International Evidence on Monetary Neutrality Under Broken Trend Stationary Models

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Abstract

We analyze the issue of the impact of multiple breaks on monetary neutrality results, using a long annual international data set. We empirically verify whether neutrality propositions remain addressable (and if so, whether they hold or not), when unit root tests are carried out allowing for multiple structural breaks in the long-run trend function of the variables. It is found that conclusions on neutrality are sensitive to the number of breaks allowed. In order to interpret the evidence for structural breaks, we utilize a notion of deterministic monetary neutrality, which naturally arises in the absence of permanent stochastic shocks to the variables. We utilize a resampling procedure to discriminate between DS and TS models with multiple breaks, based on the fact that changes in the trend function bias unit root tests towards a non-rejection, and a dynamic programming algorithm for locating breaks. We present evidence on the inability to reject plausible broken-trend stationary models that exhibit transitory dynamics around a long-run deterministic trend subject to infrequent structural breaks. This leads to interesting questions about the testing for monetary neutrality.

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1 Introduction

Economists care about long-run monetary neutrality (LRN) because most theoretical models of money predict that money is neutral in the long-run; that is, the real effects of an unanticipated, permanent change in the level of money, tend to disappear as time elapses[2]. They also care about LRN because LRN is often used as an identification assumption (i.e. the large literature using Blanchard-Quah (1989) decompositions). On the other hand, the case for monetary superneutrality has limited theoretical support[3]. As summarized by Bullard (1999), "if monetary growth causes inflation, and inflation has distortionary effects, then long-run monetary superneutrality should not hold in the data. On the contrary, a permanent shock to the rate of monetary growth should have some long-run effect on the real economy; why else should we worry about it?" (p.59). In fact, central banks around the world pursue long-run price stability, due to the distortionary effects of inflation, caused by monetary growth (see the *Federal Reserve Bank of Kansas City Symposium* "Achieving Price Stability" (1996)).

Empirical results based on the reduced-form tests of Long Run Neutrality and Long Run Superneutrality (LRSN), derived by Fisher and Seatler (1993) (henceforth, FS), depend on the order of integration of both real output and the money aggregates. A number of recent papers examine the validity of these key macro propositions using long annual data and the reduced form tests of FS[4]. In this literature, the orders of integration are identified through the application of common used test, as the Augmented Dickey-Fuller, *ADF*, of Said and Dickey (1984), the *Z* tests of Phillips-Perron (1988), and the stationarity *KPSS* tests of Kwiatkowski, et. al. (1992), to real output and the money aggregates. For instance, LRN finds empirical support in the studies of Boschen and Otrok (1994, US data), Haug and Lucas (1997, Canadian data), Serletis and Krause (1996, international data set), Wallace (1999, Mexican data), and Bae and Ratti (2000 Brazilian and Argentinean data). Utilizing the more powerful tests of Ng and Perron (2001), Noriega (2004, international data set) finds weaker support for LRN.

Several empirical studies demonstrate the prevalence of (infrequent) parameter variation in the trend function of time series models of macroeconomic variables, as well as the impact of such structural breaks on unit root testing[5]. Serletis and Koustas (1998) argue that the issue of whether neutrality results

[2]Gottardi (1994) shows, on the other hand, that the phenomenon of non neutrality is associated with the effects of monetary policy on the assets' payoffs, due to incomplete markets.

[3]In the literature on monetary growth theory, there are very few available models which embody some form of monetary superneutrality. See for instance, Sidrauski (1967), Hayakawa (1995) and Faria (2001).

[4]Bae and Jensen (1999) examine these propositions by extending FS long-run neutrality requirements to long-memory processes. An alternative econometric perspective of LRN and LRSN is presented in King and Watson (1997).

[5]Empirical examples with macro time series can be found in Perron (1989, 1992, 1997), Lumsdaine and Papell (1997), Ohara (1999), Mehl (2000), Noriega and De Alba (2001), and Gil-Alana (2002); Perron and Vogelsang (1992), Culver and Papell (1995), and Aggarwal et.al. (2000) for real exchange rates; Raj (1992) and Zelhorst and Haan (1995), for real output; Clemente, et. al. (1998) for interest rates.

hold under the presence of structural breaks -an issue that has not been resolved yet in the literature- depends on how big shocks are treated. If they are treated like any other shock, then there is no need to account for them in interpreting neutrality results. If, on the other hand, they are regarded as (infrequent) big shocks that need to be accounted for, then conclusions on neutrality may change, because such shocks may induce lower orders of integration for output and money. FS use the convention that if a variable is stationary around a *linear* trend then it is treated as trend-stationary, that is, integrated of order zero. Extending FS's idea, one can say that if a variable is stationary around a *broken* trend then it is also integrated of order zero. This is precisely the interpretation followed by Serletis and Krause (1996), and Serletis and Koustas (1998). Under their approach, however, the number of structural breaks allowed in the deterministic trend function is fixed to one. This selection may not be inconsequential[6]. Furthermore, there are several recent methods that allow the estimation of the number of breaks using sample information.

In this paper, we analyze the issue of the impact of multiple breaks on neutrality results, extending Noriega's (2004) result in one important direction. By allowing for broken trend functions, we uncover the presence of structural breaks, which alter (reduce) the order of integration of money and output, therefore modifying conclusions on LRN and LRSN. We utilize the same data set as Noriega (2004), i.e., long annual data on real output and monetary aggregates for Argentina (1884-1996), Australia (1870-1997), Brazil (1912-1995), Canada (1870-2001), Italy (1870-1997), Mexico (1932-2000), Sweden (1871-1988), and the UK (1871-2000) (The cases of Denmark and the U.S. are not analyzed in this paper, since Noriega (2004) showed that money and output are integrated of order zero for both countries).

