A Structural Analysis of Money Demand: Cross-Sectional Evidence from Japan

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Following the study of U.S. regional data by Mulligan and Sala-i-Martin (1992) and the discussion by Fujiki and Mulligan (1996a) of empirical models of the demand for money, the paper uses Japanese prefectural data to estimate the parameters of a money demand function. The cross-sectional estimates of the income elasticity for a counterpart of M_2 minus currency are in the range of 1.2–1.4 and appear stable over time. The cross-sectional income elasticities are used to estimate the interest rate elasticity of money demand from the macro time series data, and to assess changes over time in the degree of financial sophistication.

Key words: Demand for money; Monetary policy

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The authors' research has benefited from the comments of an anonymous referee. Financial support by the Bank of Japan is gratefully acknowledged.

I. Introduction

We suppose that monetary phenomena are determined by three sets of variables: (1) parameters of agents' utility or production functions; (2) technological constraints; and (3) monetary policy. We show how relationships between the money stock, GNP, and other variables depend on (1), (2), and (3), and how parameters of agents' utility or production functions can be estimated from aggregate data.

Following Mulligan and Sala-i-Martin (1992), we note that estimation of structural parameters from aggregate data requires several identifying restrictions, but that those restrictions are more plausible in a cross-regional analysis than in an aggregate time series analysis. We see at least three advantages to estimate money demand function from cross-sectional data. First, interest rates are likely to be constant across cross-sectional units. Hence, we can avoid one of the problems of time series analysis—the relevant opportunity cost of money is not so easily measured. Second, cross-sectionally, our money can be consistently measured, while time series analyses often suffer from the problem of the changes in the composition and definitions of an aggregate monetary measure. Moreover, one can suppose that the relative prices and productivities that determine the amount of substitution among various forms of "money" do not vary across regions at a point in time or vary in a way that is uncorrelated with the scale of operation. Third, we can quantify shifts of money demand function more directly than with the time series approach. Time series approaches obviously miss the structural changes that occur at the end of sample periods simply because they lack the relevant information. Our empirical model in this paper shares the same spirit as Mulligan and Sala-i-Martin (1992), but our estimation strategy is now armed with an explicit model of individual behavior and its implications for the behavior of the monetary aggregates and aggregate income discussed in Fujiki and Mulligan (1996a). We find stable cross-sectional money demand functions from the Japanese data. In particular, our best estimates of income elasticity of a counterpart of M_2 minus currency are in the range 1.2–1.4. Third, we make some additional progress on some empirical questions which were left unanswered by Mulligan and Sala-i-Martin (1992). Those include regional differences in purchasing power, a distinction between the effects of the supply of banking services and the effects of the demand for money, estimation of an interest rate elasticity from macro time series data constraining income elasticity at the value of our stable cross-sectional income elasticity, and the assessment of financial sophistication from the time effect estimates on the cross-sectional money demand function.

The organization of this paper is as follows. Section II explains our theoretical model and the reason why a cross-sectional regression of money demand is useful to recover some of the parameters of our structural model. Section III discusses our data. Section IV reports our regression results. Section V discusses our procedure for recovering an interest rate elasticity and degrees of financial sophistication. Section VI summarizes the results and suggests how knowledge of various structural parameters might be relevant for monetary policy.

II. The Production Model of Money Demand and Identification of Parameters from Aggregate Data

Here we begin with a parametric model for production by households and firms which is discussed in Section V of Fujiki and Mulligan (1996a). Useful money demand functions are derived for both types of agents. It is then shown how some of the structural parameters (i.e., parameters of the production/utility functions) can be identified from aggregate data. A brief review of previous empirical studies of Japanese money demand and their problems completes this section.

A. Parametric Model for Households and Firms

Suppose that an agent i produces his final output y_{it} at date t using an input $x_{1,it}$ as well as the quantity of transactions services T_{it} following the technology (all of the Greek parameters are positive constants):[1,2]

$$y_{it} = f(x_{1,it}, T_{it}, \lambda_f) = \left[(1 - \lambda_f) x_{1,it}^{\frac{\gamma-\beta}{\gamma}} + \lambda_f \left(\frac{\gamma - \beta}{\gamma - 1} \right) T_{it}^{\frac{\gamma-1}{\gamma}} \right]^{\frac{\gamma}{\gamma-\beta}} \tag{1}$$

$$\lambda_f \in (0, 1), \beta > 0, \gamma \in (0, \min(1, \beta))$$

$$T_{it} = \phi(m_{it}, x_{3,it} A_{it}) = A_{it} \left[(1 - \lambda_\phi) m_{it}^{(\psi_\phi - 1)/\psi_\phi} + \lambda_\phi x_{3,it}^{(\psi_f - 1)/\psi_f} \right]^{\psi_\phi/(\psi_\phi - 1)} \tag{2}$$

where m_{it} denotes real money balances, and $x_{3,it}$ denotes an input used in the production of transactions services. The production of transactions services is a CES function, with λ_ϕ and λ_f in the interval (0, 1). Transactions services are not, however, aggregated with x_1 in a homogeneous way. Notice that the exponent on the first term is $(\gamma - \beta)/\gamma$, whereas the exponent on the second term is $(\gamma - 1)/\gamma$. We will show that scale elasticities will differ from one when these two exponents differ (i.e., when $\beta \neq 1$).

Agent i's choices of money m_{it} and other inputs $x_{k,it}$ (k = 1, 2) for period t minimize the rental cost r_{it} of producing output y_{it}, where cost is:

$$r_{it} = q_{1,t} x_{1,it} + R_t m_t + q_{3,t} x_{3,it} \tag{3}$$

where $q_{k,t}$ is date t rental rates of the k-th input and R_t is the nominal interest rate at date t. Money is "rented" at a rental price equal to the nominal interest rate. This formulation can be justified on the grounds that there exists an alternative asset

1. For γ = 1, log T replaces the power function of T. For $\gamma = \beta$, log replaces the power function whose argument is a term in small square brackets to the $\psi_f/(\psi_f - 1)$ power. For ψ_ϕ or ($\gamma = \beta$) equal to 1, the corresponding CES aggregator is replaced by a Cobb-Douglas aggregator, with exponents λ_f, $(1 - \lambda_f)$ or λ_ϕ, $(1 - \lambda_\phi)$.
2. It is straightforward to allow for more nonmonetary inputs without changing the implications that are derived below. For example, one could replace x_1 with a homogeneous function of several inputs. Instead of representing a single rental rate, q_1 is interpreted as a price index for the rental rates of the several inputs.

which pays interest (in currency units) at rate R_t but does not enter into the production of transactions services.

The minimum cost achieved is a function of the production level y_{it} and the prices R_t and $q_t = (q_{1t}, q_{3t})$. This cost function, familiar from standard microeconomic theory (e.g., Deaton and Muellbauer [1980]), will be denoted $\Omega\ (y_{it}, R_t, q_t, A_{it}, \lambda_f)$:

$$\Omega\ (y_{it}, R_t, q_t, A_{it}, \lambda_f) \equiv \min_{X_{it},\, m_{it}} [q_t' X_{it} + R_t m_{it}]$$
$$\text{s.t. (1) and (2)}$$

where $X_{it} = (x_{1,it}, x_{3,it})$. The cost-minimizing choices of money and other inputs are functions of output y_{it}, the nominal interest rate R_t, the rental rates of the other inputs q_t, and the level of financial sophistication A_{it}. The Hicksian or derived demand for m_{it} is what we will call the derived demand for money:

$$m_{it} = L(y_{it}, R_t, q_t, A_{it}) = \frac{\partial \Omega(y_{it}, R_t, q_t, A_{it})}{\partial R_t}. \tag{4}$$

The second equality follows from Shephard's Lemma.[3]

In the case of households, y might correspond to "household production," which is observed neither at the micro or macro levels. To think about alternative scale variables, we define a Marshallian money demand:

$$m_{it} = M(r_{it}, R_t, q_t, A_{it}) \equiv L[\Omega^{-1}(r_{it}, R_t, q_t, A_{it}), R_t, q_t, A_{it}]. \tag{5}$$

Maintaining the analogy with standard microeconomic theory, we compute the Marshallian money demand function in two steps. First, the cost function Ω is inverted in order to obtain an "indirect production function" y as a function of r, R, q, and A. Second, the indirect production function Ω^{-1} is substituted into the derived money demand function to obtain the Marshallian money demand M.

