# Testing for Long Run Neutrality of Money in Mexico

by

**Gary L. Shelley***
**Department of Economics, Finance, and Urban Studies**
**East Tennessee State University**
**Johnson City, Tennessee 37614**

**and**

**Frederick H. Wallace**
**Departamento de Ciencias Económico-Administrativas**
**Universidad de Quintana Roo**
**Chetumal, Quintana Roo**
**Mexico**

* **Corresponding Author:**
**Dr. Gary L. Shelley**
**Dept. of Economics, Finance, and Urban Studies**
**East Tennessee State University**
**Johnson City, TN 37614**
**Phone: (423) 439-5139 or (423) 439-4202**
**Email: Shelley@mail.etsu.edu**

**Testing for Long Run Neutrality of Money in Mexico**

**JEL classification: E31, E52**

**Keywords: money, neutrality, Mexico**

## Abstract

The Fisher-Seater methodology is used to investigate long run money neutrality in Mexico from 1932-2001. Long run neutrality is rejected for the full sample period. However, evidence suggests that the rejection is the result of a severe, downward shift in the mean growth rate of real GDP occurring at 1982. Neutrality is not rejected if post-1981 real GDP is adjusted for the change in mean growth or if one uses data only through 1981. The downward shift in mean real GDP growth followed sharp upward movements in the money and inflation series. This finding of non-neutrality in Mexico arising from extreme conditions is similar to that of Boschen and Otrok (1994) for the US.

## 1. Introduction

Does a permanent change in the money supply alter only nominal variables, leaving real variables unaffected? In other words is money neutral? This issue has interested generations of economists and will probably continue to do so. Many macroeconomic models do allow for short run non-neutrality, although exactly how money might affect output and other real variables in the short run remains a contentious issue. Numerous models provide a plethora of competing possibilities, ranging from the new classical misperceptions model of Lucas (1972) to new Keynesian models of wage and nominal price rigidities. However, long run monetary neutrality is a feature of most macroeconomic models. Indeed, it is likely that the absence of long run neutrality in a modern macro model with optimizing agents would be considered remarkable.

The long run neutrality (LRN) proposition has not proven easy either to verify or to dismiss. Recently, Fisher and Seater (1993, henceforth FS) and King and Watson (1997) have shown that conclusions regarding the proposition depend critically on the time series properties of the data, specifically the orders of integration of money and real output. Since the issue of variable integration is central to testing for neutrality, yet is a relatively recent development in time series analysis, some earlier findings regarding neutrality are suspect.[1]

We add to the growing literature using the FS methodology to examine long run money neutrality. Specifically, we test for long run neutrality of money on real output (GDP) in Mexico using annual data for 1932-2001. Our study makes several contributions to the LRN literature. First, most previous studies of neutrality have focused on industrialized countries which generally impose relatively few financial

[1] See the discussion of some of the earlier tests in FS.

constraints on their citizens and whose financial markets are highly developed. During at least part of the 1932-2001 period Mexico experienced capital controls, controls on interest rates, fixed exchange rates, bank nationalization, and other forms of direct government intervention in financial markets. Does long run neutrality of money characterize a financially constrained, developing economy such as Mexico? The response is a qualified yes.

Second, we expand the range of econometric tests to which data in neutrality studies are typically subjected. The orders of integration of the variables determine the appropriate LRN test in the FS approach, yet it is well known that unit-root tests suffer from problems of low power and sensitivity to lag specification. Recognizing these problems, we base our inferences on results from several different tests and a variety of lag selection methods. It also is widely known that the presence of structural breaks may influence the results of unit-root tests. We include unit-root tests that allow for the possibility of a one-time break in a linear trend under the alternative hypothesis. Finally, although the FS test is based on relatively few assumptions, it does require that money be exogenous. Unlike previous applications of the FS methodology, the assumption of money exogeneity is tested.

Third, as Boschen and Otrok (1994) demonstrated for the U.S. and Haug and Lucas (1997) showed for Canada, inclusion of a period of aberrant macroeconomic performance in the study sample may result in a rejection of long run neutrality. We provide evidence that sudden, unusually large changes in the money process can have real effects. For Mexico the post-1981 period, rather than the Great Depression as in the Boschen and Otrok and Haug and Lucas papers, is responsible for the rejection of long run neutrality. After 1981 a severe reduction in Mexican real growth followed previously unrealized increases in money growth and inflation.

