VOLKSWIRTSCHAFTLICHE DISKUSSIONSBEITRÄGE

VWL

UNIVERSITÄT - GESAMTHOCHSCHULE - SIEGEN
FACHBEREICH WIRTSCHAFTSWISSENSCHAFTEN

Paper prepared for the Carnegie-Rochester

Conference Series on Public Policy, Fall 1993

"TESTING LONG-RUN NEUTRALITY: Empirical Evidence for G7-Countries with Special Emphasis on Germany"

by

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Discussion Paper No. 46-93, November 1993

ABSTRACT

Modern neo-Keynesian, new classical and real business cycle models typically differ in the degree to which they incorporate certain long-run or even short-run neutrality propositions. Given the importance of these neutrality propositions, it is somewhat surprising how little firm international empirical evidence on their validity is available to date. In part this must be blamed on the fact that until recently a sufficiently general and widely accepted econometric framework for empirically analysing long-run neutrality propositions was not available. King and Watson (1992) and Fisher and Seater (1993) provide such a tool, as well as empirical evidence for the post-war U.S. economy. The present paper analyses to what degree their results are confirmed by tests using data from G7-countries, with special emphasis placed on Germany.

"TESTING LONG-RUN NEUTRALITY: EMPIRICAL EVIDENCE FOR G7-COUNTRIES WITH SPECIAL EMPHASIS ON GERMANY"

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JEL classification:
E31, E43, E52

Keywords:
unit roots, vector autoregressions, long-run neutrality, superneutrality, 'Phillips curve', 'Fisher relation', 'Lucas critique'

Acknowledgements:
Part of the research reported here was carried out whilst on a visiting fellowship at the Brookings Institution, Washington, D.C. I should like to thank the Brookings Institution for its hospitality, and the Fritz Thyssen Stiftung for its financial support in sponsoring this fellowship. This work is also part of a research network on *Macroeconomic Policy and Monetary Integration in Europe* (grant No. SPES-0016-NL(a)) and of a CEPR research programme on *Financial and Monetary Integration in Europe* (grant No. SPES E89300205/RES), financially supported by grants from the Commission of the European Communities. This financial support is also gratefully acknowledged. I should like to thank Robert King for kindly supplying a distribution diskette containing data and programmes, and participants at seminars at the University of Bonn and the Swiss National Bank for useful comments and discussions. The usual disclaimer applies. All programmes and data used in the present paper are available on request.

1. Introduction

Does money matter? This question has been at the heart of much of the theoretical and empirical debate in macroeconomics over recent decades and will presumably keep economists busy for decades to come. Whilst modern rational expectations neo-Keynesian, new classical and real business cycle models typically differ in the mechanisms and degrees to which they incorporate short-run non-neutralities of money, little disagreement is to be found over the theoretical proposition that money is neutral in the long-run. The main controversy about the long-run neutrality hypothesis is empirical in nature, and several empirical approaches towards testing long-run neutrality propositions may be identified.

One set of studies looks at long-run neutrality propositions from a cross-country perspective. Using restricted least-squares regression techniques for data averaged over long-period for a cross-section of countries, the long-run neutrality of money and the 'Fisher relation' are tested in Lothian (1985) for 20 OECD countries, in Loef (1993) for 12 EC countries, and in Duck (1988, 1993) for 33 countries (16 industrialised and 17 developing countries). The cross-country evidence from these studies is largely in favour of the long-run neutrality of money, but whilst Duck (1993) finds evidence in favour of a long-run 'Fisher effect', both Lothian (1985) and Loef (1993) report a less than proportional effect of inflation on nominal interest rates, thereby rejecting the 'Fisher hypothesis'.

Another set of studies, such as Lucas (1980), Mills (1982), Geweke (1982, 1986) and Summers (1983), has attempted to test 'long-run' economic relationships by means of frequency-domain time series techniques. Lucas (1980) and Mills (1982) extract low-frequency signals about the long-run properties of U.S. data on money, prices, output and real interest rates from observable time series and analyse the predictions of the quantity theory in terms of the co-movements between such low-frequency signals. Geweke (1982, 1986) develops a method that allows the measures of linear dependence and feedback in a bivariate vector autoregressive model to be decomposed by frequency, and applies it to test the neutrality and superneutrality of money at both high and low frequencies. Summers (1983) uses band-spectrum regression techniques to obtain low-frequency estimates of the effects of inflation on real interest rates in a 'Fisher-relation'. Summers (1983) rejects the 'Fisher relation' for post-war U.S. data, but the results from the studies by Lucas (1980), Mills (1982), Geweke (1982, 1986) are consistent with both the long-run neutrality and superneutrality of money. Drawing on earlier remarks made by Lucas (1972, 1976) and Sargent (1971), McCallum (1984) criticised these neutrality results derived by frequency-domain methods as being uninformative, due to the problem of observational equivalence: without detailed knowledge about the time series properties of the money supply process, both frequency-domain and reduced form econometric methods are not able to discriminate empirically between long-run non-neutralities and the effects arising from autoregressive money supply processes in models incorporating rational expectations and short-run non-neutralities. McCallum (1984) points out that in this context a valid test of long-run neutrality may only be constructed using cross equation restrictions in a bivariate approach.

In a third class of papers, including, amongst others, Geweke (1986), Stock and Watson (1988), King and Watson (1992) and Fisher and Seater (1993), inference about long-run neutrality propositions is based on explicit tests of coefficient restrictions in bivariate vector autoregressive models. A single equation moving average representation is estimated in Mishkin's (1984) analysis of the 'Fisher effect'. In all of these papers long-run neutrality restrictions typically involve a zero-restriction on the sum of coefficients of the contemporaneous and lagged monetary variables, as opposed to the tests of short-run neutrality, such as those conducted in Sims (1972), King and Plosser (1984), Litterman and Weiss (1985), Bernake (1985), Eichenbaum and Singleton (1986), Boschen and Mills (1988) or Manchester (1989), which impose zero-restrictions on the individual coefficients of the monetary impulses. Long-run neutrality is thus a weaker test, and short-run non-neutralities may well be compatible with the long-run neutrality of money. The evidence reported in Geweke (1986), King and Watson (1992) and Fisher and Seater (1993) suggests that U.S. post-war data are consistent with the long-run neutrality of money, but the long-run superneutrality of money is largely rejected. The existence of a 'Fisher effect' of inflation on interest rates is rejected in King and Watson (1992) for U.S. data, whilst the results reported in Mishkin (1984) are in favour of a 'Fisher effect' for the U.S., U.K. and Canada, but not for data from Germany, France, the Netherlands or Switzerland. Finally, King and Watson (1992) also find little evidence against a long-run vertical Phillips curve in the U.S. post-war data.

The key point about the above most recent contributions to the literature on testing long-run neutrality is that meaningful neutrality tests can only be constructed if both monetary and real variables satisfy certain non-stationarity conditions, which are spelled out in detail in King and Watson (1992) and Fisher and Seater (1993). These studies demonstrate that straightforward neutrality tests, such as imposing the restriction that the coefficients of current and lagged monetary impulses in a regression on real economic variables sum to zero, can only be conducted if the order of integration of both series is at least one and equal for both series. Superneutrality tests, on the other hand, require that the order of integration of the monetary variables is equal to one plus the order of integration of the real economic variables. Fisher and Seater (1993) show that much of the evidence in the older literature on long-run neutrality and superneutrality violates these non-stationarity requirements, and hence has to be disregarded. King and Watson (1992) prove that the above long-run neutrality and superneutrality coefficient restrictions also carry over to richer econometric models in which allowance is made for money to react endogenously to movements in output, instead of being exogenously determined. The authors further show how consistent and efficient estimates on the validity of the long-run neutrality and superneutrality restrictions may be obtained.

The task of the present paper is to apply the King and Watson (1992) approach to post-war data from G7-countries, with special emphasis on the case of Germany. One reason for carrying out such an analysis is the fact that, except for the cross-country studies mentioned above and Mishkin's (1984) study on the 'Fisher-effect', no comparative international evidence on the validity of long-run neutrality propositions exists to date. Second, given the important results of King and Watson (1992) and Fisher and Seater (1993), it is interesting to obtain evidence on the degree to which their results are confirmed by tests using data from a number of different countries and time periods. The robustness of the King and Watson (1992) findings for the U.S. economy is further checked in the present paper by using post-war U.S. data with slightly different data definitions, sample periods, as well as other data sources. Finally, the present paper aims at providing some evidence on the extent to which the results derived by using the King and Watson (1992) framework are subject to the Lucas (1976) critique, according to which changes in monetary policy may have systematically altered the structure of the underlying econometric model and hence inference about long-run neutrality propositions. In this part of the empirical analysis I shall concentrate exclusively on the case of Germany.

The remainder of paper is organised as follows: section 2 presents a brief summary of the econometric approach developed in the King and Watson (1992) paper. Section 3 presents the empirical results of the various neutrality tests for G7-countries. Since these tests critically depend on the order of integration in the data, a variety of unit root tests are presented at the beginning of section 3. Section 4 provides some evidence on the relevance of the Lucas critique for Germany. Section 5 concludes the paper with a summary of the main findings and some suggestions for further research.

2. An Econometric Framework for Testing Long-run Neutrality

A fairly general analytical framework for testing long-run neutrality propositions, that is, the hypothesis that changes in nominal variables have no effect on real variables, is provided in King and Watson (1992) and Fisher and Seater (1993). In order to briefly illustrate their point, consider the long-run neutrality proposition that a permanent change in the money stock (m) has no long-run consequences for the level of output (y). Allowing money to both affect output and at the same time to react endogenously to movements in output, the following bivariate vector autoregressive (VAR) model[1] may be estimated if both money and output are integrated of order one:

$$\Delta m_t = \lambda_{my} \Delta y_t + \sum_{j=1}^{p} \alpha_{my}^{j} \Delta y_{t-j} + \sum_{j=1}^{p} \alpha_{mm}^{j} \Delta m_{t-j} + \varepsilon_t^m, \tag{1}$$

$$\Delta y_t = \lambda_{ym} \Delta m_t + \sum_{j=1}^{p} \alpha_{yy}^{j} \Delta y_{t-j} + \sum_{j=1}^{p} \alpha_{ym}^{j} \Delta m_{t-j} + \varepsilon_t^\eta, \tag{2}$$

where λ_{my} and λ_{ym} represent the contemporaneous effect of output on the money supply and the contemporaneous response of output to changes in the money supply respectively. A more convenient representation of this bivariate VAR system is:

$$\alpha_{mm}(L)\Delta m_t = \alpha_{my}(L)\Delta y_t + \varepsilon_t^m, \tag{3a}$$

$$\alpha_{yy}(L)\Delta y_t = \alpha_{ym}(L)\Delta m_t + \varepsilon_t^\eta, \tag{3b}$$

whereby $\alpha_{mm}(L)=1-\sum_{j=1}^{p}\alpha_{mm}^{j}L^{j}$, $\alpha_{my}(L)=\lambda_{my}+\sum_{j=1}^{p}\alpha_{my}^{j}L^{j}$, $\alpha_{yy}(L)=1-\sum_{j=1}^{p}\alpha_{yy}^{j}L^{j}$ as well as $\alpha_{ym}(L)=\lambda_{ym}+\sum_{j=1}^{p}\alpha_{ym}^{j}L^{j}$ applies. In stacked form this may be re-written as:

$$\alpha(L)X_t = \varepsilon_t, \tag{4}$$

[1] Similar bivariate equation systems have been employed by Sargent (1972) in his exposition of tests for the long-run 'Fisher effect', and by Cooley and LeRoy (1985) and Geweke (1986) in the discussion of tests regarding the long-run neutrality and superneutrality tests of money. However, these studies typically employed the levels rather than the first differences of the relevant series. King and Watson (1992) and Fisher and Seater (1993) show that neutrality tests are uninformative if the variables in question are stationary, and inefficient due to cointegration in the case of nonstationarity.

where $\alpha(L)=\sum_{j=0}^{p}\alpha^j L^j$ with

$$X_t=\begin{bmatrix}\Delta m_t\\ \Delta y_t\end{bmatrix},\ \varepsilon_t=\begin{bmatrix}\varepsilon_t^m\\ \varepsilon_t^\eta\end{bmatrix},\ \alpha_0=\begin{bmatrix}1 & -\lambda_{my}\\ -\lambda_{ym} & 1\end{bmatrix},\ \text{and}\ \alpha_j=-\begin{bmatrix}\alpha_{mm}^j & \alpha_{my}^j\\ \alpha_{ym}^j & \alpha_{yy}^j\end{bmatrix},\ j=1,2,\ldots,p.$$

In this notation the long-run multipliers are $\gamma_{my}=\alpha_{my}(1)/\alpha_{mm}(1)$ and $\gamma_{ym}=\alpha_{ym}(1)/\alpha_{yy}(1)$, whereby γ_{my} measures the long-run response of money to a one unit permanent increase in output, whilst γ_{ym} measures the long-run response of output to a permanent unit increase in money. Long-run neutrality of money implies the restriction $\gamma_{ym}=0$.

As noted by Watson and King (1992), equation (4) is econometrically unidentified and the neutrality restriction is no longer testable when money is endogenous. Thus, even if the hypothesis that ε_t^m and ε_t^η are uncorrelated is maintained, one additional restriction is required in order to identify the linear simultaneous equation model. In the literature various identifying restrictions are to be found. It is common practice in the older literature on long-run neutrality testing to assume that money is exogenous, so that $\gamma_{my}=(\lambda_{my}+\Sigma_{j=1}^{p}\alpha_{my}^j)/(1-\Sigma_{j=1}^{p}\alpha_{mm}^j)=0$, which holds for $\lambda_{my}=\alpha_{my}^1=\alpha_{my}^2=\ldots=\alpha_{my}^p=0$. A less restrictive approach is simply to assume that the model is recursive, so that either $\lambda_{my}=0$ or $\lambda_{ym}=0$.[2] Finally, long-run neutrality $\gamma_{ym}=(\lambda_{ym}+\Sigma_{j=1}^{p}\alpha_{ym}^j)/(1-\Sigma_{j=1}^{p}\alpha_{yy}^j)=0$ may be assumed in order to identify the system and estimate the remaining parameters. However, in principle it is possible to identify the above model by specifying a value of any one of the four parameters λ_{my}, λ_{ym}, γ_{my} or γ_{ym} and find the implied estimates for the other three parameters. This is in fact the approach taken in King and Watson (1992), but rather than focussing on a single identifying restriction, the authors report results for a wide range of identifying restrictions by iterating each of the four key parameters (λ_{my}, λ_{ym}, γ_{my} and γ_{ym}) within a reasonable range and each time obtaining estimates of the remaining three parameters and their standard errors.[3]

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[2] Geweke (1986), Stock and Watson (1988) and Fisher and Seater (1993) present tests for neutrality under the assumption that $\lambda_{ym}=0$, and Geweke (1986) also presents results under the restriction that $\lambda_{my}=0$.

[3] The model is estimated by simultaneous equation methods. A detailed description of the estimation proceedure is contained in the Appendix of the King and Watson (1992) paper.

The present paper follows this principle testing strategy, but for Germany results are reported for iterations in both parameter space and time space. I believe that such recursive estimates of the model are not only potentially more informative in terms of the robustness of inference about key long-run neutrality propositions, they also highlight the degree to which this type of econometric policy evaluation is subject to the Lucas (1972) critique that changes in monetary policy, occurring frequently during the post-war period, may systematically have altered the structure of the econometric model and hence the degree to which the data generating processes are consistent with certain long-run neutrality propositions.

3. Empirical Results for G7-Countries with Special Emphasis on Germany

The testing strategy developed by King and Watson (1992) critically depends on the order of integration of the data. Therefore, before presenting any results of the neutrality tests, it is important to discuss at some length the unit root properties of the data. These tests are conducted for output, various monetary aggregates, inflation rates, unemployment rates as well as nominal and real interest rates in G7-countries.

3.1. Unit Root Properties of the Data

Tables 1a to 1f display three types of unit root test statistics: (i) augmented Dickey-Fuller (ADF) 't-statistics' for both demeaned ($t^u(z)$) and detrended data ($t^\tau(z)$), (ii) Stock's (1991) 95% confidence intervals for the largest unit root (ρ), along with the estimate of the actual root of the series ($\hat{\rho}$), and (iii) three Phillips-Perron 't-statistics' ($Z(t_{\hat{\alpha}})$, $Z(t_{\alpha*})$ and $Z(t_{\tilde{\alpha}})$). The main difference between the ADF and Phillips-Perron tests thereby is that the ADF $t^\tau(z)$ and $t^u(z)$ tests adjust for autocorrelations in the first differences of the data parametrically by optionally including j lags of the differenced data as regressors:

$$\Delta z_t = \mu + (\alpha - 1)z_{t-1} + \gamma_i \sum_{i=1}^{j} \Delta z_{t-i} + u_t, \qquad t^{\mu}(z) \text{ for } H_0\text{: } \alpha=1, \qquad (5a)$$

$$\Delta z_t = \mu + (\alpha - 1)z_{t-1} + \beta t + \gamma_i \sum_{i=1}^{j} \Delta z_{t-i} + u_t, \qquad t^{\tau}(z) \text{ for } H_0\text{: } \alpha=1, \qquad (5b)$$

whilst the three Phillips-Perron tests:

$$\Delta z_t = (\hat{\alpha} - 1)z_{t-1} + \hat{u}_t, \qquad Z(t_{\hat{\alpha}}) \text{ for } H_0\text{: } \hat{\alpha}=1, \qquad (6a)$$

$$\Delta z_t = \mu^* + (\alpha^* - 1)z_{t-1} + u_t^*, \qquad Z(t_{\alpha^*}) \text{ for } H_0\text{: } \alpha^*=1, \qquad (6b)$$

$$\Delta z_t = \tilde{\mu} + (\tilde{\alpha} - 1)z_{t-1} + \tilde{\beta}(t - T/2) + \tilde{u}_t, \qquad Z(t_{\tilde{\alpha}}) \text{ for } H_0\text{: } \tilde{\alpha}=1, \qquad (6c)$$

are based on a non-parametric adjustment for this type of autocorrelation based on the Newey-West (1987) estimator, which obtains a robust variance estimate in the presence of dependent and heterogenously distributed data by prefiltering the residuals u_t from the regression:

$$\Delta z_t = \mu + \beta(t - T/2) + u_t, \qquad (7)$$

(under the restrictions μ=0 and β=0 for (6a), and β=0 for (6b) above) with a triangular lag window with weights for lag i (i=1,...,j) given by $\omega(i,j) = 1 - [i/(j+1)]$. The choice of the relevant test statistics amongst $Z(t_{\hat{\alpha}})$, $Z(t_{\alpha^*})$ and $Z(t_{\tilde{\alpha}})$ is made dependent on the Phillips-Perron 't-tests' $Z(t_{\mu^*})$, $Z(t_{\tilde{\mu}})$ and $Z(t_{\tilde{\beta}})$ for the significance of the deterministic drifts and trends in equations (6b) and (6c) respectively.[4]

Monetary Aggregates

Table 1a shows that the unit root properties of monetary aggregates vary considerably, both within and between G7-countries. All monetary aggregates are found to be stationary in second differences, so only the unit root test statistics for the levels and the first differences of the series are reported. The existence of a unit root

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[4] For further details on these tests see Perron (1988).

in the levels series is typically indicated by both the ADF t-statistics t^{μ} and the Phillips-Perron statistics $Z(t_{\alpha*})$ in case of a significant drift term, and t^{τ} and $Z(t_{\tilde{\alpha}})$ in case of a significant deterministic trend. For German M3, French M1, M2, M3 and M4, Italian M1, as well as Japanese M1 and M2 the Phillips-Perron $Z(t_{\alpha*})$ test rejects the null hypothesis of a unit root, whilst the ADF test statistic t^{μ} does not. In these cases Stock's (1991) intervals indicate that even a root substantially smaller than unity may be consistent with non-stationarity. The fact that the heteroscedasticity consistent estimates $Z(t_{\alpha*})$ indicate mean reversion may thus not be regarded as evidence against I(1) processes, and these money stock series are hence viewed as being integrated at least of order one. For the annualised growth rates (Δm) of these monetary aggregates the second row of tests in Table 1a again produces slightly different results depending on whether the null hypothesis of a unit root is judged on the basis of the ADF t^{τ} or the Phillips-Perron $Z(t_{\tilde{\alpha}})$ test statistics: the ADF t^{τ} statistics rarely rejects the hypothesis of a second unit root at a ten percent level, whilst the Phillips-Perron $Z(t_{\tilde{\alpha}})$ tests frequently reject the hypothesis of a second unit root in favour of an I(1) plus deterministic trend hypothesis at the one percent significance level. Stock's (1992) confidence intervals for discriminating a unit root from one smaller than unity are again very wide, suggesting a large degree of uncertainty about the degree of integration of the data. I therefore conclude that in these cases the money stock data are consistent with both trend stationary in growth rates (I(1) plus trend) as well as an I(2) hypothesis. The decision on the order of integration is documented in the last column of Table 1a. For the monetary aggregate M1 in the U.K., the U.S. and Japan both the ADF t^{μ} and the Phillips-Perron $Z(t_{\alpha*})$ suggest deterministic trends in growth rates. In empirically testing for long-run neutrality Fisher and Seater (1993) do not discriminate between the I(1) plus trend and the I(1) plus drift non-stationarity cases, but King and Watson (1992) note that trend stationrity in growth rates makes the neutrality and superneutrality tests hard to interpret. However, given that Stock's

(1991) intervals for the largest unit root are fairly wide, I have nevertheless decided to carry out the neutrality tests but accept that some caution should be exercised in interpreting these results. In the remaining cases the various test statistics either suggest that the data are generated by an I(1) process with drift (German MCB, M1 and M2, and Canadian M1), or are consistent with both I(2) as well as I(1) plus deterministic drift hypotheses (German M3E, Italian M2, U.K. M1, M2 and M4, U.S. M1 and M2, and Japanese M2+CD's). This suggests that it is justifyable to carry out the neutrality tests for the first group, whilst superneutrality tests should be conducted for the second group of monetary aggregates.

