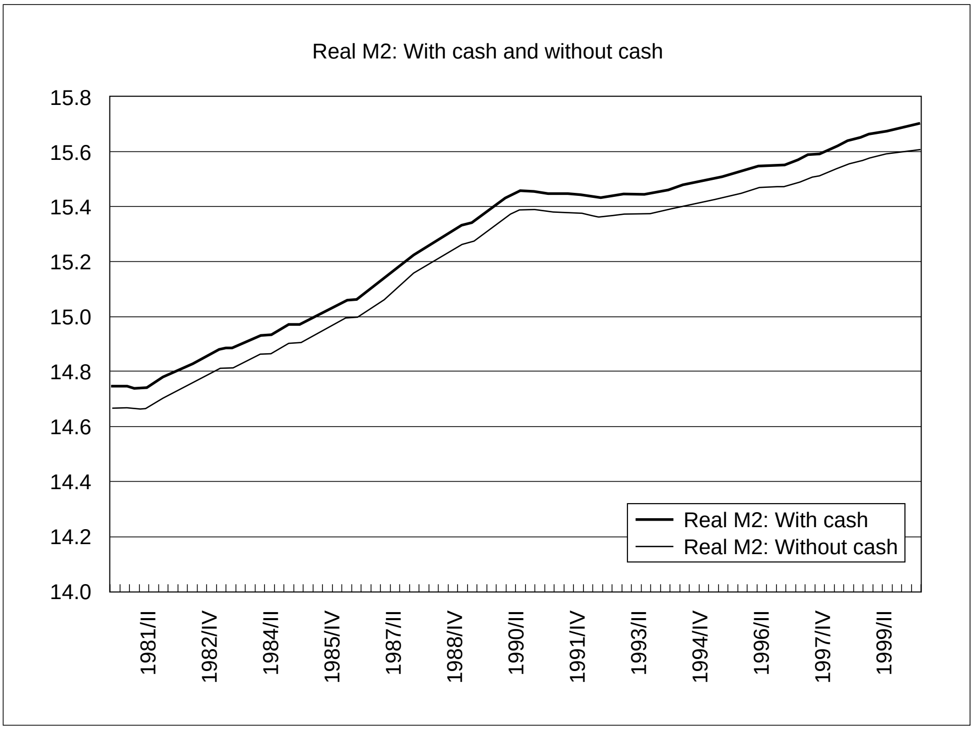


Figure 2 Natural Logarithm of Real M2 with and without Currency



D. Personal Deposits (PDs)

Personal deposits (PDs) are the sum of deposits held by the individuals at domestically licensed banks, Shinkin banks, post offices, agricultural cooperatives, fishery cooperatives, credit cooperatives, and labor credit associations surveyed at the end of March. The data for the individual deposits are available from the *Prefecture Economic Statistics and Monthly Economic Statistics* published by the BOJ.

Two important drawbacks of the personal deposit data are as follows. First, they do not contain a breakdown of the demand deposits and savings deposits. Second, they include the deposits of small businesses for the sake of business operations as long as the deposits are made in the name of an individual.

All MF1, MF2, and PD figures are deflated by the gross prefectural expenditure deflator and divided by the population in each region to obtain the per capita real money balance.

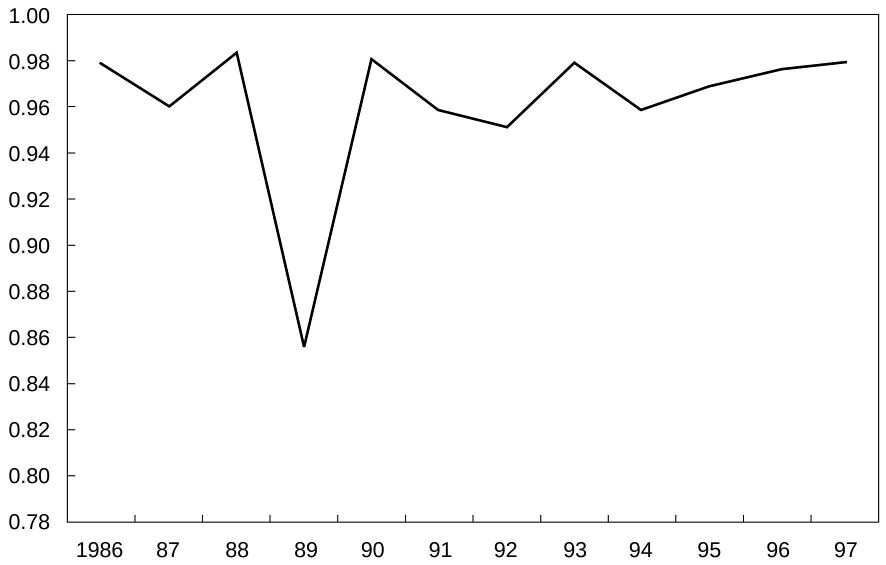
V. Empirical Results

In this section, we report the results based on panel data analysis and discuss the differences between our findings and findings based on time-series (Nakashima and Saito [2000]) or cross-sectional analysis (Fujiki [2002]).

We use prefectural data from 1985 to 1997. However, there are a number of data measurement issues raised for the sample period. First, there was a change in the definition of the banks surveyed in the deposit statistics in 1989. Due to an extension of the coverage of regional banks II in the deposit statistics in that year, the BOJ's data in the *Financial and Economic Statistics Monthly* showed an unusual increase in 1989 and the sudden collapse of the economic "bubble" in the early 1990s added large savings to the data. Second, there is an argument that people who live in suburban areas but work in large metropolitan prefectures—Tokyo, Osaka, or Kyoto—make their deposits at banks near where they work, instead of where they live. To avoid the possibility of obtaining biased results because of inconsistent data measurements in 1989 and 1990, one may just fit equation (3) for the years 1992 to 1997. To avoid the problem of people living in one prefecture but having bank deposits in other prefectures, we can exclude the data for Tokyo and its neighboring prefectures Chiba, Saitama, and Kanagawa from consideration and use the data for the remaining 43 prefectures to fit equation (3). We can also further exclude Osaka, Kyoto, and neighboring Hyogo Prefecture from consideration and perform an analysis using the data for the remaining 40 prefectures.

First, we note that the change in definitions of the coverage of banks does create some instability in the estimates. Figure 3 plots the cross-sectional estimates of the coefficient of lagged dependent variables (log(MF1)) for model equation (3) from 1986 to 1997. There is a significant drop in the coefficient in 1989. However, after 1990, it shows remarkable stability over time. Therefore, to avoid possible contamination of regression results, we concentrate on estimating the money demand equation for the period 1992–97, as the period of high interest rates in the early 1990s adds large swings to the data.