In the following section we briefly describe the FS test and the literature it has spawned. Data integration and exogeneity are addressed in section three. The fourth section presents our finding of long run non-neutrality for the full data sample. The fifth section describes the evidence for structural change in the real GDP growth process and shows that non-neutrality is the result of the post 1981 shift in mean real GDP growth. Section six demonstrates that the shift in average real GDP growth followed significant increases in both the money and inflation series. Our conclusions follow.

## 2. The Fisher-Seater Methodology

FS begin with a two variable, log-linear ARIMA model that is stationary and invertible. The model is given by equations (1) and (2) where $m_t$ and $y_t$ are log money

$$a(L)\Delta^{\langle m\rangle}m_t = b(L)\Delta^{\langle y\rangle}y_t + u_t \tag{1}$$

$$d(L)\Delta^{\langle y\rangle}y_t = c(L)\Delta^{\langle m\rangle}m_t + w_t \tag{2}$$

and log output respectively, $a_0 = d_0 = 1$, and $u_t$ and $w_t$ are mean zero, i.i.d. error vectors. The notation $\langle q\rangle$ refers to the order of integration of variable $q = m, y$.[2] The long run derivative ($LRD_{y,m}$) of output with respect to a permanent change in money is given by equation (3)

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

[2] We substitute the money and output symbols, m and y respectively, for the more general notation used by FS. Otherwise their notation is used throughout the discussion.

provided that $\lim_{k\to\infty} = \frac{\partial m_{t+k}}{\partial u_t} \neq 0$. Equation (3) shows that the long run derivative is the limit of the long run elasticity of output with respect to money. If the limit of the denominator in equation (3) is zero then there are no permanent changes in the monetary variable, thus $\langle m \rangle = 0$, and there can be no test of neutrality. For $\langle m \rangle \geq 1$, FS show that equation (3) can be written as

$$LRD_{y,m} = \frac{(1-L)^{\langle m \rangle - \langle y \rangle}\gamma(L)\big|_{L=1}}{\alpha(L)} \tag{3′}$$

where $\alpha(L)$ and $\gamma(L)$ are functions of the coefficients from equations (1) and (2).[3]

Equation (3′) demonstrates that the value of $LRD_{y,m}$ depends on $\langle m \rangle - \langle y \rangle$, the difference in orders of integration of (log) money and (log) real output. The unit-root tests, reported below, suggest that money and output in Mexico are I(1) variables, thus only the case of $\langle m \rangle - \langle y \rangle = 0$ is considered. In this instance equation (3′) reduces to

$$LRD_{y,m} = \frac{\gamma(1)}{\alpha(1)} = \frac{c(1)}{d(1)}. \tag{4}$$

Imposing the restriction that money is exogenous in the long run, FS demonstrate that $c(1)/d(1)$ can be consistently estimated as $b_k$ from the OLS regression in equation (5).

$$y_t - y_{t-k-1} = a_k + b_k(m_t - m_{t-k-1}) + e_{kt} \tag{5}$$

Equation 5 is the basic equation used in our study of neutrality as well as others which employ the FS methodology.[4]

---

[3] $\alpha(L)=d(L)/[a(L)c(L)-b(L)c(L)]$ and $\gamma(L)=c(L)/[a(L)c(L)-b(L)c(L)]$.

[4] FS also show that long run superneutrality can be tested when $\langle m \rangle = 2$ and $\langle y \rangle = 1$. The estimated model is then $y_t - y_{t-k-1} = a_k + b_k(\Delta m_t - \Delta m_{t-k-1}) + e_{kt}$. In this case long run neutrality immediately holds with these orders of integration when y is a real variable. See case (iii) in the FS discussion of superneutrality.

FS test for long run neutrality of money (M2) on real income using US data for 1869-1975. They conclude that both variables are I(1) so that estimation of equation (5) is the appropriate test. Their results do not support LRN for the US during this period. The $b_k$ coefficients are positive and significant for most values of k.