In summarising the findings of Table 1a, it is interesting to note that narrow monetary aggregates, such as M1, tend to be integrated of order one, whilst wider monetary aggregates, such as M2 and M3 are frequently found to be integrated of order one or two. The only monetary aggregates which are found to be definitely difference stationary without trend, as postulated by the King and Watson (1992) approach to testing long-run neutrality, are M1, M2 and the central bank money stock MCB in Germany and M1 in Canada. For these aggregates it makes sense to carry out the neutrality tests. The broad monetary aggregates in France (M2, M3 and M4), Italy (M2), the United Kingdom (M2, M3), the United States (M2, M3) and Japan (M2, M2+CD's), as well as the narrow monetary aggregates in the United Kingdom (M0), France (M1) and Italy (M1), are likely to contain a second unit root. For these monetary aggregates the superneutrality tests are of prime interest, whilst the neutrality tests must be interpreted with caution: as King and Watson (1992) and Fisher and Seater (1993) show the long-run neutrality restriction $\gamma_{ym}=0$ holds by construction when money is I(2) and output is I(1), whilst $\gamma_{ym}\neq0$ must be viewed as an indication that output is integrated of order two.

Output

In Table 1b all test statistics indicate that output in all G7-countries appears to be integrated of order one. Using the Stock-Watson approach, the first differences of the output series exhibit significant deterministic trends in Italy, Canada and Japan, and deterministic drifts in Germany, France, the United Kingdom and the United States. The Phillips-Perron tests $Z(t_{\mu *})$ and $Z(t_{\tilde{\beta}})$, on the other hand, points towards a deterministic drift rather than a trend for Italy and Canada, but nor for Japan. For all countries except Japan the ADF statistics t^{u} and the Phillips-Perron statistics $Z(t_{\alpha *})$ reject the null hypothesis of a unit root in output growth rates at least at the five percent level (for France at the 10% level), suggesting that output in these countries follows a random walk with a deterministic drift. Japanese output growth, on the other hand appears to be trendstationary. As discussed above, the deterministic trend in the growth rates of Japanese output make it difficult to interpret either the neutrality or the superneutrality tests for this country.

Inflation

Using the Stock-Watson approach no evidence of significant deterministic trends in the levels or first differences of G7 inflation rates is detected. The first differences of inflation rates are found to be stationary at least at the five percent level according to all test statistics. Based on the ADF t-statistics t^{u} and Stock's confidence intervals for the largest unit root the inflation series must be viewed as being integrated of order one. A slightly different picture is suggested by the Phillips-Perron unit root tests. Here the Phillips-Perron $Z(t_{\alpha *})$ statistics, which is the relevant one for testing the existence of a unit root in the presence of a deterministic drift, rejects the unit-root hypothesis for the level of inflation rates in all G7-countries[5] except Italy, pointing towards a considerable degree of mean reversion in particular in the two low inflation

[5] The same is true if these Phillips-Perron unit root tests are caried out for the inflation series used in King and Watson (1992).

countries, Germany and Japan. Since Stock's confidence intervals for the largest unit root are fairly wide in all three cases, I conclude that inflation rates in Italy are an I(1) process and that the inflation data for Germany, France, the United Kingdom, the United States, Canada and Japan are consistent with both mean-reverting I(0) as well as I(1) processes. The results from Table 1c thus suggest that it is reasonable to carry out neutrality tests for inflation conditional on integrated processes, but that except for Italy the results should be interpreted with caution.

Unemployment

Of all the time series considered in this paper the unemployment data exhibit the most homogeneous unit root pattern across G7-countries. Table 1d indicates that according to all test statistics unemployment rates in G7-countries appear to be integrated of order one. Furthermore, the first differences of unemployment rates are found to exhibit no significant drifts or trends and the hypothesis of a second order unit root is rejected by both the ADF t^{μ} and the Phillips-Perron $Z(t_{\hat{\alpha}})$ statistics at least at the five percent level for all countries except France, where it is rejected at the ten percent level only. I conclude that unemployment rates in G7-countries follow random walks without drifts. This 'hysteresis hypothesis' is consistent with previous findings in the literature,[6] and strongly encourages the application of the King and Watson (1992) approach to testing the hypothesis of a long-run vertical Phillips, bearing in mind the reservations made in the previous section about the inflation processes.

Interest Rates

As for the unemployment series, the results of the unit root tests for nominal three-month interest rates in Table 1e are relatively homogeneous across countries. Except for the United Kingdom, short-term nominal interest rates in G7-countries appear to

[6] See Blanchard and Summers (1986) and the references given there.

be integrated of order one, as indicated by both the ADF t^{μ} test and the Phillips-Perron $Z(t_{\hat{\alpha}})$ test. The hypothesis of a second order unit root is rejected by all test statistics at least at the five percent level. The first differences of nominal interest rates in most G7-countries display neither deterministic trends nor drifts according to the Phillips-Perron tests $Z(t_{\mu*})$ and $Z(t_{\tilde{\beta}})$. However, for the United Kingdom both the ADF t^{τ} test as well as the Stock-Watson test for a deterministic trend are significant at the five percent level, pointing towards trend stationarity. I conclude, with some reservations about the UK, that nominal interest rates in G7-countries follow random walks without drifts, and hence possess the nonstationarity necessary for testing the 'Fisher hypothesis' of a long-run unit effect of inflation on nominal interest rates.

Another way of looking at the conditions necessary for carrying out the neutrality tests proposed by King and Watson (1992) is that nominal interest rates and inflation must not be cointegrated. This becomes obvious when considering that the neutrality hypothesis implied by the 'Fisher relation' is that inflation has a zero long-run effect on real interest rates. For the neutrality tests to apply, real interest rates must be integrated of order one if inflation follows an I(1) process. In this case nominal interest rates (r_t) and inflation (π_t) must not be cointegrated, since otherwise their linear combination (r_t-π_t) may be stationary. The unit root tests reported in the lower part of Table 1e suggest that except for Germany the null hypothesis of a unit root in real interest rates ($\phi_t \equiv r_t - \pi_t$)[7] cannot be rejected at the 95% level when judgement is based exclusively on the ADF 't-statistics' t^{τ} and t^{μ}. However, Stock's (1991) confidence intervals are very wide in each case, suggesting a large degree of uncertainty about the unit root properties of real interest rates. This is also obvious when considering the Phillips-Perron unit root tests, which indicate that real interest rates are trendstationary I(0) processes for France and Italy and mean-reverting I(0)

[7] King and Watson (1992) carry out their unit root tests for real interest rates for the series $\phi_t \equiv r_t - \pi_{t+1}$ and do not reject the hypotesis of a unit root in real rates. However, if $\phi_t \equiv r_t - \pi_t$ is tested with their data set, the unit root hypothesis is rejected at least at the 95% level.

processes for the U.K., the U.S., Canada and Japan. I conclude that German real interest rates are likely to be stationary, whilst real interest rates in the remaining G7-countries are consistent with both I(0) and I(1) hypotheses. As in the King and Watson (1992) study, the tests of the 'Fisher relation' in the present paper may thus be subject to misspecification arising from potential cointegration between nominal interest rates r_t and inflation π_t. Due to the large degree of uncertainty concerning the order of integration of the data, I nevertheless chose to report these tests below, again keeping in mind that caution should be exercised in interpreting the results.

3.2. Empirical Evidence on the Long term Neutrality Propositions for G7-countries with Special Emphasis on Germany

3.2.1. The Neutrality of Money

The present paper follows the approach of King and Watson (1992) in reporting empirical results for a wide range of identifying parameter restrictions. Table 2 summarises the results of the neutrality tests for the various monetary aggregates in G7-countries.

Germany

For Germany, MCB, M1, M2 and real output were judged in Tables 1a and 1b as being driven by random walks with deterministic drifts. As real output and these monetary aggregates are integrated of order one, they possess the nonstationarity necessary for applying the neutrality tests.

Table 2 shows that the variances of the VAR forecast errors of German output and money (columns 3 and 4) are of similar magnitude and slightly positively related in the short-run, as indicated by the correlations between the VAR forecast errors ranging between 0.07 and 0.19 (in column 5). The long-run relation between German output and money, as measured by the stochastic trend innovations, is relatively high, ranging between 0.4 for M1 and 0.73 for the central bank money stock (MCB), the monetary aggregate targeted by the German Bundesbank between 1975 and 1987.

Given these relatively high long-run correlations, the ranges for the various identifying restrictions on λ_{my}, λ_{ym} and γ_{my} consistent with long-run neutrality at the 95% confidence level are found to be relatively narrow for MCB and M1. For example, neutrality cannot be rejected at the 5% level for any value of λ_{my} greater than 0.2 for M1, that is, for modestly accommodative monetary policy. However, the common identifying assumption of contemporaneous exogeneity (λ_{my}=0) would lead to a rejection of the neutrality hypothesis for both MCB and M1, but not for M2, M3 and M3E (see column 9). Table 2 also provides ranges for alternative identifying restrictions on the short-run impact of money on output, λ_{ym} (in column 10), and the long-run impact of output on money, γ_{my} (in column 11). [8]

A second set of evidence considered in Table 2 concerns the estimates of the behavioural parameters and their associated standard errors under the neutrality hypothesis. Imposing γ_{ym}=0 as a long-run condition, the resulting point estimates of γ_{my} are 0.2 for M2, 0.65 for M1 and 0.71 for MCB. Whilst the estimate of γ_{my} for M2 is not significantly different from zero at reasonable significance levels, the significant point estimates for M1 of 0.65 and for MCB of 0.71 suggest that monetary policy in Germany would need to be fairly accommodative in the short-run for long-run monetary neutrality to hold. Also, the estimates of the short-run impact of money on output, γ_{ym} , is found to significantly negative for MCB (at the 10% level) and M1 (at the 5% level), indicating the short-run non-neutrality of these monetary aggregates even if long-run neutrality is imposed upon the estimates.[9] This finding is also visible

[8] Note that both M3 and M3E were judged as being I(2) processes in Table 1a, so long-run neutrality should hold by construction given that output is found to be an I(1) process in Table 1b.

[9] For a discussion of the short-run non-neutralities of various German monetary aggregates in the context of Granger (1963) causality tests see Weissenberger (1984), von Hagen, (1984), and Hansen (1989).

from the graphs of the coefficients trajectories and standard error bands displayed in Figure 1 for German M1 (on the left hand side) and M2 (on the right hand side). I conclude from these graphs that for Germany the evidence about the broad monetary aggregates, such as M2, appears to be consistent with the hypothesis of the long-run neutrality of money, whilst for the narrower monetary aggregates, such as M1, some indication of both short-term and long-run non-neutrality exist.

G7-countries

The above pattern of the results for Germany repeats itself in Table 2 for the other G7-countries. For example, in France, Italy and Canada the narrow monetary aggregates, such as M1, are consistent with the hypothesis of long-run neutrality only for certain narrow ranges of the identifying restrictions on the coefficients λ_{my}, λ_{ym} and γ_{my}, whilst for the wider monetary aggregates, such as M2, M3 or M4, a much wider range of identifying restrictions is consistent with the long-run neutrality hypothesis. The narrow monetary aggregates M1 are thereby not necessarily non-neutral in the long-run: under the common identifying assumption of the long-run exogeneity of money (γ_{my}=0), M1 in France and Italy does not violate the long-run neutrality restriction (γ_{ym}=0) at the five percent significance level. Finally note that, as in the case of Germany, monetary policy is required to be modestly accommodative for M2 in the United States, M1 in Canada and M1, M2 and M2+CD's in Japan in order for long-run neutrality to hold. Except for the United States, the common identifying restriction of the long-run exogeneity of money (γ_{my}=0) would lead to a rejection of the long-run neutrality proposition in Table 2. Similar evidence is interpreted in King and Watson (1992) as empirical support for Goodfriend's (1987) conjecture that these central banks responded to changes in output in order to implement interest rate smoothing policies.

To summarise, the results reported in Table 2 suggest that a wide range of plausible identifying restrictions does not lead to a rejection of the neutrality hypothesis for various monetary aggregates in G7-countries. This qualitative finding is in line with the evidence provided in King and Watson (1992) and Fisher and Seater (1993). Similar results are also reported in Lucas (1980), Mills (1982), and in the cross-country studies of Lothian (1985) and Loef (1993). Moreover, the findings reported here clearly indicate the robustness of the neutrality proposition, both for a large number of monetary aggregates and countries. As in the King and Watson (1992) study, the present paper obtains plausible values for the behavioural parameters when long-run neutrality is imposed. To illustrate this fact, the estimated confidence ellipses for λ_{my} and λ_{ym} under the null hypothesis of the long-run neutrality are displayed in Figure 2a for M1 and in Figure 2b for the broader monetary aggregate (M2 for France, the United Sates, Canada and Japan, and M3 for Germany, Italy and the United Kingdom).

The similarity of the results across G7-countries is striking in several aspects. First, the confidence ellipses for the point estimates are all concentrated in the second and fourth quadrant. This implies that under the long-run neutrality restriction (γ_{ym}=0) money is found to react anti-cyclically to output in the short run (λ_{ym}<0) when the short-run impact of money on output is positive (λ_{ym}>0), and pro-cyclically (λ_{my}>0) when the short-run effect of money on output is negative (γ_{ym}<0). Hence, both quadrants are consistent with the attempts of central banks to stabilise the economy in the short-run, whilst monetary policy is neutral in the long-run. The short-run neutrality (λ_{my}=0) of predetermined money (λ_{ym}=0) is thereby not ruled out, since the origin lies within the confidence ellipses for both narrow (M1) and broad monetary aggregates (M2, M3) for all G7-countries. Second, the admissible range of behavioural parameters for which long-run neutrality cannot be rejected is considerably narrower for narrow money M1 than it is for the broader monetary aggregates M2 or M3.[10]

3.2.2. The Superneutrality of Money

The results for the superneutrality tests for various monetary aggregates in G7-countries obtained from bivariate vector autoregressions with changes in the money stock (Δm_t) in equations (1) and (2) replaced by changes in the growth rates ($\Delta^2 m_t$).

Germany

Table 3 and Figure 3 show that it is fairly easy to find evidence against the superneutrality of money in the German data. A rejection of the superneutrality hypothesis for M1 and MCB, which in Table 2 were found to exhibit some signs of non-neutrality, is little surprising. But even M2 and M3, for which long-run neutrality could not be rejected at the 5% level, are found not to be superneutral for a wide range of identifying parameter restrictions. For example, when M2 growth rates are predetermined (exogenous in the long-run), that is $\lambda_{\Delta my}=0$ ($\gamma_{\Delta my}=0$) holds, the estimated long-run effect of money growth on output is $\gamma_{y\Delta m}=1.10$ ($\gamma_{\Delta my}=1.15$), and superneutrality is rejected at least at the five percent level in both cases.

G7-countries

The superneutrality of money is also rejected for a wide range of identifying restrictions for the majority of monetary aggregates in the remaining G7-countries. The only major exception from this is France, for which the relatively short sample of

[10] It would be interesting in this context to differentiate between inside and outside money along the lines suggested by King and Plosser (1984) and to check the long-run and short-run neutrality of both components of the wider monetary aggregates by using the King and Watson (1992) techniques. Due to space limitations the present paper abstracts from these issues.

only 12 years may not suffice to convey the relevant long-run information about the relationship between money growth rates and real output.

3.2.3. The Long-Run Vertical Phillips Curve

The results for the Phillips curve tests are obtained from bivariate vector autoregressions with changes in inflation rates ($\Delta\pi_t$) replacing money growth rates (Δm_t) and changes in unemployment rates (Δu_t) replacing real output growth rates (Δy_t) in equations (1) and (2). These specification of the bivariate VAR, which underlies the tests for $\gamma_{u\pi}=0$ is reversed for the test $\gamma_{\pi u}=0$ below.

Germany

In Table 4 the estimated short-run correlation between unemployment and inflation is slightly positive (0.03), but the long-run correlation of -0.59 is consistent with a 'statistical' Phillips curve relation. As stressed by King and Watson (1992), this correlation may arise from a causal long-run link running from inflation to unemployment ($\gamma_{u\pi}\neq0$), or conversely, from unemployment to inflation ($\gamma_{\pi u}\neq0$). The present paper finds little evidence against the hypothesis $\gamma_{u\pi}=0$. This is illustrated for the German case in Figure 4, where except for very extreme values of $\lambda_{\pi u}$ (>6.37) and $\lambda_{u\pi}$ (<-0.06) the hypothesis of a long-run vertical Phillips curve ($\gamma_{u\pi}=0$) cannot be rejected at the five percent level. In this case $\gamma_{\pi u}<0$ holds over the entire range of $\gamma_{u\pi}$ estimates. In the case with reverse causation, displayed in Figure 5, it is easy to find evidence against the hypothesis $\gamma_{\pi u}=0$ for a wide range of identifying parameter restrictions on $\lambda_{\pi u}$ (<3.69), $\lambda_{u\pi}$ (>-0.03) and $\gamma_{\pi u}$ (>-0.04). Note that it is thereby sufficient to assume $\gamma_{u\pi}\geq0$, which includes the case of the long-run neutrality of inflation, to ensure that the long-run exogeneity of inflation, $\gamma_{\pi u}=0$, is inconsistent with the German post-war data.

G7-countries

The above findings for Germany can be generalised for the G7-countries. Whilst the short-run correlations between output and inflation in column 5 of Table 4 are found to be either positive or negative, the long-run correlations in column 8 are all negative and range between -0.4 for Italy to -0.71 for Canada. For all countries except Italy these 'statistical' Phillips curves appear to arise due to a long-run causal link from inflation to unemployment. This can be seen from the admissible range of $\lambda_{\pi u}$ in column 9 of Table 4: in carrying out these tests $\lambda_{\pi u}$ has been iterated in the interval between -3.5 and 7, and the hypothesis of a long-run vertical Phillips curve cannot be rejected at the 5% significance level for France, the United Kingdom and Japan over the entire $\lambda_{\pi u}$ range, whilst for the United States and Germany it is only rejected for extreme values of $\lambda_{\pi u}$ (>5.95 and >6.37 respectively). The reverse hypothesis $\gamma_{u\pi}$=0 is, however, rejected for these countries for a wide range of identifying parameter restrictions on $\lambda_{\pi u}$. On the other hand, for Italy the hypothesis $\gamma_{u\pi}$=0 is rejected at the 5% significance level for a fairly wide range of $\lambda_{\pi u}$ values (>1.33), whilst the hypothesis $\gamma_{\pi u}$=0 is only rejected for extreme values of $\lambda_{\pi u}$ (<-2.08). I suspect that the only recently abolished Italian automatic wage indexation scheme (the 'scala mobile'), which had no corresponding equivalent in the other G7-countries, may explain this difference in the long-run interaction between inflation and unemployment between Italy and the remaining G7-countries.

3.2.4. The 'Fisher Hypothesis of Inflation and Interest Rates

The neutrality proposition here is that permanent movements in inflation rates have no long-run effect on real interest rates. Thus, as specified in King and Watson (1992) and Fisher and Seater (1993), the 'Fisher-hypothesis' implies that inflation rates (π_t) and nominal interest rates (r_t) will move one-to-one in the long-run, or $\gamma_{r\pi}$=1. To test the existence of a long-run 'Fisher-effect', the bivariate vector autoregressions (1) and (2) are estimated with changes in inflation rates ($\Delta\pi_t$) replacing money growth rates (Δm_t) and changes in nominal interest rates (Δr_t) replacing real output growth rates (Δy_t).

Germany

In Table 5 the estimated short-run correlation between nominal interest rates and inflation is slightly positive (0.16), and the long-run correlation of 0.76 is substantially smaller than one. Figure 6 also raises some questions about the validity of the 'Fisher-effect', since significant deviations from a long-run coefficient $\gamma_{r\pi}$ of unity can be observed over a relatively wide range of parameter values. The interesting point about Figure 6 is that in order to obtain a long-run one-to-one effect of inflation on nominal interest rates ($\gamma_{r\pi}$=1), the contemporaneous effect of inflation on interest rates ($\lambda_{r\pi}$) has to fall within the range of -0.05 to 0.5, with a point estimate of $\lambda_{r\pi}$=0.18 under the neutrality hypothesis. Note that this estimate is close to the correlation suggested by the VAR estimate. Furthermore, for the 'Fisher-effect' not to be rejected at the 5% level, the short-run effect of interest rates on inflation ($\lambda_{\pi r}$) must fall within the range between -2.4 and 0.6, whilst no specific restrictions on the long-run effect $\gamma_{\pi r}$ are implied. Thus, both predetermined and long-run exogenous inflation would be consistent with the existence of a long-run 'Fisher-effect' for the German data. I therefore conclude that whilst the 'Fisher-effect' does not appear to hold for a wide range of identifying restrictions, the restrictions implied by imposing $\gamma_{r\pi}$=1 include the benchmark cases of short-run and long-run exogeneity of inflation, as well as the conventional finding of only partial short-run adjustment of nominal interest rates to inflation. I view this evidence as compatible with a long-run 'Fisher-effect' for German post-war data. Note that this results contradicts the evidence reported by Mishkin (1984), who rejects the 'Fisher-effect' for Germany.

