Testing the Long-Run Neutrality Hypothesis Using Intra-year Data

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Abstract Previous tests of the long-run neutrality hypothesis have generally relied on annual time series data. This paper analyses the long-run neutrality of money in Australia using different sources of intra-year data, which permits an examination of the effects of seasonality and the robustness of previous empirical results. A reduced form ARIMA model is used with both quarterly seasonally unadjusted and adjusted Australian real GDP and nominal money supply to test the neutrality hypothesis. Using two measures of money stock, namely M1 and M3, it is shown that the hypothesis is supported using M1 as the measure of money supply, while it is rejected using M3. Recent trends and developments in the money and credit markets in Australia provide a possible explanation of the sensitivity of the outcome to the measure of money stock employed in the analysis.

1. INTRODUCTION

The classical theory of macroeconomics, whereby changes in nominal variables have no effect on real magnitudes in the long run, has attracted considerable research interest for extended periods. Cited as the 'Classical Dichotomy', this is an area in which important issues such as the (possibly) vertical long-run Phillips Curve, the empirical Fisher Effect, and the neutrality of money are constantly debated. In the first example, the inflation rate and unemployment rate are the central variables (as in King and Watson (1992)); in the second, the real interest rate and the inflation rate are emphasized (as in Mishkin and Simon (1995)); while the third example focuses on the nominal money stock and real output (as in Fisher and Seater (1993)).

Various econometric procedures are available for testing the above neutrality hypotheses. The empirical testability of these issues is important for policy formulation and design, such as the effectiveness of monetary policy. In an attempt to evaluate one aspect of the efficacy of monetary policy in Australia, the neutrality hypothesis is the focus of this paper.

The aim of this paper is to test for long-run monetary neutrality using post-war quarterly seasonally unadjusted and adjusted data, to address the issue of seasonality in the data, and to provide an explanation for the outcome of the test. Although the hypothesis to be tested is the long-run neutrality of *money*, the framework can also be applied to the other neutrality hypotheses mentioned above.

The plan of the paper is as follows. In Section 2, the long-run neutrality of money is defined, and outcomes from previous research are highlighted. Section 3 presents the formal econometric framework. A description of the data used in the empirical study is contained in Section 4. Issues of seasonality and unit roots are addressed in Section 5. Section 6 presents the Australian evidence on the long-run neutrality of money, together with a discussion on recent trends in the variables involved. Some concluding remarks are provided in Section 7.

2. LONG-RUN NEUTRALITY OF MONEY

The key feature of the classical model, which underlies the long-run neutrality of money, is perfect foresight. In the long run, when the labour force has had sufficient time to revise errors in its expectations-formation mechanisms, there is an absence of money illusion. Consequently, the aggregate supply curve is vertical. In an expansionary monetary policy regime, such as the numerous Cash and 10-year Bond rate cuts by the Reserve Bank of

Australia in the recessionary period of the early 1990s, a perfectly inelastic supply schedule implies that the general price level will increase, with no change in equilibrium employment. Hence, an increase in nominal money supply has the effect of equi-proportionate increases in the price level, so that real money supply and real output remain unchanged.

A formal definition of the long-run neutrality (LRN) of money is that permanent, exogenous changes to the *level* of money supply have no effect on the level of real output in the long run. Related to the LRN of money is the long-run superneutrality (LRSN) of money, which occurs when permanent, exogenous changes in the *growth rate* of money supply leave the level of real output unaffected.

The LRN hypothesis has been tested widely, using datasets from various countries and time periods, and with different approaches. For example, King and Watson (1992) applied their test for LRN (under different structural assumptions) to US seasonally adjusted quarterly data, and found that LRN was supported by the data, that is, money was found to be neutral with respect to output. More recently, Fisher and Seater (1993) found that LRN was rejected by US annual data for 1869 to 1975. The seminal research by Fisher and Seater has led to numerous related publications. Boschen and Otrok (1994) and Olekalns (1996) adopted Fisher and Seater's approach to analysing the money supply - output relationship in the USA and Australia, respectively, accommodating structural breaks with split samples and dummy variables. It was revealed that the outcomes of the tests are not robust to structural breaks.