Figure 3 Cross-Sectional Estimates of the Coefficient of Lagged Dependent Variables from 1986–97



Tables 1 and 2 present the GLS estimates of the RE and the minimum distance estimation (MDE) of the FE model of MF1 for the 47 prefectures, 43 prefectures, and 40 prefectures, respectively (for details, see Appendices 1 and 2). Tables 3 and 4 present the RE and FE estimation of MF2, respectively. Tables 5 and 6 present the RE and FE estimation of PD. Practically all the model estimates have the expected signs and are statistically significant. In particular, the following points are worth noting.

Table 1 RE Estimation of MF1, 1992–97

	47 prefectures		43 prefectures		40 prefectures	
Variable	Coefficient	Standard error	Coefficient	Standard error	Coefficient	Standard error
MF1 (–1)	0.7135	0.023	0.7058	0.024	0.7274	0.021
Income	0.2785	0.101	0.3526	0.109	0.3603	0.093
Call rate	–0.0562	0.004	–0.0538	0.004	–0.0499	0.004
Constant	–0.2923	4.139	–0.5321	3.963	–0.6012	4.146

Table 2 FE Estimation of MF1, 1992–97

	47 prefectures		43 prefectures		40 prefectures	
Variable	Coefficient	Standard error	Coefficient	Standard error	Coefficient	Standard error
MF1 (–1)	0.7423	0.022	0.7312	0.021	0.7194	0.017
Income	0.7421	0.082	0.6907	0.078	0.4928	0.073
Call rate	–0.0246	0.004	–0.0253	0.003	–0.0362	0.003

Table 3 RE Estimation of MF2, 1992–97

	47 prefectures		43 prefectures		40 prefectures	
Variable	Coefficient	Standard error	Coefficient	Standard error	Coefficient	Standard error
MF2 (–1)	0.5341	0.028	0.4759	0.036	0.4816	0.040
Income	0.1046	0.051	0.1682	0.060	0.1509	0.064
Call rate	–0.0188	0.002	–0.0200	0.002	–0.0204	0.003
Constant	1.2678	1.299	1.2436	1.235	1.2655	1.156

Table 4 FE Estimation of MF2, 1992–97

	47 prefectures		43 prefectures		40 prefectures	
Variable	Coefficient	Standard error	Coefficient	Standard error	Coefficient	Standard error
MF2 (–1)	0.5728	0.035	0.5821	0.037	0.5260	0.034
Income	0.2267	0.046	0.1989	0.044	0.1342	0.048
Call rate	–0.0155	0.002	–0.0164	0.002	–0.0189	0.002

Table 5 RE Estimation of PD, 1992–97

	47 prefectures		43 prefectures		40 prefectures	
Variable	Coefficient	Standard error	Coefficient	Standard error	Coefficient	Standard error
PD (–1)	0.5701	0.023	0.5705	0.023	0.5776	0.023
Income	0.0680	0.041	0.0826	0.042	0.0828	0.042
Call rate	–0.0276	0.002	–0.0267	0.002	–0.0259	0.002
Constant	1.4446	0.616	1.3940	0.619	1.3611	0.652

Table 6 FE Estimation of PD, 1992–97

	47 prefectures		43 prefectures		40 prefectures	
Variable	Coefficient	Standard error	Coefficient	Standard error	Coefficient	Standard error
PD (–1)	0.6572	0.021	0.6737	0.022	0.6117	0.022
Income	0.0710	0.032	0.0987	0.030	0.0374	0.036
Call rate	–0.0246	0.002	–0.0234	0.001	–0.0280	0.002

Table 7 Hausman Test of the Presence of Measurement Error

Variables	RE model	FE model
MF1	—	1163.21
MF2	11.49*	28.23^
PD	21.77*	26.027

Notes: —: Hausman test statistics are negative.
^ : Call rate is deleted to avoid the singularity problem.
* : Test statistics based on instrumental variable (IV) results.

First, the data for Tokyo, Osaka, Kyoto, and their neighboring prefectures probably contain some systematic measurement errors. Table 7 presents the Hausman specification test of the presence of measurement errors by comparing the differences between the coefficients estimates based on 47 prefectures and 40 prefectures. They appear to confirm the presence of measurement errors in the seven prefectures we exclude from consideration. Both the coefficients of the lagged dependent variables and income variables for the 47 prefectures differ somewhat from the estimates for the 40 prefectures. However, the coefficients of the interest rate are remarkably stable across estimates using data for different prefectures, indicating that the substitution effects between money and other financial assets are not affected by the issue of whether people living in one prefecture could have bank accounts in a different prefecture.

Second, the income elasticity of money demand is positive and statistically significant. Based on the results of using data for 40 prefectures, the short-run income elasticity for MF1 is about 0.36 for the RE model and about 0.493 for the FE model. The long-run elasticity is 0.36/(1 – 0.728) = 1.32 for the RE model and 0.493/(1 – 0.719) = 1.75 for the FE model. The short-run income elasticity for MF2 is about 0.151 for the RE model and 0.134 for the FE model. The long-run income elasticity for MF2 is about 0.29 for the RE model and about 0.28 for the FE model. The short-run income elasticity for PD is 0.08 for the RE model and 0.037 for the FE model. The long-run income elasticity is 0.196 for the RE model and 0.1 for the FE model.[6]

Third, the coefficients of the interest rate are negative and statistically significant. The short-run semi-interest rate elasticity for MF1 is about –0.05 for the RE model and –0.036 for the FE models. The long-run semi-interest rate elasticity is about –0.18 for the RE model and –0.14 for the FE model. The short-run semi-interest rate elasticity for MF2 is about –0.02 for the RE model and –0.019 for the FE model. The long-run semi-interest rate elasticity for MF2 is about –0.04 for the RE model and –0.04 for the FE model. The short-run semi-interest rate elasticity for PD is –0.026 for the RE model and –0.028 for the FE model. The long-run semi-interest rate elasticity is –0.06 for the RE model and –0.07 for the FE model.

