On Long-Run Monetary Neutrality in Japan

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This paper comprehensively investigates long-run monetary neutrality in Japan, with due consideration to the order of integration of the money stock and real output, mainly using long-term time-series data retroactively available from the Meiji Period (1868–1912). The empirical results indicate little evidence against the long-run neutrality of money (especially defined as M2) with respect to real GNP. In addition, such findings are robust to a wide range of identifying assumptions.

Keywords: Long-run monetary neutrality; Long-term time-series data; Structural changes; Unit root tests; Bivariate structural vector autoregression (VAR)
JEL Classification: C22, C32, E40, E51

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The authors would like to thank Professors Yuzo Honda, Yukinobu Kitamura, Ryuzo Miyao, Masao Ogaki, and Etsuro Shioji, and the staff of the Monetary Affairs Department, Research and Statistics Department, and the Institute for Monetary and Economic Studies (IMES) of the Bank of Japan for their valuable comments. Regardless, the opinions expressed in this paper are those of the authors, and do not represent the official views of the Bank of Japan or IMES.

I. Introduction

This paper comprehensively investigates long-run monetary neutrality in Japan, with due consideration to the order of integration of the money stock and real output, mainly using long-term time-series data retroactively available from the Meiji Period (1868–1912).

The issue of whether or not changes in the money stock influence real variables, such as real output and the unemployment rate, has long been a major topic in monetary economics. If changes in the money stock and changes in real variables are independent, then money is deemed to be neutral, and if not then money is non-neutral. Specifically, long-run monetary neutrality is said to exist if a permanent increase or decrease in the nominal money stock does not have any long-term effect on the level of output.

The independence between real and nominal variables in the long run is a widely accepted notion in economic theory, the so-called classical dichotomy. In theoretical analyses of long-term consequences in the economy, including economic growth, long-run monetary neutrality is generally assumed.[1] In empirical analyses of business cycles, long-run monetary neutrality is often employed as an identifying restriction in the structural vector autoregression (VAR) model, proposed by Blanchard and Quah (1989). Thus, it is deemed important to reexamine the empirical validity of long-run monetary neutrality, based on actually observed data.

Empirical examinations to confirm long-run monetary neutrality have been conducted for many years. During the 1960s, many adopted the method of regressing output by the money stock.[2] In response to these efforts, Lucas (1972, 1973) and Sargent (1971, 1976) pointed out the problems with investigating the long-run neutrality of nominal variables using reduced-form models. Lucas employs a simple rational expectations macroeconomic model to demonstrate that analyses using reduced-form models lead to the erroneous conclusion that long-run monetary neutrality does not exist, even in cases where it does.[3]

The critique offered by Lucas and Sargent is closely related to the stationarity of data. To verify long-run monetary neutrality, changes in the money stock are required to contain a permanent component that is independent from any changes in output. If the money stock does not contain a unit root, it does not include any permanent changes, and therefore its long-run neutrality cannot be verified.

Yet the presence of a unit root in the money stock is also not sufficient. For example, when a central bank takes an endogenous policy response, changes in the money stock are not independent from real output. This is because in such cases simply looking at the reduced-form information does not enable us to examine the long-term effects of "pure" variations in the money stock. To verify long-run neutrality, it is essential to identify shocks that are independent from changes in output.

1. Some theoretical arguments assume monetary non-neutrality even in the long run. For example, Tobin (1965) introduces outside money into his economic growth model to show that inflation, induced by the central bank's excess supply of outside money, raises the capital-labor ratio. Obstfeld and Rogoff (1995) employ a two-country version of the dynamic equilibrium model to show that short-term changes in the money supply influence production and consumption in the long run through fluctuations in trade imbalances.
2. The simple regression of output by the money stock is known as the St. Louis equation.
3. See the Appendix for a summary of Lucas (1972). See Okina (1986) for an overview of empirical research in Japan regarding the relationship between the money stock and real output.

Among these points of dispute regarding the testing procedures, Fisher and Seater (1993) further explore the empirical issues regarding the time-series properties of the data. They demonstrate that the preconditions for testing long-run monetary neutrality can be established not by using the money stock alone, but rather based on the order of integration of both the money stock and real output.

On the basis of the argument in Fisher and Seater (1993), King and Watson (1997) address the empirical issues for identification of shocks. They employ a structural VAR model with *a priori* restrictions to identify exogenous shocks, thereby enabling verification of long-run monetary neutrality with a reduced-form model. They then examine the robustness of empirical evidence on long-run neutrality by testing a wide range of identifying restrictions.

In response to King and Watson (1997), the direction of research on long-run monetary neutrality has moved to verifying the robustness of empirical evidence using a diverse range of data.[4] The extension of data has been carried out mainly in two ways: the extension of time series retroactive to the period before World War II, and various alternative definitions of the money stock. Moreover, advanced procedures of unit root testing with structural breaks are also applied to reexamining empirical evidence in previous studies.

More specifically, research on further explicating long-run monetary neutrality includes the studies below. On the one hand, Weber (1994), Bullard and Keating (1995), Serletis and Krause (1996), and Serletis and Koustas (1998) use cross-country data.[5] Among these, Weber (1994) and Bullard and Keating (1995) both use quarterly data for the period after World War II. On the other hand, Serletis and Krause (1996) and Serletis and Koustas (1998) employ long-term time-series data that include the prewar period. These studies reveal that long-run monetary neutrality generally holds for wide-ranging periods and countries.[6] Serletis and Koustas (1998), however, conclude that long-run neutrality cannot be tested in Japan, because their unit root tests with a structural break show that the monetary aggregates do not include unit roots.

In spite of the progress in research on long-run monetary neutrality, only a very limited number of comprehensive studies are available in Japan.[7] Given that highly accurate Japanese data are readily available from the Meiji Period, it is possible to test monetary neutrality in Japan by using long-term time-series data that include various definitions of monetary aggregates. As noted in Weber (1994), longer observation periods are deemed more desirable in testing long-run monetary neutrality, since they include more information on long-run fluctuations. Moreover, considering the

4. Some studies focus on types of long-run neutrality other than monetary neutrality, such as the Fisher effect and the slope of the long-run Phillips curve. For example, Koustas and Serletis (1999) examine the Fisher effect, and King and Watson (1994) examine whether or not the long-run Phillips curve is vertical.

5. Weber (1994) uses data across the G-7 nations. Bullard and Keating (1995) use data for the 58 countries where readily available data exist. Serletis and Krause (1996) and Serletis and Koustas (1998) use data for nine countries: Australia, Denmark, Germany, Italy, Japan, Norway, Sweden, the United Kingdom, and the United States.

6. However, in addition to Weber (1994), long-run monetary neutrality is also rejected in some cases employing different definitions of the money stock: Olekalns (1996) for Australia since 1900; Coe and Nason (2003) for Australia, Canada, the United Kingdom, and the United States; Jefferson (1997) for measures of both inside money and outside money; and Serletis and Koustas (2001) for broadly defined monetary aggregates.

7. Yamada (1997) shows that monetary neutrality holds in terms of real output, by applying Fisher and Seater's (1993) procedure to the Japanese quarterly data from 1957/I to 1995/I.

limited power of the unit root test and the important role of the order of integration of the data in verifying long-run monetary neutrality, it is important to apply the more sophisticated procedures of unit root testing to reexamine the time-series property of the data. In this paper, we therefore comprehensively investigate long-run monetary neutrality in Japan, with attention to the order of integration of the variables, mostly by using long-term time-series data retroactively available from the Meiji Period.

This paper is organized as follows. Section II provides a summary of empirical frameworks for testing long-run monetary neutrality. Section III explains long-term annual data for monetary aggregates and real output. Section IV examines the time-series properties of the annual data by carrying out unit root and cointegration tests. Section V presents empirical findings on long-run monetary neutrality using long-term annual data in Japan. Section VI reports the estimation results for the postwar quarterly data to supplement the empirical evidence using the long-term annual data. Section VII summarizes our empirical findings and concludes the paper.

II. Outline of Analytical Methods

In this section, we first review the relationship between long-run monetary neutrality and the order of integration of the money stock and real output. We then move on to explaining the methodology for verifying long-run monetary neutrality by identifying exogenous monetary shocks.

A. Order of Integration and Long-Run Monetary Neutrality

Fisher and Seater (1993) formulate tests for long-run neutrality in relation to the order of integration, based on the reduced-form model below.

$$\begin{aligned} a(L)\Delta^{<m>}m_t &= b(L)\Delta^{<y>}y_t + u_t, \\ d(L)\Delta^{<y>}y_t &= c(L)\Delta^{<m>}m_t + w_t. \end{aligned} \tag{1}$$

Here, y and m are respectively the natural logarithms of output and the money stock, while u and w are the error terms in the equations, $<x>$ denotes the order of integration of a variable x, Δ is a difference operator, and $a(L)$, $b(L)$, $c(L)$, and $d(L)$ are lag polynomials. The terms a_0 and d_0 are scaled to one, and the error terms u and w are i.i.d. with zero mean.

The long-run derivative of output with respect to the money stock, $LRD_{y,m}$, is defined as the equation below:

$$LRD_{y,m} \equiv \lim_{k\to\infty} \frac{\partial y_{t+k}/\partial u_t}{\partial m_{t+k}/\partial u_t}. \tag{2}$$

Long-run monetary neutrality holds when $LRD_{y,m}$ is equal to zero. Based on this formulation of $LRD_{y,m}$, Fisher and Seater (1993) relate tests for long-run neutrality to the order of integration.

First, when the order of integration of the money stock is zero ($<m> = 0$), the shock u_t does not have any permanent effect on the money stock. The denominator of $LRD_{y,m}$ thus becomes zero ($\lim_{k\to\infty} \partial m_{t+k}/\partial u_t = 0$), and long-run neutrality cannot be tested.[8]

Next, when the order of integration of the money stock is one or higher, three cases are possible. The first case is when the order of integration of the money stock is higher than that of real output. This case always supports long-run monetary neutrality, because a shock to the money stock does not have any permanent influence on real output, and the numerator of equation (2) becomes zero ($\lim_{k\to\infty} \partial y_{t+k}/\partial u_t = 0$). The second case is when the order of integration of the money stock is equal to that of real output, and they are both one or higher. This case requires the identification of shocks to the money stock, which are independent from changes in real output. Under this case, therefore, one cannot judge whether long-run monetary neutrality holds just from the order of integration. The third case is when the order of integration of the money stock is less than that of real output. This case supports long-run monetary neutrality, if shocks to the money stock do not influence the growth rate of real output.

In the meantime, monetary superneutrality holds when changes in the growth rate of the money stock do not have any long-term effects on real output, that is to say, when $LRD_{y,\Delta m}$ is equal to zero. In this case, the conditions for testing monetary superneutrality are obtained by replacing the money stock with the money stock growth in the aforementioned four cases for long-run monetary neutrality. In other words, the order of integration of the money stock growth and real output must be equal, and must be one or higher (and thus the order of integration of the money stock must be at least two).

B. Identifying Monetary Shocks

As explained in the previous subsection, exogenous shocks[9] to the money stock must be identified to test long-run monetary neutrality when the orders of integration of the money stock and real output are the same and both one or higher. King and Watson (1997) propose an empirical framework to identify shocks with a broad range of identifying restrictions, thereby checking the robustness of long-run monetary neutrality. Following King and Watson (1997), we explain the method of identifying shocks below.

When the orders of integration of the money supply and real output are both one, the bivariate reduced-form model is described by a vector moving average, expressed as the sum of two exogenous shocks (structural shocks) of the monetary shock (ϵ_t^m) and the non-monetary shock (ϵ_t^η), as shown in equation (3).

$$\begin{aligned} \Delta y_t &= \theta_{y\eta}(L)\epsilon_t^\eta + \theta_{ym}(L)\epsilon_t^m, \\ \Delta m_t &= \theta_{m\eta}(L)\epsilon_t^\eta + \theta_{mm}(L)\epsilon_t^m. \end{aligned} \tag{3}$$

8. Fisher and Seater (1993) call the data "uninformative" concerning long-run monetary neutrality when the money stock is a stationary process and does not include the information required for testing long-run monetary neutrality.
9. Shocks to the money stock are independent from the fluctuations of real output. These shocks cannot be directly observed due to various factors including endogenous monetary policy responses.

ϵ_t^m, ϵ_t^η are unrelated, and both i.i.d. Here, the constant terms are omitted to simplify the explanation.

The long-term effects of the monetary shock on output and on the money stock can be expressed, respectively, as $\sum \theta_{ym,j}\epsilon_t^m \equiv \theta_{ym}(1)\epsilon_t^m$ and $\sum \theta_{mm,j}\epsilon_t^m \equiv \theta_{mm}(1)\epsilon_t^m$. In this case, $\gamma_{ym} = \theta_{ym}(1)/\theta_{mm}(1)$ expresses the long-term elasticity of output to monetary shock. Corresponding to $LRD_{y,m}$ in the previous subsection, long-run monetary neutrality holds when $\gamma_{ym} = 0$.

Assuming that all characteristic roots for the coefficient matrix of equation (3) lie outside the unit circle, equation (3) can be inverted to yield a VAR of equation (4) below. This requires that the money stock and real output both have a unit root but are not cointegrated.

$$\begin{aligned} \alpha_{y0}\Delta y_t &= \lambda_{ym}\Delta m_t + \sum_{j=1}^{p}\alpha_{j,yy}\Delta y_{t-j} + \sum_{j=1}^{p}\alpha_{j,ym}\Delta m_{t-j} + \epsilon_t^\eta, \\ \alpha_{m0}\Delta m_t &= \lambda_{my}\Delta y_t + \sum_{j=1}^{p}\alpha_{j,my}\Delta y_{t-j} + \sum_{j=1}^{p}\alpha_{j,mm}\Delta m_{t-j} + \epsilon_t^m. \end{aligned} \tag{4}$$

Based on the specification of equation (4), the long-term elasticity of output to monetary shock γ_{ym} and that of the money stock to output shock γ_{my} are defined as equation (5) below:

$$\begin{aligned} \gamma_{ym} &= \sum_{j=1}^{p}\alpha_{j,ym}\Bigg/\left(1 - \sum_{j=1}^{p}\alpha_{j,yy}\right), \\ \gamma_{my} &= \sum_{j=1}^{p}\alpha_{j,my}\Bigg/\left(1 - \sum_{j=1}^{p}\alpha_{j,mm}\right). \end{aligned} \tag{5}$$

The bivariate structural VAR of order p, as in equation (4), has $2^2 \times (p + 1)$ unknowns in the coefficients and 3 $(= 2 \times (2 + 1)/2)$ unknowns in the covariance matrix of the residual. Meanwhile, the bivariate reduced-form VAR of order p provides estimates of 2^2p parameter values for the coefficients and three values for the covariance matrix of the reduced disturbance.