In particular, we empirically verify whether the monetary neutrality propositions remain addressable (and if so, whether they hold or not), when unit root tests are carried out allowing for (possibly) multiple structural breaks in the long-run trend function of the variables. It is found that conclusions on monetary neutrality are sensitive not only to whether there is a break or not, but also to the *number* of breaks allowed. In order to interpret the evidence for structural breaks, we utilize a notion of deterministic monetary neutrality, which naturally arises in the absence of permanent stochastic shocks to the variables. For the UK for instance, LRN fails under linear trends, and becomes unaddressable under broken trends. However, it is interesting to note that, apart from the 1918 break, UK output's long-run trend remained unaltered, even though two big shocks hit the level (1938), and level and trend (1970) of money. The absence of breaks in real output following these two shocks in money is what we refer to as "deterministic" LRN (with respect to the break in level), and LRSN (with respect to the second break). With the data set used in this paper, we are able

[6]It is well documented by now that structural breaks in the trend function of macro series are responsible for the 'apparent' unit root behaviour which results from ignoring them in the model's specification. In a recent paper, Arestis and Biefang-Frisancho Mariscal (1999) conclude that "...unit root tests that do not account sufficiently for the presence of structural breaks are misspecified and suggest excessive persistence"(p.155).

to show the extent to which breaks affect neutrality conclusions. Results range from no effect of breaks, to reversing conclusions regarding monetary neutrality. It also offers the possibility of formulating an heuristic notion of deterministic neutrality.

We use as our starting point the results recently obtained in Noriega (2004) regarding the order of integration of money and output for the data set described above. With this information, we then analyze the behavior of these orders of integration under different trend specifications, allowing for an increasing number of structural breaks in the long-run trend function under the alternative hypothesis. Note that under broken trend-stationary models, permanent changes are deterministic, as opposed to stochastic. This allows the possibility of investigating any potential relationship between the estimated break dates and historic events. The identified breaks can be analyzed through careful examination of the particular economic and political environment surrounding them.

The rest of the paper is organized as follows. Next section briefly presents the FS reduced-form tests for LRN and LRSN of money. Then, a description of the methods for testing stationarity while allowing for an unknown number of structural breaks in the trend function of the data is provided. Our econometric methodology is based on methods of Bai (1997b), Bai and Perron (1998a, b), and Noriega and de Alba (2001). Section 3 reports the empirical results on monetary neutrality under both the traditional stochastic interpretation, and a deterministic one. Section 4 offers some discussion of the results and concludes.

2 Econometric Methodology

2.1 Tests of LRN and LRSN

For testing LRN we follow Fisher and Seater's (1993) methodology. Briefly, they show that LRN and LRSN can be tested through the significance of the slope parameters b_k in the following long-horizon (OLS) regression:

$$\left[\sum_{j=0}^{k} \Delta^{\langle y \rangle} y_{t-j}\right] = a_k + b_k \left[\sum_{j=0}^{k} \Delta^{\langle m \rangle} m_{t-j}\right] + \varepsilon_{kt}, \tag{1}$$

where y and m stand for real output and (exogenous) money; Δ represents the difference operator ($\Delta^j y_t = y_t - y_{t-j}$), $\langle y \rangle$ stands for the order of integration of y (i.e. $\langle y \rangle = 1$ means that y is integrated of order one, or $y \sim I(1)$), and ε is a mean zero uncorrelated random variable. Theoretically, $\lim_{k\to\infty} b_k \equiv b$, gives an estimate of the long-run derivative (LRD) of real output with respect to a permanent stochastic exogenous shock in both the level of money (denoted LRD_N), and the trend of money (denoted LRD_{SN}).

FS show that, in order to interpret neutrality results, the order of integration of output and money should obey certain restrictions. For instance, the order of integration of money should be at least equal to one ($\langle m \rangle \geq 1$) for LRN to make sense, otherwise there are no stochastic permanent changes in money that can affect real output. The following table summarizes values of the LRD under different possibilities on the order of integration of the variables.

Table 1

The LRD and the Order of Integration of Money and Output

	LRD_N			LRD_{SN}		
$\langle y \rangle$	$\langle m \rangle = 0$	$\langle m \rangle = 1$	$\langle m \rangle = 2$	$\langle m \rangle = 0$	$\langle m \rangle = 1$	$\langle m \rangle = 2$
0	undefined	$\equiv 0$	$\equiv 0$	undefined	undefined	$\equiv 0$
1	undefined	b	$\equiv 0$	undefined	undefined	b

Source: Adapted from Fisher and Seater (1993).

When $\langle m \rangle \geq \langle y \rangle + 1 \geq 1$ the long-run derivative is zero, providing direct evidence of neutrality. When $\langle m \rangle = \langle y \rangle = 1$, LRN is testable through b. In this case, LRD_N measures whether the permanent movements in output are associated with permanent movements in money. If for instance b is significantly different from zero, then LRN does not hold.

Superneutrality, however, is not addressable when there are no permanent changes in the growth rate of money. In other words, superneutrality requires $\langle m \rangle \geq 2$. When $\langle m \rangle = 2$ and $\langle y \rangle = 0$, $LRD_N = LRD_{SN} = 0$, i.e., both LRN and LRSN hold, since one cannot associate permanent shocks to the growth rate of money to nonexistent permanent changes in output (further discussion of several cases of interest can be found in FS). Therefore, proper determination of the orders of integration of y and m is crucial in assessing LRN and LRSN of money.