Fourth, there are some differences between the RE and FE estimation, although they are not substantial. Which model provides a more reliable inference? Unfortunately, the Hausman (1978) specification test of RE versus FE specification cannot be implemented, because the estimated covariance matrix is negative. Therefore, to check the reliability of the RE versus FE inference, we rely on the prediction principle (Hsiao and Sun [2000]). We reestimate the RE and FE models for the period 1992–96 and use the estimated coefficients to predict the outcomes for 1997. Figures 4–9 plot the actual and predicted values for the 40 prefectures in 1997. It is quite remarkable how

6. One might argue that since high-net-worth individuals hold large amounts of financial assets such as large savings deposits, income elasticity of MF2 should be larger compared with MF1. However, our result shows that long-run income elasticity of MF1 is far larger than that of MF2. One interpretation of this evidence might be that a substantial part of demand deposits is held by firms, while savings deposits are presumably held by individuals. Hence, if our dynamic panel approach is correct, relatively high-income elasticity of MF1 could be due to the demand for money by firms. The idea is consistent with the evidence that personal deposits, which exclude deposits made by firms, show the smallest income elasticity of money demand. Information on the distribution of demand deposits held by firms might provide such evidence.

Figure 4 Post-Sample Actual and RE Predicted Values of 1997 MF1 for the 40 Prefectures

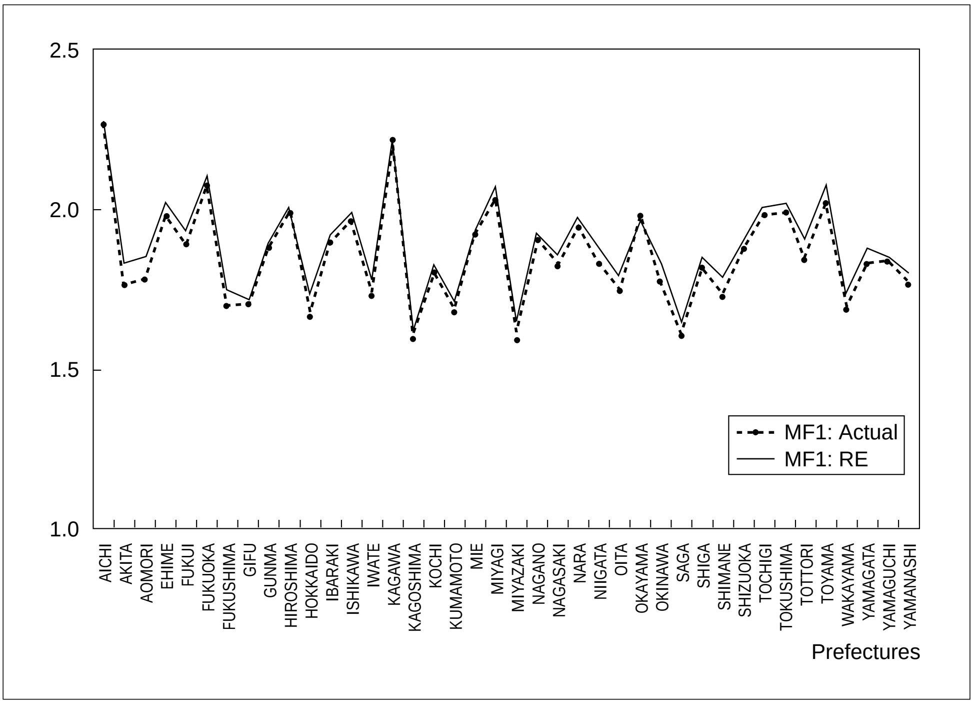


Figure 5 Post-Sample Actual and FE Predicted Values of 1997 MF1 for the 40 Prefectures

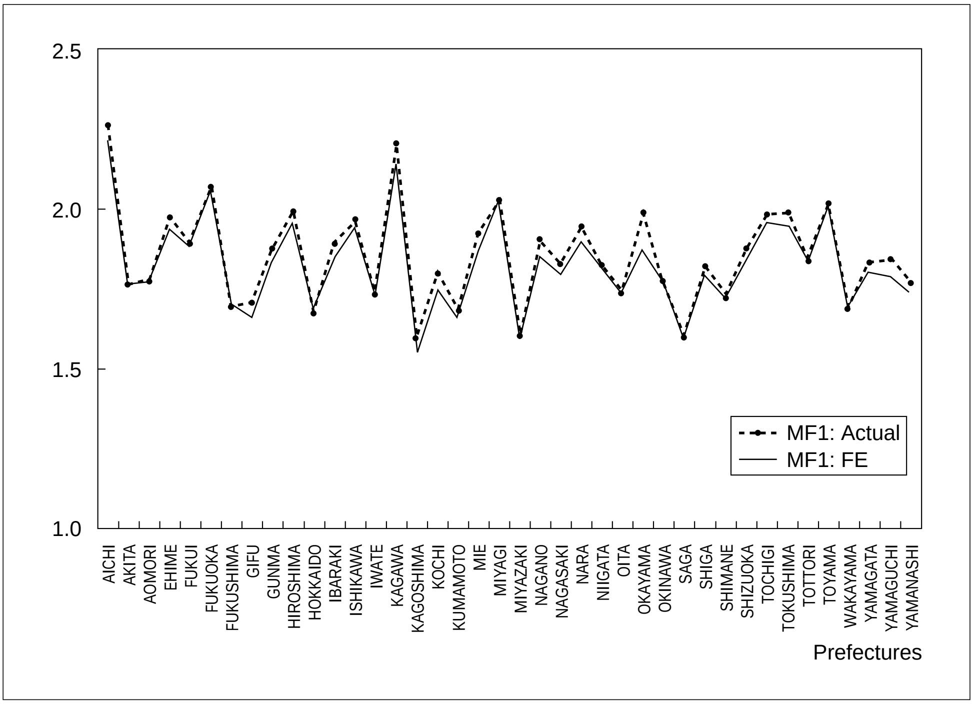


Figure 6 Post-Sample Actual and RE Predicted Values of 1997 MF2 for the 40 Prefectures

