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NEW EVIDENCE ON LONG-RUN MONETARY NEUTRALITY*

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ABSTRACT

This paper re-examines the issue of long-run monetary neutrality by using fractional integration and allowing for possible structural breaks in six countries. We use an extension of Fisher and Seater´s (1993) reduced-form test recently proposed by Bae et al. (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.

Keywords: Money neutrality; Long memory; Structural breaks.

JEL Classification: E40; E51; C32.

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1. 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.[1] 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 of long-run neutrality and superneutrality using traditional unit-root tests to the 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 papers aforementioned, 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 techniques we employ fractional integration. Note that most of the above mentioned papers employ classic methods such as DF (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.

[1] For example, King and Watson (1992) using US quarterly data found that long run neutrality was supported while Fisher and Seater (1993) found that long run neutrality was rejected using US annual data. Serletis and Krause (1996) found that the long run neutrality hypotheses were also supported using the Fisher and Seater' conditions.

(2005) test the long-run neutrality using a fractional approach in various countries.[2] 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 the long-run neutrality in five out of the six countries considered.

Second, we also allow for structural breaks in a fractionally integrated framework. In the above-mentioned literature, only a few papers consider structural breaks in the I(1)/I(0) approach (see, for example, Boschen and Otrok, 1994; Serletis and Krause, 1996; Serletis and Koustas, 1998 and Noriega et al., 2005 among others). Moreover, structural breaks and fractional integration are issues that are closely related. Granger and Hyung (1999), Gourieroux and Jasiak (2001), Diebold and Inoue (2001) are some of the papers relating these two concepts.

This paper is organized as follows. The following section provides a brief description of the Fisher and Seater´s (1993) conditions of long-run monetary neutrality, along with the Bae et al.`s (2005) extension to the fractional case. In Section 3 we briefly describe the econometric approach employed in the paper for fractional integration and structural breaks. In Section 4, 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 5 contains some concluding comments.

[2] Compared to Bae et al. (2005), this paper present two main differences. First, we used the testing methodoly recently proposed by Gil-Alana (2007) instead of using domain maximum likelihood estimation procedure proposed by Sowell (1992). Second, we use data for Argentina, Australia, Mexico, the UK and the US while Bae et al. (2005) use data for Argentina, Canada, Italy, Sweden, the UK and the US.

2. The Fisher - Seater conditions and the Bae et al.'s (2005) 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:

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

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

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 i.i.d., with zero mean and variance-covariance matrix V with elements σ_{uu}, σ_{ww} and σ_{uw}.

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 $LRD_{y,m}$),

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

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 (see for example 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 among others) using standard I(0)/I(1) procedures. However, in a recent paper Bae et al. (2005) propose a fractionally integrated model to analyse the same topic. They present the extension of Fisher and Seater (1993) framework to the fractional case. This extension can be found in Table 1 (page 262) in Bae et al. (2005). They present seven different cases where the relative order of integration for m (d_m) and y (d_y) are between 0 and 1 along with the economic interpretation.

3. The econometric approach

In this section we present a procedure suggested by Gil-Alana (2007) 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 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 by now that there is just a single break in the data. Following Gil-Alana (2007) we assume that y_t is the observed time series, generated by the model

$$y_t = \alpha_1 + \beta_1 t + x_t; \quad (1-L)^{d_1} x_t = u_t, \quad t = 1,\ldots,T_b \tag{4}$$

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

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) and T_b is the time of the break that is supposed to be unknown. Note that the model in equations (4) and (5) can also be written as:

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

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

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

The approach taken in this article 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,

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

for 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: $RSS(T_b) \;=\; \arg\min_{\{i,j\}} RSS(T_b; d_{1o}^{(i)}, d_{2o}^{(j)})$. Then, the estimated break date, $\hat{T}_k$, is such that $\hat{T}_k \;=\; \arg\min_{i=1,\ldots,m} RSS(T_i)$, where the minimisation is over all partitions $T_1, T_2, \ldots, T_m$, 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 (2007) show that the procedure performs well even in relatively small samples.

