

Money Demand in Japan and the Liquidity Trap*

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Abstract

In this paper we estimate long-run money demand for Japan with two functional forms that allow for the liquidity trap, and compare the empirical results for these functional forms with those for the standard log-level functional form. Estimating different functional forms leads to nonlinear cointegration. We compare the out-of-sample prediction performance of the three functional forms. Our empirical results indicate that the functional forms which allow for the liquidity trap are better than the log-level functional form based on the prediction performance.

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1 Introduction

The theory of money demand implies that the money demand function is or is almost infinitely elastic at low or zero nominal interest rates. This feature of the money demand function has important implications for monetary policy. For example, the quantity of money that the central bank prints does not have any effect on inflation or output. Keynes and monetarists were interested in this problem which has been called the liquidity trap or the zero interest bound. Because of very low short-term interest rates in Japan today and the lowest short-term interest rates in the United States in 45 years, many researchers are interested in this problem again. For example, see Krugman (1998), Orphanides and Wieland (2000), Jung, Terashashi, and Watanabe (2001), Woodford (2003), Eggertsson and Woodford (2003), and Eggertsson (2004).¹ Therefore, it is important to incorporate the liquidity trap feature in estimating the money demand function. However, in the recent literature which uses cointegration to estimate long-run money demand, the log-level (semi-log) functional form has typically been used (see, e.g., Stock and Watson (1993), and Ball (2001)). The log-level form with log money and the level of the interest rate does not incorporate the liquidity trap feature. A notable exception is Hoffman and Rasche (1991), who use a log-log form.

In this paper we estimate long-run money demand for Japan with two functional forms that allow for the liquidity trap: the log-log form and the form implied by the money in the utility function with the constant elasticity of substitution (the MUFCEs form for short). We compare the results with the log-level form. These functional forms are motivated by theory. We compare the empirical results for these two functional forms with those for the standard log-level functional form. Because of very low short-term interest rates observed in Japan since 1995, this task is important. Different functional forms lead to nonlinear cointegration as discussed by Bae and de Jong (2004), and we use their Nonlinear Cointegration Least Square estimation technique.

Anderson and Rasche (2001) and Bae and de Jong (2004) estimate and compare different functional forms of long-run money demand for the United States. Miyao (2003) uses structural break tests to study the stability of long-run money demand with the log-linear and log-log functional forms. His empirical results indicate a structural break for the log-linear form but no structural break for the log-log form. Fujiki and Watanabe (2004) use the stability of long-run money demand with the log-log form and confirmed Miyao's finding that the log-log form is stable. Our paper is complementary to

¹The functional form of Taylor-type interest rules used in these papers implicitly depends on the form of the relationship between velocity and the interest rate as in Taylor (1999), among other factors. Therefore, the functional form depends on the shape of the money demand function.

these two papers, but it is different from them in that we use the MUFCEs form in addition to the log-linear and log-log forms, compare different forms in terms of out-of sample prediction performance, and take into account nonlinear cointegration.

Our empirical results indicate that the log-log and MUFCEs functional forms that allow for the liquidity trap are better than the log-level functional form in terms of the out-of-sample prediction. The results were qualitatively similar between the log-log and the MUFCEs forms and between cointegration and nonlinear cointegration techniques.

2 Functional Forms of Money Demand

This section discusses the three main functional forms of money demand that are estimated in this paper. The difference between the three forms arises because there are various plausible ways in which the nominal interest rate enters the money demand function.

Much of the empirical work on money demand has estimated a conventional money demand function of the following functional form²

$$\ln\left(\frac{M^d}{P}\right) = \beta_0 + \beta_1 \ln(Y) + \beta_2 i, \quad (1)$$

where M^d denotes nominal money balances; P denotes the price level; Y is a “scale” variable that proxies for the volume of transactions such as real GDP or consumption; and i is the nominal interest rate which measures the opportunity cost of holding money. The parameter β_1 is the income-elasticity of money demand and β_2 is the “semi-elasticity” of money-demand with respect to the interest rate.

