An Empirical Study on Long-run Neutrality of Money in Japanese Economy

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(Received on May 8, 2007)

Abstract

Previous tests of the long-run neutrality of money hypothesis have generally relied on seasonally adjusted data and overlooked the important issues of seasonality. This paper analyses the long-run neutrality of money in Japan using quarterly seasonally unadjusted data, which permits an examination of the effects of seasonality and the robustness of previous empirical results. Fisher and Seater (1993) methodology is used with both seasonally unadjusted and adjusted Japanese real GDP and nominal money supply to test the long-run neutrality of money hypothesis. Using two measures of money stock, namely M1 and M2, it is shown that the hypothesis is supported using M2 as the measure of money supply, while it is rejected using M1.

1. Introduction

Monetary economists long have thought that government injections of money into a macroeconomy have a certain neutral effect. The main idea is that changes in the money stock eventually change nominal prices and nominal wages, ultimately leaving important real variables, like real output, real consumption expenditures, real wages, and real interest rates, unaffected. Since economic decision making is based on real factors, the long-run effect of injecting money into the macro economy is often described as *neutral* — in the end, real variables do not change and so economic decision making is also unchanged. How long such a process takes, and what might happen in the meantime, are hotly debated questions. Although there are many classical hypotheses to the efficacy of monetary policy, one hypothesis that is widely accepted among the economists and policymakers is the long-run neutrality of money. A formal definition of the long-run neutrality (LRN) of money is that a permanent, unexpected (exogenous) change to the level of money supply has no effect on the level of real output in the long run. Under LRN, changes in the money supply may or may not have short-run real effects. Related to the LRN of money is the long-run super neutrality (LRSN) of money,

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which occurs when permanent, exogenous changes in the growth rate of money supply leave the level of real output unaffected.

These long-run neutrality prepositions have been studied extensively, both theoretically and empirically and various econometric procedures are available for testing these hypotheses. Modern theoretical foundation was provided by Friedman (1969a, 1969b). Much of the existing empirical literature has been motivated by the work of McCallum (1984) who, drawing on the remarks made by Sargent (1971) and Lucas (1972, 1976), shows that the neutrality results obtained from frequency-domain methods are uninformative because of the problem of observational equivalence. McCallum (1984) claims that without knowing the time series properties of the money supply, both frequency-domain and reduced-form estimates are unable to discriminate empirically between long-run non-neutralities and the effects arising from autoregressive money supply specifications of the type developed by Lucas (1972). He lucidly points out that a valid test of the neutrality hypothesis can only be conducted using cross-equation restrictions in a bivariate vector autoregression (VAR).

In response to this, Fisher (1988), Fisher and Seater (1993), and also in a series of papers by King and Watson (1992, 1994, 1997) have advanced the dominant approaches in this line of research. The neutrality of money hypothesis has been tested for numerous countries using their methodologies. For example, Boschen and Otrok (1994), Boschen and Otrok (1994), Olekalns (1996), Serletis and Krause (1996), Haug and Lucas (1997), Coe and Nason (2003), Shelley and Wallace (2006) used Fisher and Seater method, while Weber (1994), Jefferson (1997), Serletis and Koustas (1998, 2001) employed King and Watson methods for testing the LRN and LRSN of money.

In spite of the progress in research on long-run neutrality of money, only a very limited number of comprehensive studies are available in Japan. Yamada (1997) shows that monetary neutrality holds in terms of real output, by applying Fisher and Seater's (1993) procedure to the Japanese seasonally adjusted quarterly data from 1957q1 to 1995q1. Using two types of data sets for century-long annual data covering 119 years from 1885 through 2003 as well as post-war seasonally adjusted quarterly data over the period 1955q2 – 2003q4, Oi *et al.* (2004) have found evidence supporting long-run neutrality, especially for the case of M2 as a measure of money stock. They used King and Watson's (1997) procedure.

But, an overlooked important issue in all previous papers in Japan is that of seasonality. Even most of the previous researches used seasonally adjusted data in testing long-run neutrality hypotheses. But in quarterly observe data, seasonality has been an issue that has attracted considerable research interest in the modeling of economic time series. Recently, some published papers have argued that modeling, instead of removing, seasonality may be beneficial for economic analysis, for example, Hylleberg (1992), Leong *et al.* (2000).

Traditionally, seasonality has been explicitly removed using seasonal adjustment procedures, such as using dummy variables or prefiltering the series using period-to-period differences or the ARIMA X-11 methodology. These methodologies have important drawbacks: Ghysels and Perron (1993), and Franses (1996) shows that seasonal adjustment procedures can remove cyclical fluctuations in the data, as seasonality and business cycles are typically correlated, and can thereby distort the data. Moreover, seasonal adjustments can affect the power of unit root and cointegration tests. If seasonality is deterministic, removing it with the aid of dummy variables has no effect whatsoever on unit root tests (Dickey *et al.*, 1984). However, when seasonal effects are stochastic, standard filters can greatly affect the power of unit root tests. Ghysels (1990) shows that removing seasonality using the X-11 method or the "variation in *s* periods" induces excess persistence in the series and consequently reduces the power of unit root tests to reject nonstationarity. Olekalns (1994) extends this result to the cases in which dummies or *band-pass* filters are used to remove seasonality. Abeysinghe (1994) shows that removing stochastic seasonality with dummy variables leads to the spurious regression problem.

The issues of long-run neutrality have been the heart of debates over the real effects of macroeconomic policies and are still very controversial topics among macroeconomic researchers. The empirical testability of these hypotheses is important for policy formulation and design, such as the effectiveness in monetary policy, whiles the determination of the order of integration, seasonal or nonseasonal, is crucial for these existing methods. So this paper presents the econometric treatment of seasonality for the purpose of testing for the long-run neutrality of money. Fisher and Seater (1993) approach is used to investigate the long-run neutrality of money in Japan over the extended period to 2006.[1] Consequently, both quarterly seasonally unadjusted and adjusted time series data on real GDP, nominal monetary variables M1 and M2 are used in this paper to examine the effects of seasonality and the robustness of previous empirical results in Japan.

1) We also used King and Watson (1997) method. The results (not shown) are qualitatively same as those obtained from Fisher and Seater (1993) method.

2. Overview of Long-Run Neutrality Tests

There has long been interest in testing long-run propositions about the link between money and real or nominal variables, which are at the heart of classical macroeconomics. In this section, the history of such tests is reviewed and the approach taken in this paper is highlighted.

One notable early strand of research on these issues was at the Federal Reserve Bank of St. Louis during the 1960s. The St. Louis researchers began the empirical study of the relationship between nominal income and the money stock in a dynamic regression framework,

$$\Delta Y_t = B(L)\Delta m_t + C(L)x_t + \varepsilon_t \tag{1}$$

where Y_t is log of nominal income, m_t is log of money stock, x_t are other variables that affect nominal income, ε_t is an error term, $B(L) = \sum_{i=0}^{m} B_i L^i$ and $C(L) = \sum_{i=0}^{f} C_i L^i$ in the lag operator L and $\Delta = 1 - L$ indicates a first difference.