2.2 Unit roots and structural breaks in money and output

Fisher and Seater's tests of monetary neutrality rely on the presence of stochastic permanent changes in money and output. For instance, if there are no such changes in either variable, then LRN is unaddressable (the LRD_N is undefined). On the other hand if there is a stochastic permanent change in the level of money, while output follows a stationary process, then LRN holds (since $LRD_N = 0$).

The presence of permanent stochastic changes in money and output, as indeed in many other macro variables, depends, however, on the way the trend function is treated, i.e., the modelling of the long-run. The most common approaches in the literature include simple linear trends (Nelson and Plosser (1982)), broken trends (Perron (1989, 1997)), polynomial trends (Schmidt and Phillips (1992)), the Hodrick-Prescott filter (Hodrick and Prescott (1997), Cogley and Nason (1995)), and smooth transition trend models (Leybourne, et. al. (1998), Sollis, et. al. (1999)).[7] Among these, models allowing for structural breaks (broken trend models) have become very popular in the literature, both theoretical and applied (Lanne, et. al. (2003), Sen (2003), Perron and Shu

[7] See Pollock (2001) for the analysis of three different approaches to the estimation of econometric trends.

(2002), Maddala and Kim (1998), Ben-David and Papell (1995, 1998), Stock (1994)). As Perron (2003) has pointed out, "changes in the trend function bias unit root tests towards a non-rejection and they need to be explicitly accounted for prior to performing unit root tests" (p.5). We utilize a resampling procedure based on this idea -unit root testing is carried out allowing for up to four structural breaks in the trend function of the variables[8].

In particular, following Rudebusch (1992) and Diebold and Senhadji (1996), we simulate the distribution (and obtain the empirical density) of the t-statistic for the null of a unit root, under the hypotheses that the true models are both a Broken Trend Stationary (BTS) model with up to four structural breaks, and a Difference-Stationary (DS) model, both estimated from the data. We then compare the position where the sample estimate of the t-statistic for testing a unit root lies relative to the empirical densities under the estimated BTS DGPs and DS DGP.

We now present the procedure for testing the presence of a unit root with an unknown number of structural breaks in the deterministic trend function. Let us denote by Y_t the logarithm of the observed series (output or money). The first step is to estimate (by OLS) the following BTS models with $0 \leq m \leq 4$ structural breaks in both level and slope of trend, and DS model, respectively:

$$\Delta Y_t = \mu + \beta t + \sum_{i=0}^{m} \theta_i DU_{it} + \sum_{i=0}^{m} \gamma_i DT_{it} + \alpha Y_{t-1} + \sum_{i=1}^{k} a_i \Delta Y_{t-i} + \varepsilon_t, \qquad (2)$$

$$\Delta Y_t = \sum_{i=1}^{k} a_i \Delta Y_{t-i} + \varepsilon_t, \qquad (3)$$

for $t = 1, 2, \ldots, T$, where T is the sample size and ε_t is an iid process; DU_{it} and DT_{it} are dummy variables allowing changes in the trend's level and slope respectively, that is, $DU_{it} = \mathbf{1}(t > T_{b_i})$ and $DT_{it} = (t - T_{b_i})\mathbf{1}(t > T_{b_i})$, where $\mathbf{1}(\cdot)$ is the indicator function and T_{b_i} is the unknown date of the i^{th} break. We use the convention that $\theta_0 = \gamma_0 = 0$. Under the BTS model of Equation (2), $\alpha < 0$, so that Y_t generates stationary fluctuations around a deterministic linear trend, perturbed by m structural breaks. This is a generalization to m breaks of the Innovational Outlier Model, used by Perron (1989) and others.[9] Under the DS model (3), Y_t does not generate stationary cycles, due to the presence of a unit root ($\alpha = 0$). Note that in the DS model, no deterministic components are considered. The reason is that interest centers on the autoregressive parameter and its associated t-statistic estimated from (2), both of which are invariant

[8] Noriega (2003) analyzes the hit rate of a rule for estimating the number of breaks in a time series, based on a unit root test. The procedure starts with testing DS vs. TS (allowing no changes in the trend function). If the unit root hypothesis can not be rejected (due to the bias towards non-rejection), then one change in the trend function is allowed, and then a unit root test is performed again. The procedure ends when a rejection occurs (the bias disappears). His Monte Carlo results show a good performance of this rule, relative to a parameter constancy rule or the BIC.

[9] The only difference is that (2) does not include a pulse variable, called $D(TB)_t$ by Perron (1989). This is also the approach in Zivot and Andrews (1992).

with respect to the parameters $\mu, \beta, \theta_i, \gamma_i$, for any sample size[10]. Note that the location (T_{b_i}), type (level, trend, or level and trend), and number (m) of breaks, as well as the autoregressive order (k) in the above equations are unknown.

In order to determine the location of breaks, we use the criterion which selects the break dates, from all possible combinations of m break dates, that minimize the residual sum of squares from (2). The occurrence of a break has to be restricted to the following intervals. For $m = 1$, $k+1+h \leq T_{b_1} \leq T-mh$; for two breaks, $k+1+h \leq T_{b_1} \leq T-mh$ and $T_{b_1}+h \leq T_{b_2} \leq T-(m-1)h$; for the three breaks case, $k+1+h \leq T_{b_1} \leq T-mh$, $T_{b_1}+h \leq T_{b_2} \leq T-(m-1)h$, and $T_{b_2}+h \leq T_{b_3} \leq T-(m-2)h$, etc., where h represents the smallest possible size for an interval or segment[11]. This criterion is called $\min RSS$.[12] Note that this criterion implies simultaneous determination of m breaks via a global search.

