XII

Volume XII, Number 2, November 2009

Journal of Applied Economics

Juncal Cunado
Luis A. Gil-Alana
Fernando Perez de Gracia

New evidence on long-run monetary neutrality

UNIVERSIDAD DEL CEMA
MCMLXXVIII
UCEMA

Edited by the Universidad del CEMA

Print ISSN 1514-0326
Online ISSN 1667-6726

Journal of Applied Economics. Vol XII, No. 2 (November 2009), 229-248

NEW EVIDENCE ON LONG-RUN MONETARY NEUTRALITY

Juncal Cunado, Luis A. Gil-Alana and Fernando Perez de Gracia*
Universidad de Navarra

Submitted December 2007; accepted June 2009

This paper re-examines the issue of long-run monetary neutrality by using fractional integration and allowing for a possible structural break in six countries: the United States, the United Kingdom, Mexico, Brazil, Australia and Argentina. We use an extension of Fisher and Seater's (1993) reduced-form test recently proposed by Bae, Jensen and Murdock (2005). The results show that long-run monetary neutrality holds for five countries when no structural breaks are taken into account, and for all countries if one break is allowed.

JEL classification codes: E40, E51, C32
Key words: money neutrality, long memory, structural breaks

I. Introduction

In the recent years, neutrality of money has received increasing attention from academic researchers. The empirical evidence concerning the monetary neutrality hypothesis is mixed. Many papers have analyzed long-run neutrality (e.g., King and Watson 1992, 1997; Fisher and Seater 1993; Boschen and Otrok 1994; Haug and Lucas 1997; Serletis and Koustas 1998, 2001; Shelley and Wallace 2004; Noriega 2004; Coe and Nason 2004; Bae et al. 2005; Noriega and Soria 2005; and Noriega et al. 2005, among many others). Most of these papers test the existence

* Fernando Perez de Gracia (corresponding author): University of Navarra, Department of Economics, Campus Universitario 31080, Pamplona, SPAIN ; tel: 00 34 948 425 625, fax: 00 34 948 425 626, e-mail: fgracia@unav.es. Juncal Cunado: email jcunado@unav.es. Luis A. Gil-Alana: email alana@unav.es. The authors thank Jorge Streb and three anonymous referees for helpful comments and suggestions on an earlier version of this paper. Juncal Cunado and Luis A. Gil-Alana gratefully acknowledge financial support from the Ministerio de Ciencia y Tecnología (ECO2008-03035 ECON Y FINANZAS, Spain). Fernando Pérez de Gracia also acknowledges research support from the Plan Especial de Investigacion of the Universidad de Navarra. Finally, we also thank Antonio Noriega for providing us the data set used in this paper. An earlier version of this paper can be found in Fundación de las Cajas de Ahorros (FUNCAS) - Working Paper no. 324.

of long-run neutrality (i.e., a permanent change in the money stock has no long-run consequences for real variables) and superneutrality (i.e., permanent changes in the growth rate of money have no real effects other than on real money balances, see McCallum 1990) using traditional unit-root tests on monetary aggregates and output with a long span of data. The tests are then applied on the reduced form of the Fisher and Seater (1993) conditions.

In line with the above papers, we also test for long-run monetary neutrality considering the reduced form of Fisher and Seater (1993). Two main differences of the present work with the former ones are the following. First, instead of using classic approaches based on I(1)/I(0) integration and cointegration we employ fractionally integrated techniques. Note that most of the above mentioned papers employ classic methods such as ADF (Dickey and Fuller 1979), PP (Phillips and Perron 1988), KPSS (Kwiatkowski et al. 1992), or some of the recent developments based on these procedures (Elliot et al. 1996; Ng and Perron 2001; etc.). These methods are too restrictive in the sense that they only consider the cases of I(0) stationarity and I(1) nonstationarity, and do not take into account fractional orders of integration. Bae et al. (2005) test long-run neutrality using a fractional approach in various countries. They apply the time domain maximum likelihood estimation procedure of Sowell (1992) in an Autoregressive Fractionally Integrated Moving Average (ARFIMA) model using annual data of money and real output for six countries. Their results support long-run neutrality in five out of the six countries considered. Second, we also allow for a single structural break in the fractionally integrated framework. In the above-mentioned literature, only a few papers consider structural breaks –endogenous or exogenous- in the I(1)/I(0) approach (for example, Boschen and Otrok 1994; Serletis and Krause 1996; Serletis and Koustas 1998; Noriega and Soria 2005; and Noriega et al. 2005). Moreover, structural breaks and fractional integration are issues that are intimately related (Gourieroux and Jasiak 2001; Diebold and Inoue 2001; Granger and Hyung 2004; etc.). Other authors like Lobato and Savin (1998) argue that structural breaks may be responsible for the long memory in return volatility processes. Engle and Smith (1999) investigate the relationship between structural breaks and long memory using a simple unit root process which occasionally changes over time. Other authors, such as Beran and Terrin (1996) and Bos et al. (1999) proposed Lagrange Multiplier tests for fractional integration with breaks.[1] Unlike these papers that assume a fixed fractional differencing

[1] Jensen and Liu (2006) also found spurious long memory behaviour in Markov switching models where the regime durations are time dependent.

parameter in the context of breaks, we use a methodology that permits us to consider different orders of integration at each subsample, implying different degrees of persistence.

In the present study, we focus on a single break to explain the stochastic nature of the series. The reason is the following. Though historical annual data such as those studied here may contain more than one single break, for the validity of the type of long-memory (fractional integration) model we use here it is necessary that the data span a sufficiently long period of time to detect the dependence across time of the observations; given the sample size of the series employed here, the inclusion of two or more breaks would result in relatively short sub-samples, thereby invalidating the analysis based on fractional integration. Moreover, other recent empirical studies on macro series in the United States and the United Kingdom come to the conclusion that a single break is sufficient to describe the behaviour of many series. Thus, for example, Boschen and Otrok (1994), Serletis and Krause (1996), Haugg and Lucas (1997), Serletis and Koustas (1998) and Shelley and Wallace (2004) among others test the long-run monetary neutrality hypothesis considering a single structural break.

This paper is organized as follows. The following section provides a brief description of the Fisher and Seater (1993) conditions of long-run monetary neutrality, along with the Bae et al. (2005) extension to the fractional case. In Section III we briefly describe the econometric approach employed in the paper for fractional integration and structural breaks. In Section IV, the long-run neutrality hypothesis is tested for six economies using long annual data already used by Noriega (2004) and Noriega et al. (2005). Finally, Section V contains some concluding comments.