The St. Louis researchers were motivated to study such distributed lag models by Friedman's (1969) argument that there was a lag in the effect of monetary actions on the macro economy. In the well-known work of Anderson and Jordan (1968), they sought to determine the nature of the lags in the effects of monetary policy in estimating the B coefficients in Equation (1). They found that there was a less than one-for-one short-run effect of money on nominal income, i.e., $B_0 < 1$. To calculate the effect of a sustained (long-run) change in the level of money on the path of nominal income, they calculated dynamic multipliers as follows:

$$\frac{\partial Y_{t+s}}{\partial m_t} = \sum_{i=0}^{s} B_i$$

Later St. Louis analysis, Andersen and Karnosky (1972), used this regression framework to test LRN as follows. They imagined a permanent change in the level of money have no long-run effect in the level of nominal income if

$$\lim_{s\to\infty} \frac{\partial Y_{t+s}}{\partial m_t} = \sum_{i=0}^{s} B_i = B(1) = 1$$

They also implemented the comparable test for LRN using log of real income (y_t) as

$$\lim_{s\to\infty} \frac{\partial y_{t+s}}{\partial m_t} = \sum_{i=0}^{s} B_i = B(1) = 0$$

The St. Louis approach was controversial. Notably, Ando and Modigliani (1990) criticized the St. Louis regression for not recognizing that nominal income and the money stock were simultaneously determined. However, Sims (1972) provided some support in a bivariate context for the St. Louis regression, building on Granger's (1969) earlier work on testing for causality. Theoretically, Sims established that it was only legitimate to run the regression

$$A(L)Y_t = B(L)m_t + e_{Y,t}$$

for the purposes of the St. Louis researchers if the reverse regression

$$C(L)m_t = D(L)Y_t + e_{m,t}$$

displayed D coefficients that were zero. Looking at nominal income and money empirically, he found evidence that the D coefficients were statistically insignificant.

The interpretation of the St. Louis regressions was also called into question by the analyses of Sargent (1971) and Lucas (1972). Studying an economy in which only unanticipated monetary changes had real effects and which otherwise displayed the LRN properties, Lucas (1972) showed that restrictions on sums of coefficients did not provide a way of testing the classical propositions. To illustrate Lucas's point, consider an economy in which the behavior of real and nominal income is given by

$$y_t = \phi(m_t - E_{t-1}m_t) + e_{y,t}$$
$$Y_t = m_t - \phi\theta(m_t - E_{t-1}m_t) + e_{Y,t}$$

where ϕ is a positive parameter and $0 < \theta < 1$. That is: unanticipated monetary expansions raise real income and raise nominal income less than one-for-one. If the money supply is given by the first-order autoregression

$$m_t = \rho m_{t-1} + e_{m,t},$$

it then follows that the rational expectations solutions for real and nominal output are

$$y_t = \phi m_t - \phi\rho m_{t-1} + e_{y,t} \tag{2a}$$
$$Y_t = (1-\phi\theta)m_t + \rho\phi\theta m_{t-1} + e_{Y,t} \tag{2b}$$

Under Lucas's assumption that the money supply process is stationary ($|\rho| < 1$), the sum of coefficients in the real output equation is inconsistent with neutrality ($\phi(1-\rho) > 0$). These implications occur despite the fact that the model is one with a strong form of neutrality. On the basis of this finding, Lucas and Sargent argued that if long-run variation in money is not a part of the environment that shapes the behavioral responses of economic agents, then a reduced form analysis — such as dynamic regressions or vector autoregressions — can never provide answers about the effect of long-run variations in money.

Concern about causality and the Lucas critique cast a shadow over applied research on long-run (LR) tests for nearly two decades. However, Fisher and Seater (1993) pointed out that the pessimism was not necessarily justified if economists are concerned about whether LR hypotheses held in a particular history and the historical data contained long-run variation in money. Fisher and Seater (1993) employ a bivariate structural vector autoregression (SVAR) model to test long-run neutrality and supper neutrality under the identification assumption that nominal variable is exogenous in the long-run. They demonstrate the restrictions implied by these prepositions are conditional on the order of integration of the variables comprised in the analysis. Specifically, neutrality tests are possible only if nominal as well as real variables are at least integrated of same order. Moreover, supper neutrality tests are possible only if the order of integration of the nominal variable is equal to one plus the order of integration of the real variable. Based on these two ideas, integration and identification, King and Watson (1992, 1997) provide a way of testing long-run neutrality by means of assuming distinct structural assumptions about the economy, while paying special attention to the integration and cointegration properties of the data. Their methodology does not assume money erogeneity, and proposes that the impact of money on output, or output on money be tested under alternative identifying restrictions in two variable SVAR models. Although all these test procedures differ in the details, each was based on the core idea that long-run propositions are testable if there is suitable long-run variation in money. As an example, suppose that the money stock is assumed to be a random walk ($\rho = 1$) in the Lucas model just considered. This assumption means that all changes in money are unanticipated and permanent. Evaluating the expressions above at $\rho = 1$, it then follows that the sum of coefficients on the monetary variables in (2a) is zero and the sum in (2b) is one as suggested by prior neutrality tests.

Neutrality tests based on SVAR are more complicated than the simple Lucas example on three dimensions. First, they allow real output to be potentially affected by real shocks in the long-run. Second, they allow for money growth to respond to its own lags and to lags of output growth. Third, they allow for short-run interactions of real and nominal variables. All of these considerations are reflected in the following structural vector autoregression,

$$\begin{aligned} \pi_{yy}\Delta y_t &= \pi_{ym}\Delta m_t + \sum_{i=1}^{p}\alpha_{yy,i}\Delta y_{t-i} + \sum_{i=1}^{p}\alpha_{ym,i}\Delta m_{t-i} + \varepsilon_{y,t} \\ \pi_{mm}\Delta m_t &= \pi_{my}\Delta y_t + \sum_{i=1}^{p}\alpha_{my,i}\Delta y_{t-i} + \sum_{i=1}^{p}\alpha_{mm,i}\Delta m_{t-i} + \varepsilon_{m,t} \end{aligned} \tag{3}$$