Boschen and Otrok (1994) reexamine the FS results for the US. They find the same coefficient estimates as FS and likewise conclude that money was not neutral over the 1869-1975 period. However, Boschen and Otrok also examine neutrality over two sub-periods 1869-1929 and 1940-1992 and conclude that LRN held for both periods. Furthermore, neutrality holds for the 1869-1992 period (values of $k > 3$) when equation (5) is modified by the inclusion of an intercept dummy for the 1930-1939 Great Depression. They surmise that the large number of bank failures, which accompanied the money supply shocks during the depression period, may have had real effects that do not normally occur with purely monetary events.

Haug and Lucas (1997) investigate the Boschen-Otrok conjecture about the effects of bank failures by testing for neutrality of money in Canada using the FS methodology. Although Canada suffered sharp declines in money and real output during the Great Depression, it did not experience the disruption of widespread bank failures seen in the United States. For Canada from 1914-1994, Haug and Lucas find that the $b_k$ from estimation of equation (5) are significant for values of k from 10 to 13, thus LRN cannot be accepted of Canada. However, Haug and Lucas also test a version of equation (5) that includes a dummy variable for the depression period and find support for neutrality with point estimates of $b_k$ that are insignificant for all k. They interpret this finding as indicating that the Great Depression was an "anomalous period in terms of the money-output relationship" (p. 759) even in the absence of widespread bank failures.

Olekalns (1996) uses the FS approach to investigate monetary neutrality in Australia for 1900-1994. He concludes that neutrality holds unambiguously with respect to real GDP when M1 is the money measure but that values of $b_k$ are significantly different from zero for broader money, M3. Olekalns suggests that the neutrality results may be sensitive to how money is measured.

Coe and Nason (2002) apply the FS methodology to data for Canada, Australia, the United States, and the United Kingdom and find weak support for LRN. Like Olekalns, they conclude that results are sensitive to the particular measure of money used in estimating equation (5). Using Andrews (1989) inverse power function, Coe and Nason find that the power of the FS test is often low so that it may be difficult to assess the validity of the neutrality proposition.

The previously cited studies focus on developed countries.[5] Just two published studies apply the Fisher-Seater methodology to data from developing nations. Bae and Ratti (2000) find that money is long-run neutral, but not long-run superneutral for Brazil and Argentina. Wallace (1999) examines the LRN proposition with annual data for Mexico, 1932-1992. Both M1 and M2 are found to be I(1) as is real output. With few exceptions Wallace finds that the $b_k$ coefficients are not significantly different from zero when estimating equation (5) with either money measure.[6]

---

[5] Using a different methodology King and Watson (1997) and Serlitis and Koustas (1998, 2001) find support for the LRN proposition in the developed countries they study.
[6] As explained in more detail below, the data set used by Wallace differs from the one we use.

## 3. Integration and Exogeneity

Data used in this study are from either the National Institute for Statistics, Geography, and Information of Mexico (known by its Spanish acronym of INEGI) or the Central Bank (Banco de México). A complete description is provided in Appendix A.

The orders of integration of the money and output series determine the appropriate form of the FS test. Unfortunately, it is well known that unit-root tests have low power and that results can vary with the type of test used and on the number of lags included in the test equation. With this in mind, the results of several procedures are examined in order to draw conclusions regarding variable integration.

We first apply the familiar Phillips-Perron (PP) and Augmented Dickey-Fuller (ADF) tests. Three lags are included in the PP test equations, as indicated by the procedure of Newey and West (1994b).[7] For the ADF tests, four different methods are used to select the appropriate lag length for the test equation: Lagrange multiplier (LM), Akaike Information Criterion (AIC), Bayesian Information Criterion (BIC), and a general to simple (GS) approach. For the LM specification, beginning with zero, lags are added to the ADF equation until an LM test indicates that serial correlation is eliminated at the 10% significance level. In the GS method, we begin with 6 lags and then iteratively drop each final lag if it is insignificant at the 10% level.

