Strong Money Demand and Nominal Rigidity: Evidence from the Japanese Money Market under the Low Interest Rate Policy

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Abstract. This paper empirically explores some theoretical implications of the strong money demand that is driven by extremely low nominal interest rates. In particular, it examines whether nominal prices move inertially when nominal interest rates are extremely low, using the Japanese money market data recorded between 1985 and 1999. The paper presents the following empirical findings. First, the cointegration relationship dictated by the long-run real money demand function changed structurally around 1995 when the Bank of Japan started to guide overnight call rates below 0.5%. In the period between 1995 and 1999, money demand was extremely interest-elastic, and was not responsive to the aggregate production activity. Second, even with due consideration for the long-run relationship, nominal prices did not respond to changes in nominal money supply under the low interest rate policy. These findings are fairly consistent with theoretical implications available from monetary models where nominal rigidity arises due to strong money demand. In conclusion, the paper explores several implications for monetary policy given the above empirical findings.

JEL classification:

Keywords: real money balance, zero interest rate policy, nominal rigidity.

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1. Introduction Since September 1995, the Bank of Japan, hereafter the BOJ, had developed the low interest rate policy that had never been experienced before (see Figure 1). While the BOJ initially guided overnight call rates, or overnight interbank rates, below 0.5%, it lowered overnight rates even from 0.25% in September 1998. In February 1998, the BOJ implemented the so-called zero interest rate policy whereby the targeted overnight rate was set at almost zero percent. It was in August 2000 that the zero interest rate policy was lifted. The targeted rate, however, had remained to be well below 0.5% since then.

The above monetary policy characterized by extremely low policy-targeted interest rates, however, did not necessarily yield favorable effects on the Japanese macroeconomy, which effects were expected initially when the policy was implemented. In particular, a limited impact on nominal prices kept real interest rates rather high, thereby exercising negative effects on household and corporate spending activities, and prolonging a recession. Growing frustrations at such limited impacts shared among politicians, policy makers, and economist, both at home and abroad, urged the BOJ to implement even more expansionary policy such as expanding money supply aggressively and targeting higher inflation.

Did the above limited impact of monetary policy on nominal prices arise from the passive BOJ monetary policy as pointed out unanimously among influential journalists, critics, and economists at home and abroad? In what follows, this paper pursue another possibility. That is, we demonstrate that the inertial movement of nominal prices was caused not by a lack of sufficient money supply, but by the extremely low interest rate policy itself. As discussed below, the standard quantity theory of money is unlikely to hold when nominal interest rates are extremely low, and accordingly the relationship between money supply and nominal prices tends to break down.

From a macroeconomic point of view, the phenomenon that nominal prices never respond to current money supply or the arrival of information as to future money supply is called *nominal rigidity*. In the standard macroeconomic argument, the source of nominal rigidity is quite often identified as frictional factors in goods and labor markets. For example, Ball and Mankiw [3] discuss, from both theoretical and empirical studies, how frictional factors are responsible for nominal rigidity.

Nominal rigidity, however, has been discussed in a completely different context. That is, strong money demand absorbs quickly and vigorously an additional supply of money, thereby nullifying effects on nominal prices. As is well-known, Keynes [11] argues for this case, the liquidity trap where money is perfectly substitute for long-term bonds when nominal interest rates are low, and holding long-term bonds will result in capital losses with a heavy probability.

While the above liquidity trap arises from a substitution for long-term bonds, the same phenomenon emerges when lower nominal interest rates tend to substitute money for money market instruments. When nominal interest rates are very low, money demand is boosted by low costs of money-holding. In particular, when nominal interest rates are almost zero, money demand is infinitely interest-elastic, and consequently the real money balance can take any value.

In a modern macroeconomic context, Farmer [6] presents the case where the externality of money circulation tends to promote money demand. That is, the more actively currency circulates, the more frequently trading takes place, thereby leading to even stronger money demand. He demonstrates that nominal prices never respond to the arrival of information concerning money supply when such an exchange externality is strong enough. While he presents a calibration study based on this model, a judgement of the empirical relevance of this implication for US economy, in particular about the magnitude of exchange externality, differs among macroeconomists.

In whichever case, strong money demand is responsible for nominal rigidity, ant it limits impacts of monetary policy on macroeconomy. The low interest rate policy implemented by the BOJ also brought about 'strong money demand.' According to Figure 2, the size of monetary aggregates represented by M1 and M2+CD relative to nominal GDP increased remarkably in the latter half of the 1990s. In particular, the ratio based on M1, which monetary aggregate reflects transaction motive to larger extent, had swelled towards 40 % since the middle of the 1990s, while it had been fairly stable between 25% and 30% until then.

This paper is organized as follows. Section 2 discusses how strong money demand driven by high interest elasticity is responsible for nominal rigidity using a simple monetary model. Section 3 empirically demonstrates that the low interest rate policy led to a structural change in an interest elasticity of a money demand function, and that it weakened substantially the relationship between money supply and nominal prices. Finally, Section 4 explores several policy implications given our empirical findings about nominal rigidity.

2. Strong money demand and nominal rigidity In a typical case of a money demand function, when nominal interest rates are lower, not only cheaper costs of money-holding, but also larger interest elasticity boost money demand. In the extreme case of zero nominal interest rates, interest elasticity of money demand becomes infinite, and the real money balance can take any value. In this section, we explore some implications of highly interest-elastic money demand for nominal prices using a simple monetary model. As demonstrated below, high interest elasticity may yield nominal rigidity.

Suppose that the real money demand is a function of the real aggregate output and the nominal interest rate, or

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

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

The nominal interest rate is assumed to be determined by the Fisher equation, or to be 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 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:

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

The property of the equilibrium path of p_t depends on the magnitude of interest elasticity or γ. In the standard case where the real money balance is decreasing in the nominal interest rate or $\gamma < 0$, we obtain the following forward-looking path:

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

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

On the other hand, if interest elasticity is positive and smaller than two, then the above forward-looking path is no longer available, and the nominal price is determined in a backward way according to the weighted average of p_t and m_t, or $p_{t+1} = (1-\gamma)p_t + \gamma m_t$. In this case, the current nominal price is adjusted gradually from the lagged level in response to only the current money supply. If γ is close enough to zero, then the current nominal price does not respond even to the current money supply. This case corresponds to the monetary model with exchange externality built by Farmer [6] [1] .

On the border where interest elasticity is infinite or $\gamma = 0$, we obtain

$$p_{t+1} = p_t.$$

That is, nominal prices do not respond at all to either the current or future money supply, and perfect nominal rigidity emerges. Even allowing for time-varying real interest rates r_t and real outputs y_t, we still obtain the above border property, or

$$p_{t+1} = p_t - r_t.$$

In the above case, the current nominal price never responds to either money supply or outputs. If real interest rates are uncorrelated with real outputs, then the real money balance $m_t - p_t$ is also irrespective of real outputs.

