Long-Run Monetary Neutrality: Evidence from High Inflation Countries

Erdinc TELATAR – Tarkan CAVUSOGLU***

Abstract

The aim of this paper is to investigate issues of long-run neutrality and long-run superneutrality of money using data of high inflation countries (Argentina, Brazil, Ecuador, Mexico, Uruguay and Turkey). It is found that money is long-run neutral but not superneutral with respect to real output for Argentina and Uruguay indicating that money growth has a negative effect on real output. The long-run superneutrality holds for Brazil, Mexico and Turkey. The long-run neutrality is rejected for Ecuador.

Keywords: *long-run neutrality, inflation*

JEL Classifications: E31, E51

1. Introduction

The issue of neutrality of money, which originates from the quantity theory of money, has been the heart of debates over the real effects of macroeconomic policies. There are two branches of the studies on the neutrality proposition differing with respect to the definition of policy instrument. One emphasizes a permanent change in the level of money, while the other in the rate of growth of money. The hypothesis which states that a permanent change of money has no effect on real variables is called the long-run neutrality of money (LRN). The other quantity-theoretical proposition is that of long-run superneutrality of money (LRSN), indicating that a change in the growth rate of nominal money has no effect on real variables.

* Erdinc TELATAR, Hacettepe University, Department of Economics, Ankara, Turkey; e-mail: etelatar@hacettepe.edu.tr

** Tarkan CAVUSOGLU, Hacettepe University, Department of Public Finance, Ankara, Turkey; e-mail: tarkan.cavusoglu@hacettepe.edu.tr

The studies by King and Watson (1992) and Fisher and Seater (1993) on US data have developed reduced-form models to test the hypotheses based on the theory of non-stationary regressors and have provided little evidence of the LRN. The main contribution of these studies to empirical literature is that both monetary and real variables must satisfy certain non-stationarity conditions in order to conduct meaningful neutrality tests. In particular, they show that if both monetary and real variables are integrated of order one, the long-run neutrality can be tested. On the other hand, superneutrality tests require that the order of monetary variables is equal to one plus the order of integration of the real variables. Fisher and Seater (1993) demonstrate that much of the older studies violate these conditions and hence be disregarded.[1]

Empirical studies using Fisher and Seater's (1993) framework have been conducted on the data for developed countries, particularly for the United States, the European countries, Canada and Japan. For example, Boschen and Otrok (1994) find supportive evidence for the LRN for the US using dummy variables for the Great Depression. Similarly, Weber (1994) reports some evidence in favor of the LRN for G-7 countries, and Haug and Lucas (1997) and Koustas (1998) for Canada. Serleties and Koustas (1998) and Serleties and Krause (1996) conclude that Backus and Kheo's (1992) long annual international data are generally supportive of the LRN. There are a negligible number of studies in regard to developing countries; Bae and Ratti (2000) for Argentina and Brazil and Wallace (1999) for Mexico, which concluded that the LRN of money could not be rejected.

The aim of the paper is to reexamine the LRN and LRSN of money using data of Argentina, Brazil, Ecuador, Mexico, Turkey and Uruguay. These are developing countries and share similar economic and political history with high and volatile rates of inflation, money growth and output growth. Taking into account that the results of the LRN of money depends crucially on the univariate time series properties of the data, we perform a number of unit root tests recently developed to obtain robust results. The results on the LRN of money are supported by data for all of the analyzed countries, except for Ecuador. However, we find mixed results for the LRSN of money; it is rejected for Argentina and Uruguay but not rejected for Brazil Mexico and Turkey. A rise in the money growth decreases the level of real output in Argentina and Uruguay, which is the opposite of the Tobin effect where a rise in the growth rate of money lowers the real interest rate and increases capital (Tobin, 1965).

[1] For a detailed review of the older literature, see Weber (1994).

Section 2 presents an overview of the econometric methodology developed by Fisher and Seater (1993). The data and the univariate time series properties of the data are described in Section 3. The empirical results of LRN and LRSN of money with respect to output are reported in Section 4. Section 5 presents concluding remarks.

2. The Methodology

Our testing procedure of LRN and LRSN is based on Fisher and Seater's (1993) bivariate ARIMA representation in which money and real output are modelled with a log-linear system as follows:

$$\begin{aligned} a(L)\Delta^{\langle m\rangle} m_t &= b(L)\Delta^{\langle y\rangle} y_t + u_t \\ d(L)\Delta^{\langle y\rangle} y_t &= c(L)\Delta^{\langle m\rangle} m_t + w_t \end{aligned} \tag{1}$$

In the system, $a(L)$, $b(L)$, $c(L)$ and $d(L)$ are polynomials in the lag operator L while $\Delta^{\langle m\rangle}$ and $\Delta^{\langle y\rangle}$ denote the orders of integration of the difference-stationary variables money (m_t) and real output (y_t), respectively; u_t and w_t are independently and identically distributed disturbance terms with zero mean and Σ covariance matrix of which the elements are σ_{uu} σ_{ww} and σ_{uw}.