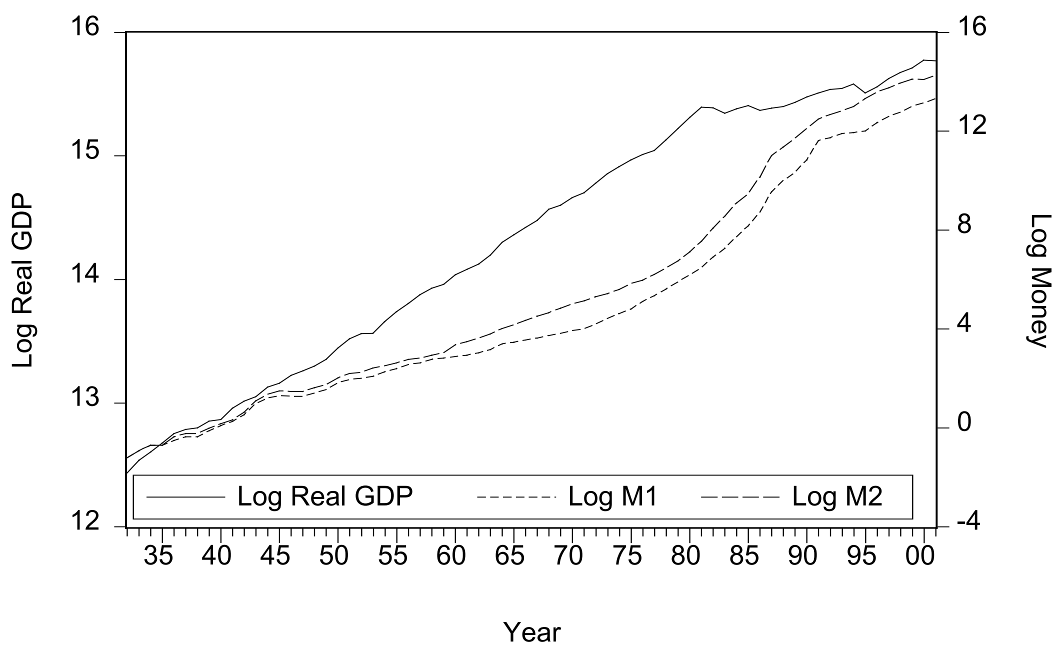
In addition, we employ a test proposed by Elliott, Rothenberg, and Stock (1996) that is more powerful than the ADF and PP tests when the alternative is trend-stationarity. This test is similar to the usual augmented Dickey-Fuller test, except that the logged series are detrended or "quasi-differenced" in a way that is efficient under the

[7] Experimentation showed that changes in lag length did not affect the conclusions of the PP tests.

alternative hypothesis. Because of its equivalence to generalized least squares, this method is referred to as DFGLS.

We also use the two tests proposed by Kwiatkowski, et. al. (1992) which are denoted KPSS( $\mu$ ) and KPSS( $\tau$ ). The KPSS( $\mu$ ) test provides a test of stationarity versus an alternative of trend-stationarity. The KPSS( $\tau$ ) test offers an alternative to the usual unit-root testing strategy because the null hypothesis is trend stationarity and the alternative is a random walk with drift. The KPSS tests tend to be sensitive to lag length employed in the Newey-West serial correlation correction. However, our test statistics stabilized with the use of eight lags.[8]

**Figure 1**
**Logarithms of Real Output, M1, and M2**
**Mexico: 1932-2001**




[8] In their original paper KPSS also found test statistics for annual data tended to stabilize with 8-lags. Unit-root tests for all series in levels are conducted with both a constant and a trend included in the test equations.

The real GDP and money series are logged. As shown in Figure 1, these variables display the upward movement typical of money and output in most countries. The KPSS ($\mu$) test rejects stationarity of all variables. Tests for trend stationarity versus an alternative of a random walk with drift then are appropriate for these series.

The PP test, the DFGLS test, and all ADF tests fail to reject a unit root in each series using a 10% critical value. Furthermore the KPSS ($\tau$) test rejects trend-stationarity in favor of a unit root for each series at the 2.5% significance level. Given the variety of tests and lag truncation schemes used, the strength and consistency of the conclusions regarding log real GDP and money are striking. Unit root test results are provided in Appendix B.