Although this specification of money demand has been widely used in the empirical literature on money demand, there are two important classes of models that give rise to other specifications. The first class of models is based on the inventory-theoretic approach to money demand pioneered by Allais (1947, Vol. 1, pp.235-241), Baumol (1952) and Tobin (1956). Consider an individual who receives an income Y in the form of bonds. There is a fixed transactions cost b of converting interest-bearing bonds into cash. Let K denote the real value of bonds converted into cash each time there is a conversion. The total transaction costs γ incurred by the individual are given by

$$\gamma = b\left(\frac{Y}{K}\right) + i\left(\frac{K}{2}\right), \quad (2)$$

where the first term represents conversion costs and the second term represents the interest cost on average money holdings ($K/2$) over the period. Minimizing the transaction costs with respect to K

²See for example, Stock and Watson (1993).

yields the following square-root law for optimal real money balances

$$\frac{M^d}{P} = \frac{K}{2} = \frac{1}{2} \left(\frac{2bY}{i} \right)^{1/2}. \quad (3)$$

Expressing Equation (3) in logarithmic form, we obtain the following log-linear money demand function

$$\ln \left(\frac{M^d}{P} \right) = \beta_0 + \beta_1 \ln(Y) + \beta_2 \ln(i), \quad (4)$$

where the parameters β_1 and β_2 represent the constant income- and interest-elasticities of money demand that are implied to be 1/2 by the model.³

Miller and Orr (1966) extend the Allais-Baumol-Tobin analysis to the case in which cash flows are stochastic while maintaining the assumption of a fixed transaction cost in converting bonds to money. In the basic version of their model, cash flows follow a stationary random walk without drift so that in a small time interval ($1/t$) the cash flow either increases or decreases by m dollars with equal probabilities. The optimal rule for money holdings is a “trigger-target” rule. Whenever cash balances reach the lower bound (the trigger) of zero, z dollars are converted from bonds to cash; when cash balances reach the upper bound of h , $(h - z)$ dollars of cash are converted to bonds. Miller and Orr show that the optimal size of average cash balances is given by

$$\frac{M^d}{P} = \frac{4}{3} \left[\frac{3b}{4i} \sigma^2 \right]^{1/3}, \quad (5)$$

where $\sigma^2 = m^2 t$ is the daily variance of the changes in the cash balances. The Miller-Orr model also implies a constant interest elasticity of money demand but the value is 1/3 rather than 1/2.

In more recent work Bar-Ilan (1990) extends the inventory-theoretic model further to allow for the possibility of overdrafting by relaxing the assumption that the “trigger” be restricted to zero. Money balances may fall below zero and, when they do, the individual has to pay a penalty at a rate $p > 0$ for using the overdrafting facility. It is shown that for any finite nominal interest i and penalty rate p the optimal trigger point is negative. Only in the special case when the penalty rate of using the credit is infinitely high relative to the interest rate does the model yield the Allais-Baumol-Tobin result. Since credit and money are very close substitutes, even small increases in the cost of holding money relative to credit (a higher i/p ratio) results in substitution of credit for money, thereby yielding a higher interest elasticity of money demand than the earlier models.

Another important class of models that have implications for the functional form of the money demand function are those where real balances enter the utility function directly. This approach was

³Miller and Orr (1966) also use the inventory theoretic approach to modelling the optimal amount of money holdings.

pioneered by Sidrauski (1967) and Brock (1974) and has since been widely used to study a variety of issues in monetary economics. Money enters the utility function because it helps economize on the time spent transacting and hence higher real balances are associated with higher leisure and hence higher utility.⁴ Suppose that the representative household maximizes the lifetime utility function

$$U = \sum_{t=0}^{\infty} \beta^t u(c_t, m_t), \quad 0 < \beta < 1 \quad (6)$$

by choosing time paths for consumption (c_t) and real balances (m_t) subject to an appropriate economy wide budget constraint. The first order conditions for maximizing utility yield

$$\frac{u_m(c_t, m_t)}{u_c(c_t, m_t)} = 1 - \frac{1}{(1+r_t)(1+\pi_{t+1})} = \frac{i_t}{1+i_t}, \quad (7)$$

where r_t is the real return on capital, ϕ_{t+1} is the expected inflation rate and i_t denotes the nominal interest rate. Equation (7) equates the marginal rate of substitution between real balances and consumption to the relative price of holding money. If the household holds one less dollar of money, it foregoes the opportunity to earn an interest payment i_t . Since this payment would be received next period, it is discounted by the nominal interest rate to obtain its present value.