G7-countries

Sargent (1973), Summers (1983), King and Watson (1992) provide evidence which on the whole rejects the Fisherian theory of inflation and interest rates for post-war U.S. data, whilst Mishkin (1984) views post-war U.S. data as being consistent with a long-run 'Fisher-effect'. The evidence reported in Table 5 for U.S. post-war data is qualitatively very similar to that of King and Watson (1992): for the 'Fisher-effect' to hold the contemporaneous effect of inflation on interest rates ($\lambda_{r\pi}$) has to be larger than 0.4 with a point estimate of $\lambda_{r\pi}$=0.69, whilst the short-run (long-run) effect of interest rates on inflation has to be less than -1.2 (less than -0.3). Both predetermined and long-run exogenous inflation processes are ruled out by these identifying restrictions. I conclude in keeping with King and Watson (1992) that the post-war U.S. data largely reject the Fisherian theory of interest rates. Even stronger evidence against the validity of the 'Fisher-hypothesis' can be found in Table 5 for the United Kingdom, where across almost the entire range of identifying restrictions on $\lambda_{r\pi}$ the estimates of the long-run effect of inflation on interest rates ($\gamma_{r\pi}$) are found to be both significantly larger than zero and smaller than unity. For the remaining G7-countries the evidence in Table 5 is much more supportive of a long-run 'Fisher-effect'. In most cases only modest contemporaneous effects of inflation on interest rates are required (with $\lambda_{r\pi}$ ranging from 0.1 for Japan to 0.25 for Italy) in order to not be able to reject the existence of a long-run 'Fisher-effect', $\gamma_{r\pi}$=1, at the 5% significance level. Also, imposing a 'Fisher-effect' on the data ($\gamma_{r\pi}$=1) requires that in many countries both the short-run and long-run effect of interest rates on inflation, $\lambda_{\pi r}$ and $\gamma_{\pi r}$, have to be only modestly negative. Long-run exogeneity of inflation ($\gamma_{\pi r}$=0) is thereby not being ruled out by the data for France, Canada, and Japan. I view this as evidence in favour of the 'Fisher hypothesis' for these countries. Mishkin (1984), on the contrary, rejects the 'Fisher-effect' for France but not for the United Kingdom and Canada.

To summarise, the present paper supplements the studies of Fisher and Seater (1993) and King and Watson (1992), and provides similar evidence for the G7-countries. The major results of King and Watson (1992) are thereby confirmed in tests using data from G7-countries, even though some minor modifications occur. It is found that for most countries and for a wide number of different monetary aggregates the data are consistent with the long-run neutrality of money for a wide range of identifying restrictions, whilst the long-run superneutrality of money is largely rejected by the data. King and Watson's (1992) finding of a long-run vertical Phillips curve is also supported by the evidence from G7-countries, even though for Italy the data are more consistent with reversed causation and short-run non-neutralities running from unemployment to inflation, rather than from inflation to unemployment. Finally, whilst in keeping with King and Watson (1992) the 'Fisher-relation' is rejected for U.S. postwar data in the present paper,[11] some evidence for the existence of a 'Fisher-effect' of inflation on nominal interest rates is found for the remaining G7-countries, with the exception of the United Kingdom.

3.3. Long-run Neutrality and the 'Lucas Critique': The Case of Germany

According to the famous 'Lucas critique', the structure of econometric models is in general not invariant to changes in policy objectives, operating procedures or policy constraints over time. The structural parameters of these models are therefore not policy invariant, meaning that they will change whenever policy is changed. As a result, reduced form econometric models used for quantitative policy evaluation tend to exhibit structural breaks if policy changes are of the once-and-for-all type, or, more generally, will vary as policies evolve over time. Such policy induced structural

[11] For the U.S. post-war data recursive estimates like the ones reported below for Germany suggest that the 'Fisher relation' broke down in the early 1970's.

change of econometric models has hitherto received little attention in studies concerned with testing long-run neutrality propositions. I view this as a major short-coming of this literature, and in the following section an attempt is made to gain some insight into the pattern of time-variation with which certain neutrality propositions evolve over time. In studying these issues I shall concentrate on the case of Germany.

The motivation for this analysis comes from the tremendous changes in the policy environment under which German monetary policy in particular had to operate over the past three decades. Initially the Bundesbank's monetary policy was constrained by the exchange rate target vis-à-vis the U.S. Dollar under the Bretton Woods system. During the late 1960's this system became increasingly flexible and finally collapsed in August 1971. An interim period followed, during which the newly-established European currency snake system was tied to the U.S. Dollar by the Smithsonian agreement (snake in the 'tunnel'), until group-floating started in March 1973. In late 1974, the German Bundesbank was the first central bank to announce a formal monetary target in terms of the growth rate of a monetary aggregate, the central bank money stock (MCB), for a period as long as a year. Between 1975 and 1978 the Bundesbank announced its money growth target as a fix-point target, and from 1979 onwards switched to announcing a target range for MCB growth.[12] Also in 1979, the European Monetary System (EMS) came into operation.[13] Due to massive target overshooting during 1986 and 1987 the Bundesbank switched to announcing target ranges for M3 growth in 1988. Finally, Bundesbank policy had to adapt to the fall of the Berlin wall in November 1989 and German economic and monetary unification in July 1990.

All the afore-mentioned changes in the monetary policy environment are not unlikely to have had some effects, at least temporary ones, on the operating

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[12] For a review of the experience of Germany and other G7 countries with monetary target announcements, see Weber (1990).

[13] It is frequently argued that the rules of the EMS did not place a constraint on the Bundesbank, which is typically believed to have been the anchor of the exchange rate mechanism of the EMS. For a review of this literature and an alternative view of the EMS see Weber (1991).

characteristics and transmission of German monetary policy on output. Furthermore, as stressed in the real business cycle literature, real economic shocks, such as the oil price hikes of 1973 and 1979 and the German unification shock of 1989/90, are likely to have had an effect on the degree to which monetary policy accommodates (or stabilises) shock-induced changes in output. In any case, the bivariate money-output VAR system is likely to display some degree of time-variability at certain policy relevant points in time. In the context of the present paper, I am interested only in those events which influence the long-run neutrality characteristics of the German money-output system.

3.3.1. The Neutrality of Money and the 'Lucas-Critique'

Figure 7 displays the estimates for the long-run neutrality proposition of German M1 and output, using recursive estimation methods. The estimates are obtained by holding the initial period (1962:1) fixed and successively extending the sample end from a minimum sample length (1966:4) to the full sample (1992:4). I extended the sample to 1992:4, instead of 1990:4 as in Table 2, in order to test for the effects of German unification on the estimates. The results of the recursive estimates are displayed on an annual grid, with the sample ending in the last quarter of the corresponding year, so the third to the last line in the 3-dimensional graphs corresponds to the estimates reported for the sample period in Table 2 (1962:1-1990:4). Rather than reporting coefficient estimates and standard error bands for the long-run effect of money on output (λ_{ym}), I have chosen to depict the 't-statistics' of the γ_{ym}-coefficients implied by various identifying restrictions on the coefficients λ_{my} (Figure 7a), λ_{ym} (Figure 7b) and γ_{my} (Figure 7c). I also report recursive estimates of the confidence ellipses for the results imposing long-run neutrality (Figure 7d).

Several findings are worth mentioning. First, for any choice of the sample period up to the year 1980 the estimates of the long-run effect of money on output never significantly violate the long-run neutrality coefficient restriction (γ_{ym}=0) for the entire range of λ_{my}, λ_{ym} and γ_{my} coefficients considered. Indications of long-run non-neutrality emerge only in the early 1980's:[14] this effect is most visual in the 3-dimensional plane of Figure 7a, where for each coefficient λ_{my} the corresponding estimate of γ_{ym} is considerable higher in the early 1980's as compared to the late 1970's. For values of λ_{my} smaller than 0.2 significant violations of the long-run neutrality condition γ_{ym}=0 are found in the post-1980's sample, as indicated by the significant t-statistics (τ>2 implies γ_{ym}>0) in the 2-dimensional smaller insets on the right hand side of Figure 7a. Second, in Figure 7a and 7b German unification is found to have only a minor effect on the range of short-run and long-run policy response parameters (λ_{my} and γ_{my}) for which long-run neutrality (γ_{ym}=0) is rejected at the 5% significance level. Figure 7c, however, indicates a strong effect of German unification on the admissible range of coefficients for which the short-run non-neutrality of money (λ_{ym}≠0) is compatible with long-run neutrality (γ_{ym}=0). Third, Figure 7d points out that if long-run neutrality of M1 is imposed on the estimates, then the admissible strength of a response of German M1 to changes in output has increased over the years, whereas the range within which short-run non-neutrality is compatible with long-run neutrality has declined. The point estimates of λ_{ym} thereby suggest that the degree of accommodation has increased permanently after the second round of oil price shocks in 1979/80 and only transitorily after the unification shock in 1990/91.

3.3.2. The Superneutrality of Money and the 'Lucas-Critique'

Figure 8 displays the estimates for the long-run superneutrality proposition of German M1 and output using recursive estimation methods. Again, several points are worth noting. First, compared to the neutrality proposition the superneutrality

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[14] Similar findings about the sensitivity of neutrality tests with respect to the inclusion of data from the 1980's are reported for the United States in Litterman and Weiss (1985) and Eichenbaum and Singleton (1986).

proposition breaks down earlier in time: only for a choice of the sample period up to the year 1974 do the estimates of the long-run effect of money growth on output never significantly violate the long-run superneutrality restriction ($\gamma_{y\Delta m}$=0) for the entire range of $\lambda_{\Delta my}$, $\lambda_{y\Delta m}$ and $\gamma_{\Delta my}$ coefficients. As indicated by Figure 8a, the German data for the post-1974 period are consistent with long-run superneutrality only for values of $\lambda_{\Delta my}$ smaller than 0.5, and predetermined money ($\lambda_{\Delta my}$=0) is compatible with superneutrality. In Figure 8c long-run exogenous money ($\gamma_{\Delta my}$=0) is also compatible with long-run superneutrality ($\gamma_{y\Delta m}$=0) for the entire sample period. Second, as in the case of the neutrality proposition above, German unification is found in Figures 8a and 8c to only have a minor effect on the range of short-run and long-run policy response parameters ($\lambda_{\Delta my}$ and $\gamma_{\Delta my}$) for which long-run superneutrality ($\gamma_{y\Delta m}$=0) has to be rejected at the 5% significance level. However, Figure 8b and 8d indicate that German unification has primarily reduced the admissible range within which short-run deviations from superneutrality ($\lambda_{y\Delta m}\neq 0$) are compatible with long-run superneutrality ($\gamma_{y\Delta m}$=0).

In summarising the above results it must be said that inference about both the long-run neutrality and superneutrality of money is not invariant with respect to time: for Germany significant deviations from long-run superneutrality emerged in the early 1970's, whilst the long-run neutrality of money appears to be violated only after the early 1980's and admittedly only for extremely accommodative monetary policy. Rather than changing policy reactions, it appears to have been changes in the short-term effectiveness of monetary policy in the aftermath of major real economic shocks which have caused deviations from long-run neutrality and superneutrality. For the long-run vertical Phillips curve and the 'Fisher-effect' of inflation on interest rates similar variations of the estimates over time were not detected.

4. Conclusions and Suggestions for Further Research

In this paper I have examined the degree to which long-run neutrality propositions are consistent with quarterly post-war data from G7-countries. Using the bivariate approach to testing long-run neutrality propositions developed in King and Watson (1992), it has been analysed to what extent their results for the U.S. economy are confirmed by data from other industrialized economies. The main findings of the present paper are that, as in King and Watson (1992), little evidence can be found in the data against the long-run neutrality of various monetary aggregates in the majority of G7-countries. Evidence against the superneutrality of money is, however, relatively easy to detect. The present paper also finds that the data from G7-countries do not reject the hypothesis of a long-run vertical Phillips curve, consistent with the findings of King and Watson (1992) for the U.S. economy. The only substantially different conclusion is reached with respect to the validity of the 'Fisher-hypothesis' of a unit long-run effect of inflation on nominal interest rates, which is rejected by the data from the U.S. and U.K. economies, but not for the remaining G7-countries. Controversial evidence on the long-run 'Fisher-effect' is also reported in Sargent (1973), Summers (1983) and Mishkin (1984).

The above conclusions are subject to some reservations: first, due to problems concerning the availability of compatible international macroeconomic data, the longest sample periods considered in the present study cover only 35 years of quarterly data, and frequently much shorter samples had to be used. Obviously, only limited 'long-run' information is conveyed by such short data spans, and this may seriously undermine the power of both the unit root and the long-run neutrality tests conducted in this paper. Second, the applicability of the neutrality tests developed by King and Watson (1992) critically depends on the order of integration in the data, which typically is subject to a large degree of uncertainty due to the low power of most unit root tests. Furthermore, the sub-set of time series which possess the postulated unit root properties may be small, so that in many cases long-run neutrality tests may only be applicable to a limited extent. This has been found to be the case for tests regarding the long-run neutrality of a number of monetary aggregates in G7-countries. Similar caution about the results obtained was found to be justified in neutrality tests involving inflation rates and real interest rates. Third, like any form of econometric policy evaluation, the neutrality tests proposed by King and Watson (1992) are likely to be subject to the famous 'Lucas critique', according to which the structure of econometric models is in general not invariant to changes in policy objectives, operating procedures or policy constraints over time. The present paper shows for the case of Germany that inference about certain long-run neutrality propositions displays a considerable degree of time-variation, with major changes occurring at policy sensitive points in time in the aftermath of major real economic shocks, such as the oil price hikes and the German unification shock. Changes in the degree to which the data are consistent with certain long-run neutrality propositions have thereby been detected by using a recursive estimation approach.

Inspite of these limitations, the King and Watson (1992) methodology provides a powerful tool for analysing the long-run predictions of economic theories within a fairly general framework. In future research this approach may fruitfully be applied to other fields of economics in which long-run neutrality or homogeneity propositions play a key role: in the context of the EMS, for example, the issue of German dominance and asymmetry may be analysed in terms of a long-run unit effect of German inflation on inflation rates in the remaining EMS countries. The link between trade volumes and nominal exchange rates and/or nominal exchange rate variability, which is a major unresolved puzzle in the empirical literature on international trade, may also be analysed along these lines. Finally, the long-term sterilisation of foreign exchange interventions is a third interesting economic problem where the application of the King and Watson (1992) methodology may provide useful insights.

Time Series and Data Sources

All data are quarterly seasonally adjusted data. In case the original data were not seasonally adjusted, seasonal adjustment was carried out using the multiplicative adjustment procedure in MicroTSP 7.0. The time series and data sources used were:

Monetary Aggregates:

Germany:	MCB, M1, M2, M3, M3E	International Monetary Fund (IMF), *International Financial Statistics*, various issues.
France:	M1, M2, M3, M4	IMF, *International Financial Statistics*, various issues.
Italy:	M1, M2	IMF, *International Financial Statistics*, various issues.
U.K.:	M0, M1, M3	IMF, *International Financial Statistics*, various issues.
	M4	Organisation for Economic Development and Cooperation (OECD), *Main Economic Indicators*, various issues.
U.S.A.:	M1, M2, M3	IMF, *International Financial Statistics*, various issues.
Canada.:	M1	OECD, *Main Economic Indicators*, various issues.
	M2	IMF, *International Financial Statistics*, various issues.
Japan:	M1, M2, M2+CD's	IMF, *International Financial Statistics*, various issues.

Output (Real GDP, GNP):

All G7 Countries:	IMF, *International Financial Statistics*, various issues.

Consumer Price Indices:

All G7 Countries:	IMF, *International Financial Statistics*, various issues.

Unemployment Rates:

All G7 Countries:	OECD Main Economic Indicators, various issues.

Three Month Interest Rates:

Germany	Money Market Rate	OECD, *Main Economic Indicators*, various issues.
France	Money Market Rate	OECD, *Main Economic Indicators*, various issues.
Italy	Money Market Rate	IMF, *International Financial Statistics*, various issues.
U.K.	Treasury Bill Rate	IMF, *International Financial Statistics*, various issues.
U.S.A.	Treasury Bill Rate	IMF, *International Financial Statistics*, various issues.
Canada	Treasury Bill Rate	IMF, *International Financial Statistics*, various issues.
Japan	CD's Rate	OECD, *Main Economic Indicators*, various issues.

References

Ahmed, Shagil, (1993), "Does Money Affect Output?," *Federal Reserve Bank of Philadelphia Business Review*, July/August, pp. 13-28.

Bernake, Ben S. (1986), "Alternative Explanations of the Money-Income Correlation, *Carnegie Rochester Conference Series on Public Policy*, 25, pp. 49-100.

Blanchard, Olivier J. and Laurence H. Summers, (1986), "Hysteresis and the European Unemployment Problem," in: *NBER Macroeconomics Annual 1986*, MIT Press, Cambridge, pp. 15-77.

Boschen, John F. and Leonard O. Mills (1988), "Tests of the Relation Between Money and Output in Real Business Cycle Models," *Journal of Monetary Economics*, 22, pp. 355-374.

Cooley, Thomas F. and Stephen F. LeRoy (1985), "Atheoretical Macroeconometrics: A Critique," *Journal of Monetary Economics*, 16, pp. 283-308.

Dickey, David D. and Wayne A. Fuller, (1981), "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root", *Econometrica*, Vol. 49, No. 4, pp. 1057-1072.

Duck, Nigel W. (1988), "Money, Output and Prices: An Empirical Study Using Long-term Cross Country Data," *European Economic Review*, 32, pp. 1603-1619.

Duck, Nigel W. (1993), "Some International Evidence on the Quantity Theory of Money," *Journal of Money, Credit and Banking*, 25, pp. 1-12.

Eichenbaum, Martin and Kenneth J. Singleton (1986), "Do Equilibrium Real Business Cycle Theories Explain Postwar U.S. Business Cycles?, NBER Macroeconomics Annual 1986, MIT Press, Cambridge, MA, pp. 91-134.

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

Fuller, Wayne A., (1976), *Introduction to Statistical Time Series*, John Wiley & Sons, New York.

Geweke, John, (1982), "Measurement of Linear Dependence and Feedback Between Multiple Time Series," *Journal of the American Statistical Association*, 77, pp. 304-313.

Geweke, John, (1986), "The Superneutrality of Money in the United States: An Interpretation of the Evidence," *Econometrica*, 54, pp. 1-21.

Hagen, Jürgen von, (1984), "The Causal Role of Money in West Germany - Some Contradicting Comments and Evidence," *Weltwirtschaftliches Archiv*, 120.

Hansen, Gerd, (1989), "Testing for Money Neutrality," *European Journal of Political Economy*, 5, pp. 89-112.

King, Robert G. and Charles I. Plosser, (1984), "Money, Credit, and Prices in a Real Business Cycle,", *American Economic Review*, 74, pp. 363-380.

King, Robert G., Charles I. Plosser, James H. Stock and Mark W. Watson, (1992), "Stochastic Trends and Economic Fluctuations", *American Economic Review*, 82, pp. 819-840.

King, Robert G. and Mark W. Watson, (1992), "Testing Long Run Neutrality", National Bureau of Economic Research Working Paper No. 4156.

Litterman, Robert B. and Laurence Weiss (1985), "Money, Real Interest Rates, and Output: A Reinterpretation of Postwar U.S. Data," *Econometrica*, 53, pp. 129-156.

Loef, Hans-E. (1993), "Long-Run Monetary Relationships in the EC Countries," *Weltwirtschaftliches Archiv* 129, pp. 33-54.

Lothian, James R. (1985), "Equilibrium Relationships Between Money and Other Economic Variables," *American Ecconomic Review*, 75, pp. 828-835.

Lucas, Robert E. Jr (1972), "Econometric Testing of the Natural Rate Hypothesis," in Eckstein, O., ed., *The Economics of Price Determination*, Board of Governors of the Federal Reserve System, Washington D.C.

Lucas, Robert E. Jr (1973), "Some International Evidence on Output-Inflation Trade-offs," *American Ecconomic Review*, 63, pp. 326-334.

Lucas, Robert E. Jr (1976), "Econometric Policy Evaluation: A Critique," *Carnegie Rochester Conference Series on Public Policy*, 1, pp. 326-334.

Lucas, Robert E. Jr (1980), "Two Illustrations of the Quantity Theory of Money," *American Ecconomic Review*, 70, pp. 1005-1014.

Manchester, Joyce (1989), "How Money Affects Real Output", *Journal of Money, Credit and Banking*, 21, pp. 16-32.