Perron (1989) and Rappoport and Reichlin (1989) have suggested that macroeconomic variables may be trend-stationary with a break in trend rather than integrated series. Perron demonstrated that standard unit-root tests too often fail to reject the null hypothesis of a unit-root if the true data generating process is trend-stationary with a structural break.. A substantial body of literature has developed a variety of unit-root tests with an alternative of trend-break stationarity. Papers include Christiano(1992), Zivot and Andrews (1992), Banarjee, Lumsdaine, and Stock (1992), Perron and Vogelsang (1992), Perron (1997), Vogelsang and Perron (1998), and Murray and Nelson (2000). The tests may allow for a break in the intercept or the slope of the trend function (or both). However, Sen (2003) argues that selection of the model form requires use of pre-test information and selection of an incorrect model may lead to power distortions. Because of these problems, Sen recommends that one allow for breaks in both the intercept and slope (mixed model).

Examination of the plot of the real GDP series reveals a potential breakpoint at 1981-1982. This feature makes Mexican real output a candidate for trend-break testing. Also, evidence of structural change in the Mexican output series, presented later, requires us to address the possibility that our series are trend-break stationary. For each data series we estimate both additive outlier (AO) and innovative outlier (IO) models with endogenous breakpoints. We follow Sen's recommendation by using the most general form represented by the "mixed" model. Details of the procedures and the results are provided in Appendix C. Using a 10% critical value, tests conducted under both the AO and IO specifications fail to reject the presence of a unit-root in any of the three (log) money series or in (log) real GDP, thus verifying our earlier conclusion that each series contains at least one unit root.

We continue by testing each log-differenced series for the presence of a trend or additional unit roots. It is not immediately obvious from variable plots if there is consistent upward or downward movement in the growth rates of real GDP and the money measures; therefore, we use three alternative strategies to determine if trend-stationarity is a valid alternative for the unit-root tests. First, each growth rate is regressed on a constant and a time trend with the standard error on the estimated trend coefficient adjusted for serial correlation using the Newey-West correction. Second, each growth rate is regressed on a constant, a time trend, and a set of lagged growth rates to remove serial correlation. Third, we examine the results of the KPSS($\mu$) test for stationarity. For real GDP growth both regression methods reject the null of a zero coefficient on the trend. Also, the KPSS($\mu$) test statistic is very close to the 5% critical value. Therefore, both a constant and a trend are included in the unit-root test equations for real GDP growth. For the three money series at least one regression method fails to reject a zero trend coefficient at a 5% significance level, and in each case the KPSS($\mu$) test fails to reject stationarity. Consequently no trend is included in the test equations for money growth rates .

The PP test and all variations of the ADF test reject unit-roots in real GDP growth, M1 growth, and M3 growth at a 1% significance level thus indicating that M1 and M3 growth are stationary and that real GDP growth is either stationary or trend-stationary.[9] Evidence regarding M2 growth is mixed. However, because the PP test and both the LM and BIC versions of the ADF test reject a second unit root in M2 growth, we conclude that M2 growth is stationary. Further, as mentioned above, the KPSS($\mu$) test fails to reject stationarity of M2 growth. Finally, this conclusion is consistent with results for the other two money growth measures.

The Fisher-Seater methodology requires that money be exogenous. The validity of this assumption is assessed by testing if real GDP growth Granger-causes money growth. The AIC and BIC criteria are used to determine the appropriate lag length for the test equations, with a maximum of 8 lags considered. Thus the samples for all Granger-causality tests begin in 1941. Both criteria select one lag for the three money measures.[10] Granger-causality is rejected at a 10% marginal significance level for both M2 and M3 growth. However, Granger-causality cannot be rejected at the 5% marginal significance level for M1 growth. Given the finding regarding M1, neutrality results for M2 and M3 are emphasized later in the paper.

[9] As reported below real GDP growth in Mexico contains a level shift rather than a trend.
[10] LM tests for serial correlation of orders 1-6 do not indicate any serial correlation in estimates of the test equations.