Fujiki and Mulligan (1996a) obtained the results that, under certain assumptions, we can derive log-linear approximations to these two types of money demand functions if the production functions are (1) and (2):

$$\log m_{it} = \log L(y_{it}, R_t, q_{it}, A_t) \approx \beta \log y_{it} - \gamma \log R_t +$$

$$\pi_\phi\ (\psi_\phi - \gamma) \log \frac{q_{3,it}}{R_t} + \gamma \log q_{1,it} - (1-\gamma) \log A_{it} + (\text{constant}) \tag{6}$$

3. The first and second derivatives of the cost function with respect to (q, R) exist almost everywhere.

$$\log m_{it} = \log M(r_{it}, R_t, q_{it}, A_t) \approx \beta \log r_{it} - \gamma \log R_t +$$

$$\pi_\phi (\psi_\phi - \gamma) \log \frac{q_{3,it}}{R_t} + (\gamma - \beta) \log q_{1,it} - (1-\gamma) \log A_{it} + (\text{constant}) \qquad (7)$$

where rental rates of the inputs other than money have been subscripted by i to allow for different agents to use different inputs.

For those agents that are households, r is equal to income which will be denoted I. For firms, the level of production y can be interpreted as sales. It will be assumed that income, rental rates, and technology are lognormally distributed across households and that sales, rental rates, and technology are lognormally distributed across firms. In particular, we suppose for households: $\log I_{it} \sim N[\mu_{i,t}(h), \sigma^2_{it}(h)]$, $\log q_{j,it} \sim N[\mu_{j,t}(h), \sigma^2_{jt}(h)]$ where $j = 1, 3$, and $\log A_{it} \sim N[\mu_{A,t}(h), \sigma^2_{At}(h)]$ and for firms: $\log y_{it} \sim N[\mu_{y,t}(f), \sigma^2_{yt}(f)]$,

$$\log q_{j,it} \sim N[\mu_{j,t}(f), \sigma^2_{jt}(f)] \text{ where } j = 1, 3, \log A_{it} \sim N[\mu_{A,t}(f), \sigma^2_{At}(f)].$$

B. National Aggregates

Here we consider aggregation of the derived money demand functions of firms and then the aggregation of the Marshallian demand functions of households.[4] Later we obtain national aggregates by adding up two aggregate money demand function.

Let $N_t(f)$ and $N_t(h)$ denote the number of firms and households in the economy at date t, respectively. $y_t(f)$ and $m_t(f)$ are the average sales and real money balances of firms at date t (i.e., the sum of sales and money balances divided by the number of firms). $I_t(h)$ and $m_t(h)$ are date t average household income and real money balances. Using some properties of the lognormal distribution, we arrive at two aggregate money demand functions from (6) and (7): $m_t(f)$ for firms and $m_t(h)$ for households. Afterward, we derive macro money demand functions in terms of money and income *per capita*, by keeping track of the number of firms and households per capita. Define N_t to be the size of the population at date t. $\eta_t(f) \equiv N_t(f)/N_t$ and $\eta_t(h) \equiv N_t(h)/N_t$ denote the number of firms and households per capita. Let ν_t denote aggregate sales as a fraction of aggregate household income, i.e., $\nu_t \equiv [N_t(f)/N_t(h)]/[Y_t(f)/I_t(h)]$. Using the aggregate firm money demands and the aggregate household money demands from the previous subsection—together with a loglinear approximation of $\log(m_t(f) + m_t(h))$—we derive an expression (10) for real money balances per capita:

4. We consider the firm's derived demand function—as opposed to Marshallian demand—for two reasons. First, the derived demand (8) follows the empirical literature by relating money balances to sales of the firm. Second, because sales is the scale variable, firms' derived demand can be readily combined with households' Marshallian demands to arrive at a national money demand equation that resembles those found in the macro literature. The derived demand for households, on the other hand, is not as useful because household production is unobserved. Fortunately, the two types of money demand functions have some similarities; the similarities can be exploited to derive aggregate relationships that are functions of production parameters such as β and γ.

$$\log\left(\frac{M_t}{P_t N_t}\right) = \beta \log y_t(h) - \gamma \log R_t + \pi_\phi(\psi_\phi - \gamma)\left[\omega \log \frac{q_{3,t}(f)}{R_t} +\right.$$

$$\left.(1-\omega)\log \frac{q_{3,t}(h)}{R_t}\right] + \omega\gamma \log q_{1,t}(f) + (1-\omega)(\gamma - \beta)\log q_{1,t}(h) -$$

$$(1-\gamma)\Big[\omega \log A_t(f) + (1-\omega)\log A_t(h)\Big] +$$

$$\omega \log \eta_t(f) + (1-\omega)\log \eta_t(h) + \beta\omega\left[\log \nu_t + \log \frac{\eta_t(h)}{\eta_t(f)}\right] +$$

$$\frac{1}{2}\beta(\beta-1)[\omega\, \sigma^2_{yt}(f) + (1-\omega)\, \sigma^2_{lt}(h)] +$$

$$\text{(other covariances weighted by } \omega,\ 1-\omega). \qquad (8)$$

Beginning with the first four terms (the first line) of equation (8), we see that, like its micro counterparts, the per capita demand for money depends on average household income, the nominal interest rate, and the ratio q_3/R with elasticities β, $-\gamma$ and $\pi_\phi(\psi_\phi - \gamma)$. When the average price q_3 is different for households and firms, however, the geometric mean of the two q_3/R ratios (one for firms, one for households) enters the aggregate equation. The weight ω can be approximated by the share of the money stock held by firms (as opposed to households).[5] Terms reflecting averages of the price of x_1 and the level of financial technology enter the aggregate money demand equation separately for firms and households. Per capita money demand also depends on the number of firms and households per capita as well as the ratio of aggregate sales to household income (ν_t). These three terms, roughly speaking, represent the degree of vertical integration in the economy. The more stages involved in the production process, the greater the demand for money. This vertical integration result follows from the assumption that *a firm* and *a household* are the demanders of money. Economies of scale ($\beta < 1$) cannot be exploited by pooling money holdings across firms or across households while diseconomies of scale ($\beta > 1$) cannot be avoided by subdividing money holdings within the firm or within the household. Finally, for given average sales and average income, the dispersion of income and sales across agents affects aggregate money demand to the extent that there are economies (or diseconomies) of scale in the holding of money.

C. Identification of Parameters from Aggregate Data

The aggregate money demand equation (8) indicates that, with enough data, one could obtain consistent estimates of some of the structural parameters of the model such as β, γ, and perhaps ψ_ϕ. The scale elasticity is interesting for economic theory as various models of the demand for money differ on the presence and extent of scale economies. Faig (1988) shows that β is also relevant for evaluating the efficiency of

5. ω derives from an approximation to $\log[m_t(f) + m_t(h)]$.

the inflation tax relative to other distortionary taxes. γ and ψ_ϕ reflect the own price elasticity of money demand and are therefore indicative of the welfare cost of inflation and relevant for computing the optimal monetary policy (see Lucas [1994] for a discussion and for references). However, estimation of (8) requires that (i) one has a time series on the prices q_1 and q_3 and the level of financial sophistication *or* (ii) all cross-price elasticities are zero or (iii) the cross-prices and the level of financial technology are uncorrelated with household income and the nominal interest rate. Condition (iii) is certainly violated if we estimate equation (8) in levels. Financial technology has grown over time as has household income. Or, if one prefers to think of financial sophistication as endogenous, the rental rate of financial technology (which might be modeled as q_3 in our setup), such as the computer, has fallen over time. One solution to this problem might be to estimate (10) in differences. Perhaps high-frequency movements in income are not associated with high-frequency movements in financial technology. However, the same might not be true for short-term movements in the nominal interest rate. We can imagine that economy-wide stocks of financial technology (which we might model as the good x_3) are fixed in the short run. A rapid increase in the nominal interest rate will increase the demand for the technology which, because stocks are fixed, must result in an increase (but less than proportional) in the rental rate q_3. In other words, q_3 will be correlated with R at high frequencies.[6]