We follow Noriega and De Alba (2001) in the determination of the type of breaks and lag length allowed in the Innovational Outlier Model (2).[13] We first fix an arbitrary maximum value for k, labeled $k\max$. Then we estimate equation (2) with OLS for each of the three types of Innovational Outlier models (change in level only, change in level and trend, and change in trend only), over all possible values of T_{b_i}, and choose, for each model, the break date(s) for which the residual sum of squares (RSS) is minimized, as explained above. The Akaike Information Criterion (AIC) is then calculated for each of the three regressions corresponding to the estimated break dates. If the coefficient on the $k\max^{th}$ lag is not significant for the model which yields the smallest AIC, then we estimate the three versions of equation (2) again, over all possible values of T_{bi} with $k\max-1$ lags of the differenced dependent variable. Again, we choose the break date corresponding to the smallest RSS, and compute the AIC for the three regressions corresponding to the newly estimated break dates. Continuing in this fashion, we select the combination '*model type/lag length*' which corresponds to the model which yields the smallest value of the AIC (amongst the three models) *and* a corresponding significant lag (called $\hat{k}$), using a two-sided 10% test based on the asymptotic normal distribution. Note that if there are no significant lags, then $\hat{k} = 0$, which implies an AR(1) model for equation (2). If this is the case, the selection of the model follows simply from the lowest value of the AIC.

Due to the well known fact that structural change bias unit root tests towards non-rejection, we allow for (up to four) structural breaks in the trend function of the variables when testing for a unit root. In order to discriminate between DS and BTS, we utilize a resampling procedure as the one used by

[10] See for example Perron (1989, p.1393).

[11] This representation for h is based on the dynamic programming algorithm introduced by Bai and Perron (1998b) to obtain global minimizers of the RSS.

[12] This criterion for estimating break points is discussed in Bai (1997a,b), and Bai and Perron (1998a,b).

[13] Perron (1993) argues that, although a model allowing changes in both level and slope of trend is the most general one (it encompasses models with breaks in level alone, or with breaks in slope alone), there are power gains by estimating a model without irrelevant regressors. For example, model (2) with $\theta_i = 0$ would be more appropriate if it were apparent from the data that the type of break involved no change in level but only in trend.

Rudebusch (1992) and Diebold and Senhadji (1996). In particular, we simulate the distribution of the t-statistic for the null hypothesis of a unit root ($\alpha = 0$ in (2)), called $\widehat{\tau}$, under the hypotheses that the true models are the BTS models (2) and the DS model (3), both estimated from the data[14]. That is, under the BTS (DS) model we use the estimated parameters from (2)((3)), and the first $k+1$ observations as initial conditions ($\Delta Y_2, ..., \Delta Y_{k+1}$) to generate 10,000 samples of ΔY_t, $t = 2, ..., T$, with randomly selected residuals (with replacement) for each ΔY_t, $t = k+2, ..., T$ from the estimated BTS (DS) model. For each sample thus generated, regression equation (2) is run and the corresponding 10,000 values of $\widehat{\tau}$ are used to construct the empirical density function of this statistic under the BTS (DS) model, labeled $f_{BTS_m}(\widehat{\tau})$, $m = 0, ..., 4$ ($f_{DS}(\widehat{\tau})$).[15]

We then obtain and report the position where the sample estimate of the t-statistic for testing a unit root ($\overset{\wedge}{\tau}_s$) from the estimation of equation (2), lies relative to the empirical (simulated) densities. These positions are calculated as the probability mass to the left of $\overset{\wedge}{\tau}_s$, denoted $p_{BTS_m} \equiv \Pr[\widehat{\tau} \leq \widehat{\tau}_s \mid f_{BTS_m}(\widehat{\tau})]$ and $p_{DS} \equiv \Pr[\widehat{\tau} \leq \widehat{\tau}_s \mid f_{DS}(\widehat{\tau})]$.

We use in this section the above discussed convention that if a variable is stationary around a broken trend then it is integrated of order zero. We discuss below the implication of such convention. Results are given in Table A1 of the appendix. The first column indicates the number of breaks allowed in the trend function, m. The second column refers to the estimated lag length, $\hat{k}$. In the empirical applications $k\max$ is set at 5. The next columns report the estimated break dates. The type of break allowed in the trend function is reported in parenthesis. Column labeled AC reports the $p-$values for the Lagrange Multiplier test of the null hypothesis that the disturbances are serially uncorrelated against the alternative that they are autocorrelated of order one. The next column reports the value of the $t-$statistic for testing the null hypothesis of a unit root, estimated from equation (2). The probability mass to the left of this estimate, under each of the simulated DS and BTS specifications, is presented in the last two columns of the table.

In order to illustrate results of our testing procedure, let us analyze some particular countries. In the case of the U.K., Table A1 shows that for real output, the unit root can not be rejected against the alternative of trend-stationarity without breaks. When we allow for a drop in level and an increase in slope of trend in 1918, the unit root is strongly rejected[16], while the alternative is not: one would not be able to reject the estimated BTS model at even the 20% significance level. Note that, in fact, for all broken trend cases considered ($1 \leq m \leq 4$), the DS model is strongly rejected, while the various alternatives are not. We decide in favour of the BTS model with one break, however, since it is the inclusion of this single break which is sufficient to eliminate the unit root behaviour of output, and yields the most parsimonious specification. For

[14] A similar apprach is used by Kuo and Mikkola (1999) for the US/UK real exchange rate series.

[15] The 10,000 fitted regressions utilize the estimated value of k, under the BTS (DS) model. All calculations were carried out in GAUSS 3.6.