Fisher and Seater (1993) define LRN and LRSN in terms of specific values of the long-run derivative (LRD) of z with respect to a permanent change in x:

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

where

$$x_t \equiv \Delta^i m_t,$$

$$z_t \equiv \Delta^i y_t.$$

The limit of the ratio indicates the ultimate effect of an exogenous monetary disturbance on the real output relative to the ultimate effect of that disturbance on the money itself. In this framework, LRN can be tested when $\langle m\rangle = \langle y\rangle = 1$ implying $\text{LRD}_{y,m} = c(1)/d(1)$ whereas LRSN can be tested when $\langle m\rangle = 2$ and $\langle y\rangle = 1$ implying $\text{LRD}_{y,\Delta m} = c(1)/d(1)$. In the latter, the rejection of the LRSN hypothesis does not rule out the validity of LRN. In other words, LRN is necessary but not sufficient for LRSN.

In Fisher and Seater's (1993) framework, tests of LRN and LRSN are based on estimating LRD, that is $c(1)/d(1)$. Under the assumption of the long-run exogeneity of the money stock, denoted by $b(1) = \sigma_{uw} = 0$, the coefficient $c(1)/d(1)$ is estimated as the frequency-zero regression coefficient in the regression of $\Delta^{\langle y \rangle} y$ on $\Delta^{\langle m \rangle} m$. This coefficient estimate is obtained by $\lim_{k\to\infty} b_k$ where b_k is the OLS estimate in the following regression:

$$\left[\sum_{j=0}^{k} \Delta^{\langle y \rangle} y_{t-j}\right] = a_k + b_k \left[\sum_{j=0}^{k} \Delta^{\langle m \rangle} m_{t-j}\right] + e_{kt} \tag{2}$$

When the orders of the integration of the variables are $\langle m \rangle = \langle y \rangle = 1$, the relevant regression to test LRN is

$$(y_t - y_{t-k-1}) = a_k + b_k (m_t - m_{t-k-1}) + e_{kt} \tag{3}$$

and when $\langle m \rangle = 2$ and $\langle y \rangle = 1$, the relevant regression to test LRSN becomes

$$(y_t - y_{t-k-1}) = a_k + b_k (\Delta m_t - \Delta m_{t-k-1}) + e_{kt} \tag{4}$$

The testing procedure exploits the estimates of b_k up to a certain lag-length k within confidence bands corrected by using Newey and West's (1987) procedure. The bands excluding the zero constitute evidence of the non-neutrality of money with respect to the real output. The next section provides the univariate time series properties of the data.

3. The Data and Unit Root Tests

3.1. The Data

The money stock and the real output variables in the data set are annual observations obtained mostly from the IMF International Financial Statistics (IFS) database except for Turkey of which the data are from the State Institute of Statistics in Turkey. The data set covers the periods 1960 – 2002 for Argentina, 1948 – 2002 for Brazil, Mexico and Uruguay, 1948 – 2001 for Ecuador, and 1949 – 2002 for Turkey. The figures of the broad monetary aggregate *money plus quasi-money* and the real output measure *GDP volume index* with 1995 base year are taken from the IFS database. The missing observations of real GDP figures in IFS have been completed according to the growth rates of the real GDP figures given in the Oxford Latin American Economic History Database on the web page <http://oxlad.qeh.ox.ac.uk/>.

In order to provide an overview of the data, Table 1 (see Appendix) gives a summary of the growth rates of the money stock and the real output of the six countries together with their rates of price inflation. With respect to the five-year averages given in Table 1, it is observed that the six countries in question have been living with chronic inflation for years. Ecuador, Mexico and Turkey have begun to be exposed to persistent two-digit inflation since the 1970s whereas the experiences of Argentina, Brazil and Uruguay were older. However, Ecuador seems to be the sole country that keeps inflation moderate at two-digit rates compared to others. Although the two-digit rates of inflation that began to appear first in 1973 in Ecuador range between 10 to 96 per cent, the average rate of consumer price inflation was 14 per cent between 1973 – 1982 while it increased to 33 per cent from 1983 onwards. While Mexico, Turkey and Uruguay have seldom experienced triple-digit rates, Argentina in the 1975 – 1991 period and Brazil in 1980 – 1994 period experienced persistent triple-digit inflation rates, where even four-digit rates were recorded meanwhile in both countries. However, in Argentina, after a peak rate of consumer price inflation of 3 080 per cent in 1989 followed by a rate 2 314 per cent in 1990, inflation has been reduced drastically to one-digit rates from 1994 onwards. Moreover, negative annual rates of –1 per cent were achieved consecutively in the years 1999, 2000 and 2001. Similarly in Brazil, the years of four-digit inflation, among which there was a peak of 2 704 per cent whole sale price inflation in 1990, have been followed by years of modest inflation from 1995 onwards and inflation was reduced to a minimum of 4 per cent in 1998.