This has led some studies to use cross-sections of regional aggregates to identify the production parameter β.[7] The idea is that, within Japan, all agents have fairly equal access to financial technology. Thus it is assumed that the exogenous level of financial technology A, the rental price of financial machines (q_3), and the nominal interest rate are all constant in a cross-section of regions. γ and ψ_ϕ can then be estimated in levels using aggregate time series data by imposing the condition that β correspond to its estimate from the cross-sections. We expect consistent estimates as long as q_3 and A are uncorrelated with R in the time series. The basic specification of money demand function used in this paper is:

$$\ln\left(\frac{\text{money}}{\text{CPI}}\right)_{it} = b0 \cdot (T) + b1 \cdot \ln\left(\frac{\text{income}}{\text{CPI}}\right)_{it} + b2 \cdot Z_{it} + \varepsilon_{it} \qquad (9)$$

where T is a time effect, Z shows the vector of prefectural variables discussed later, $b0$, $b1$, and $b2$ are the vector of coefficients to be estimated, subscript i means prefecture, and subscript t means time. Both money and income are normalized by the population.[8] Readers may complain that our cross-sectional approach cannot recover interest rate elasticity, and therefore misses one of the most interesting structural parameters. However, under the assumptions listed above, we can recover the interest

6. This problem may also occur with seasonal data because the stock of machines such as computers may not vary across seasons, but the demand for their services might.
7. Mulligan and Sala-i-Martin (1992) are one example.
8. The timing of the usage of each variable in 1970 is as follows. Income is from fiscal 1970, money data are as of March 1970, the CPI is the average for 1970, the population is as of October 1, 1970 and the national deflator is the mean of regional CPI in 1970.

rate elasticity from macro data, after controlling for the value of income elasticity obtained from cross-sectional data.

Before moving on the discussion of our prefectural data, we review former literature of empirical investigations of Japanese money demand, and the problems inherent in these studies.

D. Review of the Literature

To the best of our knowledge, the study of Japanese money demand has always used time series data. Yoshida (1990) summarized the history of conventional studies of money demand, fitting the error correction model to Japanese quarterly macro data. Yoshida and Rasche (1990) estimated a vector error correction model of Japanese M_2+CD demand (with quarterly data). They found that the equilibrium real income elasticity of M_2+CDs was about 1.2 throughout the period 1956/I–1985/II. It is well known that the deregulation of Japanese money markets started in the middle of 1985, and Yoshida and Rasche investigated whether the equilibrium income elasticities had changed after 1985 due to the deregulation of interest rates. They added a time dummy variable that took on the value of one after 1985/III and took zero before 1985/III to their vector error correction model, and found that the time dummy absorbed all the effect of the deregulation of interest rates, and the rest of the parameters of the money demand function were unchanged from the parameters estimated from the data 1956/I–1985/II. Therefore, they claimed that their equilibrium income elasticity was stable even after adding the observations from 1985/III to 1989/II. Rasche (1990) estimated a vector error correction model of Japanese M_1 demand (with quarterly data). He advocated estimates of 1.0 for the long-run income elasticity and –0.5 to –0.6 for the interest rate elasticity.

Japanese conventional studies on money demand have the following drawbacks due to the fact that they have used time series data. First, there is no agreement about the relevant interest rate. Second, due to the problem of the financial innovation, it is difficult to settle on a consistent definition of money over long periods of time. Third, the problem of the stability of money demand function has never been resolved. For the sake of the demonstration of those problems, we display some regression estimates of the income elasticity of money demand. For this illustration, we follow the tradition of using quarterly data on real M_2+CDs deflated by the GNP deflator, real GNP, the overnight call rate, and the annual average of the interest rate of interest-bearing bank debentures (with a maturity of five years) to estimate a money demand function. The equation to be estimated is:

$$\begin{aligned}\ln(\text{real } M_2\text{+CDs})_t = b0 + b1 \cdot \ln(\text{real GNP})_t \\ + b2 \cdot \ln(\text{interest rate})_t + \varepsilon_t \end{aligned} \qquad (10)$$

where M_2+CDs is the average of the amount of M_2+CDs within the quarter. The results of level estimation in Table 1 suggest that the income elasticities are 1.2 to 1.5, and that short-term interest rate elasticities are something like –0.03 and long-term interest rate elasticities are around –0.07. However, unusually low values for Durbin-Watson statistics suggest a misspecification of these regressions. For

Table 1 Time Series Estimates of the Income Elasticity

Equation form	Income elasticity	Interest elasticity	Adj. R-2 (D.W.)
log level, call rate	1.476 (0.018)	–0.039 (0.012)	0.9932 (0.164)
log level, bond rate	1.454 (0.022)	–0.079 (0.022)	0.9935 (0.176)
	1.4929 (0.016)		0.9927 (0.153)
log level with trend call rate	1.224 (0.125)	–0.033 (0.014)	0.9937 (0.145)
log level with trend bond rate	1.259 (0.131)	–0.061 (0.026)	0.9937 (0.156)
	1.196 (0.126)		0.9933 (0.1357)
log, differenced call rate, no constant	1.011 (0.110)	–2.247 (1.125)	(1.403)
log, differenced bond rate, no constant	1.017 (0.110)	–2.376 (1.811)	(1.437)
	1.012 (0.112)		(1.394)
log, differenced call rate, with constant	0.5871 (0.1359)	–1.878 (1.107)	0.1996 (1.017)
log, differenced bond rate, with constant	0.5912 (0.1399)	–1.111 (1.641)	0.1818 (0.954)
	0.5784 (0.136)		0.1860 (0.947)

Note: Numbers in parentheses are standard error of the estimates following White.
Estimated sample periods are 1967/III–1993/I.
Data: M_2+CDs, seasonally adjusted, average stock within quarter; GNP, seasonally adjusted; call rate, overnight, collateralized, average within periods, median of offer rate; bond rate; and annual yield of five-year interest-bearing bank debentures.

example, once first differenced data are used, we see dramatic changes: income elasticities fall to 1.0 (without constant term) or 0.6 (with constant term). On the other hand, the absolute value of interest rate elasticities increases to 2.2–2.3 (without constant term) and 1.8–1.1 (with constant term).

We also estimate the long-run income elasticity of money demand by recovering the cointegrating vector of ln(real M_2+CDs) and real income. To this end, we first figure out the existence of the unit root for ln(real M_2+CDs) and ln(real GNP), and later find the cointegration relation between those two variables. Using various cointegration regression techniques, we estimate the long-run income elasticity of money demand to be 1.5–1.6 for the sample period of 1967/II–1993/I. It will be quite interesting to compare those results to our cross-sectional income elasticity.[9]

9. For details, see Fujiki and Mulligan (1996b).

III. Data

We have compiled four series of estimates of prefectural money stock for 46 prefectures and three series of scale variables to estimate a money demand function from cross-sectional data.[10] This section reviews the definition of our measure of prefectural money aggregates, prefectural income statistics, and various conditioning variables used later in the analysis.[11]

A. Prefectural Money Aggregates

We compute two types of prefectural money stocks, MF_1 and MF_2, from deposit survey data. Prefectural deposit statistics are available from 1955 to 1990 (surveyed at the end of March). Since these statistics are based on a different survey from the money supply statistics, we do our best to make the deposit data comparable with current Japanese money supply statistics. First, MF_1 is our counterpart of M_1 minus currency. Due to the limited availability of a breakdown by prefecture of demand deposits versus savings deposits, MF_1 is defined as the demand deposits held by individuals and corporations at major Japanese banks. This series is available from 1960 to 1990.[12] Second, MF_2 is our counterpart of M_2 minus currency. The definition of MF_2 is the sum of the deposits at major Japanese banks and community and rural banks. MF_2 consists of both demand deposits and savings deposits. This series is available between 1955 and 1990.

B. Prefectural Income Statistics

The Research Institute of Economy at Japan's Economic Planning Agency publishes the prefectural accounts, which are the prefectural version of the GNP statistics. Prefectural income is the prefectural counterpart of GNP. Because our measure of money includes deposits held by corporations, we prefer a measure of income that includes the activity of corporations. To this end, we use prefectural income.