II. The Fisher-Seater conditions and the Bae-Jensen-Murdock extension

Following Fisher and Seater (1993) we consider a bivariate ARMA model where m_t and y_t are log of nominal money supply and log of real output respectively. The stationary bivariate Vector Autoregressive (VAR) model is given by the equations:

$$a(L)(1-L)^{d_m} m_t \ = \ b(L)(1-L)^{d_y} y_t \ + \ u_t, \tag{1}$$

$$d(L)(1-L)^{d_y} y_t \ = \ c(L)(1-L)^{d_m} m_t \ + \ w_t, \tag{2}$$

where $a(L)$, $b(L)$, $c(L)$ and $d(L)$ are polynomials in the lag operator L (i.e., $Lx_t = x_{t-1}$), $a_0 = d_0 = 1$ and $\Delta = (1 - L)$; d_m and d_y refers respectively to the orders of integration of money supply and real output, which in most cases are assumed to be 0 or 1. The error vector $(u_t, w_t)^T$ is assumed to be independently and identically distributed, with zero mean and variance-covariance matrix V with elements σ_{uu}, σ_{ww} and σ_{uw}.[2]

As in Fisher and Seater (1993), the neutrality of money is obtained through the long-run derivative (*LRD*) or long-run elasticity of output with respect to permanent changes in money (represented by $LRD_{y,m}$),

$$LRD_{y,m} = \lim_{k \to \infty} \frac{\partial y_{t+k} / \partial u_t}{\partial m_{t+k} / \partial u_t}. \tag{3}$$

Equation (3) shows that the long-run derivative is the limit of the long-run elasticity of output with respect to money. According to Fisher and Seater (1993), there is evidence of monetary neutrality when $d_m \geq d_y + 1 \geq 1$, and the long-run derivative is zero.

In order to test for long-run monetary neutrality, most of the papers examine the orders of integration of log of nominal money supply and log of real output using standard I(0)/I(1) procedures (Fisher and Seater 1993; Boschen and Mills 1995; King and Watson 1997; Serletis and Koustas 1998; Noriega 2004; Noriega and Soria 2005; and Noriega et al. 2005). However, in a recent paper Bae et al. (2005) propose a fractionally integrated model to analyse the same topic. This is a relevant issue since the standard I(0)/I(1) procedures have very low power if the alternatives are of a fractional form. (Diebold and Rudebusch 1991; Hassler and Wolters 1994). Bae et al. (2005) present the extension of Fisher and Seater (1993) framework to the fractional case. This extension can be found in their Table 1, reproduced here for clarity of exposition. They present seven different cases where the relative orders of integration for m (d_m) and y (d_y) are between 0 and 1 or even above 1 along with the economic interpretation.

III. The econometric approach

In this section we present a recent procedure suggested by Gil-Alana (2008) that enables us to examine the stationarity/nonstationarity nature of the series of interest in a very general framework. Firstly, instead of restricting ourselves to the standard I(0) (stationarity) or I(1) (nonstationarity) cases, we consider the possibility of fractional orders of integration. Secondly, this framework also allows for the

[2] Following Fisher and Seater (1993) and Bae et al. (2005) among others, we assume the long-run exogeneity of money and independence between money and output shocks, i.e., $b(1) = \sigma_{uw} = 0$.

Table 1. Long-run neutrality restrictions by Bae, Jensen and Murdock (2005)

Case	Relative order of integration	Economic meaning	LRN→$LRD_{y,m}$
(i)	$0 < d_m < 1$	No permanent stochastic changes to *m*	Undefined
(ii)	$0 < d_y < 1 \leq d_m$	Permanent stochastic changes to *m*, no permanent stochastic changes to *y*	$\equiv 0$
(iii)	$1 \leq d_m = d_y$	Permanent stochastic changes to *m* and *y*	c(1)/d(1)
(iv)	$1 \leq d_m = d_{\Delta y}$	Permanent stochastic changes to m and Δy	c*(1)/d(1)
(v)	$1 \leq d_y < d_m$	Permanent stochastic changes to $(1–L)^{dm-1}m$, no permanent stochastic changes to $(1–L)^{dm-1}y$	$\equiv 0$
(vi)	$1 \leq d_m < d_y$	No permanent stochastic changes to $(1–L)^{dy-1}m$	Undefined
(vii)	$1 \leq d_{\Delta y} < d_m$	Permanent stochastic changes to *m*, no permanent stochastic changes to Δy	$\equiv 0$

Note: see Bae et al. (2005), Table 1, page 262. $c^*(L)$ = (1-*L*)-1*c*(*L*). LRN stands for long-run neutrality and $LRD_{y,m}$ stands for long-run derivative of *y* with respect to *m*.

inclusion of deterministic terms, like intercepts or linear trends. Finally, the possibility of structural breaks at unknown points in time is also taken into account.

For the purpose of simplicity, we suppose that there is just a single break in the data. Following Gil-Alana (2008) we assume that y_t is the observed time series, generated by the model

$$y_t = \hat{a}_1 + \hat{a}_1 t + x_t; \quad (1-L)^{d_1} x_t = u_t, \quad t = 1,...,T_b, \tag{4}$$

$$y_t = \alpha_2 + \beta_2 t + x_t; \quad (1-L)^{d_2} x_t = u_t, \quad t = T_b+1,...,T, \tag{5}$$

where the α's and the β's are the coefficients corresponding respectively to the intercept and the linear trend; d_1 and d_2 may be real values, u_t is I(0)[3], and T_b is the time of the break that is supposed to be unknown. Note that in the context of stochastic (fractional and non-fractional) differentiation, the inclusion of a linear time trend does not alter the interpretation of the results since it tends to disappear in the long run. Thus, for example, a model with a linear trend, $d = 1$, and white noise disturbances, becomes, for $t > 1$, a random walk model with a drift, and if the model includes only an intercept, it becomes a pure random walk model. In fractional contexts,

[3] An I(0) process is defined as a covariance stationary process with spectral density that is positive and finite at any frequency.

with values of d constrained between 0 and 1, the time trend becomes a constant in the long run and it tends to zero if $d > 1$.