The money stock growth rates in these economies have almost gone along with the rates of inflation, except in Ecuador where there seems to be no observable correlation between these two rates. The economic growth rates in most of these countries have been relatively high on the average but not steady over time. Especially in Argentina, Brazil, Ecuador and Mexico, the growth performances of the economies have deteriorated sharply since the 1980s, which coincides with the years of persistently soaring inflation rates.

3.2. Unit Root Tests

The univariate time series characteristics of the money stock and real output variables play a crucial role in testing the LRN and LRSN hypotheses. The precise inference on the orders of the integration of the variables is a prerequisite for a correctly specified bivariate model of the money and output. In this sense, Table 2 in Appendix exhibits some basic unit root tests to determine the stationarity characteristics of the series in question. The first is a conventional unit root test (ADF) suggested by Dickey and Fuller (1979) whereas the second is a modified version of the ADF test (DF-GLS) proposed by Elliot, Rothenberg and Stock (1996). The GLS estimation with local detrending is argued to outperform the Dickey-Fuller (DF) test in terms of small-sample size and power. The third test suggested by Kwiatkowski, Phillips, Schmidt, and Shin (1992) is a Lagrange Multiplier test (KPSS) of the null hypothesis of stationarity against the alternative of a unit root, unlike the most unit root tests that treat non-stationarity as the null hypothesis. Such characteristic of the KPSS is argued to enable a confirmatory analysis in unit root testing. The last two unit root test statistics denoted by $PP - Z_t$ and MSB are respectively the Phillips-Perron (PP) (1988) and Bhargava (1986) tests modified by Ng and Perron (2001). The modifications are argued to provide improvements in power characteristics and size distortions of the tests.

According to the figures in Table 2, the statistics computed for Argentina indicate that the money stock series is generated by an I(2) process while the real output series is I(1) at 1 per cent significance level, except for the trend included $PP - Z_t$ and MSB statistics which are significant at 5 per cent level. Such orders of integration allow testing the validity of the LRSN hypothesis for Argentina.

The unit root statistics computed for Brazil's money stock data are in favor of an I(2) data generating process at 1 per cent significance level, except for the trend included KPSS statistic indicating an I(1) process. However, in the replication of the tests with the modified version of the Schwartz Information Criteria, the money stock series appears to be I(2), as expected, at 1 per cent significance level both with and without a deterministic trend included in the analysis (not reported). The real output series of Brazil seems to be I(1) at 1 per cent significance level as we ignore the inconsistent results of the 'no trend' augmented Dickey-Fuller (ADF) and KPSS statistics. The inconsistency is because real output is stationary in levels at 5 per cent significance level in the ADF test and non-stationary in first-differences at 1 per cent significance level in the KPSS test. The final decision for Brazil that the money stock is I(2) and the real output is I(1) allows testing the LRSN hypothesis.

Ecuador's money stock series exhibits an I(1) characteristic at 1 per cent significance level according to all but the trend included $PP - Z_t$ and MSB statistics which are significant at 5 per cent significance level. The real output series are decided to be I(1) despite the rejection of the stationarity hypothesis at 5 per cent significance level in the KPSS test as it turns out to fail to reject stationarity when the estimation method in KPSS test is changed to a kernel-based one exploiting the Newey-West type of bandwith selection (not reported). As a result of the decision on the I(1) characterized money stock and real output series, it is possible to test the LRN hypothesis for Ecuador.

The test results for Mexico indicate that the money stock series is I(2) at 1 per cent significance level with respect to the first three of the five tests statistics. The last two statistics PP – Z_t and MSB exhibit such an inconsistent outcome with the kernel-based estimation methods that the second difference of the series appears non-stationary while the first difference is stationary (as seen in Table 2). However, the series turns out to be consistently I(2) under autoregressive spectral density estimation methods (not reported). On the other hand, the real output series of Mexico exhibits an I(1) characteristic at 1 per cent significance level, despite the inconsistency appearing in the 'no trend' case of the ADF and KPSS statistics which cannot be addressed anyway. Without loss of generality, the findings that the money stock is I(2) and the real output is I(1) constitute evidence on the testability of the LRSN hypothesis in the Mexico case.

The unit root tests computed for Turkey support an I(2) money stock series. The rejection of the non-stationarity hypothesis at 5 per cent significance level for the first difference of the money stock series in the trend included DF-GLS test statistics is reversed to non-rejection when the lag-length selection method is changed to Akaike Information Criteria or to the modified versions of the information criteria. The real output series in Turkey is found to be I(1) except for the case where the KPSS statistic with no trend indicates non-stationarity of the first-differenced series at 5 per cent significance level. Overall, the relevant orders of integration pointed out by the unit root tests allow testing LRSN of money in the Turkish economy.

Finally, in the Uruguay case, the money stock and the real output data are found to be generated by I(2) and I(1) processes, respectively. Although the 'no trend' case of the PP – Z_t statistic appears to indicate stationarity of the first-differenced money stock series at 5 per cent significance level, the autoregressive spectral density estimation of the statistic (not reported) support the money stock series to be I(2) in accordance with the findings of the other test statistics computed. Orders of integration determined for the money stock and the real output series of Uruguay provide evidence for testing the LRSN hypothesis.