One thing to be noticed is that the implication associated with infinite interest elasticity does not depend at all on whether the Fisher equation holds or not. In other words, the above absence of the relationship between money supply and nominal prices is fairly independent of the relevance of intertemporal equilibrium conditions.

[1] The essence of the monetary model presented by Farmer [6] may be captured by the following money demand function:

$$m_t - p_t = \frac{1}{\gamma} i_t + \lambda(\bar{m}_t - p_t),$$

where $\bar{m}_t$ equals the amount of nominal money that circulates in a macroeconomy. The second term of the right hand side represents exchange externality or the positive effect of money balance on transaction; the larger λ is, the stronger exchange externality is. Because $m_t = \bar{m}_t$ in equilibrium, the coefficient on the nominal interest rate is equal to $\frac{1}{\gamma}\frac{1}{1-\lambda}$ in a reduced form. In this case, the interest elasticity of the reduced form may be positive when exchange externality is strong enough. More concretely, even if γ is negative, the reduced-form interest elasticity is positive under the condition that $1 < \lambda < 1 - \frac{1}{\gamma}$. Beaudry and Devereux [4] construct a similar model in the two-country setup, thereby exploring implications of nominal rigidity for real exchange rates.

In a realistic situation, the theoretical implication identical to the borderline case may not necessarily require γ to be exactly equal to zero. High interest elasticity means that the money demand function is almost horizontal. Then, almost horizontal demand functions imply that no change in nominal prices in response to external shocks may not result in a large welfare loss. In the presence of small costs associated with price changes in goods markets, no change in nominal prices may be preferred to changes in nominal prices.

Figure 3 illustrates the above point. Suppose that the current position of the real money balance is located at point A, where interest elasticity is rather high. Now, there is an increase in money supply. In the case of flexible prices, then, the real money balance is still at point A, and nominal prices increase. However, even if nominal prices are intact, and the real money balance moves to point B, the welfare loss represented by the area of the triangle put between the two dotted vertical lines, is still fairly small. Then, small frictional costs may keep nominal prices from increasing.

The preceding argument suggests that if the real money balance is highly elastic with respect to nominal interest rates, then nominal prices are likely to be rigid. In addition, real outputs tend to have no impact on real money demand as far as real interest rates are not correlated with real outputs. As the consequence of strong money demand with high interest elasticity, nominal rigidity emerges. The next section explores this implication of strong money demand using the Japanese money market data.

3. Estimation Methods and Results

3.1. **Tests of structural breaks and data** In this section, we empirically examine how the shape of a Japanese money demand function changed due to the low interest rate policy implemented by the BOJ in 1995, in particular, whether interest elasticity became substantially larger than before. As discussed in the previous section, highly interest-elastic money demand serves as a necessary condition for nominal rigidity where nominal prices never respond to current changes in money supply. In the final subsection, we test how irresponsive to money supply changes nominal prices were under the low interest rate policy.

First of all, we specify a money demand equation as follows:

$$m_t - p_t = \text{constant} + \alpha y_t + \beta i_t + \epsilon_t, \tag{2}$$

where α implies income elasticity, while β denotes interest semi-elasticity. The last term ϵ_t

represents a stochastic shock on money demand.

Considering the contemporaneous and intertemporal correlation between the innovation on the explanatory variables (y_t and i_t) and the disturbance in money demand (ϵ_t), we estimate equation (2) by the dynamic OLS proposed by Saikkonen [17], Stock and Watson [18], and others, and the fully modified OLS proposed by Phillips and Hansen [15] and others. The main purpose of adopting two estimation methods is not to compete with each other, but to carefully examine the robustness of test and estimation results.

For the purpose of tests of the presence of structural breaks, the test proposed by Hansen [8] is adopted for a pure structural change where constancy in the whole set of parameters is tested against parameter instability, while the test proposed by Kuo [12] is constructed for a partial structural change where constancy in subsets of parameters are examined. In both tests of structural changes, the null hypothesis of cointegration with parameter stability is tested against the alternative of cointegration with parameter instability[2] .

As Hansen [8] suggests, the asymptotic distribution for tests of structural changes, derived by Hansen [8] and Kuo [12] is applicable to not only the fully modified OLS, but also the dynamic OLS. In what follows, the estimation results based on the dynamic OLS are presented in this subsection, and are then compared carefully with those based on the fully modified OLS.

In our estimation, the sample period is set to be between 1985 and 1999. The principal reason for excluding the period before 1985 is that Japanese money markets had been regulated strictly until the mid-1980s. It was only since the mid-1980s that commercial banks and securities companies had been allowed to issue various kinds of money market instruments at market rates. Therefore, money market rates were unlikely to properly reflect market conditions before 1985.

We build the set of monthly data as follows. As nominal monetary aggregates, we choose M1, M2+CD, and currency. These data are compiled by the BOJ. Among the above three monetary aggregates, we are particularly interested in the estimation results based on M1, because M1 reflects to greater extent the transaction motive of money-holding than the

[2] While our investigation of structural breaks regards cointegration with parameter constancy as the null hypothesis using test statistics based on the limiting distribution, alternative methods have been explored in depth in the econometric literature on structural breaks. For example, Andrews, Lee, and Ploberger [1] examine the finite sample property of a test of cointegration with structural breaks. In a different context, Gregory and Hansen [7] construct a test of no cointegration against cointegration with possible structural breaks.

other two aggregates.

The consumer price index constructed by the Statistics Bureau is used for nominal prices, while the industrial production index documented by the Ministry of International Trade and Industry is adopted for real aggregate outputs. Overnight call rates, reported by the BOJ, are used as nominal interest rates. All data are recorded in monthly averages. As for nominal monetary aggregates, consumer prices, and industrial production, our dataset is based on the variables that are seasonally-adjusted officially by the above reporting agencies. As reported in Table 1, unit root tests for the real money balance, real outputs, and nominal interest rates fail to reject the presence of unit roots in most cases.

3.2. **Test results of structural breaks** We employ the LM-type test (Lagrange multiplier test) using the estimation of either the dynamic OLS or the fully modified OLS for tests of cointegration with parameter stability against pure or partial structural changes. The first step in the procedure of tests for pure structural changes is to choose a break point T, and to construct a set of time-varying parameters $(\alpha_t, \beta_t, \text{constant}_t)$ for equation (2) as follows:

$$\text{if } t < T, \text{ then } \begin{bmatrix} \alpha_t \\ \beta_t \\ \text{constant}_t \end{bmatrix} = \begin{bmatrix} \alpha^1 \\ \beta^1 \\ \text{constant}^1 \end{bmatrix},$$

$$\text{if } t \geq T, \text{ then } \begin{bmatrix} \alpha_t \\ \beta_t \\ \text{constant}_t \end{bmatrix} = \begin{bmatrix} \alpha^2 \\ \beta^2 \\ \text{constant}^2 \end{bmatrix}.$$

Second, we compute the LM test statistics to test $(\alpha^1, \beta^1, \text{constant}^1) = (\alpha^2, \beta^2, \text{constant}^2)$ based on the estimation results of either the dynamic OLS or the fully modified OLS. The thus-computed LM test statistics is usually called F statistics by convention.