Accordingly, $2^2 = 4$ identifying restrictions must be placed to identify the structural shocks ϵ_t^m and ϵ_t^η. Standardizing α_{y0} and α_{m0} to unity and using the assumption that the structural shocks are mutually uncorrelated, in other words, that the diagonal elements of the covariance matrix are zero, the number of *a priori* restrictions is reduced to one.

King and Watson (1997) adopt two types of *a priori* restrictions: short-term and long-term restrictions. First, the short-term restriction specifies a contemporaneous relationship between endogenous variables and shocks by imposing restrictions on the short-term elasticity, such as $\lambda_{ym} = 0$ or $\lambda_{my} = 0$. The former restriction indicates short-run neutrality whereby output does not react contemporaneously to the shock to the money stock. In contrast, the latter restriction indicates the situation whereby the money stock does not contemporaneously accommodate changes in output, and output becomes the predetermined variable.

Second, the long-term restriction specifies a long-term relationship between endogenous variables and shocks by imposing restrictions on the long-term elasticity, for example, $\gamma_{my} = 0$. This is equivalent to the situation whereby the money stock does not accommodate shocks to output and therefore the general price level remains unchanged (assuming constant velocity of money). Additionally, long-run neutrality ($\gamma_{ym} = 0$) is applicable as an identifying restriction.

In the sections below, we first explain the data and estimation periods and then examine the preconditions for testing long-run neutrality based on the order of integration of each variable, as explained in the previous subsection. In this process, we conduct unit root tests with a structural break of recent years, and determine the order of integration. We then test long-run monetary neutrality by employing the procedure in identifying monetary shocks explained above.

III. Data

As for the time-series data for the money stock and real output in Japan, we compile a data set with the longest time series currently available, covering 119 years from 1885 through 2003. We use three types of the money stock: cash currency in circulation (hereafter denoted as CASH), M1, and M2. Given the limitations of the prewar period data, we use the time series for the amounts outstanding at the end of each year.

Fujino (1994) is a well-known source for Japan's long-term time-series data for the money stock.[10] The data in Fujino (1994) are consistent with the money stock data currently compiled by the Bank of Japan, retroactively available from 1955, and are thus directly linkable. Fujino (1994) does not provide any estimates for the period from 1941 through 1951, but Asakura and Nishiyama (1974) do. Utilizing these two data sets, we compiled three time series for the money stock: CASH, M1, and M2. Specifically, we utilized the series presented in Fujino (1994) for the periods 1885–1940 and 1952–54, the series presented in Asakura and Nishiyama (1974) for the 11 years from 1941 through 1951,[11] and connected these with the Bank of Japan's money stock statistics since 1955.[12]

10. Fujino (1994) estimates the prewar money stock (CASH, M1, and M2) statistics based on the *Banking Bureau Annual* published by the former Ministry of Finance. Fujino states that the *Banking Bureau Annual* "has a very high level of accuracy overall, and is not only exceptionally accurate compared with other Japanese economic statistics but peerless compared with banking statistics from other nations worldwide."

11. Asakura and Nishiyama (1974) do not adjust the data for financial institutions' cash on hand or for deposits between financial institutions, and all of their series take larger values than those in Fujino (1994). However, the ratio between the two differs significantly before and after World War II. Accordingly, we make a linear interpolation between the prewar and postwar ratios, apply this to reduce the figures in Asakura and Nishiyama (1974) so that they become consistent with those in Fujino (1994), and then use them to interpolate the missing observations during World War II. We should be cautious in that the reliability of the data source during the period from 1941 to 1950 is not that high. Fujino (1994) chooses to set the year 1951 as the initial period for his retroactive estimation of postwar money stock data, considering the sparse availability of basic statistics during and soon after World War II.

12. For M2, the time series includes CDs from 1979.

Next, the long-term output time series are sourced from Ohkawa, Takamatsu, and Yamamoto (1974), who estimate real GNP from the Meiji Period forward. From 1955, these statistics are directly linkable to the real GNP series in the System of National Accounts (SNA). In this paper, we prepare our output data using Ohkawa, Takamatsu, and Yamamoto (1974) and the SNA data. For the period from 1885 to 1954, we use the gross national output series from Ohkawa, Takamatsu, and Yamamoto (1974). This is linked with the real GNP series in the SNA for the period from 1955 to 2003. As for the SNA data, we use the previous 68SNA basis data for the period from 1955 to 1979, and the current 93SNA basis data for the period after 1980, for which 93SNA basis retroactive data series are available. We thus compile a data set for the money stock and real output covering the period from 1885 to 2003. It should be noted, however, that the data from the two years 1943 and 1945 are missing, and we make a linear interpolation for these missing observations.

Our data sets should also prove useful from the perspective of supplementing the conclusions on long-run monetary neutrality in Serletis and Koustas (1998), who utilize long-term international time-series data.

Serletis and Koustas (1998) use the prewar international long-term time-series data in Backus and Kehoe (1992) to test long-run monetary neutrality. They conclude that Japan's money stock (M2) does not have a unit root when considering a structural break, and that therefore they cannot test long-run monetary neutrality in Japan.[13] Although the data set in Backus and Kehoe (1992) covers a wide range of countries, there is room for some improvement, at least in the data for Japan.[14] Specifically, we think two points below are very important.

The first point is that M2 is the only measure of the money stock they consider. Weber (1994), Olekalns (1996), Coe and Nason (2003), Jefferson (1997), and Serletis and Koustas (2001), all of whom examined long-run monetary neutrality focusing on the definition of the money stock using postwar data sets, report that long-run monetary neutrality can sometimes be rejected when different types of money stock data are adopted. For this reason, the use of money stock data aside from M2 is considered important.[15]

The second point is that the data for the period from 1941 to 1951 are missing and that data after 1963 are not included in the sample. When time-series models are used, the data should be continuous, and need to cover periods when important events occurred, such as the sudden expansion of monetary aggregates during and after World War II and the liberalization of financial markets since the 1980s.

Considering these problems in the previous studies, we extend the coverage of the data set used in this paper. In addition to M2, we include CASH and M1 for

13. During the sample period, CDs were not issued.
14. The data set in Backus and Kehoe (1992) cover real GNP/GDP, general prices, and the money stock for 10 nations (Australia, Canada, Denmark, Germany, Italy, Japan, Norway, Sweden, the United Kingdom, and the United States) prior to World War II.
15. Weber (1994) examines long-run monetary neutrality using postwar quarterly data for the G-7 nations and reports that unit root test results are sensitive to the definitions of the money stock. More precisely, he points out that the narrowly defined money stock tends to be I(1), while the broadly defined money stock tends to be I(2). As explained later, however, using the quarterly postwar data for the money stock in Japan, CASH, M1, and M2 are all generally judged to be I(1).

the analyses. We also extend the data set up through 2003, and we interpolate the missing observations from 1941 to 1951 for the period around World War II by using an alternative data source.

IV. Preliminary Tests

In this section, we examine the time-series properties of the data with CASH, M1, and M2 as the money stock and real GNP as real output. We first conduct unit root tests on the four series. We then conduct cointegration tests on the money stock and real output. As noted in Section II, to verify long-run monetary neutrality we first need to show that the money stock and real output are both I(1) and that they are not cointegrated.

A. Unit Root Tests

Table 1 presents the results of the conventional unit root tests. We calculate three types of test statistics, the widely used augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests as well as the weighted symmetric (WS) test.[16,17] The null hypothesis that a unit root exists at the log transformed level cannot be rejected for all the

Table 1 Unit Root Tests (Sample Period: 1887–2003)

	CASH	M1	M2	Real GNP
[A] Level				
ADF	–2.076 (9)	–2.216 (7)	–2.154 (7)	–1.912 (1)
PP	–6.140	–5.315	–4.364	–4.755
WS	0.804 (12)	0.157 (8)	0.343 (5)	–1.515 (1)
[B] First difference				
ADF	–3.290* (1)	–2.838 (1)	–2.411 (12)	–4.052*** (6)
PP	–46.151***	–56.260***	–82.683***	–86.338***
WS	–2.850 (8)	–3.004* (1)	–2.598 (12)	–4.213*** (6)
[C] Second difference				
ADF	–4.884*** (8)	–5.969*** (4)	–6.216*** (4)	–5.176*** (10)
PP	–147.207***	–162.129***	–154.648***	–115.029***
WS	–5.041*** (8)	–6.126*** (4)	–6.377*** (4)	–5.293*** (10)

Notes: 1. Constant and trend terms are included in both ADF and WS tests.
2. *** and * indicate statistical significance at the 1 and 10 percent levels, respectively.
3. Figures in parentheses for ADF and WS tests show the lag order. The lag order is chosen by the general-to-specific rule: the lag order is iteratively reduced until the last coefficient becomes significant within the maximum lag order of 12.

16. While the ADF and PP tests are both widely used, Maddala and Kim (1998) recommend using alternative testing procedures because of their low power. Accordingly, in this paper we adopt the WS test, which provides improvements over the ADF and PP tests, as discussed in Maddala and Kim (1998). Maddala and Kim (1998) state that for low-frequency data the PP test is more powerful than the ADF test, so the application of the PP test is considered significant for the annual long-term time-series data used in this paper.
17. Unit root tests are conducted based on the specifications with constant and linear trend terms. The lag orders for the ADF and WS tests are decided, following Ng and Perron's (1995) general-to-specific rule, by the procedure of sequentially reducing the lag order from the maximum of 12 until the last term becomes statistically significant at the 10 percent level.

variables in all the tests. However, when taking the first difference the null hypothesis is rejected, except for the ADF test for M1 and the ADF and WS tests for M2. Moreover, when taking the second difference the null hypothesis is rejected for all the variables in all the tests. The results in Table 1 indicate that the order of integration of both CASH and real GNP is one (hereafter I(1)), and that the order of integration of M1 and M2 is either I(1) or I(2).[18,19] Because of this, as explained in Section II, if real output is I(1) and monetary aggregate series is I(2), then we need to test not for monetary neutrality, but rather for superneutrality.

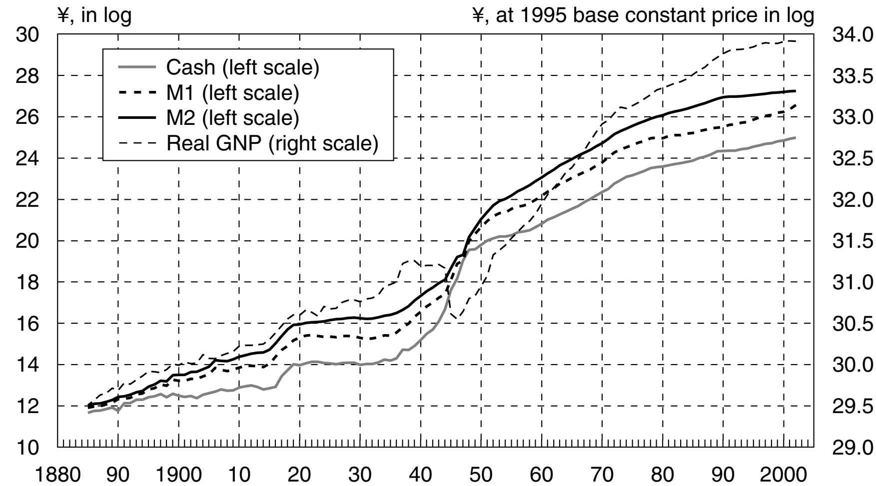
In the meantime, the movements of the time-series data plotted in Figure 1 indicate that structural changes are likely to occur around the period of World War II. Perron (1989) points out that when we conduct unit root tests ignoring potential structural changes in a deterministic trend, we are likely to judge that orders of integration are higher than they really are.

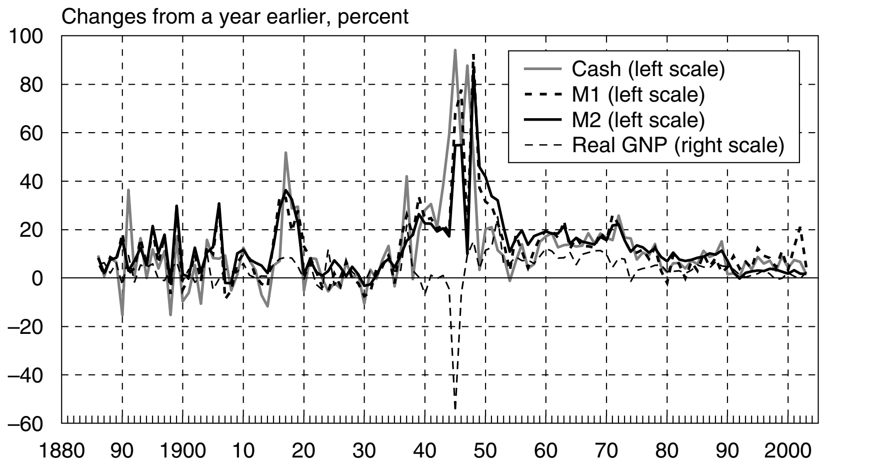
For more rigorous testing, we also apply the testing procedure in Perron (1997) for unit root tests with structural change.[20] More specifically, we test for unit roots against two alternative hypotheses regarding the stationary process: one incorporates a level shift only, and the other incorporates both a level shift and a trend break.[21] We see from Table 2 that when considering a possible structural break real GNP is still judged as I(1), while CASH and M2 are either I(1) or I(2) and M2 is I(2).[22]

In the meantime, Serletis and Koustas (1998) conclude that M2 in Japan is I(0), and thus long-run monetary neutrality cannot be tested, by using data ranging from 1885 to 1962, but with missing observations for the World War II period of 1941 to 1951. However, we see from the lower panel of Table 2 that when we use data for the same period but interpolate the World War II period observations real GNP is judged as I(1) and M2 as I(2), while CASH and M1 are likely to be I(0).