The next step to the determination of the orders of integration is to construct and estimate the relevant regressions represented by Equation (3) or Equation (4) in order to test the LRN hypothesis for Ecuador and the LRSN hypotheses for the rest of the six countries in question.

4. The Test Results of Long-Run Neutrality

Tests of the LRN and LRSN hypotheses are performed with the slope coefficients b_k estimated for $k = 1,\ldots,10$ in the regressions given in Equation (3) and Equation (4) respectively. Unlike OLS estimates used in the relevant literature, the $\hat{b}_k$ coefficients in the study are Generalized Method of Moments (GMM) estimates that provide efficiency gains in case of the money stock endogeneity, without requiring information about the exact distribution of the disturbances. The estimates of b_k are evaluated within 95 per cent confidence bands[2] to perform the tests of the LRN and LRSN hypotheses following the Fisher and Seater (1993) approach.

The coefficient estimates given in Table 3 (see Appendix) for the six countries in question are based on the orthogonality conditions imposed through GMM instruments, i.e., a constant, the explanatory variable of the relevant regression and its one-period lag. The validity of these instruments is tested by the J-statistic[3] suggested by Hansen (1982). According to the computations presented in Table 3, the validity of the GMM instruments in the estimations cannot be rejected for all of the six countries, except for the J-statistic which is hardly significant only at $k = 9$ for Turkey at 5 per cent significance level but even insignificant at 1 per cent significance level. The estimated slope coefficients $\hat{b}_k$ for Brazil, Mexico and Turkey are found to be statistically insignificant, reflecting the tumultuous characteristics of these economies due to discretionary policies. The rate of growth of the monetary expansion appears to have no effect on the real output in these economies. However, for Argentina up to $k = 8$ and for Uruguay for all k, the estimated slope coefficients are negative and statistically significant at 5 per cent significance level while for Ecuador the estimated coefficients are all significant with positive signs. These findings imply that an increase in the level of money stock is capable of increasing the real output in Ecuador while an increase in the rate of growth of money stock decreases the level of real output in Argentina and Uruguay.

The plots of the 95 per cent confidence bands constructed around $\hat{b}_k$ by using the t-distribution with n/k degrees of freedom are shown in Figure 1 to 6 (in Appendix) as tests for the long-run neutrality of money. For Argentina (see Figure 1), upper bound

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[2] Confidence bands are corrected by Newey and West's (1987) heteroscedasticity and autocorrelation consistent covariance estimator.

[3] The J-statistic has an asymptotically χ^2-distribution with as many degrees of freedom as the number of over-identifying restrictions.

of the confidence band lies below the zero line at all lags except for $k = 5$ and $k \geq 9$, indicating the rejection of the LRSN hypothesis.[4] A similar outcome is reported by Bae and Ratti (2000) for Argentina. The Uruguay case (see in Figure 6) is the other evidence against the superneutrality of money with respect to the real output, as the upper bound of the confidence band lies below the zero line at all lags.

For Brazil, Mexico and Turkey (see Figure 2, 4 and 5 respectively), LRSN hypotheses could not be rejected since the confidence bands in each case includes zero at all lags, ignoring the case where the lower bound of the confidence band at $k = 9$ is above the zero line for Mexico. Hence, the data of these countries support both the LRN and LRSN of money. The study by Wallace (1999) covering the period 1932 – 1992 also presents supporting findings for the LRN of money in Mexico, however, asserting only the LRN as a testable hypothesis with respect to ADF tests on the data. On the other hand, our non-rejection of the LRSN hypothesis for Brazil is in conflict with the conclusion of Bae and Ratti (2000) for Brazil, which may be attributed to the differences in the sample size.

The LRN hypothesis tested for Ecuador (see Figure 3) is rejected because the lower bound of the confidence band is well above the zero line at all lags with statistically significant and positive $\hat{b}_k$ coefficients. Similar long-run non-neutrality evidence is depicted by Fisher and Seater (1993) for the United States and by Olekalns (1996) for Australia with a broad monetary aggregate.

Conclusion

In this paper we have investigated the issues of the long-run neutrality and long-run superneutrality of money using the annual data for Argentina, Brazil, Ecuador, Mexico, Uruguay and Turkey. This is an interesting issue for these countries because they are all developing countries that have high and volatile inflation, money growth and output growth rates. We base our testing procedure on the Fisher and Seater's (1993) bivariate ARIMA representation in which the money and real output are modelled with a log-linear system. We pay explicit attention to the order of integration of the variables by applying a number of recent unit root tests since meaningful neutrality tests crucially depend on such properties.

[4] The insignificant $\hat{b}_k$ coefficients at $k = 9$ and $k = 10$ are ignored while deciding the rejection of the LRSN hypothesis, because the power of the tests is likely to be low for larger values of k, which correspond to small degrees of freedom, as noted by Boschen and Otrok (1994, p. 1472).