C. Conditioning Variables

Our theoretical model in Fujiki and Mulligan (1996a) suggests that we need to control the level of technology for each prefecture in order to estimate money demand function cross-sectionally. To this end, we introduce three conditioning variables following Mulligan and Sala-i-Martin (1992): population density, the percentage of net prefectural product explained by the primary industry, and prefectural fixed effect.[13] Even after controlling for these variables, the error term of the money demand function might be correlated with the level of income. Hence, we prepared the ratio of job offers to applicants for the instrumental variable of real income. We will discuss such variables in turn.

10. We excluded Okinawa Prefecture.

11. A detailed explanation of the data set can be found in Section III of Fujiki and Mulligan (1996b).

12. Note that due to mergers and to the eventual inclusion in All Banks of some banks that were initially excluded, MF_1 has discontinuities in 1968, 1976, 1983, 1984, and 1989.

13. Mulligan and Sala-i-Martin (1992) discussed the reason why those conditioning variables are relevant for our money demand function.

First, we compute the population density in order to take differences in the industry structures among regions into account. We will call this variable *PD* hereafter. These data use the population of each prefecture as of the beginning of October of each year.

Second, we compute the percentage share of primary industry (agriculture, forestry, and fisheries) of the prefecture's net product to trace the differences in financial technology between rural area and urban areas. We will call this variable *SPI* hereafter.

Third, we introduced prefectural fixed effect. Mulligan and Sala-i-Martin (1992) claim that if the financial sophistication, the structure of banking industries, and the differences in the levels of prices in regions are fairly persistent, then prefectural fixed effects should serve adequately as a proxy for them. In the case of Japan, since the structure of the banking industry has been fairly stable due to the regulation of the allocation of branches by the Ministry of Finance, the introduction of prefectural fixed effect might make sense.

Finally, we introduced the ratio of job offers to applicants. The Japanese public employment security office publishes the data for the number of job applicants and the number of job offers at the office. The job offers made at the public employment office are valid for two consecutive months. We use the ratio of these two variables in some of our analysis. The ratio of job offers to applicants is one of the important indicators of labor market conditions, hence it is highly correlated with business conditions and correlated with short-run fluctuations of income. Note that most of the job offers recorded in the public employment office are made by small businesses, so the ratio does not fully capture the labor market situation. Banks are unlikely to make job offers at the public employment office, therefore our ratio variable probably does not reflect regional differences in the financial sophistication of the banking sector. If the error term of money demand reflects technology shocks in the banking sector, then our instruments are also likely to be uncorrelated with these kinds of shocks. Hence the ratio can be used as an instrument for prefectural income in a money demand equation.

D. Regional Consumer Price Difference Index

We employ the regional consumer price difference index (1960–90, calendar year average) and the regional retail price difference index (1955–59, calendar year), which measure the cross-prefectural differences in the level of retail prices in the capital cities of each prefecture at one point in time. Using a time series of the CPI index of Tokyo (1934–1936 average base, calendar year), we construct the deflator for each prefecture in every year.

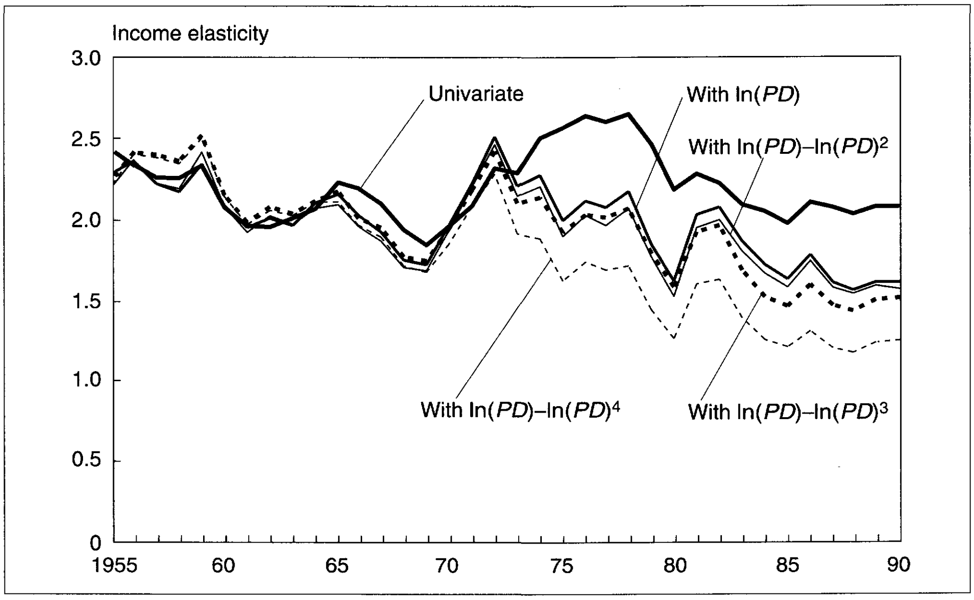
IV. Results of Regressions

This section reports the results of regressions using MF_2 as the regressor.[14]

A. Univariate Regressions

We begin with cross-sectional estimates of the elasticity of money demand to prefectural income without adding any other conditioning variables. The thick, solid line in Figure 1 shows the estimated income elasticities in each fiscal year using ln(prefectural MF_2 per capita/prefectural CPI) for the dependent variable and ln(prefectural income per capita/prefectural CPI) for the independent variable. As

Figure 1 Cross-Sectional Prefectural Income Elasticity of MF_2



can be seen in Table 2, all estimates differ from zero to a statistically significant degree, and we find that the income elasticities are well above one and moving around 1.9 to 2.6. When we pool whole sample data and add time effects, the constrained income elasticity is 2.15 (s.e. = 0.032), and we cannot reject the null hypothesis of a temporally constant income elasticity. Compared with the results of Mulligan and Sala-i-Martin (1992) (see p. 322, Figure 9), these estimates of income elasticities from Japanese data are higher than those of the U.S. counterparts.

14. Fujiki and Mulligan (1996a) report the results of regressions using MF_1 as the regressor. The estimated value of income elasticities of MF_1 and MF_2 become almost equivalent when *SPI* and ln(*PD*) are added as the conditioning variables.

Table 2 Regression Results (Univariate Regression)

Sample	Coefficient on ln(income)	R-2
1955	2.424 (0.171)	0.819
60	2.073 (0.140)	0.832
65	2.227 (0.192)	0.753
70	1.958 (0.143)	0.808
75	2.565 (0.309)	0.609
80	2.182 (0.302)	0.542
85	1.973 (0.265)	0.556
90	2.076 (0.256)	0.599
55–90	2.155 (0.032)	0.947

Note: Numbers in parentheses are standard error of estimators. All regressions include constant term or time effect, which are not reported here. F-value for the test of constancy of income elasticity = 0.952. Figure 1 plots above estimators.

B. Adding Population Density

The four other lines in Figure 1 show the estimated income elasticities holding polynomials of ln(*PD*) constant. Figure 1 shows that the introduction of ln(*PD*) reduces the level of income elasticities of money demand after 1974 compared to the case without ln(*PD*). The estimated coefficients of the polynomials of ln(*PD*) are not significantly different from zero for the case of including ln(*PD*) alone and ln(*PD*) and $\ln(PD)^2$. However, the inclusion of $\ln(PD)^3$ and $\ln(PD)^4$ yields statistically significant coefficients of the polynomials of ln(*PD*). The relationship between real MF_2 per capita and the polynomials of ln(*PD*) is nonlinear. Table 3 summarizes the long-run pooling regressions holding time effect, income, and the polynomials of ln(*PD*). We cannot reject the null hypothesis of a constant income elasticity throughout the sample periods compared to the cross-sectional regression. We might guess that the income elasticity of money demand is about 1.9–2.0 holding ln(*PD*) constant.