Then, F statistics are computed for all data points. There are two types of tests based on those computed F statistics. When the timing of a structural break is treated as unknown, it is possible to adopt the Sup F test based on the highest F statistics among them. On the other hand, when the parameters $(\alpha_t, \beta_t, \text{constant}_t)$ follow a martingale process under the alternative hypothesis, it is possible to use the Mean F test based on the average of those F statistics. For a partial structural change, the above procedure is applied only to a subset of $(\alpha_t, \beta_t, \text{constant}_t)$. Our empirical investigation considers as a partial structural change, the constancy of either the intercept, income elasticity (α), or interest semi-elasticity (β).

Critical values based on the limiting distribution are available from Hansen [8] for a pure structural change, and from Kuo [12] for a partial structural change. As mentioned before, these critical values are applicable to both the dynamic OLS and the fully modified OLS. In the case of the dynamic OLS, it is necessary to choose the number of leads and lags, denoted by k. We set k to be either two or three, basically following the preceding studies including Stock and Watson [18]. According to the Schwarz information criterion, $k = 2$ is most preferred for out dataset.

As an example of the above procedure, Figure 4 plots F statistics for each data point, together with 5% critical values of the Sup F test of the case of the constancy of β based on the dynamic OLS estimation with $k = 2$. As this figure clearly shows, the highest F statistics at January 1995 well exceeds 5% critical values of the Sup F test. This result therefore implies that the constancy of interest semi-elasticity is rejected strongly.

Tests based on dynamic OLS The upper panel of Table 2 summarizes the test results for parameter stability for the case of the dynamic OLS with $k = 2$. In terms of pure structural changes, both the Sup F and the Mean F fail to reject parameter stability for the three monetary aggregates. As shown by the two rows denoted by (4) in the upper panel, however, the constancy of interest semi-elasticity is rejected for all monetary aggregates at the 10% level of significance by the Mean F test, and is rejected at the 5% level by the Sup F test. In addition, F statistics is the highest at January 1995 for all monetary aggregates.

The lower panel of Table 2, on the other hand, reports the test results for the case of the dynamic OLS with $k = 3$. The overall results are fairly similar to those of $k = 2$ with the following exceptions. First, even the Mean F rejects the constancy of interest semi-elasticity for both M2+CD and currency at the 5% level of significance. Second, F statistics is the highest at September 1994 for all monetary aggregates..

As Hansen [8] emphasizes, it would be inappropriate to conclude that, based on the rejection of the Sup F test, there are two cointegrating regimes separated by the data point with the highest F statistics. It is particularly so for the case without any prior knowledge about break points. In other words, the rejection result would allow for various kinds of alternative hypotheses. Prior to the empirical investigation, however, we had the legitimate expectation that a structural break should occur around 1995 when the BOJ implemented the low interest rate policy. Given this expectation, one of the most natural possibilities would be that a structural break occurred at the data point with the highest F statistics, as far as the Sup F test indicates that it is close enough to the year 1995. We pursue this possibility hereinafter.

Table 3 reports the results of the residual based tests that assume that there are two cointegrating regimes separated by the data point with the highest F statistics. The reported tests are based on the dynamic OLS estimation, while critical values are available from Phillips and Ouliaris [14]. Most tests fail to reject no cointegration for the full sample, or the period between 1985 and 1999. For the first sub-sample period, however, the Phillips-Perron tests reject no cointegration, and accordingly support cointegration for the period between 1985 and 1994.

On the other hand, the cointegrating relationship may not exist for the second sub-sample period. We have two reservations for this negative result. First, if critical values are based on not Phillips and Ouliaris [14], but the augmented Dickey Fuller test (Dickey and Fuller [5]), no cointegration is rejected at the 10% level of significance. Second, provided that the second sub-sample is rather small, the above residual based test may be subject to a lack of the statistical power.

As Table 4 reports the parameters of money demand functions estimated by the dynamic OLS. There are some interesting observations about parameter estimates. First, money demand was much more elastic with respect to nominal interest rates in the second period than in the first period. In particular, the absolute value of interest semi-elasticity of M1 increased dramatically. In the case of $k = 2$, the estimated β changes from -0.038 to -0.216. With a more recent sample period, as shown later, interest semi-elasticity is even more elastic. As reported in Table 8 (the second panel), the dynamic OLS estimate of β with $k = 2$ is -0.592 for the sub-sample between July 1995 and December 1999.

Second, compared with the existing empirical results of M1 demand functions, money demand is remarkably interest-elastic in the above second sub-sample. For example, Miyao [13] reports that interest semi-elasticity is -0.07 for the sample period between 1980 and 1996. Applying several estimation methods on the postwar US data, Ball [2] demonstrates that interest semi-elasticity is around -0.05. Among the existing empirical literature, we cannot find any results of point estimates of interest semi-elasticity ranging from -0.2 to -0.6.

Third, income elasticity is often estimated to be imprecise with large standard errors. In any sample period, full or sub, the estimated α is not significantly different from zero for M1 demand functions. This result about income elasticity contrasts with that based on the fully modified OLS. As reported in Table 8 (the fourth panel), the fully-modified-OLS-estimated α is with much smaller standard errors, and close to unity for the first sub-sample. On the other hand, the fully modified OLS estimate is not statistically different from zero for the second sub-sample either.

Tests based on fully modified OLS We now report the test and estimation results based on the fully modified OLS[3] , and carefully compare those with the above results based on the dynamic OLS. Table 5 reports the test results of parameter instability. As shown in the two rows denoted by (1), both the Sup F and the mean F tests indicates that there are pure structural changes for all monetary aggregates at the 5% level of significance. More concretely, the data point with the highest F statistics is July 1995 for both M1 and currency, and December 1994 for M2+CD. In terms of M1 demand functions, the fully modified OLS indicates that the break point occurred six months after what is implied by the dynamic OLS. According to the test of partial structural changes, the instability of interest semi-elasticity is the most responsible for the pure structural change among the three parameters, constant, α, and β.

Table 6 reports the results of residual based test for the presence of cointegration by assuming two cointegrating regimes separated by the data point with the highest F statistics of pure structural changes. The results are very similar to those from the dynamic OLS; cointegration is supported for the first sub-sample, while it is not for the second sub-sample. Then, we apply the same reservations as before to the negative result of the second sub-sample.

Table 7 reports parameter estimates of money demand functions by the fully modified OLS. In addition, Table 8 compares parameter estimates among the standard OLS, the dynamic OLS, and the fully modified OLS, for M1 demand functions with a break at July 1995. There are two noticeable differences in the estimation results between the fully modified OLS and the dynamic OLS. First, interest semi-elasticity for the second sub-

[3] Professor Bruce Hansen kindly provided us with the GAUSS code of the fully modified OLS estimation.

sample is even more negative. In the case of M1, for example, the fully modified OLS estimate is -0.415, while the dynamic OLS estimate with $k = 2$ is -0.592. Second, income elasticity for the first sub-sample is estimated with smaller standard errors. In the case of M1, income elasticity is close to unity in the first sub-sample. In the second sub-sample, like in the dynamic OLS case, income elasticity is not statistically different from zero.