The results show that the data are generally supportive for the long-run neutrality of money except for Ecuador. Superneutrality of money is rejected for Argentina and Uruguay. The data from Brazil, Mexico and Turkey support the long-run superneutrality of money. In the cases of Argentina and Uruguay, we find that a rise in the rate of growth of money has a negative effect on output. Ratti and Bae (2000) report similar results for Argentina and concludes that this finding is the opposite of the Tobin effect and consistent with the prediction of cash-in-advance models in which inflation is a tax on investment or labor. Findings for Brazil, Mexico and Turkey are consistent with the McCallum's (1990) definition of superneutrality stating that a permanent change in the growth rate of money has no effect on real variables in the long-run except for real cash balances. These results imply that economic agents in Brazil, Mexico and Turkey have learnt how to live with and to hedge against high inflation, making monetary policy ineffective in the long-run.

Appendix

Table 1

Average Rates of Money Growth (μ_t), Economic Growth (ρ_t) and Inflation (π_t)

Period	Argentina			Brazil			Ecuador		
	μ_t	ρ_t	π_t	μ_t	ρ_t	π_t	μ_t	ρ_t	π_t
1950 – 1955	–	2.9	20.2	17.9	6.8	16.8	9.1	5.8	1.9
1956 – 1960	–	3.1	42.1	29.6	6.9	22.9	7.5	4.6	–0.2
1961 – 1965	22.3	4.5	23.3	65.1	8.6	61.6	4.8	5.4	4.0
1966 – 1970	31.0	4.2	19.7	35.3	6.4	25.7	11.9	4.4	5.0
1971 – 1975	80.2	3.9	72.2	36.9	10.1	22.4	24.5	11.6	13.6
1976 – 1980	206.5	1.5	211.2	52.8	7.2	57.1	26.0	6.5	11.7
1981 – 1985	344.8	–1.9	382.4	180.1	1.5	166.8	–2.1	2.2	28.1
1986 – 1990	813.5	–0.2	1 191.6	953.2	3.0	1 003.3	–0.4	2.2	47.0
1991 – 1995	53.0	5.8	43.0	1 284.1	3.1	1 161.7	20.2	3.5	39.7
1996 – 2002	8.8	–0.3	3.5	12.5	2.4	12.0	2.1	1.1	41.4

Period	Mexico			Turkey			Uruguay		
	μ_t	ρ_t	π_t	μ_t	ρ_t	π_t	μ_t	ρ_t	π_t
1950 – 1955	16.2	6.7	8.7	19.5	8.4	4.7	9.4	4.1	8.6
1956 – 1960	10.0	6.2	5.9	17.7	4.7	14.5	20.5	0.0	23.4
1961 – 1965	12.4	7.2	1.9	13.8	4.8	5.2	37.7	0.9	30.8
1966 – 1970	12.2	6.9	3.6	18.4	6.0	6.5	46.8	2.3	65.1
1971 – 1975	18.1	6.6	12.2	27.1	5.8	20.1	70.4	1.6	71.2
1976 – 1980	54.1	6.7	21.4	44.1	2.5	50.2	86.2	4.5	56.7
1981 – 1985	55.6	2.0	62.4	57.0	4.9	40.0	48.8	–2.6	46.0
1986 – 1990	80.1	1.8	75.8	54.7	5.6	54.4	91.5	3.9	79.0
1991 – 1995	27.8	1.6	18.0	79.4	3.3	76.9	49.7	4.0	62.3
1996 – 2002	14.8	4.0	15.5	78.0	3.0	62.9	22.7	–0.5	12.5

Notes: (1) The data on the economic growth rate of Argentina begin from 1952.
(2) For Ecuador, the data on the rate of inflation begin from 1952 and data on the money growth rate are not available for 2002.
(3) The average rates of inflation are based on the Consumer Price Index (1995 = 100) for Argen-tina, Ecuador, Mexico and Uruguay, the Wholesale Price Index (1995 = 100) for Brazil and the GNP Deflator (1987 = 100) for Turkey.

