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Abstract. Based on a standard model of money demand, this paper first shows that a relationship between money supply and prices may be substantially weakened when money demand is highly interest-elastic, and then presents empirical evidence for this implication using the Japanese money market data for the sample period, 1985–1999.

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1. Introduction Many empirical studies on Japanese money demand functions, including Miyao [4], Fujiki and Watanabe [3], Bae, Kakkar, and Ogaki [1], and Nakashima and Saito [5] document that the interest-rate *semi-elasticity* has been extremely high since the middle of the 1990s, when overnight money market rates were below 0.5% per year. The first three articles adopt the log-log specification, in which the interest-rate semi-elasticity is inversely proportional to nominal interest rates, while the last article finds a substantial decrease in the interest-rate semi-elasticity using the semi-log specification with structural breaks. For example, Nakashima and Saito [5] find that the interest-rate semi-elasticity ranges from -0.039 to -0.037 up to 1995 (or 1996), while it takes a value between -0.678 and -0.459 after 1995 (or 1996). Miyao [4], and Bae, Kakkar, and Ogaki [1] show that the semi-elasticity implied by the log-log specification is comparable with what Nakashima and Saito [5] report.

However, the literature of monetary economics has not explored in depth the empirical implications of such highly interest-elastic money demand. This paper first derives from a standard model of Cagan [2] possible empirical implications of highly interest-elastic money demand in terms of a money-price relationship (Section 2), and then presents evidence for these implications using the Japanese money market data for the sample period from 1985 to 1999 (Section 3).

2. Interest-elastic money demand and its implications This section briefly reviews implications of a money demand function for the quantity theory of money, and explores possible impacts of interest-rate semi-elasticity on a money–price relationship.

Suppose that the demand for real money balances is characterized as a function of real aggregate output and the nominal interest rate in themanner of Cagan [2], or in terms of the following semi-log specification:

$$m_t - p_t = \theta y_t + \frac{1}{\gamma} i_t, \tag{1}$$

where m_t is the logarithm of the nominal money stock at time t; p_t is the logarithm of nominal prices; and i_t is the nominal interest rate. The two parameters θ and γ denote the income elasticity and the inverse interest-rate semi-elasticity, respectively. In the above specification, the closer the absolute value of γ is to zero, the higher the degree of interest-rate semi-elasticity.

The nominal interest rate, assumed to be determined by the Fisher equation, is equal to the sum of the real interest rate r_t and the expected inflation $p^e_{t+1} - p_t$ where p^e_{t+1} denotes the expected future price. For the moment and for simplicity, it is further assumed that $r_t = 0$ and $y_t = 0$. Then, equation (1) reduces to the following rational expectations model:

$$p_t = \frac{1}{1-\gamma} p_{t+1} + \frac{-\gamma}{1-\gamma} m_t$$

In the standard case in which the real money balance is a decreasing function of the nominal interest rate ($\gamma < 0$), we obtain the following forward-looking path:

$$p_t = -\gamma \sum_{\tau=0}^{\infty} \left[\left(\frac{1}{1-\gamma} \right)^{\tau+1} m_{t+\tau} \right]. \tag{2}$$

With the above path, the current nominal price reflects both the current and the future money supplies, and nominal prices respond flexibly to changes in the money supply. If the money supply increases permanently by an amount Δm, then nominal prices rise by the same magnitude. In the case of a permanent change in the money supply, therefore, there is a one-to-one correspondence between the money supply and nominal prices; that is, the standard quantity theory holds firm.

Nominal prices are, however, less responsive to nonpermanent changes in the money supply, as γ is closer to zero and money demand is more interest-elastic. According to the coefficient of the future money supply in equation (2), $-\gamma \left(\frac{1}{1-\gamma}\right)^{\tau+1}$, as γ is closer to zero, less weight is put on the current and immediate future money supply and more on the distant future money supply. Therefore, transitory changes in the money supply are not significantly reflected in current nominal prices when money demand is extremely interest-elastic.

Even if changes in the money supply are permanent with γ close to zero, then nominal prices may be less responsive to the money supply in the following cases. First, if the Central Bank cannot make a firm commitment to permanent increases in the money supply, policy shocks on the money supply turn out to be transitory; therefore, as discussed above, current nominal prices do not respond to the money supply with such limited commitment when negative γ is close to zero. Second, when market participants are myopic and consider only current and immediate future money supplies, we have the same implication for a money–price relationship, as in the first case.

The above discussion suggests that, when money demand is extremely interest-elastic, nominal prices are likely to be less responsive to monetary expansion. Then, when monetary policy is aggressive, the relative size of money demand $(m_t - (p_t + y_t))$ would increase as a result of a breakdown of one-to-one correspondence between the nominal money supply and nominal prices.