As the preceding estimation results suggest, both the dynamic OLS and the fully modified OLS indicate the structural change in money demand functions as well as the presence of the highly interest-elastic money demand under the low interest rate policy. In particular, the latter OLS offers more convincing evidence for almost horizontal M1 demand functions. First, the implied break point, July 1995, is closer to September 1995 when the BOJ started to develop the low interest rate policy. Second, the estimated interest semi-elasticity is even more negative. Third, as discussed in Section 2, almost zero impacts of real outputs on the real money balance are consistent with the implication of highly interest-elastic money demand.

3.3. **Short-run responses to changes in money supply** The empirical results of the previous subsection strongly suggest that money demand was extremely interest elastic, or money demand functions were almost horizontal under the low interest rate policy. As shown in Section 2, highly interest-elastic demand serves as a necessary condition for nominal rigidity; almost horizontal money demand functions may make nominal prices irresponsive to changes in money supply. In this subsection, we empirically examine whether such nominal rigidity indeed emerged due to the above-detected highly interest-elastic money demand.

To differentiate the effect of money supply on nominal prices between the two periods, we estimate the following equation:

$$\begin{aligned}\Delta p_t &= \text{constant} + \gamma_0^c I_{date<break} + \gamma_0^m I_{date<break} \Delta m_t + \gamma_1^m I_{date \geq break} \Delta m_t \\ &+ \gamma_0^y I_{date<break} \Delta y_t + \gamma_1^y I_{date \geq break} \Delta y_t \\ &+ \mu_0 I_{date<break} \epsilon_{t-1} + \mu_1 I_{date \geq break} \epsilon_{t-1} + \xi_t, \qquad (3)\end{aligned}$$

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

dynamic OLS estimation of equation (2), serves as an error correction term.

If nominal rigidity was motivated by the low interest rate policy, then we expect $\gamma_0^m > 0$ and $\gamma_1^m = 0$. In addition, we may have $\gamma_0^m \neq \gamma_1^m$ in the emergence of nominal rigidity. With respect to coefficients on error correction terms, we expect $\mu_0 > 0$ and $\mu_1 > 0$ if there is a rather quick recovery to the long-run equilibrium.

The estimation of equation (3) requires instrumental variable estimation to control for simultaneous biases. We include as instrumental variables, constant terms, lagged changes in money supply, and lagged nominal price increases. The number of lags is controlled from one to four. Table 9-1 reports the estimation results for the case of the dynamic OLS with $k = 2$. In this estimation, we set January 1995 as a break for all monetary aggregates. Because the estimated constant term of equation (3), if included, is not significantly different from zero, this table reports the case without any constant term.

The most important finding is that γ_0^m is significantly positive, while γ_1^m is not significantly different from zero. The contrast between γ_0^m and γ_1^m is the most remarkable in the case of M1 money demand; $\gamma_0^m = \gamma_1^m$ is rejected statistically as shown by the last column of Table 4. On the other hand, both γ_0^y and γ_1^y are insignificant, and the coefficients on error correction terms are also insignificant. As shown in Table 9-2, the estimation results do not change at all between the fully modified OLS case and the dynamic OLS case.

The above findings clearly suggest that changes in money supply were reflected immediately, at least partially, in nominal prices in the before-break period, but nominal prices were irresponsive to money supply changes in the after-break period. In other words, they present conclusive evidence for the emergence of nominal rigidity under the low interest rate policy.

4. Conclusion This paper has demonstrated that the low interest rate policy implemented by the BOJ during the second half of the 1990s had yielded nominal rigidity. More concretely, the long-run money demand function changed structurally, in particular in terms of interest elasticity, when the BOJ started to develop the above policy in 1995. It was under this monetary policy that money demand became highly interest elastic, and that nominal prices never responded to changes in money supply.

These findings carry several important implications for monetary policy. First, a monetary policy characterized by extremely low interest rates may weaken substantially the relationship between monetary aggregates and nominal prices. Stating otherwise, nominal prices may fail to aggregate the information concerning the demand and supply condition of money markets. In such a low interest rate situation, the conventional monetary policy whereby the timing of changes in monetary policy is linked basically with the current condition of inflation, may not work properly. On this ground, the BOJ announcement of April 2000 that linked the lifting of the zero interest rate policy with the resolution of concerns over a deflationary pressure, is not necessarily justifiable.

Second, it is fairly difficult for any expansionary monetary policy to trigger inflation, if money demand is extremely strong at almost zero interest rates. Without any additional measure, therefore, a monetary policy accompanied by zero interest rates is unlikely to yield any favorable results on business cycles, and it tends to simply end in nominal rigidity. To avoid the above policy ineffectiveness, a central bank should implement unorthodox policy measures to guide the expectations about nominal prices borne by market participants in a desirable direction. In any case, however, such measures would require of a central bank, great policy sophistication.

More fundamentally, what we have learned from the Japanese experience of monetary policy during the latter half of the 1990s is that it is extremely difficult to control business cycles under low nominal interest rates. It may be more reasonable to control real interest rates directly and business cycles eventually, by keeping nominal interest rates to some extent far from zero percent, and thereby avoiding nominal rigidity. In other words, 2% interest rate together with 4% inflation may be preferred to zero interest rate combined with zero inflation.

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Table 1: Unit Root Tests on $m_t - p_t$, y_t, and i_t
1985:8-1999:12

	Test Statistics			
	$ADF - t$	$ADF - Z$	$PP - Z_t$	$PP - Z_\alpha$
real M1	1.36	0.79	1.19	0.76
real M2+CD	-2.59	-3.09	-4.94**	-2.57
real Currency	-0.49	-0.70	-0.87	-1.34
y_t	-2.06	-4.64	-2.02	-4.71
i_t	-1.89	-3.66	-0.77	-1.35

1. The ADF_t and ADF_Z indicate the augmented Dickey Fuller t and Z statistics respectively (Dickey and Fuller, 1979).
2. The $PP - Z_t$ and $PP - Z_\alpha$ indicate the Phillips-Perron Z_t and Z_α statistics respectively (Phillips and Perron 1988).
3. * and ** indicate the 5% and 1% levels of significance respectively.