Table 2

Unit Root Tests

		ADF		DF–GLS		KPSS		PP–Z_t		MSB		
Country	Variable	no trend	trend	no trend	trend	no trend	trend	no trend	trend	no trend	trend	Decision
Argentina	M	–0.93(3)	–2.22(3)	–0.69(3)	–2.13(3)	10.8(3)**	0.49(3)**	1.26(9.8)	–1.55(10.3)	1.22(9.8)	0.31(10.3)	
	ΔM	–1.86(2)	–1.74(2)	–1.61(2)	–1.75(2)	1.92(2)**	1.65(2)**	–1.83(3.07)	–1.88(3)	0.27(3.07)	0.25(3)	$M \rightarrow I(2)$
	$\Delta^2 M$	–6.42(1)**	–6.42(1)**	–6.32(1)**	–6.54(1)**	0.03(1)	0.01(0)	–3.24(1.54)**	–3.25(1.53)*	0.15(1.54)**	0.15(1.53)*	
	Y	–2.12(0)	–1.47(0)	–0.27(0)	–1.39(0)	81.2(0)**	1.87(0)**	0.32(2.24)	–1.23(2.43)	0.98(2.24)	0.29(2.43)	$Y \rightarrow I(1)$
	ΔY	–4.58(0)**	–4.79(0)**	–4.42(0)**	–4.75(0)**	0.34(0)	0.08(0)	–2.81(1.01)**	–2.94(0.944)*	0.15(1.01)**	0.16(0.944)*	
Brazil	M	–0.64(1)	–2.41(1)	–0.61(1)	–2.23(1)	16.6(1)**	2.66(1)**	2.35(14.4)	–0.79(13.7)	1.19(14.4)	0.51(13.7)	
	ΔM	–1.98(0)	–2.73(4)	–1.88(0)	–2.78(4)	5.30(0)**	0.07(4)	–1.86(2.3)	–2.03(2.37)	0.27(2.3)	0.23(2.37)	$M \rightarrow I(2)$
	$\Delta^2 M$	–6.18(0)**	–6.14(0)**	–6.22(0)**	–6.26(0)**	0.09(0)	0.07(0)	–3.57(0.756)**	–3.57(0.743)**	0.14(0.756)**	0.14(0.743)**	
	Y	–3.11(0)*	0.14(0)	–0.04(4)	–0.06(0)	37.9(4)**	23.4(0)**	5.08(1.16)	0.23((0.778)	3.83(1.16)	0.65((0.778)	$Y \rightarrow I(1)$
	ΔY	–5.79(0)**	–6.79(0)**	–5.82(0)**	–6.80(0)**	1.13(0)**	0.13(0)	–3.52(0.385)**	–3.61(0.0746)**	0.14(0.385)**	0.14(0.0746)**	
Ecuador	M	–0.87(1)	–2.69(3)	0.57(1)	–2.83(3)	153(1)**	0.26(3)**	1.44(2.92)	–1.82(3.16)	1.27(2.92)	0.26(3.16)	
	ΔM	–5.37(0)**	–5.34(0)**	–5.39(0)**	–5.43(0)**	0.16(0)	0.12(0)	–3.45(0.114)**	–3.46(0.18)**	0.14(0.114)**	0.14(0.18)**	$M \rightarrow I(1)$
	$\Delta^2 M$	–9.07(1)**	–8.98(1)**	–8.60(1)**	–9.01(1)**	0.01(1)	0.01(1)	–2.96(1.97)**	–2.97(1.95)*	0.17(1.97)**	0.17(1.95)*	
	Y	–2.01(0)	0.12(0)	0.76(1)	–0.73(1)	237(1)**	6.55(1)**	3.65(1.75)	–0.33(1.51)	2.86(1.75)	0.45(1.51)	$Y \rightarrow I(1)$
	ΔY	–5.86(0)**	–6.33(0)**	–5.58(0)**	–6.04(0)**	0.60(0)*	0.19(0)*	–3.41(0.571)**	–3.40(0.169)*	0.15(0.571)**	0.15(0.169)*	