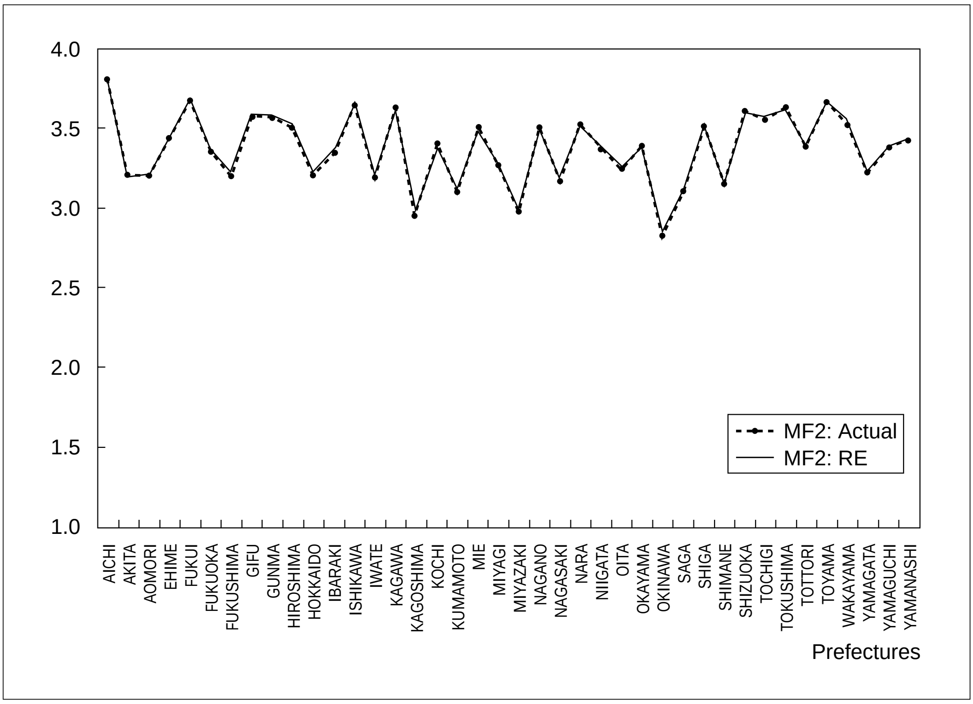


Figure 7 Post-Sample Actual and FE Predicted Values of 1997 MF2 for the 40 Prefectures

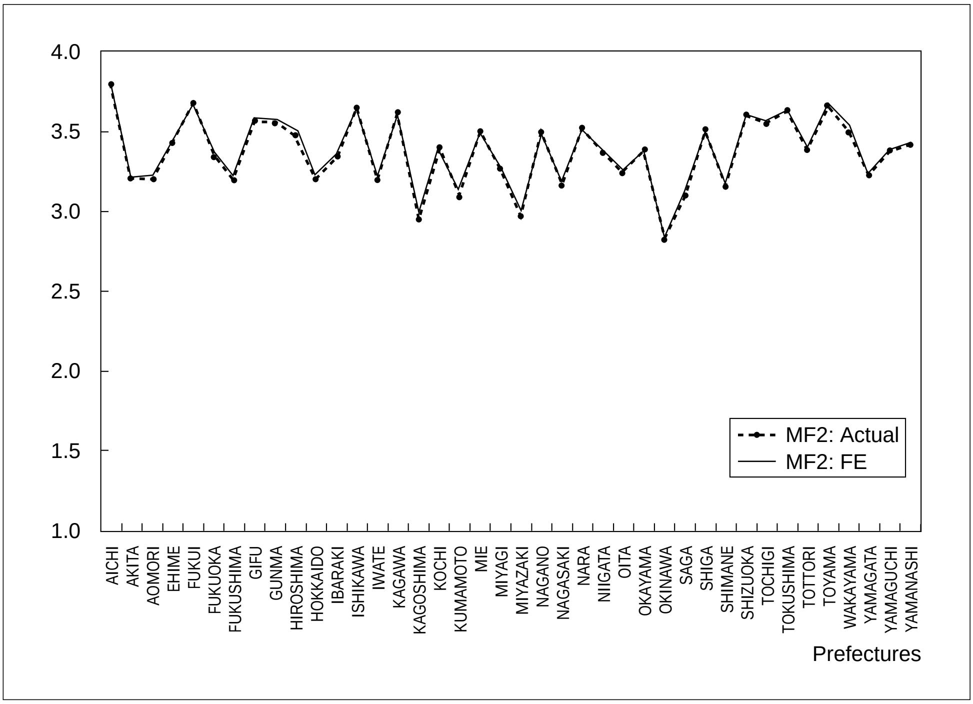


Figure 8 Post-Sample Actual and RE Predicted Values of 1997 PD for the 40 Prefectures

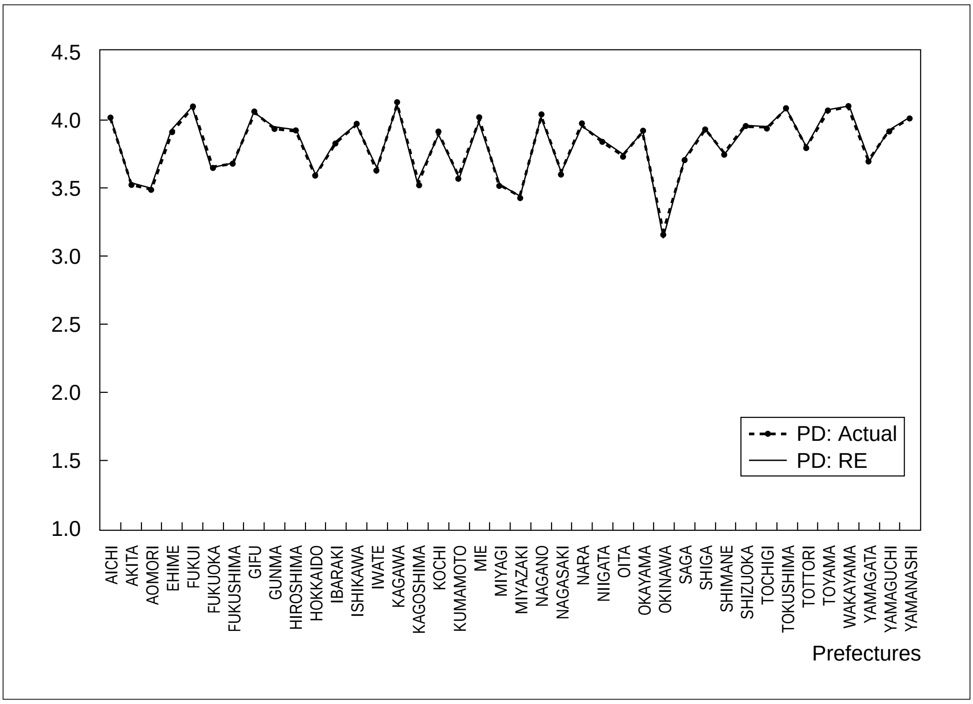
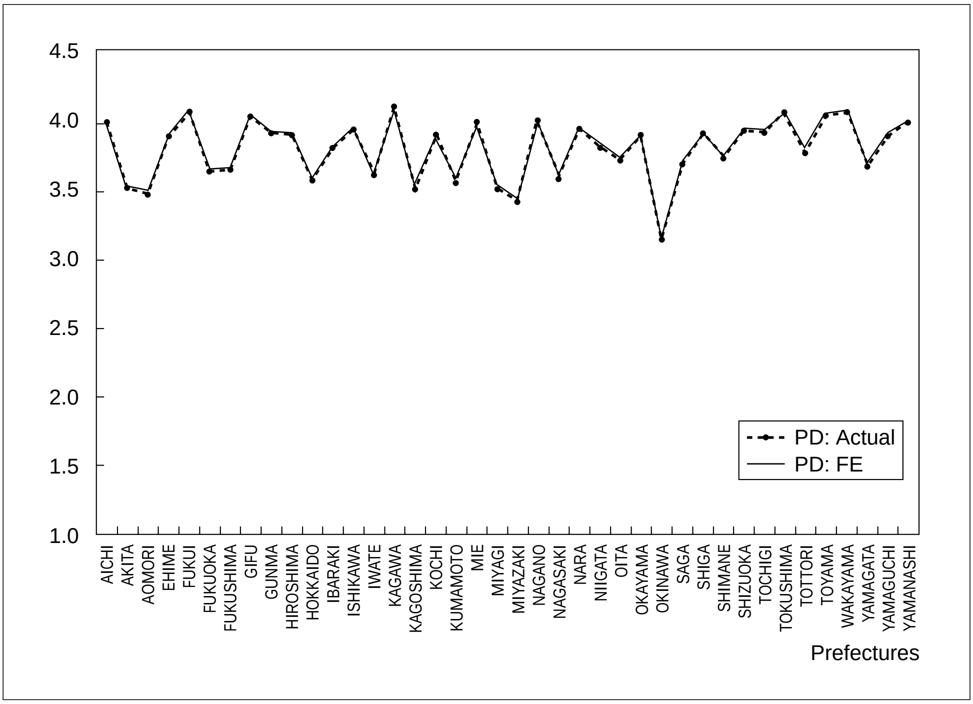