Table 2: Tests for Parameter Instability of Money Demand Equations by Dynamic OLS

$k=2$	M1	M2+CD	Currency	5% c.v.
Sup F				
(1)	6.86	7.34	7.08	14.8
	(1995:1)	(1995:1)	(1995:1)	
(2)	7.89	7.17	7.66	10.09
	(1994:2)	(1994:2)	(1994:2)	
(3)	7.87	7.13	7.62	9.78
	(1994:2)	(1994:2)	(1994:2)	
(4)	19.83**	20.58**	20.09**	9.78
	(1995:1)	(1995:1)	(1995:1)	
Mean F				
(1)	1.07	1.20	1.15	6.17
(2)	1.56	1.79	1.72	2.44
(3)	1.57	1.80	1.73	2.52
(4)	2.02*	2.43*	2.24*	2.52

$k=3$	M1	M2+CD	Currency	5% c.v.
Sup F				
(1)	7.99	8.44	8.21	14.8
	(1994:9)	(1994:9)	(1994:9)	
(2)	9.03	8.94	8.76	10.09
	(1994:2)	(1996:12)	(1994:2)	
(3)	9.02	8.06	8.73	9.78
	(1994:2)	(1994:2)	(1994:2)	
(4)	23.19**	25.25**	24.24**	9.78
	(1994:9)	(1994:9)	(1994:9)	
Mean F				
(1)	1.20	1.36	1.30	6.17
(2)	1.62	1.88	1.80	2.44
(3)	1.63	1.89	1.81	2.52
(4)	2.41*	2.90**	2.68**	2.52

1. Tests are based on the dynamic OLS proposed by Stock and Watson (1993) with the number of leads and lags equal to two or three.
2. Critical values refer to Kuo (1998) for a partial structural change, and to Hansen (1992) for a pure structural change.
3. In each panel, the first raw denoted by (1) refers to testing the whole cointegrating vector, (2) to tesing the intercept, (3) to testing the coefficient on y_t, and (4) to testing the coefficient on i_t.
4. Data points with the highest F statistics are reported in parentheses.
5. * and ** indicate the 10% and 5% levels of significance respectively.

Table 3: Residual Based Tests for Cointegration by Dynamic OLS

No. of Leads and Lags	m_t	Test Statistics $ADF-t$	$PP-Z_t$	$PP-Z_\alpha$
$k=2$	M1			
	1985:8-1999:12	-0.95	-1.06	-3.94
	1985:8-1994:12	-3.22*	-6.66**	-68.72**
	1995:1-1999:12	-1.44	-1.70	-6.18
	M2+CD			
	1985:8-1999:12	-1.92	-3.06	-16.84
	1985:8-1994:12	-2.14	-3.79**	-24.89**
	1995:1-1999:12	-1.81	-1.82	-7.43
	Currency			
	1985:8-1999:12	-2.56	-3.86**	-28.79**
	1985:8-1994:12	-3.18*	-6.38**	-61.69**
	1995:1-1999:12	2.28	-1.87	-8.45
$k=3$	M1			
	1985:8-1999:12	-0.90	-1.02	-3.73
	1985:8-1994:12	-3.67**	-6.43**	-64.71**
	1995:1-1999:12	-1.39	-1.55	-5.70
	M2+CD			
	1985:8-1999:12	-1.91	-3.07	-16.93
	1985:8-1994:12	-2.14	-3.71**	-24.18**
	1995:1-1999:12	-1.88	-1.86	-8.04
	Currency			
	1985:8-1999:12	-2.51	-3.79**	-27.99**
	1985:8-1994:12	-3.12	-6.22**	-58.26**
	1995:1-1999:12	0.31	-2.08	-10.18

1. The ADF_t indicates the augmented Dickey Fuller t and Z statistics respectively (Dickey and Fuller, 1979).
2. The $PP-Z_t$ and $PP-Z_\alpha$ indicate the Phillips-Perron Z_t and Z_α statistics respectively (Phillips and Perron 1988).
3. * and ** indicate the 10% and 5% levels of significance respectively.
4. Critical values are based on Phillips and Ouliaris (1990).

Table 4: Parameter Estimates of Money Demand Equations by Dynamic OLS

No. of Leads and Lags	m_t	Point Estimates (Standard Errors) constant	α	β
$k = 2$	M1			
	1985:8-1999:12	3.633(56.52)	1.330(12.28)	-0.065(0.323)
	1985:8-1994:12	4.052(22.29)	1.205(4.950)	-0.038(0.201)
	1995:1-1999:12	14.18(3.037)	-0.931(0.666)	-0.216(0.061)
	M2+CD			
	1985:8-1999:12	3.779(2.340)	1.554(0.508)	-0.025(0.013)
	1985:8-1994:12	2.759(2.614)	1.783(0.580)	-0.030(0.023)
	1995:1-1999:12	13.27(3.216)	-0.498(0.698)	-0.067(0.064)
	Currency			
	1985:8-1999:12	0.515(1.664)	1.705(0.362)	-0.053(0.009)
	1985:8-1994:12	-0.108(2.003)	1.828(0.444)	-0.042(0.018)
	1995:1-1999:12	12.71(2.158)	-0.922(0.468)	-0.138(0.043)
$k = 3$	M1			
	1985:8-1999:12	3.680(7.345)	1.319(1.597)	-0.064(0.399)
	1985:8-1994:8	3.938(4.600)	1.230(1.021)	-0.038(0.041)
	1994:9-1999:12	15.02(2.820)	-1.114(0.610)	-0.206(0.036)
	M2+CD			
	1985:8-1999:12	3.732(2.513)	1.564(0.545)	-0.024(0.013)
	1985:8-1994:8	2.660(2.912)	1.804(0.646)	-0.029(0.026)
	1994:9-1999:12	13.58(2.448)	-0.564(0.530)	-0.066(0.031)
	Currency			
	1985:8-1999:12	0.672(2.707)	1.671(0.588)	-0.052(0.014)
	1985:8-1994:8	-0.073(3.654)	1.820(0.811)	-0.041(0.032)
	1994:9-1999:12	13.30(2.370)	-1.050(0.513)	-0.149(0.030)

1. The estimation method is based on the dynamic OLS proposed by Stock and Watson (1993) with the number of leads and lags equal to two or three.
2. Standard errors are based on AR(2) spectral estimators for $k = 2$, and AR(3) estimators for $k = 3$.

Table 5: Tests for Parameter Instability of Money Demand Equations by Fully Modified OLS

	M1	M2+CD	Currency	5% c.v.
Sup F				
(1)	74.74**	31.87**	44.93**	14.8
	(1995:7)	(1994:12)	(1995:7)	
(2)	6.21	25.03**	9.33	10.09
	(1991:5)	(1989:2)	(1994:2)	
(3)	6.32	25.06**	9.32	9.78
	(1991:5)	(1989:2)	(1997:10)	
(4)	33.90**	26.31**	17.28**	9.78
	(1995:6)	(1996:12)	(1995:7)	
Mean F				
(1)	17.61**	20.62**	12.15**	6.17
(2)	1.62	5.49**	0.95	2.44
(3)	1.64	5.39**	0.93	2.52
(4)	7.83**	3.91**	3.29**	2.52

1. Tests are based on the fully modified OLS proposed by Hansen (1993).
2. Critical values refer to Kuo (1998) for a partial structural change, and to Hansen (1992) for a pure structural change.
3. In each panel, the first raw denoted by (1) refers to testing the whole cointegrating vector; (2) to tesing the intercept, (3) to testing the coefficient on y_t, (4) to testing the coefficient on i_t,
4. Data points with the highest F statistics are reported in parentheses.
5. * and ** indicate the 10% and 5% levels of significance respectively.