Mexico	M	–0.36(3)	–1.68(0)	–0.54(3)	–1.83(3)	11.7(3)**	2.79(3)**	4.64(2.79)	–0.73(2.53)	2.43(2.79)	0.64(2.53)	M→I(2)
	ΔM	–1.83(2)	–1.68(2)	–1.74(2)	–1.89(2)	4.48(2)**	1.45(2)**	–3.14(1.37)**	–3.36(1.04)*	0.16(1.37)**	0.14(1.04)*	
	Δ^2M	–9.64(1)**	–9.61(0)**	–12.0(0)**	–9.03(1)**	0.03(0)	0.01(1)	–1.66(2.68)	–1.69(2.71)	0.29(2.68)	0.29(2.71)	
	Y	–3.63(0)**	–2.63(0)	0.49(1)	–0.57(1)	144(1)**	11.4(1)**	5.58(2.55)	0.09(2.44)	4.00(2.55)	0.64(2.44)	Y→I(1)
	ΔY	–4.79(0)**	–8.24(0)**	–4.84(0)**	–5.74(0)**	1.47(0)**	0.11(0)	–3.34(0.277)**	–3.49(0.662)**	0.15(0.277)**	0.14(0.662)**	
Turkey	M	1.59(1)	–0.88(1)	–0.72(2)	–1.99(2)	1.51(2)**	1.24(2)**	12.1(6.73)	1.21(6.31)	4.08(6.73)	1.11(6.31)	M→I(2)
	ΔM	–2.23(0)	–3.17(0)	–1.93(0)	–3.24(0)*	11.8(0)**	0.31(0)**	–1.63(2.75)	–2.43(1.55)	0.31(2.75)	0.18(1.55)	
	Δ^2M	–10.0(0)**	–9.96(0)**	–9.17(0)**	–9.59(0)**	0.07(0)	0.06(0)	–3.14(0.575)**	–3.26(0.561)**	0.15(0.575)**	0.15(0.561)*	
	Y	–2.54(0)	–2.63(0)	0.89(2)	–1.48(0)	166(2)**	2.53(0)**	5.56(2.19)	–0.79(1.28)	3.53(2.19)	0.33(1.28)	Y→I(1)
	ΔY	–7.57(0)**	–8.24(0)**	–3.13(1)**	–8.19(0)**	0.73(1)*	0.04(0)	–3.18(0.66)**	–3.46(0.258)**	0.16(0.66)**	0.14(0.258)**	
Uruguay	M	–0.54(2)	–2.75(2)	–1.42(4)	–2.07(2)	14.9(4)**	1.89(2)**	4.07(7.99)	–0.98(6.99)	2.31(7.99)	0.51(6.99)	M→I(2)
	ΔM	–1.92(1)	–1.67(1)	–1.57(1)	–1.74(1)	6.21(1)**	2.89(1)**	–2.01(3.08)*	–2.59(2.87)	0.25(3.08)	0.19(2.87)	
	Δ^2M	–11.9(0)**	–6.33(2)**	–12.0(0)**	–12.1(0)**	0.06(0)	0.02(0)	–3.14(0.926)**	–3.12(0.768)*	0.16(0.926)**	0.16(0.768)*	
	Y	–1.44(1)	–3.03(1)	–0.40(1)	–3.15(1)	67.9(1)**	0.36(1)**	0.91(4.38)	–2.29(4.68)	1.05(4.38)	0.19(4.68)	Y→I(1)
	ΔY	–3.98(0)**	–3.99(0)*	–4.04(0)**	–4.01(0)**	0.15(0)	0.12(0)	–2.88(0.904)**	–2.96(0.9)*	0.13(0.904)**	0.15(0.9)*	

Notes: (1) The figures in parentheses denote the lag length determined by the Schwartz Information Criteria in the ADF, DF-GLS and KPSS tests whereas these figures in PP-Z_t and MSB tests denote the Andrews bandwith.

(2) KPSS statistics are computed with the GLS detrended autoregressive spectral density estimation method while the PP-Z_t and MSB statistics are computed with the Bartlett Kernel estimation method.

(3) For the KPSS statistic, the null hypothesis is stationarity against the alternative of a unit root.

(4) ** and * denote statistical significance at 1 per cent and 5 per cent levels, respectively.

Table 3

GMM Estimates of the Slope Coefficients

	Argentina[II]		Brazil[II]		Ecuador[I]		Mexico[II]		Turkey[II]		Uruguay[II]	
	$\hat{b}_k$	J-statistic	$\hat{b}_k$	J-statistic	$\hat{b}_k$	J-statistic	$\hat{b}_k$	J-statistic	$\hat{b}_k$	J-statistic	$\hat{b}_k$	J-statistic
k = 1	–0.0331* (0.0101)	1.0071	–0.0247 (0.0144)	1.2114	0.1656* (0.0506)	0.7604	–0.0239 (0.0235)	0.0002	–0.1392 (0.0847)	0.8944	–0.1327* (0.0453)	0.8426
k = 2	–0.0348* (0.0081)	1.1102	–0.0184 (0.0159)	0.4053	0.1533* (0.0488)	1.4480	–0.0070 (0.0302)	0.0059	–0.0630 (0.0767)	0.2173	–0.0986* (0.0367)	0.4219
k = 3	–0.0505* (0.0116)	0.6448	–0.0105 (0.0234)	0.1642	0.1578* (0.0463)	0.4946	–0.0056 (0.0385)	0.2896	–0.0450 (0.0845)	0.1930	–0.1287* (0.0554)	1.3478
k = 4	–0.0548* (0.0162)	0.7027	–0.0016 (0.0251)	0.0502	0.1645* (0.0475)	0.4794	–0.0124 (0.0396)	0.9594	0.0741 (0.0821)	0.2472	–0.0882* (0.0407)	0.5578
k = 5	–0.0269 (0.0173)	2.2054	0.0016 (0.0279)	0.0730	0.1809* (0.0499)	0.7577	0.0618 (0.0480)	0.6131	–0.0126 (0.0609)	0.0220	–0.1263* (0.0461)	0.9919
k = 6	–0.0419* (0.0180)	1.8760	0.0042 (0.0299)	0.0509	0.1945* (0.0486)	0.6721	0.0716 (0.0578)	1.3575	–0.0467 (0.1253)	0.9754	–0.1418* (0.0476)	0.4992
k = 7	–0.0556* (0.0226)	0.9711	0.0116 (0.0384)	0.1577	0.2098* (0.0465)	0.2607	0.0676 (0.0743)	2.1020	0.0086 (0.0816)	2.7829	–0.1951* (0.0618)	1.4086
k = 8	–0.0578* (0.0252)	0.4366	0.0202 (0.0443)	0.8103	0.2336* (0.0426)	0.1118	0.1047 (0.0749)	2.0492	–0.0362 (0.0731)	0.8175	–0.1931* (0.0479)	0.5020
k = 9	–0.0504 (0.0265)	0.2120	0.0244 (0.0509)	0.8448	0.2585* (0.0366)	0.3328	0.2086* (0.0873)	1.4600	–0.0227 (0.1395)	4.0042*	–0.1906* (0.0554)	0.0869
k = 10	–0.0397 (0.0245)	0.1527	0.0193 (0.0578)	0.6571	0.2819* (0.0302)	0.0291	0.1148 (0.0770)	2.1952	0.0051 (0.0544)	2.1216	–0.1421* (0.0530)	0.5470