[16] These results are in line with those obtained by Duck (1992).

the money aggregate, the DS model is rejected only after two breaks are included in the trend function, one in level (1939), and another in level and trend (1970). Note that for the case of four breaks, the unit root is also rejected, while the alternative displays the same probabilities as the two breaks model; they can not be rejected at even the 15% level. Again, for reasons of parsimony, we reject the unit root null in favour of a model with two structural breaks. Figure 1 shows graphs of output and money for the U.K., together with their corresponding fitted broken trends.

A different picture arises for the case of Argentina. For real output, the DS model is rejected for all broken trend specifications, while under the BTS model, the probability closer to the middle of the distribution corresponds to the case of 3 structural breaks. For $M2$, the probabilities indicate that a stochastic permanent change can not be rejected. In fact, for the cases $m = 1, 2$, the rejection of the DS model is towards an explosive root, i.e., on the right tail of the empirical distribution. The values of the AR parameter for these cases are $\widehat{\alpha} = 1.07, 1.08$, respectively. For the cases $m = 3, 4$ the results indicate that the DS model can not be rejected. For the other two Latin American economies, Brazil and Mexico ($M1$), the picture is similar: after allowing for breaks, money remains stochastically nonstationary, while output becomes a broken trend stationary process.

Table 2 summarizes the empirical results. The countries in the sample have been grouped according to the effect of breaks on the order of integration of the variables. The first group includes Australia, Canada, Mexico ($M2$), Sweden, and the U.K. For this group, the inclusion of breaks has reduced the order of integration of both money and output (with the exception of Canadian output, which was already known to be $I(0)$ without breaks). A second group comprises the Latin American countries, Argentina, Brazil, and Mexico ($M1$), for which real output seems to follow a broken trend stationary model, while money remains stochastically nonstationary. Finally, for Italy the inclusion of breaks in the trend function does not alter the order of integration for money and output, already established by Noriega (2004) as I(1).

Table 2
Summary of Results

Country	Series	Order of Integration $m = 0^*$	Order of Integration $m > 0$	LRN m=0*	LRN m > 0 S	LRN m > 0 D	LRSN m=0*	LRSN m > 0 S	LRSN m > 0 D
Australia	Y	I(1)	I(0)+3b						
	M2	I(1)	I(0)+2b	F	NA	H	NA	NA	H
Canada	Y	I(0)	I(0)						
	M2	I(1)	I(0)+1b	HC	NA	H	NA	NA	NA
Mexico	Y	I(1)	I(0)+3b						
	M2	I(2)	I(0)+2b	HC	NA	F	F	NA	F
Sweden	Y	I(1)	I(0)+3b						
	M2	I(1)	I(0)+3b	H	NA	F	NA	NA	F
UK	Y	I(1)	I(0)+1b						
	M4	I(1)	I(0)+2b	F	NA	H	NA	NA	H
Argentina	Y	I(1)	I(0)+3b						
	M2	I(1)	I(1)	F	HC	NA	NA	NA	NA
Brazil	Y	I(1)	I(0)+2b						
	M2	I(2)	I(2)	HC	HC	NA	F	HC	NA
Mexico	Y	I(1)	I(0)+3b						
	M1	I(1)	I(1)	F	HC	NA	NA	NA	NA
Italy	Y	I(1)	I(1)						
	M2	I(1)	I(1)	F	NA	NA	NA	NA	NA

S, D stands for Stochastic and Deterministic respectively.
F, H, HC and NA stands for Fails, Holds, Holds by Construction and Not Addressable respectively.
*These results are taken from Noriega (2004).

3 Neutrality and Superneutrality Results

The effect of structural breaks on the order of integration of money and output poses an interesting question about the testing for monetary neutrality: ¿Should we derive conclusions on LRN based only on the stochastic version of the FS test? In order to interpret the evidence for structural breaks presented above, we utilize here a notion of *deterministic* monetary neutrality, which naturally arises in the absence of permanent stochastic shocks to the variables. We propose to broaden the concept of stochastic neutrality to one which allows a way of analyzing the long-run deterministic behaviour of money and output, and not only the stochastic one.

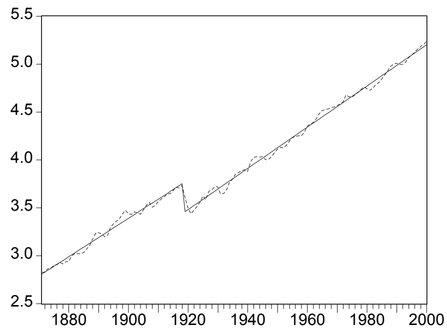
Based on results in the previous section, we present in Table 2 conclusions on LRN and LRSN. The columns under the headings LRN and LRSN show whether the neutrality propositions hold, fail, or are not addressable. Results are reported for the cases of no breaks ($m = 0$), and up to 4 breaks ($m > 0$).

When allowing for breaks, we offer two distinct interpretations of the results: one based on the stochastic (S) version of neutrality tests (Fisher and Seater (1993))[17], and one based on a deterministic (D) version.

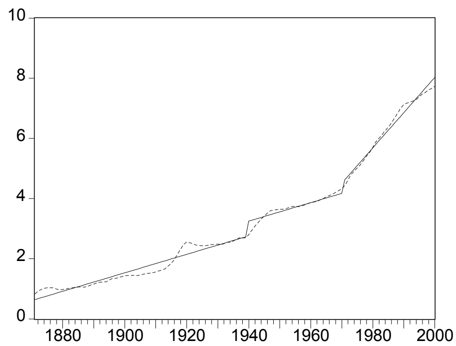
Take for instance the U.K. If breaks are not allowed, the FS test indicates that (stochastic) neutrality fails. Allowing for breaks, both variables are found to follow a stationary process around a broken trend, which means that (stochastic) LRN is not addressable, since there are no stochastic permanent changes in the variables. However, under a deterministic interpretation, some form on LRN seems to hold. According to our results, the long-run behaviour of U.K. output is well characterized as a linear trend, perturbed by a single break in 1918. On the other hand, U.K. money underwent (at least) two structural breaks, one in level in 1938, and the other in level and trend in 1970. Note that these two breaks had no effect on the long-run behaviour of output, which was found to follow a linear trend from 1918 onwards (see Figure 1).