On the other hand, the model in equations (4) and (5) can also be written as:

$$(1 - L)^{d_1} y_t = \alpha_1 \tilde{1}_t(d_1) + \beta_1 \tilde{t}_t(d_1) + u_t, \quad t = 1, ..., T_b, \tag{6}$$

$$(1 - L)^{d_2} y_t = \alpha_2 \tilde{1}_t(d_2) + \beta_2 \tilde{t}_t(d_2) + u_t, \quad t = T_b + 1, ..., T, \tag{7}$$

where $\tilde{1}_t(d_i) = (1 - L)^{d_i} 1$, and $\tilde{t}_t(d_i) = (1 - L)^{d_i} t$, $i = 1,2$.[4]

The approach adopted here is based on the least square principle. First, we choose a grid for the values of the fractionally differencing parameters d_1 and d_2, for example, $d_{io} = 0, 0.01, 0.02, \ldots, 2$, $i = 1, 2$. Then, for a given partition $\{T_b\}$ and given d_1, d_2-values, $(d_{1o}^{(j)}, d_{2o}^{(j)})$, we estimate the α's and the β's by minimising the sum of squared residuals,

$$\min_{w.r.t.\{\alpha_1,\alpha_2,\beta_1,\beta_2\}} \sum_{t=1}^{T_b} \left[(1-L)^{d_{1o}^{(j)}} y_t - \alpha_1 \tilde{1}_t(d_{1o}^{(j)}) - \beta_1 \tilde{t}_t(d_{1o}^{(j)}) \right]^2$$
$$+ \sum_{t=T_b+1}^{T} \left[(1-L)^{d_{2o}^{(j)}} y_t - \alpha_2 \tilde{1}_t(d_{2o}^{(j)}) - \beta_2 \tilde{t}_t(d_{2o}^{(j)}) \right]^2$$

in case of uncorrelated u_t, or, alternatively, using GLS for weakly autocorrelated disturbances. Let $\hat{\alpha}(T_b;\, d_{1o}^{(1)}, d_{2o}^{(1)})$ and $\hat{\beta}(T_b;\, d_{1o}^{(1)}, d_{2o}^{(1)})$ denote the resulting estimates for partition $\{T_b\}$ and initial values $d_{1o}^{(1)}$ and $d_{2o}^{(1)}$. Substituting these estimated values in the objective function, we obtain RSS(T_b; $d_{1o}^{(1)}$, $d_{2o}^{(1)}$), and minimising this expression for all values of d_{1o} and d_{2o} in the grid we obtain: $\mathrm{RSS}(T_b) = \arg\min_{\{i,j\}} \mathrm{RSS}(T_b; d_{1o}^{(i)}, d_{2o}^{(j)})$. Then, the estimated break date, $\hat{T}_k$, is such that $\hat{T}_k = \arg\min_{i=1,\ldots,m} \mathrm{RSS}(T_i)$, where the minimisation is over all partitions T_1, T_2, …, T_m, such that $T_i - T_{i-1} \geq |\varepsilon T|$. The regression parameter estimates are the associated least-squares estimates of the estimated k-partition, i.e., $\hat{\alpha}_i = \hat{\alpha}_i(\{\hat{T}_k\}), \hat{\beta}_i = \hat{\beta}_i(\{\hat{T}_k\})$, and their corresponding differencing parameters, $\hat{d}_i = \hat{d}_i(\{\hat{T}_k\})$, for $i = 1$ and 2. Several Monte Carlo experiments conducted in Gil-Alana (2008) show that the procedure performs well even in relatively small samples. This model can be easily extended to allow for multiple breaks.[5]

---

[4] In what follows, we assume that $(1-L)^{d_i} y_t = \tilde{1}_t(d_i) = \tilde{t}(d_i) = 0$, for $t \leq 0$. This is a standard assumption in the applied work, and is related with the Type II definition as opposed to the Type I definition of fractional integration (Robinson and Marinucci 2001; Gil-Alana and Hualde 2009).

[5] See again Gil-Alana (2008).

IV. Data and results

In this section we examine if there is evidence of the monetary neutrality hypothesis using annual international data. We use the same dataset as in Noriega (2004) and Noriega et al. (2005). The data include information on real output and monetary aggregates for a group of six countries: Argentina, Australia, Brazil, Sweden, the United Kingdom and the United States.[6] The starting dates are 1869 for the United States; 1870 for Australia; 1871 for the United Kingdom; 1884 for Argentina; 1912 for Brazil, and 1932 for Mexico. The ending years are 1995 (Brazil), 1996 (Argentina), 1997 (Australia) and 2000 for Mexico, the United States and the United Kingdom. The original dataset elaborated by Noriega (2004) included ten economies (Argentina, Australia, Brazil, Canada, Denmark, Italy, Mexico, Sweden, the United Kingdom and the United States). However, in this paper we focus only on the six economies for which the available data present relatively long series. Note that for the analysis of fractional integration, as is the case in the present paper, we need a sufficiently long span of data, especially if structural breaks are taken into account.

We first suppose that there are no breaks and look at the orders of integration of the series from a fractional viewpoint. We compute the estimates of d based on Sowell's (1992) maximum likelihood in the time domain, along with Robinson's (1994) parametric tests in the frequency domain. In both cases the results are essentially the same. An advantage of Robinson's (1994) approach is that it allows us to test the fractional differencing parameter for any real value, including thus stationary ($d < 0.5$) and nonstationary ($d \geq 0.5$) cases. In those cases where d was found to be above 0.5 with Robinson's (1994) method, we first differentiated the data before implementing Sowell's (1992) approach adding then 1 to obtain the proper estimates of d. Table 2 refers to the estimates of the monetary aggregates while Table 3 displays the results for the real output. In both tables we assume that the disturbances are white noise and autocorrelated, in the latter case using the exponential model of Bloomfield (1973),[7] and we do so for the three cases of no deterministic components, an intercept and an intercept with a linear time trend.

---

[6] For a detailed description of the source of the variables and the sample period see Table 1 in Noriega (2004). Noriega (2004) uses the Backus and Kehoe (1992) dataset for Australia; Bae and Ratti's (2000) data for Argentina and Brazil, and Friedman and Schwartz's (1992) for the United Kingdom and the United States.

[7] The model of Bloomfield (1973) is a non-parametric approach of modelling the I(0) disturbances that produces autocorrelations decaying exponentially as in the AR(MA) case.