The demand for money can be derived from Equation (7) by positing a specific utility function for the representative household. The following constant elasticity of substitution (CES) utility function has often been used

$$u(c_t, m_t) = \left[\alpha c_t^{1-\beta} + (1-\alpha) m_t^{1-\beta} \right]^{1/(1-\beta)}, \quad (8)$$

where $0 < \alpha < 1$ and $\beta > 0$, $\beta \neq 1$. With these preferences, the marginal rate of substitution between real balances and consumption is given by

$$\frac{u_m}{u_c} = \left(\frac{1-\alpha}{\alpha} \right) \left(\frac{c_t}{m_t} \right)^\beta. \quad (9)$$

Equating the marginal rate of substitution to the relative price of real money balances, we obtain the following demand for money (in log form)

$$\ln(m_t) = \frac{1}{\beta} \ln \left(\frac{1-\alpha}{\alpha} \right) + \ln(c_t) - \frac{1}{b} \ln \left(\frac{i_t}{1+i_t} \right). \quad (10)$$

This model implies a unit consumption elasticity of money demand. The interest elasticity of money demand implied by this model is

$$\frac{\partial \ln(m_t)}{\partial \ln(i_t)} = -\frac{1}{b} \frac{1}{1+i_t}, \quad (11)$$

which is a decreasing function of the nominal interest rate in terms of absolute value.

⁴Alternatively, one could assume that real balances help reduce transactions costs, so that higher real balances lead to a greater proportion of income being spent on consumption. In this case real balances enter the individual's budget constraint rather than the utility function. See Brock (1974) and Feenstra (1986) for details.

3 Estimation Results of Money Demand for Japan

3.1 Cointegration Methods

In this section the following three functional forms of the long-run money demand are estimated by cointegration methods. Since the nominal interest rate shows a persistent serial correlation, the assumption that r_t is I(1) is generally accepted as a good approximation. We regard the long-run money demand function as a cointegrating regression.

$$m_t = \beta_0 + \beta_1 i_t + u_t \quad (12)$$

$$m_t = \beta_0 + \beta_1 \ln(i_t) + u_t \quad (13)$$

$$m_t = \beta_0 + \beta_1 \ln\left(\frac{i_t}{1+i_t}\right) + u_t \quad (14)$$

where $m_t (= \ln \frac{M_t}{P_t Y_t})$ is the logarithm of the real money balance and i_t is the nominal interest rate. Note that we impose the restriction of the unit income elasticity of the money demand. We allow i_t and u_t to be temporally dependent and u_t to be serially correlated.

In Equations (13) and (14) the money demand becomes a nonlinear function of the interest rate. To use the conventional linear cointegration methods, such as “Fully Modified OLS” (FMOLS) and “Dynamic OLS” (DOLS), we must have different assumptions for different functional forms; i.e. i_t , $\ln(i_t)$ and $\ln\left(\frac{i_t}{1+i_t}\right)$ must be assumed to be I(1) for Equations (12), (13) and (14), respectively. However, if i_t is I(1), $\ln(i_t)$ and $\ln\left(\frac{i_t}{1+i_t}\right)$ cannot be I(1) in any meaningful sense and vice versa. Because of this internal inconsistency, estimation results from the conventional linear cointegration methods might not be directly comparable with each other. Therefore, along with the conventional linear cointegration methods, we also consider a nonlinear cointegration method which has been proposed recently by Bae and de Jong (2004). In their “Nonlinear Cointegration Least Square” (NCLS) estimation technique, it is possible to estimate different functional forms under the one assumption that i_t is I(1). However, the NCLS estimation method used in this paper has no asymptotic justification for Equation (14). A theory has not been fully developed yet. Therefore, we also report bootstrap confidence intervals along with asymptotic ones.

Since the NCLS estimation technique is relatively new, we illustrate how to implement the NCLS estimation technique for the estimation of β_1 in Equation (13). Let k_n be an integer-valued positive sequence that diverges to infinity at a slower rate than n such that $k_n n^{-\frac{p-2}{3p+\eta}} \rightarrow 0$ for some $\eta > 0$ and

p ,⁵ and $n_j = \left\lfloor \frac{n_j}{k_n} \right\rfloor$ for $j = 0, 1, 2, \dots, k_n$. Let $z_t = \ln(i_{n_{j-1}+1})$ for $n_{j-1}+1 \leq t \leq n_j$ for $j = 1, 2, \dots, k_n$. Then the NCLS estimator $\tilde{\beta}_1$ is defined as an IV estimator that uses z_t as the instrumental variable for $\ln(i_t)$. Note that although $\tilde{\beta}_1$ is a consistent estimator, it cannot be used for statistical inference unless the limiting Brownian processes associated with i_t and u_t are orthogonal, which is unlikely in the case of the long-run money demand function. Therefore, the following fully modified type NCLS estimation technique is used. The estimation procedure is as follows.