Clearly, this model can be extended to allow for multiple breaks. One then considers the following specification:

$$y_t = \alpha_j + \beta_j t + x_t; \quad (1 - L)^{d_j} x_t = u_t, \quad t = T_{j-1}+1,\ldots,T_j,$$

for j = 1, …, m+1, $T_0 = 0$ and $T_{m+1} = T$, and m stands for the number of breaks. The break dates $(T_1, \ldots, T_m)$ are explicitly treated as unknown and for i = 1, …, m, we have $\lambda_i = T_i/T$, with $\lambda_1 < \ldots < \lambda_m < 1$. Following the same lines as in the previous case, for each j-partition, $\{T_1, \ldots T_j\}$, denoted $\{T_j\}$, the associated least-squares estimates of α_j, β_j and the d_j are obtained by minimising the sum of squared residuals in the d_i-differenced models, i.e., $\sum_{j=1}^{m+1} \sum_{t=T_{j-1}+1}^{T_j} \left[(1-L)^{d_i} y_t - \alpha_i \tilde{1}_t - \beta_i \tilde{t}_t\right]^2$, where $\hat{\alpha}_i(T_j)$, $\hat{\beta}_i(T_j)$ and $\hat{d}(T_j)$ denote the resulting estimates. Substituting them in the new objective function and denoting the sum of squared residuals as $RSS_T(T_1, \ldots, T_m)$, the estimated break dates $(\hat{T}_1, \hat{T}_2, \ldots, \hat{T}_m)$ are obtained by: $\min_{(T_1, T_2, \ldots, T_m)} RSS_T(T_1, \ldots, T_m)$ where the minimisation is again obtained over all partitions $(T_1, \ldots, T_m)$.

In the present study, however, we focus instead 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 US and the UK 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.

4. 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, Mexico, the UK and the US.[3] The starting dates are 1869 for the US; 1870 for Australia; 1871 for the UK; 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 US and the UK.

[3] 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's (1992) dataset for Australia; Bae and Ratti's (2000) data for Argentina and Brazil, and Friedman and Schwartz's (1992) for the UK and the US.

We first suppose that there are no breaks in the data and look at the orders of integration of the series from a fractional viewpoint. We present the estimates of d based on maximum likelihood in the frequency domain for the monetary aggregates (see Table 1) and for real output (see Table 2). In both tables we assume that the disturbances are white noise and autocorrelated, in the latter case using the model of Bloomfield (1973)[4], and we do so for the three cases of no deterministic components, an intercept and an intercept with a linear trend.

Starting with the monetary aggregates (Table 1) 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 (UK 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 US, 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 US 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, and, if we allow for an intercept and/or a linear time trend the unit-root hypothesis is rejected in all countries except in the US and Australia.

[4] 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 case.

Table 2 reports the estimates of d for the real output across countries. A first look at this table 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 US and the UK (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 US and the UK 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) disturbances. Figure 1 summarizes the estimates of d_m and d_y for each country, and the numerical values are displayed in Table 3.

It is observed that for all cases d_m is higher than d_y. We also see that the US is the only country where the two orders of integration are strictly below unity. For the UK, 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 holds for Argentina, Brazil and the UK. 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 (see case (ii) in Bae et al.´s Table 1). Futhermore, in the cases of Australia and Mexico the long-run neutrality also holds, $d_y - d_m < 0$ (see case (v) in Bae et al.´s Table 1). Finally, for the US $d_m = 0.82$, and the long-run derivative is then not defined (see case (i) in Bae et al.´s Table 1).

Next we are concerned with the effect that a structural break in the data might have had on the above results. For this purpose we employ the procedure described in Section 3. Table 4 refers to the monetary aggregates and 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 take place at 1883 for Australia; around 1920 for the US and the UK; 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 3, 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 UK and Brazil do we observe a decrease in the degree of persistence during the second subsample.[5]

Table 5 refers to the 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 UK; 1931 for US, 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 US, the UK, Australia and Argentina. The inclusion of a structural break reduces then the estimates of d_m and d_y in Brazil, the UK and the US while both estimates increase in the case of Australia.

Figures 2 and 3 and Tables 6 and 7 are similar to Figure 1 and Table 3 above referring now to the first and second subsamples respectively. The results for the first subsample show that long-run neutrality holds for Argentina, Mexico, the US and the UK (see case (ii) of Bae et al.´s). For Brazil the values of 1.96 for d_m and 1.70 for d_y also suggests that long-run neutrality holds (see case (v) in Bae et al.´s paper). Finally, long-run neutrality is rejected in the case of Australia. Table 7 reports the estimates of d_m and d_y in the second subsample period across countries. The values of the estimates for Australia, Argentina and Mexico support case (ii) in Bae et al.'s (2005) paper. Brazil, the US and the UK are consistent with case (v) in that paper, supporting thus the long-run neutrality hypothesis for the six countries examined.