McCallum, Bennet T. (1984), "On Low-Frequency Estimates of Long-Run Relationships in Macroeconomics", *Journal of Monetary Economics*, 14, pp. 3-14.

Mills, Terrence C. (1982), "Signal Extraction and Two Illustrations of the Quantity Theory," *American Ecconomic Review*, 72, pp. 1162-1168.

Mishkin, Frederic S. (1984), "The Real Interest Rate: A Multi-Country Empirical Study," *Canadian Journal of Economics*, 1, pp. 283-311.

Nelson, Charles R. and Charles I. Plosser, (1982), "Trends and Random Walks in Macroeconomic Time Series: Some Evidence and Implications," *Journal of Monetary Economics*, 10, pp. 139-162.

Newey, Whitney K. and Kenneth D. West, (1985), "A Simple Positive Definite Heteroscedastisity and Autocorrelation Consistent Covariance Matrix," *Econometrica*, 55, pp.703-708.

Perron, Pierre, (1988), "Trends and Random Walks in Macroeconomic Time Series: Further Evidence from a New Approach," *Journal of Economic Dynamics and Control*, 12, pp. 297-332.

Phillips, A.W.H., (1958), "The Relation between Unemployment and the Rate of Change of Money Wage Rates in the United Kingdom, 1861-1957," *Economica*, pp. 283-299.

Phillips, Peter C. B., (1987), "Time Series Regression with a Unit Root, *Econometrica*, Vol. 55, No. 2, pp. 277-301.

Phillips, Peter C. B. and Pierre Perron, (1988), "Testing for a Unit Root in Time Series Regression," *Biometrica*, Vol. 75, No. 2, pp. 335-346.

Plosser, Charles I., (1989), "Money and Business Cycles: A Real Business Cycle Interpretation," Rochester Center for Economic Research Working Paper No. 210.

Sargent, Thomas J. (1971), "A Note on the Accelerationist Controversy," *Journal of Money, Credit and Banking*, 3, pp. 50-60.

Sargent, Thomas J. (1973), "Interest Rates and Prices in the Long Run: A Study of the Gibson Paradox," *Journal of Money, Credit and Banking*, 5, pp. 383-449.

Sims, Charles A. (1972), "Money, Income and Causality,", *American Economic Review, Papers and Proceedings*, 62, pp. 540-552.

Stock, James H., (1991), "Confidence Intervals for the Largest Unit Root in U.S. Macroeconomic Time Series," *Journal of Monetary Economics*, 28, pp. 435-459.

Stock, James H. and Mark W. Watson, (1988a), "Interpreting the Evidence on Money-Income Causality", *Journal of Econometrics*, Volume 40, Number 1, pp. 161-182.

Stock, James H. and Mark W. Watson, (1988b), "Variable Trends in Economic Time Series", *Journal of Economic Perspectives*, Vol. 2, Number 3, pp. 147-174.

Summers, Lawrence H. (1983), "The Non-adjustment of Nominal Interest Rates: A Study of the Fisher Effect," in: James Tobin, ed., *Macroeconomics, Prices and Quantities: Essays in Memory of Arthur Okun*, The Brookings Institution, Washington D.C.

Weber, Axel A., (1990), "The Credibility of Monetary Target Announcements: An Empirical Evaluation", Centre for Economic Research Discussion Paper No. 9031, Tilburg.

Weber, Axel A., (1991), "Credibility, Reputation and the European Monetary System", *Economic Policy*, 12, pp. 57-102.

Weissenberger, E., (1984), "The Causal Role of Money in West Germany," *Weltwirtschaftliches Archiv*, 119, pp. 64-83.

Table 1a
Unit-Root Test Statistics for Money in G7-Countries

Country	Variable /Period	Augmented Dickey-Fuller Tests Detrended Data				Augmented Dickey-Fuller Tests Demeaned Data				Phillips-Perron Tests			Decision
		$t^{\tau}(z)$	sig. level	$\hat{\rho}$	Stock's ρ intervals	$t^{\mu}(z)$	sig. level	$\hat{\rho}$	Stock's ρ intervals	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha*})$	$Z(t_{\tilde{\alpha}})$	
Germany	MCB 64:1-90:2	-0.88	(-)	0.99	(0.98,1.05)	-1.85	-	1.00	(0.87,1.04)	11.6	-2.20	-0.75	I(1)+
	ΔMCB	-3.57	(**)	0.33	(0.64,1.02)	-3.03	**	0.47	(0.70,1.00)	-2.37(**)	-8.79***	-9.19(***)	drift
	M1 62:1-90:4	-1.80	(-)	0.96	(0.89,1.04)	-0.60	-	1.00	(0.97,1.04)	9.49	-0.74	-1.67	I(1)+
	ΔM1	-4.06	/***)	0.37	(,0.93)	-4.06	***	0.39	(- ,0.88)	-3.04(***)	-6.47***	-6.48(***)	drift
	M2 72:1-90:4	-3.21	*	0.90	(0.54,1.04)	-0.88	(-)	1.00	(0.94,1.07)	6.87	-0.97	-2.77	I(1)+
	ΔM2	-3.09	(*)	0.47	(0.61,1.05)	-3.08	**	0.51	(0.58,1.01)	-2.84(***)	-6.32***	-6.40(***)	drift
	M3 72:1-90:4	-2.47	-	0.96	(0.73,1.06)	-2.45	(-)	0.99	(0.71,1.04)	9.84	-4.53***	-2.63	I(2),
	ΔM3	-2.96	-	0.36	(0.62,1.05)	-2.10	(-)	0.72	(0.75,1.04)	-1.33	-5.13(***)	-6.36***	I(1)+t
	M3E 76:1-90:4	-2.33	(-)	0.91	(0.66,1.08)	-2.26	-	0.99	(0.66,1.05)	11.53	-2.87**	-2.36	I(2),
	ΔM3E	-3.33	*	-0.19	(0.37,1.05)	-2.41	(-)	0.28	(0.60,1.04)	-1.62(*)	-5.90***	-6.24(***)	I(1)+d
France	M1 80:1-92:4	1.08	(-)	1.04	(1.05,1.14)	-2.52	*	0.97	(0.51,1.05)	3.76	-3.64***	0.89	I(2),
	ΔM1	-3.13	*	-0.28	(0.33,1.06)	-1.29	(-)	0.71	(0.84,1.10)	-4.90(***)	-7.20(***)	-8.95***	I(1)+t
	M2 80:1-92:4	0.50	(-)	1.01	(1.04,1.14)	-2.22	-	0.98	(0.59,1.06)	3.12	-4.70***	0.77	I(2),
	ΔM2	-2.62	-	0.14	(0.50,1.09)	-0.59	(-)	0.93	(0.94,1.11)	-2.69(***)	-4.35(***)	-7.72***	I(1)+t
	M3 80:1-92:4	0.56	(-)	1.02	(1.05,1.13)	-1.86	-	0.99	(0.71,1.08)	9.45	-3.96***	0.51	I(2),
	ΔM3	-1.88	-	0.21	(0.75,1.11)	-0.72	(-)	0.84	(0.91,1.11)	-1.88(*)	-6.31(***)	-7.70***	I(1)+t
	M4 80:1-92:4	0.91	(-)	1.04	(1.05,1.14)	-1.98	-	0.99	(0.69,1.08)	9.55	-4.03***	0.09	I(2),
	ΔM4	-1.94	(-)	0.09	(0.71,1.11)	-0.70	-	0.83	0.93,1.11)	-1.91(*)	-6.30(***)	-7.65***	I(1)+t
Italy	M1 66:1-92:4	1.13	-	1.01	(1.02,1.06)	-2.45	(-)	1.00	(0.80,1.02)	8.90	-3.45***	1.07	I(2),
	ΔM1	-3.13	*	0.47	(0.71,1.03)	-1.98	(-)	0.73	(0.86,1.03)	-2.55(**)	-6.61(***)	-7.58***	I(1)+t
	M2 66:1-92:4	0.96	(-)	1.01	(1.02,1.06)	-2.10	-	1.00	(0.83,1.03)	11.31	-2.59*	0.97	I(2),
	ΔM2	-1.97	-	0.64	(0.88,1.05)	-1.18	(-)	0.81	(0.94,1.04)	-3.02(***)	-9.00***	-9.46(***)	I(1)+d

Table 1a continued

UK	M0 71:2-92:4 ΔM0	-1.13 -1.60	(-) (-)	0.97 0.61	(0.92,1.09) (0.83,1.09)	-1.07 -1.36	- -	0.99 0.72	(0.90,1.08) (0.85,1.08)	6.05 -4.44(***)	-1.78 -7.50***	-0.30 -7.75(***)	I(2), I(1)+d
	M1 59:1-86:4 ΔM1	-0.92 -5.03	(-) ***	0.98 -0.49	(0.97,1.05) (- ,0.77)	2.34 -3.05	- (**)	1.01 0.37	(- , -) (0.72,1.01)	6.82 -10.1(***)	3.32 -12.4(***)	-1.28 -14.1***	I(1)+ trend
	M3 65:1-86:4 ΔM3	-2.75 -2.82	- (-)	0.94 0.49	(0.72,1.05) (0.69,1.04)	0.78 -2.52	(-) *	1.00 0.60	(1.01,1.06) (0.73,1.02)	9.19 -3.75(***)	1.42 -8.14***	-2.64 -8.51(***)	I(2), I(1)+d
	M4 65:1-92:4 ΔM4	-3.12 -2.18	- (-)	0.94 0.72	(0.72,1.03) (0.86,1.04)	-0.94 -2.46	(-) -	1.00 0.70	(0.95,1.04) (0.81,1.02)	13.1 -1.68(*)	0.22 -6.54***	-2.68 -6.54(***)	I(2), I(1)+d
USA	M1 59:1-92:4 ΔM1	-2.40 -3.63	- **	0.97 0.42	(0.87,1.03) (0.71,1.01)	2.02 -2.56	(-) (-)	1.00 0.66	(- , -) (0.83,1.02)	10.77 -3.69(***)	2.96 -8.41(***)	-2.57 -9.04***	I(1)+ trend
	M2 61:1-92:4 ΔM2	-0.27 -2.77	(-) (-)	1.00 0.68	(1.01,1.05) (0.81,1.02)	-1.62 -2.54	- -	1.00 0.71	(0.91,1.03) (0.82,1.02)	11.40 -1.74	-1.58 -9.06***	0.70 -6.35(***)	I(2), I(1)+d
	M3 61:1-92:4 ΔM3	0.30 -1.73	(-) -	1.00 0.86	(1.02,1.05) (0.91,1.04)	-1.94 -1.27	- (-)	1.00 0.90	(0.88,1.03) (0.94,1.04)	9.44 -1.32	-1.93 -3.72***	1.73 -4.33(***)	I(2), I(1)+d
Canada	M1 62:1-92:4 ΔM2	-0.84 -3.15	(-) (***)	0.98 0.32	(0.98,1.05) (0.75,1.02)	-1.08 -3.03	- **	1.00 0.36	(0.95,1.04) (0.75,1.00)	6.58 -7.93(***)	-1.19 -11.0***	-0.59 -11.1(***)	I(1)+ drift
	M2 72:1-92:4 ΔM2	-1.95 -3.40	(-) **	0.97 0.30	(0.84,1.06) (0.54,1.03)	-2.83 -2.34	* (-)	0.99 0.62	(0.66,1.02) (0.75,1.03)	6.20 -2.34(**)	-3.90*** -5.59***	-1.84 -6.68(***)	I(1)+d I(1)+t
Japan	M1 59:1-92:4 ΔM1	-0.59 -3.45	(-) **	1.00 0.39	(1.01,1.04) (0.74,1.02)	-3.16 -1.98	** (-)	0.99 0.77	(0.76,1.00) (0.89,1.03)	6.54 -5.30(***)	-5.36*** -9.71(***)	-0.39 -11.4***	I(1)+ trend
	M2 59:1-92:4 ΔM2	0.61 -3.29	(-) *	1.00 0.53	(1.02,1.05) (0.77,1.02)	-3.41 -1.17	- (-)	1.00 0.90	(0.72,0.96) (0.95,1.03)	8.52 -2.49(**)	-6.36*** -6.40(***)	0.54 -9.20***	I(2), I(1)+t
	M2CD 81:1-92:4 ΔM2CD	-0.76 -0.83	(-) (-)	0.84 0.78	(0.74,1.13) (0.94,1.14)	-1.60 -0.72	- -	0.99 0.80	(0.76,1.10) (0.89,1.12)	6.36 -2.13(**)	-1.73 -5.87***	-0.10 -6.64(***)	I(2), I(1)+d

Key to Table: See Table 1b

Table 1b
Unit-Root Test Statistics for Output in G7-Countries

Country	Variable /Period	Augmented Dickey-Fuller Tests Detrended Data				Augmented Dickey-Fuller Tests Demeaned Data				Phillips-Perron Tests			Decision
		$t^{\tau}(z)$	sig. level	$\hat{\rho}$	Stock's ρ intervals	$t^{\mu}(z)$	sig. level	$\hat{\rho}$	Stock's ρ intervals	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha *})$	$Z(t_{\tilde{\alpha}})$	
Germany	Y 64:1-92:4	-2.12	-	0.95	(0.87,1.04)	-2.15	-	0.99	(0.86,1.03)	5.88	-2.32	-2.16	I(1)+
	ΔY	-3.90	(***)	-0.03	(- , -)	-3.49	***	0.13	(0.70,0.96)	-10.23(**)	-12.1***	-12.5(***)	drift
France	Y 72:1-92:4	-2.51	-	0.92	(0.73,1.05)	-1.65	-	0.99	(0.86,1.05)	5.55	-2.01	-2.65	I(1)+
	ΔY	-2.90	-	0.39	(0.65,1.04)	-2.70	*	0.49	(0.71,1.03)	-5.32(***)	-7.44***	-7.71(***)	drift
Italy	Y 62:1-92:4	-1.02	-	0.98	(0.97,1.04)	-3.25	**	0.99	(0.73,0.98)	5.67	-3.02**	-1.18	I(1)+
	ΔY	-4.45	**	-0.10	(- ,0.88)	-3.55	(***)	0.26	(0.68,0.94)	-8.41(***)	-10.2***	-10.7(***)	drift
UK	Y 59:1-92:4	-2.49	-	0.91	(0.86,1.03)	-2.03	-	0.99	(0.91,1.03)	5.23	-2.16	-2.49	I(1)+
	ΔY	-3.80	(**)	-0.01	(- ,1.01)	-3.45	***	0.12	(0.82,1.02)	-12.1(***)	-13.8***	-14.2(***)	drift
USA	Y 59:1-92:4	-2.10	-	0.95	(0.88,1.03)	-1.40	-	1.00	(0.93,1.03)	6.20	-1.72	-2.04	I(1)+
	ΔY	-4.31	(***)	0.29	(- ,0.90)	-4.14	***	0.35	(- ,0.89)	-6.75(***)	-8.91***	-9.03(***)	drift
Canada	Y 59:1-92:4	0.07	-	1.00	(1.01,1.04)	-2.26	-	0.99	(0.86,1.02)	7.09	-2.67*	0.20	I(1)+
	ΔY	-4.02	***	0.25	(- ,0.93)	-3.33	(**)	0.44	(0.74,0.97)	-6.53(***)	-9.55***	-9.96(***)	drift
Japan	Y 59:1-92:4	-1.87	-	0.99	(0.91,1.04)	-4.10	***	0.99	(- ,0.89)	7.79	-5.92***	-2.00	I(1)+
	ΔY	-3.92	**	0.27	(- ,0.95)	-2.16	-	0.72	(0.74,0.97)	-5.71(***)	-10.9(***)	-12.7***	trend

Key to Table: $t^{\tau}(z)$ in column 3 and $t^{\tau}(z)$ in column 7 are the augmented Dickey-Fuller tests for detrended data and demeaned data respectively. Their significance levels are taken from Table 8.5.2. of Fuller (1976), p. 373. A rejection of the null hypothesis of a unit root at the 1% significance level is marked with ***, at the 5% level with ** and at the 10% level with *. Significance levels in brackets indicate that the corresponding coefficient of the deterministic trend in column 3 (drift in column 7) was not significantly different from zero at the 5% level. Stock's (1992) 95% confidence intervals for the largest unit root ρ were calculated from the ADF statistics using Stock's Tables A1 and A2 and the procedure described in Appendix B of his paper. In addition to the confidence belts for ρ the estimated roots $\hat{\rho}$ are displayed. The three Phillips-Perron unit root tests reported are discussed in Perron (1988). $Z(t_{\hat{\alpha}})$ tests the null hypothesis of a unit root against an AR(1) regression without deterministic drift or trend, whilst $Z(t_{\alpha *})$ includes a drift and $Z(t_{\tilde{\alpha}})$ both a drift and trend. Estimates were obtained using the Newey-West estimator. A rejection of the null hypothesis of a unit root at the 1% significance level is marked with ***, at the 5% level with ** and at the 10% level with *. Significance levels in brackets indicate that the Phillips-Perron test $Z(t_{\mu *})$ for a drift and $Z(t_{\tilde{\beta}})$ for a trend were not significant at the 5% level. Critical values for the Phillips-Perron tests are taken from Tabele 8.5.2. in Fuller (1976), p. 371, and Tables I to III in Dickey and Fuller (1981), p. 1062. All ADF statistics are based on regressions including six lagged differences of the variable, and the Newey-West estimators for the Phillips-Perron used a lag-window of order six.

Table 1c
Unit-Root Test Statistics for Consumer Price Inflation Rates in G7-Countries

Country	Variable /Period	Augmented Dickey-Fuller Tests Detrended Data				Augmented Dickey-Fuller Tests Demeaned Data				Phillips-Perron Tests			Decision
		$t^{\tau}(z)$	sig. level	$\hat{\rho}$	Stock's ρ intervals	$t^{\mu}(z)$	sig. level	$\hat{\rho}$	Stock's ρ intervals	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\tilde{\alpha}})$	
Germany	ΔP 60:1-92:4	-2.36	(-)	0.78	(0.86,1.03)	-2.37	-	0.78	(0.85,1.03)	-4.01(***)	-8.18***	-8.16(***)	I(1),
	Δ^2P	-5.63	(***)	-1.92	(- ,0.72)	-5.66	***	-1.91	(- ,0.71)	-20.8***	-20.8(***)	-20.7(***)	I(0)
France	ΔP 60:1-92:4	-1.13	(-)	0.94	(0.96,1.04)	-1.25	-	0.93	(0.94,1.04)	-1.53	-3.32**	-3.31(*)	I(1)
	Δ^2P	-5.51	(***)	-1.17	(-- ,0.74)	-5.47	**	-1.14	(- ,0.74)	-17.8***	-17.8(***)	-17.9(***)	
Italy	ΔP 60:1-92:4	-1.41	(-)	0.91	(0.94,1.04)	-1.63	-	0.91	(0.91,1.03)	-1.86	-3.91(***)	-4.10(***)	I(1)
	Δ^2P	-6.78	(***)	-1.81	(- , -)	-6.71	***	-1.73	(- , -)	-20.7***	-20.7(***)	-20.9(***)	
UK	ΔP 60:1-92:4	-2.03	(-)	0.83	(0.89,1.04)	-2.23	-	0.82	(0.86,1.02)	-3.30(***)	-6.54***	-6.60(***)	I(1),
	Δ^2P	-6.43	(***)	-1.94	(- , -)	-6.31	***	-1.87	(- , -)	-24.6***	-24.6(***)	-24.9(***)	I(0)
USA	ΔP 60:1-92:4	-2.42	(-)	0.85	(0.85,1.03)	-2.45	-	0.86	(0.84,1.02)	-1.69	-3.89***	-3.98(***)	I(1),
	Δ^2P	-4.29	(***)	-0.55	(- ,0.92)	-4.27	***	-0.52	(- ,0.88)	-17.1***	-17.1(***)	-17.1(***)	I(0)
Canada	ΔP 60:1-92:4	-1.63	(-)	0.89	(0.93,1.04)	-1.84	-	0.89	(0.89,1.03)	-1.91	-4.37***	-4.62(***)	I(1),
	Δ^2P	-5.45	(***)	-1.43	(- ,0.76)	-5.40	***	-1.35	(- ,0.75)	-20.4***	-20.4(***)	-20.6(***)	I(0)
Japan	ΔP 60:1-92:4	-3.26	(*)	0.64	(0.76,1.02)	-2.83	*	0.70	(0.79,1.01)	-4.89(***)	-8.11***	-8.62(***)	I(1),
	Δ^2P	-6.25	(***)	-1.96	(- , -)	-6.21	***	-3.70	(- , -)	-27.8***	-27.8(***)	-27.9(***)	I(0)