In this structural VAR, there are two structural shocks $\varepsilon_{y,t}$ and $\varepsilon_{m,t}$ with mean 0 and $\mathrm{cov}(\varepsilon,\varepsilon') = \Sigma_\varepsilon$, where $\varepsilon = (\varepsilon_{y,t}, \varepsilon_{m,t})'$. The former refers to real productivity shocks while the latter refers to monetary shocks. In addition, both variables are treated endogenously, which mitigates the causality problems previously discussed in the context of the St. Louis regression. Short-run interactions of y_t and m_t are governed by the π coefficients, while the dynamic interactions are governed by the α coefficients. The above system can be written as

$$\begin{bmatrix} \pi_{yy}-\sum_{i=1}^{p}\alpha_{yy,i}L^i & -(\pi_{yy}+\sum_{i=1}^{p}\alpha_{ym,i}L^i) \\ -(\pi_{my}+\sum_{i=1}^{p}\alpha_{my,i}L^i) & \pi_{mm}-\sum_{i=1}^{p}\alpha_{mm,i}L^i \end{bmatrix}\begin{bmatrix}\Delta y_t \\ \Delta m_t\end{bmatrix}=\begin{bmatrix}\varepsilon_{y,t} \\ \varepsilon_{m,t}\end{bmatrix}$$

where L is lag operator, that is, $L^i x_t = x_{t-i}$. Assuming both y and m are I(1) and not cointegrated, the long-run responses of y_t and m_t to the shocks $\varepsilon_{y,t}$ and $\varepsilon_{m,t}$, are the solutions to the above system, that can be approximated as

$$\lim_{k\to\infty}\begin{bmatrix} \partial y_{t+k}/\partial\varepsilon_{y,t} & \partial y_{t+k}/\partial\varepsilon_{m,t} \\ \partial m_{t+k}/\partial\varepsilon_{y,t} & \partial m_{t+k}/\partial\varepsilon_{m,t} \end{bmatrix}=\varphi\begin{bmatrix} \pi_{mm}-\sum_{i=1}^{p}\alpha_{mm,i} & \pi_{ym}+\sum_{i=1}^{p}\alpha_{ym,i} \\ \pi_{my}+\sum_{i=1}^{p}\alpha_{my,i} & \pi_{yy}-\sum_{i=1}^{p}\alpha_{yy,i} \end{bmatrix}$$

where $\varphi=\dfrac{1}{(\pi_{yy}-\sum_{i=1}^{p}\alpha_{yy,i})(\pi_{mm}-\sum_{i=1}^{p}\alpha_{mm,i})-(\pi_{my}+\sum_{i=1}^{p}\alpha_{my,i})(\pi_{ym}+\sum_{i=1}^{p}\alpha_{ym,i})}$. The long-run effect of monetary shock ε_m is

$$\lim_{k\to\infty}\begin{bmatrix} \dfrac{\partial y_{t+k}/\partial\varepsilon_{m,t}}{\partial m_{t+k}/\partial\varepsilon_{m,t}} \end{bmatrix}=\varphi\begin{bmatrix} \pi_{ym}+\sum_{i=1}^{p}\alpha_{ym,i} \\ \pi_{yy}-\sum_{i=1}^{p}\alpha_{yy,i} \end{bmatrix}$$

and the long-run effect of real shock ε_y is

$$\lim_{k\to\infty}\begin{bmatrix} \dfrac{\partial y_{t+k}/\partial\varepsilon_{y,t}}{\partial m_{t+k}/\partial\varepsilon_{y,t}} \end{bmatrix}=\varphi\begin{bmatrix} m_{mm}+\sum_{i=1}^{p}\alpha_{mm,i} \\ \pi_{my}-\sum_{i=1}^{p}\alpha_{my,i} \end{bmatrix}.$$

Based on the above specifications, the long-run neutrality of money with respect to real output can be tested using the following long-run multiplier:

$$\gamma_{ym}=\lim_{k\to\infty}\frac{\partial y_{t+k}/\partial\varepsilon_{m,t}}{\partial m_{t+k}/\partial\varepsilon_{m,t}}=\frac{\pi_{ym}+\sum_{i=1}^{p}\alpha_{ym,i}}{\pi_{yy}-\sum_{i=1}^{p}\alpha_{yy,i}} \tag{4}$$

The expression in equation (4) is at the heart of Fisher and Seater (1993) and King and Watson (1997) tests. Fisher and Seater (1993) term the ratio as long-run derivative (LRD) of y with respect to m, and King and Watson (1997) called the long-run elasticity of output (y) with respect to permanent exogenous change in money (m).

If both money and real GDP are integrated of order one, I(1), (or same order) then the implication of long-run neutrality of money is that

$$\mathrm{LRD} = \gamma_{ym} = \lim_{k\to\infty} \frac{\partial y_{t+k} / \partial \varepsilon_{m,t}}{\partial m_{t+k} / \partial \varepsilon_{m,t}} = \frac{\pi_{ym} + \sum_{i=1}^{p} \alpha_{ym,i}}{\pi_{yy} - \sum_{i=1}^{p} \alpha_{yy,i}} = 0 \tag{5}$$

Long-run supper neutrality (LRSN) can be tested when money is I(2) and real output is I(1). Using the growth of money (Δm_t) in stead of money (m_t) in SVAR model in Equation (3), the LRD can be obtained as

$$\gamma_{y\Delta m} = \lim_{k\to\infty} \frac{\partial y_{t+k} / \partial \varepsilon_{\Delta m,t}}{\partial \Delta m_{t+k} / \partial \varepsilon_{\Delta m,t}} = \frac{\pi_{ym} + \sum_{i=1}^{p} \alpha_{y\Delta m,i}}{\pi_{yy} - \sum_{i=1}^{p} \alpha_{yy,i}} \tag{6}$$

Note that the LRD is not defined for the case in which $\lim_{k\to\infty}(\partial m_{t+k} / \partial \varepsilon_{m,t}) = 0$. Therefore, a necessary condition for testing LRN is that there have been permanent shocks to money supply. Money therefore must be at least first order integrated, or I(1), to apply the test. Equations (4) and (6) also show that if there have no permanent shocks to real output, then $\lim_{k\to\infty}(\partial y_{t+k} / \partial \varepsilon_{m,t})$ and $\lim_{k\to\infty}(\partial y_{t+k} / \partial \varepsilon_{\Delta m,t})$ equal to zero respectively, that is, the LRD is equal to zero. Therefore, if the out put is I(0), LRN and LRSN cannot be rejected.

3. Identifying Restrictions

By its nature, the long-run multiplier γ_{ym} is a structural parameter, which requires identifying assumptions to estimate it. These identifying assumptions are most easily discussed if we follow the actual practice used in some of the prior literature. To begin, suppose that we estimate a reduced form VAR as

$$\Delta y_t = \sum_{i=1}^{p} b_{yy,i} \Delta y_{t-i} + \sum_{i=1}^{p} b_{ym,i} \Delta m_{t-i} + u_{y,t}$$

$$\Delta m_t = \sum_{i=1}^{p} b_{my,i} \Delta y_{t-i} + \sum_{i=1}^{p} b_{mm,i} \Delta m_{t-i} + u_{m,t}$$