[5] 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.

5. 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 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) 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 UK, whereas the US monetary neutrality is not addressable. However, when we take into account one possible structural break, we find that in five cases (Argentina, Brazil, Mexico, the UK and the US) the long-run monetary neutrality holds in the first subsamples, and for all countries in the second subsamples.

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 US if no break is taken into account or why this hypothesis is not satisfied in the case of Australia for the first subsample. 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 UK are the breaks in output and money close in time, though for the US and Australia (in the early part of the sample) and for Mexico (during the 1980s) the breaks are not too far apart.

Finally, the paper can be extended in several directions. First, the orders of integration of the series can be estimated in a multivariate model. Note that in the present work as well as in Bae et al. (2005) the estimation of d_m and d_y is univariate. Multivariate models of fractional integration have been recently developed by Gil-Alana (2003a, b) and Nielsen (2005), the latter proposing a time domain version of Gil-Alana's tests. These approaches can be viewed as reduced form models that permit us to identify structural fractional VAR models through standard identification restrictions. The bivariate fractional VAR approach can be then extended to the case of structural breaks though the theoretical model in this context still needs to be fully developed.

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TABLE 1: Estimates of d based on maximum likelihood in the frequency domain: MONETARY AGGREGATES				
Country	Disturbances	No regressors	An intercept	A linear trend
UNITED STATES	White noise	**1.10 (1.00, 1.25)**	1.64 (1.43, 1.92)	1.63 (1.43, 1.91)
	Bloomfield	**1.00 (0.87, 1.25)**	**0.90 (0.83, 1.21)**	**0.82 (0.51, 1.20)**
UNITED KINGDOM	White noise	1.27 (1.18, 1.41)	1.94 (1.75, 2.22)	1.96 (1.77, 2.20)
	Bloomfield	1.20 (1.07, 1.44)	1.36 (1.21, 1.62)	1.43 (1.27, 1.71)
MEXICO	White noise	1.21 (1.09, 1.40)	1.48 (1.33, 1.69)	1.49 (1.34, 1.73)
	Bloomfield	**1.18 (0.98, 1.56)**	1.28 (1.10, 1.66)	1.43 (1.18, 1.93)
BRAZIL	White noise	**1.08 (0.96, 1.28)**	1.82 (1.62, 2.69)	1.79 (1.60, 2.66)
	Bloomfield	**0.90 (0.71, 1.15)**	1.49 (1.34, 1.70)	1.51 (1.37, 1.70)
AUSTRALIA	White noise	**1.02 (0.92, 1.16)**	1.23 (1.13, 1.38)	1.23 (1.13, 1.38)
	Bloomfield	**1.01 (0.85, 1.25)**	**1.12 (0.99, 1.31)**	**1.13 (0.99, 1.33)**
ARGENTINA	White noise	1.14 (1.05, 1.28)	1.84 (1.62, 2.23)	1.85 (1.62, 2.23)
	Bloomfield	**1.15 (0.99, 1.42)**	1.23 (1.13, 1.37)	1.26 (1.15, 1.42)

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.