Figure 9 Post-Sample Actual and FE Predicted Values of 1997 PD for the 40 Prefectures



well both models predict the outcomes. Table 8 provides the root mean square prediction error of these four models. Again the difference is not significant, although it does appear to favor RE specification slightly.

Table 8 Root Mean Square Prediction Error Comparison

	Variables	RE	FE
47 prefectures	MF1	0.2464	0.3991
	MF2	0.1298	0.0947
	PD	0.0577	0.0699
43 prefectures	MF1	0.1984	0.3217
	MF2	0.0986	0.0802
	PD	0.0571	0.0750
40 prefectures	MF1	0.2563	0.2396
	MF2	0.0928	0.1235
	PD	0.0569	0.0875

Using the information from the panel data, we find that there appears to be a stable relation between Japan's demand for real balance and real income and the nominal interest rate, even during the period of low interest rates, whether we use an RE or FE specification. Table 9 summarizes the estimated income elasticity and semi-interest rate elasticity based on data for 40 prefectures. They are of similar magnitude between the RE and FE specifications. On the other hand, Nakashima and Saito (2000), using monthly aggregate time-series data, find that there was a structural break in 1995 and there did not appear to be a stable relation between money demand and income for the period 1995 to 1999. Moreover, they find that money demand was extremely interest rate-elastic, implying the existence of a liquidity trap. Unfortunately, our annual panel data contain too little time dimension-related information to directly test for a structural break in 1995. However, if there was indeed a structural break in 1995, then one would expect that estimates based on 1992–96 data probably would not predict the outcomes of 1997 well. But Figures 4–9 show that the predictions for 1997 are borne out remarkably well. This may be viewed as indirect evidence in support of a stable disaggregated money demand function. Furthermore, although we find that money demand is responsive to interest rate changes, they are not of the magnitude of Nakashima and Saito (2000). Their estimated semi-interest rate elasticity for M1 is in the range of –0.415 to –0.592. Ours is much smaller: the long-run semi-interest rate elasticity for MF1 is about –0.14 based on the FE model and –0.18 based on the RE model.

Table 9 Estimated Income Elasticity and Semi-Interest Rate Elasticity

Elasticities of interest		MF1		MF2		PD	
		Short run	Long run	Short run	Long run	Short run	Long run
Income elasticity	RE	0.36	1.32	0.15	0.29	0.08	0.196
	FE	0.49	1.75	0.13	0.28	0.037	0.1
Semi-interest rate elasticity	RE	–0.05	–0.18	–0.02	–0.04	–0.026	–0.06
	FE	–0.04	–0.14	–0.019	–0.04	–0.028	–0.07

Compared to the study that also uses panel data, Fujiki (2002) obtains employee income elasticities of MF1 of about one, while our estimated short-run income elasticity is significantly below one and the implied long-run income elasticity is above one. However, there is a significant difference in the two model specifications. First, cross-sectional estimates use a static model while our model is a dynamic one. Secondly, cross-sectional estimates do not use the call rate as an explanatory variable. We find that both the coefficients of the lagged dependent variable and the call rate are highly significant.

A referee has suggested using the gross prefectural product to approximate regional economic activity because the prefectural income data represent income received by residents of each specific area, regardless of the location of the economic activity that generates the income. Tables 10, 11, 12, and 13 present the RE and FE estimates of a regional MF1 and MF2 demand model using gross prefectural product instead of gross prefectural income. The results are very similar, again appearing to support a stable relationship between disaggregated money demand and economic activity.