In this reduced form VAR, $u_{y,t}$ and $u_{m,t}$ are error terms with mean 0 and $\mathrm{cov}(u,u') = \Sigma_u$, where $u = (u_{y,t}, u_{m,t})'$. Comparing this expression with (3), we note a structural relationship between reduced form shocks u_t and structural shocks ε_t that takes the form

$$\begin{bmatrix} u_{y,t} \\ u_{m,t} \end{bmatrix} = \begin{bmatrix} \pi_{yy} & -\pi_{ym} \\ -\pi_{my} & \pi_{mm} \end{bmatrix}^{-1} \begin{bmatrix} \varepsilon_{y,t} \\ \varepsilon_{m,t} \end{bmatrix}$$

with covariance matrix

$$\begin{bmatrix} \mathrm{var}(u_{y,t}) & \mathrm{cov}(u_{y,t},u_{m,t}) \\ \mathrm{cov}(u_{y,t},u_{m,t}) & \mathrm{var}(u_{m,t}) \end{bmatrix}$$
$$= \begin{bmatrix} \pi_{yy} & -\pi_{ym} \\ -\pi_{my} & \pi_{mm} \end{bmatrix}^{-1} \begin{bmatrix} \mathrm{var}(\varepsilon_{y,t}) & \mathrm{cov}(\varepsilon_{y,t},\varepsilon_{m,t}) \\ \mathrm{cov}(\varepsilon_{y,t},\varepsilon_{m,t}) & \mathrm{var}(\varepsilon_{m,t}) \end{bmatrix} \begin{bmatrix} \pi_{yy} & -\pi_{ym} \\ -\pi_{my} & \pi_{mm} \end{bmatrix}'^{-1}$$

and the corresponding parameters are

$$\begin{bmatrix} b_{yy,i} & b_{ym,i} \\ b_{my,i} & b_{mm,i} \end{bmatrix} = \begin{bmatrix} \pi_{yy} & -\pi_{ym} \\ -\pi_{my} & \pi_{mm} \end{bmatrix}^{-1} \begin{bmatrix} \alpha_{yy,i} & \alpha_{ym,i} \\ \alpha_{my,i} & \alpha_{mm,i} \end{bmatrix}; \quad i = 1, 2, \ldots, p$$

The bivariate structural VAR of order p, as in equation (3), has $2^2 \times (p + 1)$ unknowns in the coefficients and 3 (= 2 × (2+1)/2) unknowns in the covariance matrix of the residual. Meanwhile, the bivariate reduced-form VAR of order p provides estimates of $2^2 \times p$ parameter values for the coefficients and three values for the covariance matrix of the reduced form errors. Accordingly, $2^2 = 4$ identifying restrictions must be placed to identify the structural shocks $\varepsilon_{y,t}$ and $\varepsilon_{m,t}$. Standardizing π_{yy} and π_{mm} to unity and using the assumption that the structural shocks are mutually uncorrelated, i.e., $\mathrm{cov}(\varepsilon_{y,t}, \varepsilon_{m,t}) = 0$, one additional identifying restriction is required. For this reason, to compute the γ_{ym}, the different researchers imposed different identifying assumptions.

Fisher and Seater (1993) assumed that money is long-run exogenous by imposing $(\pi_{my} + \sum_{i=1}^{p} \alpha_{my,i}) = 0$. That is, the real shock $\varepsilon_{y,t}$ is assumed not to affect the variation in money in the long-run. Formally, given this assumption, the second equation of (3) governing the long-run response of money implies that: the long-run response of money does not depend on the long-run response of output. Using this assumption, Fisher and Seater demonstrate that γ_{ym} can be consistently estimated as the slope coefficient in the following OLS regression:

$$(y_t - y_{t-k-1}) = \alpha_k + \beta_k (m_t - m_{t-k-1}) + \varepsilon_{kt} \tag{7}$$

That is, LRD can be approximated by $\lim_{k\to\infty} \beta_k$. Standard practice is to estimate β_k using OLS for each value of k taking values of one through a predetermined upper limit. With T observations, the 95-percent confidence intervals then are constructed for the β_k's from a t-distribution with T/k degrees of freedom using standard errors corrected for serial correlation by the Newey-West procedure. Long-run neutrality of money is rejected if zero lies outside the confidence interval as k become large.

For the LRSN of money, which is testable if m is I(2) and y is I(1), the estimator of γ_{ym} is the slope coefficient in

$$(y_t - y_{t-k-1}) = \alpha_k + \beta_k (\Delta m_t - \Delta m_{t-k-1}) + \varepsilon_{kt} \tag{8}$$

King and Watson (1997) emphasize the relationship between test results and identifying issues. They analyze long-run neutrality propositions across a range of possible identifications of their bivariate structural VAR system, in an effort to understand the robustness of various conclusions to differing assumptions. They used one of the following identifying restrictions in their LR tests.

(i) the impact elasticity of y with respect to m is known (i.e., π_{ym} is known),

(ii) the impact elasticity of m with respect to y is known (i.e., π_{my} is known),

(iii) the long-run elasticity of y with respect to m is known (i.e., γ_{ym} is known),

(iv) the long-run elasticity of m with respect to y is known (i.e., γ_{my} is known).[2)]

First, the short-term restriction specifies a contemporaneous relationship between endogenous variables and shocks by imposing restrictions on the short-term elasticity, such as suppose $\pi_{ym} = 0$ or $\pi_{my} = 0$. 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, consider $\gamma_{my} = 1$. This assumption is consistent with long-run price stability under the assumption of stable volatility. Additionally, long-run neutrality ($\gamma_{ym} = 0$) is applicable as an identifying restriction. Varying the values of π_{ym}, π_{my} and γ_{my}, they found the evidence of supporting long-run neutrality of money for real output in the postwar U.S. data.

4. Tests for Seasonality

Seasonality can be deterministic and/or stochastic. For quarterly data, deterministic seasonality assumes that the data generating process for the variable y_t is

$$y_t = \gamma_1 s_{1t} + \gamma_2 s_{2t} + \gamma_3 s_{3t} + \gamma_4 s_{4t} + \varepsilon_t$$

where s_{st} (= 1 in season s, 0 elsewhere, for s = 1, 2, 3, 4) is a seasonal dummy variable. Including seasonal dummy variables in a regression model is appropriate for variables with deterministic seasonality. The absence of these dummy variables will lead to the standard problem of

2) $\gamma_{my} = \dfrac{\pi_{my} + \sum_{i=1}^{p} \alpha_{my,i}}{\pi_{mm} - \sum_{i=1}^{p} \alpha_{mm,i}}$ denotes the long-run elasticity of money (m) with respect to permanent exogenous change in output (y).

bias associated with the exclusion of relevant explanatory variables. Stochastic seasonality extends the unit root hypothesis to seasonal time series. An integrated seasonal process is a process that contains unit roots at the seasonal frequencies, and appropriate differencing filters are required for seasonally integrated processes.

One method to test for seasonality is due to Miron (1994), and involves the use of an auxiliary regression of the form

$$(1-L)y_t = \gamma_1 s_{1t} + \gamma_2 s_{2t} + \gamma_3 s_{3t} + \gamma_4 s_{4t} + \varepsilon_t \tag{9}$$

where ε_t is assumed to be a stationary and invertible ARMA process. If the logarithm of y_t is taken as the dependent variable, the equation involves the regression of the growth rate of the variable on a set of seasonal dummy variables. The explanatory power of the regressors would indicate the extent of the seasonality in the series.