18. Following the proposal in Maddala and Kim (1998), we conduct our confirmatory analysis using stationarity as null to examine the robustness of our conclusions about unit roots, based on KPSS tests (Kwiatkowski *et al.* [1992]) for level stationarity or trend stationarity. All the log-transformed series for CASH, M1, M2, and real GNP reject the null hypothesis of level stationarity or trend stationarity at the 1 percent level, confirming that they are I(1) or higher. All the first differenced series do not reject the null of level stationarity at the 10 percent level. The first difference of real GNP does not reject the null of trend stationarity at the 10 percent level, and that of CASH and M1 does not at the 1 percent level, while that of M2 rejects the null of trend stationarity at the 1 percent level. Thus, the possibility that M2 is I(2) remains.
19. The results for unit root tests are robust against the effects of recent data under zero nominal short-term interest rates. To check this point, we conduct the same unit root tests by excluding data after 1995, and gain results that are qualitatively the same as the results for full sample estimations.
20. Perron (1997), which is employed in this paper, extends the testing procedure in Perron (1989) by endogenously selecting structural break points from the data, and computing asymptotic distributions for finite samples. Christiano (1992) points out the bias in test statistics when a structural break point is given exogenously. In response, Zivot and Andrews (1992) and Banerjee, Lumsdaine, and Stock (1992) propose unit root testing procedures of endogenously selecting a structural break point from the data. In the meantime, Soejima (1995) examines the unit root testing procedures with a structural break using Japanese data.
21. A level shift and a trend break imply that a structural beak occurs in the constant term and the slope of a linear trend, respectively.
22. Of course, the structural break point is not necessarily restricted only once during the period around World War II. As examined extensively in Soejima (1995), it is highly likely that a structural change occurs at the first oil crisis in the money stock and real output. We leave the issues of endogenously determining the number and timing of structural changes for future research, given that this line of research on unit root testing procedures is currently advancing. See, for example, Lee and Strazicich (2003) for unit root testing procedures with two structural changes.

Figure 1 Money Stock and Output





Note: Data are compiled as follows.

- For monetary aggregate figures, the present money stock statistics are used from 1955 forward. Prior to 1955, we utilized the estimates in Fujino (1994), which are consistent with the present time series. For the period 1941–51, which is missing from the data in Fujino (1994), we use the data in Asakura and Nishiyama (1974). Finally, for the years 1943 and 1945, which are missing from Asakura and Nishiyama (1974), we use a linear extrapolation.
- For real GNP, we use the present series from 1955 forward. The data prior to 1955 are sourced from Ohkawa, Takamatsu, and Yamamoto (1974).

Sources: Asakura and Nishiyama (1974); Fujino (1994); Ohkawa, Takamatsu, and Yamamoto (1974); Bank of Japan, "Financial and Economic Statistics Monthly"; Cabinet Office, "National Accounts."

Table 2 Unit Root Tests with a Structural Change

[1] Full Sample Period (1887–2003)

	CASH	M1	M2	Real GNP
[A] Level				
Model-1	2.587 (11) [1947]	–3.941 (2) [1935]	–4.358 (8) [1943]	–4.468 (8) [1944]
Model-2	2.593 (9) [1947]	0.534 (8) [1948]	–0.843 (8) [1947]	–1.123 (12) [1957]
[B] First difference				
Model-1	–3.178 (8) [1941]	–2.185 (2) [1948]	–1.801 (6) [1949]	–8.547*** (6) [1944]
Model-2	–5.297** (8) [1941]	–4.861 (7) [1943]	–5.342** (7) [1943]	–5.601** (11) [1957]
[C] Second difference				
Model-1	–5.720*** (9) [1946]	–5.710*** (7) [1947]	–6.377*** (9) [1947]	–5.288** (11) [1969]
Model-2	–4.802 (9) [1948]	–5.309** (5) [1946]	–5.832** (5) [1946]	–10.588*** (7) [1944]

[2] Subsample Period (1887–1962)

	CASH	M1	M2	Real GNP
[A] Level				
Model-1	–2.347 (9) [1941]	–2.292 (2) [1935]	–1.807 (2) [1936]	–0.543 (1) [1932]
Model-2	–5.877*** (11) [1936]	–5.181* (11) [1940]	–4.234 (11) [1940]	–2.641 (1) [1936]
[B] First difference				
Model-1	–3.826 (11) [1941]	–4.265 (11) [1943]	–4.080 (1) [1935]	–3.477 (6) [1932]
Model-2	–5.277** (11) [1942]	–4.033 (7) [1941]	–4.253 (1) [1937]	–9.416*** (11) [1943]
[C] Second difference				
Model-1	–13.451*** (0) [1917]	–4.634 (5) [1930]	–4.779* (5) [1931]	–8.779*** (11) [1944]
Model-2	–4.152 (9) [1937]	–4.331 (9) [1943]	–4.469 (9) [1943]	–4.655 (7) [1936]

Notes: 1. "Model-1" and "Model-2" test the null hypothesis $\alpha = 1$ by

$$\text{Model-1: } x_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha x_{t-1} + \sum_{i=1}^{k} \Delta x_{t-i} + e_t,$$

$$\text{Model-2: } x_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_b)_t + \alpha x_{t-1} + \sum_{i=1}^{k} \Delta x_{t-i} + e_t,$$

where t is the time trend and e is the error term. Here, the value of DU_t is zero before the structural change and one thereafter; the value of $D(T_b)$ is one at the period after the structural change and zero at all other times; and the value of DT_t is zero before the structural change, and then becomes the dummy variable t.

2. ***, **, and * indicate statistical significance at the 1, 5, and 10 percent levels, respectively, based on Perron (1997).
3. Dates inside brackets show the time points when the structural change occurs. These time points are selected to maximize the absolute value of t, the structural change parameter.
4. Figures in parentheses show the lag order. The lag order is chosen by the general-to-specific rule: the lag order is iteratively shortened until the last coefficient becomes significant within the maximum lag order of 12.

The above results for unit root tests with a structural change suggest that test results for long-term time series of the money stock in Japan are sensitive to the handling of missing observations for the World War II period. It should be noted that such a structural change, if any, influences the data generation process gradually over a decade. This makes it difficult to detect the influence of a structural change when supplementary data are used to fill in the missing observations. In contrast, if we exclude the missing observations and simply concatenate observations for the prewar and postwar periods, then a large gap emerges in the time-series movement of the data and the influence from the structural change seems to be overly identified.

B. Cointegration Tests

As explained in Section II, long-run monetary neutrality cannot be tested when the money stock and real output are cointegrated. We next test for cointegration between real GNP and CASH, M1, and M2, respectively.

Table 3 presents the results of the Engle-Granger test for cointegration between real GNP and CASH, M1, and M2. This table indicates that the null hypothesis that there is no cointegration between the money stock and real GNP cannot be rejected even at the 10 percent level for all three measures of the money stock, CASH, M1, and M2.

We further conduct the cointegration test of Gregory and Hansen (1996) that gives rigorous consideration to structural changes. This residual basis test considers three types of structural shift: (1) level shift; (2) level shift with a trend; and (3) regime change (level shift combined with change of the cointegration vector).

Table 4 summarizes the results of cointegration tests with a structural change for the full sample period. On the whole, these test results reverse the results of the Engle-Granger test without considering a structural change. These tests show cointegration between real GNP and both CASH and M1 for all three types of structural changes. In the case of level shift and trend shift, however, the probability of cointegration with M1 is somewhat lower. As for M2, the results indicate that cointegration exists under the regime change case, but does not exist under the two other cases.

Table 3 Cointegration Tests

	CASH	M1	M2
Full sample (1887–2003)	–2.047 (1)	–1.904 (1)	–1.814 (1)

Notes: 1. We conduct the ADF test on the regression results for real GNP on the constant term and the money variables (the Engle-Granger test). The table does not include any results that are significant at the 10 percent level or less.
2. Figures in parentheses show the lag order. The lag order is chosen by the general-to-specific rule: the lag order is iteratively shortened until the last coefficient becomes significant within the maximum lag order of 12.

Table 4 Cointegration Tests with a Structural Change (Sample Period: 1887–2003)

	CASH	M1	M2
[A] Level shift (*C*)			
ADF^*	–6.700*** [1943]	–5.288*** [1943]	–3.530 [1929]
Z_t^*	–6.713*** [1943]	–5.195*** [1943]	–4.024 [1940]
Z_α^*	–65.509*** [1943]	–43.063** [1943]	–27.440 [1939]
[B] Level shift and trend break (*C/T*)			
ADF^*	–6.224*** [1943]	–5.162*** [1943]	–3.855 [1962]
Z_t^*	–6.182*** [1943]	–5.044** [1943]	–4.159 [1942]
Z_α^*	–57.394*** [1943]	–40.790** [1943]	–29.068 [1942]
[C] Regime change (*C/S*)			
ADF^*	–6.698*** [1943]	–6.323*** [1944]	–7.014*** [1944]
Z_t^*	–6.712*** [1943]	–6.325*** [1944]	–7.044*** [1944]
Z_α^*	–65.408*** [1943]	–61.027*** [1944]	–70.176*** [1944]

Notes: 1. Using the methods presented in Gregory and Hansen (1996), we conduct residual-basis cointegration tests on level shift (*C*), level shift with trend break (*C/T*), and regime change (*C/S*), respectively, under the following specifications.

(*C*) $x_{1t} = \mu_1 + \mu_2 D_t + \alpha x_{2t} + e_t.$

(*C/T*) $x_{1t} = \mu_1 + \mu_2 D_t + \beta t + \alpha x_{2t} + e_t.$

(*C/S*) $x_{1t} = \mu_1 + \mu_2 D_t + \alpha_1 x_{2t} + \alpha_1 x_{2t} D_t + e_t.$

Here, D_t has a value of zero before the structural change, and then becomes the dummy variable *t*.

2. Dates inside brackets show the time points when the structural change occurs.
3. *** and ** indicate statistical significance at the 1 and 5 percent levels, respectively.

C. Summary of Time-Series Properties of the Data

To sum up the results of the unit root and cointegration tests, it is difficult to draw definite conclusions on the time-series properties of the data. The test results differ depending on the testing procedures, especially whether or not consideration is given to a structural change. Although not ideal, it is deemed highly probable that the combination of M2 and real GNP satisfies the preconditions for testing long-run monetary neutrality: they are both I(1) and are not cointegrated. It should be noted, however, that the possibility that M2 is I(2) cannot be denied, so we also test long-run monetary superneutrality for M2.

In the analysis below, we mainly examine long-run monetary neutrality in M2, and employ CASH and M1 in a supplementary manner to check the robustness of the benchmark estimation results for M2.

V. Estimation Results

We now follow the method used in King and Watson (1997), which is explained in Section II, to test for long-run monetary neutrality based on the long-term Japanese time-series data from 1885 to 2003. Specifically, we make four sets of calculations using a bivariate (the money stock and real output) structural VAR model under the four different identifying restrictions: (1) the short-term elasticity of the money stock to real output (λ_{my}) is known; (2) the elasticity of real output to the money stock (λ_{ym}) is known; (3) the long-term elasticity of the money stock to real output (γ_{my}) is known; and (4) the long-term elasticity of real output to the money stock (γ_{ym}) is known.

As noted above, our subsequent empirical investigation focuses on M2, which most satisfies the preconditions of the time-series properties for testing long-run monetary neutrality. Considering the restrictions of the lag order, we set the beginning of the sample period as 1890 and use data prior to the initial period as lags in regressions.

A. Benchmark Estimation Results for M2

Figure 2 presents the benchmark estimation results for M2 using the sample period from 1890 to 2003 with a lag order of two. In this estimation, we add a dummy variable that takes a value of one after 1941.