Table 2. Estimates of *d* based on maximum likelihood in the frequency domain

		Monetary aggregates			Real income		
Country	Disturbances	NR	AI	LT	NR	AI	LT
Argentina	White noise	1.14	1.84	1.85	**0.97**	**0.93**	**1.12**
(1884-1996)		(1.05, 1.28)	(1.62, 2.23)	(1.62, 2.23)	**(0.86, 1.12)**	**(0.84, 1.09)**	**(0.96, 1.29)**
	Bloomfield	**1.15**	1.23	1.26	**0.92**	**0.75**	**0.85**
		(0.99, 1.42)	(1.13, 1.37)	(1.15, 1.42)	**(0.73, 1.21)**	**(0.68, 1.13)**	**(0.69, 1.08)**
Australia	White noise	**1.02**	1.23	1.23	**1.00**	**1.01**	**1.01**
(1870-1997)		**(0.92, 1.16)**	(1.13, 1.38)	(1.13, 1.38)	**(0.90, 1.12)**	**(0.92, 1.15)**	**(0.93, 1.15)**
	Bloomfield	**1.01**	**1.12**	**1.13**	**1.08**	**1.02**	**1.04**
		(0.85, 1.25)	**(0.99, 1.31)**	**(0.99, 1.33)**	**(0.89, 1.41)**	**(0.86, 1.40)**	**(0.85, 1.38)**
Brazil	White noise	**1.08**	1.82	1.79	**0.94**	1.40	1.43
(1912-1995)		**(0.96, 1.28)**	(1.62, 2.69)	(1.60, 2.66)	**(0.81, 1.13)**	(1.17, 1.72)	(1.22, 1.72)
	Bloomfield	**0.90**	1.49	1.51	**0.84**	**0.97**	**0.91**
		(0.71, 1.15)	(1.34, 1.70)	(1.37, 1.70)	**(0.61, 1.17)**	**(0.86, 1.25)**	**(0.64, 1.25)**
Mexico	White noise	1.21	1.48	1.49	**0.97**	1.33	1.34
(1932-2000)		(1.09, 1.40)	(1.33, 1.69)	(1.34, 1.73)	**(0.79, 1.20)**	(1.15, 1.57)	(1.20, 1.50)
	Bloomfield	**1.18**	1.28	1.43	**0.76**	**1.22**	**1.13**
		(0.98, 1.56)	(1.10, 1.66)	(1.18, 1.93)	**(0.37, 1.26)**	**(0.95, 1.64)**	**(0.92, 1.41)**
United Kingdom	White noise	1.27	1.94	1.96	**0.98**	**1.16**	**1.17**
(1871-2000)		(1.18, 1.41)	(1.75, 2.22)	(1.77, 2.20)	**(0.88, 1.12)**	**(0.99, 1.42)**	**(0.99, 1.42)**
	Bloomfield	1.20	1.36	1.43	**0.94**	0.81	0.72
		(1.07, 1.44)	(1.21, 1.62)	(1.27, 1.71)	**(0.75, 1.18)**	(0.73, 0.97)	(0.57, 0.98)
United Sates	White noise	**1.10**	1.64	1.63	**0.99**	**0.98**	**0.98**
(1869-2000)		**(1.00, 1.25)**	(1.43, 1.92)	(1.43, 1.91)	**(0.86, 1.14)**	**(0.80, 1.24)**	**(0.82, 1.24)**
	Bloomfield	**1.00**	**0.90**	**0.82**	**0.87**	0.74	0.49
		(0.87, 1.25)	**(0.83, 1.21)**	**(0.51, 1.20)**	**(0.59, 1.18)**	(0.68, 0.83)	(0.21, 0.86)

Notes: NR stands for No Regressors, AI for An Intercept, and LT for Linear Trend. In bold those cases where the unit root null hypothesis cannot be rejected at 5% level. The values in parenthesis refer to the 95% confidence intervals for the values of *d*.

Starting with the monetary aggregates (see Table 2) we observe that if the disturbances are white noise, the estimates of d are strictly above 1 in all cases, the values ranging from 1.02 (Australia with no regressors) to 1.96 (United Kingdom with a linear trend), and the unit-root null hypothesis (i.e., $d = 1$) is rejected in practically all cases in favor of higher orders of integration, the only exceptions being the United States, Brazil and Australia in the case of no deterministic terms.

If we permit autocorrelation throughout the model of Bloomfield (1973), the values of d are generally smaller than in the white noise case, though again above 1 in the majority of cases. The exceptions are now the United States with an intercept / linear trend and Brazil with no deterministic components. As a conclusion, we can summarize the results presented in this table by saying that the order of integration for the monetary aggregates seems to be in most cases above 1. If the disturbances are autocorrelated, the values are slightly smaller, though, if we allow for an intercept and/or a linear time trend the unit-root hypothesis is rejected in favor of $d > 1$ in all countries except in the United States and Australia.

We can interpret the above results in the following way: given that in most cases d is strictly higher than 1, shocks affecting the monetary aggregates will be not only permanent, but shocks affecting the growth rates will take longer time to disappear than in the standard I(1) (ARIMA) case usually employed in the time series literature.

Table 2 also reports the estimates of d for the real output across countries. The estimates of d indicates that the unit-root cannot be rejected in many cases. Thus, if we do not consider deterministic terms, the unit-root is included in all the intervals. Including an intercept and/or a linear trend, the unit root cannot be rejected for the United States and the United Kingdom (in case of white noise u_t), for Mexico and Brazil (with autocorrelated disturbances) and for Australia and Argentina for the two types of disturbances. Finally, the lowest degrees of persistence are achieved in the cases of the United States and the United Kingdom with autocorrelated disturbances. In these cases, d is strictly smaller than 1 and statistically significantly different from 1. On the other extreme we have the cases of Mexico and Brazil with white noise u_t, with values of d strictly above 1 in all cases.

The results presented so far seem to indicate that the order of integration of the monetary aggregates is 1 or above 1, while the one corresponding to real output is 1 or smaller than 1. In order to have more concise results about the orders of integration of the series, we have selected the best specification for each series according first to the significance of the coefficients related with the deterministic components, and then using likelihood criteria (LR tests) to choose between the white noise and the weak dependence structure for the I(0) disturbance term. The estimates of d_m and d_y for each country are displayed in Table 3.

It is observed that for all cases d_m is higher than d_y. We also see that the United States is the only country where the two orders of integration are strictly below unity. For the United Kingdom, Brazil and Argentina, the order of integration of money is above 1, while d_y (the order of integration of ouput) is below 1. Finally, for Australia and Mexico, the two degrees of integration are strictly above 1. The orders of integration of money supply and real output, displayed in Table 3, suggest that long-run neutrality clearly holds for the United Kingdom. This may also be the case for Argentina and Brazil though here the unit root cannot statistically be rejected for the output series. According to Fisher and Seater (1993) and Bae et al. (2005), when $d_m \geq 1$ and $d_y \in (0,1)$ long-run monetary neutrality holds since real output will be unaffected in the long-run by a change in money (case (ii) in Table 1). Futhermore, in the cases of Australia and Mexico the long-run neutrality also holds, $d_y - d_m < 0$ (case (v) in Table 1). Finally, for the United States $d_m = 0.82$, and the long-run derivative is then not defined (case (i) in Table 1). Note however that for some of these countries, the results presented in Tables 2 and 3 indicate that there is considerable overlap in the confidence bands. Thus, the results reported so far should be taken with caution. For example, for the Australian economy, the null hypothesis of $d_y = d_m$ cannot be rejected at the 5% level, although the non-significance of the ratio $c(1)/d(1)$ (case (iii) in Table 1) suggests that long-run neutrality (LRN) still holds.