If seasonal components are stochastic, the variable can have a unit root not only in its long-run behavior, but also at its seasonal frequencies. To test the unit root properties of variables with a substantial seasonal pattern Hylleberg *et al.* (1990) have proposed a generalization to the Dickey-Fuller approach, denoted as HEGY-test and applied in Engle *et al.* (1993). The procedure is based on the assumption that the individual time series is generated by a finite autoregressive process:

$$\Delta_4 y_t = (1-L^4)y_t = \varepsilon_t$$

with ε_t as a zero mean white noise. Where Δ_4 being the differencing filter defined as $\Delta_4 = (1 - L^4)$ with L being the usual lag operator. If the process y_t is integrated and seasonally integrated then the differencing operator $(1 - L^4) = (1 - L)(1 + L)(1 + L^2) = (1 - L)(1 + L)(1 - iL)(1 + iL) = 0$ has two real roots 1 and −1 and two complex roots i and $-i$. The real root 1 is usual zero frequency (non-seasonal) root which can be removed by applying the filter $(1 - L)$, while the real root −1 corresponds to unit root at ½ cycle per quarter (semi-annual unit root) that is eliminated by the filter $(1 + L)$. The complex roots i and $-i$ correspond to unit roots at ¼ cycle (annual unit roots) that can be filtered by means of $(1 + L^2)$.

To analyze the unit roots at all seasonal frequencies and zero frequency considering the deterministic seasonality in y_t, the process is rearranged as

$$\Delta_4 y_t = \alpha + \delta t + \gamma_1 s_{1t} + \gamma_2 s_{2t} + \gamma_3 s_{3t} + \pi_1 y_{1,t-1} + \pi_2 y_{2,t-1} + \pi_3 y_{3,t-2} + \pi_4 y_{3,t-1} + \varepsilon_t \tag{10}$$

where t is deterministic time trend, s_{1t}, s_{2t}, and s_{3t} are seasonal dummy variables shown as above, and

$$y_{1,t} = (1+L)(1+L^2)y_t$$
$$y_{2,t} = -(1-L)(1+L^2)y_t$$

$y_{3,t} = -(1-L)(1+L)y_t$

where the signs in defining $y_{2,t}$ and $y_{3,t}$ are used just for convenience. Given a quarterly integrated series, $y_{1,t}$ hence just contains the usual long-run trend, while $y_{2,t}$ and $y_{3,t}$ consist of the nonstationary bi-annual and annual cycles, respectively.

In order to determine the differencing filters that are necessary for the stationarity of the variable, Hylleberg et al. (1990) suggest the following procedure. Estimate equation (4) by OLS and test the following hypotheses:

(i) H_0: $\pi_1 = 0$, H_1: $\pi_1 < 0$;

(ii) H_0: $\pi_2 = 0$, H_1: $\pi_2 < 0$;

(iii) H_0: $\pi_3 = \pi_4 = 0$, H_1: $\pi_3 \neq 0$; or/and $\pi_4 \neq 0$.

Standard *t*-tests are used for the first two hypotheses (i) and (ii), while an *F*-test is used for the third hypothesis (iii). If the hypothesis (i) is not rejected, there is a non-seasonal or zero frequency unit root in the series, which requires the filter $(1 - L)$ for stationarity. For the hypothesis (ii), the null hypothesis is consistent with a semi-annual unit root, implying that any shocks to the variable will lead to permanent changes in the seasonal pattern of the variable at the semi-annual level. This requires the filter $(1 + L)$ for stationarity. A non-rejection of the third null hypothesis (iii) implies that the series has at least one of the two unit roots in the annual frequency, requiring the filter $(1 + L^2)$ for stationarity. With annual unit roots, a shock to the variable will change permanently the seasonal pattern of the variable at the annual level. The empirical rejection of the three null hypotheses implies that the series has no non-seasonal, semi-annual or annual unit roots, respectively, for quarterly data. The asymptotic distributions of the HEGY statistics are non-standard, and are functionals of Wiener processes. Hylleberg *et al.* (1990) compute the appropriate critical values for these tests by Monte Carlo simulations.

5. Data and Description

We investigate the long-run neutrality of money hypotheses using postwar data for Japan. Both seasonally unadjusted and adjusted quarterly time series over the period 1955:q2 – 2006:q1 are used in this study. Quarterly real GDP, and nominal money supplies M1 and M2 + CD, are used as the real output and nominal money supply series. We used Bank of Japan (BOJ) monthly monetary statistics for M1 and M2 + CD. Monthly series were used to form the quarterly data. That is, the observation for the following month is used as a proxy for the end-of-quarter money. For example, our M1 figure for the first quarter of 1990 is the value of

M1 on April, 1990. Whereas, quarterly real GDP series was taken from the System of National Accounts (SNA). The following symbolic notations are used through out this article.

y_t = log(real GDP)

$m1_t$ = log(M1)

$m2_t$ = log(M2 + CD)

Although $m2$ is the major indicator of board monetary aggregate, consideration of two measures of money supply, namely $m1$ and $m2$, serves as a sensitivity analysis of the potential effects of money on real output. It is shown in Bullard (1994) and Olekalns (1996) that the outcome of the test of LRN is sensitive to measure the money involved.

6. Time Series Properties of Data

Identification of the orders of integration of money and real gdp is an important issue before testing long-run neutrality using the methodology discussed in above. Since we used seasonally unadjusted data, it is also important to check wether the series are seasonally integrated. In this section we employed HEGY test to investigate the seasonal integration, while DF-GLS and KPSS tests were used to identify the order of nonseasonal integrations in the data used in this paper.

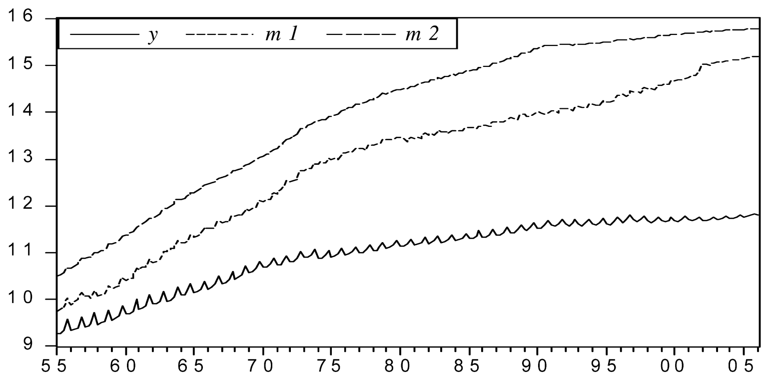


Figure 1. Logarithm of quarterly unadjusted real GDP (*y*) and nominal money supplies (*m*1 & *m*2)

Figure 1 contains the plots of the real GDP y, nominal money supply $m1$ and $m2$ in level. All three variables are trending upward. It can also be observed in Figure 1 that real GDP exhibits substantial seasonal fluctuations. To guess whether the seasonality is deterministic or stochastic, we used the auxiliary regression in Equation (9) proposed by Miron (1994). Regression (9) is performed for the real GDP and the two monetary variables $m1$ and $m2$, for the full sample 1955:2 to 2006:1, as well as two subset samples, each with an approximately equal number of observations. Consideration of split samples allows an examination of whether the seasonal fluctuations have changed over time, namely, whether there is stochastic seasonality. Regression results for the estimates of the parameters in the auxiliary regression, and their R^2 values, are presented in Table 1. The high R^2 values suggest that real GDP exhibits substantial seasonal fluctuations. As expected, the explanatory powers of the seasonal dummy variables for the two monetary variables $m1$ and $m2$ are comparatively low.