Table 10 RE Estimation of MF1 Using Gross Prefectural Product per Capita (GPPP), 1992–97

	47 prefectures		43 prefectures		40 prefectures	
Variable	Coefficient	Standard error	Coefficient	Standard error	Coefficient	Standard error
MF1 (–1)	0.6930	0.024	0.6818	0.025	0.7038	0.023
GPPP	0.3101	0.119	0.3850	0.126	0.3511	0.116
Call rate	–0.0587	0.003	–0.0571	0.004	–0.0541	0.003
Constant	–1.8618	12.591	–2.4640	12.145	–2.2322	11.772

Table 11 FE Estimation of MF1 Using GPPP, 1992–97

	47 prefectures		43 prefectures		40 prefectures	
Variable	Coefficient	Standard error	Coefficient	Standard error	Coefficient	Standard error
MF1 (–1)	0.6935	0.025	0.6998	0.025	0.6816	0.019
GPPP	0.6500	0.116	0.6277	0.113	0.4633	0.097
Call rate	–0.0306	0.004	–0.0298	0.003	–0.0428	0.003

Table 12 RE Estimation of MF2 Using GPPP, 1992–97

	47 prefectures		43 prefectures		40 prefectures	
Variable	Coefficient	Standard error	Coefficient	Standard error	Coefficient	Standard error
MF2 (–1)	0.5707	0.026	0.4715	0.040	0.4767	0.045
GPPP	–0.0054	0.053	0.1177	0.072	0.0914	0.082
Call rate	–0.0202	0.002	–0.0220	0.002	–0.0226	0.002
Constant	1.5367	3.513	0.8552	3.210	1.0357	2.774

Table 13 FE Estimation of MF2 Using GPPP, 1992–97

	47 prefectures		43 prefectures		40 prefectures	
Variable	Coefficient	Standard error	Coefficient	Standard error	Coefficient	Standard error
MF2 (–1)	0.5520	0.038	0.5494	0.040	0.4879	0.038
GPPP	0.2027	0.061	0.2532	0.060	0.1801	0.064
Call rate	–0.0160	0.002	–0.0164	0.002	–0.0196	0.002

VI. Conclusions

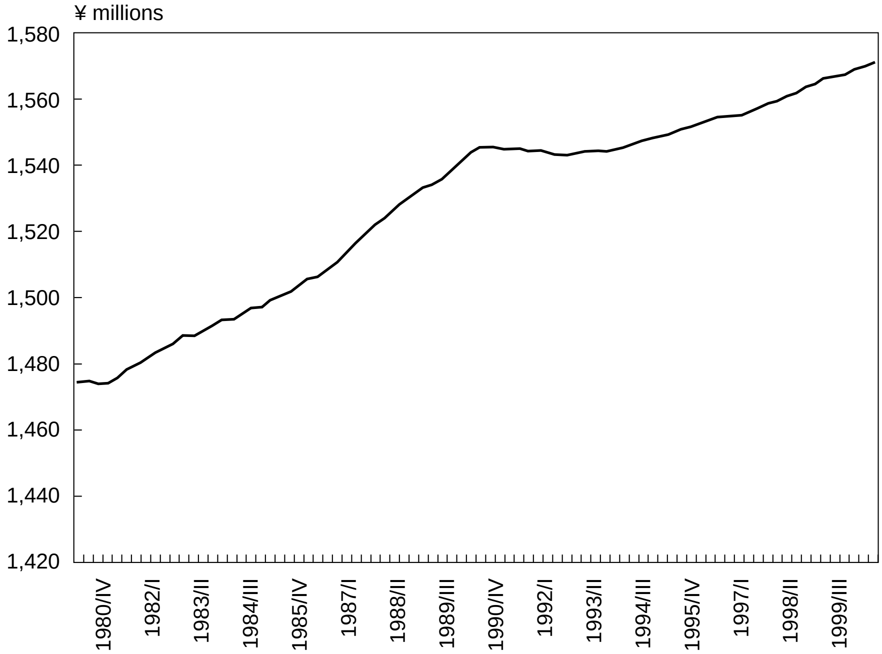
In this paper, we used Japanese prefectural data from 1992–97 to estimate the money demand equations. Contrary to the findings relying on an aggregate time-series, we found that there was a stable money demand equation for Japan even during the period of low interest rates. Based on the results of the RE dynamic panel data model, the estimated short-run income elasticity is about 0.493 and the long-run income elasticity is about 1.32 for MF1, 0.151 and 0.29, respectively, for MF2, and 0.08 and 0.196, respectively, for PD. The estimated short-run semi-interest rate elasticity is about –0.05 and the long-run semi-elasticity is about –0.18 for MF1, –0.02 and –0.04, respectively, for MF2 and –0.026 and –0.06, respectively, for PD.

The conflicting evidence between the analysis based on aggregate time-series data and disaggregated panel data could be due to many reasons. First, our analysis is in fact an analysis of the demand for deposits of various types, because panel data on holdings of currency are not available. However, in the Japanese economy currency is widely used, especially by households. Second, there could be an issue of aggregation. Third, there could be an issue of simultaneity between the aggregate money and income. Fourth, the most troublesome issue concerning the analysis of aggregate time-series data is the lack of sample variability. The minimum and maximum values of the logarithm are 14.943 and 15.4925 for real GDP, 13.6094 and 17.7069 for real M1, and 14.738 and 15.7089 for real M2, respectively, for the quarterly data over the period 1980/IV–2000/IV. With sample observations clustered together, any regression results are possible depending on the period covered or variability of a particular pair of observations. We plan to investigate the discrepancy between aggregate and disaggregate time-series in the future. However, if there indeed exists a stable real money demand equation, then the following elementary argument presumably should hold: "The monetary authorities can issue as much money as they like. Hence, if the price level were truly independent of money issuance, then the monetary authorities could use the money they create to acquire indefinite quantities of goods and assets. This is manifestly impossible in equilibrium. Therefore, money issuance must ultimately raise the price level, even if nominal interest rates are bounded at zero" (Bernanke [2000]).

Then why did monetary authorities fail to stimulate aggregate demand and prices in the 1990s? If the estimate provides any guidance, it is not because of the ineffectiveness of the low interest rate policy, but perhaps because the money supply did not increase as much as desired by the monetary authorities. Figure 10 plots M2 from 1980/I–2000/IV. It is obvious that the growth rate of M2 in the 1990s failed to maintain the same rate as in the 1980s. In the 1980s, the average growth rate was about 9.34 percent, yet the inflation rate (GDP deflator) was only 1.98 percent (with a real GDP growth rate of 4.13 percent). In the 1990s, the average growth rate of M2 was only 2.69 percent, with an inflation rate of 0.14 percent (and a real GDP growth rate of 1.38 percent). This significant drop in the growth rate of the money supply was mainly due to the reluctance of commercial banks to make loans to small and medium-sized enterprises because of the erosion of their capital base due to the accumulation of nonperforming assets after the economic "bubble" burst in the early 1990s. In fact, the growth rate of high-powered money was about 5.67 percent in the 1990s (compared to 8.08 percent in the 1980s). It is the ineffectiveness of the transmission of the growth of high-powered money to the growth of M2 that led to the slowdown of growth in the money supply. Moreover, the purchasing of long-term bonds is likely to push the interest rate further down and money demand is sensitive to interest rate changes.