3. Empirical results

3.1. **Data** For our estimation, the sample period is August 1985 to March 1999. The principal reason for excluding the period before 1985 is that Japanese money markets were strictly regulated until the mid-1980s. It is only since the mid-1980s that commercial banks and securities companies have been allowed to issue various types of money market instruments at market rates. Therefore, money market rates were unlikely to have properly reflected market conditions before 1985. Our sample period thus starts from August 1985, when the uncollateralized call market was established.

A major reason for omitting the period after March 1999 is that an almost infinitely elastic money demand at zero interest rates has been self-evident as a result of either the zero interest-rate policy of February 1999 or the quantity-easing policy of March 2001. However, February and March of 1999 are included in the sample period because the Bank of Japan (BOJ) publicly announced a firm commitment to a zero interest-rate policy in April 1999. In addition, the inclusion of data for years with nominal interest rates at the lower bound (0%) for relatively long periods would cause serious econometric problems. For the sample period before April 1999, nominal interest rates stayed at low levels, but were still above zero rates over time.

We build the set of monthly data as follows. As nominal monetary aggregates, we choose M1 because M1 reflects the transaction demand for money to a greater extent than do other monetary aggregates.

The consumer price index constructed by the Statistics Bureau is used for nominal prices, and the industrial production index documented by the Ministry of International Trade and Industry is adopted for real aggregate output. Overnight call rates, reported by the BOJ, are used as nominal interest rates. All data are recorded as monthly averages. As for both nominal monetary aggregates and industrial production, our data set is based on variables that are officially seasonally adjusted by the above reporting agencies. The consumer price index is seasonally adjusted by the X11 method based on the sample period, 1970–2005.

Unit root tests for the real money balance, real output, and nominal interest rates (call rates) fail to reject unit roots for levels, but do reject unit roots for first differences in all cases.

3.2. **Short-run responses to changes in money supply** As discussed in the previous section, highly interest-elastic demand may make nominal prices unresponsive to changes in the money supply when the Central Bank cannot make a firm commitment to permanent changes in money or the market participants do not have long-run expectations about the money supply. In this subsection, we empirically examine whether such a phenomenon indeed emerged as a result of the highly interest-elastic money demand.

To differentiate the effect of money supply on nominal prices between the pre-break and post-break periods, we estimate the following equation:

$$
\begin{aligned}
\Delta p_t &= \text{constant} + \lambda_0^c I_{date<break} + \lambda_0^m I_{date<break}\Delta m_t + \lambda_1^m I_{date\geq break}\Delta m_t \\
&+ \lambda_0^y I_{date<break}\Delta y_t + \lambda_1^y I_{date\geq break}\Delta y_t \\
&+ \mu_0 I_{date<break}\epsilon_{t-1} + \mu_1 I_{date\geq break}\epsilon_{t-1} + \xi_t, \qquad (3)
\end{aligned}
$$

where I is the indicator function dependent on the condition defined in subscripts. For example, if a data point is before a break, then $I_{date<break}$ is one, otherwise zero. The final term, ξ_t, represents a stochastic disturbance. In addition, the lagged ϵ_t, defined by the semi-log specification estimated by Nakashima and Saito [5], serves as the error correction term, given that a cointegration relationship holds among real money balances, outputs, and nominal interest rates. [1]

If a weak short-run relationship between nominal prices and the money supply was created by the low interest-rate policy, then we expect $\lambda_0^m > 0$ and $\lambda_1^m = 0$. In addition, we may have $\lambda_0^m \neq \lambda_1^m$. With respect to the coefficients on the error correction terms, we expect $\mu_0 > 0$ and $\mu_1 > 0$ if there is a quick recovery to long-run equilibrium.

The estimation of equation (3) requires instrumental variable estimation to control for simultaneity biases. We include, as instrumental variables, constant terms, lagged changes in money supply, and lagged real output increases. The number of lags is controlled from one to four. Table 1 reports the empirical results. [2] In this estimation, we set June 1996 for M1 as a break point following the result of Nakashima and Saito [5].

The most important finding is that λ_0^m is significantly positive, whereas λ_1^m is not significantly different from zero. The contrast between λ_0^m and λ_1^m is remarkable; $\lambda_0^m = \lambda_1^m$ is rejected statistically, as is shown by the last column of Table 1. On the other hand,

[1] More precisely, Nakashima and Saito [5] estimate, by dynamic OLS, $m_t - p_t = 3.812 + 1.277y_t - 0.039i_t$, for the sample period 1985:8-1996:5; and $m_t - p_t = 8.823 + 0.253y_t - 0.459i_t$ for the sample period, 1996:6-1999:3; and, by the fully modified OLS, $m_t - p_t = 4.092 + 1.197y_t - 0.037i_t$, for the sample period, 1985:8-1995:7; and $m_t - p_t = 7.810 + 0.485y_t - 0.678i_t$, for the sample period, 1995:8-1999:3.