Estimates of the γ_s (s = 1, 2, 3, 4) coefficients can be used to observe the pattern of seasonality. For the full and subsamples, the estimated coefficients of the seasonal dummy variables are very similar for real GDP and nominal monetary variable $m2$. This suggests that the pattern of seasonality has remained relatively constant over the entire sample for these two series, and suggests that seasonal unit roots are likely to be absent. But the sign and significance level of first two dummy coefficients of nominal monetary variable $m1$ are not consistent with full and subsamples. So the monetary variable $m1$ may be suspected to have seasonal unit root.

Table 1. Seasonality in real GDP *y* and nominal money supplies *m*1 and *m*2

Variable	Sample	R^2	γ_1	γ_2	γ_3	γ_4
y	1955:2–2006:1	0.88	−0.0390 (−3.680)	−0.0502 (−5.193)	0.0723 (10.976)	0.1724 (22.657)
	1955:2–1980:3	0.95	−0.0587 (−3.492)	−0.0625 (−3.998)	0.0772 (9.398)	0.2113 (22.146)
	1980:4–2006:1	0.90	−0.0555 (−3.434)	−0.0240 (−1.585)	0.0579 (4.527)	0.1092 (7.711)
$m2$	1955:2–2006:1	0.79	0.0257 (3.990)	0.0549 (17.555)	0.0414 (8.943)	0.0826 (21.185)
	1955:2–1980:3	0.84	−0.0091 (−0.957)	0.0477 (13.334)	0.0283 (4.760)	0.0790 (14.872)
	1980:4–2006:1	0.75	0.0412 (4.674)	0.0562 (8.956)	0.0428 (5.673)	0.0743 (11.498)
$m1$	1955:2–2006:1	0.68	0.0272 (2.516)	0.0339 (5.192)	0.0154 (2.233)	0.1247 (19.104)
	1955:2–1980:3	0.83	−0.02574 (−1.473)	0.0406 (4.749)	0.0151 (1.686)	0.1503 (18.500)
	1980:4–2006:1	0.64	0.0100 (0.603)	−0.0103 (−0.667)	−0.0364 (−2.333)	0.0514 (3.238)

Note: Figures in parentheses show t-values

When seasonally unadjusted data are used, the unit root tests cannot detect seasonal unit roots or stochastic seasonality. Therefore, when the outcomes of unit root tests suggest that a series should be first-differenced, the series may not necessarily be stationary due to the possible presence of seasonal unit roots. As such, testing for seasonal unit roots in quarterly observed data is of paramount importance.

Table 2. HEGY tests for seasonal integration

Variable	H_0: $\pi_1 = 0$	H_0: $\pi_2 = 0$	H_0: $\pi_3 = \pi_4 = 0$
y	−1.5039	−2.654**	19.513***
$m2$	0.3298	−7.2433***	124.00***
$m1$	−1.2096	−6.3419***	98.4274***

Note: The critical values are taken from HEGY for 200 observations.: ***, ** and * indicate the significant level at 1%, 5% and 10% respectively

Thus, the three series considered in this paper are tested for possible seasonal unit roots using the HEGY test using Equation (10). The results of HEGY test under three hypotheses are presented in Table 2. It is found that for all three variables hypotheses (ii) and (iii) are rejected at highly significance level, while hypothesis (i) is not rejected. These results imply that all three variables are not seasonally integrated, namely, the seasonally unadjusted quarterly series do not contain seasonal unit roots at the semi-annual or the annual frequencies and the variables do, however, have unit roots at the zero or non-seasonal frequencies. These results are consistent with the results of previous tests of LRN.

Conventional unit root tests are also used to confirm the I(1) nature of the three series. Two unit root tests are performed: (i) the asymptotically most powerful DF-GLS test of Elliott

Table 3. Unit Root Tests for Levels and First Differences Series

variable	Unadjusted			Adjusted		
	lag	DF-GLS	KPSS	lag	DF-GLS	KPSS
y	4	−0.5093	0.4288***	4	−0.3252	0.4248***
$m2$	8	−1.2084	0.4535***	5	−0.6521	0.4515***
$m1$	8	−0.6932	0.4185***	2	−0.2376	0.4114***
Δy	2	−41.7683***	0.0515	2	−4.1079***	0.0975
$\Delta m2$	6	−3.3507**	0.0681	1	−3.3731**	0.0925
$\Delta m1$	7	−2.9668**	0.1282*	1	−6.4649***	0.1876**

Note: Constant and trend are included in the models.: ***, ** and * indicate the significant level at 1%, 5% and 10%, respectively.

et al. (1996) and (ii) the Kwiatkowski *et al.* (1992) LM test (KPSS). The null hypothesis of the former test is that a variable has a unit root while that of the later test is that a variable is stationary. A common strategy is to present results of both DF-GLS and KPSS tests, and show that the results are consistent (e.g., that the former reject the null while the later fail to do so and vice-versa). The lag length is selected by the Akaike Information Criteria (AIC). The results are shown in Table 3. We perform the same test on the first differenced variables and found that all variables are I(1). That is, both tests indicate that the variables are integrated at the zero frequency.

As real GDP is not seasonally integrated, the seasonal variations are likely to be deterministic. In order to test the LRN hypothesis using real GDP as the output variable, three seasonal dummy variables and a trend term are included in Equation (7) to capture the seasonal variations. Equation (7) is modified as follows:

$$(y_t - y_{t-k-1}) = \alpha + \delta t + \gamma_1 s_{1t} + \gamma_2 s_{2t} + \gamma_3 s_{3t} + \beta_k (m_t - m_{t-k-1}) + \varepsilon_t \tag{11}$$

Shelley and Wallace (2006) shown that the inclusion of trend and seasonal dummy in the FS equation does not bias the estimates of the β_k coefficients.

The Fisher and Seater (1993) methodology requires money be exogenous. The validity of this assumption is assessed by testing if real GDP growth Granger-causes money growth. We used the VAR model for testing the Granger-causality where the null hypothesis is that real GDP growth does not Granger-causes money growth. The lag length was determined using the AIC criteria, with a maximum of 15 lags considered. The criteria select 12 and 10 lags for $m1$ and $m2$ respectively in case of seasonally unadjusted series. For $m1$ and $m2$ lags 3 and 9 respectively are used when the series are seasonally adjusted. Table 4 contains the results for both seasonally unadjusted and adjusted data.