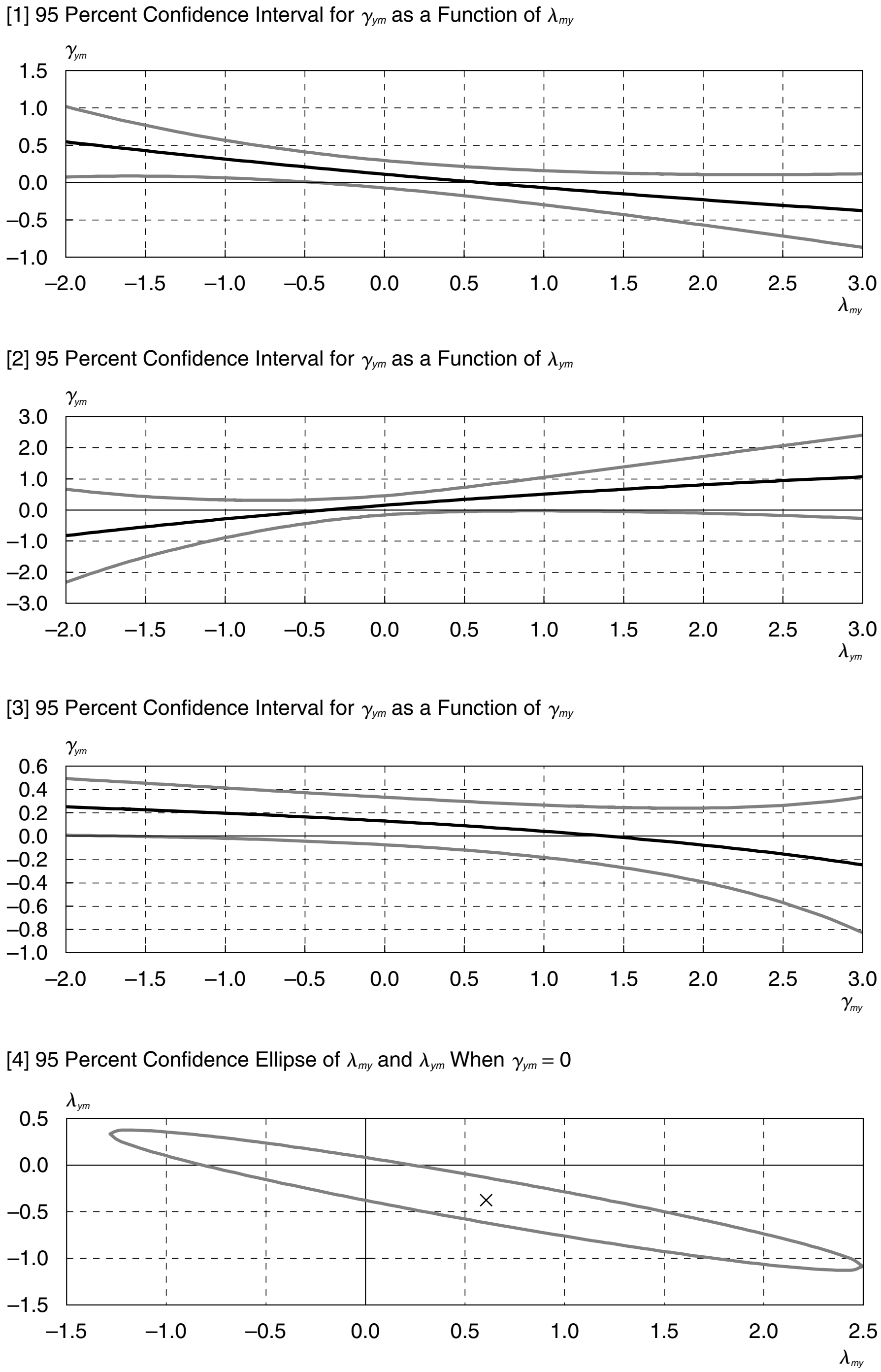
First, panels [1] through [3] show the estimates of γ_{ym} (the black lines) and their 95 percent confidence interval (the gray lines) for the elasticities of λ_{my}, λ_{ym}, and γ_{my} as the identifying restrictions ranging from –2 to 3, respectively. In these panels, if $\gamma_{ym} = 0$ (the horizontal line at the value of zero) is included in the 95 percent confidence interval, then long-run monetary neutrality cannot be rejected at the 5 percent significance level. For example, in panel [1] when $\lambda_{my} = 0$, the estimate of γ_{ym} is 0.112 and the 95 percent confidence interval is $-0.072 \leq \gamma_{ym} \leq 0.296$, so long-run monetary neutrality is not rejected.

Panels [1]–[3] as a whole show that long-run monetary neutrality holds for a wide range of identifying restrictions. The point estimates of γ_{ym} are respectively a decreasing function in λ_{my}, an increasing function in λ_{ym}, and a decreasing function in γ_{my}. The point estimates of γ_{ym} take a value of zero around the identifying restrictions of $\lambda_{my} = 0.6$, $\lambda_{ym} = -0.4$, and $\gamma_{my} = 1.45$, respectively. As for the results of testing for long-term monetary neutrality, $\gamma_{ym} = 0$ stays within the 95 percent confidence interval for the identifying restrictions of $\lambda_{my} \geq -0.4$, $-2 \leq \lambda_{ym} \leq 3$, and $\gamma_{my} > -1.6$, respectively, and the long-run monetary neutrality hypothesis is not rejected.[23]

Closer examinations of panels [1]–[3] give the observations below. Looking first at panel [1], as noted by King and Watson (1997), the value of λ_{my} depends on the supply process of the money stock. When a central bank provides reserves to smooth interest rates, m is adjusted to accommodate money demand shocks resulting from changes in y. Therefore, λ_{my} corresponds to the short-term elasticity of money demand, and λ_{my} is expected to take a positive value.

23. The estimation results using identification restrictions outside the values shown in panels [1]–[3] in Figure 2 generally support long-run monetary neutrality. First, using λ_{my} as the identifying restriction (panel [1]), $\gamma_{ym} = 0$ comes back to the confidence interval when $\lambda_{my} \leq -2.65$, and, thus, long-run monetary neutrality turns out not to be rejected. The power of the test, however, becomes weak when $\lambda_{my} \leq -3.2$ or $\lambda_{my} \geq 5.5$, since the confidence interval becomes wider than two. Second, using λ_{ym} as the identifying restriction (panel [2]), $\gamma_{ym} = 0$ always stays within the confidence interval regardless of the value of λ_{ym}, thus long-run monetary neutrality is not rejected under any values of the identifying restriction. Nevertheless, the power of the test becomes weak when $\lambda_{ym} \leq -1.55$ or $\lambda_{ym} \geq 2.25$, since the confidence interval becomes broader than two. Third, using γ_{my} as the identifying restriction (panel [3]), $\gamma_{ym} = 0$ returns within the confidence interval when $\gamma_{my} \leq -6.85$, and thus long-run monetary neutrality turns out not to be rejected. However, the confidential interval becomes broader than two when $\gamma_{my} \geq 3.75$, and thus the testing power becomes weak.

Figure 2 Long-Run Neutrality of M2 (Lag = 2; Sample Period: 1890–2003)



Note: Black lines in panels [1]–[3] and "×" in panel [4] are estimated values. Gray lines indicate the 95 percent confidence intervals.

Panel [2] shows that long-run monetary neutrality holds even when short-run monetary neutrality does not, implying that short-run non-neutrality is deemed consistent with long-run neutrality. Standard macroeconomic models assume that an expansion in monetary aggregates, at least over the short term, does not reduce real output. This corresponds to the case of $\lambda_{ym} \geq 0$, and even under this assumption long-run monetary neutrality is not rejected.

In panel [3], γ_{my} is a parameter that indicates the long-term reaction of m to the exogenous and permanent shock to y. If the velocity of money is constant, then price levels also remain constant when $\gamma_{my} = 1$. Given the downward trend in the velocity of M2 and the upward trend in general price levels, the value of γ_{my} is likely to be greater than one. Even under such assumptions for γ_{my}, long-run monetary neutrality is still not rejected.

In reverse to panels [1]–[3], panel [4] in Figure 2 employs long-run monetary neutrality ($\gamma_{ym} = 0$) as an identifying restriction and plots the 95 percent confidence ellipse for the short-term elasticities of λ_{my} and λ_{ym}. When the identifying restriction of $\gamma_{ym} = 0$ is used, the point estimates of (λ_{my}, λ_{ym}) are (0.607, –0.378) and the 95 percent confidence ellipse includes the origin and slopes downward to the right. Here, it should be noted that the upper and lower bounds of the value of λ_{ym} are somewhat broader than those of λ_{my}.

B. Robustness of the Estimation Results for M2

Next, we check the robustness of the estimation results shown above by changing the lag order and estimating without the postwar dummy variable. The results are summarized in Table 5. The upper panel presents the estimation results for γ_{ym} using the three identifying restrictions $\lambda_{my} = 0$, $\lambda_{ym} = 0$, and $\gamma_{my} = 1$, respectively. Conversely, the lower panel presents the estimation results for λ_{my}, λ_{ym}, and γ_{my} using the identifying restriction $\gamma_{ym} = 0$. The table also shows the estimation results for lag orders of three and four in addition to the lag order of two in the benchmark estimation.[24] Moreover, the table provides the estimates with and without a constant dummy that takes a value of one from 1941 on.

We see from Table 5 that the results generally support long-run monetary neutrality, except for certain cases. We see a tendency of expanding the standard error as the lag order increases. We confirm that the inclusion or non-inclusion of a postwar dummy variable does not influence the test results on long-run monetary neutrality.

We next examine the robustness of the period for setting the postwar dummy variable in detail. We estimate γ_{ym} by sequentially shifting the beginning period of the postwar dummy variable from 1941 to 1951, using identifying restrictions $\lambda_{my} = 0$, $\lambda_{ym} = 0$, and $\gamma_{my} = 1$, respectively (estimations corresponding to Table 5 [1]). The

24. The lag order could also be selected based on the information criteria. However, the lag order selection may have a major impact on the estimation results since we focus on the long term, in the sense that the period is long enough to materialize all the effects. As pointed out by Faust and Leeper (1997), the lag order itself should also be considered as a restriction when using long-term identifying restrictions. For these reasons, we employ the more conservative approach of examining the robustness by computing for each lag order.

Table 5 Long-Run Neutrality of M2: Robustness to Lag Length and Sample Period

[1] Estimates of γ_{ym} with Various Identifying Restrictions ($\lambda_{my} = 0$, $\lambda_{ym} = 0$, and $\gamma_{my} = 1$)

Sample period	Dummy variable	Lag order	Identifying restriction		
			$\lambda_{my} = 0$	$\lambda_{ym} = 0$	$\gamma_{my} = 1$
1890–2003	—	2	0.122 (0.089)	0.161 (0.155)	0.060 (0.107)
1890–2003	—	3	0.118 (0.090)	0.167 (0.171)	–0.059 (0.090)
1890–2003	—	4	0.148 (0.109)	0.207 (0.240)	–0.067 (0.117)
1890–2003	PW	2	0.112 (0.093)	0.154 (0.156)	0.043 (0.113)
1890–2003	PW	3	0.108 (0.095)	0.159 (0.173)	–0.076 (0.094)
1890–2003	PW	4	0.141 (0.117)	0.203 (0.247)	–0.087 (0.124)

[2] Estimates of λ_{my}, λ_{ym}, and γ_{my} with the Identifying Restriction of Long-Run Neutrality ($\gamma_{ym} = 0$)

Sample period	Dummy variable	Lag order	Identifying restriction: $\gamma_{ym} = 0$		
			λ_{my}	λ_{ym}	γ_{my}
1890–2003	—	2	0.689 (0.192)	–0.411 (0.228)	1.576 (1.096)
1890–2003	—	3	0.664 (0.173)	–0.545 (0.336)	–0.145 (1.627)
1890–2003	—	4	0.733 (0.186)	–0.587 (0.350)	0.105 (1.467)
1890–2003	PW	2	0.607 (0.183)	–0.378 (0.226)	1.401 (1.101)
1890–2003	PW	3	0.564 (0.161)	–0.488 (0.326)	–0.402 (1.607)
1890–2003	PW	4	0.644 (0.174)	–0.537 (0.342)	–0.101 (1.437)

Notes: 1. Figures in parentheses are the standard errors.
2. Specifications with a dummy variable (PW) are estimated adding a constant dummy variable that takes the value of one from 1941 on.

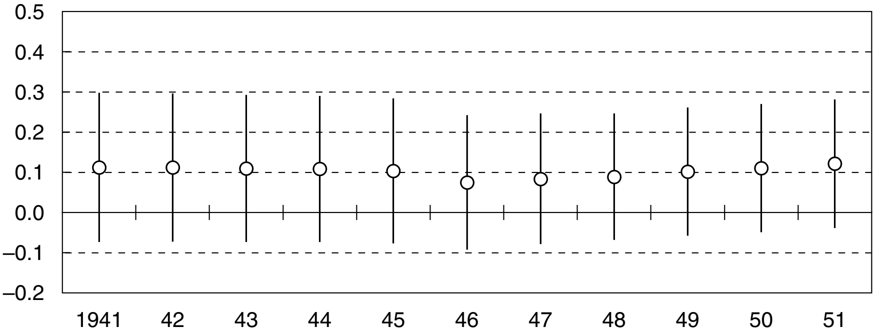
results are presented in Figure 3, which shows the point estimates as a circle and their 95 percent confidence intervals as vertical lines. We see from the figure that long-run monetary neutrality is not rejected in most cases under the identifying restrictions $\lambda_{my} = 0$, $\lambda_{ym} = 0$, or $\gamma_{my} = 1$.

To examine the effects of including data under the recent zero interest rates, we estimate γ_{ym} by sequentially extending the end of the sample period from 1994 to 2003.[25] Figure 4 shows the point estimates of γ_{ym} as well as their confidence intervals, as shown in Figure 3. We see from the figure that the point estimates of γ_{ym} are not so sensitive to the inclusion of data for the period of zero interest rates, and long-run monetary neutrality generally holds. Nevertheless, we should be a bit cautious in concluding that the data for the period of zero interest rates are unlikely to influence the estimation results, when using long-term historical data. To address this point, we examine the effects of the data for the period of zero interest rates by using quarterly data in Section VI.

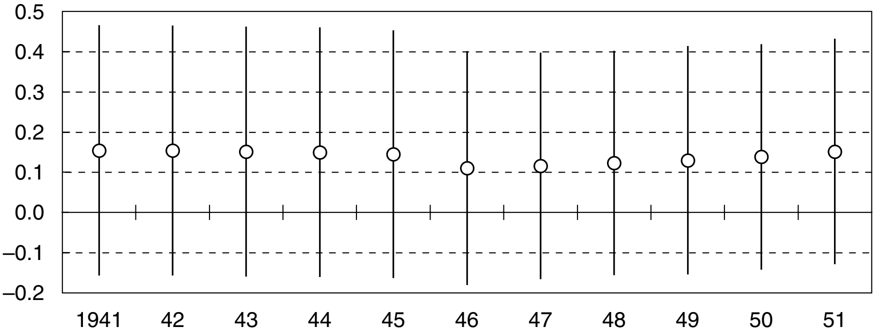
25. It is often pointed out that the money demand function becomes unstable under zero nominal interest rates. For the details on this point, see Bank of Japan (2003), Nakashima and Saito (2002), Fujiki (2002), and Fujiki and Watanabe (2004).