TABLE 2: Estimates of d based on maximum likelihood in the frequency domain: REAL INCOME				
Country	Disturbances	No regressors	An intercept	A linear trend
UNITED STATES	White noise	**0.99 (0.86, 1.14)**	**0.98 (0.80, 1.24)**	**0.98 (0.82, 1.24)**
	Bloomfield	**0.87 (0.59, 1.18)**	0.74 (0.68, 0.83)	0.49 (0.21, 0.86)
UNITED KINGDOM	White noise	**0.98 (0.88, 1.12)**	**1.16 (0.99, 1.42)**	**1.17 (0.99, 1.42)**
	Bloomfield	**0.94 (0.75, 1.18)**	0.81 (0.73, 0.97)	0.72 (0.57, 0.98)
MEXICO	White noise	**0.97 (0.79, 1.20)**	1.33 (1.15, 1.57)	1.34 (1.20, 1.50)
	Bloomfield	**0.76 (0.37, 1.26)**	**1.22 (0.95, 1.64)**	**1.13 (0.92, 1.41)**
BRAZIL	White noise	**0.94 (0.81, 1.13)**	1.40 (1.17, 1.72)	1.43 (1.22, 1.72)
	Bloomfield	**0.84 (0.61, 1.17)**	**0.97 (0.86, 1.25)**	**0.91 (0.64, 1.25)**
AUSTRALIA	White noise	**1.00 (0.90, 1.12)**	**1.01 (0.92, 1.15)**	**1.01 (0.93, 1.15)**
	Bloomfield	**1.08 (0.89, 1.41)**	**1.02 (0.86, 1.40)**	**1.04 (0.85, 1.38)**
ARGENTINA	White noise	**0.97 (0.86, 1.12)**	**0.93 (0.84, 1.09)**	**1.12 (0.96, 1.29)**
	Bloomfield	**0.92 (0.73, 1.21)**	**0.75 (0.68, 1.13)**	**0.85 (0.69, 1.08)**

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.

FIGURE 1: Estimates of d_m and d_y in case of no structural break

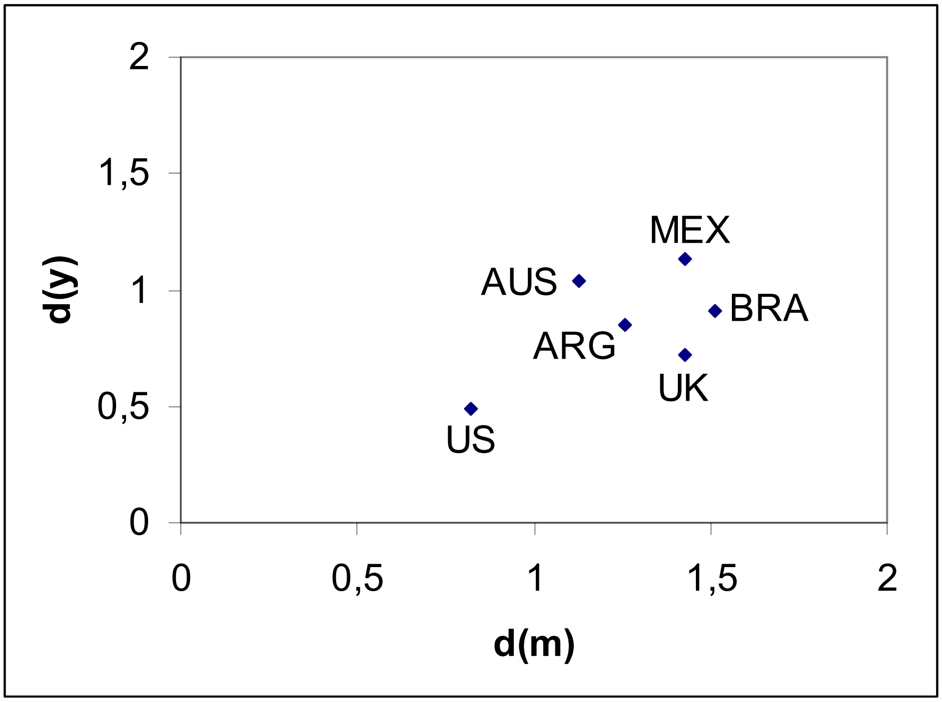


TABLE 3: Estimates of d_m and d_y in case of no structural break		
Country	d_m (money)	d_y (output)
UNITED STATES	0.82*	0.49
UNITED KINGDOM	1.43	0.72
MEXICO	1.43	1.13*
BRAZIL	1.51	0.91*
AUSTRALIA	1,13*	1.04*
ARGENTINA	1.26	0.85*

*: The unit root hypothesis cannot be rejected at the 5% level.