Key to Table: See Table 1b

Table 1d
Unit-Root Test Statistics for Unemployment Rates in G7-Countries

Country	Variable /Period	Augmented Dickey-Fuller Tests Detrended Data				Augmented Dickey-Fuller Tests Demeaned Data				Phillips-Perron Tests			Decision
		$t^{\tau}(z)$	sig. level	$\hat{\rho}$	Stock's ρ intervals	$t^{\mu}(z)$	sig. level	$\hat{\rho}$	Stock's ρ intervals	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha*})$	$Z(t_{\tilde{\alpha}})$	
Germany	u 64:1-92:4	-2.74	-	0.96	(0.79,1.03)	-1.08	(-)	0.99	(0.66,1.03)	0.29	-1.10	-1.64	I(1)
	Δu	-3.67	(**)	0.68	(0.66,1.02)	-3.70	***	0.68	(- ,0.93)	-3.66***	-3.80(***)	-3.78(**)	
France	u 70:1-92:4	-1.69	(-)	0.96	(0.90,1.06)	-0.76	-	1.00	(0.95,1.06)	2.10	-0.89	-1.39	I(1)
	Δu	-2.92	(-)	0.53	(0.69,1.04)	-2.93	*	0.53	(0.67,1.01)	-4.60***	-5.21(***)	-5.21(***)	
Italy	u 62:1-92:4	-3.07	-	0.86	(0.77,1.03)	-0.18	(-)	1.00	(1.00,1.04)	1.78	0.16	-3.14*	I(1)
	Δu	-3.76	(***)	0.00	(- ,0.99)	-3.64	***	0.04	(- ,0.93)	-12.1***	-12.3(***)	-12.4(***)	
UK	u 62:1-92:4	-2.87	-	0.97	(0.80,1.03)	-0.93	(-)	1.00	(0.95,1.04)	0.58	-0.82	-2.03	I(1)
	Δu	-3.68	(**)	0.74	(0.68,1.01)	-3.69	***	0.74	(- ,0.94)	-3.35***	-3.48(***)	-3.50(**)	
USA	u 62:1-92:4	-2.75	(-)	0.94	(0.80,1.03)	-2.28	-	0.96	(0.85,1.02)	-0.31	-1.93	-2.44	I(1)
	Δu	-4.01	(***)	-0.52	(- ,0.92)	-4.04	***	0.51	(- ,0.89)	-5.31***	-5.31(***)	-5.28(***)	
Canada	u 62:1-92:4	-2.96	-	0.94	(0.78,1.03)	-1.28	(-)	0.98	(0.69,1.02)	0.48	-0.89	-2.60	I(1)
	Δu	-3.69	(**)	0.52	(0.69,1.02)	-3.68	***	0.54	(- ,0.94)	-5.90***	-5.94(***)	-5.94(***)	
Japan	u 62:1-92:4	-2.38	(-)	0.92	(0.86,1.04)	-1.20	-	0.98	(0.95,1.04)	0.46	-1.02	-1.97	I(1)
	Δu	-3.66	(**)	0.15	(0.68,1.01)	-3.67	***	0.15	(- ,0.94)	-11.9***	-12.0(***)	-12.0(***)	

Key to Table: See Table 1b

Table 1e
Unit-Root Test Statistics for Nominal and Real Three-Month Interest Rates in G7 Countries

Country	Variable /Period	Augmented Dickey-Fuller Tests Detrended Data				Augmented Dickey-Fuller Tests Demeaned Data				Phillips-Perron Tests			Decision
		$t^{\tau}(z)$	sig. level	$\hat{\rho}$	Stock's ρ intervals	$t^{\mu}(z)$	sig. level	$\hat{\rho}$	Stock's ρ intervals	$Z(t_{\hat{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\tilde{\alpha}})$	
Germany	r 62:1-92:4	-3.09	(-)	0.87	(0.77,1.03)	-2.89	*	0.89	(0.78,1.01)	-0.74	-2.88(*)	-3.01	I(1)
	Δr	-4.77	(***)	0.17	(- ,0.84)	-4.79	***	0.17	(- ,0.81)	-8.09***	-8.10(***)	-8.10(**)	
France	r 72:1-92:4	-2.50	(-)	0.88	(0.73,1.05)	-2.58	*	0.88	(0.72,1.03)	-0.18	-2.51	-2.44	I(1)
	Δr	-3.78	(**)	0.14	(- ,1.01)	-3.77	***	0.15	(- ,0.89)	-5.51***	-5.52(***)	-5.54(***)	
Italy	r 73:1-92:4	-2.33	(-)	0.88	(0.76,1.06)	-2.60	*	0.88	(0.71,1.03)	0.00	-2.81(*)	-2.69	I(1)
	Δr	-3.60	(**)	-0.15	(0.52,1.03)	-3.36	**	-0.01	(0.54,0.94)	-6.34***	-6.36(***)	-6.45(***)	
UK	r 59:1-92:4	-3.54	**	0.77	(0.72,1.01)	-2.44	(-)	0.92	(0.84,1.02)	-0.64	-2.48	-3.04	I(1),
	Δr	-5.12	(***)	-0.18	(- ,0.80)	-5.09	***	-0.17	(- ,0.80)	-10.3***	-10.3(***)	-10.4(**)	I(0)+t
USA	r 59:1-92:4	-2.16	(-)	0.92	(0.89,1.03)	-2.24	-	0.94	(0.86,1.02)	-0.82	-2.14	-1.90	I(1)
	Δr	-5.38	(***)	-0.03	(- ,0.78)	-5.26	***	0.01	(- ,0.77)	-9.77***	-9.77(***)	-9.84(***)	
Canada	r 59:1-92:4	-2.91	(-)	0.88	(0.79,1.02)	-2.12	-	0.95	(0.87,1.02)	-0.61	-2.14	-2.51	I(1)
	Δr	-5.19	(**)	-0.04	(- ,0.81)	-5.17	***	-0.02	(- ,0.78)	-8.75***	-8.76(***)	-8.76(***)	
Japan	r 79:2-92:1	-2.88	-	0.77	(0.46,1.08)	-1.85	(-)	0.88	(0.71,1.08)	-0.55	-2.12	-2.73	I(1)
	Δr	-3.42	(**)	0.08	(0.21,1.05)	-3.45	***	0.08	(0.21,0.88)	-3.35***	-3.34(***)	-3.46(***)	
Germany	r-π 62:1-92:4	-4.49	***	0.37	(- ,0.88)	-3.46	(***)	0.59	(0.70,0.96)	-4.47(***)	-7.52(***)	-8.19***	I(0)+t
France	r-π 72:1-92:4	-2.88	-	0.57	(0.69,1.05)	-0.16	(-)	0.99	(0.99,1.06)	-1.22(-)	-3.31(-)	-6.25***	I(0),I(1)
Italy	r-π 73:1-92:4	-3.13	*	0.51	(0.59,1.04)	-1.14	(-)	0.91	(0.91,1.06)	-2.50(-)	-3.18(**)	-6.43***	I(0),I(1)
UK	r-π 59:1-92:4	-2.11	(-)	0.79	(0.88,1.03)	-1.85	-	0.83	(0.90,1.03)	-7.76***	-8.18(***)	-8.55(***)	I(0),I(1)
USA	r-π 59:1-92:4	-2.33	(-)	0.82	(0.86,1.03)	-2.30	-	0.83	(0.85,1.02)	-4.18***	-5.32(***)	-5.39(***)	I(0),I(1)
Canada	r-π 59:1-92:4	-2.30	(-)	0.79	(0.88,1.03)	-1.92	-	0.86	(0.89,1.03)	-3.33(***)	-5.11***	-5.94(***)	I(0),I(1)
Japan	r-π 79:2-92:1	-1.95	(-)	0.50	(0.72,1.11)	-1.96	-	0.52	(0.68,1.08)	-4.86(***)	-9.71***	-9.78(***)	I(0),I(1)

Key to Table: See Table 1b.

Note: All real interest rates (r-π) were found to be stationary in first differences without significant drifts or trends

Table 2
The Neutrality of Money in G7-Countries

Country	Money / Period	VAR Estimates			Structural Model Estimates			γ_{ym}=0 in 95% conf. interval			Estimates imposing γ_{ym}=0		
		σ^2_y	σ^2_m	cor_{ym}	σ^2_y	σ^2_m	cor_{ym}	λ_{my}	λ_{ym}	γ_{my}	λ_{my}	λ_{ym}	γ_{my}
Germany	MCB 64:1-90:2	4.44	3.07	0.07	5.74	6.28	0.73	≥0.25	≤-0.49	≥0.27	0.71 (0.38)	-1.49 (0.87)	0.80 (0.30)
	M1 62:1-90:4	4.93	4.14	0.07	5.15	6.85	0.40	≥0.20	≤-0.21	≥-0.46	0.65 (0.29)	-0.89 (0.42)	0.54 (0.41)
	M2 72:1-90:4	3.87	4.41	0.19	2.99	10.6	0.51	-	-	-	0.20 (0.38)	0.01 (0.28)	1.80 (1.28)
	M3 72:1-90:4	3.98	2.49	-0.01	3.92	9.29	0.26	-	-	-	0.36 (0.41)	-0.94 (0.99)	0.62 (1.26)
	M3E 76:1-90:4	3.81	3.06	-0.04	3.76	5.34	0.01	-	-	-	-0.06 (0.37)	0.04 (0.59)	0.01 (0.83)
France	M1 80:1-92:4	1.99	5.17	-0.09	4.32	17.6	-0.80	≤1.43	≥-0.22	≤1.32	-1.23 (1.25)	0.15 (0.19)	-3.25 (1.79)
	M2 80:1-92:4	1.91	3.62	-0.22	7.70	48.0	-0.94	≤1.60	≥-0.43	-	-1.83 (1.64)	0.50 (0.68)	-5.89 (3.15)
	M3 80:1-92:4	1.99	2.97	-0.16	5.07	16.5	-0.86	-	-	-	-1.88 (2.00)	0.91 (1.33)	-2.78 (1.89)
	M4 80:1-92:4	2.02	3.16	-0.20	4.61	14.9	-0.81	-	-	-	-1.58 (1.62)	0.65 (0.96)	-2.62 (1.97)
Italy	M1 67:1-92:4	5.06	5.06	0.21	6.44	23.7	0.66	≥-0.10	≤0.23	≤-1.51	0.91 (0.62)	-0.56 (0.59)	2.43 (1.47)
	M3 62:1-92:4	5.24	6.01	0.08	7.85	22.2	0.67	-	-	-	0.95 (0.82)	-0.70 (0.69)	1.90 (1.35)
UK	M0 71:2-86:4	5.51	5.47	0.04	4.56	13.0	-0.32	-	-	-	-0.63 (0.69)	0.67 (0.59)	-0.92 (1.48)
	M1 59:1-86:4	5.49	11.6	0.01	3.76	21.2	-0.23	-	-	-	0.88 (0.87)	-0.19 (0.19)	-1.28 (2.29)
	M3 65:1-86:4	5.43	7.23	0.05	4.08	19.17	-0.20	-	-	-	-0.18 (0.55)	0.14 (0.30)	-0.98 (2.24)
	M4 65:1-92:4	4.41	5.12	0.05	18.7	5.15	0.42	-	-	-	0.28 (0.31)	-0.17 (0.22)	0.12 (0.13)
USA	M1 59:1-92:4	3.45	3.77	0.12	5.39	11.8	-0.21	-	-	-	0.42 (0.44)	-0.25 (0.39)	-0.47 (0.96)
	M2 61:1-92:4	3.37	2.72	0.03	5.55	11.2	0.48	≤-.4,≥.1	≤-.1,≥.7	-	1.35 (1.23)	-2.12 (2.01)	0.97 (0.89)
	M3 61:1-92:4	3.40	2.52	0.12	6.82	31.4	0.56	-	-	-	0.80 (1.20)	-1.50 (2.86)	2.59 (3.84)
Canada	M1 62:1-92:4	3.90	7.50	0.01	7.48	10.6	0.70	≥0.75	≤-0.17	≥0.29	2.06 (1.00)	-0.45 (0.20)	0.99 (0.40)
	M2 70:1-92:4	3.58	5.97	-0.02	6.40	16.9	0.53	-	-	-	-0.48 (0.63)	0.16 (0.22)	1.14 (1.11)
Japan	M1 59:1-92:4	4.34	8.17	0.10	14.7	31.3	0.88	≥0.20	≤-0.01	≥1.22	1.04 (0.60)	0.08 (0.05)	1.88 (0.51)
	M2 59:1-92:4	4.24	4.32	0.06	15.2	28.8	0.90	≥0.25	≤-0.18	≥1.14	0.85 (0.51)	-0.80 (0.52)	1.71 (0.48)
	M2CD 81:2-92:4	2.32	3.03	0.17	3.45	7.85	0.85	≥0.93	≤-0.51	≥0.81	2.64 (2.22)	-2.20 (2.82)	1.94 (0.77)

Table 3
The Superneutrality of Money in G7-Countries

Country	Money / Period	VAR Estimates			Structural Model Estimates			$\gamma_{y\Delta m}=0$ in 95% conf. interval			Estimates imposing $\gamma_{y\Delta m}=0$		
		σ_y^2	$\sigma_{\Delta m}^2$	$cor_{y\Delta m}$	σ_y^2	$\sigma_{\Delta m}^2$	$cor_{y\Delta m}$	$\lambda_{\Delta my}$	$\lambda_{y\Delta m}$	$\gamma_{\Delta my}$	$\lambda_{\Delta my}$	$\lambda_{y\Delta m}$	$\gamma_{\Delta my}$
Germany	MCB 64:1-90:2	4.39	3.16	0.01	5.17	0.93	0.07	≥.03,≤1	-1.6 -2	-	0.39 (0.17)	-0.75 (0.28)	0.01 (0.05)
	M1 62:1-90:4	5.03	4.36	-0.01	5.79	1.89	-0.41	≤0.50	≥-0.60	≥0.08	0.15 (0.16)	-0.21 (0.18)	-0.13 (0.09)
	M2 72:1-90:4	3.85	4.68	0.18	3.61	2.06	0.67	≥0.38	≤-0.14	≥0.12	0.87 (0.36)	-0.50 (0.27)	0.39 (0.14)
	M3 72:1-90:4	3.99	2.53	-0.03	3.94	0.81	-0.12	≤0.75	≥-1.49	-	0.22 (0.16)	-0.58 (0.32)	-0.02 (0.07)
	M3E 76:1-90:4	3.87	3.17	-0.03	3.83	0.93	-0.21	-	-	-	-0.10 (0.20)	0.12 (0.26)	-0.05 (0.10)
France	M1 80:1-92:4	1.95	5.34	-0.10	3.25	1.70	-0.18	-	-	-	0.25 (0.81)	-0.07 (0.09)	-0.09 (0.24)
	M2 80:1-92:4	1.88	3.62	-0.23	3.28	1.09	-0.14	-	-	-	0.23 (0.58)	-0.17 (0.12)	-0.05 (0.15)
	M3 80:1-92:4	2.00	2.98	-0.15	3.15	0.68	0.17	-	-	-	0.13 (0.36)	-0.16 (0.12)	0.04 (0.09)
	M4 80:1-92:4	2.02	3.15	-0.19	3.15	0.65	0.12	-	-	-	0.10 (0.36)	-0.16 (0.11)	0.03 (0.08)
Italy	M1 67:1-92:4	5.01	6.09	0.18	6.21	2.41	0.42	≥0.23	≤-0.06	≥-0.09	0.77 (0.30)	-0.44 (0.23)	0.16 (0.11)
	M3 62:1-92:4	5.28	5.80	0.04	6.29	1.45	0.12	≥-.4,≤.9	≥-.6,≤.2	-	0.26 (0.22)	-0.18 (0.16)	0.03 (0.07)
UK	M0 71:2-86:4	5.43	5.30	0.18	4.45	1.54	0.77	≥0.25	≤-0.17	≥0.11	0.56 (0.19)	-0.46 (0.16)	0.22 (0.07)
	M1 59:1-86:4	5.41	11.9	-0.04	4.09	2.88	-0.41	≤1.40	≥-0.27	≤-0.10	0.45 (0.42)	-0.11 (0.07)	-0.29 (0.18)
	M3 65:1-86:4	5.32	7.49	0.06	4.00	2.61	0.35	≥-0.10	≤0.04	≥-0.33	0.60 (0.35)	-0.27 (0.16)	0.23 (0.20)
	M4 65:1-92:4	4.97	4.46	0.01	5.05	1.94	0.51	≥0.13	≤-0.21	≥0.00	0.52 (0.23)	-0.65 (0.25)	0.20 (0.11)
USA	M1 59:1-92:4	3.45	3.86	0.10	5.57	1.35	-0.31	≤0.73	≥-.5,≤.3	≤0.08	0.29 (0.20)	-0.14 (0.14)	-0.08 (0.07)
	M2 61:1-92:4	3.42	2.74	0.01	5.88	1.08	-0.42	≤0.60	≥-0.81	≤0.04	0.24 (0.15)	-0.37 (0.20)	-0.08 (0.05)
	M3 61:1-92:4	3.40	2.52	0.12	5.93	1.08	-0.07	≥-.1,≤.6	≥-.9,≤.3	-	0.22 (0.13)	-0.26 (0.22)	-0.01 (0.06)
Canada	M1 62:1-92:4	3.56	7.56	-0.02	7.40	1.55	0.16	≤1.95	≥-0.05	-	0.98 (0.36)	-0.22 (0.06)	-0.33 (0.07)
	M2 70:1-92:4	3.58	6.20	-0.01	6.84	2.17	0.51	-	-	-	-0.37 (0.33)	0.12 (0.09)	0.16 (0.10)
Japan	M1 59:1-92:4	4.38	8.26	0.05	13.3	2.26	0.35	≥-0.20	≤0.03	-	0.49 (0.26)	-0.11 (0.06)	0.06 (0.06)
	M2 59:1-92:4	4.30	4.28	0.02	14.0	1.72	0.51	≥0.05	≤-0.08	≥0.07	0.47 (0.19)	-0.47 (0.17)	0.06 (0.04)
	M2CD 62:1-92:4	2.47	2.99	0.11	4.20	1.09	-0.29	-	-	-	0.63 (0.37)	-0.36 (0.24)	-0.08 (0.18)

Table 4
The Long-run Vertical Phillips Curve in G7-Countries

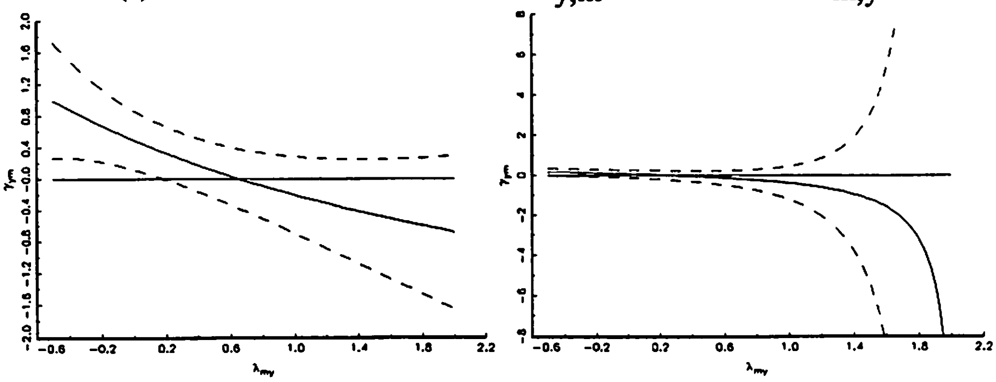
Country	Causality/ Period	VAR Estimates			Structural Model Estimates			$\gamma_{u\pi} / \gamma_{\pi u}$=0 in 95% confidence interval			Estimates imposing $\gamma_{u\pi} / \gamma_{\pi u}$=0		
		σ_u^2	σ_π^2	$cor_{u\pi}$	σ_u^2	σ_π^2	$cor_{u\pi}$	$\lambda_{\pi u}$	$\lambda_{u\pi}$	$\gamma_{\pi u} / \gamma_{u\pi}$	$\gamma_{\pi u}$	$\lambda_{u\pi}$	$\gamma_{\pi u} / \gamma_{u\pi}$
Germany	u←π 64:1-92:4	0.16	1.58	0.03	0.59	0.74	-0.59	≤6.37	≥-0.06	-	2.05 (2.06)	-0.02 (0.02)	-0.75 (0.35)
	π←u							≥3.69	≤-0.03	≤-0.04	10.7 (5.19)	-0.11 (0.06)	-0.47 (0.22)
France	u←π 70:1-92:4	0.15	1.83	0.21	0.33	1.04	-0.49	-	-	-	2.91 (2.76)	0.00 (0.02)	-1.56 (1.01)
	π←u							≥2.66	≤0.00	≤0.09	10.1 (4.33)	-0.06 (0.04)	-0.16 (0.10)
Italy	u←π 62:1-92:4	0.37	3.31	0.09	0.38	1.46	-0.40	≤1.33	≥0.00	≤0.79	-1.66 (1.62)	0.03 (0.02)	-1.54 (1.03)
	π←u							≥-2.08	≤0.04	≤0.05	2.04 (2.09)	-0.02 (0.03)	-0.10 (0.07)
UK	u←π 62:2-86:4	0.15	4.11	0.20	0.74	1.43	-0.57	-	-	-	15.1 (4.50)	-0.02 (0.01)	-1.09 (0.51)
	π←u							no	no	no	43.0 (20.1)	-0.08 (0.06)	-0.30 (0.14)
USA	u←π 62:1-92:4	0.25	1.64	-0.19	0.52	1.06	-0.62	≤5.95	≥-0.14	-	2.21 (1.65)	-0.08 (0.03)	-1.27 (0.47)
	π←u							≥5.38	≤-0.13	≤-0.10	13.7 (8.24)	-0.25 (0.11)	-0.30 (0.11)
Canada	u←π 62:1-92:4	0.30	1.82	0.05	0.64	0.97	-0.71	≤2.48	≥-0.05	≤-0.53	0.51 (1.08)	-0.01 (0.02)	-1.07 (0.32)
	π←u							≥3.41	≤-0.08	≥-0.23	6.87 (2.59)	-0.19 (0.08)	-0.47 (0.14)
Japan	u←π 62:1-92:4	0.09	4.21	-0.01	0.12	1.76	-0.57	-	-	-	-5.66 (8.61)	0.00 (0.00)	-8.55 (3.95)
	π←u							≥1.27	≤0.00	≤0.00	24.4 (14.8)	-0.01 (0.01)	-0.04 (0.02)