Figure 3 Long-Run Neutrality of M2: Robustness to Period for Setting Postwar Dummy Variable

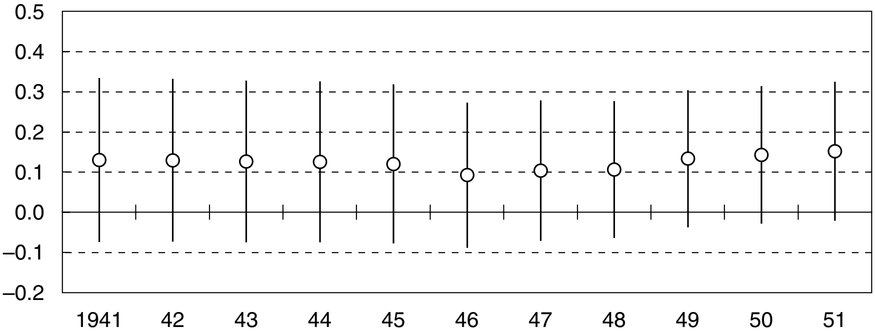
[1] 95 Percent Confidence Interval for γ_{ym} When $\lambda_{my} = 0$



[2] 95 Percent Confidence Interval for γ_{ym} When $\lambda_{ym} = 0$

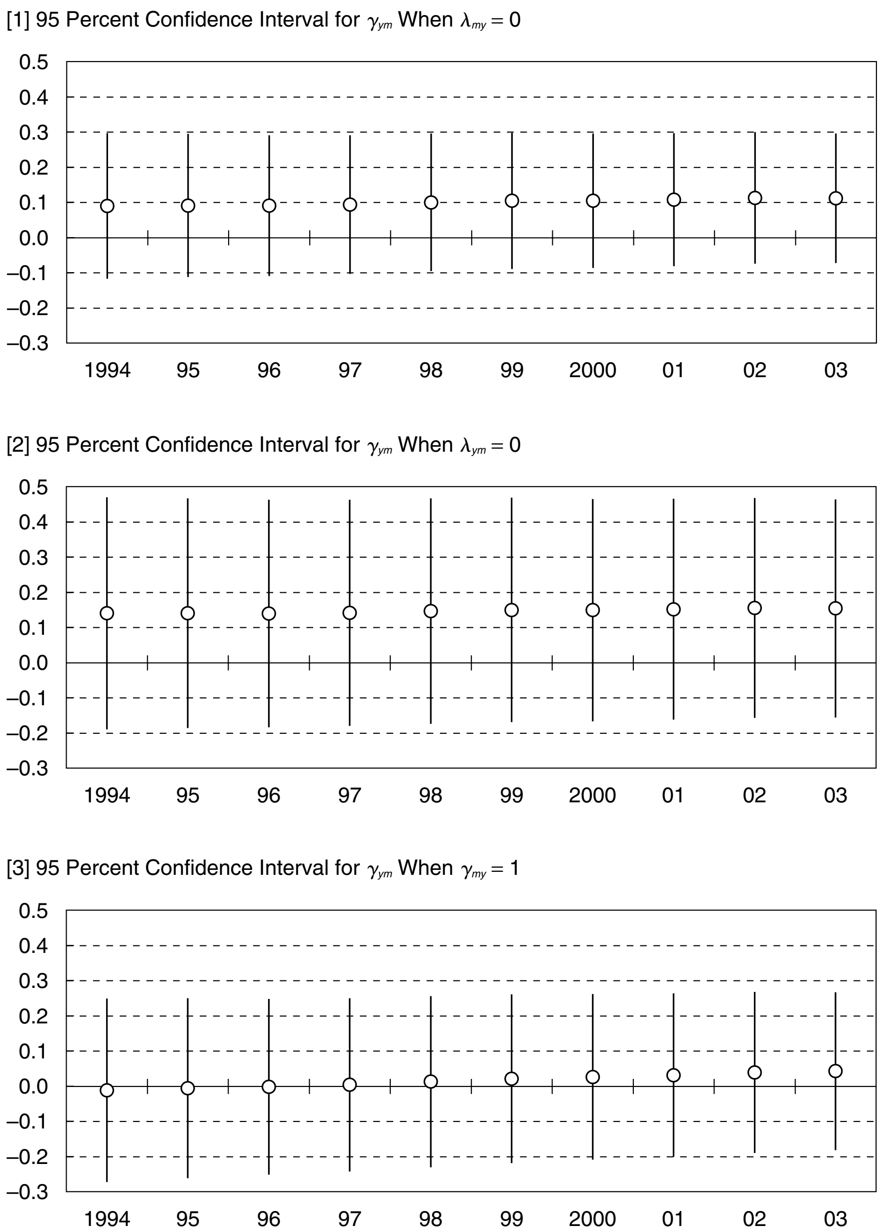


[3] 95 Percent Confidence Interval for γ_{ym} When $\gamma_{my} = 1$



Notes: 1. Horizontal axis indicates start of estimation sample period; all samples run through 2003.
2. The "○" in the figures indicates the estimated value of γ_{ym}, and the vertical line shows the 95 percent confidence interval.

Figure 4 Long-Run Neutrality of M2: Robustness to Data of the Zero Nominal Interest Rate Period



Notes: 1. Horizontal axis corresponds to the end of the sample period.
2. The "○" in the figures indicates the estimated value of γ_{ym}, and the vertical line shows the 95 percent confidence interval.

C. Estimation Results for CASH and M1

Next, let us briefly review the estimation results for CASH and M1. Like Figure 2, Figure 5 presents the estimates of γ_{ym} (the black line) and their 95 percent confidence intervals (the gray lines) for the elasticities of λ_{my}, λ_{ym}, and γ_{my} as the identifying restrictions ranging from –1 to 2, respectively, as well as the estimated values of λ_{my} and λ_{ym} and their 95 percent confidence ellipses when $\gamma_{ym} = 0$ is adopted as the identifying restriction.

The results for CASH give rather limited support to long-run monetary neutrality, although these are somewhat unstable, with low reliability, compared with the estimation results for M2. In particular, the estimates of γ_{ym} are significantly negative when $\lambda_{my} > 0.30$ and when $\gamma_{my} > -0.80$. Moreover, the estimates of γ_{ym} rise suddenly when λ_{ym} approaches 1.0 and the estimation accuracy declines.

The M1 results include no cases where long-run monetary neutrality is significantly rejected within the range of the identifying restrictions presented in the figure, and generally support long-run monetary neutrality. Moreover, the estimation reliability is generally higher than that for CASH, although it does decline somewhat under the identifying restriction λ_{ym}.

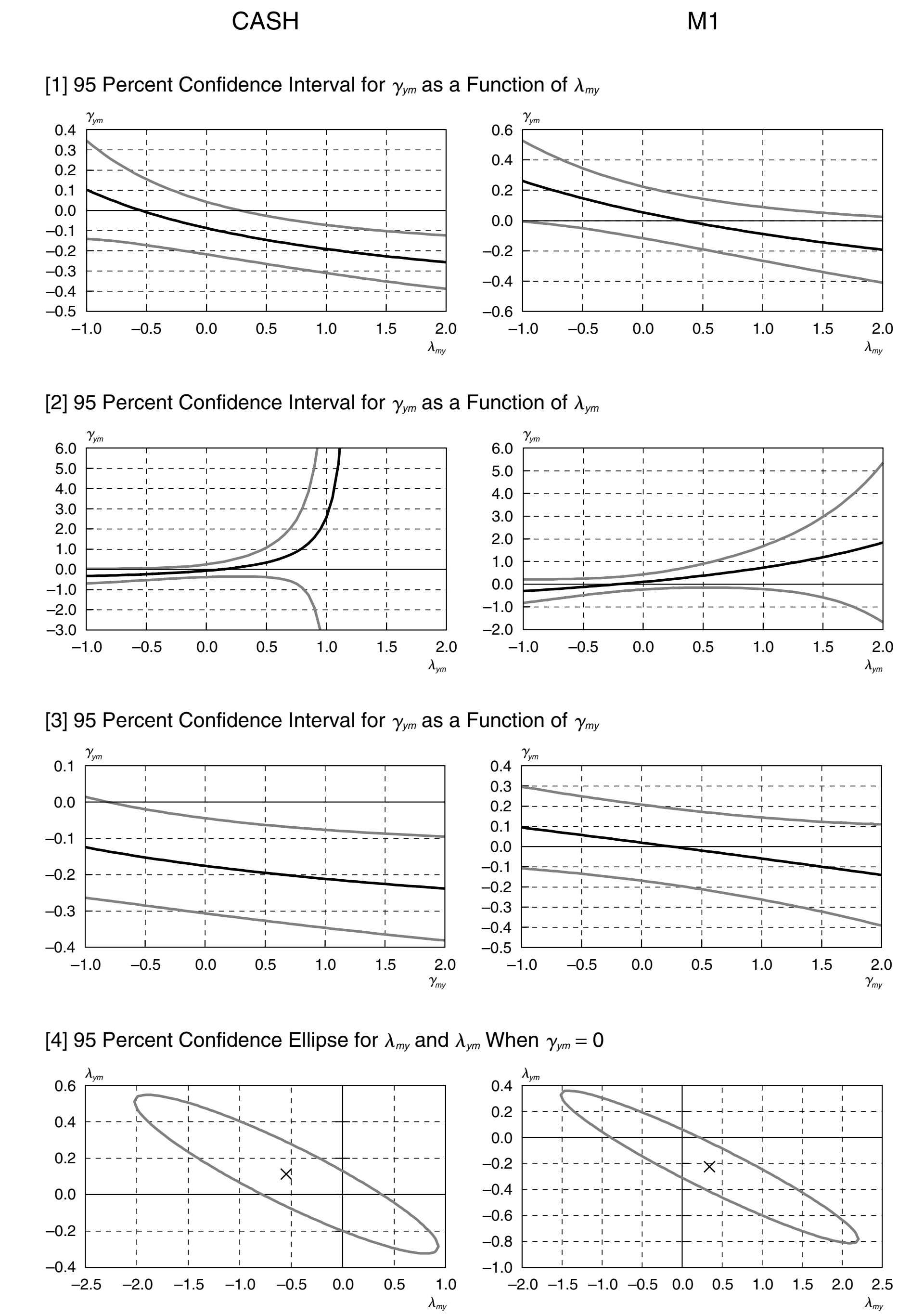
D. Estimation Results regarding the Superneutrality of M2

As a final estimation using long-term annual data, Figure 6 and Table 6 present the estimation results regarding superneutrality for M2. In this figure and table, the restricting parameters are changed to $\lambda_{\Delta my}$, $\lambda_{y\Delta m}$, $\gamma_{\Delta my}$, and $\gamma_{y\Delta m}$ in order to examine the effects of changes in the growth of the money stock, but the setup is otherwise the same as under Figure 2 and Table 5.

Figure 6 shows that superneutrality is supported under a very limited range of parameters. Specifically, when the identifying restriction is $\lambda_{\Delta my}$ or $\lambda_{y\Delta m}$, superneutrality is not rejected only within the very narrow ranges $-0.65 < \lambda_{\Delta my} < 0.10$ and $-0.10 < \lambda_{y\Delta m} < 0.10$. When the identifying restriction is $\gamma_{\Delta my}$, superneutrality is rejected just in the range $0.15 < \gamma_{\Delta my} < 2.20$, but the standard error is generally large and the estimation accuracy is low.

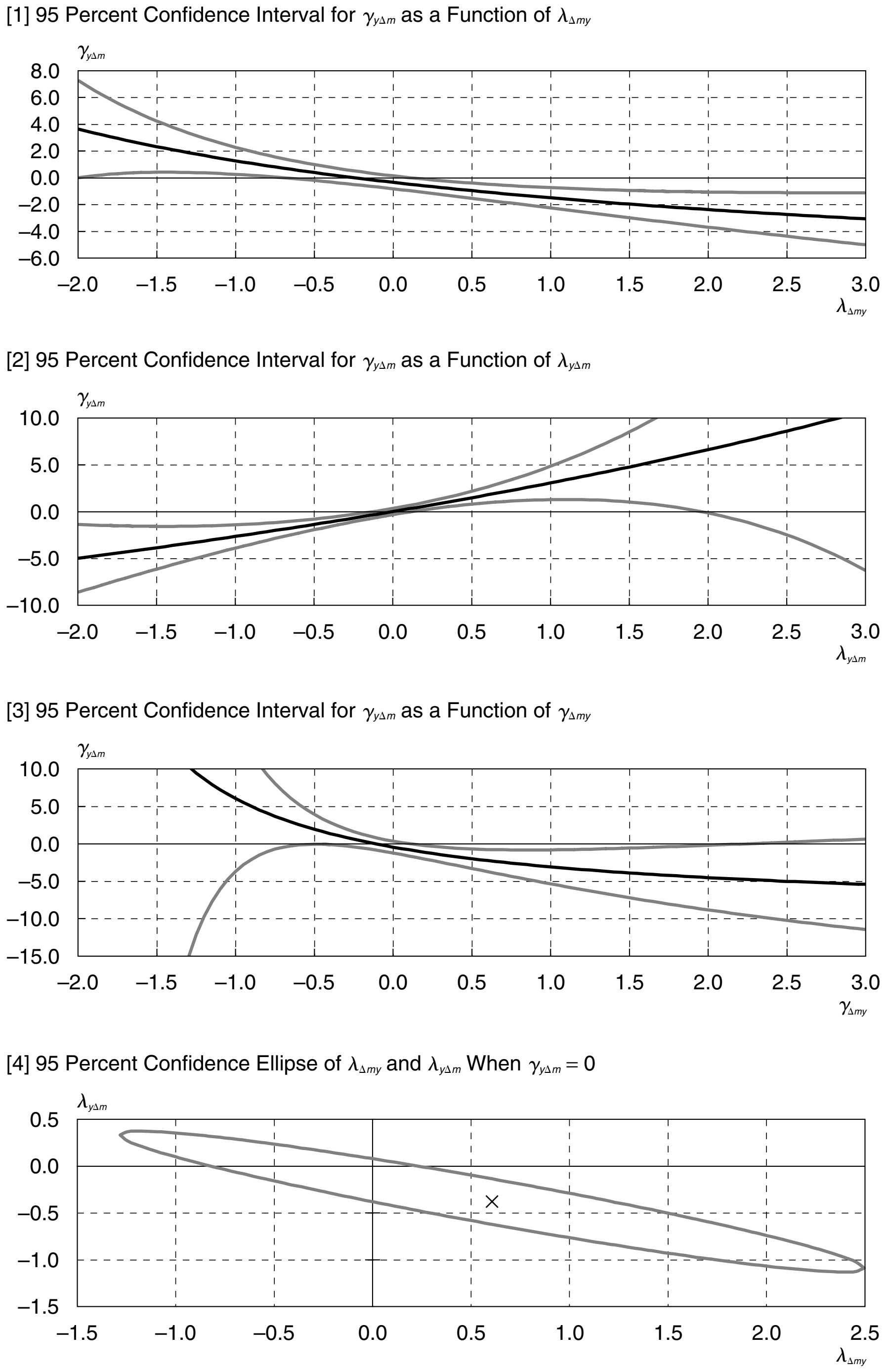
The results of the robustness check on the lag order and the prewar dummy presented in Table 6 show that when the identifying restriction is $\lambda_{\Delta my}$ or $\lambda_{y\Delta m}$, or conversely when $\lambda_{\Delta my}$ and $\lambda_{y\Delta m}$ are estimated using superneutrality ($\gamma_{y\Delta m} = 0$) as the identifying restriction, there are many cases that support superneutrality. However, when $\gamma_{y\Delta m}$ is estimated using $\gamma_{\Delta my}$ as the identifying restriction, or $\gamma_{\Delta my}$ is estimated using superneutrality as the identifying restriction, the estimation accuracy is low and superneutrality is rejected in a large number of cases. Also, similar to the neutrality test results, the superneutrality estimation accuracy tends to decline as the lag order increases.