Table 3 The Results of Pooling Regressions Holding Constant Polynomials of ln(*PD*)

Dependent variable ln(MF2/CPI) per capita

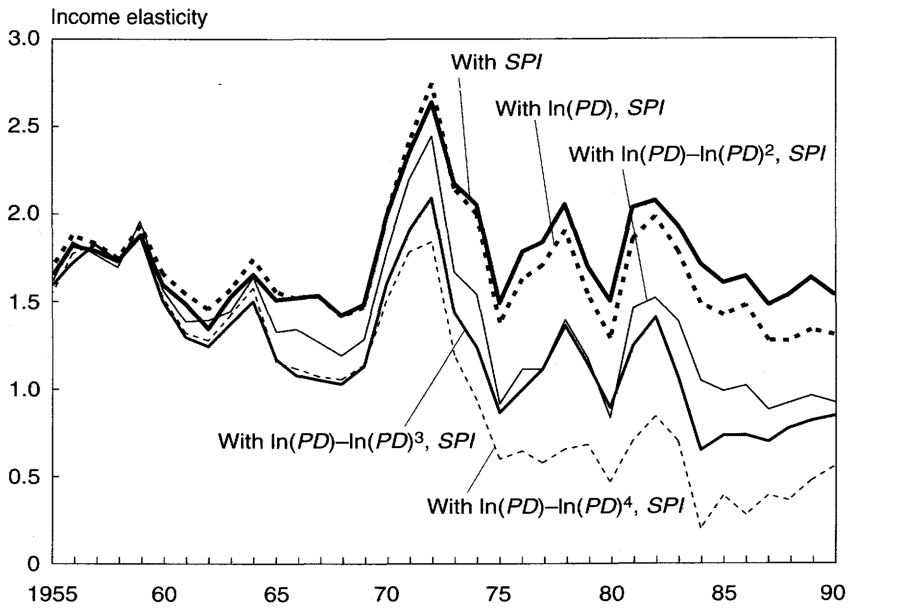
List of explanatory variables					
Income	**ln(*PD*)**	**ln(*PD*)^2**	**ln(*PD*)^3**	**ln(*PD*)^4**	**R-2 (F)**
2.155 (0.032)					0.947 (0.95)
1.977 (0.046)	0.043 (0.008)				0.948 (0.69)
1.932 (0.046)	0.080 (0.010)	0.026 (0.004)			0.949 (0.52)
1.985 (0.058)	−0.016 (0.011)	0.110 (0.007)	0.047 (0.003)		0.955 (0.55)
1.916 (0.043)	−0.193 (0.019)	0.049 (0.008)	0.132 (0.008)	0.033 (0.002)	0.958 (0.67)

Note: Sample periods are 1955–90. Numbers in parentheses are standard error of estimators. (F) shows the F-value of the null hypothesis that the cross-sectional income elasticities are constant over the sample periods. These F-values are too small to reject the null hypothesis of constant income elasticity over 1955–90.

C. Adding the Share of Primary Industry to Prefectural Net Product

We add the share of primary industry to prefectural net product (*SPI* hereafter) to the explanatory variables. Figure 2 shows the cross-sectional income elasticity holding

Figure 2 Cross-Sectional Prefectural Income Elasticity of MF2



SPI and polynomials of ln(*PD*). The solid line in Figure 2 shows the income elasticities conditioned on prefectural income and *SPI*. The coefficients of *SPI* were negative except for the sample periods of 1970–72, and income elasticities conditioned on *SPI* take generally smaller values than those of univariate regressions, apart from the 1970–72 periods. (For these three years, *SPI* is not significantly different from zero.) Observe that compared to the results reported in Figure 2, the use of *SPI* and polynomials of ln(*PD*) together as the conditioning variables generally make the income elasticities smaller than those of holding the polynomials of ln(*PD*) only, except for the periods of the early 1970s. However, note that the addition of *SPI* for the regressions holding ln(*PD*), $\ln(PD)^2$, $\ln(PD)^3$ and $\ln(PD)^4$ yields an income elasticity of money demand that is not significantly different from zero after 1974. Hence, ignoring the results obtained from the sample period 1970–72, and the results after 1974 for the regressions holding *SPI*, ln(*PD*), $\ln(PD)^2$, $\ln(PD)^3$ and $\ln(PD)^4$, we see that our income elasticities swing around 0.8–2.0. To construct a plausible range of our income elasticities, we conduct long-run pooling regressions. The results of long-run pooling regressions using the entire sample holding *SPI*, polynomials of ln(*PD*), and the time effect constant are summarized in Table 4. All the pooling regressions are justified in a sense that we cannot reject the null hypothesis of constancy of cross-sectional income elasticity over time as can be seen in the F-values in Table 4. We conclude that our income elasticities are something like 1.2–1.4 judging from the results reported in Table 4. Surprisingly, our findings are consistent with the time series vector error correction model by Yoshida and Rasche (1990), which found a stable income elasticity of 1.2, which is significantly larger than one based on the results of estimation using 1956–1985 quarterly data.

Table 4 The Results of Pooling Regressions Holding Constant SPI and Polynomials of ln(*PD*)

Dependent variable $\ln(MF_2/CPI)$ per capita

List of explanatory variables						
Income	***SPI***	**ln(*PD*)**	**ln(*PD*)^2**	**ln(*PD*)^3**	**ln(*PD*)^4**	**R-2 (F)**
1.678 (0.057)	–0.014 (0.001)					0.950 (0.63)
1.597 (0.061)	–0.013 (0.001)	0.027 (0.008)				0.951 (0.51)
1.384 (0.063)	–0.017 (0.001)	0.083 (0.009)	0.043 (0.004)			0.954 (0.43)
1.284 (0.058)	–0.023 (0.001)	–0.037 (0.011)	0.015 (0.007)	0.059 (0.003)		0.961 (0.46)
1.222 (0.056)	–0.023 (0.001)	–0.210 (0.018)	0.094 (0.008)	0.142 (0.007)	0.032 (0.002)	0.965 (0.62)

Note: Sample periods are 1955–90. Numbers in parentheses are standard error of estimators. (F) shows the F-value of the null hypothesis that the cross-sectional income elasticities are constant over the sample periods.
These F-values are too small to reject the null hypothesis of constant income elasticity over 1955–90.

In addition to the F-test for the long-run pooling regressions, we report two additional pieces of evidence in order to validate the stability of our cross-sectional income elasticity over time. First, following Yoshida and Rasche (1990), we introduced the time dummy that takes the value one after 1985 and zero before 1985 for the regressions presented in Table 4. We test if this time dummy can capture the effect of the deregulation of interest rates after 1985. None of the regressions reported in Table 4 suggest that the time dummy is statistically significantly different from zero; namely, we did not find the evidence for the upward shift of all of the time dummies after 1985. This is the opposite of the result reported in Yoshida and Rasche (1990). Second, some readers may wonder whether our error terms in the regressions are serially correlated. We therefore conduct an additional analysis of a model:

$$\ln\left(\frac{\text{money}}{\text{CPI}}\right)_{it} = b0 \cdot (T) + b1 \cdot \ln\left(\frac{\text{income}}{\text{CPI}}\right)_{it} + b2 \cdot Z_{it} + \varepsilon_{it},$$
$$\varepsilon_{it} = \rho\varepsilon_{i,t-1} + u_t \tag{11}$$

where Z is SPI, $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, and $\ln(PD)^4$. Since the GLS estimation of parameters requires differencing, the set of time dummies cannot be estimated directly. Hence we regressed each variable on the time effect and used the standard statistical package to estimate equation (11). The results of estimation suggest that the estimate of ρ is 0.02 and that for $b1$ is 1.19 (s.e. = 0.019). Compared to the OLS result of $b1$ = 1.22 holding the same conditioning variables constant, the modification of error term changes our results based on OLS only slightly.[15]

D. Cross-Prefectural Deviations from Purchasing Power Parity

The original specification in Mulligan and Sala-i-Martin (1992) was:

$$\ln(\text{money})_{it} = b0 \cdot (T) + b1 \cdot \ln(\text{income})_{it} + b2 \cdot Z_{it} + \varepsilon_{it}. \tag{12}$$