To date, there does not seem to have been any analyses of LRN using both seasonally unadjusted and adjusted quarterly data. Apart from providing an alternative set of data, use of intra-year data permits an examination of the effects of seasonality and the robustness of previous empirical results. The Fisher and Seater (1993) approach is adopted and augmented to accommodate seasonality in the present paper. This approach relies on a reduced form time series model, which is appealing because no structural modelling is required.

3. THE ECONOMETRIC FRAMEWORK

Following the approach developed by Fisher and Seater (hereafter FS), it is assumed that the relationship between money supply and real output can be represented by a stationary and invertible bivariate log-linear ARIMA model. The focus variables are m_t and y_t, which are the natural logarithms of nominal money supply and real output, respectively. It is standard to use natural logarithmic transformations of variables because first differences approximate percentage changes, which are valuable for an analysis of superneutrality.

The model is given as follows:

$$a(L)(1-L^{<m>})m_t = b(L)(1-L^{<y>})y_t + u_t$$
$$d(L)(1-L^{<y>})y_t = c(L)(1-L^{<m>})m_t + w_t \quad (1)$$

where L is the lag operator, a(L), b(L), c(L) and d(L) are distributed lag polynomials, and <m> and <y> are the orders of integration of the money stock and output, respectively. The vector $(u_t\ w_t)'$ is assumed to be iid, with the moments given by $(0,\ \Sigma)$. For the distributed lags a(L) and d(L), it is convenient to set the initial values $a_0 = d_0 = 1$, while the parameters b_0 and c_0 are not restricted. Equation (1) can be rewritten as

$$(1 - a_1L - a_2L^2 - ... - a_kL^k)(1 - L^{<m>})m_t =$$
$$(b_0 - b_1L - b_2L^2 - ... - b_kL^k)(1 - L^{<y>})y_t + u_t$$

$$(1 - d_1L - d_2L^2 - ... - d_kL^k)(1 - L^{<y>})y_t =$$
$$(c_0 - c_1L - c_2L^2 - ... - c_kL^k)(1 - L^{<m>})m_t + w_t.$$

This system of equations reflects the interaction between m_t and y_t, with k being the lag length. All variables are difference stationary due to the $(1-L^I)$ filter, where I is given by <m> or <y>.

For the test of LRN and LRSN, the parameters of the second equation in (1) are of interest, in which the stationary values of y over time are explained by stationary values of m over time. Thus, a test for the LRN and LRSN of money can be derived from the normalised coefficients of the endogenous explanatory variables. This leads to the concept of the Long-Run Derivative (LRD), which is a measure of the dynamics of the partial effects of m_t on y_t. With $z_t \equiv (1 - L^j)y_t$ and $x_t \equiv (1 - L^i)m_t$, where i and j equal 0 or 1, the general form of the LRD is $LRD_{z,x} = \lim_{k \to \infty} \frac{\partial z_{t+k} / \partial u_t}{\partial x_{t+k} / \partial u_t}$

where $\lim_{k \to \infty} \partial x_{t+k} / \partial u_t \neq 0$. Apart from the mathematical requirement that this inequality has to hold for the LRD to be defined, the inequality implies that there must be permanent stochastic shocks to the money supply, otherwise LRN and LRSN cannot be considered.

The numerator of the LRD provides a measure of the ceteris paribus effect of a shock to the money supply on real output. This shock can be regarded as an exogenous monetary disturbance. Similarly, the denominator of the LRD provides a measure of the ceteris paribus effect of a shock to the money supply on the money supply variable itself.

Thus, the LRD can be expressed as

$$LRD_{z,x} = \lim_{k \to \infty} \frac{\partial z_{t+k}}{\partial x_{t+k}}$$

which is interpreted as the long-run impact on real output of a monetary disturbance, that is, unanticipated changes in the nominal money stock. The specific value of the LRD depends on <m> and <y>, namely, the orders of integration of m_t and y_t.