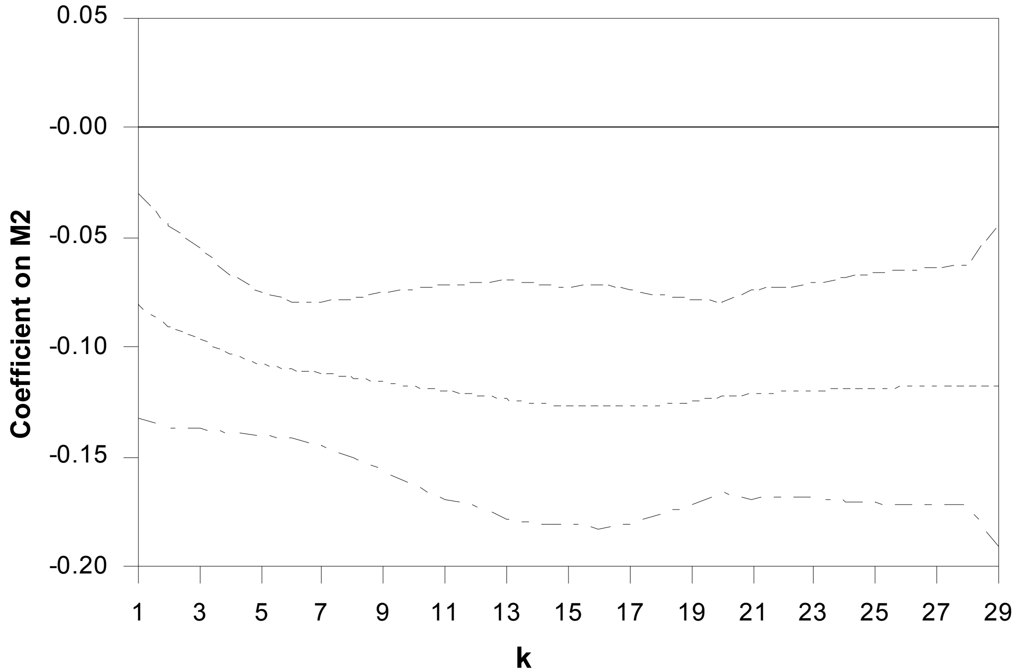
## 4. Long run neutrality tests

Since the Granger-causality tests indicate M2 and M3 are exogenous with respect to real GDP and each of the money series and real GDP are integrated of order one, we proceed with neutrality tests under the FS methodology. Equation (5) is separately estimated for values of k = 1…29. Each coefficient, $b_k$, represents the estimated response of the change in log real GDP to the change in logged money over k+1 periods. The values of $b_k$ for logged M2 and the 95% confidence interval are plotted in Figure 1.[11] As the figure shows, all the estimated $b_k$ are negative, ranging from -.08 to -.13, and the upper limit of the confidence interval lies below the zero line thus indicating that all the coefficients are significant. For Mexico, M2 was not neutral over this period. Furthermore, increases in M2 were associated with declines in real GDP. Although the Granger-causality test results suggest that M1 is not exogenous, thus violating an assumption of the FS methodology, the same non-neutrality result and finding of a negative effect on output are obtained. Results with M3 are virtually identical to those for M2.[12]

This result is counter to that of Wallace who examines the 1932-1992 period in Mexico. His finding that LRN holds is attributable to the use of out-of-sample data from 1925-1931 in the creation of the differenced output and money variables used in the estimation of equation (5). Prior to 1932 Mexico was on a gold standard during which money was likely to be endogenous. Reestimation of equation (5) using Wallace's real GDP series (which differs slightly from the one we use) and only data from 1932 on yields results very similar to ours with LRN clearly rejected.

[11] The degrees of freedom are T/k where T is the number of observations. In this instance T=70.
[12] Results using M1 and M3 are available from the authors.