Table 5
The 'Fisher-Effect' in G7-Countries

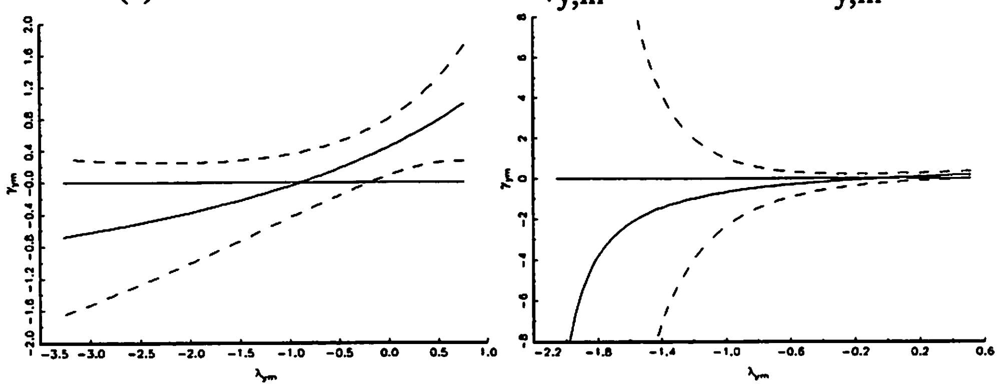
Country	Causality/ Period	VAR Estimates			Structural Model Estimates			$\gamma_{r\pi}=1$ in 95% confidence interval			Estimates imposing $\gamma_{r\pi}=1$		
		σ^2_r	σ^2_π	$cor_{r\pi}$	σ^2_r	σ^2_π	$cor_{r\pi}$	$\lambda_{\pi r}$	$\lambda_{r\pi}$	$\gamma_{\pi r}$	$\lambda_{\pi r}$	$\lambda_{r\pi}$	$\gamma_{\pi r}$
Germany	$r\leftarrow\pi$ 62:1-92:4	0.78	1.65	0.16	0.78	1.15	0.76	≥-2.4,≤.6	≥-.05,≤.5	-	-0.49 (0.56)	0.18 (0.10)	-0.12 (0.22)
France	$r\leftarrow\pi$ 72:1-92:4	0.78	1.84	0.31	1.08	0.98	0.79	≤-.3,≥1.7	≤-.3,≥.2	-	-2.89 (1.56)	0.41 (0.14)	-0.38 (0.90)
Italy	$r\leftarrow\pi$ 73:1-92:4	1.37	3.39	0.63	1.15	1.15	0.55	≤0.17	≥0.25	≤-0.28	-1.24 (1.41)	0.35 (0.07)	-2.99 (4.68)
UK	$r\leftarrow\pi$ 59:2-92:4	1.10	3.79	0.07	1.18	1.88	0.73	≥1.90	≤-0.15	≥1.58	17.1 (19.7)	-2.23 (3.74)	8.00 (11.9)
USA	$r\leftarrow\pi$ 59:1-92:4	0.75	1.72	0.23	0.94	1.14	0.70	≤-1.2,≥2.3	≤-.5,≥.4	≤-.3,≥1.1	-4.83 (3.47)	0.69 (0.26)	-4.23 (8.79)
Canada	$r\leftarrow\pi$ 59:1-92:4	0.96	2.10	0.24	1.16	1.05	0.68	≤-.5,≥2.1	≤-.5,≥.2	≤.3,≥.8	-2.01 (1.19)	0.43 (0.15)	-0.54 (0.77)
Japan	$r\leftarrow\pi$ 79:2-92:1	0.53	1.80	0.43	0.71	0.71	0.82	≤0.62	≥0.10	-	-2.32 (2.35)	0.25 (0.09)	-0.98 (2.40)

Figure 1: Neutrality Test for German Output and Money
(M1 left hand side, M2 right hand side)

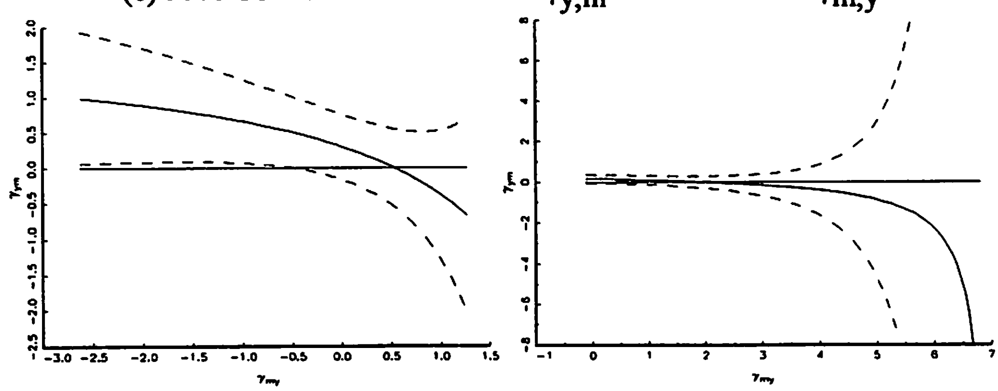
(a) 95% Confidence Intervals for $\gamma_{y,m}$ as a function of $\lambda_{m,y}$



(b) 95% Confidence Intervals for $\gamma_{y,m}$ as a function of $\lambda_{y,m}$



(c) 95% Confidence Intervals for $\gamma_{y,m}$ as a function of $\gamma_{m,y}$



(a) 95 % Confidence Elipses for $\lambda_{y,m}$ and $\lambda_{m,y}$ imposing $\gamma_{y,m}=0$

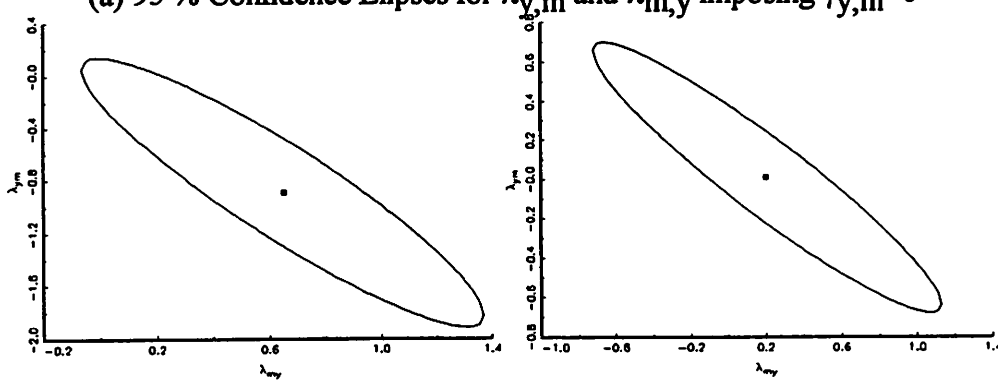
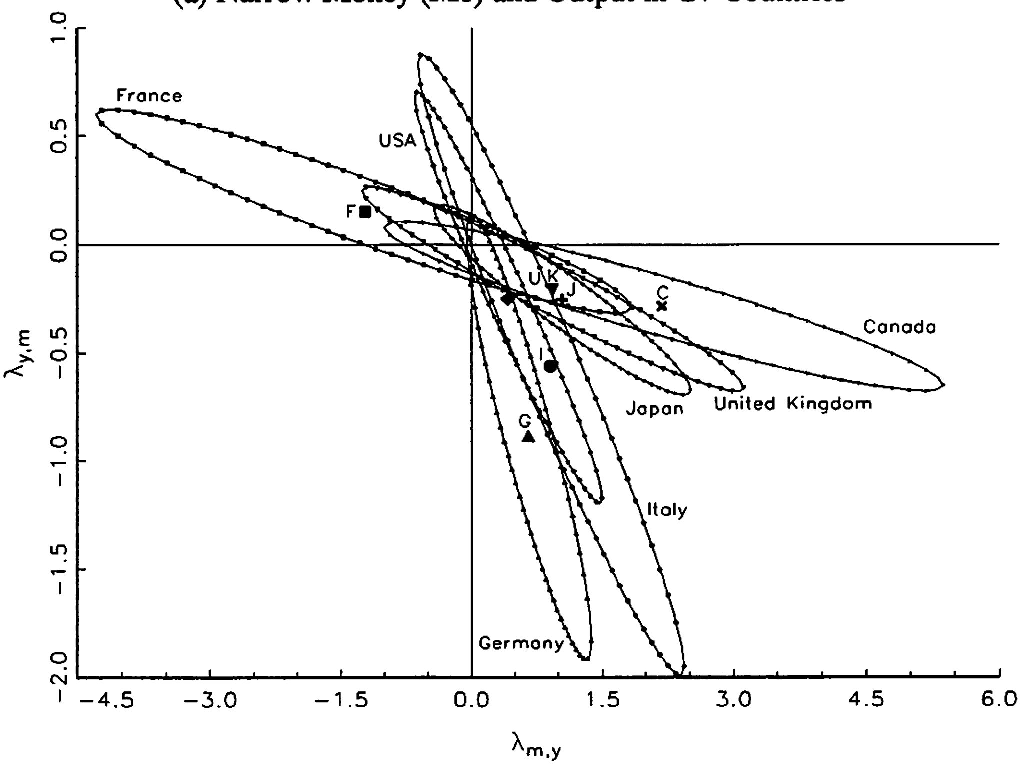


Figure 2: Estimated Confidence Elipses for λ_{ym} and λ_{my} under the Nullhypothesis of the Long-run Neutrality of Money (γ_{ym}=0)

(a) Narrow Money (M1) and Output in G7 Countries



(b) Broad Money (M2, M3) and Output in G7 Countries

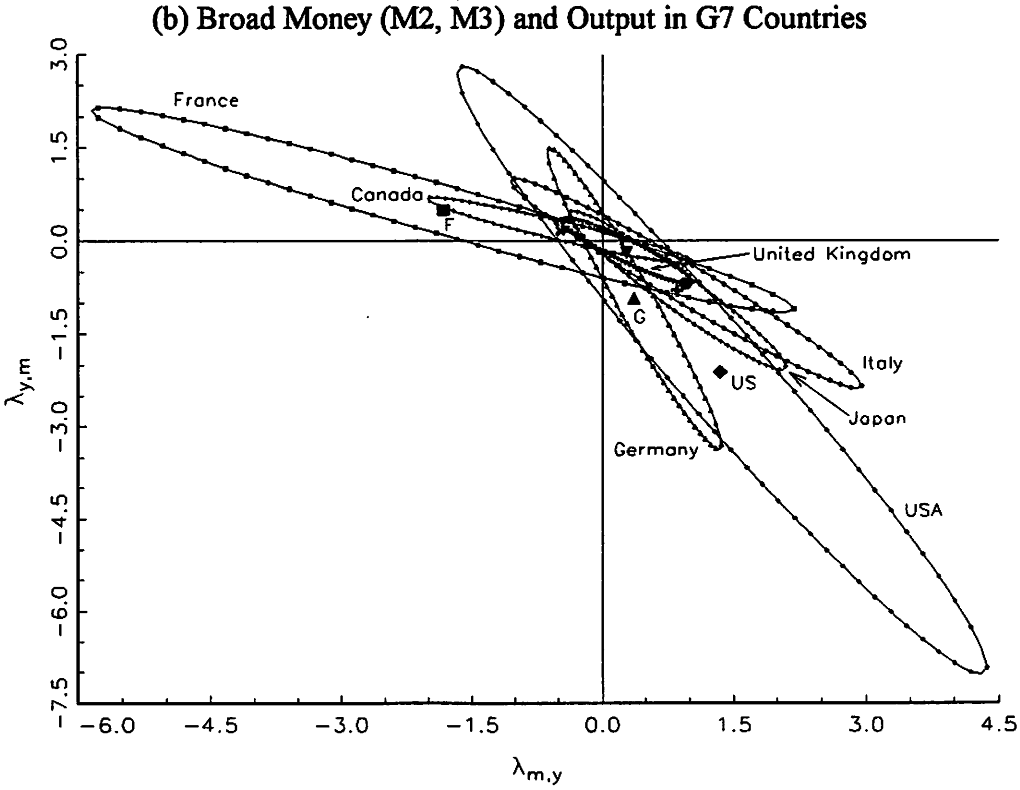
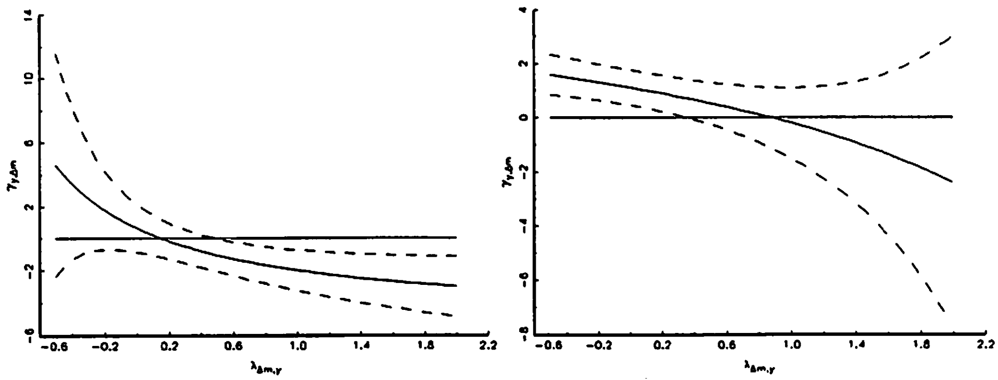
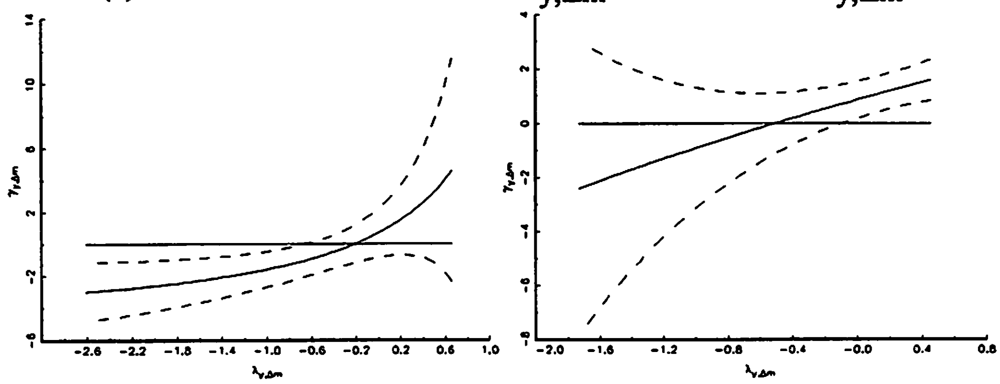


Figure 3: Superneutrality Test for German Output and Money
(M1 left hand side, M2 right hand side)

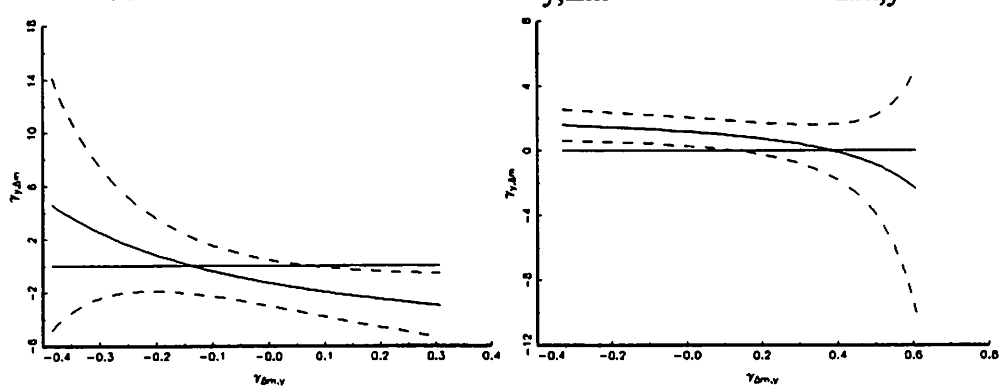
(a) 95% Confidence Intervals for $\gamma_{y,\Delta m}$ as a function of $\lambda_{\Delta m,y}$



(b) 95% Confidence Intervals for $\gamma_{y,\Delta m}$ as a function of $\lambda_{y,\Delta m}$



(c) 95% Confidence Intervals for $\gamma_{y,\Delta m}$ as a function of $\gamma_{\Delta m,y}$



(d) 95 % Confidence Elipses for $\lambda_{y,\Delta m}$ and $\lambda_{\Delta m,y}$ imposing $\gamma_{y,\Delta m}$=0

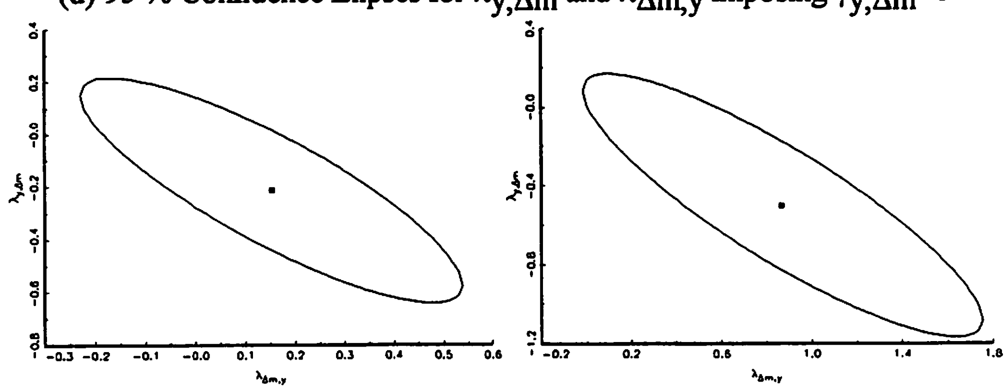


Figure 4: Test for a German Long-run Vertical Phillips Curve with Causation Running from Inflation to Unemployment

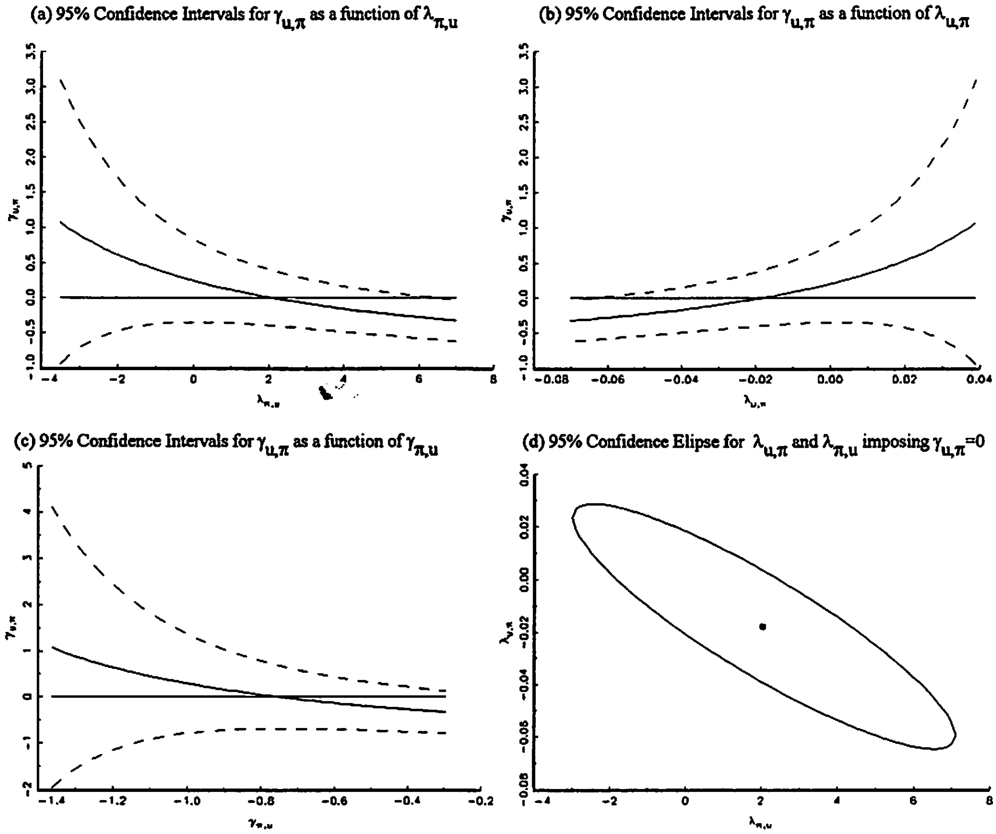


Figure 5: Test for a German Long-run Vertical Phillips Curve with Causation Running from Unemployment to Inflation