Of particular interest in examining the LRN and LRSN of money are the cases <m> = <y> = 1 and <m> = 2, <y> = 1. These cases are important because, when <m> ≥ 1 and <y> ≥ 1, there are permanent changes in both m_t and y_t. The case <y> ≥ 1 is a possibility, and FS show that propositions regarding how permanent changes in the growth rate of money are ultimately reflected in the growth rate of output have LRN rather than LRSN interpretations. Using the impulse response representation of the ARIMA system, FS determined the values of the LRD under the LRN and LRSN hypotheses to derive implications that are empirically testable. Specifically, when <x> - <z> = 0, or when <m> = <y> = 1, $LRD_{z,x}$ = c(1)/d(1). The LRN of money can now be defined in terms of the LRD: money (any nominal variable m) is long-run neutral with respect to output (any real variable y) if $LRD_{y,m} = 0$; similarly, LRSN will prevail if $LRD_{y,\Delta m} = 0$.

The ARIMA approach to LRN is appealing because it involves a time series model and does not require any structural assumptions; specifically, the parameters of the distributed lags c(L) and d(L) need not be estimated. When the error terms u_t and w_t in the ARIMA model are uncorrelated, or when money is exogenous, c(1)/d(1) is the frequency-zero coefficient in a regression of $(1-L^{<y>})y_t$ on $(1-L^{<m>})m_t$. Hence, the term c(1)/d(1) can be approximated by $\lim_{k \to \infty} \beta_k$, where β_k is the slope coefficient in the following regression:

$$\left[\sum_{j=0}^{k}(1-L^{<y>})y_{t-j}\right] = \alpha_k + \beta_k\left[\sum_{j=0}^{k}(1-L^{<m>})m_{t-j}\right] + \varepsilon_{kt}. \quad (2)$$

Following the above specification, $\hat{\beta}_k$ is, in fact, the Bartlett estimator of the frequency-zero regression coefficient, where the Bartlett estimator is the infinite limit of the slope coefficient . If <m> = <y> = 1, which is the case applicable for testing LRN, the estimator is the slope coefficient in

$$(y_t - y_{t-k-1}) = \alpha_k + \beta_k(m_t - m_{t-k-1}) + \varepsilon_{kt}. \quad (3)$$

For the LRSN of money, which is testable if <m> = 2 and <y> = 1, the estimator of c(1)/d(1) is the slope coefficient in

$$(y_t - y_{t-k-1}) = \alpha_k + \beta_k[(1-L)m_t - (1-L)m_{t-k-1}] + \nu_{kt}.$$

4. DATA AND DESCRIPTION

Quarterly seasonally unadjusted Australian real GDP (expenditure based), and nominal money supplies M1 and M3, are used as the real output and nominal money supply series. The period 1975Q2 to 1995Q4 is chosen as the sample so that the effects of the 1973/1974 oil price shock have subsided to a certain extent, thereby avoiding the need to model structural breaks, as in Boschen and Otrok (1994) and Olekalns (1996). King and Watson (1992) used quarterly seasonally adjusted data for their study of LRN. It has been shown that seasonal adjustment procedures such as the Census X-11 can remove cyclical fluctuations in the data, as seasonality and business cycles are typically correlated, and can thereby distort the data (see, for example, Franses (1996)). Consequently, both quarterly seasonally unadjusted and adjusted time series data are used in this paper to examine the effects of seasonality and the robustness of previous empirical results.

Consideration of two measures of money supply, namely M1 and M3, serves as a sensitivity analysis for the potential effects of money on real output. It is shown in Olekalns(1996) that the outcome of the test for LRN is sensitive to the measure of money employed. In addition, the two monetary aggregates have not moved together in the period under consideration, so that using one measure of money supply as a sole representative of m_t would not seem to be warranted.