It appears that the challenge faced by the monetary authorities to find a way to increase the money supply cannot be resolved through monetary means alone. Complementary fiscal policies must be implemented. If the U.S. experience could be applied to Japan, the policy option of raising taxes for high-income families may

Figure 10 Quarterly M2 Data, 1980/I to 2000/IV



Note: M2 is seasonally adjusted.

deserve serious study. Raising the taxes of high-income families within bounds may have a negligible discouraging effect on consumption and investment. After all, the Clinton administration imposed a 10 percent surcharge on high-income families but U.S. consumption and investment remained strong in the 1990s. With the increased revenue from the income tax surcharge, the government could retire the bad loans held by financial institutions. Ideally, with their improved balance sheets, commercial banks would be more willing to lend to small and medium-sized enterprises, and this would lead to an increase in the money supply and get Japan out of deflation. However, taxing the wealthy in Japan might mean taxing the elderly, and could further discourage consumption if uncertainty regarding the social security system were an important factor. Thus, it appears that a case can be made for conducting serious empirical study of the discouraging effects on consumption and investment of a tax surcharge on high-income families.

APPENDIX 1: SPECIFICATION AND ESTIMATION IN THE GLS FRAMEWORK FOR THE RE MODEL

We start with a model

$$y_{it} = \rho y_{i,t-1} + \underset{\sim}{\beta}'\underset{\sim}{x}_{it} + \underset{\sim}{\gamma}'\underset{\sim}{z}_i + v_{it}, \quad i = 1, \ldots, N, \; t = 2, \ldots, T, \qquad \text{(A.1)}$$

where $\underset{\sim}{x}_{it}$ is a $k_1 \times 1$ vector of time-variant explanatory variables, and $\underset{\sim}{z}_i$ is a $k_2 \times 1$ vector of time-invariant explanatory variables including the constant term, $v_{it} = \alpha_i + u_{it}$. The error term u_{it} and the prefecture-specific effects α_i satisfy

$$\begin{aligned}
&E\alpha_i = Eu_{it} = 0, \quad E\alpha_i\underset{\sim}{z}'_{it} = 0', \quad E\alpha_i\underset{\sim}{x}'_{it} = \underset{\sim}{0}', \\
&E\alpha_i u_{it} = 0, \\
&E\alpha_i\alpha_j = \sigma_\alpha^2 \text{ if } i = j, \\
&\qquad\quad = 0 \text{ if otherwise} \\
&Eu_{it}u_{js} = \sigma_u^2 \text{ if } i = j, \; t = s, \\
&\qquad\quad = 0 \text{ if otherwise}
\end{aligned}$$

and ρ, $\underset{\sim}{\beta}$, and $\underset{\sim}{\gamma}$ are parameters of interest. For the model in this paper, $\underset{\sim}{x}_{it}$ includes prefectural income and the call rate, and z_i is an intercept term.

To complete the system, we let

$$y_{i0} = \underset{\sim}{\bar{\pi}}'\underset{\sim}{\bar{x}}_i + \underset{\sim}{\bar{\gamma}}'\underset{\sim}{z}_i + v_{i0}, \quad i = 1, \ldots, N, \qquad \text{(A.2)}$$

where y_{i0} is the initial observation for i, $\underset{\sim}{\bar{x}}_i = \frac{1}{T}\sum_{t=1}^{T}\underset{\sim}{x}_{it}$. The GLS estimates for equations (A.1) and (A.2) are given by

$$\hat{\delta}_{GLS} = \left(\sum_{i=1}^{N} X_i' V^{-1} X_i\right)^{-1} \left(\sum_{i=1}^{N} X_i' V^{-1} y_i\right),$$

where $\delta = (\underset{\sim}{\bar{\pi}}_i', \underset{\sim}{\bar{\gamma}}', \rho, \underset{\sim}{\beta}', \underset{\sim}{\gamma}')$,

$$X_i = \begin{bmatrix} \underset{\sim}{x}_i & \underset{\sim}{z}_i & 0 & 0 & 0 \\ 0 & 0 & y_{i0} & \underset{\sim}{x}'_{i1} & \underset{\sim}{z}'_i \\ \vdots & \vdots & \vdots & \vdots & \vdots \\ \vdots & \vdots & \vdots & \vdots & \vdots \\ 0 & 0 & y_{i,T-1} & \underset{\sim}{x}'_{iT} & \underset{\sim}{z}'_i \end{bmatrix},$$

$$V = \begin{bmatrix} \sigma_{v_0}^2 & r_{01} & \cdots & \cdots r_{0T} & \\ r_{01} & \sigma_u^2 + \sigma_\alpha^2 & \sigma_\alpha^2 & & \sigma_\alpha^2 \\ \vdots & \sigma_\alpha^2 & \ddots & \ddots & \\ r_{0t} & \sigma_\alpha^2 & & & \sigma_u^2 + \sigma_\alpha^2 \end{bmatrix},$$

since $V(v_{it}) = \sigma_u^2 + \sigma_\alpha^2$, $E(v_{it}v_{is}) = \sigma_\alpha^2$ for $t = 1, 2, \ldots, T$, and $\underset{\sim}{y}_i = (y_{i0}, y_{i1} \ldots . y_{iT})'$.

To obtain the initial values for the implementation of the GLS estimation, we first take the first difference of equation (A.1), obtaining

$$y_{it} - y_{i,t-1} = \rho(y_{i,t-1} - y_{i,t-2}) + \underset{\sim}{\beta}'(\underset{\sim}{x}_{it} - \underset{\sim}{x}_{i,t-1}) + u_{it} - u_{i,t-1}. \tag{A.3}$$

Since by assumption $y_{i,t-2}$ is not correlated with $u_{it} - u_{i,t-1}$ but is correlated with $y_{i,t-1} - y_{i,t-2}$, we use $y_{i,t-2}$ as an instrument for $y_{i,t-1} - y_{i,t-2}$ and estimate $\underset{\sim}{\beta}$ and ρ by the instrumental variable method.