1. Calculate the residual, \hat{u}_t , from a regression by the NCLS estimation method.
2. Get a HAC estimate for the long-run covariance matrix of $(u_t, \Delta i_t)$, $\hat{\Omega}$, by using $(\hat{u}_t, \Delta i_t)$.
3. Calculate \hat{m}_t^\dagger in a way analogous to the FMOLS,

$$\hat{m}_t^\dagger = m_t - \hat{\Omega}'_{21} \hat{\Omega}_{22}^{-1} \Delta i_t.$$

4. The fully modified version of the NCLS estimator $\hat{\beta}_1$ is defined as the NCLS estimator that is calculated using the modified dependent variable \hat{m}_t^\dagger instead of m_t .

Note that now the usual “ t and F -statistics” are valid because they achieve the correct significance level conditionally on Δi_t .

3.2 Data and Empirical Results

For the estimation of the Japanese long-run money demand function the quarterly data set from 1976:1 to 2002:4 is used.⁶ Since the data frequency is quarterly, we add quarterly seasonal dummies in the regression. M2+CD, the Consumer Price Index (CPI), both the Gross Domestic Product (GDP) and the Private Consumption (CON), and the lending rate of “City Banks” are used for money, price, output, and nominal interest rate, respectively. Due to the financial “Big Bang” in Japan, we include foreign banks’ accounts in M2+CD beginning in 1998:2.

Table 1 reports coefficient estimates, and its asymptotic and bootstrap confidence intervals. Since no asymptotic and bootstrap confidence intervals contain zero, coefficient estimates are statistically

⁵The following assumption needs to be made regarding p . Let u_t and Δi_t be linear processes given by

$$u_t = \sum_{i=0}^{\infty} \phi_{1,i} \varepsilon_{1,t-i} \quad \Delta i_t = \sum_{i=0}^{\infty} \phi_{2,i} \varepsilon_{2,t-i}$$

where $\varepsilon_t = (\varepsilon_{1,t}, \varepsilon_{2,t})$ is a sequence of independent and identically distributed (i.i.d.) random variables with mean zero. $E|\varepsilon_{j,t}|^p < \infty$ for some $p > 2$ for $j = 1, 2$; for details, see Bae and de Jong (2004).

⁶Shinichi Nishiyama in the Bank of Japan kindly provided the data of M2+CD and the lending rate of “City Bank”. For CPI, GDP, and the Private Consumption the Datastream was used.

significant in all combinations of functional forms and estimation methods. Asymptotic and bootstrap confidence intervals are generally similar, though bootstrap one is larger than asymptotic one. When we compare the estimation results across the different functional forms, there are significant differences, as expected. However, the estimation results are robust across the different estimation methods, including the NCLS estimator.

To further address the question of which functional form is most appropriate for the Japanese long-run money demand, we investigate out-of-sample prediction performances for the three different functional forms. Table 2 reports the sum of squared error for two different methods of out-of-sample prediction performance. Equations (13) and (14), which are nonlinear functions of the interest rate, clearly outperform Equation (12), the linear one, in all estimation methods except DOLS. These prediction performance results support empirically our conviction that the nonlinear functional forms, such as Equations (13) and (14), are more appropriate for the Japanese long-run money demand.

4 Conclusions

In this paper we estimated long-run money demand for Japan with two functional forms that allow for the liquidity trap and compared the empirical results for these functional forms with those for the standard log-level functional form. Estimating different functional forms leads to nonlinear cointegration. However, we found that the empirical results are robust to estimation methods that assume linear and nonlinear cointegration. We then compared the out-of-sample prediction performance of the three functional forms. Our empirical results indicated that the functional forms which allow for the liquidity trap are better than the log-level functional form.