Table 6: Residual Based Tests for Cointegration by Fully Modified OLS

	$ADF-t$	$PP-Z_t$	$PP-Z_\alpha$
M1			
1985:8-1999:12	-2.53	-2.28	-8.26
1985:8-1995:6	-3.32*	-6.89**	-73.06**
1995:7-1999:12	-1.45	-1.57	-5.27
M2+CD			
1985:8-1999:12	-1.57	-2.29	-7.82
1985:8-1994:11	-1.75	-2.77	-13.81
1994:12-1999:12	-3.66**	-3.59**	-9.66
Currency			
1985:8-1999:12	-3.30	-4.00**	-30.43**
1985:8-1995:6	-3.44**	-6.49**	-63.78**
1995:7-1999:12	-2.99	-3.12	-16.31

1. The ADF_t indicates the augmented Dickey Fuller t and Z statistics respectively (Dickey and Fuller, 1979), while the $PP-Z_t$ and $PP-Z_\alpha$ indicate the Phillips-Perron Z_t and Z_α statistics respectively (Phillips and Perron 1988).
2. * and ** indicate the 10% and 5% levels of significance respectively.
3. Critical values are based on Phillips and Ouliaris (1990).

Table 7: Parameter Estimates of Money Demand Equations by Fully Modified OLS

	constant	α	β
M1			
1985:8-1999:12	-2.812(4.776)	2.744(1.040)	-0.075(0.029)
1985:8-1995:6	4.100(0.238)	1.194(0.052)	-0.037(0.002)
1995:7-1999:12	12.26(1.790)	-0.501(0.390)	-0.415(0.072)
M2+CD			
1985:8-1999:12	5.726(1.092)	1.130(0.238)	-0.026(0.006)
1985:8-1994:11	3.707(0.574)	1.582(0.128)	-0.036(0.005)
1994:12-1999:12	13.14(2.445)	-0.460(0.529)	-0.138(0.037)
Currency			
1985:8-1999:12	0.133(1.235)	1.790(0.269)	-0.055(0.007)
1985:8-1995:6	-0.008(0.514)	1.804(0.114)	-0.040(0.004)
1995:7-1999:12	8.081(2.040)	0.100(0.444)	-0.421(0.082)

1. Estimators are based on Hansen(1992)'s fully modified 0LS.
2. Standard errors are in parentheses.

Table 8: Parameter Estimates of M1 Demand Equations Based on Alternative Specifications with a Break at July 1995

	Period	Point Estimates (Standard Errors)		
		constant	α	β
Standard OLS	1985:8-1999:12	3.627(0.424)	1.330(0.092)	-0.064(0.002)
	1985:8-1995:6	4.285(0.122)	1.152(0.027)	-0.035(0.001)
	1995:7-1999:12	9.957(0.999)	0.002(0.217)	-0.475(0.037)
Dynamic OLS($k = 2$)	1985:8-1999:12	3.633(56.52)	1.330(12.28)	-0.065(0.323)
	1985:8-1995:6	4.108(16.17)	1.193(3.568)	-0.037(0.138)
	1995:7-1999:12	7.278(2.080)	0.592(0.454)	-0.592(0.078)
Dynamic OLS($k = 3$)	1985:8-1999:12	3.680(7.345)	1.319(1.597)	-0.064(0.399)
	1985:8-1995:6	3.995(6.625)	1.217(1.456)	-0.037(0.052)
	1995:7-1999:12	6.17(1.887)	0.831(0.412)	-0.609(0.070)
Fully Modified OLS	1985:8-1999:12	-2.812(4.776)	2.744(1.040)	-0.075(0.029)
	1985:8-1995:6	4.100(0.239)	1.195(0.053)	-0.037(0.002)
	1995:7-1999:12	12.26(1.790)	-0.501(0.390)	-0.415(0.072)

Table 9-1: Parameter Estimates of Error Correction Type Models Based on Dynamic OLS ($k = 2$)

m_t	Lags	γ_0^m	γ_1^m	γ_0^y	γ_1^y	μ_0	μ_1	F
M1								
	1	0.274	0.022	-0.021	0.014	0.056	-0.004	3.679
		(0.148)	(0.071)	(0.580)	(0.083)	(0.363)	(0.034)	(0.056)
	2	0.263	0.027	-0.007	0.036	0.041	0.001	7.644
		(0.087)	(0.060)	(0.133)	(0.049)	(0.077)	(0.019)	(0.006)
	3	0.176	0.014	0.060	0.028	-0.015	-0.005	5.341
		(0.059)	(0.054)	(0.044)	(0.038)	(0.028)	(0.013)	(0.022)
	4	0.144	0.020	0.033	0.020	0.004	-0.002	3.782
		(0.051)	(0.051)	(0.034)	(0.035)	(0.023)	(0.012)	(0.053)
M2+CD								
	1	0.270	0.122	-0.016	0.072	0.035	0.074	0.727
		(0.076)	(0.181)	(0.054)	(0.118)	(0.024)	(0.177)	(0.395)
	2	0.259	0.095	-0.028	0.043	0.031	0.022	1.099
		(0.070)	(0.156)	(0.048)	(0.079)	(0.021)	(0.130)	(0.295)
	3	0.225	0.110	-0.004	0.064	0.018	0.055	0.604
		(0.062)	(0.152)	(0.036)	(0.045)	(0.016)	(0.080)	(0.437)
	4	0.230	0.108	-0.011	0.067	0.021	0.053	0.753
		(0.060)	(0.149)	(0.032)	(0.036)	(0.014)	(0.067)	(0.386)
Currency								
	1	0.220	0.309	-0.014	0.102	0.077	0.127	0.055
		(0.187)	(1.324)	(0.190)	(0.340)	(0.081)	(0.560)	(0.814)
	2	0.029	0.044	0.056	0.042	0.009	0.022	0.044
		(0.024)	(0.074)	(0.057)	(0.041)	(0.017)	(0.044)	(0.833)
	3	0.031	0.046	0.048	0.050	0.011	0.021	0.054
		(0.022)	(0.071)	(0.039)	(0.039)	(0.014)	(0.043)	(0.815)
	4	0.029	-0.003	0.036	0.024	0.012	0.000	0.333
		(0.020)	(0.058)	(0.032)	(0.034)	(0.012)	(0.037)	(0.564)