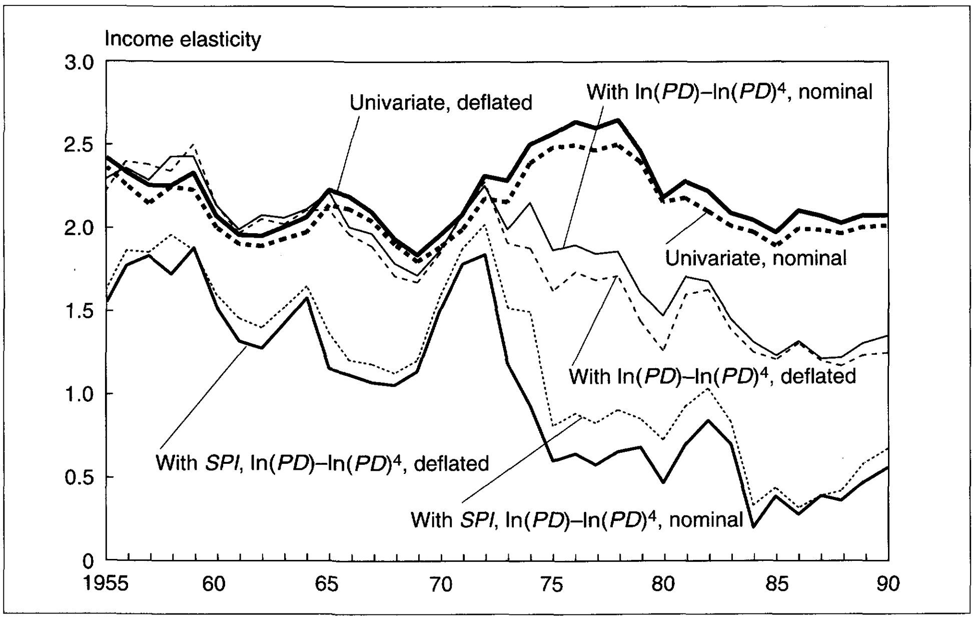
Their specification added one more assumption, namely, that the level of price in each region is the same, or at least orthogonal to the level of income to equation (9). This assumption is unnecessary for our analysis, since we can use a regional consumer price index. According to our computation, the correlation of log of CPI and log of real prefectural income by the cross-sectional estimates is 0.06–0.14, which is statistically different from zero (except for the outlier of 0.03 in 1958 that is not significantly different from zero). Mulligan and Sala-i-Martin (1992) showed that if the partial correlation of real income and price was s, then the coefficient of income, without taking into account the differences in the regional price level, $\hat{\beta}$, and the true income elasticity, β, would be related as:

15. We estimate this model by the PANEL procedure in LIMDEP 6.0.

$$\hat{\beta} = \beta + \frac{(1-\beta)s}{(1+s)}.$$

We know that the bias due to the failure to adjust for differences in the level of regional prices must be relatively small since $s/(1 + s)$ is at most 0.1. Nonetheless, we have made some progress on one of the open questions in Mulligan and Sala-i-Martin (1992). To see this in detail, Figure 3 plots the results of estimation with and without regional CPI. According to our computations, the average ratio of the cross-sectional income elasticity estimated using nominal data to that estimated with real data is 0.963 for the case of univariate regression, 1.042 for the case of holding $\ln(PD)$, $(\ln(PD))^2$, $(\ln(PD))^3$, and $(\ln(PD))^4$ constant, and 1.189 for the case of holding SPI, $\ln(PD)$, $\ln(PD)^2$, $(\ln(PD))^3$, and $(\ln(PD))^4$ constant. Namely, the bias due to the omission of CPI for the case of univariate regression is only about 5.7 percent. As we increase our conditioning variables, the bias increases and is in the opposite direction to the case of univariate regression. Note that as long as we use the aggregate time series deflator, for the case of univariate regression, the aggregate deflator will be absorbed by the constant term. Therefore, our results are valid for all other kinds of aggregate time series deflators.

Figure 3 Cross-Sectional Prefectural Income Elasticity of MF2



V. The Demand for Money vs. the Supply of Banking

The goal of this paper has been to use cross-prefectural data to estimate certain structural parameters, by which we mean parameters of utility or production functions of agents in the economy. The first step in doing so is to aggregate the demand functions of the agents to arrive at a relationship between the aggregate quantity of money in the economy, aggregate income earned by residents of the economy, and other aggregate variables. However, we have attempted to estimate this aggregate relationship using deposits at banks in the prefecture—which need not be identical to money held by residents of the prefecture. Denoting money held by prefecture i residents at date t as $\ln(\text{money})^*_{it}$ and the percentage difference between money held by residents of the prefecture and money deposited in banks of the prefecture as η_{it}, we have an obvious relationship between the variable required by the theory and the measure of money that is available to us:

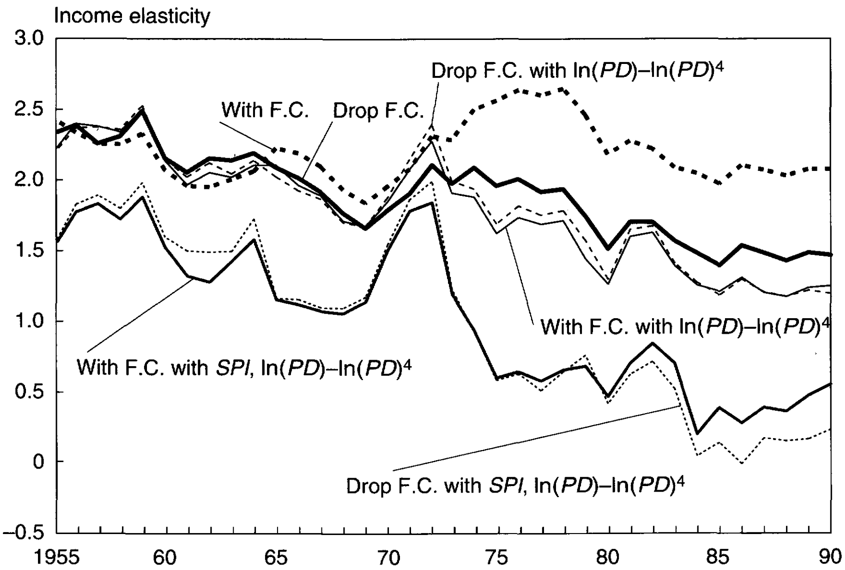
$$\ln(\text{money})_{it} = \ln(\text{money})^*_{it} + \eta_{it}. \tag{13}$$

Of course, our estimates of the income elasticity are still consistent if η_{it} is uncorrelated with prefectural income. However, there are reasons to believe that η_{it} is positively correlated with prefectural income in our 46-prefecture cross-sections. The obvious example is Tokyo. Tokyo specializes in banking and other financial activities, so we expect that many of the deposits in Tokyo banks are not in fact owned by either corporate or household residents of Tokyo. In addition to having relatively large positive values for η_{it}, Tokyo is also the richest prefecture. We propose three ways to purge our estimates of the income elasticity of the effects of the measurement error η_{it}: (1) dropping financial centers from the data set; (2) estimating prefectural fixed effects; and (3) using instrumental variables methods.

A. Dropping Financial Centers

We suspect that Tokyo, Osaka, and Kanagawa are the most important financial centers in Japan. By dropping these three prefectures, we drop three of the richest and three that we suspect have large positive values for η_{it}. Figure 4 displays cross-prefectural estimates of the income elasticity without the three financial centers and the estimates obtained with all 46 prefectures. A univariate regression does show a large difference between the income elasticity with and without financial centers; however, with conditioning variables, the difference between the two types of estimates are small. The pooled estimate of the income elasticity without Tokyo, Osaka, and Kanagawa is 1.91 (s.e. = 0.043) holding $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, and $\ln(PD)^4$ constant and 1.22 (s.e. = 0.056) holding SPI, $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, and $\ln(PD)^4$ constant. Compared with the pooled estimate of 1.92 (s.e. = 0.06) holding $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, and $\ln(PD)^4$ constant and 1.25 (s.e. = 1.25) holding SPI, $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, and $\ln(PD)^4$ constant, it appears that the inclusion of Tokyo, Osaka, and Kanagawa keeps the estimates of the income elasticity with conditioning variables virtually unchanged.