Figure 1

United Kingdom, Y, I(0) + 1 break: 1918 (LT)



United Kingdom, M4, I(0) +2 breaks: 1939 (L), 1970 (LT)



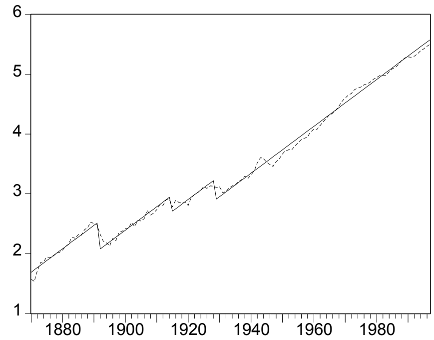
[17] Serletis and Krause (1996) and Serletis and Koustas (1998) utilize this interpretation when analyzing their empirical findings.

We say that money is deterministically neutral (DN) in the U.K. since output fluctuates in a stationary fashion around a linear trend with no breaks from 1918 to 2000. Furthermore, since the 1970 money-break was in level *and trend*, we say that U.K. money is also deterministically superneutral (DSN), at least over a horizon of 33 years.

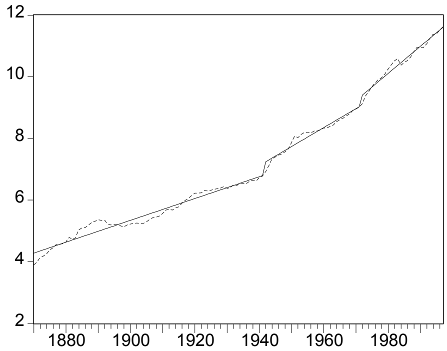
Similarly, for Australia we conclude that money is DSN (for over 50 years), since the trend of output remain unaltered, after the occurrence of two monetary breaks in level and slope of trend. See Figure 2. For Canada, a deterministic interpretation of LRN can be applied, since the big drop in level of money in 1920 had no effect on the long-run trend of output, which fluctuated stationarily around a linear trend for over 80 years after the (permanent) monetary break.

Figure 2

Australia, Y, I(0) + 3 breaks: 1891 (L), 1914 (L), 1928 (L)



Australia, M2, I(0) + 2 breaks: 1941 (LT), 1972 (LT)



In the case of Mexico, changes in the level and trend of $M2$ have been followed by changes in the trend function of output in the opposite direction:

the reduction in the growth rate of money in 1945 was followed by an increase in output growth in 1953; the noticeable increase in the rate of monetary growth in 1977 was followed by a severe slowdown in output growth, starting in 1981. Finally, the drastic slowdown of money growth in 1986 was followed by a recovery of the output growth rate starting in 1994. Although our analysis does not suggest any form of causality among breaks in the two variables, it seems hard to argue that money can be regarded as neutral in the case of Mexico. Similar arguments apply to Sweden.[18]

Note that, for all countries in this first group, the presence of deterministic changes in the trend function of the variables would lead to the (preliminary) conclusion that (stochastic) LRN is not addressable. Under the deterministic interpretation, on the other hand, LRN holds for Australia, Canada, and the U.K.

For the group of Latin American economies, deterministic neutrality is not addressable, due to the prevalence of stochastic (unit root) nonstationarity in the money series. However, LRN holds, if it is assumed that the stochastic permanent changes in money are uncorrelated with the structural breaks found in real output. Results for Argentina and Mexico ($M1$) again stand in contrast with previous results.

4 Discussion and Concluding Remarks

This paper empirically documents the impact of (endogenously determined) changes in the long-run trend of money and real output, on the Fisher and Seater (1993) tests of LRN and LRSN, using a long annual international data set. We present evidence on the inability to reject plausible broken-trend stationary models that exhibit transitory dynamics around a long-run deterministic trend subject to infrequent structural breaks. This is particularly true for the output series, whose orders of integration reduced after the inclusion of breaks (with the exception of Italy).[19]

It was found that conclusions on monetary neutrality are sensitive to the number of breaks allowed in the long-run trend of the relevant variables, and generally conflict with previous results reported in the literature. Noriega (2004) found mixed evidence in favour of LRN, holding for only half of the countries

[18] Some recent theoretical research has identified different sources of monetary non-neutrality. See for instance Gottardi (1994, pure portfolio effect under incomplete markets), and Bental and Eden (1996 inventories targeting with uncertain and sequential trade).

[19] According to Blanchard (1997), the dichotomy between an invariant steady-state path and fluctuations around it underlies the core of usable macroeconomics. For time series output data, this dichotomy is challenged by the nonstationarity of the unit root type. For the majority of countries in our sample, the unit root hypothesis is rejected in favour of broken trend stationary models. Hence, our results show that the dichotomy does hold, except for infrequent changes in long-run behaviour. Some recent research is building in this direction. Startz (1998) introduces a transmission mechanism in a two sector model of growth which allows sufficiently large shocks (to either technology or preferences) to induce multiple growth states, the theoretical counterpart of the broken trend models of Perron (1989). See also Durlauf (1993), Cooper (1994), Lau (1997), Acemoglu and Scott (1997) and Meng (2003).

in the sample. Our findings are also mixed. Allowing for breaks and under the stochastic interpretation, LRN is not addressable for half of the countries, while under the deterministic interpretation it holds for Australia, Canada, and the U.K. Furthermore, deterministic LRSN holds for Australia and the U.K.