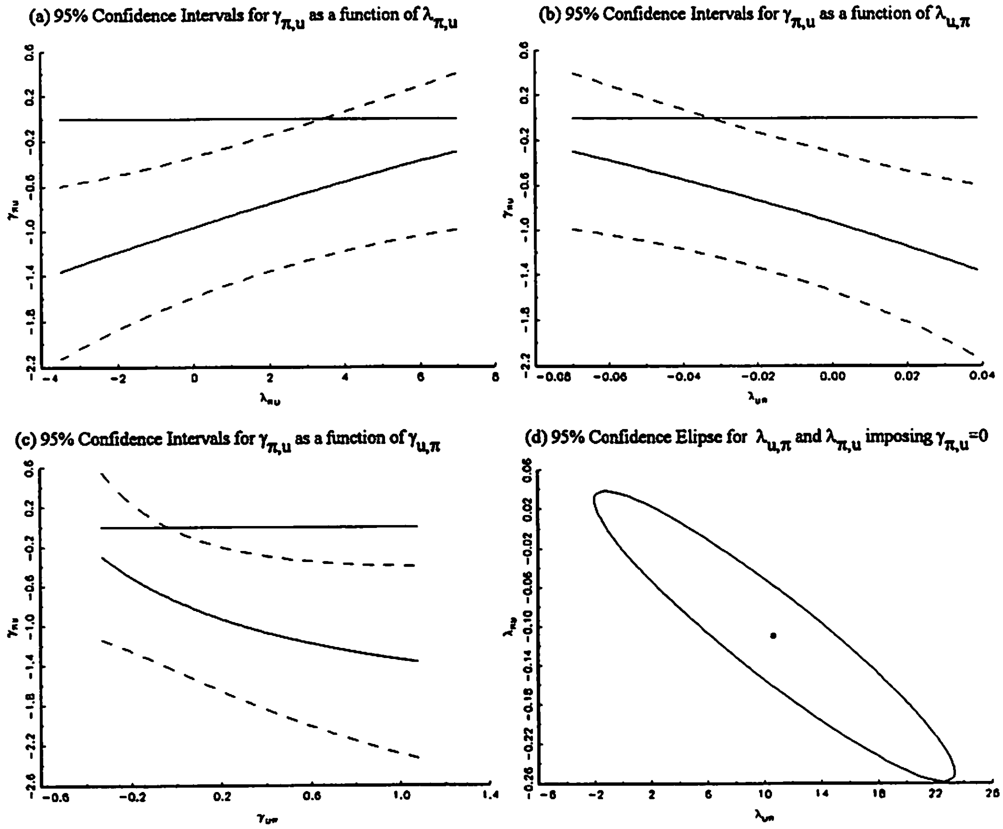


Figure 6: Test for a Long-run 'Fisher-Effect' of Inflation on Nominal Interest Rates in Germany

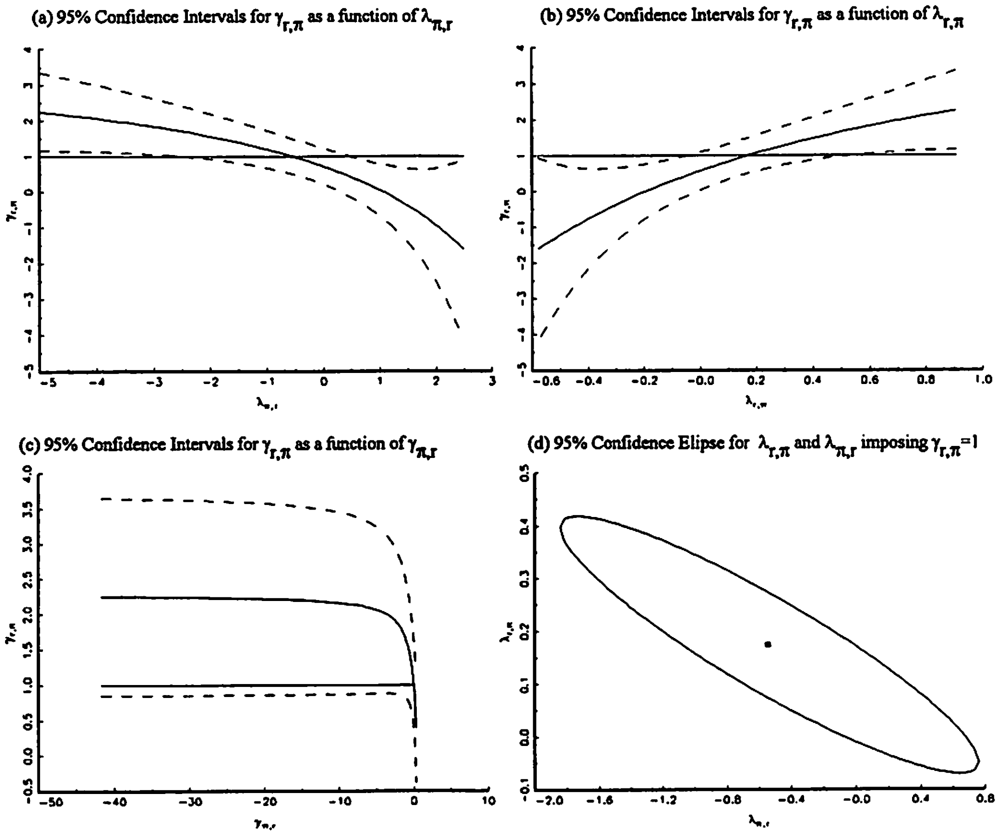
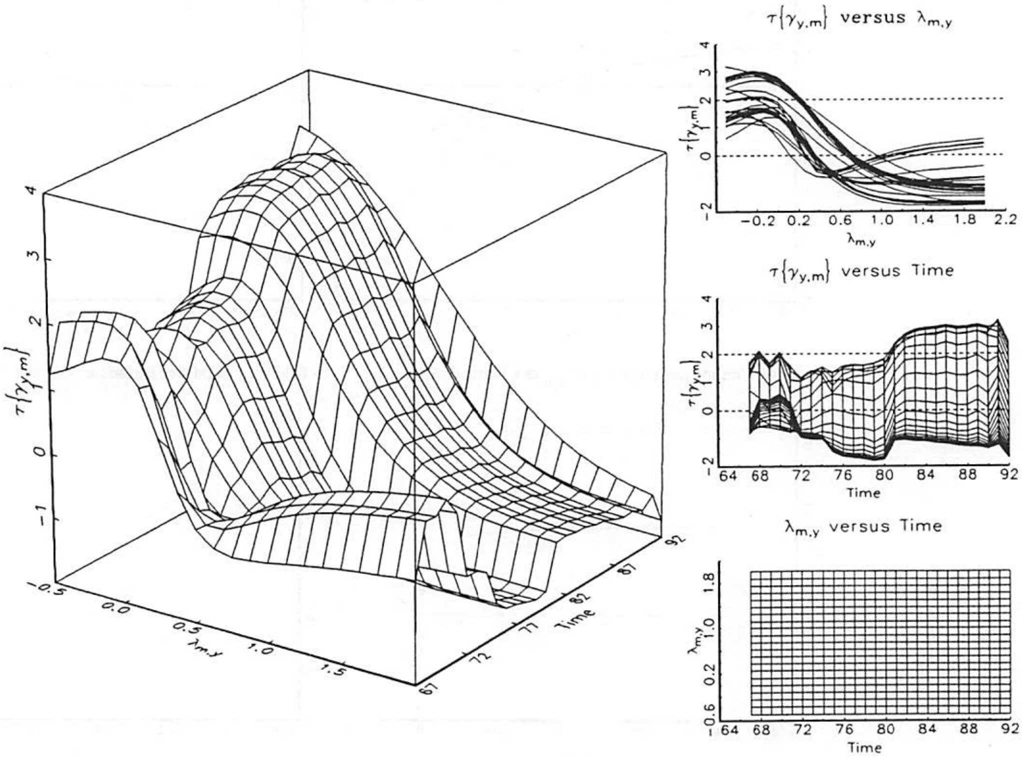


Figure 7: Time-profile of Neutrality Tests for the German Monetary Aggregate M1 and Output

(a) T-test for $\gamma_{y,m} \neq 0$ as a function of $\lambda_{m,y}$ and time



(b) T-test for $\gamma_{y,m} \neq 0$ as a function of $\lambda_{y,m}$ and time

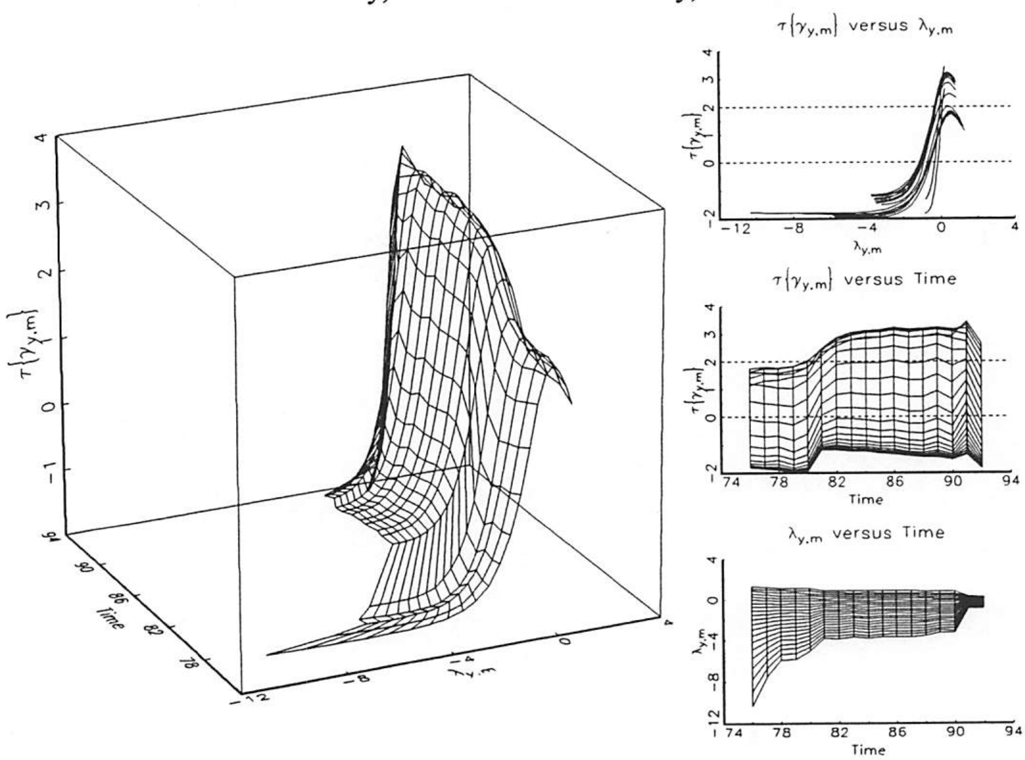
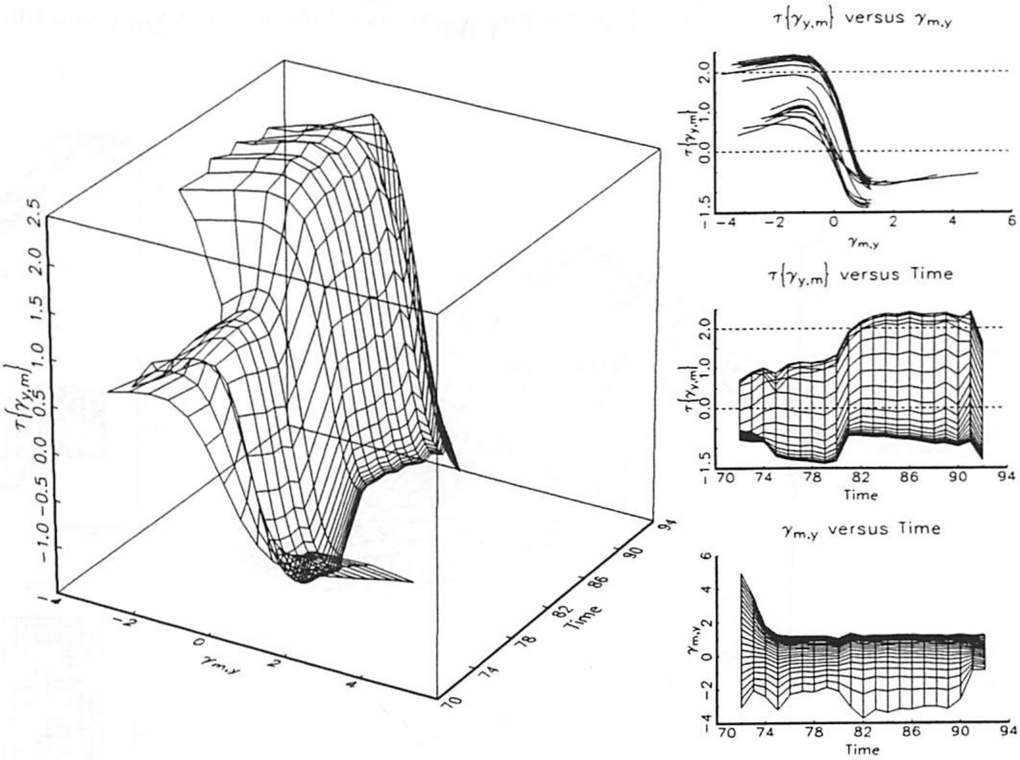


Figure 7 continued

(c) T-test for $\gamma_{y,m} \neq 0$ as a function of $\gamma_{m,y}$ and time



(d) Confidence Ellipses for $\lambda_{m,y}$ and $\lambda_{y,m}$ when $\gamma_{y,m}=0$

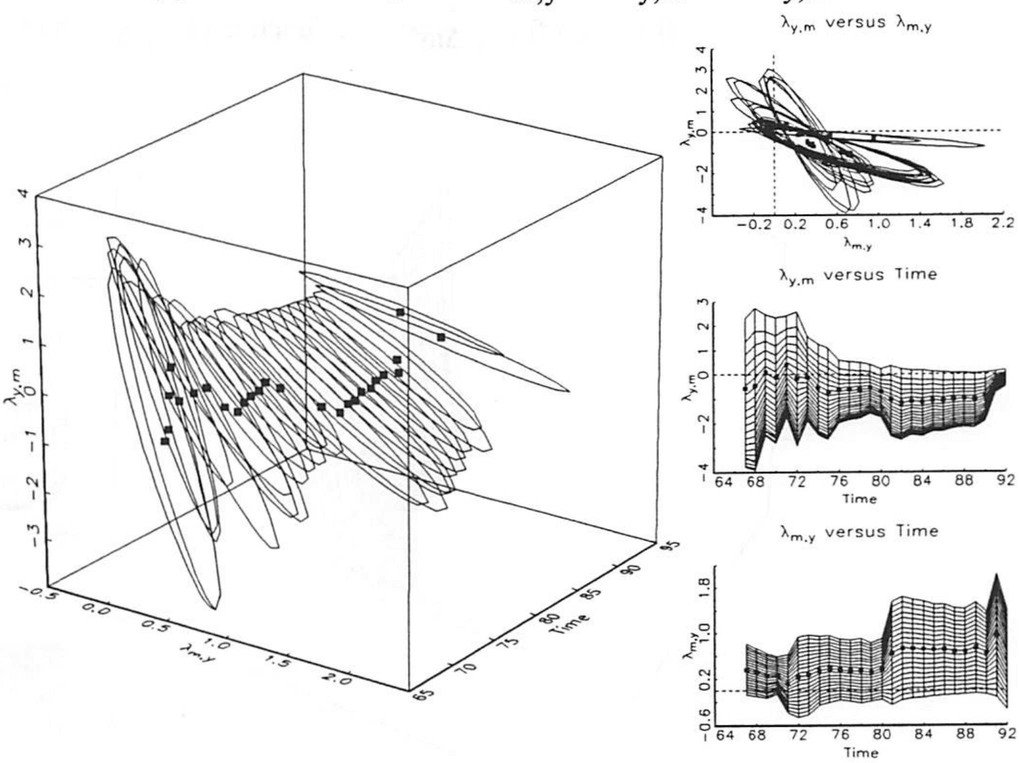
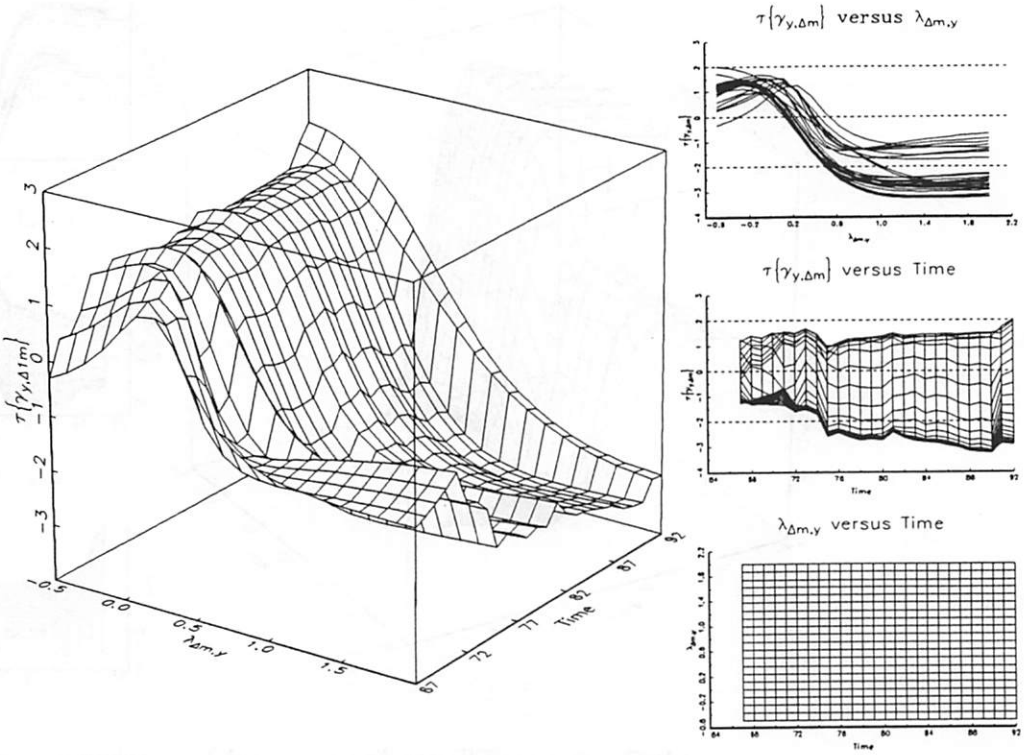


Figure 8: Time-profile of Superneutrality Tests for the German Monetary Aggregate M1 and Output

(a) T-test for $\gamma_{y,\Delta m} \neq 0$ as a function of $\lambda_{\Delta m,y}$ and time



(b) T-test for $\gamma_{y,\Delta m} \neq 0$ as a function of $\lambda_{y,\Delta m}$ and time