Figure 4 Cross-Sectional Prefectural Income Elasticity of MF2



For two reasons, it is not clear that the income elasticity estimate without the three financial centers is superior to that obtained with all 46 prefectures. First, we might suspect that, among the prefectures that are not "financial centers," richer families and larger businesses are more likely to search across prefectural borders for banking services. Richer prefectures might, for example, be more likely to have citizens and firms whose travels to Tokyo or another financial center facilitate banking there. This is a reason why the error term in a money demand equation might be *negatively* correlated with income per capita. Second, the theory tells us that the level of financial technology decreases the demand for money because more transactions services can be obtained for a given money balance. If superior financial technology exists in the richer financial centers, this is yet another reason why the error term in a money demand equation might be *negatively* correlated with income per capita.

B. Prefectural Fixed Effects

If we suppose that the percentage difference between money held by residents of the prefecture and money deposited in banks of the prefecture, η_{it}, differs across prefectures in a way that might be correlated with income and other variables, but that cross-prefectural differences are constant over time, then the introduction of prefectural fixed effects is the correct way to purge our estimates of the effects of the supply of banking. A pooled regression of log real MF2 per capita on log real income per capita, *SPI*, $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, and $\ln(PD)^4$and sets of year and prefectural dummies, yields an income elasticity of 0.55 (s.e. = 0.032). Unlike our previous estimates, this is dramatically less than one.

It is well known that the fixed effect cleans out the permanent component of unobservables, and the fixed effect estimator picks up short-run correlations between dependent and independent variables (see Mairesse [1990] for such examples). However, the addition of prefectural fixed effect could induce a downward bias to our income elasticity because of the correlation of explanatory variables and transformed residuals if explanatory variables are measured with error. Suppose that our prefectural income has i.i.d measurement errors. This is not such a bad assumption, since the method of compiling the data differs across prefectures. Let us assume that we observe $\ln(\text{income})_{it}$, and the relation with true data $\ln(\text{income})^*_{it}$ is something like:

$$\ln(\text{income})_{it} = \ln(\text{income})^*_{it} + \psi_{it} \tag{14}$$

and ψ_{it} is i.i.d mean zero disturbance. The regression we wish to estimate is equation (15):

$$\ln(\text{money})^*_{it} = b0 + b1 \cdot \ln(\text{income})^*_{it} + b2 \cdot Z_{it} + e_{it}. \tag{15}$$

However, we can run regressions something like equation (16),

$$(\ln(\text{money})^*_{it} + \eta_{it}) = b0 + b1 \cdot (\ln(\text{income})^*_{it} - \psi_{it}) + b2 \cdot Z_{it} + e_{it}. \tag{16}$$

Suppose that η_{it} is constant over time. In this case, even after taking the difference from prefectural times series mean, we found a negative correlation between observable income and measurement error, and a fixed effect approach would bias downward the fixed effect estimator, as Griliches and Hausman (1986) have pointed out. Note that permanent unobservables could also be removed by taking the first difference of variables. Griliches and Hausman took advantage of the relationship between the fixed effect estimator and an estimator based on the first differenced data, and presented a formula to compute a consistent estimator from these two estimator. Following their method, we arrive at a consistent income elasticity estimate of 0.63 (with s.e. = 0.041) holding *SPI*, $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, and $\ln(PD)^4$ constant. Since the income elasticity with fixed effects is 0.55, the downward bias due to measurement error due to the introduction of fixed effect is 14 percent. We get a consistent income elasticity estimate of 0.89 (with s.e. = 0.032) holding $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, and $\ln(PD)^4$ constant. Since the income elasticity holding $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, and $\ln(PD)^4$ with fixed effects is 0.78 (s.e. = 0.032), the downward bias is 13 percent in this case. So far, we do not find convincing evidence against the fixed effect estimators.

C. Instrumental Variables Methods

A third way to purge our estimates of the income elasticity of a measurement error that is correlated with income is to search for a variable that is correlated with income but uncorrelated with a prefecture's propensity to attract a high percentage of its deposits from out-of-prefecture depositors. Moreover, the instrument must be uncorrelated with the error term of money demand functions. The instrument that we propose is the ratio of job offers to the number of job applications. We employ this measure for each prefecture for each of the dates 1962–90. Our instrumental variables estimate (using the job offer ratio, *SPI*, $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, $\ln(PD)^4$, year dummies, and prefectural fixed effects in a first-stage regression for log prefectural income per capita and including *SPI*, $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, $\ln(PD)^4$, and year dummies as second-stage regressors) of the income elasticity is 1.47 (s.e. = 0.18). When we exclude prefectural fixed effects for both first- and second-stage regression, estimated income elasticities by instrumental variable methods is 1.07 (s.e. = 0.19). These can be compared with our OLS estimate (holding *SPI*, $\ln(PD)$, $\ln(PD)^2$, $\ln(PD)^3$, $\ln(PD)^4$, and year dummies constant) of 1.04 (s.e. = 0.066) for the same sample period (1962–90).

Throughout the exercise done in this section, we conclude that our mismeasurement of deposits held by prefecture residents cannot explain our finding of an income elasticity that is greater than one. The only way to pull down income elasticities far below one is the introduction of prefectural fixed effects, and so far we do not have a good explanation of why the income elasticity obtained from the regression with fixed effects is so low.

VI. Extensions

We have documented the stability of cross-sectional income elasticities in sections IV and V. We now discuss how we can benefit from our stable cross-sectional estimates. We show two applications.

A. Recovering the Interest Elasticity

Lucas (1988) argued that in a cash-in-advance economy, with homothetic preferences, the equilibrium condition gives the same form of the standard money demand function which used consumption instead of income and the scale variable elasticity was one. Poole (1988) expressed his dissatisfaction with time series estimates of income elasticity, and he assumed an income elasticity of one. Although Poole (1988) and Lucas (1988) assumed that the income elasticity of M_2 was one, we can use the income elasticity estimated from the cross-sectional data, since Fujiki and Mulligan (1996a) suggest that cross-sectional estimates are valid for aggregate per capita time series. Define:

$$\begin{aligned}\text{Lucas residual} &= \ln(M_2 \text{ per capita/CPI}) \\ &\quad - \varepsilon_y \cdot \ln(\text{GNP per capita/CPI}) \qquad (17)\end{aligned}$$

where ε_y is income elasticity of money demand. We see in Table 4 that our best estimates of ε_y are between 1.2 and 1.4, and we know that some part of the Lucas residual can be explained by interest rate movements. Let us recover the interest rate elasticity by regressing the Lucas residual on an interest rate. More specifically, we use a long-term bond rate, the annual yield on five-year interest-bearing bank debentures, to make our results comparable to the results in Table 1. We pick two income elasticities, 1.384 (holding *SPI*, ln(*PD*), and ln$(PD)^2$ constant) and 1.222 (holding *SPI*, ln(*PD*), ln$(PD)^2$, ln$(PD)^3$, and ln$(PD)^4$ constant) from Table 3. Table 5 shows the results of regressing the Lucas residual on a long-term interest rate. Since some readers might suspect the plausibility of the effect of interest rate movements

Table 5 The Results of Estimated Interest Rate Elasticity

Sample periods	$\varepsilon_y = 1.384$	$\varepsilon_y = 1.222$
1956–90	–0.161 (0.074)	–0.374 (0.075)
1975–90	–0.199 (0.051)	–0.282 (0.068)
1980–90	–0.201 (0.060)	–0.265 (0.079)

Note: Numbers in parentheses are standard error of estimators.
Estimated equations are [ln(MF_2) – ε_y ln(income)] on constant and ln(interest rate).

on money demand in early periods of our sample due to the interest rate regulation in Japan, we report our results of estimation using the entire 1955–90 sample, 1975–90 sample, and 1980–90 sample. Table 5 suggests that the magnitude of our estimates of interest rate elasticities is more than 0.2, and takes on the value of –0.2 to –0.3 if we drop the samples of earlier days. These estimates have larger magnitudes than the time series estimates obtained from the level regressions in Table 1.