Table 3. Estimates of d_m and d_y in case of no structural break

Country	d_m (money)	d_y (output)
Argentina	1.26	0.85*
(1884-1996)	(1.15, 1.42)	(0.69, 1.08)
Australia	1.13*	1.04*
(1870-1997)	(0.99, 1.33)	(0.85, 1.38)
Brazil	1.51	0.91*
(1912-1995)	(1.37, 1.70)	(0.64, 1.25)
Mexico	1.43	1.13*
(1932-2000)	(1.18, 1.93)	(0.92, 1.41)
United Kingdom	1.43	0.72
(1871-2000)	(1.27, 1.71)	(0.57, 0.92)
United States	0.82*	0.49
(1869-2000)	(0.51, 1.20)	(0.21, 0.86)

Notes: * the unit root hypothesis cannot be rejected at the 5% level. 95% confidence bands in parenthesis.

Next we are concerned with the effect that a structural break in the data might have had on the above results. We estimate the date of the structural break using the procedure developed by Gil-Alana (2008) described in Section III. Figure 1 presents the graphs of monetary aggregates and real income data with a structural break for each country, which has been endogenously determined by the model.

Figure 1. Monetary aggregates and real income

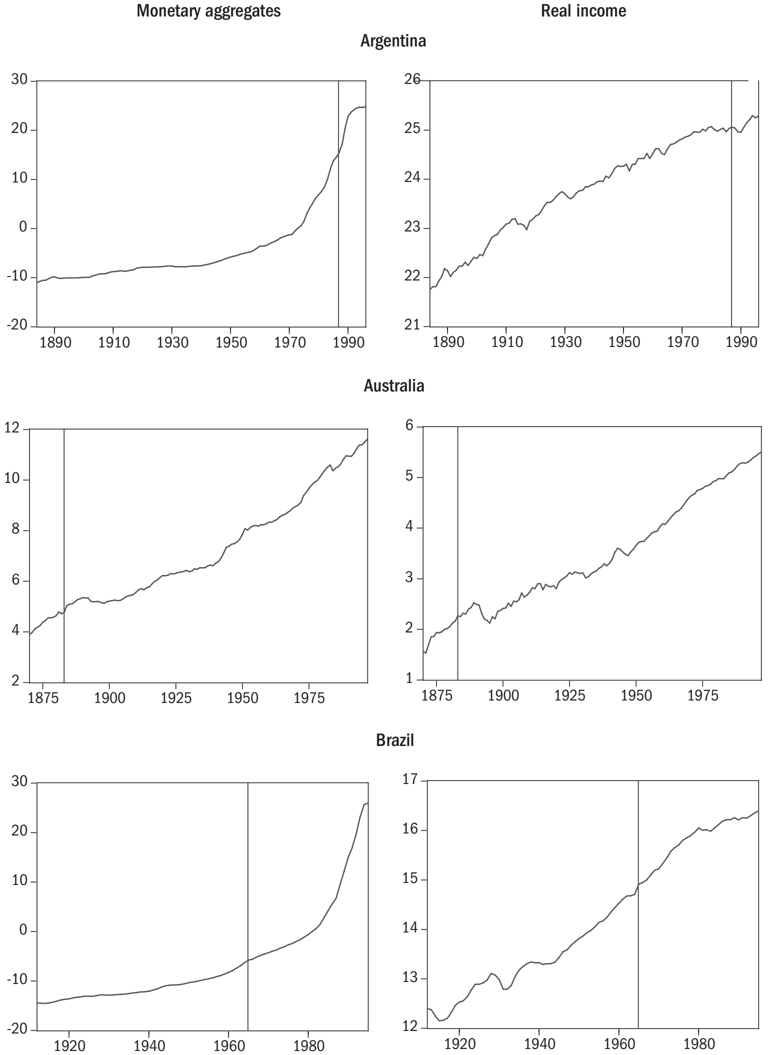

Figure 1 (continued). Monetary aggregates and real income

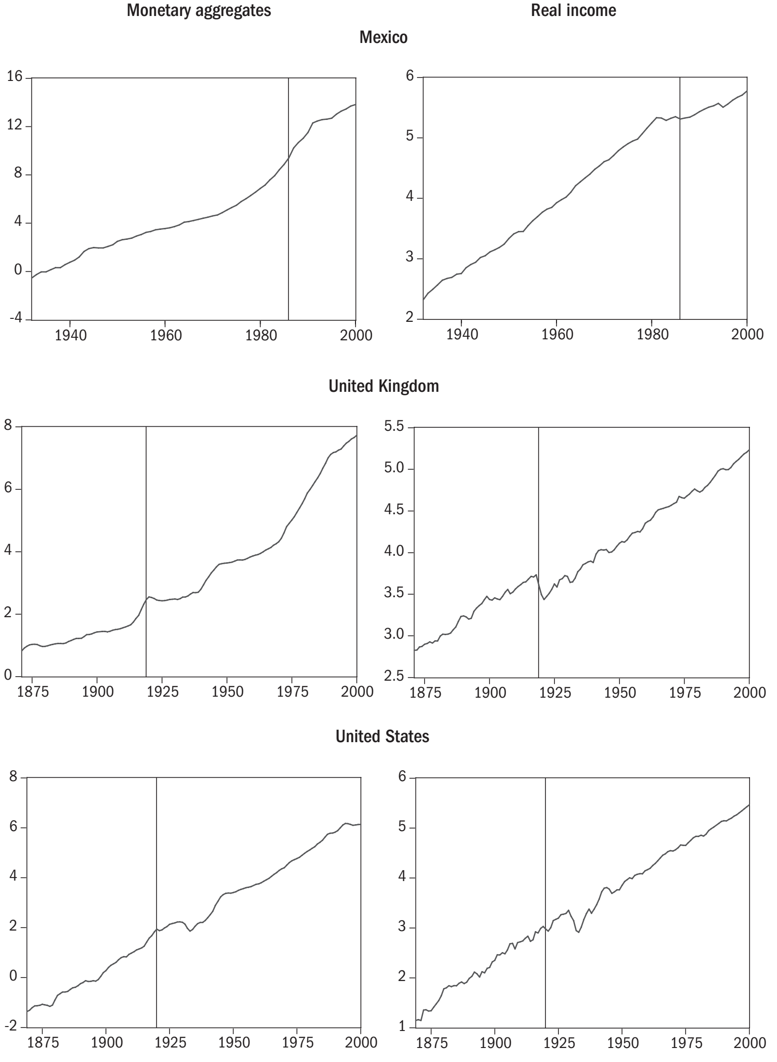

Table 4 displays the estimates of the fractional differencing parameters and the coefficients associated to the deterministic terms for each subsample along with the time of the break across countries. The break dates for monetary aggregates take place at 1883 for Australia; around 1920 for the United States and the United Kingdom; at 1965 for Brazil, and during the 1980s for Mexico and Argentina. Note that though we do not explicitly provide confidence intervals for the fractional differencing parameters in the procedure presented in Section III, they can be obtained by means of using alternative methods of fractional integration for each subsample. Across the tables we display the 95% confidence intervals corresponding to Robinson's (1994) univariate test, which is a Lagrange Multiplier procedure and it should thus approximate the maximum likelihood intervals. The orders of integration are substantially above 1 in all except one case (Australia, first subsample), and only for the United Kingdom and Brazil do we observe a decrease in the degree of persistence during the second subsample.[8]

Table 4 also presents results for real output. We observe that in all countries except in Mexico, the break date occurs now at the early part of the sample. It is at 1891 for Australia; at 1913 for Argentina; 1918 for the United Kingdom; 1931 for United States, and at 1982 for Mexico. For the first subsamples, the orders of integration are smaller than 1, (the exception here is Brazil), and the values increase during the second subsample for the United States, the United Kingdom, Australia and Argentina. The inclusion of a structural break reduces then the estimates of d_m and d_y in Brazil, the United Kingdom and the United States, while both estimates increase in the case of Australia. A crucial point here is to determine if the structural break should be included in the model or not. For this purpose we implemented some standard F-test statistics, testing the null of no break against the alternative of a single structural break, in the latter case using different regressions for each subsample. The results supported the existence of a break in all cases at the 5% level. Nevertheless, these results should also be taken with caution given the small number of observations used in cases like Australia (first subsample) and Mexico and Argentina (second subsamples).[9]