1. The error correction model is spedified 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 $z_t = (m-p)_t - (\text{constant} + \alpha y_t + \beta i_t)$ using the estimationn result of the dynamic OLS with $k = 2$.
2. The number of lags for instrumental varialbles is cotroled from one to four. Insutrumental variables include constant, Δm_t, and Δy_t.
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 9-2: **Parameter Estimates of Error Correction Type Models Based on Fully Modified OLS**

m_t	Lags	γ_0^m	γ_1^m	γ_0^y	γ_1^y	μ_0	μ_1	F
M1								
	1	0.242	0.037	0.028	0.013	0.026	-0.003	1.850
		(0.168)	(0.056)	(0.457)	(0.053)	(0.289)	(0.032)	(0.175)
	2	0.244	0.037	-0.006	0.030	0.043	0.006	5.207
		(0.096)	(0.055)	(0.154)	(0.036)	(0.092)	(0.019)	(0.023)
	3	0.163	0.024	0.063	0.040	-0.015	-0.001	4.221
		(0.058)	(0.051)	(0.046)	(0.032)	(0.029)	(0.015)	(0.041)
	4	0.142	0.029	0.038	0.035	0.003	0.002	3.269
		(0.052)	(0.048)	(0.035)	(0.030)	(0.023)	(0.014)	(0.079)
M2+CD								
	1	0.251	-0.026	-0.003	0.097	0.025	0.138	1.990
		(0.103)	(0.345)	(0.069)	(0.198)	(0.027)	(0.370)	(0.160)
	2	0.242	-0.020	-0.012	0.085	0.025	0.124	2.903
		(0.095)	(0.256)	(0.061)	(0.096)	(0.023)	(0.185)	(0.090)
	3	0.235	0.023	-0.009	0.060	0.021	0.048	2.389
		(0.062)	(0.147)	(0.036)	(0.038)	(0.014)	(0.054)	(0.124)
	4	0.229	0.034	-0.011	0.054	0.020	0.023	2.197
		(0.055)	(0.132)	(0.030)	(0.029)	(0.012)	(0.031)	(0.140)
Currency								
	1	0.195	0.044	0.015	-0.041	0.069	0.050	4.145
		(0.127)	(0.147)	(0.125)	(0.240)	(0.051)	(0.182)	(0.043)
	2	0.032	0.042	0.044	0.001	0.016	0.013	0.052
		(0.022)	(0.046)	(0.052)	(0.071)	(0.015)	(0.036)	(0.818)
	3	0.030	0.026	0.051	0.036	0.015	-0.005	0.008
		(0.020)	(0.037)	(0.038)	(0.035)	(0.013)	(0.016)	(0.928)
	4	0.029	-0.002	0.040	0.021	0.015	-0.003	0.819
		(0.019)	(0.033)	(0.031)	(0.033)	(0.012)	(0.015)	(0.366)

1. The error correction model is spedified 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 $z_t = (m-p)_t - (\text{constant} + \alpha y_t + \beta i_t)$ using the estimationn result of the fully modified OLS.
2. The number of lags for instrumental varialbles is controlled from one to four. Insutrumental variables include constant, Δm_t, and Δy_t.
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.

Figure 1: Overnight Call Rates (Interbank Rates, %)

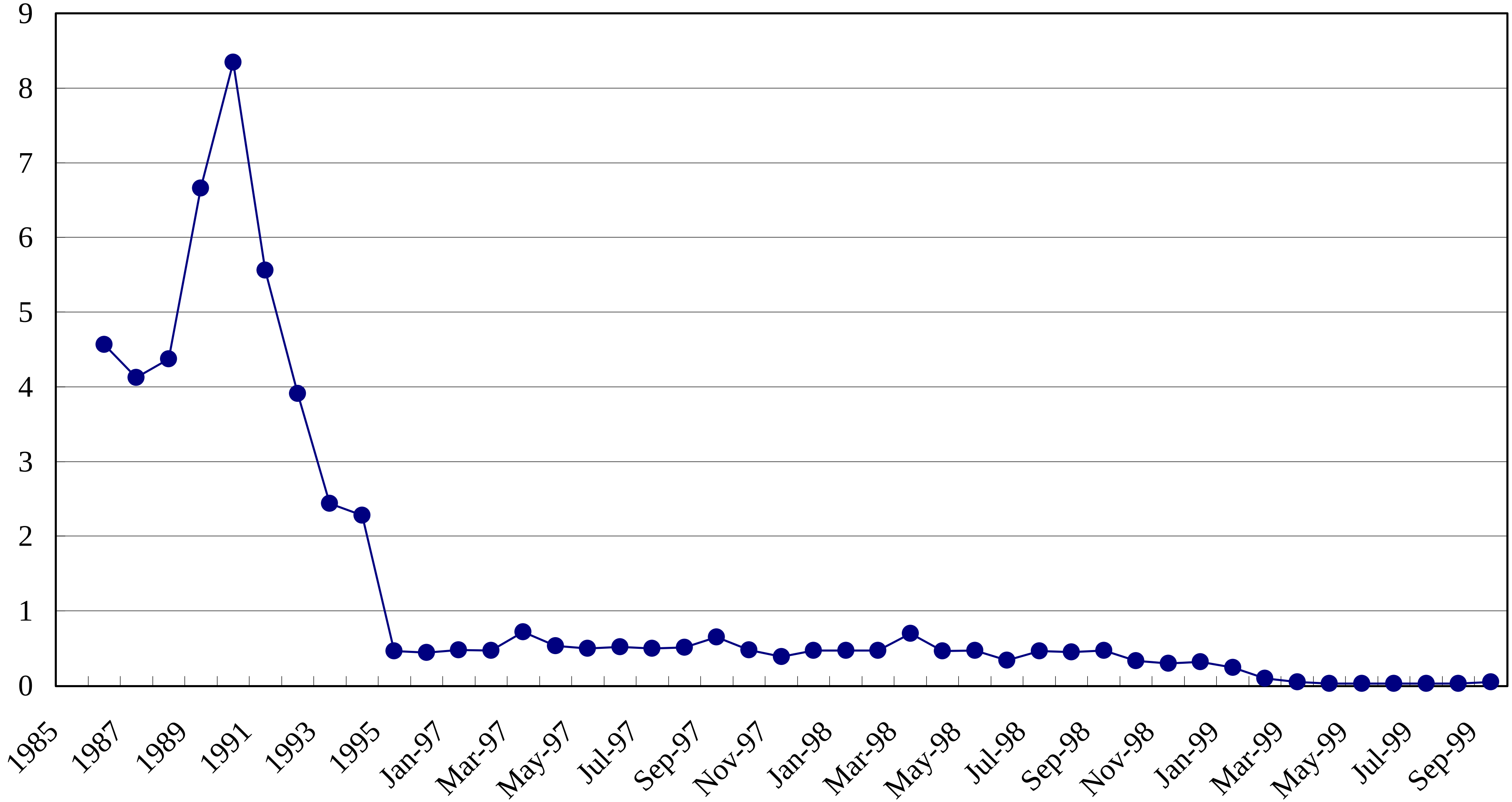
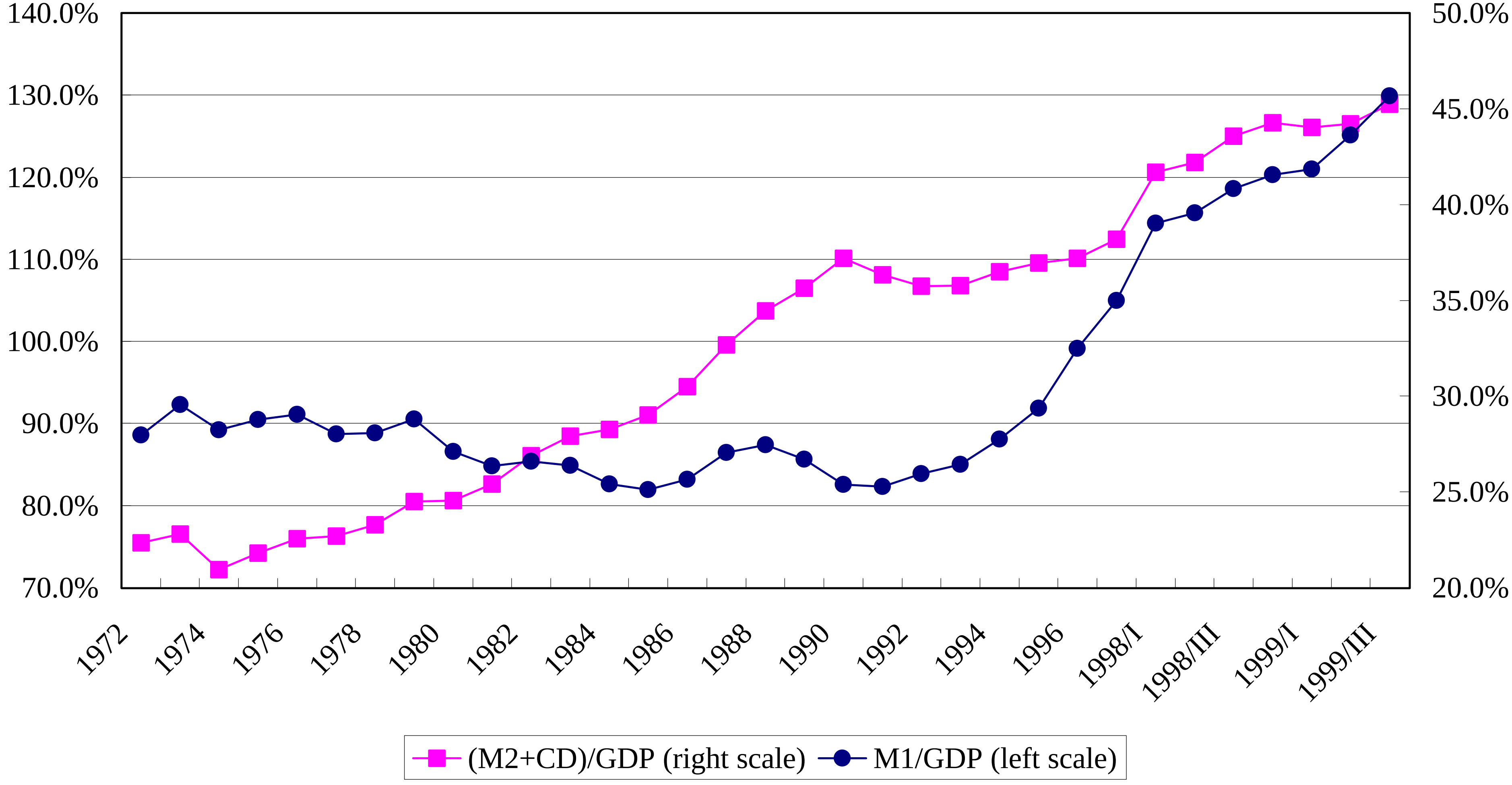