Table 4. Granger-causality tests for erogeneity of money

Hypothesis	χ^2-value (unadjusted)	χ^2-value (adjusted)
y does not causes $m1$	17.051 (12) [0.1477]	0.420 (3) [0.9360]
y does not causes $m2$	12.072 (10) [0.2803]	7.138 (9) [0.6228]

Note: ( ) and [ ] contain lags and p-values respectively

As table shows, the null hypothesis of real GDP does not Granger-causes money cannot be rejected for each of money and real GDP series. So we conclude that for both cases, seasonally unadjusted and adjusted, money is exogenous.

7. Empirical Results of Long-Run Neutrality of Money with Seasonality

Since the Granger-Causality tests indicate $m1$ and $m2$ are exogenous with respect to y and each of the money series and real GDP are integrated of order one, we can proceed with the neu-

Table 5. Estimation results for the long-run neutrality of money: $m1$ & $m2$

k (lags)	n (observations)	For m2		For m1	
		β_k	Standard error (Newly-West)	β_k	Standard error (Newly-West)
1	203	1.1567	0.2252	0.4527	0.0696
2	202	0.6985	0.1626	0.3381	0.0719
3	201	0.2708	0.0797	0.0882	0.0444
4	200	0.5015	0.1226	0.2358	0.0611
5	199	0.479	0.1202	0.2133	0.0519
6	198	0.408	0.1141	0.2086	0.0522
7	197	0.2839	0.0894	0.1247	0.0409
8	196	0.3582	0.1122	0.183	0.0479
9	195	0.4000	0.1116	0.1701	0.0447
10	194	0.2943	0.1097	0.162	0.045
11	193	0.2312	0.1029	0.1161	0.0402
12	192	0.2697	0.1143	0.1452	0.0445
13	191	0.2588	0.1169	0.1389	0.0447
14	190	0.2341	0.119	0.136	0.0457
15	189	0.1982	0.1187	0.1121	0.0449
16	188	0.2242	0.1239	0.1307	0.0466
17	187	0.2118	0.1261	0.1211	0.0467
18	186	0.1927	0.129	0.1185	0.0471
19	185	0.1618	0.1299	0.1013	0.0473
20	184	0.1833	0.1351	0.114	0.0478
21	183	0.1711	0.1385	0.1069	0.0477
22	182	0.1562	0.1428	0.107	0.0479
23	181	0.1381	0.1446	0.0971	0.0477
24	180	0.1626	0.1492	0.1107	0.0473
25	179	0.1583	0.1516	0.1095	0.0467
26	178	0.1499	0.1593	0.1102	0.0463
27	177	0.141	0.1546	0.1037	0.0459
28	176	0.165	0.1578	0.1155	0.0454
29	175	0.161	0.159	0.1125	0.0449
30	174	0.151	0.160	0.1149	0.0446

trality tests under the Fisher and Seater (1993) methodology. Equation (11) for two nominal monetary variables $m1$ and $m2$ are separately estimated when y is dependent variable for values of $k = 1, 2, \ldots, 30$. Each coefficient, β_k, presents the estimated response of the change in log real GDP to the change in logged money over $k + 1$ periods which are shown in Table 5. Test outcomes for LRN can be observed by examining a plot of the estimates of β_k and its 95% confidence interval against the lag length k. As such, the residuals from the regression for the various lags may be non-spherical, possibly leading to biased t-ratios and outcomes of the LRN tests. Following Fisher and Seater, the 95% confidence intervals derived for the estimated coefficient of money supply are obtained using standard errors that are adjusted using the Newey-West (1987) procedure. The t-distribution with n/k degrees of freedom is used to construct the confidence intervals. Long-run neutrality of money is rejected if zero lies outside the confidence interval as k become large.

In Figure 2(a), we first present the graph of the coefficient β_k and its 95% confidence intervals when the $(k + 1)$ difference of nominal money $m2$ is the independent variable in Equation (11). The value of k is ranging 1 to 30 over the period 1952q2 – 2006q1. As can be seen from the graph, the estimated coefficients are significantly positive for the values of k less than 14 quarters, suggesting a short-run positive effect of monetary policy using $m2$. But these

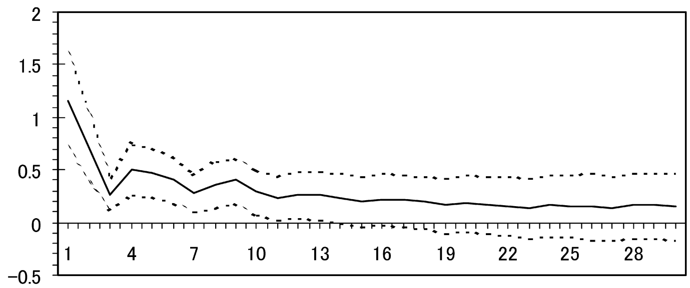


Figure 2(a). Long-run neutrality of money using unadjusted *m*2

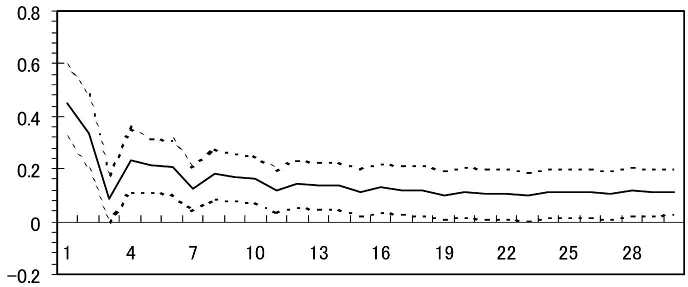


Figure 2(b). Long-run neutrality of money using unadjusted *m*1

short-run effects disappear after about three years. That is, as the lag length increases, there is an obvious downward trend in the plot of the estimates. The confidence interval bands include 0 for high values of $k \geq 14$. This suggests that $m2$ does not affect real GDP in the long run. So it is observed that the LRN hypothesis is supported by the nominal money $m2$. Since money supply $m2$ is not integrated of order 2, the LRSN of $m2$ cannot be analyzed.

Figure 2(b) shows the β_k coefficient plot and confidence intervals against the lag length k when nominal money $m1$ is the independent variable in Equation (11) over the period 1952q2 – 2006q1. For all values of k, 95% confidence interval lies over the zero line, suggesting a positive correlation between $m1$ and real GDP in the short-run as well as in the long-run. Hence LRN of money using $m1$ is rejected by the data. This result indicates the sensitivity of the LRN results to the type of money supplies considered. Since the LRN hypothesis is rejected using $m1$ as the measure of money supply, the LRSN hypothesis using $m1$ is also rejected, that is, the hypothesis that permanent stochastic changes to the growth rate of $m1$ ultimately leave the level of real GDP unchanged is rejected.