Table 5 displays the estimates for the two subsamples. The results for the first subsample show that long-run neutrality holds for Argentina, Mexico, the United States and the United Kingdom (see case (ii) of Bae et al.) though in all these cases the unit root cannot be rejected for d_y at conventional significance levels. For Brazil, the null hypothesis $d_y = d_m$ cannot be rejected at the 5% significance level, although the non-significance of the ratio $c(1)/d(1)$ in equation (3) (case (iii) in Table 1)

---

[8] In order to deal with the problem of breaks occurring in the extremes of the samples, we have constrained the analysis to the $[0.1T, 0.9T]$ interval of the samples. In case of Argentina, the break occurs at the extreme- right-point of the interval.

[9] The F-tests produced the following results: 5.87 (United States); 6.03 (Australia); 4.45 (United Kingdom); 4.50 (Argentina); 3.98 (Brazil); 5.55 (Mexico). Critical value: 2.60 (5%).

Table 4. Estimates of the parameters in the model with a single break: Monetary aggregates and real income

Country	Time break	First sub-sample			Second sub-sample		
		d_1	α_1	β_1	d_2	α_2	β_2
				Monetary aggregates			
Argentina	1987	**1.75**	**-11.205**	**-382.81**	**2.51**	0.2085	**3.8053**
(1884-1996)		**(1.58, 1.94)**	**(-58.51)**	**(-4.63)**	**(1.19, 2.98)**	(0.93)	**(4.82)**
Australia	1883	**0.71**	**3.851**	**0.0681**	**1.28**	**4.2135**	**0.0551**
(1870-1997)		**(0.40, 1.56)**	**(74.74)**	**(8.66)**	**(1.12, 1.41)**	**(14.63)**	**(2.84)**
Brazil	1965	**1.96**	**-14.378**	-0.0495	**1.85**	-29.3970	0.4329
(1912-1995)		**(1.64, 2.18)**	**(-195.0)**	(-0.48)	**(1.42, 2.64)**	(-0.72)	(0.58)
Mexico	1986	**1.55**	**-1.8626**	**0.2372**	**2.02**	-9.1703	**0.3533**
(1932-2000)		**(1.38, 1.71)**	**(-18.14)**	**(2.77)**	**(1.21, 2.97)**	(-1.28)	**(2.74)**
United Kingdom	1919	**2.04**	**0.7296**	**0.0963**	**1.86**	**3.4958**	-0.0187
(1871-2000)		**(1.79, 2.29)**	**(32.26)**	**(3.04)**	**(1.64, 2.11)**	**(1.84)**	(-0.49)
United States	1920	**1.45**	**-1.4256**	**0.0639**	**1.89**	0.2266	0.0312
(1869-2000)		**(1.12, 1.83)**	**(-30.13)**	**(2.08)**	**(1.60, 2.13)**	(0.078)	(0.57)

Table 4 (continued). Estimates of the parameters in the model with a single break: Monetary aggregates and real income

Country	Time break	First sub-sample			Second sub-sample		
		d_1	α_1	β_1	d_2	α_2	β_2
				Real income			
Argentina	1913	**0.71**	**21.704**	**0.0494**	**0.83**	**22.2430**	**0.0270**
(1884-1996)		**(0.58, 1.04)**	**(390.8)**	**(10.51)**	**(0.69, 1.35)**	**(185.43)**	**(8.38)**
Australia	1891	**0.44**	**1.5566**	**0.045**	**0.92**	**1.5951**	**0.0306**
(1870-1997)		**(0.19, 0.78)**	**(38.14)**	**(14.35)**	**(0.65, 1.07)**	**(16.53)**	**(8.76)**
Brazil	1930	**1.70**	**12.429**	-0.0252	**1.25**	**11.6921**	**0.0549**
(1912-1995)		**(1.04, 2.42)**	**(187.66)**	(-0.34)	**(1.01, 1.58)**	**(41.01)**	**(3.81)**
Mexico	1982	**0.88**	**2.265**	**3.8344**	**0.80**	**0.0601**	**0.0278**
(1932-2000)		**(0.78, 1.28)**	**(91.00)**	**(15.01)**	**(0.69, 1.57)**	**(25.98)**	**(5.85)**
United Kingdom	1918	**0.85**	**2.8034**	**0.0194**	**1.03**	**2.6378**	**0.0198**
(1871-2000)		**(0.77, 1.25)**	**(109.00)**	**(8.61)**	**(0.87, 1.48)**	**(13.31)**	**(4.97)**
United States	1931	**0.72**	**1.127**	**0.0341**	**1.18**	0.4930	**0.0383**
(1869-2000)		**(0.63, 1.02)**	**(19.11)**	**(11.82)**	**(1.06, 1.62)**	(0.78)	**(3.87)**