[2] Because the estimated constant term of equation (3), if included, is not significantly different from zero, this table reports the case without a constant term.

both λ_0^y and λ_1^y are insignificant. The coefficients on the error correction terms are also insignificant, suggesting that the path returned to long-run equilibrium quite slowly.

As shown in Table 2, the empirical results do not change substantially even if break points are based on August 1995 for M1 as a break point, following another result of Nakashima and Saito [5].

Our findings clearly suggest that nominal prices responded immediately, although only partially, to changes in the money supply in the pre-break period, but not at all in the post-break period.

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Table 1: Parameter Estimates of Error Correction Type Models (June 1996 for M1 as a break point)

m_t	Lags	γ_0^m	γ_1^m	γ_0^y	γ_1^y	μ_0	μ_1	F
M1								
	1	0.258	0.056	-0.043	-0.035	0.016	0.003	6.273
		(0.090)	(0.074)	(0.354)	(0.251)	(0.187)	(0.034)	(0.013)
	2	0.267	0.081	-0.135	0.047	0.069	-0.014	7.669
		(0.085)	(0.072)	(0.143)	(0.135)	(0.071)	(0.024)	(0.006)
	3	0.198	0.084	-0.010	0.019	0.001	-0.019	3.830
		(0.051)	(0.052)	(0.047)	(0.083)	(0.023)	(0.017)	(0.052)
	4	0.190	0.100	-0.014	0.045	0.005	-0.024	2.747
		(0.048)	(0.053)	(0.036)	(0.076)	(0.019)	(0.017)	(0.099)

1. The error correction type model is specified as

$$\begin{aligned}\Delta p_t &= \gamma_0^m I_{year<break}\Delta m_t + \gamma_1^m I_{year\geq break}\Delta m_t \\ &+ \gamma_0^y I_{year<break}\Delta y_t + \gamma_1^y I_{year\geq break}\Delta y_t + \mu_0 I_{year<break} z_{t-1} + \mu_1 I_{year\geq break} z_{t-1},\end{aligned}$$

   where z_t is defined as $(m-p)_t - (\text{constant} + \alpha y_t + \beta i_t)$ using the estimation result of the dynamic OLS with the number of lagged variables equal to 3 in Nakashima and Saito (2007), or $m_t - p_t - 3.812 - 1.277 y_t + 0.039 i_t$ for the sample period 1985:8-1996:5 and $m_t - p_t - 8.823 - 0.253 y_t + 0.459 i_t$ for the sample period 1996:6-1999:3.

2. Instrumental variables include constant, lagged Δm_t, and lagged Δy_t. The number of lags for instrumental variables is controlled from one to four.

3. Standard errors are in parentheses.

4. The last column reports the F statistics of $\gamma_0^m = \gamma_1^m$. P values of the F statistics are in Parentheses.

Table 2: Parameter Estimates of Error Correction Type Models (August 1995 for M1 as a break point)

m_t	Lags	γ_0^m	γ_1^m	γ_0^y	γ_1^y	μ_0	μ_1	F
M1								
	1	0.241	0.060	-0.024	-0.030	0.040	0.004	9.378
		(0.069)	(0.058)	(0.065)	(0.180)	(0.303)	(0.018)	(0.002)
	2	0.249	0.062	-0.029	0.030	0.115	0.001	9.911
		(0.090)	(0.075)	(0.062)	(0.140)	(0.148)	(0.017)	(0.001)
	3	0.193	0.059	-0.013	-0.004	-0.019	-0.006	6.296
		(0.047)	(0.046)	(0.032)	(0.070)	(0.035)	(0.010)	(0.013)
	4	0.183	0.067	-0.016	0.038	-0.019	-0.007	5.366
		(0.043)	(0.046)	(0.030)	(0.062)	(0.028)	(0.010)	(0.022)

1. See the footnotes to Table 1.

2. z_t is defined as $(m-p)_t - (\text{constant} + \alpha y_t + \beta i_t)$ using the estimation result of the fully modified OLS in Nakashima and Saito (2007), or $m_t - p_t - 4.092 - 1.197 y_t + 0.037 i_t$ for the sample period 1985:8-1995:7 and $m_t - p_t - 7.810 - 0.485 y_t + 0.678 i_t$ for the sample period 1995:8-1999:3.