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Table 1: Estimation Results for β_1^1 (Seasonal Dummy)

$y_t = \text{GDP}$				
	i_t			
	SOLS	DOLS ²	FMOLS	NCLS ^{3]}
Asymptotic ⁴	-0.0425 (-0.0569, -0.0280)	-0.0477 (-0.0637, -0.0317)	-0.0448 (-0.0589, -0.0307)	-0.0440 (-0.0581, -0.0299)
Bootstrap ⁵	(-0.0657, -0.0263)	(-0.0709, -0.0247)	(-0.0680, -0.0264)	(-0.0669, -0.0247)
	$\ln(i_t)$			
Asymptotic	-0.1646 (-0.2177, -0.1114)	-0.1715 (-0.2326, -0.1104)	-0.1700 (-0.2219, -0.1180)	-0.1653 (-0.2171, -0.1134)
Bootstrap	(-0.2431, -0.0977)	(-0.2577, -0.0857)	(-0.2516, -0.0982)	(-0.2450, -0.0888)
	$\ln\left(\frac{i_t}{1+i_t}\right)$			
Asymptotic	0.1706 (0.1155, 0.2257)	0.1774 (0.1139, 0.2408)	0.1760 (0.1222, 0.2299)	0.1712 (0.1174, 0.2250)
Bootstrap	(0.1008, 0.2515)	(0.0881, 0.2669)	(0.1016, 0.2607)	(0.0865, 0.2475)
$y_t = \text{Consumption}$				
	i_t			
Asymptotic	-0.0452 (-0.0645, -0.0258)	-0.0518 (-0.0725, -0.0311)	-0.0480 (-0.0668, -0.0292)	-0.0471 (-0.0659, -0.0282)
Bootstrap	(-0.0766, -0.0232)	(-0.0811, -0.0219)	(-0.0792, -0.0238)	(-0.0782, -0.0218)
	$\ln(i_t)$			
Asymptotic	-0.1764 (-0.2477, -0.1050)	-0.1864 (-0.2646, -0.1082)	-0.1831 (-0.2524, -0.1138)	-0.1774 (-0.2468, -0.1080)
Bootstrap	(-0.2821, -0.0866)	(-0.2964, -0.0772)	(-0.2930, -0.0887)	(-0.2896, -0.0776)
	$\ln\left(\frac{i_t}{1+i_t}\right)$			
Asymptotic	0.1828 (0.1089, 0.2568)	0.1929 (0.1117, 0.2741)	0.1897 (0.1179, 0.2616)	0.1838 (0.1119, 0.2557)
Bootstrap	(0.0896, 0.2920)	(0.0793, 0.3072)	(0.0915, 0.3034)	(0.0712, 0.2856)

- 1) Figures in parenthesis are the 95% confidence interval.
- 2) The number of leads and lags is 3.
- 3) The bandwidth parameter, $\left[\frac{n}{k_n}\right]$, is 5. Results are robust between 4 and 8.
- 4) An HAC estimator with Bartlett kernel is used with the bandwidth parameter of 4.
- 5) The overlapping block bootstrap method is used with the block size of 5.
Bootstrap sample size is 10,000.

Table 2: Prediction Performance Results¹

	SOLS	DOLS	FMOLS	NCLS
GDP				
stepwise one-step ahead forecast ²				
i_t	0.0863	0.0704	0.0696	0.0755
$\ln(i_t)$	0.0359	0.1341	0.0393	0.0381
$\ln(\frac{i_t}{1+i_t})$	0.0354	0.1390	0.0394	0.0379
two separated sample ²				
i_t	0.1326	0.0708	0.0930	0.1059
$\ln(i_t)$	0.0343	0.4753	0.0466	0.0373
$\ln(\frac{i_t}{1+i_t})$	0.0339	0.5082	0.0489	0.0381
Consumption				
stepwise one-step ahead forecast ²				
r_t	0.0923	0.0663	0.0688	0.0761
$\ln(r_t)$	0.0303	0.1694	0.0357	0.0342
$\ln(1 + \frac{1}{r_t})$	0.0297	0.1770	0.0361	0.0341
two separated sample ²				
r_t	0.1584	0.0801	0.1009	0.1172
$\ln(r_t)$	0.0301	0.6321	0.0467	0.0349
$\ln(1 + \frac{1}{r_t})$	0.0292	0.6763	0.0499	0.0360

1) For estimation the same specifications as Table 1 are used.

2) Prediction period is from 1997:1 to 2002:4.