Notes: the values in parenthesis for d_1 and d_2 refer to the 95% confidence intervals for the fractional differencing parameters. For α_1, β_1, α_2 and β_2 they are t-values. In bold those cases where the unit root null hypothesis cannot be rejected at 5% level.

suggests that LRN holds (case (v) in Table 1). Finally, for Australia d_m is smaller than 1 though the unit root cannot be rejected, and thus long-run neutrality may also hold for this country. Table 5 also reports the estimates of d_m and d_y in the second subsamples across countries. The values of the estimates for Australia, Argentina and Mexico support case (ii) in Table 1, though once more $d_y = 1$ cannot be rejected. Brazil, the United States and the United Kingdom are consistent with case (v) in Table 1, supporting thus the long-run neutrality hypothesis for the six countries examined.

Table 5. Estimates of d_m and d_y

Country	d_m (money)	d_y (output)
	First subsample	
Argentina	1.75	0.71*
(1884-1996)	(1884, 1986)	(1884, 1912)
Australia	0.71*	0.44
(1870-1997)	(1870, 1882)	(1870, 1890)
Brazil	1.96	1.70
(1912-1995)	(1912, 1964)	(1912, 1929)
Mexico	1.55	0.88*
(1932-2000)	(1932, 1985)	(1932, 1981)
United Kingdom	2.04	0.85*
(1871-2000)	(1871, 1918)	(1871, 1917)
United States	1.45	0.72*
(1869-2000)	(1869, 1919)	(1869, 1930)
	Second subsample	
Argentina	2.51	0.83*
(1884-1996)	(1987, 1996)	(1913, 1996)
Australia	1.28	0.92*
(1870-1997)	(1883, 1997)	(1891, 1997)
Brazil	1.85	1.25
(1912-1995)	(1965, 1995)	(1930, 1995)
Mexico	2.02	0.80*
(1932-2000)	(1986, 2000)	(1982, 2000)
United Kingdom	1.86	1.03*
(1871-2000)	(1919, 2000)	(1918, 2000)
United States	1.89	1.18
(1869-2000)	(1920, 2000)	(1931, 2000)

Notes: the sample period for each country is in brackets. * the unit root hypothesis cannot be rejected at the 5% level. The 95% confidence bands are displayed in Table 4.

V. Concluding comments

In this paper we have re-examined the issue of long-run monetary neutrality in a group of six countries using fractional integration techniques and allowing for a single structural break that is endogenously determined by the model. Most of the previous empirical evidence is based on the reduced-form test of Fisher and Seater (1993), which is conducted via classic methods of I(0)/I(1) hypotheses. In this paper, we employ an extension of Fisher and Seater's (1993) model recently proposed by Bae et al. (2005) to the fractional case.

When we suppose that there is no break in the data, the results with fractional integration suggest that long-run monetary neutrality holds for Argentina, Australia, Brazil, Mexico and the United Kingdom, whereas US monetary neutrality is not addressable. However, these results are fairly ambiguous due to the overlapping in the confidence intervals for the fractional differencing parameters. When we take into account one structural break, we find that the long-run neutrality hypothesis holds in the two subsamples, and for all countries examined.

The results presented in this work still leave several questions unanswered. Thus, for example, we should investigate why the neutrality hypothesis is not addressable in the United States if no break is taken into account, which may be related with the importance of the break for this country. Another remarkable result is the fact that the break dates do not coincide either across countries or within each country for the two series examined. Thus, only for the United Kingdom are the breaks in output and money close in time, though for the United States and Australia (in the early part of the sample) and for Mexico (during the 1980s) the breaks are not too far apart. Note that the estimated break in US real income is in 1932 closely to the Great Depression period,[10] while the detected break in the monetary aggregate is previous to 1929 crisis. In the United Kingdom case, the monetary aggregates and the real income's breaks are related with the end of the First World War. Finally, in the other economies, as we can see, the detected breaks do not seem to be related with any significant or relevant economic event. Anyway, asynchronous breaks when the break dates in money and output do not line up and their effects on the long-run neutrality derivate is another issue that should be investigated more deeply.

---

[10] Earlier papers on long run neutrality (Boschen and Otrok, 1994; Bae and Ratti, 2000) also considered breaks in the United States during the Great Depression.

Finally, the paper can be extended in other directions. First, the orders of integration of the series can be estimated in a multivariate model.[11] Note that in the present work as well as in Bae et al. (2005) the estimation of d_m and d_y is univariate. Further investigation about the relationship between fractional integration and structural breaks should be elaborated even in univariate models. Note that some of the papers analyzing this relationship and referenced in this article come to the conclusion that in the presence of breaks the long memory issue becomes a spurious phenomenon. Other authors however have proposed statistics for testing such a relationship based on the invariance property of the long memory parameter to temporal aggregation (Ohanissian et al. 2008).

References

Backus, David K., and Patrick J. Kehoe (1992), International evidence on the historical properties of business cycles, *American Economic Review* **82**: 864-888.

Bae, Sang-Kun, and Ronald A. Ratti (2000), Long-run neutrality, high inflation and bank insolvencies in Argentina and Brazil, *Journal of Monetary Economics* **46**: 581-604.

Bae, Sang-Kun, Mark J. Jensen, and Scott G. Murdock (2005), Long-run neutrality in a fractionally integrated model, *Journal of Macroeconomics* **27**: 257-274.

Beran, Jan, and Norma Terrin (1996), Testing for a change of the long memory parameter, *Biometrika* **83**: 627-638.

Bos, Charles, Phillips H.Franses, and Marius Ooms (2001), Inflation, forecast intervals and long memory regression models, *International Journal of Forecasting* **18**: 243-264.

Boschen, John F., and Leonard O. Mills (1995), Tests of long-run neutrality using permanent monetary and real shocks, *Journal of Monetary Economics* **35**: 25-44.