**Figure 2**
**FS Regression Results**
**Output and M2 (1932-2001)**




## 5. Non-Neutrality and Structural Change in Real Output.

Why the clear rejection of neutrality? After real output sharply declined in 1932, real GDP grew consistently for fifty years until the next decline occurred in 1982 (see Figure 1). The pace of real growth then slowed over the last twenty years of the sample. The severity and persistence of this slowdown raises the possibility that significant structural changes occurred between the years 1981 and 1982.[13]

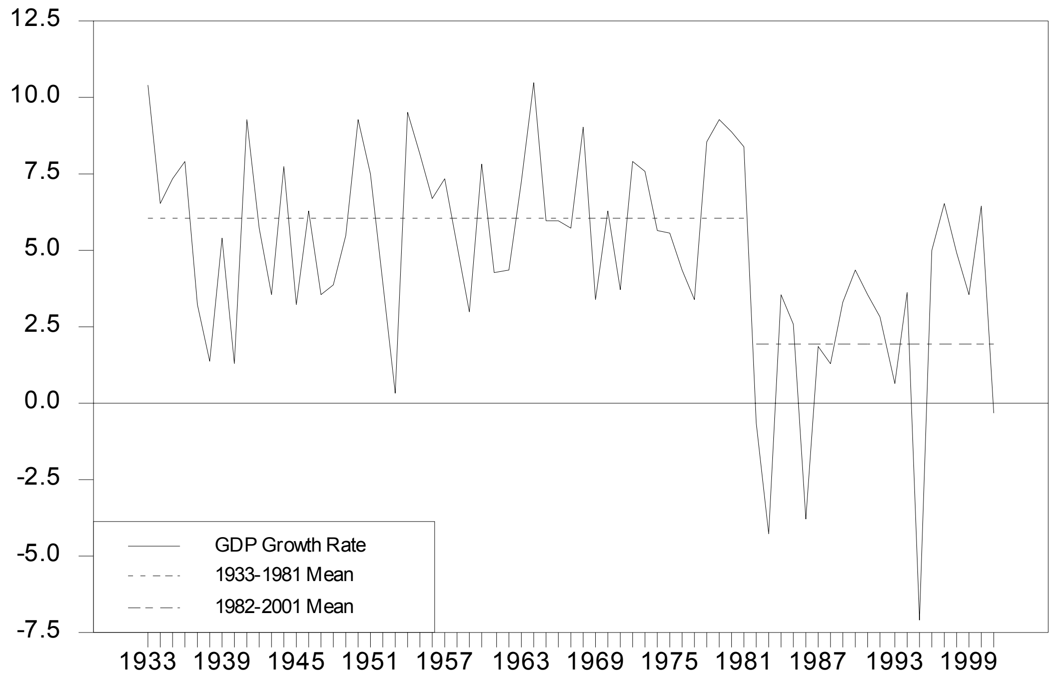
Bai and Perron (1998, 2003) provide a method for identifying and estimating multiple structural changes in regression coefficients. Break-dates are chosen to globally minimize the sum of squared residuals of the regression model. We use the Bai-Perron methodology to test the possibility that the mean of Mexican real GDP has undergone

[13] Recall that earlier test results indicate that any such structural change was not of the trend-break variety.

one or more changes.[14] Results indicate that a shift in mean real GDP growth occurred after 1981. A SupF test indicates that this is the only shift significant at a 5% marginal significance level.[15]

As can be seen in Figure 3, the size of the post-1981 (downward) shift is striking. Real GDP grew at an overall average annual rate of 4.84% from 1933 through 2001. However, from 1933 to 1981 the annual growth rate was 6.04% compared to a much slower annual growth rate of 1.88% from 1982 to 2001. The two horizontal line segments indicate the average growth rates for 1933-1981 and 1982-2001.

**Figure 3**
**Mexican Real GDP Growth (1933-2001)**




[14] We gratefully acknowledge use of the computer program available at Perron's website. We use a trimming value of 15% for all estimates. For our data sample, this requires any potential regime to be of at least 10 years duration.
[15] The SupF is the maximum F-statistic among those for all potential breakpoints. The test for a second potential break takes the initial breakpoint of 1981 as a given.

Further evidence of structural change between 1981 and 1982 is provided by a Jarque-Bera test for normality of the real GDP growth series. For the 1933-2001 sample, normality is rejected using the overall Jarque-Bera statistic and the individual statistics for skewness and kurtosis. However, this result reverses when one tests the 1933-1981 and 1982-2001 sub-periods separately. All three statistics fail to reject normality for the 1933-1981 sub-sample. Even for the relatively short 1982-2001 sub-sample one cannot reject normality at a 5% marginal significance level using any of the statistics. Thus, it appears that real GDP growth follows different normal distributions for the two sub-samples.