Figure 2: The Ratio of Monetary Aggregates Relative to Nominal GDP (seasonally adjusted for quarterly data from 1998/I)



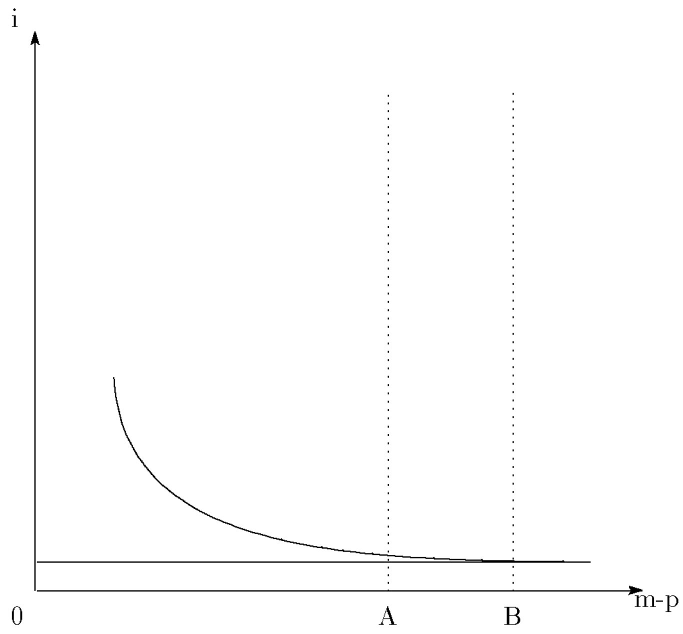


Figure 3: **Highly Interest-Elastic Money Demand**

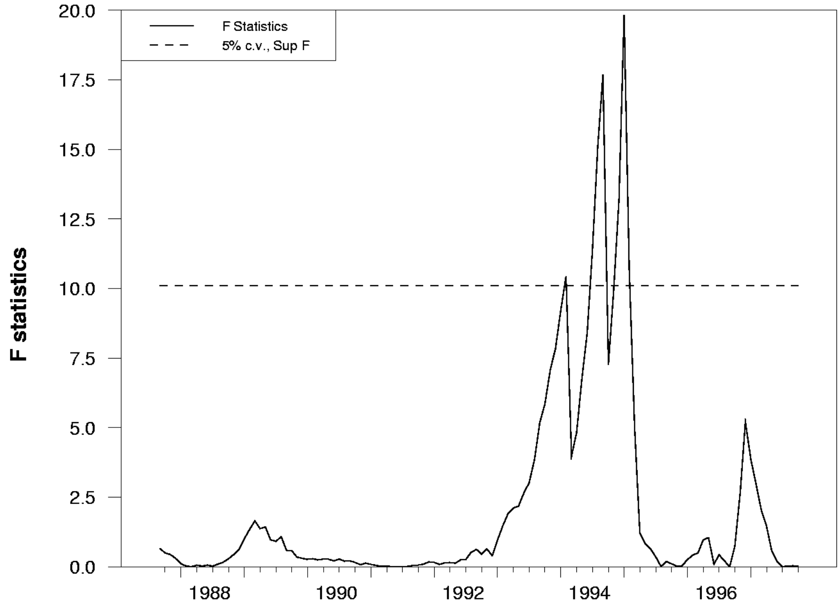


Figure 4: **Testing Structural Breaks (M1, Dynamic OLS with $k = 2$)**