Results presented in this paper lead to interesting questions about the testing for monetary neutrality, and more generally about the joint long-run behaviour of output and money. For instance, are there any connections between the breaks found in money and those found in output? (specially in the cases of Sweden and Mexico). For the case of Mexico, even though the standard FS test indicates that $M2$ is neutral, the identified changes in money growth (1976, 1985) could be behind the last two breaks found in the long-run trend of real output (1981, 1994). As far as the money aggregates is concerned, 1970 saw the end of the "Stabilizing Development" economic strategy, supplanted by a populist development strategy, which included the second (for $M1$) and third (for $M2$) largest upward trend shift in money of the century[20]. The resulting inflation and the rigid exchange rate policy, lead to a 76% peso devaluation between 1976 and 1977. In theory (see Marty 1994), anticipated inflation would lead people to economize on real balances, affecting the payment matrix and, therefore, the allocation of resources. After the second break in 1985, $M2$ growth rapidly declined, lowering the inflation rate, thus inducing the representative agent to devote less leisure time in acquiring commodities. This again affects the allocation of resources. Therefore, it is theoretically possible that these breaks had an influence on those registered for real output in 1981 and 1994. If this was the case, then it could be argued that permanent deterministic breaks in aggregate money are neither neutral nor superneutral for Mexico.[21]

These phenomena can be analyzed using the recently developed theory of co-breaking, introduced in Hendry and Mizon (1998), and Clements and Hendry (1999, chapter 9), or the techniques for testing for common features (Engle and Kozicki (1993), Vogelsang and Franses (2001)). The reduced rank technique developed by Krolzig and Toro (2000) yields information on how breaks are related through economic variables and across time. We hope to report results in this direction in a separate paper.

Our results suggest that a distinction should be made between reactions to deterministic and stochastic shocks. The FS test measures the correlation between permanent stochastic shocks in money and output data. Our findings suggest that it could be useful to broaden the notion of LRN by allowing for deterministic and stochastic LRN.

Finally, the use of smooth transition models (Leybourne, et. al. (1998),

[20] The first one ocurred around 1931, see Noriega and De Alba (2001).

[21] It is interesting to note that the 1953 upward trend break in real output coincides with the reduction in $M2$ volatility. When monetary volatility went up again in the late 70s, real output experienced a persistent slowdown, from 1981 onwards. A related phenomenon is documented in Ramey and Ramey (1995), who find that countries with high public expenditure volatility have lower growth. Santaella (1998) analized possible causes of the Mexico production slowdown starting in 1982. He argues that the evolution of both public spending and the inflation rate are consistent with the hypthesis that macroeconomic instability caused by expansionary policies is behind the slowdown in real GDP. See Santaella for further discussion.

Sollis, et. al. (1999)) to test for a unit root would help refining our results, since the break dates in these models are not restricted to be instantaneous, but allowed to occur along a number of periods in a smooth way. This might shed some light on the issue of co-breaking for Sweden and Mexico, by establishing more accurately the beginning and end of the breaks found for money and real output.

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6 Appendix

Table A1

TS Models allowing for Structural Breaks

Argentina, Y

1884-1996

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	0					.73	-2.23	.876	.472
1	0	1902 (L)				.93	-3.98	.795	.030
2	0	1902(L)	1980(L)			.72	-5.97	.767	.000
3	5	1912(T)	1917(LT)	1980(L)		.88	-5.73	**.516**	**.001**
4	5	1896(L)	1913(LT)	1929(LT)	1980(LT)	.53	-6.39	.913	.001

Argentina, M2

1884-1996

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	3					.69	-0.107	.830	.956
1	2	1989(LT)				.69	8.32	.038	1.00
2	5	1974(L)	1988(LT)			.38	6.01	.563	1.00
3	5	1930(LT)	1974(LT)	1988(LT)		.90	-1.55	.848	.813
4	5	1930(LT)	1970(LT)	1979(LT)	1988(LT)	.95	-3.19	.771	.403

Australia, Y

1870-1997

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	4					.95	-1.28	.922	.877
1	4	1889(L)				.77	-3.11	.879	.156
2	2	1891(LT)	1930(LT)			.96	-6.69	.899	.000
3	4	1891(L)	1914(L)	1928(L)		.75	-7.43	**.871**	**.000**
4	4	1891(L)	1914(LT)	1928(L)	1962(L)	.76	-8.65	.987	.000

Table A1 (continued)

TS Models allowing for Structural Breaks

Australia, M2

1870-1997

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	1					.98	-0.25	.924	.965
1	1	1933(T)				.94	-3.54	.877	.209
2	1	1941(LT)	1971(LT)			.99	-4.55	**.811**	**.064**
3	1	1892(LT)	1941(LT)	1972(LT)		.63	-6.09	.857	.007
4	2	1892(LT)	1941(LT)	1972(LT)	1983(LT)	.24	-5.43	.893	.034

Brazil, Y

1912-1995

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	1					.85	-2.73	.793	.229
1	1	1928 (LT)				.65	-4.14	.831	.030
2	3	1928(L)	1970(LT)			.87	-4.44	**.825**	**.055**
3	5	1928(L)	1940(L)	1980(LT)		.23	-6.07	.856	.000
4	4	1928(LT)	1947(T)	1970(L)	1980(T)	.36	-7.96	.810	.000