Boschen and Otrok (1994) omitted the period of the Great Depression from the FS dataset and found that the result was overturned, namely, the LRN hypothesis was supported instead of rejected. They concluded that structural breaks could bias the outcomes of the tests for LRN and LRSN. As the sample used in this paper begins in 1975, there are no obvious structural breaks in the dataset, so such potential biases in the outcome of the test are likely to be avoided.

5. SEASONALITY AND TESTING FOR LONG-RUN NEUTRALITY

Figures 1 and 2 contain the plots of the three variables. It can also be observed in Figure 1 that real GDP exhibits substantial seasonal fluctuations. In intra-year observed data, seasonality has been an issue that has attracted considerable research interest in the modelling of economic time series. Traditionally, seasonality has been explicitly removed using seasonal adjustment procedures. However, many published papers have argued that modelling, instead of removing, seasonality may be beneficial for economic analysis (see, for example, Hylleberg (1992)). This section presents the econometric treatment of seasonality for purposes of testing for the long-run neutrality of money.

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

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

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 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.

When intra-year data are used, the ADF test cannot detect seasonal unit roots or stochastic seasonality. Therefore, when the outcome of an ADF test suggests 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 intra-year data is of paramount importance. The most commonly used test for seasonal unit roots is the HEGY test of Hylleberg *et al.* (1990).

To motivate the HEGY test, consider the seasonal differencing filter $(1-L^4) = (1-L^2)(1+L^2) = (1-L)(1+L)(1-iL)(1+iL)$, where i is a complex number. This implies that the seasonal difference operator assumes four roots of length 1 on the unit circle. The roots are 1, -1, i and - i, where the first is the usual zero frequency (non-seasonal) root, and the other three are seasonal unit roots. The root -1 corresponds to unit roots at 1/2 cycle per quarter or 2 cycles per year (semi-annual unit roots), and the roots i and - i correspond to unit roots at 1/4 cycle per quarter or one cycle per year (annual unit roots).

Given that real GDP contains deterministic seasonality while the monetary variables do not, two versions of the HEGY tests are considered. They are based on the following regressions:

$$a_t = \alpha + \delta t + \pi_1 b_t + \pi_2 c_t + \pi_3 d_t + \pi_4 e_t + \varepsilon_t \quad (5)$$

$$a_t = \alpha + \delta t + \gamma_1 S_{1t} + \gamma_2 S_{2t} + \gamma_3 S_{3t} + \pi_1 b_t + \pi_2 c_t + \pi_3 d_t + \pi_4 e_t + \varepsilon_t \quad (6)$$

where t is a deterministic time trend, S_{1t}, S_{2t}, and S_{3t} are seasonal dummy variables, and

$a_t = (1 - L^4)x_t$

$b_t = (1 + L + L^2 + L^3)x_{t-1}$

$c_t = -(1 - L + L^2 - L^3)x_{t-1}$

$d_t = -(1-L^2)x_{t-2}$

$e_t = -(1-L^2)x_{t-1}$

ε_t is an i.i.d $(0, \sigma_\varepsilon^2)$ error term.

In order to determine the differencing filters that are necessary for the stationarity of the variables, the HEGY hypotheses are formalised as follows:

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

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

3) H_0: $\pi_3 = \pi_4 = 0$,

H_1: $\pi_3 \neq 0$ and / or $\pi_4 \neq 0$.

Standard t-tests are used for the first two hypotheses, while an F-test is used for the third hypothesis. If the first hypothesis is not rejected, there is a non-seasonal or zero frequency unit root in the series, which requires the filter (1 - L) for stationarity. In the second case, 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 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. For the series to have no unit roots at all, the two separate null hypotheses, $\pi_3 = 0$ and $\pi_4 = 0$, must be rejected.

The three series considered in this paper are tested for possible unit roots using the HEGY test, the results of which are presented in Table 1. It is found that all three variables are not seasonally integrated, namely, the quarterly series do not contain unit roots at the semi-annual or the two annual frequencies. The variables do, however, have unit roots at the zero or non-seasonal frequencies, which is consistent with the results of previous tests of LRN.