Notes: (1) * Denotes statistical significance at 5 per cent level.
(2) Figures in parentheses under the estimates of slope coefficients are heteroskedasticity and autocorrelation consistent standard errors.
(3) J-statistic has a χ^2-distribution with degrees of freedom equal to 1.
(4) The country estimates in the table are based on the following regressions:

I $(y_t - y_{t-k-1}) = \hat{a}_k + \hat{b}_k(m_t - m_{t-k-1}) + \hat{e}_{kt}$

II $(y_t - y_{t-k-1}) = \hat{a}_k + \hat{b}_k(\Delta m_t - \Delta m_{t-k-1}) + \hat{e}_{kt}$

Figure 1
Argentina (1960 – 2002)

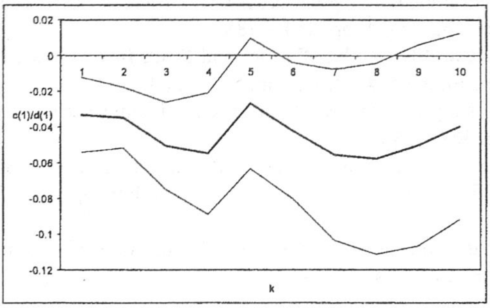


Figure 2
Brazil (1948 – 2002)

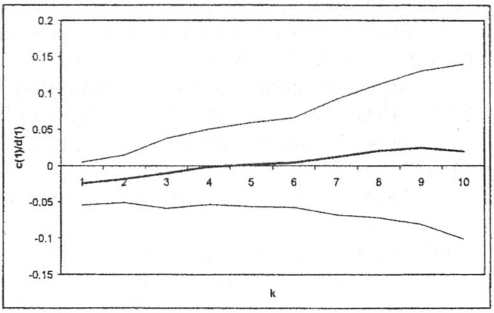


Figure 3
Ecuador (1948 – 2001)

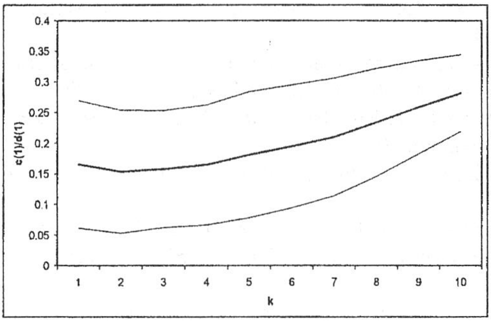


Figure 4
Mexico (1948 – 2002)

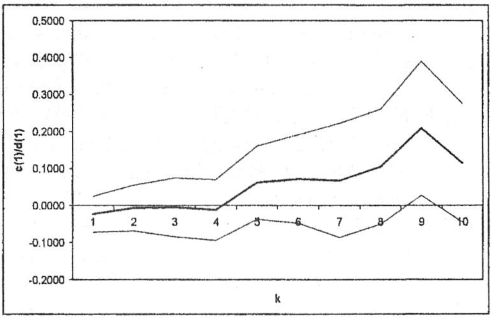


Figure 5
Turkey (1949 – 2002)

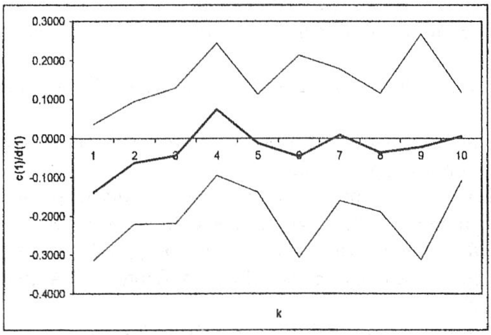
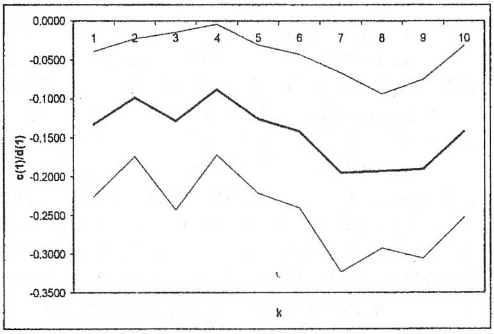


Figure 6
Uruguay (1948 – 2002)



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