Figure 5 Long-Run Neutrality of CASH and M1 (Lag = 2; Sample Period: 1890–2003)



Note: Black lines in panels [1]–[3] and "×" in panel [4] are estimated values. Gray lines indicate the 95 percent confidence intervals.

Figure 6 Long-Run Superneutrality of M2 (Lag = 2; Sample Period: 1890–2003)



Note: Black lines in panels [1]–[3] and "×" in panel [4] are estimated values. Gray lines indicate the 95 percent confidence intervals.

Table 6 Long-Run Superneutrality of M2: Robustness to Lag Length and Sample Period

[1] Estimates of $\gamma_{y\Delta m}$ with Various Identifying Restrictions ($\lambda_{\Delta my} = 0$, $\lambda_{y\Delta m} = 0$, and $\gamma_{\Delta my} = 1$)

Sample period	Dummy variable	Lag order	Identifying restriction		
			$\lambda_{\Delta my} = 0$	$\lambda_{y\Delta m} = 0$	$\gamma_{\Delta my} = 1$
1890–2003	—	2	–0.344 (0.246)	0.028 (0.170)	–3.061 (1.130)
1890–2003	—	3	–0.321 (0.297)	0.076 (0.242)	–2.017 (0.496)
1890–2003	—	4	–0.267 (0.340)	0.170 (0.430)	–2.322 (0.709)
1890–2003	PW	2	–0.329 (0.245)	0.033 (0.169)	–3.059 (1.146)
1890–2003	PW	3	–0.256 (0.296)	0.085 (0.239)	–2.015 (0.503)
1890–2003	PW	4	–0.256 (0.337)	0.170 (0.422)	–2.313 (0.720)

[2] Estimates of $\lambda_{\Delta my}$, $\lambda_{y\Delta m}$, and $\gamma_{\Delta my}$ with the Identifying Restriction of Long-Run Superneutrality ($\gamma_{y\Delta m} = 0$)

Sample period	Dummy variable	Lag order	Identifying restriction: $\gamma_{y\Delta m} = 0$		
			$\lambda_{\Delta my}$	$\lambda_{y\Delta m}$	$\gamma_{\Delta my}$
1890–2003	—	2	0.689 (0.192)	–0.411 (0.228)	1.576 (1.096)
1890–2003	—	3	0.664 (0.173)	–0.545 (0.336)	–0.145 (1.627)
1890–2003	—	4	0.733 (0.186)	–0.587 (0.350)	0.105 (1.467)
1890–2003	PW	2	0.607 (0.183)	–0.378 (0.226)	1.401 (1.101)
1890–2003	PW	3	0.564 (0.161)	–0.488 (0.326)	–0.402 (1.607)
1890–2003	PW	4	0.644 (0.174)	–0.537 (0.342)	–0.101 (1.437)

Notes: 1. Figures in parentheses are the standard errors.
2. Specifications with a dummy variable (PW) are estimated adding a constant dummy variable that takes the value of one from 1941 on.

VI. Testing Using Quarterly Postwar Data

In this section, we test long-run monetary neutrality by using quarterly data over the postwar period. By doing so, we confirm the robustness of our results so far using long-term annual century-long data.

The data used in this section are seasonally adjusted quarterly time-series data starting from 1955 from the data set as explained in Section III. We use the Bank of Japan monetary statistics[26] for cash, M1, and M2, and the SNA statistics for real GDP. The sample period runs over approximately 49 years from 1955/II through 2003/IV, and is seasonally adjusted using X-12-ARIMA.[27]

26. Regarding the money stock data, we mainly use the average outstanding data, which are available for most of the sample period. However, we use the end-quarter data for periods when the average outstanding data are not available: 1955–62 for CASH and M1, and 1955–66 for M2. Moreover, from April 1998 the coverage of the money supply statistics is widened to include foreign banks operating in Japan (new basis), so we adjust for the differential to make the data continuous. Additionally, from 1979 the M2 data include CDs.

27. In compiling the seasonally adjusted series for the money stock, we take the following two steps. First, we conduct the seasonal adjustments using X-12-ARIMA on three series: the end-quarter data for the old basis, the average outstanding data for the old basis, and the average outstanding for the new basis. Second, we concatenate the three series after adjusting for the differentials between them. This is because it is likely that three series have different properties of seasonal fluctuations, and thus it is deemed necessary to concatenate the three series after adjusting for seasonal fluctuations. The ARIMA models used for the seasonal adjustments were (2, 1, 1)(2, 1, 2) for cash, (2, 0, 0)(2, 1, 2) for M1, (2, 1, 2)(1, 1, 0) for M2, and (2, 1, 0)(2, 1, 0) for real GDP, and adjustments were made for changes in the number of business days.

A. Preliminary Tests

We first conduct unit root tests and cointegration tests between real GDP and the money stocks to examine the time-series properties of the data. The results of the unit root tests are presented in Tables 7 and 8, and the results of the cointegration tests in Tables 9 and 10. In the same approach adopted for the empirical analysis using the annual data in Section IV, we conduct two sets of unit root and cointegration tests with or without considering a structural change using the quarterly data. The results in Tables 7 through 10 for the quarterly data correspond to those in Tables 1 through 4 for the annual data.

Summarizing the test results, whether or not a structural change is considered, the results indicate that M1 and M2 are I(1), are not cointegrated with real GDP,

Table 7 Unit Root Test (Postwar Quarterly Data)

[1] Full Sample (1955/III–2003/IV)

	CASH		M1		M2		Real GDP	
[A] Level								
ADF	–0.712	(4)	–2.626	(9)	0.089	(4)	–2.174	(10)
PP	–0.045		–2.151		0.321		–0.782	
WS	1.211	(4)	0.146	(1)	4.517	(1)	2.069	(12)
[B] First difference								
ADF	–4.043***	(1)	–3.603**	(4)	–3.341*	(8)	–3.118	(9)
PP	–62.268***		–94.140***		–36.255***		–252.750***	
WS	–3.790***	(1)	–3.767***	(4)	–2.949*	(8)	–2.995*	(9)
[C] Second difference								
ADF	–6.852***	(6)	–6.540***	(10)	–6.697***	(7)	–8.110***	(8)
PP	–240.905***		–177.302***		–165.253***		–250.415***	
WS	–6.892***	(6)	–6.358***	(10)	–6.806***	(7)	–7.998***	(8)

[2] Post-First Oil Crisis Sample (1976/IV–2003/IV)

	CASH		M1		M2		Real GDP	
[A] Level								
ADF	–2.732	(3)	–0.354	(1)	–0.715	(8)	–0.804	(3)
PP	–7.444		0.048		–1.016		–0.752	
WS	–0.429	(7)	–0.336	(1)	2.224	(1)	–0.258	(3)
[B] First difference								
ADF	–3.522**	(1)	–6.147***	(0)	–3.036	(6)	–3.978***	(2)
PP	–56.145***		–68.553***		–29.012**		–120.878***	
WS	–3.718**	(1)	–6.277***	(0)	–3.221**	(6)	–4.141***	(2)
[C] Second difference								
ADF	–5.042***	(6)	–5.317***	(10)	–4.148***	(7)	–4.948***	(8)
PP	–122.067***		–93.950***		–106.279***		–131.162***	
WS	–5.127***	(6)	–5.709***	(8)	–3.365***	(7)	–5.021***	(8)

Notes: 1. Constant and trend terms are included in both ADF and WS tests.
2. ***, **, and * indicate statistical significance at the 1, 5, and 10 percent levels, respectively.
3. Figures in parentheses show the lag order. The lag order is chosen by the general-to-specific rule: the lag order is iteratively shortened until the last coefficient becomes significant within the maximum lag order of 12.

Table 8 Unit Root Test with a Structural Change (Postwar Quarterly Data)

	CASH	M1	M2	Real GDP
[A] Level				
Model-1	1.020 (2) [1974/II]	1.889 (1) [1973/II]	–4.426 (1) [1989/IV]	–0.882 (2) [1972/IV]
Model-2	–5.642*** (2) [1971/IV]	–3.881 (1) [1970/II]	–3.206 (1) [1970/II]	–2.983 (2) [1967/IV]
[B] First difference				
Model-1	–5.524*** (1) [1974/II]	–9.212*** (0) [1973/I]	–4.919** (0) [1973/I]	–6.545*** (2) [1972/IV]
Model-2	–5.586*** (1) [1974/III]	–8.656*** (0) [1988/III]	–5.392** (2) [1972/IV]	–6.554*** (2) [1970/I]
[C] Second difference				
Model-1	–22.316*** (0) [1973/I]	–13.809*** (1) [1980/III]	–16.132*** (0) [1972/III]	–12.634*** (2) [1974/I]
Model-2	–22.234*** (0) [1965/IV]	–14.155*** (1) [1963/II]	–16.548*** (0) [1989/IV]	–12.706*** (2) [1974/I]

Notes: 1. "Model-1" and "Model-2" test the null hypothesis $\alpha = 1$ by

$$\text{Model-1: } x_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha x_{t-1} + \textstyle\sum_{i=1}^{k} \Delta x_{t-i} + e_t,$$

$$\text{Model-2: } x_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_b)_t + \alpha x_{t-1} + \textstyle\sum_{i=1}^{k} \Delta x_{t-i} + e_t,$$

where t is the time trend and e is the error term. Here, the value of DU_t is zero before the structural change and one thereafter; the value of $D(T_b)$ is one in the period after the structural change and zero at all other times; and the value of DT_t is zero before the structural change, and then becomes the dummy variable t.

2. *** and ** indicate statistical significance at the 1 and 5 percent levels, respectively, based on Perron (1997).
3. Dates inside brackets show the time points when the structural change occurs. These time points are selected to maximize the logarithm of the t-value of the structural change parameter.
4. Figures in parentheses show the lag order. The lag order is chosen by the general-to-specific rule: the lag order is iteratively shortened until the last coefficient becomes significant within the maximum lag order of 12.

Table 9 Cointegration Test (Postwar Quarterly Data)

	CASH	M1	M2
Full sample (1955/IV–2003/IV)	–2.661 (10)	–2.269 (10)	–2.847 (10)
Post-first oil crisis (1976/IV–2003/IV)	–0.700 (0)	–2.475 (3)	–2.028 (3)

Notes: 1. We conduct the ADF test on the regression results for real GDP on the constant term and the money variables (the Engle-Granger test). The table does not include any results that are significant at the 10 percent level or less.

2. Figures in parentheses show the lag order. The lag order is chosen by the general-to-specific rule: the lag order is iteratively shortened until the last coefficient becomes significant within the maximum lag order of 12.

Table 10 Cointegration Test with a Structural Change (Postwar Quarterly Data) (Sample Period: 1955/III–2003/IV)

	CASH		M1		M2	
[A] Level shift (*C*)						
ADF^*	–3.237	[1962/III]	–2.711	[1973/III]	–3.470	[1963/II]
Z_t^*	–2.711	[1964/III]	–3.524	[1973/I]	–2.572	[1993/III]
Z_α^*	–12.066	[1964/III]	–22.454	[1973/I]	–12.609	[1971/IV]
[B] Level shift and trend break (*C/T*)						
ADF^*	–2.692	[1989/IV]	–2.871	[1990/I]	–3.285	[1987/I]
Z_t^*	–1.334	[1993/III]	–1.665	[1993/II]	–2.313	[1986/III]
Z_α^*	–4.812	[1993/II]	–6.799	[1993/II]	–9.248	[1984/III]
[C] Regime change (*C/S*)						
ADF^*	–4.033	[1976/I]	–4.051	[1976/I]	–4.068	[1974/II]
Z_t^*	–3.569	[1973/IV]	–3.815	[1973/II]	–4.015	[1974/III]
Z_α^*	–23.896	[1973/IV]	–26.877	[1973/II]	–28.958	[1974/III]

Notes: 1. Using the methods presented in Gregory and Hansen (1996), we conduct residual-basis cointegration tests on level shift (*C*), level shift with trend break (*C/T*), and regime change (*C/S*), respectively, under the following specifications. The table does not include any results that are significant at the 10 percent level or less.