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 are included in equation (3) to capture the seasonal variations. Equation (3) is modified as follows:

$$(y_t - y_{t-k-1}) = \alpha_k + \gamma_1 S_{1t} + \gamma_2 S_{2t} + \gamma_3 S_{3t} + \beta_k (m_t - m_{t-k-1}) + \varepsilon_{kt.} \quad (7)$$

6. AUSTRALIAN EVIDENCE

Equation (7) is estimated for M1 and M3. Test outcomes for LRN can be observed by examining a plot of the estimates of β_k against the lag length k. The plots are presented in Figures 1 and 2 for both types of money supply considered, with the starting value of k being 2 and the terminal value set at 30. Of primary interest in the estimation of equation (7) is the estimated value of β_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 FS, the 95 percent 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 (n = 83) degrees of freedom is used to construct the confidence intervals.

Using M1, it is observed that the LRN hypothesis is supported by the data. 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. This suggests that M1 does not affect real GDP in the long run. With reference to the LRD, its zero value is warranted empirically. Since money supply M1 is not integrated of order 2, the LRSN of M1 cannot be analysed.

On the other hand, the LRN of money using M3 is rejected by the data. As the lag length increases, there is an upward trend in the plot of the estimates, as well as the lower confidence interval band, suggesting a positive correlation between M3 and real GDP. The LRD is non-zero when M3 is used, indicating the sensitivity of the results to the type of money supply considered. As the LRN hypothesis is rejected using M3 as the measure of money supply, the LRSN hypothesis using M3 is also rejected, that is, the hypothesis that permanent stochastic changes to the growth rate of M3 ultimately leave the level of real GDP unchanged is rejected.

These results are consistent with those of Olekalns (1996), where annual Australian data for the period 1900/01 to 1993/94 are used. The LRN hypothesis is supported using M1 as the money supply, while the use of M3 leads to the LRN hypothesis being rejected for Australia.

The present analysis uses quarterly seasonally unadjusted data for the three variables. As real GDP is highly seasonal, dummy variables are introduced to capture these fluctuations. Failing to accommodate seasonality in real GDP will lead to estimated standard errors of $\hat{\beta}_k$ which include a seasonal component, thereby affecting inferences regarding the long-run neutrality of money.

Due to the differences in the outcomes obtained when M1 and M3 are used as the money supply m_t, recent trends in these aggregates and their relations with real GDP are examined. Referring to the plots of the two money supplies (Figures 2 and 3), in levels and in growth rates (log-differences), it can be observed that the growth in the broader measure of money supply (M3) has been exceptionally slow in the early 1990s. In contrast, growth in the narrow monetary aggregate (M1) has been relatively rapid. The slow growth in M3 and the rapid growth in M1 in Australia in the early 1990s has parallelled international experience, especially the USA and Japan, for the same time period.

Growth in money supply M3 declined sharply over the period 1989 to 1991. This decline can be explained by the decline in private final demand, in particular, falls of large magnitudes in investment spending by both the private and public sectors in the early 1990s. The fall in aggregate investment expenditure and spending had not been experienced since the early 1960s. This led to weak demand for new credit (and consequently M3) needed to finance new investment and capital equipment. Although there have been supply-side arguments for the decline in the growth rate of M3, such as the 'credit crunch' of the late 1980s, the sluggish performance of M3 in Australia in the early 1990s can most satisfactorily be explained by demand-side disturbances.

A different picture emerges with the more narrow definition of money. M1 has grown noticeably over the period in which growth in M3 has declined. The rise in the growth of M1 was expected as a result of the easing of monetary policy by the Reserve Bank of Australia in the midst of the recession. M1 has traditionally been highly sensitive to changes in the interest rate, and the fall in the official cash rate from about 15 percent in mid-1990 to about 12 percent in mid-1991 affected M1 with little delay. Other factors that are likely to have affected the trend in M1 include the shocks to business confidence as a result of the failures of various non-bank financial institutions, such as the Farrow group of Building Societies in Victoria in the early 1990s. The failures of these societies has led to cash being preferred over deposits with non-bank financial intermediaries.