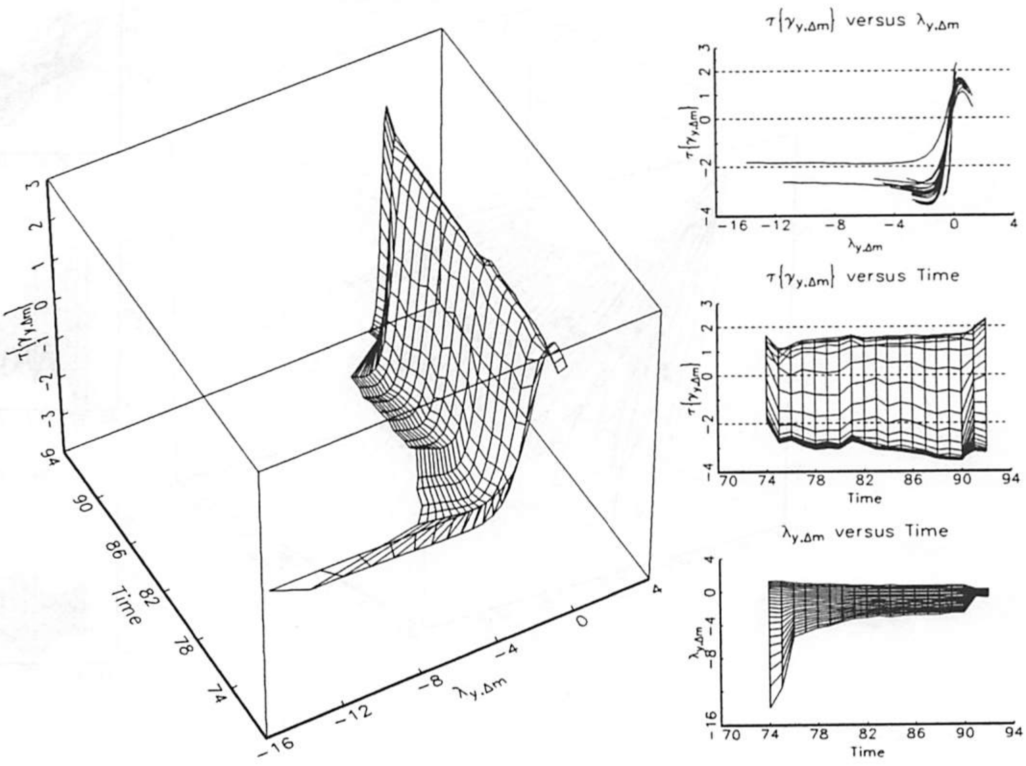
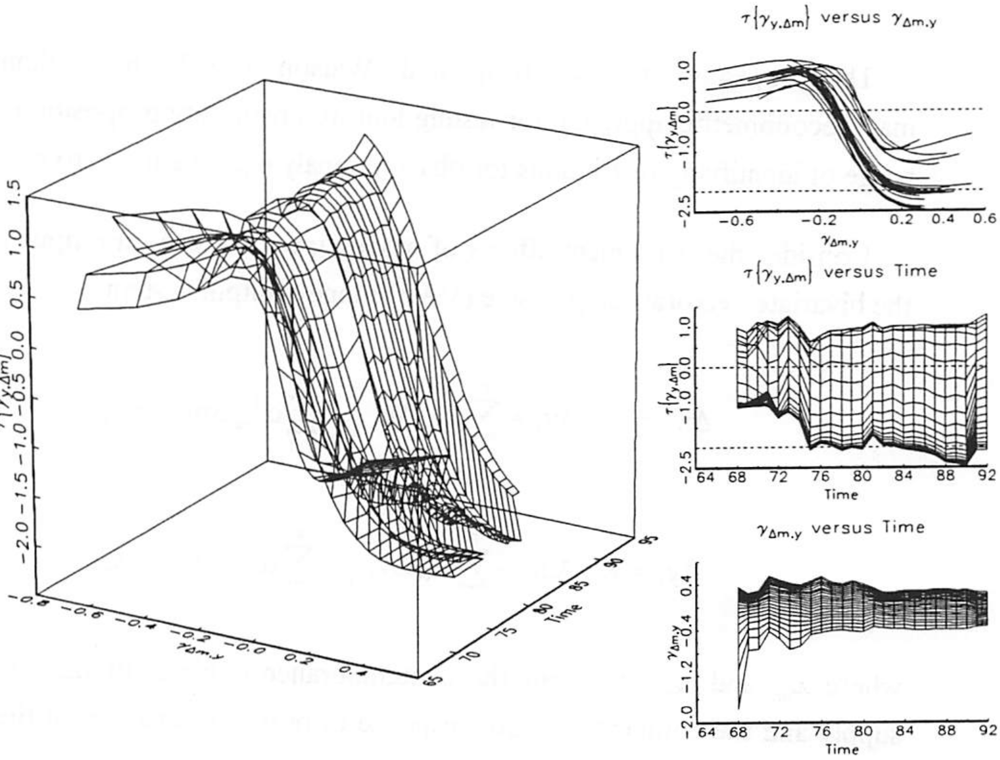
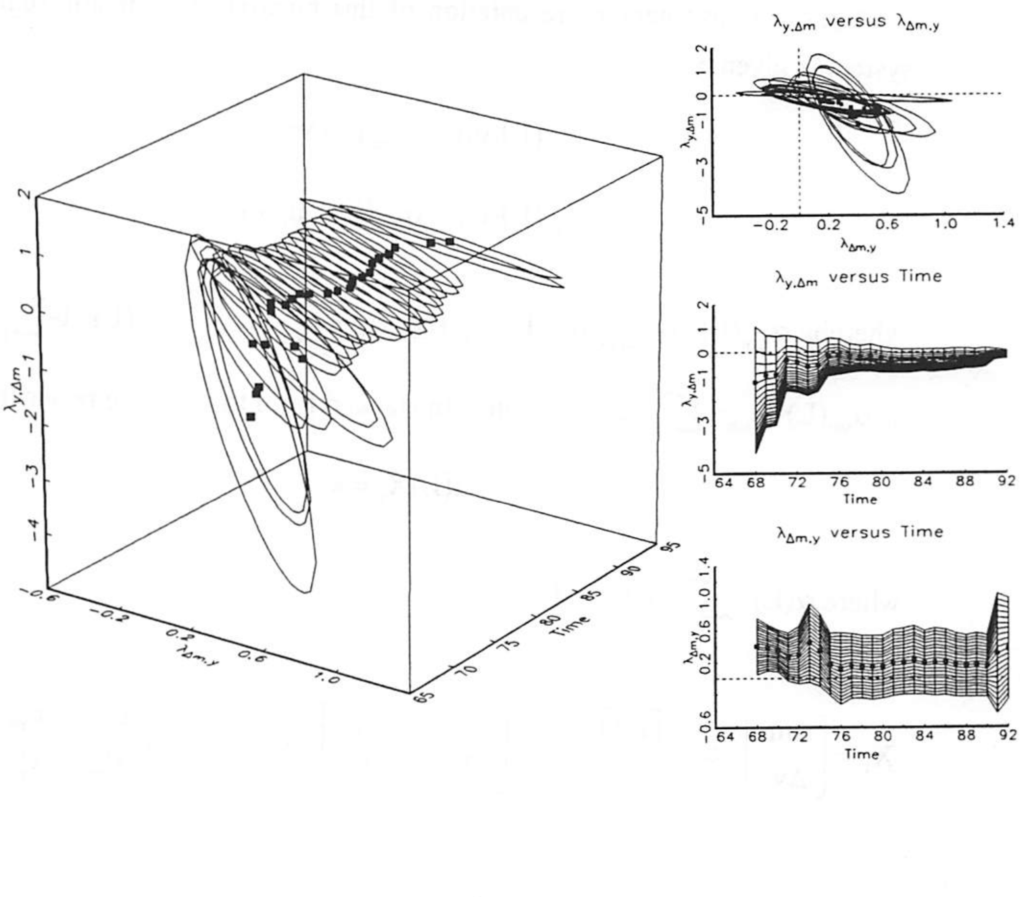


Figure 8 continued

(c) T-test for $\gamma_{y,\Delta m} \neq 0$ as a function of $\gamma_{\Delta m,y}$ and time



(d) Confidence Ellipses for $\lambda_{\Delta m,y}$ and $\lambda_{y,\Delta m}$ when $\gamma_{y,\Delta m}=0$



Technical Appendix

This appendix follows King and Watson (1992) in outlining the basic macroeconometric approach for testing long-run neutrality propositions within a wide range of identifying restrictions for observationaly equivalent macro models.

Consider the permanent effects of money growth (Δm) on output growth (Δy) in the bivariate vectorautoregressive (VAR) money-output system:

$$\Delta m_t = \lambda_{my}\Delta y_t + \sum_{j=1}^{p}\alpha_{my}^{j}\Delta y_{t-j} + \sum_{j=1}^{p}\alpha_{mm}^{j}\Delta m_{t-j} + \varepsilon_t^m, \tag{A1}$$

$$\Delta y_t = \lambda_{ym}\Delta m_t + \sum_{j=1}^{p}\alpha_{yy}^{j}\Delta y_{t-j} + \sum_{j=1}^{p}\alpha_{ym}^{j}\Delta m_{t-j} + \varepsilon_t^\eta, \tag{A2}$$

where λ_{my} and λ_{ym} represent the contemporaneous effect of output on the money supply and the contemporaneous response of output to changes in the money supply respectively.

A more convenient representation of this bivariate vector autoregressive (VAR) system is given by:

$$\alpha_{mm}(L)\Delta m_t = \alpha_{my}(L)\Delta y_t + \varepsilon_t^m, \tag{A3a}$$

$$\alpha_{yy}(L)\Delta y_t = \alpha_{ym}(L)\Delta m_t + \varepsilon_t^\eta, \tag{A3b}$$

whereby $\alpha_{mm}(L)=1-\sum_{j=1}^{p}\alpha_{mm}^{j}L^j$, $\alpha_{my}(L)=\lambda_{my}+\sum_{j=1}^{p}\alpha_{my}^{j}L^j$, $\alpha_{yy}(L)=1-\sum_{j=1}^{p}\alpha_{yy}^{j}L^j$ as well as $\alpha_{ym}(L)=\lambda_{ym}+\sum_{j=1}^{p}\alpha_{ym}^{j}L^j$ applies. In stacked form this may be re-written as:

$$\alpha(L)X_t = \varepsilon_t, \tag{A4}$$

where $\alpha(L)=\sum_{j=0}^{p}\alpha^j L^j$ and

$$X_t = \begin{bmatrix}\Delta m_t \\ \Delta y_t\end{bmatrix},\ \varepsilon_t = \begin{bmatrix}\varepsilon_t^m \\ \varepsilon_t^\eta\end{bmatrix},\ \alpha^0 = \begin{bmatrix}1 & -\lambda_{my} \\ -\lambda_{ym} & 1\end{bmatrix},\ \text{and}\ \alpha^j = -\begin{bmatrix}\alpha_{mm}^j & \alpha_{my}^j \\ \alpha_{ym}^j & \alpha_{yy}^j\end{bmatrix},\quad j=1,2,\ldots,p.$$

In this notation the long-run multipliers are $\gamma_{my}=\alpha_{my}(1)/\alpha_{mm}(1)$ and $\gamma_{ym}=\alpha_{ym}(1)/\alpha_{yy}(1)$, whereby γ_{my} measures the long-run response of money *m* to a one unit permanent increase in output *y*, whilst γ_{ym} measures the long-run response of output *y* to a permanent unit increase in money *m*. Long-run neutrality of money thereby implies the restriction $\gamma_{ym}=0$.

A1. Identification

As noted by Watson and King (1992), equation (4) is econometrically unidentified and the neutrality restriction is no longer testable when money is endogenous. Thus, even if we maintain the hypothesis that ε_t^m and ε_t^η are uncorrelated, one additional restriction is required in order to identify the linear simultaneous equation model. In the literature various identifying restrictions are to be found. Common practice in the older literature on long-run neutrality testing is to assume that money is exogenous, so that $\gamma_{my}=(\lambda_{my}+\Sigma_{j=1}^{p}\alpha_{my}^{j})/(1-\Sigma_{j=1}^{p}\alpha_{mm}^{j})=0$, which holds for $\lambda_{my}=\alpha_{my}^{1}=\alpha_{my}^{2}=...=\alpha_{my}^{p}=0$. A less restrictive approach is to simply assume that the model is recursive, so that either money or output are predetermined, that is $\lambda_{my}=0$ or $\lambda_{ym}=0$ holds.[1] Finally, long-run neutrality $\gamma_{ym}=(\lambda_{ym}+\Sigma_{j=1}^{p}\alpha_{ym}^{j})/(1-\Sigma_{j=1}^{p}\alpha_{yy}^{j})=0$ may be assumed in order to identify the system and estimate the remaining parameters. Thus, it is in principle possible to identify the above model by specifying a value of any one of the four parameters λ_{my}, λ_{ym}, γ_{my} or γ_{ym} and find the implied estimates for the other three parameters. This is in fact the approach taken in King and Watson (1992), but rather than focussing on a single identifying restriction, the authors report results for a wide range of identifying restrictions which imply observationally equivalent estimated models.

A2. Estimation

Under each identifying restriction Gaussian maximum likelihood estimates are constructed using standard ordinary least squares (OLS) and instrumental variable (IV) regression, as will be discussed in this section.

[1] Geweke (1986), Stock and Watson (1988) and Fisher and Seater (1993) present tests for neutrality under the assumption that $\lambda_{ym}=0$, and Gewecke (1986) also presents results under the restriction that $\lambda_{my}=0$.

A2.1. Estimation when λ_{my}, and λ_{ym} are known.

In the case of λ_{my} being known, equation (A1) can be estimated directly by restricted least squares and regressing $\Delta m_t-\lambda_{my}\Delta y_t$ onto $\Delta m_{t-1},\Delta m_{t-2},\ldots,\Delta m_{t-p},\Delta y_{t-1},\Delta y_{t-2}$, ..., Δy_{t-p}. Equation (A2) can not be estimated by ordinary least squares because one of the regressors, Δm_t, is potentially correlated with the residuals ε_t^{η}. In order to account for this problem instrumental variables must be used, with the appropriate set of instruments being given by Δm_{t-1}, $\Delta m_{t-2},\ldots$, Δm_{t-p}, Δy_{t-1}, $\Delta y_{t-2},\ldots$, Δy_{t-p} together with the estimated residuals from equation (A1). These residuals are a valid instrument because of the assumption that ε_t^{m} and ε_t^{η} are uncorrelated.

In the case of λ_{ym} assumed known the above procedure is reversed, equation (A2) is estimated by regressing $\Delta y_t-\lambda_{ym}\Delta m_t$ onto Δm_{t-1}, $\Delta m_{t-2},\ldots$, Δm_{t-p}, Δy_{t-1}, $\Delta y_{t-2},\ldots$, Δy_{t-p}, whilst equation (A1) is estimated by instrumental variables with the appropriate set of instruments being Δm_{t-1}, $\Delta m_{t-2},\ldots$, Δm_{t-p}, Δy_{t-1}, $\Delta y_{t-2},\ldots$, Δy_{t-p} together with the estimated residuals from equation (2).

In both cases a complication arises in the calculation of the standard errors of the long-run multipliers, because, rather than being estimated directly, these long-run multipliers are nonlinear functions of the estimated regression coefficients of the model. Hence, their standard errors are calculated by using delta method arguments, based on a Taylor approximation of the standard errors around their expected values.

A2.2. Estimation when γ_{my} or γ_{ym} are known.

When instead of the impact elasticities λ_{my} or λ_{ym} one of the long-run elasticities.γ_{my} or γ_{ym}. is assumed known, a similar but modified estimation procedure can be adopted.

When a value of $\gamma_{my}=\alpha_{my}(1)/\alpha_{mm}(1)=(\lambda_{my}+\Sigma_{j=1}^{p}\alpha_{my}^{j})/(1-\Sigma_{j=1}^{p}\alpha_{mm}^{j})$ is used to identify the model, it is convenient to rewrite equation (A1) as:

$$\begin{aligned}\Delta m_t &= \left(\lambda_{my}-\sum_{j=1}^{p}\alpha_{my}^{j}\right)\Delta y_t+\sum_{j=1}^{p}\alpha_{mm}^{j}\Delta m_{t-1}+\sum_{j=0}^{p-1}\tilde{\alpha}_{my}^{j}\Delta^2 y_{t-j}+\sum_{j=1}^{p-1}\tilde{\alpha}_{mm}^{j}\Delta^2 m_{t-j}+\varepsilon_t^{m}\\ &=\alpha_{my}(1)\Delta y_t+\beta_{mm}\Delta m_{t-1}+\sum_{j=0}^{p-1}\tilde{\alpha}_{my}^{j}\Delta^2 y_{t-j}+\sum_{j=1}^{p-1}\tilde{\alpha}_{mm}^{j}\Delta^2 m_{t-j}+\varepsilon_t^{m}\end{aligned} \tag{A5}$$

whereby the regression coefficients $\tilde{\alpha}^j_{my}$ and $\tilde{\alpha}^j_{mm}$ in equation (A5) may be expressed in terms of the coefficients from equation (A1) as $\tilde{\alpha}^j_{my}=-\Sigma^p_{i=j+1}\alpha^i_{my}$ $\forall$ j=0,1,2,...,p-1 and $\tilde{\alpha}^j_{mm}=-\Sigma^p_{i=j+1}\alpha^i_{mm}$ $\forall$ j=1,2,...,p-1. Since the long-run multiplier γ_{my} may be expressed as $\gamma_{my}=\alpha_{my}(1)/(1-\beta_{mm})$, which implies $\alpha_{my}(1)=\gamma_{my}-\beta_{mm}\gamma_{my}$, equation (A5) may be modified to:

$$\Delta m_t - \gamma_{my}\Delta y_t = \beta_{mm}\left(\Delta m_{t-1} - \gamma_{my}\Delta y_t\right) + \sum_{j=0}^{p-1}\tilde{\alpha}^j_{my}\Delta^2 y_{t-j} + \sum_{j=1}^{p-1}\tilde{\alpha}^j_{mm}\Delta^2 m_{t-j} + \varepsilon^m_t \tag{A6}$$

which due to the potential correlation between Δy_t and the error term ε^m_t can be estimated by instrumental variables using Δm_{t-1}, Δm_{t-2},..., Δm_{t-p}, Δy_{t-1}, Δy_{t-2},..., Δy_{t-p} as instruments. Equation (A2) is then estimated by instrumental variables with the appropriate set of instruments being Δm_{t-1}, Δm_{t-2},..., Δm_{t-p}, Δy_{t-1}, Δy_{t-2},..., Δy_{t-p} together with the estimated residuals from equation (A6). In case the model was identified by setting $\gamma_{ym}=\alpha_{ym}(1)/\alpha_{yy}(1)=(\lambda_{my}+\Sigma^p_{j=1}\alpha^j_{my})/(1-\Sigma^p_{j=1}\alpha^j_{mm})$, this procedure was reversed.

As above, the calculation of the standard errors of the long-run multipliers are calculated by using standard formula from delta method arguments. A second complication similar to Pagan's (1984) "generated regressor problem" arises because one of the equations is estimated using instruments that are the residuals from another equation. To understand the problem, stack the observations from each equation as:

$$\Delta M = \Delta X_1\delta_1 + \varepsilon_1, \tag{A7a}$$

$$\Delta Y = \Delta X_2\delta_2 + \varepsilon_2, \tag{A7b}$$

where M and Y are the Tx1 vectors of endogenous variables and X_1 and X_2 are the Tx(2*p+1) matrices of regressors in equations (A6) and (A2). Denote the matrix of instruments for the first equation by Z (consisting of Δm_{t-1},,...,Δm_{t-p}, Δy_{t-1},,..., Δy_{t-p}) and the instruments of the second equation by $\hat{U}=[\hat{\varepsilon}_1\, Z]$, and let $U=[\varepsilon_1\, Z]$. Since $\hat{\varepsilon}_1 = \varepsilon_1 - X_1(\hat{\delta}_1 - \delta_1)$ it is true that $\hat{U} = U - [X_1(\hat{\delta}_1 - \delta_1)\ 0]$. Now let the asymptodic covariance matrix be $V_1 = \sigma^2_{\varepsilon_1}\,\mathrm{plim}[T(Z'X_1)^{-1}(Z'Z)(X_1'Z)]$, then one may write:

$$\begin{aligned} T^{1/2}\left(\hat{\delta}_2 - \delta_2\right) &= \left(T^{-1}\hat{U}'X_2\right)^{-1}\left(T^{-1/2}\hat{U}'\varepsilon_2\right) \\ &= \left(T^{-1}\hat{U}'X_2\right)^{-1}\left(T^{-1/2}U'\varepsilon_2\right) - \left(T^{-1}\hat{U}'X_2\right)^{-1}\begin{bmatrix} T^{1/2}\left(\hat{\delta}_1 - \delta_1\right)'\left(T^{-1}X_1'\varepsilon_2\right) \\ 0 \end{bmatrix} \end{aligned} \tag{A8}$$

Potential problems arise due to the second term on the right hand side of (A8). Since $T^{1/2}(\hat{\delta}_1 - \delta_1)$ converges in distribution, the second term can only be disregarded asymptotically when $\mathrm{plim} T^{-1} X_1' \varepsilon_1 = 0$, that is, when the regressors in (A7a) are uncorrelated with the error terms in (A7b). When the long-run elasticities γ_{my} or γ_{ym} are assumed known, X_1 will contain the contemporaneous value of Δy_t.and this condition is violated. The necessary modification consists of an adjustment of the covariance matrix of δ_2 to account for the second term. It can thereby be shown that $T^{1/2}(\hat{\delta}_1 - \delta_1)$ converges to a random variable with $N(0,V_2)$ distribution, where $V_1 = \sigma^2_{\varepsilon_2}\, \mathrm{p\,lim}[T(\hat{U}'X_2)^{-1}(\hat{U}'\hat{U})(X_2'\hat{U})^{-1}] + \mathrm{p\,lim}[T(\hat{U}'X_2)^{-1} D (X_2'\hat{U})^{-1}]$ with D as a matrix with all elements equal to zero except that $D_{11} = (\varepsilon_2' X_1) TV_1 (X_1' \varepsilon_2)$, where $TV_1 = \sigma^2_{\varepsilon_1}(Z'X_1)^{-1}(Z'Z)(X_1'Z)^{-1}$.

A3. Iteration

The testing strategy in King and Watson (1992) now consists of providing information on the link between the individual behavioral parameters (λ_{my}, λ_{ym} and γ_{my}) and the long run neutrality parameter (γ_{ym}). For this purpose King and Watson (1992) iterate each of the four parameters (λ_{my}, λ_{ym}, γ_{my} and γ_{ym}) within a reasonable range and each time obtaining estimates of the remaining three parameters and their standard errors. Graphs of the neutrality coefficient estimates (γ_{ym}) and standard error bands against each of the other behavioral parameters (λ_{my}, λ_{ym} and γ_{my}) are then used to show for which values of the behavioural parameters the neutrality proposition is rejected at the five percent significance level. The present paper follows this principle testing stragegy. Due to space limitations, the results for the non-German G7 countries are presented in tables rather than graphs.

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3–89 Rüdiger Pethig, Calculus of Consent: A Game–theoretic Perspective. Comment

4–89 Rüdiger Pethig, Problems of Irreversibility in the Control of Persistent Pollutants

5–90 Klaus Schöler, On Credit Supply of PLS–Banks

6–90 Rüdiger Pethig, Optimal Pollution Control, Irreversibilities, and the Value of Future Information

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