These results are consistent with those of Yamada (1996) and Oil *et. al.* (2004), where Japanese seasonally adjusted quarterly data are used. The LRN hypothesis is supported using $m2$ as the money supply, while the use of $m1$ leads to the LRN hypothesis being rejected for Japan.

8. Sensitivity Test

To examine the sensitivity of the above test results, we have used the natural logarithms of quarterly seasonally adjusted real GDP, M1 and M2 in for LRN over the same period in Fisher and Seater method. Unit root tests show (in Table 3) that all variables are integrated of order one, I (1). Granger-causality test (in Table 4) support the exogeneity of money for both $m1$ and $m2$, which is the key assumption in Fisher and Seater model. Since the data are seasonally adjusted, we exclude the seasonal dummy variable from Equation (11) and use the following OLS regression model.

$$(y_t - y_{t-k-1}) = \alpha + \delta t + \beta_k (m_t - m_{t-k-1}) + \varepsilon_t \tag{12}$$

Estimated coefficient and corresponding 95% confidence interval using $m2$ and $m1$ as nominal money supply in Equation (12) for different lag lengths are presented in Figures 3(a) and 3(b) respectively. The outcomes of the tests using seasonally adjusted $m2$ and $m1$ are qualitatively the same as those obtained using seasonally unadjusted data, namely, that the LRN hypothesis is

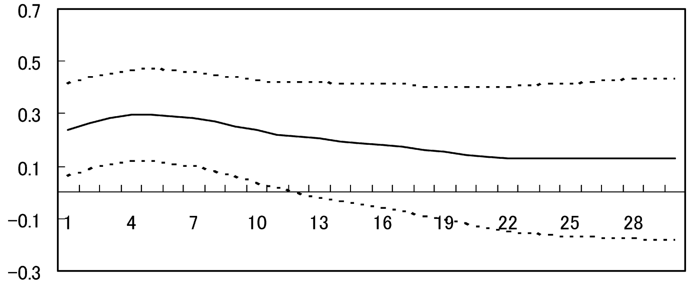


Figure 3(a). Long-run neutrality of money using adjusted *m*2

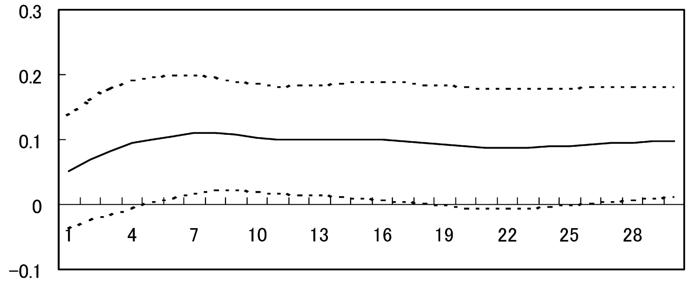


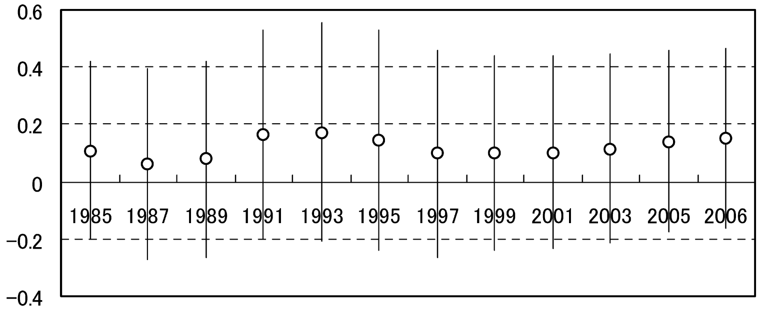
Figure 3(b). Long-run neutrality of money using adjusted *m*1

not rejected using $m2$ but is rejected using $m1$. But, the result of $m1$ is somewhat different in the short-run ($k \leq 5$) and in the medium-run ($20 \leq k \leq 26$). It shows short-run neutrality as well as mixed effects in the long-run.

Thus, the test results for LRN presented in this paper are robust to the standard seasonal adjustment transformation.

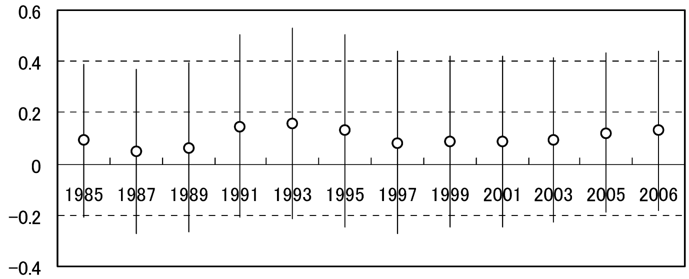
Finally, to check the robustness of our results in terms of $m2$, we examine the effects of including the data under the bubble economy (before 1989), after the burst of the bubble (after 1989) and the recent zero interest rates (after 1995). We estimate the long-run multiplier (γ_{ym}) by sequentially extending the end of the sample period (2 years period basis) from 1985 to 2006. Figures 4(a) and 4(b) shows the point estimate of γ_{ym} as well as their confidence intervals for seasonally unadjusted and adjusted cases.

We can see from the figures that the point estimate for γ_{ym} are not so sensitive to the inclusion of data for the periods of the bubble economy, the burst of the bubble and the recent zero interest rate, and thus generally support the long-run monetary neutrality of money in terms of $m2$.



Note: Horizontal axis corresponds to the end of the sample period. The "o" in the figure indicates the estimated value of γ_{my} and the vertical line shows the 95% confidence interval.

Figure 4(a). Long-run neutrality of *m*2 for unadjusted case: Robustness to the data at the end of the sample periods from 1985 to 2006.



Note: Horizontal axis corresponds to the end of the sample period. The "o" in the figure indicates the estimated value of γ_{ym} and the vertical line shows the 95% confidence interval.

Figure 4(b). Long-run neutrality of *m*2 for adjusted case: Robustness to the data at the end of the sample periods from 1985 to 2006.

9. Conclusion

Fisher and Seater's (1993) seminal research on the long-run neutrality of money is adopted to test quarterly seasonally unadjusted Japanese data over the period 1952q2 – 2006q1. The seasonal variations in the real GDP variable have been modeled explicitly. The long-run monetary neutrality hypothesis is supported using M2 as the measure of money supply, that is, changes in M2 have no effects on changes in real output in the long-run. Since M2 is not integrated to a sufficiently high order, there are no permanent stochastic changes to the growth rate, and hence the superneutrality hypothesis using M2 cannot be analyzed for the dataset. However, the long-run neutrality and superneutrality hypotheses are rejected using M1, in that changes in M1 significantly affect changes in real output.

Seasonally adjusted real GDP, M1 and M2 have been used to check the sensitivity of the test results for LRN over the same period. The results are qualitatively same as those obtained from seasonally unadjusted data. Thus, the test results for LRN presented in this paper are robust and the LRN prepositions in Japan indicate the sensitivity of the outcome to the type of money supply.

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