Brazil, M2

1912-1995

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	5					.02	-0.94	.999	.900
1	4	1987(LT)				.14	0.24	.143	.101
2	4	1968(T)	1987(LT)			.99	-3.56	.999	.046
3	5	1944(LT)	1965(LT)	1987(L)		.64	-9.68	.999	.000
4	5	1940(T)	1958(T)	1981(LT)	1987(LT)	.83	-7.95	.988	.000

Table A1 (continued)

TS Models allowing for Structural Breaks

Brazil, ΔM2

1913-1995

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	5					.17	-5.79	.999	.000
1	5	1987(LT)				.86	-0.28	.967	.962
2	5	1965(LT)	1987(LT)			.95	-0.02	.906	.978
3	5	1944(L)	1965(LT)	1987(LT)		.84	0.49	.789	.983
4	4	1923(T)	1944(L)	1965(LT)	1989(LT)	.32	3.64	.024	.999

Canada, M2

1870-2001

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	1					.60	-1.54	.882	.779
1	1	1920(L)				.95	-3.80	**.650**	**.052**
2	1	1875(LT)	1920(L)			.97	-3.57	.716	.087
3	5	1920(LT)	1940(L)	1969(L)		.53	-3.73	.672	.158
4	5	1920(L)	1940(LT)	1959(T)	1980(LT)	.95	-1.44	.749	.806

Italy, Y

1870-1997

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	1					.79	-1.83	.856	.678
1	2	1945(L)				.48	-2.86	.921	.27
2	5	1938(LT)	1945(LT)			.37	-1.42	.774	.929
3	5	1897(L)	1938(LT)	1945(LT)		.84	-0.85	.767	.965
4	5	1917(LT)	1929(L)	1939(LT)	1945(LT)	.51	-0.62	.732	.983

Table A1 (continued)
TS Models allowing for Structural Breaks

Italy, M2
1870-1997

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	1					.44	-2.63	.720	.222
1	1	1937(LT)				.43	-3.65	.902	.191
2	1	1914(LT)	1937(LT)			.43	-3.79	.946	.218
3	3	1914(L)	1939(LT)	1989(T)		.33	-7.24	.935	.000
4	2	1914(L)	1936(LT)	1946(LT)	1987(T)	.15	-7.39	.918	.000

Sweden, Y
1871-1988

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	1					.92	-2.63	.783	.281
1	3	1958(L)				.90	-3.91	.904	.041
2	1	1916(LT)	1930(LT)			.17	-4.49	.802	.058
3	5	1916(LT)	1930(LT)	1975(LT)		.09	-5.15	**.687**	**.017**
4	4	1892(T)	1916(LT)	1939(LT)	1968(LT)	.24	-9.55	.920	.000

Sweden, M2
1871-1988

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	2					.81	-1.59	.916	.795
1	4	1918(L)				.99	-2.40	.952	.451
2	4	1912(LT)	1918(LT)			.43	-3.29	.744	.273
3	1	1912(LT)	1918(LT)	1970(L)		.41	-5.45	**.776**	**.006**
4	3	1894(T)	1916(LT)	1935(T)	1970(LT)	.14	-9.20	.959	.0000

Table A1 (continued)

TS Models allowing for Structural Breaks

Mexico, Y

1932-2000

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	1					.97	-0.46	.946	.917
1	0	1981(L)				.95	-2.96	.749	.266
2	3	1953(T)	1981(LT)			.63	-4.09	.970	.151
3	5	1953(T)	1981(T)	1994(LT)		.76	-8.16	**.603**	**.000**
4	3	1953(T)	1981(T)	1985(L)	1994(LT)	.46	-9.25	.785	.000

Mexico, M1

1932-2000

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	3					.95	-1.40	.867	.737
1	1	1991(L)				.38	3.44	1.00	1.00
2	4	1971(T)	1991(LT)			.87	-2.52	.702	.187
3	4	1944(LT)	1971(T)	1991(LT)		.17	0.64	.598	.974
4	1	1942(LT)	1974(T)	1982(LT)	1991(T)	.73	-10.1	.929	.000

Mexico, M2

1932-2000

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	3					.73	-2.09	.818	.507
1	1	1987(LT)				.69	4.19	1.00	1.00
2	1	1976(T)	1985(LT)			.78	-4.74	**.880**	**.014**
3	2	1945(LT)	1977(T)	1986(LT)		.40	-5.15	.949	.017
4	5	1945(LT)	1959(LT)	1977(LT)	1986(LT)	.99	-5.15	.833	.055

Table A1 (continued)

TS Models allowing for Structural Breaks

United Kingdom, Y

1871-2000

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	3					.93	-1.67	.933	.742
1	1	1918(LT)				.91	-9.14	**.764**	**.000**
2	1	1902(L)	1918(L)			.78	-9.77	.763	.000
3	1	1902(L)	1918(L)	1979(LT)		.65	-10.32	.791	.000
4	1	1902(L)	1918(L)	1945(LT)	1973(L)	.61	-10.79	.839	.000

United Kingdom, M4

1871-2000

m	$\widehat{k}$	$Tc1$	$Tc2$	$Tc3$	$Tc4$	AC	$\widehat{\tau}_s$	p_{ET_m}	p_{ED_m}
0	2					.93	-0.94	.839	.902
1	2	1970(LT)				.91	-3.12	.853	.227
2	1	1939(L)	1970(LT)			.14	-5.53	**.833**	**.002**
3	1	1913(L)	1939(LT)	1970(LT)		.48	-7.57	.917	.000
4	1	1913(L)	1939(LT)	1967(T)	1989(LT)	.92	-9.66	.866	.000