(*C*) $x_{1t} = \mu_1 + \mu_2 D_t + \alpha x_{2t} + e_t$.

(*C/T*) $x_{1t} = \mu_1 + \mu_2 D_t + \beta t + \alpha x_{2t} + e_t$.

(*C/S*) $x_{1t} = \mu_1 + \mu_2 D_t + \alpha_1 x_{2t} + \alpha_1 x_{2t} D_t + e_t$.

Here, D_t has a value of zero before the structural change, and then becomes the dummy variable *t*.

2. Dates inside brackets show the time points when the structural change occurs.

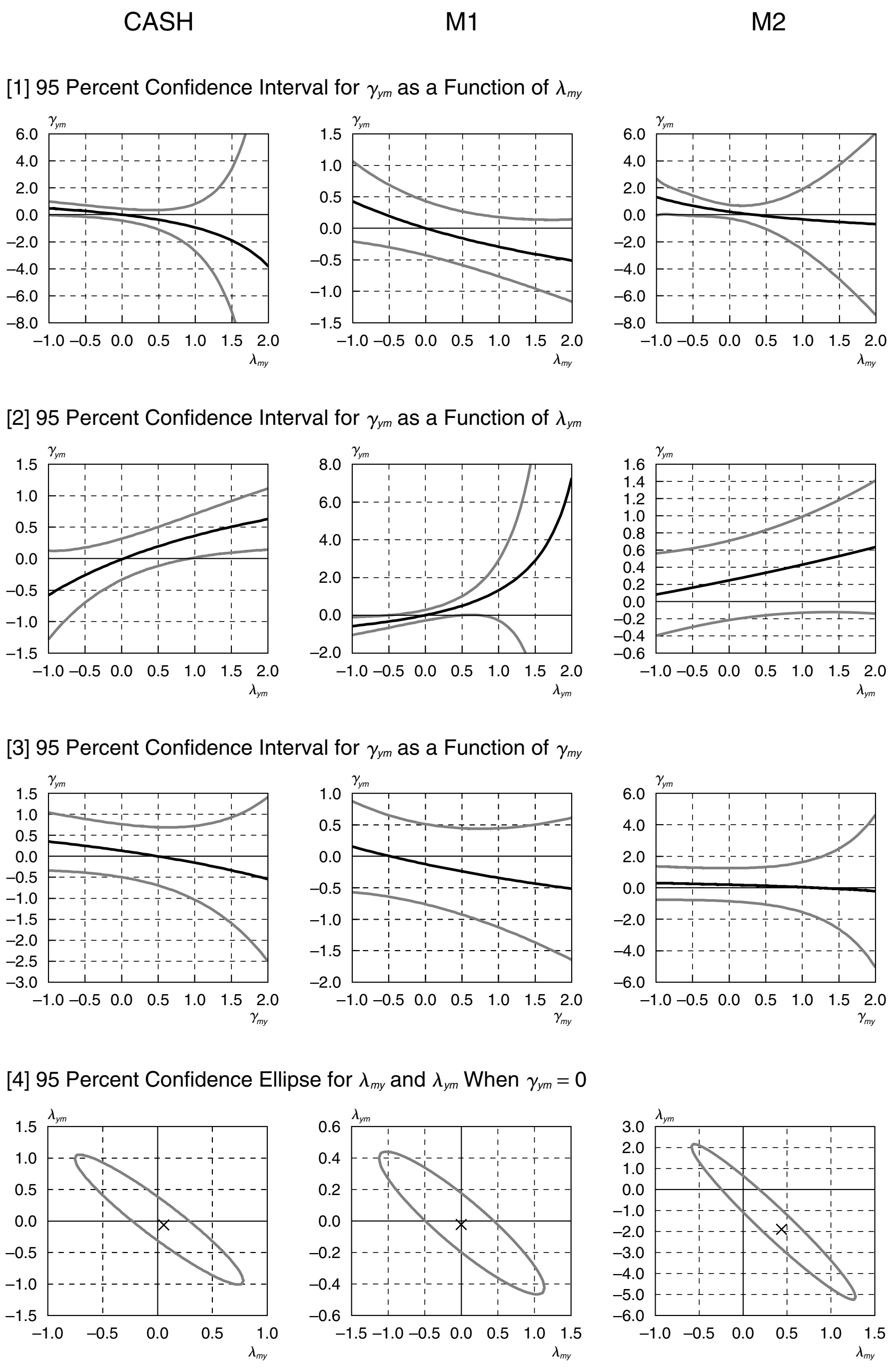
and thus meet the preconditions for testing long-run neutrality. However, when the beginning of the sample period for M1 is set to 1976/IV, M1 is cointegrated with real output. Turning to CASH, when a structural change is not considered CASH is I(1) and is not cointegrated with real GDP, but when a structural change is considered CASH is likely to be I(0).

Based on the above test results, when quarterly postwar data are used, M2 fully meets the time-series conditions to test for long-run neutrality, while M1 and CASH generally meet these testing requirements.

B. Benchmark Estimation Results

Figure 7 presents the benchmark estimation results for CASH, M1, and M2, respectively, for the sample period from 1957/III to 2003/IV and the lag order of four, and with a dummy variable (a constant dummy with a value of one) that takes the value of one after the first oil crisis (from 1973/IV on). Like Figure 2, the first three panels in Figure 7 show the estimates of γ_{ym} (the black lines) and their 95 percent confidence intervals (the gray lines) when the identifying restrictions for λ_{my}, λ_{ym}, and γ_{my} range from –1 to 2, respectively. Conversely, the fourth panel adopts long-run monetary neutrality ($\gamma_{ym} = 0$) as the identifying restriction and plots the 95 percent confidence ellipse for the two short-term elasticities λ_{my} and λ_{ym}.

Figure 7 Long-Run Neutrality Using Postwar Quarterly Data (Lag = 4; Sample Period 1957/III–2003/IV)



Note: Black lines in panels [1]–[3] and "×" in panel [4] are estimated values. Gray lines indicate the 95 percent confidence intervals.

On the whole, the estimation results for CASH, M1, and M2 all generally support long-run monetary neutrality.[28,29] Looking at the estimation results in the first three rows, the estimates of γ_{ym} are a decreasing function of λ_{my} and γ_{my} and an increasing function of λ_{ym}, and its 95 percent confidence intervals generally include $\gamma_{ym} = 0$ for the identifying restrictions ranging from –1 to 2. However, when λ_{my} is used as the identification restriction for CASH, λ_{ym} for M1, and either λ_{my} or γ_{my} for M2, the standard error increases and the estimation accuracy declines as the value of the identification restriction increases. Additionally, looking at the estimation results in the fourth row, the confidence ellipses all include the origins and all slope downward to the right. As for M2, however, the range of the confidence ellipse for λ_{ym} is somewhat broad.

It should be noted that the estimation results for the postwar quarterly data show a wider range of confidence intervals and are weak in testing power compared with those for the annual century-long data. This suggests that longer observation periods are more desirable in testing long-run monetary neutrality, since they include more information on long-run fluctuations.

C. Robustness Check for M2

Table 11 presents the results of the robustness check for the lag order and the sample period. We see from the table that the estimation results for the full sample with the first oil crisis dummy are robust to the selection of the lag order, and support long-run monetary neutrality. We, however, also see from the table that long-run monetary neutrality is rejected when the first oil crisis dummy is excluded in the full sample estimation and when the sample period is shortened after the first oil crisis.

Next, to examine the effects of the recent data observed under zero interest rates, Figure 8 plots the results for sequentially estimating γ_{ym} by extending the end of the sample period from 1994/IV using $\lambda_{my} = 0$, $\lambda_{ym} = 0$, and $\gamma_{my} = 1$ as identifying restrictions. We see from the figure that the point estimates of γ_{ym} remain almost unchanged regardless of the end of the sample period after 1994, and thus generally support long-run monetary neutrality. Nevertheless, a closer look at the figure shows that long-run monetary neutrality is just rejected when using $\lambda_{my} = 0$ as an identifying restriction when the end of the sample period is set after 2001/III.

28. We also examine the robustness of the estimation results regarding the lag order and exclusion of the post-first oil crisis dummy. When the post-first oil crisis dummy variable is excluded from the full sample estimation, long-run monetary neutrality tends to be rejected for all three indicators for the money stock: CASH, M1, and M2. However, when the beginning of the sample is set at 1976/IV, long-run monetary neutrality is generally supported, regardless of the lag order.
29. To compare the estimation results in King and Watson (1997) for the U.S. postwar period data, we conduct similar estimations by setting the end of the sample period at 1990. The estimation results are almost equivalent to Figure 7, which are close to the estimation results in the United States. It should, however, be noted that the test power is weak as evidence in the broadened significant intervals, especially for the case when γ_{my} is used as an identifying restriction.

Table 11 Long-Run Neutrality of M2 (Postwar Quarterly Data): Robustness to Lag Order and Sample Period

[1] Estimates of γ_{ym} with Various Identifying Restrictions ($\lambda_{my} = 0$, $\lambda_{ym} = 0$, and $\gamma_{my} = 1$)

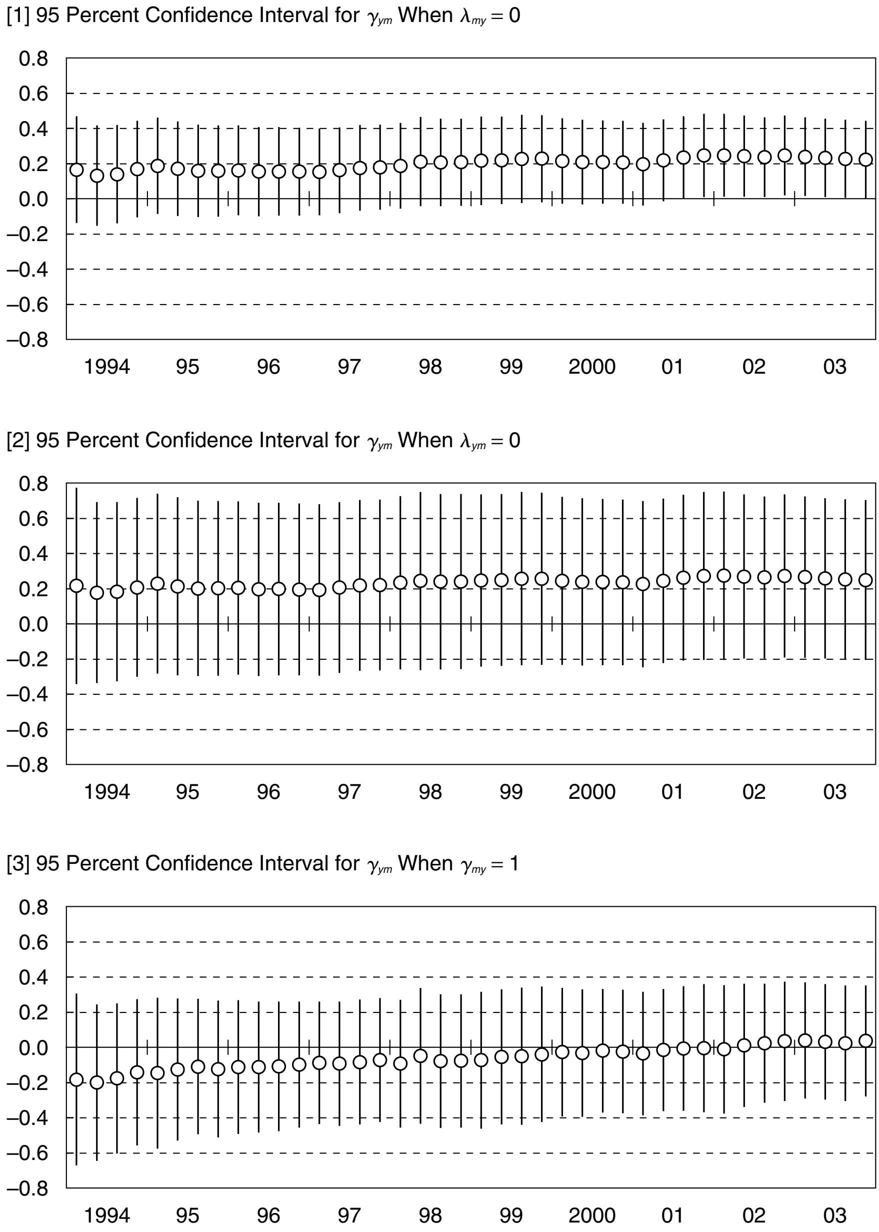
Sample period	Dummy variable	Lag order	Identifying restriction		
			$\lambda_{my} = 0$	$\lambda_{ym} = 0$	$\gamma_{my} = 1$
1957/III–2003/IV	—	4	0.429 (0.077)	0.434 (0.254)	0.427 (0.084)
1957/III–2003/IV	—	6	0.399 (0.081)	0.396 (0.270)	0.413 (0.083)
1957/III–2003/IV	—	8	0.410 (0.088)	0.407 (0.298)	0.427 (0.090)
1957/III–2003/IV	PO	4	0.224 (0.112)	0.249 (0.230)	0.037 (0.160)
1957/III–2003/IV	PO	6	0.185 (0.105)	0.186 (0.222)	0.088 (0.137)
1957/III–2003/IV	PO	8	0.192 (0.105)	0.193 (0.229)	0.092 (0.138)
1976/IV–2003/IV	—	4	0.412 (0.084)	0.409 (0.191)	0.403 (0.094)
1976/IV–2003/IV	—	6	0.404 (0.093)	0.401 (0.214)	0.395 (0.104)
1976/IV–2003/IV	—	8	0.406 (0.082)	0.401 (0.194)	0.404 (0.092)

[2] Estimates of λ_{my}, λ_{ym}, and γ_{my} with the Identifying Restriction of Long-Run Neutrality ($\gamma_{ym} = 0$)

Sample period	Dummy variable	Lag order	Identifying restriction: $\gamma_{ym} = 0$		
			λ_{my}	λ_{ym}	γ_{my}
1957/III–2003/IV	—	4	0.950 (0.179)	–3.753 (2.128)	2.011 (0.349)
1957/III–2003/IV	—	6	0.737 (0.119)	–3.230 (2.060)	2.154 (0.396)
1957/III–2003/IV	—	8	0.730 (0.119)	–3.127 (2.206)	2.117 (0.411)
1957/III–2003/IV	PO	4	0.316 (0.055)	–1.425 (0.778)	1.164 (0.735)
1957/III–2003/IV	PO	6	0.304 (0.051)	–1.296 (0.923)	1.583 (0.992)
1957/III–2003/IV	PO	8	0.345 (0.057)	–1.441 (1.054)	1.669 (1.108)
1976/IV–2003/IV	—	4	0.871 (0.195)	–3.290 (1.865)	2.031 (0.402)
1976/IV–2003/IV	—	6	0.942 (0.229)	–3.618 (2.885)	2.069 (0.488)
1976/IV–2003/IV	—	8	0.958 (0.213)	–3.669 (2.510)	2.124 (0.428)

Notes: 1. Figures in parentheses are the standard errors.
2. Specifications with a dummy variable (PO) are estimated adding a constant dummy variable that takes the value of one from 1973/IV on.