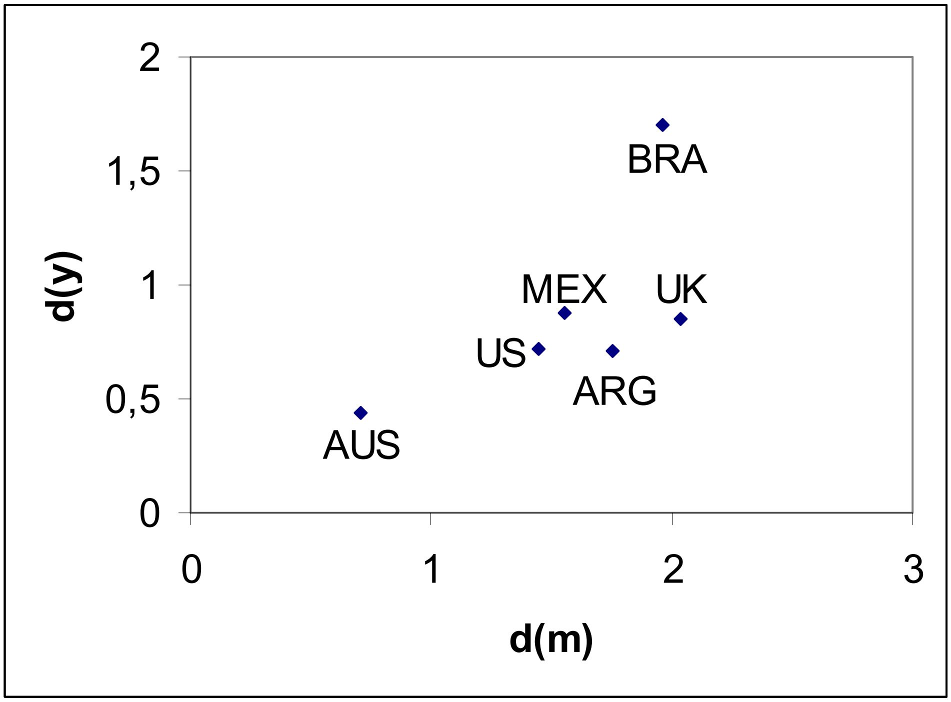
TABLE 4: Estimates of the parameters in the model with a single break: MONETARY AGGREGATES							
Country	Time break	First sub-sample			Second sub-sample		
		d_1	α_1	β_1	d_2	α_2	β_2
UNITED STATES	**1920**	**1.45 (1.12, 1.83)**	**-1.4256 (-30.13)**	**0.0639 (2.08)**	**1.89 (1.60, 2.13)**	0.2266 (0.078)	0.0312 (0.57)
UNITED KINGDOM	**1919**	**2.04 (1.79, 2.29)**	**0.7296 (32.26)**	**0.0963 (3.04)**	**1.86 (1.64, 2.11)**	**3.4958 (1.84)**	-0.0187 (-0.49)
MEXICO	**1986**	**1.55 (1.38, 1.71)**	**-1.8626 (-18.14)**	**0.2372 (2.77)**	**2.02 (1.21, 2.97)**	-9.1703 (-1.28)	**0.3533 (2.74)**
BRAZIL	**1965**	**1.96 (1.64, 2.18)**	**-14.378 (-195.0)**	-0.0495 (-0.48)	**1.85 (1.42, 2.64)**	-29.397 (-0.72)	0.4329 (0.58)
AUSTRALIA	**1883**	**0.71 (0.40, 1.56)**	**3.851 (74.74)**	**0.0681 (8.66)**	**1.28 (1.12, 1.41)**	**4.2135 (14.63)**	**0.0551 (2.84)**
ARGENTINA	**1987**	**1.75 (1.58, 1.94)**	**-11.205 (-58.51)**	**-382.81 (-4.63)**	**2.51 (1.19, 2.98)**	0.2085 (0.93)	**3.8053 (4.82)**

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.