Second, we substitute estimated $\underset{\sim}{\beta}$ and ρ into

$$\bar{y}_i - \gamma\bar{y}_{i,-1} - \underset{\sim}{\beta}'\bar{\underset{\sim}{x}}_i = \underset{\sim}{\gamma}'\underset{\sim}{z}_i + \alpha_i + \bar{u}_i, \tag{A.4}$$

to estimate $\underset{\sim}{\gamma}$ using the ordinary least squares (OLS) method, where $\bar{y}_i$, $\bar{\underset{\sim}{x}}_i$, and $\bar{u}_i$ are averages taking over T for prefecture i.

We then can estimate σ_u^2 based on equation (A.3):

$$\sigma_u^2 = \frac{\sum_{i=1}^{N}\sum_{t=2}^{T}\left[(y_{it} - y_{i,t-1}) - \hat{\rho}(y_{i,t-1} - y_{i,t-2}) - \hat{\underset{\sim}{\beta}}'(\underset{\sim}{x}_{it} - \underset{\sim}{x}_{i,t-1})\right]^2}{2N(T-1)}$$

and σ_α^2 is estimated by

$$\sigma_\alpha^2 = \frac{\sum_{i=1}^{N}(\bar{y}_i - \hat{\rho}\bar{y}_{i,-1} - \hat{\underset{\sim}{\beta}}'\bar{\underset{\sim}{x}}_i)^2}{N} - \frac{1}{T}\hat{\sigma}_u^2.$$

To obtain estimates of $\sigma_{v_0}^2$ and the covariance between v_{i0} and v_{it}, we can first use the OLS procedure to estimate equation (A.2) cross-sectionally, then use the estimated error sum of the squares to estimate the initial variance $\sigma_{v_0}^2$. To estimate the covariance between v_{i0} and v_{it}, we first plug in the estimated ρ, β, and γ into equation (A.1) to estimate v_{it}, then estimate the covariances by

$$r_{0t} = \mathrm{cov}(v_{i0}, v_{it}) = \frac{\sum_{i=1}^{N}(v_{it} - \bar{v}_i)v_{i0}}{N}.$$

APPENDIX 2: MINIMUM DISTANCE ESTIMATION (MDE) FOR THE FE MODEL

We take the first difference of equation (A.1) to eliminate α_i, obtaining

$$\Delta y_{it} = \gamma\Delta y_{i,t-1} + \underset{\sim}{\beta}'\Delta\underset{\sim}{x}_{it} + \Delta u_{it}, \quad \begin{array}{l} t = 2, 3, \ldots, T \\ i = 1, 2, \ldots, N \end{array}. \tag{A.5}$$

Equation (A.5) is well defined for $t = 2, \ldots, T$ but not for $t = 1$, since $y_{i,-1}$ is not available. The marginal distribution of Δy_{i1} conditional on $\Delta\underset{\sim}{x}_i$, can be written as

$$\Delta y_{i1} = b^* + \underset{\sim}{\pi}'\Delta \underset{\sim}{x}_i + \underset{\sim}{v}_{i1} \tag{A.6}$$

where $\underset{\sim}{\pi}$ is a $(T-1) \times k_1 \times 1$ vector of unknown coefficients that in general varies independently of the variations of β and ρ, and $\Delta\underset{\sim}{x}_i = (\Delta\underset{\sim}{x}'_{i2}, \ldots, \Delta\underset{\sim}{x}'_{iT})'$.[7] We consider $\underset{\sim}{x}_i$ to be strictly exogenous and the likelihood function is given by

$$(2\pi)^{-\frac{NT}{2}} |\Omega|^{-\frac{N}{2}} \exp\left\{-\frac{1}{2}\sum_{i=1}^{N} \Delta\underset{\sim}{u}_i^{*\prime}\Omega^{-1}\Delta\underset{\sim}{u}_i^{*}\right\} \tag{A.7}$$

where

$$\Delta\underset{\sim}{u}_i^* = [\Delta y_{i1} - b^* - \underset{\sim}{\pi}'\Delta\underset{\sim}{x}_{i1}, \Delta y_{i2} - \gamma\Delta y_{i1} - \underset{\sim}{\beta}'\Delta\underset{\sim}{x}_{i2}, \ldots, \Delta y_{iT} - \gamma\Delta y_{iT-1} - \underset{\sim}{\beta}'\Delta\underset{\sim}{x}_{iT}]',$$

and

$$\Omega = \sigma_u^2 \begin{bmatrix} w & -1 & 0 & \cdots & 0 & \\ -1 & 2 & -1 & 0 & \cdots & \\ 0 & -1 & 2 & -1 & \cdots & \\ \vdots & & & & \ddots & -1 \\ 0 & & & -1 & 2 & \end{bmatrix} = \sigma_u^2\Omega^*,$$

where $w = \dfrac{1}{\sigma_u^2}\operatorname{var}(\Delta y_{i1})$.

The MLE estimator is highly nonlinear. A simple but less efficient estimator of equations (A.5) and (A.6) is to estimate $\underset{\sim}{\theta} = (\gamma, \underset{\sim}{\beta}')'$ by minimum distance estimation (MDE):

$$\hat{\underset{\sim}{\theta}} = \begin{pmatrix} \hat{\gamma} \\ \underset{\sim}{\beta} \end{pmatrix} = \left[\sum_{i=1}^{N} \Delta Z_i'\Omega^{*-1}\Delta Z_i\right]^{-1}\left[\sum_{i=1}^{N} \Delta Z_i'\Omega^{*-1}\Delta\underset{\sim}{y}_i\right],$$

where

$$\Delta Z_i = \begin{bmatrix} 1 & \Delta\underset{\sim}{x}_i' & 0 & 0 & \\ 0 & 0 & \Delta y_{i1} & \Delta\underset{\sim}{x}_{i2}' & \\ & \vdots & \vdots & \vdots & \vdots \\ 0 & 0 & \Delta y_{i,T-1} & \Delta\underset{\sim}{x}_{iT}' & \end{bmatrix}.$$

In our estimation, to avoid the singularity problem, we use $\Delta\bar{\underset{\sim}{x}}_i'$ instead, where $\Delta\bar{\underset{\sim}{x}}_i'$ contains averages of each explanatory variable over time.

The variable covariance matrix for $\hat{\gamma}$ is estimated by $\operatorname{cov}(\hat{\underset{\sim}{\theta}}) = \hat{\sigma}_u^2 \cdot \left[\sum_{i=1}^{N}\Delta Z_i'\Omega^{*-1}\Delta Z_i\right]^{-1}$.

7. Please refer to Hsiao, Pesaran, and Tahmiscioglu (2001) for details of specification and a discussion of the strictly exogenous and weakly exogenous assumptions of x_i.

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