Boschen, John F., and Charles Otrok (1994), Long-run neutrality and superneutrality in an ARIMA framework: Comment, *American Economic Review* **84**: 1470–1473.

Bloomfield, Peter (1973), An exponential model in the spectrum of a scalar time series, *Biometrika* **60**: 217-226.

Coe, Patrick J., and James M. Nason (2004), Long-run monetary neutrality and long-horizon regressions, *Journal of Applied Econometrics* **19**: 355-373.

Dickey, David., and Wayne A. Fuller (1979), Distributions of the estimators for autoregressive time series with a unit root, *Journal of the American Statistical Association* **74**: 427-431.

Diebold, Francis X., and Atsushi Inoue (2001), Long memory and regime switching, *Journal of Econometrics* **105**: 131-159.

Diebold, Francis X., and Glen Rudebusch (1991), On the power of Dickey-Fuller tests against fractional alternatives, *Economic Letters* **35**: 155-160.

Elliot, Graham, Thomas Rothenberg, and James H. Stock (1996), Efficient tests for an autoregressive unit root, *Econometrica* **64**: 813-836.

Engle, Robert F., and Aaron D. Smith (1999), Stochastic permanent breaks, *Review of Economics and Statistics* **81**: 553-574.

[11] Jensen (2009) extends the current framework to a multivariate setting in the Fisher effect hypothesis.

Fisher, Mark E.,and John J. Seater (1993), Long-run neutrality and superneutrality in an ARIMA framework, *American Economic Review* **83**: 402–415.

Friedman, Milton, and Anna J. Schwartz (1982), *Monetary trends in the United States and the United Kingdom. Their relation to income, prices and interest rates 1867-1975*, Chicago, University of Chicago Press.

Gil-Alana, Luis. A. (2003a), Multivariate tests of nonstationary hypotheses, *South African Statistical Journal* **37**: 1-28.

Gil-Alana, Luis A. (2003b), A fractional multivariate long memory model for the US and the Canadian real output, *Economics Letters* **81**: 355-359.

Gil-Alana, Luis A. (2008), Fractional integration and structural breaks at unknown periods of time, *Journal of Time Series Analysis* **29**: 163-185.

Gil-Alana, Luis A., and Javier Hualde (2009), Fractional integration and cointegration. An overview and an empirical application, *Palgrave Handbook of Econometrics*, vol.2, part 3.

Gourieroux, Christian, and Joann Jasiak (2001), Memory and infrequent breaks, *Economics Letters* **70**: 29-41.

Granger, Clive W.J., and Namwon Hyung (2004), Occasional structural breaks and long memory with an application to the S&P 500 absolute stock returns, *Journal of Empirical Finance* **11**: 399-421.

Haug, Alfred A., and Robert F. Lucas (1997), Long-run neutrality and superneutrality in an ARIMA framework: Comment, *American Economic Review* **87**: 756-759.

Hassler, Uwe, and Jurgen Wolters (1994), On the power of unit root tests against fractional alternatives, *Economic Letters* **45**: 1-5.

Jensen, Mark J. (2009), The long-run Fisher effect. Can it be tested?, *Journal of Money, Credit and Banking* **41**: 221-231.

Jensen, Mark J., and Ming Liu (2006), Do long swings in the business cycle lead to strong persistence in output?, *Journal of Monetary Economics* **53**: 597-611.

Kwiatkowski, Denis, Peter Phillips, Peter Schmidt, and Yongaheol Shin, (1992), Testing the null hypothesis of stationary against the alternative of a unit root: How sure are we that economic time series have a unit root?, *Journal of Econometrics* **54**: 159-178.

King, Robert G., and Mark Watson (1992), Testing long-run neutrality, Working Paper 4156, NBER.

King, Robert G., and Mark Watson (1997), Testing long-run neutrality, *Federal Reserve Bank of Richmond Economic Quarterly* **83**: 69-101.

Lobato, Ignacio N., and N. Savin (1998), Real and spurious long memory properties of stock market data, *Journal of Business and Economic Statistics* **16**: 261-268.

McCallum, Bennett T. (1990), Inflation: Theory and evidence, in B.M. Friedman and F.H. Hahn, eds., *Handbook of Monetary Economics*, vol. 2, Amsterdam, North-Holland.

Nielsen, Morten O. (2004), Efficient inference in multivariate fractionally integrated time series models, *Econometrics Journal* **7**: 63-97.

Nielsen, Morten O. (2005), Multivariate Lagrange Multiplier tests for fractional integration, *Journal of Financial Econometrics* **3**: 372-398.

Ng, Serena, and Pierre Perron (2001), Lag length selection and the construction of unit root test with good size and power, *Econometrica* **69**: 1519-1554.

Noriega, Antonio E. (2004), Long-run monetary neutrality and unit-root hypothesis: Further international evidence, *North American Journal of Economics and Finance* **15**: 179-197.

Noriega, Antonio E., and Luis M. Soria (2005), Structural breaks, order of integration and the neutrality hypothesis: Further evidence, mimeo.

Noriega, Antonio E., Luis M. Soria, and Ramón Velazquez (2005), International evidence on monetary neutrality under broken trend stationary models, mimeo.

Ohanissian, Arek, Jeffrey R. Russell, and Ruey S. Tsay (2008), True or spurious long memory? A new test, *Journal of Business and Economics Statistics* **26**, 161-175.

Phillips, Peter C.B., and Pierre Perron (1988), Testing for a unit root in a time series regression, *Biometrika* **75**: 335-346.

Robinson, Peter M. (1994), Efficient tests of nonstationary hypotheses, *Journal of the American Statistical Association* **89**: 1420-1437.

Robinson, Peter M., and David Marinucci (2001), Narrow band analysis of nonstationary processes, *Annals of Statistics* **29**: 947-986.

Serletis, Apostolos, and David Krause (1996), Empirical evidence on the long-run neutrality hypothesis using low-frequency international data, *Economics Letters* **50**: 323–327.

Serletis, Apostolos, and Zisimos Koustas (1998), International evidence on the neutrality of money, *Journal of Money, Credit, and Banking* **30**: 1–25.

Serletis, Apostolos, and Zisimos Koustas (2001), Monetary aggregation and the neutrality of money, *Economic Inquiry* **39**: 124-138.

Shelley, Gary L., and Frederick H. Wallace (2004), Testing for long-run neutrality in Mexico, mimeo.

Sowell, Fallaw (1992), Maximum likelihood estimation of stationary univariate fractionally integrated time series models, *Journal of Econometrics* **53**: 165-188.