Technical changes in the financial sector during 1991 may also have contributed to the disparity between growths in M1 and M3. During that period, several leading banks developed or expanded cheque-linked savings account facilities, in response to competitive pressure in the retail deposit market. As a result, there were shifts between categories of bank deposits in the official statistics. These shifts left M3 unchanged to a certain extent, as M3 includes *all* bank deposits, while increasing M1. Overall, the declining growth in M3 in the early 1990s can be explained by weak demand and income. As with M1, strong growths were likely to be caused by the numerous interest rate cuts by the Reserve Bank of Australia in the recessionary periods.

The outcome of the tests of neutrality is that neutrality holds with respect to M1, while the hypothesis is rejected with respect to M3. With the growth in real GDP in Australia slowing down in the early 1990s, to the extent that the year ended percentage change was negative in 1990/91 - 1991/92, the strong growth in M1 is likely to have led to the non-rejection of the neutrality hypothesis. The fall in the growth rate of M3 is consistent with weak growth in real GDP over the sample considered, thereby leading to the rejection of the neutrality hypothesis.

In the Appendix, the natural logarithms of seasonally adjusted real GDP, M1 and M3 are used to examine the sensitivity of the test results for LRN to the seasonal adjustment of the time series over the same period. The outcomes of the tests using seasonally adjusted M1 and M3 are qualitatively the same as those obtained using seasonally unadjusted data, namely, that the LRN hypothesis is not rejected using M1 but is rejected using M3. Thus, the test results for LRN presented in this paper are robust to the standard seasonal adjustment transformation.

7. CONCLUDING REMARKS

This paper has considered the long-run neutrality hypothesis. Fisher and Seater's (1993) seminal research on the long-run neutrality of money is adopted to test quarterly seasonally unadjusted Australian data, with the seasonal variations in the real GDP variable being modelled explicitly. The long-run monetary neutrality hypothesis is supported using M1 as the measure of money supply, that is, changes in M1 have no effect on changes in real output. Since M1 is not integrated of a sufficiently high order, there are no permanent stochastic changes to the growth rate, and hence the superneutrality hypothesis using M1 cannot be analysed for the dataset. However, the long-run neutrality and superneutrality hypotheses are rejected using M3, in that changes in M3 significantly affect changes in real output. These results for Australia indicate the sensitivity of the outcome to the type of money supply used. Recent demand-side disturbances and the easing of monetary policy, which affected the two monetary aggregates, are likely causes of the disparity.

ACKNOWLEDGMENTS

The authors wish to thank Philip Hans Franses, Les Oxley, Adrian Pagan, Herman van Dijk and seminar participants at the Econometric Institute, Erasmus University Rotterdam and the University of Western Australia for helpful comments and suggestions. The first author wishes to acknowledge the financial support of an Australian Postgraduate Award and an Individual Research Grant from the Faculties of Economics & Commerce, Education and Law at the University of Western Australia, and the second author acknowledges the financial support of the Australian Research Council. An earlier version of this paper was presented at the Australasian Meeting of the Econometric Society, Perth, Australia, July 1996.

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TABLE 1: HEGY Tests for Seasonal Integration

	RGDP	M1	M3
Lag	0	1	2
H_0: $\pi_1 = 0$	-2.745	-3.102	-2.052^{w}
H_0: $\pi_2 = 0$	-3.096	-3.422	-2.778^{w}
H_0: $\pi_3 = 0$	-6.414	-2.219	-1.422^{w}
H_0: $\pi_4 = 0$	-3.944	-1.798	-3.131^{w}
H_0: $\pi_3 = \pi_4 = 0$	40.333	4.069	5.750^{w}

Note: A time trend and three seasonal dummy variables are included in the HEGY regression for RGDP. A time trend is included in the HEGY regressions for M1 and M3. 'w' indicates White's heteroscedasticity-consistent test statistics.