Figure 8 Long-Run Neutrality of M2 (Postwar Quarterly Data): Robustness to Data of the Zero Nominal Interest Rate Period



Notes: 1. Horizontal axis corresponds to the end of the sample period.
2. The "○" in the figures indicates the estimated value of γ_{ym}, and the vertical line shows the 95 percent confidence interval.

VII. Conclusions

In this paper, we have comprehensively investigated long-run monetary neutrality in Japan in three respects. First, we have compiled two types of data sets for century-long annual data as well as postwar quarterly data. Second, we have carefully examined the time-series properties of the data, especially their orders of integration. Third, we have fully analyzed the robustness of the empirical results against wide-ranging identifying restrictions in the bivariate structural VAR model of the money stock and real output.

Based on these investigations, we have found robust evidence supporting long-run monetary neutrality, especially in the case of M2 as a measure of the money stock. In particular, we have found very solid evidence of long-run monetary neutrality from long-term time-series data retroactively available from the Meiji Period.

Of course, it is important to note that the estimation results shown in this paper rely on two relatively strong assumptions. First, the testing framework for long-run monetary neutrality in Fisher and Seater (1993) depends crucially on the assumption of the time-series properties of the data series. In light of this point, we pay due consideration to properly account for the time-series properties of the money stock and real output by employing the longest time series currently available since the Meiji Period. We, however, should be somewhat cautious about the limited power of unit root testing procedures. The results of the order of integration of the money stock and real output are still sensitive to the handling of structural breaks, such as the World War II period for century-long annual data and the first oil crisis for the postwar quarterly data.

Second, the bivariate VAR model of King and Watson (1997) has a limitation that it is able to identify just two macroeconomic shocks, i.e., monetary and non-monetary shocks. The testing results on long-run monetary neutrality are not necessarily the same if the economy has three or more sources of macroeconomic shocks. Thus, it is conceivable to extend the analytical framework into a multivariate model that includes other variables in addition to the money stock and real output.[30] The estimation results in the paper show that long-run monetary neutrality is supported in both century-long annual data and postwar quarterly data. Among these, the estimation results show higher precision when using annual data. This suggests that longer observation periods are more desirable in testing long-run monetary neutrality, since they include more information on long-run fluctuations. Thus, for the testing based on the postwar quarterly data with a shorter period but higher-frequency observations, it is deemed worthwhile to consider the extension into a multivariate model with three or more variables.

30. However, Bullard (1999) surveys recent empirical studies on long-run monetary neutrality across countries, and points out that the analytical results using multivariate models generally support those found using bivariate models.

APPENDIX: OUTLINE OF LUCAS'S MODEL

In this appendix, we explain that even when long-run monetary neutrality actually holds, empirical analyses using a simple reduced-form model may lead to a false conclusion of monetary non-neutrality.

The economy is depicted using a Lucas-type aggregate supply curve and a monetarist aggregate demand function. The money stock follows the autoregression process presented as equation (A.1).

$$\begin{aligned} y_t &= \theta(p_t - E_{t-1}p_t), \\ p_t &= m_t - \delta y_t, \\ m_t &= \rho m_{t-1} + \epsilon_t^m. \end{aligned} \tag{A.1}$$

Here y, m, and p are the logarithms of real output, the money stock, and price level, respectively. The money stock follows a stationary process ($\rho \neq 1$), and ϵ^m is a shock to the money stock. Equation (A.1) is structured so that only unexpected changes in the money stock influence output. Thus, permanent changes in the money stock do not influence output, and so long-run monetary neutrality holds.

When equation (A.1) is solved for output, we can derive a distributional lag model for the money stock as presented in equation (A.2).

$$y_t = \frac{\theta}{1+\theta\delta}(m_t - \rho m_{t-1}). \tag{A.2}$$

Even though equation (A.1) shows long-run monetary neutrality, the reduced-form model presented as equation (A.2) suggests that a single-unit permanent increase in the money supply will result in an output increase of $\theta(1 - \rho)/(1 + \theta\delta)$ units. In other words, if a simple reduced-form model is used for testing monetary neutrality, this may lead to an erroneous conclusion of non-neutrality. When the money stock has a unit root (when $\rho = 1$), however, the results cannot be misread.

There are actually diverse shocks in the economy, so it is not possible to directly observe ϵ^m, a pure shock to the money stock. If we also give consideration to the existence of other shocks, endogenous policy responses, and other factors, then even if the money stock has a unit root, it is still inappropriate to use a simple reduced-form equation to test long-run neutrality, and we need to identify the shock ϵ^m.

References

Asakura, Kokichi, and Chiaki Nishiyama, eds. *Nihon Keizai no Kaheiteki Bunseki: 1886–1970* (Monetary Analysis of the Japanese Economy: 1886–1970), Sobunsha Publishing Co., 1974 (in Japanese).

Backus, David K., and Patrick J. Kehoe, "International Evidence on the Historical Properties of Business Cycles," *American Economic Review*, 82 (4), 1992, pp. 864–888.

Banerjee, Anindya, Robin L. Lumsdaine, and James H. Stock, "Recursive and Sequential Tests of the Unit-Root and Trend-Break Hypotheses: Theory and International Evidence," *Journal of Business & Economic Statistics*, 10 (3), 1992, pp. 271–287.

Bank of Japan, "The Role of the Money Stock in Conducting Monetary Policy," *Bank of Japan Quarterly Bulletin*, 11 (2), 2003, pp. 151–202.

Blanchard, Oliver Jean, and Danny Quah, "The Dynamic Effects of Aggregate Demand and Supply Disturbances," *American Economic Review*, 79 (4), 1989, pp. 655–673.

Bullard, James, "Testing Long-Run Monetary Neutrality Propositions: Lessons from the Recent Research," *Review*, 81 (6), Federal Reserve Bank of St. Louis, 1999, pp. 57–77.

———, and John W. Keating, "The Long-Run Relationship between Inflation and Output in Postwar Economies," *Journal of Monetary Economics*, 36 (3), 1995, pp. 477–496.

Christiano, Laurence J., "Searching for a Break in GNP," *Journal of Business & Economic Statistics*, 10 (3), 1992, pp. 237–250.

Coe, Patrick J., and James M. Nason, "The Long-Horizon Regression Approach to Monetary Neutrality: How Should the Evidence Be Interpreted?" *Economic Letters*, 78 (3), 2003, pp. 351–356.

Faust, Jon, and Eric M. Leeper, "When Do Long-Run Identifying Restrictions Give Reliable Results?" *Journal of Business and Economic Statistics*, 15 (3), 1997, pp. 345–353.

Fisher, Mark E., and John J. Seater, "Long-Run Neutrality and Superneutrality in an ARIMA Framework," *American Economic Review*, 83 (3), 1993, pp. 402–415.

Fujiki, Hiroshi, "Money Demand near Zero Interest Rate: Evidence from Regional Data," *Monetary and Economic Studies*, 20 (2), Institute for Monetary and Economic Studies, Bank of Japan, 2002, pp. 25–42.

———, and Kiyoshi Watanabe, "Japanese Demand for M1 and Demand Deposits: Cross-Sectional and Time-Series Evidence from Japan," *Monetary and Economic Studies*, 22 (3), Institute for Monetary and Economic Studies, Bank of Japan, 2004, pp. 47–78 (this issue).

Fujino, Shozaburo, *Nihon no Mane Sapurai* (The Money Supply of Japan), Keisoshobo, 1994 (in Japanese).

Gregory, Allan W., and Bruce E. Hansen, "Residual-Based Test for Cointegration in Models with Regime Shifts," *Journal of Econometrics*, 70 (1), 1996, pp. 99–126.

Jefferson, Philip N., "On the Neutrality of Inside and Outside Money," *Economica*, 64 (256), 1997, pp. 567–586.

King, Robert G., and Mark W. Watson, "The Post-War U.S. Phillips Curve: A Revisionist Econometric History," *Carnegie-Rochester Conference Series on Public Policy*, 41, 1994, pp. 157–219.

———, and ———, "Testing Long Run Neutrality," *Economic Quarterly*, 83 (2), Federal Reserve Bank of Richmond, 1997, pp. 69–101.

Koustas, Zisimos, and Apostolos Serletis, "On the Fisher Effect," *Journal of Monetary Economics*, 44 (1), 1999, pp. 105–130.

Kwiatkowski, Denis, Peter C. B. Phillips, Peter Schmidt, and Yongcheol Shin, "Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root," *Journal of Econometrics*, 54 (1–3), 1992, pp. 159–178.

Lee, Junsoo, and Mark Strazicich, "Minimum LM Unit Root Tests with Two Structural Breaks," *Review of Economics and Statistics*, 85 (4), 2003, pp. 1082–1089.

Lucas, Robert E., Jr., "Expectations and the Neutrality of Money," *Journal of Economic Theory*, 4 (2), 1972, pp. 103–124.

———, "Some International Evidence on Output-Inflation Trade-Offs," *American Economic Review*, 63 (5), 1973, pp. 326–334.

Maddala, G. S., and In-Moo Kim, *Unit Roots, Cointegration, and Structural Change*, Cambridge: Cambridge University Press, 1998.

Nakashima, Kiyotaka, and Makoto Saito, "Strong Money Demand and Nominal Rigidity: Evidence from the Japanese Money Market under the Low Interest Rate Policy," mimeo, 2002.

Ng, Serena, and Pierre Perron, "Unit Root Tests in ARMA Models with Data-Dependent Methods for the Selection of the Truncation Lag," *Journal of the American Statistical Association*, 90 (429), 1995, pp. 268–281.

Obstfeld, Maurice, and Kenneth Rogoff, "Exchange Rate Dynamics Redux," *Journal of Political Economy*, 103 (2), 1995, pp. 624–660.

Ohkawa, Kazushi, Nobukiyo Takamatsu, and Yuzo Yamamoto, *Choki Keizai Tokei: 1. Kokumin Shotoku* (Estimates of Long-Term Economic Statistics of Japan since 1868: 1. National Income), Toyo Keizai, Inc., 1974 (in Japanese).

Okina, Kunio, "Relationship between Money Stock and Real Output in the Japanese Economy: Survey on the Empirical Tests of the LSW Proposition," *Monetary and Economic Studies*, 4 (1), Institute for Monetary and Economic Studies, Bank of Japan, 1986, pp. 41–77.

Olekalns, Nilss, "Some Further Evidence on the Long-Run Neutrality of Money," *Economic Letters*, 50 (3), 1996, pp. 393–398.

Perron, Pierre, "The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis," *Econometrica*, 57 (6), 1989, pp. 1361–1401.

———, "Further Evidence on Breaking Trend Function in Macroeconomic Variables," *Journal of Econometrics*, 80 (2), 1997, pp. 355–385.

Sargent, Thomas J., "A Note on the Accelerationist Controversy," *Journal of Money, Credit and Banking*, 3 (3), 1971, pp. 721–725.

———, "Observational Equivalence of Natural and Unnatural Rate Theories of Macroeconomics," *Journal of Political Economy*, 84 (3), 1976, pp. 631–640.

Serletis, Apostolos, and Zisimos Koustas, "International Evidence on the Neutrality of Money," *Journal of Money, Credit and Banking*, 30 (1), 1998, pp. 1–25.

———, and ———, "Monetary Aggregation and the Neutrality of Money," *Economic Inquiry*, 39 (1), 2001, pp. 124–138.

———, and David Krause, "Empirical Evidence on the Long-Run Neutrality Hypothesis Using Low-Frequency International Data," *Economic Letters*, 50 (3), 1996, pp. 323–327.

Soejima, Yutaka, "A Unit Root Test with Structural Change for Japanese Macroeconomic Variables," *Monetary and Economic Studies*, 13 (1), Institute for Monetary and Economic Studies, Bank of Japan, 1995, pp. 53–68.

Tobin, James, "Money and Economic Growth," *Econometrica*, 33 (4), 1965, pp. 671–684.

Weber, Axel A., "Testing Long-Run Neutrality: Empirical Evidence for G7 Countries with Special Emphasis on Germany," *Carnegie-Rochester Conference Series on Public Policy*, 41, 1994, pp. 67–117.

Yamada, Kazuo, "Nihon ni Okeru Kahei no Choki Churitsusei (Long-Run Monetary Neutrality in Japan)," *Osaka University Economics Journal*, 46 (3), 1997, pp. 46–54 (in Japanese).

Zivot, Eric, and Donald W. K. Andrews, "Further Evidence on the Great Crash, the Oil-Price Shock, and Unit-Root Hypothesis," *Journal of Business & Economic Statistics*, 10 (3), 1992, pp. 251–270.