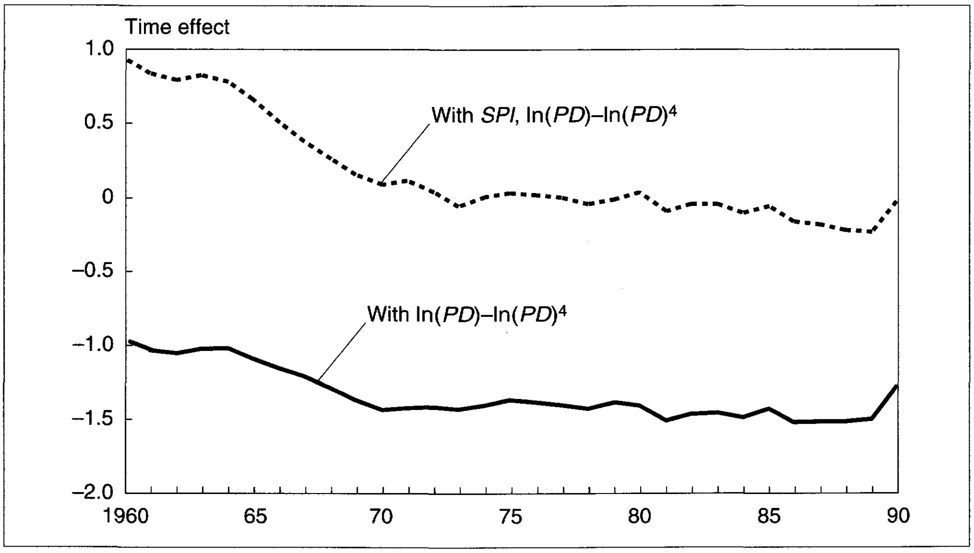
B. Recovering Technical Progress from Long-Run Pooling Regressions

The other interesting exercise is to recover the technical progress of cash management technology. Let us think about a simple model of transaction technology of prefecture *i* at time *t*:

$$(\text{Money stock}_{it}) = \alpha_t\, (\text{scale variable}_{it})^{\beta_t}. \tag{18}$$

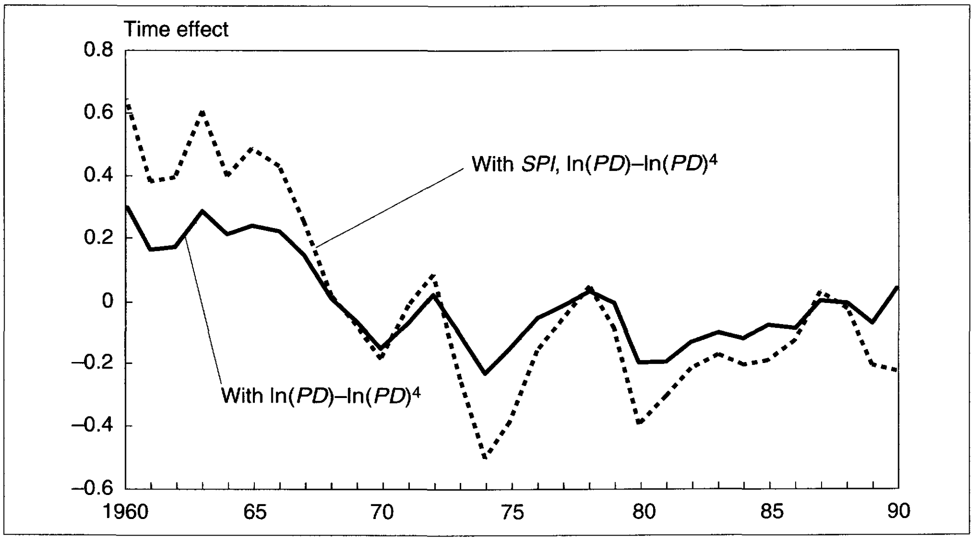
Our cross-sectional univariate model is the same as this model at a certain time. Based on the results of the previous section, we suspect that the cross-sectional income elasticities are stable over time. Hence it is natural to think that α_t captures the effect of interest rate and technical progress in a sense that α_t decreases over time. In the final part of Section V.A, we find that pooling regression for MF_2 cannot support the hypothesis of the shift of the intercept term after 1985. We think that MF_1 would be the most relevant monetary measure to examine the implications of this model, since MF_1 would be likely to reflect the effect of the changes in the financial sophistication more than MF_2 would, and take prefectural income as a scale variable. Figure 5 is our time effect obtained from the pooling regression of log MF_1 on prefectural income, *SPI*, holding ln(*PD*), ln$(PD)^2$, ln$(PD)^3$, and ln$(PD)^4$ constant and ln(*PD*), ln$(PD)^2$, ln$(PD)^3$, and ln$(PD)^4$ constant using the 1960–90 sample. We

Figure 5 Time Effect of ln(MF_1) on Income



regress these estimates of the time effects on a constant term and an interest rate to obtain the residuals. Here we use a short-term interest rate (overnight call rate) as the relevant opportunity cost of MF_1. Since the time effect captures the aggregate shock at one time, holding income constant, and if we clean out the effect of the interest rate, the residual must reflect the neutral technical progress of transaction technology common to all regions. We find a steady reduction of residuals, or the evidence of technical progress in the transaction technology from 1960 to 1990, particularly in the 1960s, in Figure 6.

Figure 6 Time Effect Net of Interest Rate Effect



VII. Summary and Policy Implications

We find that our cross-sectional estimates of income elasticities with conditioning variable are stable over time. In particular, our best estimates of income elasticity of our counterpart of M_2 minus currency are 1.2–1.4. We take advantage of the stable income elasticities obtained from cross-sectional estimates, and extend the analysis of Lucas (1988) to obtain the interest rate elasticities from aggregate time series data. Our estimates of interest rate elasticity based on our finding of income elasticity of 1.2–1.4 are –0.2 to –0.3 for real M_2 per capita.

Several papers examine the shift of money demand functions using time series models, typically checking if the fit of money demand function worsens in the 1980s or not. According to our cross-sectional estimation, our income elasticities for MF_2 are remarkably stable. We show that our approach can get information about the technical progress from the time effect of pooling regression, since our income elasticities are stable. People casually claim that financial sophistication means that they need to leave a smaller part of their income at banks. But they do not explicitly say whether the financial innovation can be captured by the income elasticities (slope effect) or the time effect (intercept effect), or both (money demand shifts every year). If we believe that financial innovation takes place, we must look at the intercept term, since we know that the slope term has been stable.

Given our arguments about the validity of the identifying restrictions required to obtain consistent estimates of the structural parameters, our findings are consistent with a variety of findings from other data sets. For example, Yoshida and Rasche (1990) found equilibrium income elasticity of M_2 to be 1.2 using time series data from 1956 to 1985, and Rasche (1990) found that Japanese M_1 income elasticity was 1.0. Using U.S. time series data for the period 1874–1975, Friedman and Schwartz (1982) found a gross M_2 income elasticity of 1.2. Lucas (1988) and Meltzer (1963) found a gross M_1 income elasticity of 1.0, but a potentially important source of bias in their findings is the rapid improvement in financial technology that has occurred since World War II. This is consistent with the pattern of our annual estimates of the level of financial technology reported in Figure 5. Although we estimate a gross scale elasticity of the demand for money, our estimates are not inconsistent with some kind of scale economies in the holding of money such as those proposed by Baumol (1952) and Tobin (1956).[16]

Our finding that we have stable income elasticities leads to several strong implications for the money targeting rules. First, our results suggest that the scale elasticity may be stable but not one. Friedman's k% rule assumes that money and income grow at the same speed, hence a primitive k% rule might not be attractive. Second, when a central bank looks carefully at the velocity as an important indicator, our findings lead to a natural guideline for the trend of velocity. If the income elasticity of money demand is 1.2–1.4, we see that the secular decline of velocity in Japanese M_2 is explained by the growth of real income. Our theory predicts that,

16. See, for example, the discussions in Karni (1973) and Mulligan (1994).

holding interest rates and the level of financial technology constant, one percentage point of real income growth reduces velocity by 0.2–0.4 percentage point. Third, when central bankers wish to control the level of money stock using the interest rate as an instrument, the precision of control depends crucially on the accuracy of the income elasticity if the money demand function is used as a guideline for setting the interest rate. Therefore, central bankers can benefit from our relatively stable and precise estimates of income elasticity.

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