FIGURE 1: Log of Australian Real GDP

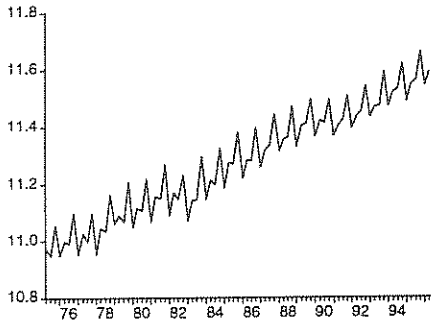
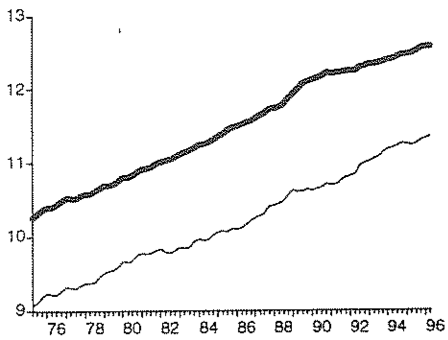


FIGURE 2: Log of Australian M1 and M3



Note: The bold line represents the natural logarithm of nominal M3 and the fine line represents the natural logarithm of nominal M1.

FIGURE 3: Long-Run Neutrality of Money: M1

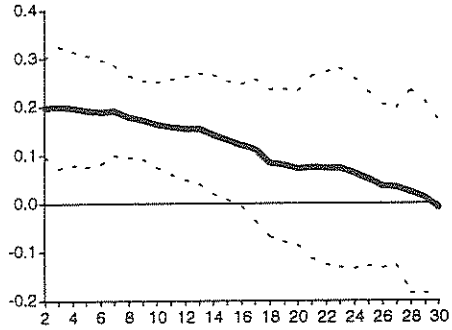
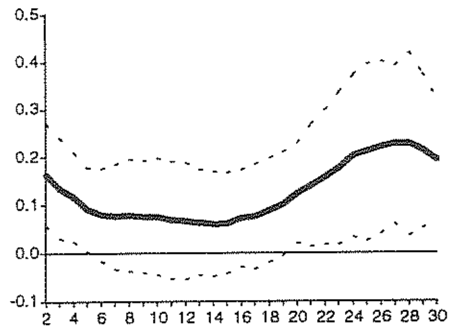
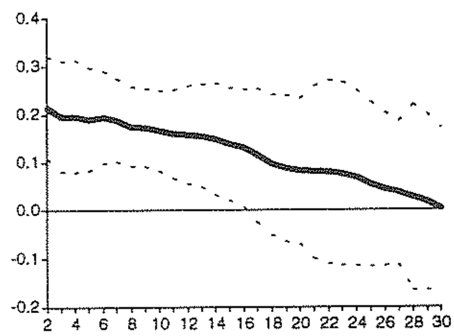


FIGURE 4: Long-Run Neutrality of Money: M3

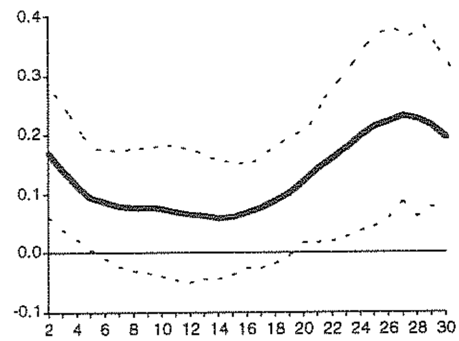


APPENDIX

1) Long-Run Neutrality of Money: M1s



2) Long-Run Neutrality of Money: M3sa



Note: The solid lines are plots of the estimated coefficients over the lag length (k). The dotted lines represent the 95 percent confidence intervals.