TABLE 5: Estimates of the parameters in the model with a single break: REAL INCOME							
Country	Time break	First sub-sample			Second sub-sample		
		d_1	α_1	β_1	d_2	α_2	β_2
UNITED STATES	**1931**	**0.72** **(0.63, 1.02)**	**1.127** **(19.11)**	**0.0341** **(11.82)**	**1.18** **(1.06, 1.62)**	0.493 (0.78)	**0.0383** **(3.87)**
UNITED KINGDOM	**1918**	**0.85** **(0.77, 1.25)**	**2.8034** **(109.00)**	**0.0194** **(8.61)**	**1.03** **(0.87, 1.48)**	**2.6378** **(13.31)**	**0.0198** **(4.97)**
MEXICO	**1982**	**0.88** **(0.78, 1.28)**	**2.2650** **(91.00)**	**3.8344** **(15.01)**	**0.80** **(0.69, 1.57)**	**0.0601** **(25.98)**	**0.0278** **(5.85)**
BRAZIL	**1930**	**1.70** **(1.04, 2.42)**	**12.429** **(187.66)**	-0.0252 (-0.34)	**1.25** **(1.01, 1.58)**	**11.6921** **(41.01)**	**0.0549** **(3.81)**
AUSTRALIA	**1891**	**0.44** **(0.19, 0.78)**	**1.5566** **(38.14)**	**0.0450** **(14.35)**	**0.92** **(0.65, 1.07)**	**1.5951** **(16.53)**	**0.0306** **(8.76)**
ARGENTINA	**1913**	**0.71** **(0.58, 1.04)**	**21.704** **(390.8)**	**0.0494** **(10.51)**	**0.83** **(0.69, 1.35)**	**22.243** **(185.43)**	**0.0270** **(8.38)**

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.

FIGURE 2: Estimates of d_m and d_y in case of the first subsample

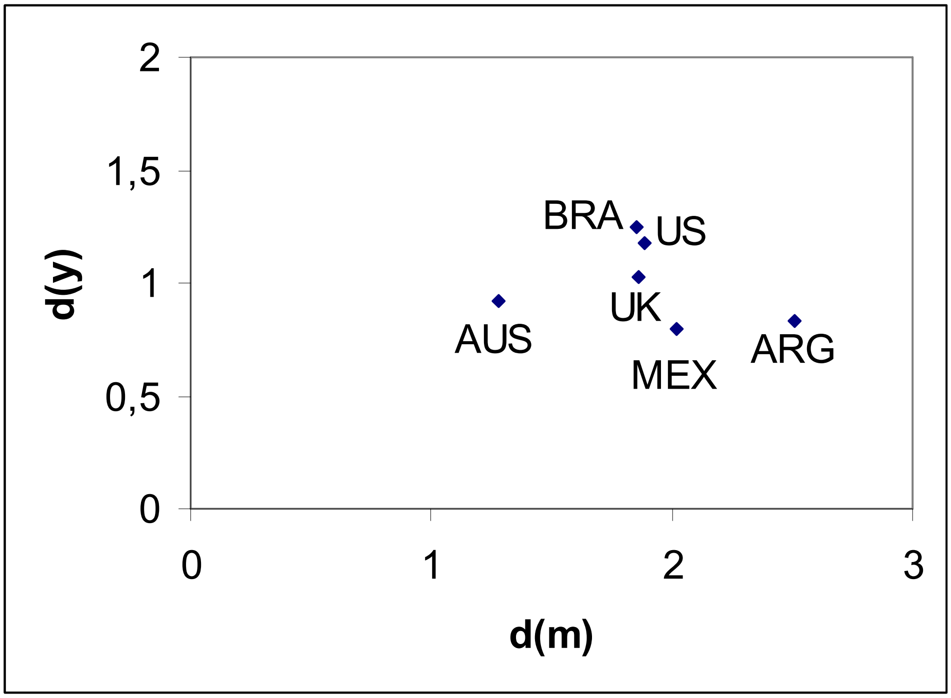


d(m) refers to the order of integration of money, while d(y) is the order of integration of output.

TABLE 6: Estimates of d_m and d_y in case of the first subsample		
Country	d_m (money)	d_y (output)
UNITED STATES	1.45	0.72^*
UNITED KINGDOM	2.04	0.85^*
MEXICO	1.55	0.88^*
BRAZIL	1.96	1.70
AUSTRALIA	0.71^*	0.44
ARGENTINA	1.75	0.71^*

*: The unit root hypothesis cannot be rejected at the 5% level.

FIGURE 3: Estimates of d_m and d_y in case of the second subsample



d(m) refers to the order of integration of money, while d(y) is the order of integration of output.

TABLE 7: Estimates of d_m and d_y in case of the second subsample		
Country	d_m (money)	d_y (output)
UNITED STATES	1.89	1.18
UNITED KINGDOM	1.86	1.03*
MEXICO	2.02	0.80*
BRAZIL	1.85	1.25
AUSTRALIA	1.28	0.92*
ARGENTINA	2.51	0.83*

*: The unit root hypothesis cannot be rejected at the 5% level.

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