**Table 1**
**Normality Tests of Real GDP Growth**

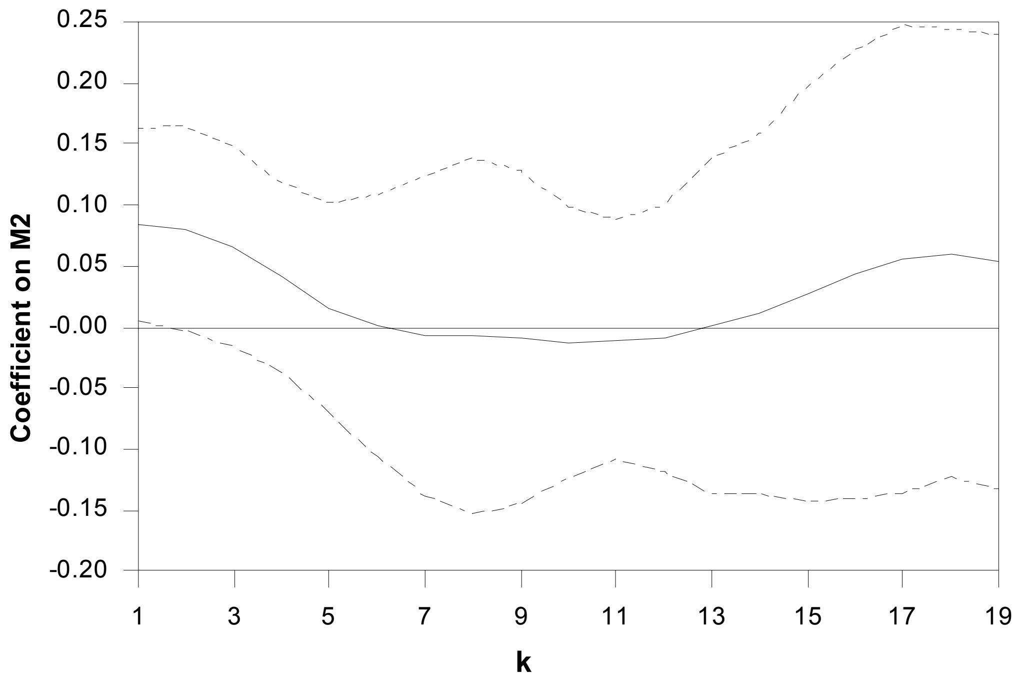
| | **1933-2001** | **1933-1981** | **1982-2001** |
|---|---|---|---|
| **Jarque-Bera** | 21.4289<br>(.00002) | 1.2457<br>(.5364) | 4.7567<br>(.0927) |
| **Skewness** | -1.0101<br>(.00081) | -.2317<br>(.5210) | -1.1107<br>(.0608) |
| **Kurtosis** | 1.8365<br>(.0031) | -.6288<br>(.4035) | .8793<br>(.5043) |

What is the effect of the post-1981 shift in mean real GDP growth on the FS test results? We begin our investigation of this question by conducting the FS test using only data through 1981.[16] Results using M2 as the monetary variable are shown in Figure 4.[17] Zero is contained within the 95% confidence interval for all values of k>1 thus indicating

[16] Unit-root tests using data only through 1981 are generally consistent with our earlier conclusion that output and money are I(1). Results of these tests are available from the authors on request.
[17] Again the results using M1 and M3 are very similar to those reported for M2 and are available from the authors. Due to the reduced number of observations only variable differences corresponding to k = 1…19 are used in the estimations.

that there were no long run effects of permanent monetary shocks on real output for the1932-1981 sub-sample. Money was LRN. Further, the result indicates that the post-1981 behavior of output and money caused the earlier non-neutrality finding.

**Figure 4**
**FS Regression Results**
**Output and M2 (1932-1981)**




Next using a simple experiment, we explore the possibility that the non-neutrality finding in the full sample can be entirely explained by the 1981-1982 level shift in real GDP growth.[18] Suppose that the level shift did not occur. Then (log) real GDP for each year from 1982 through 2001 could be expressed as equation (6) where the (hypothetical)

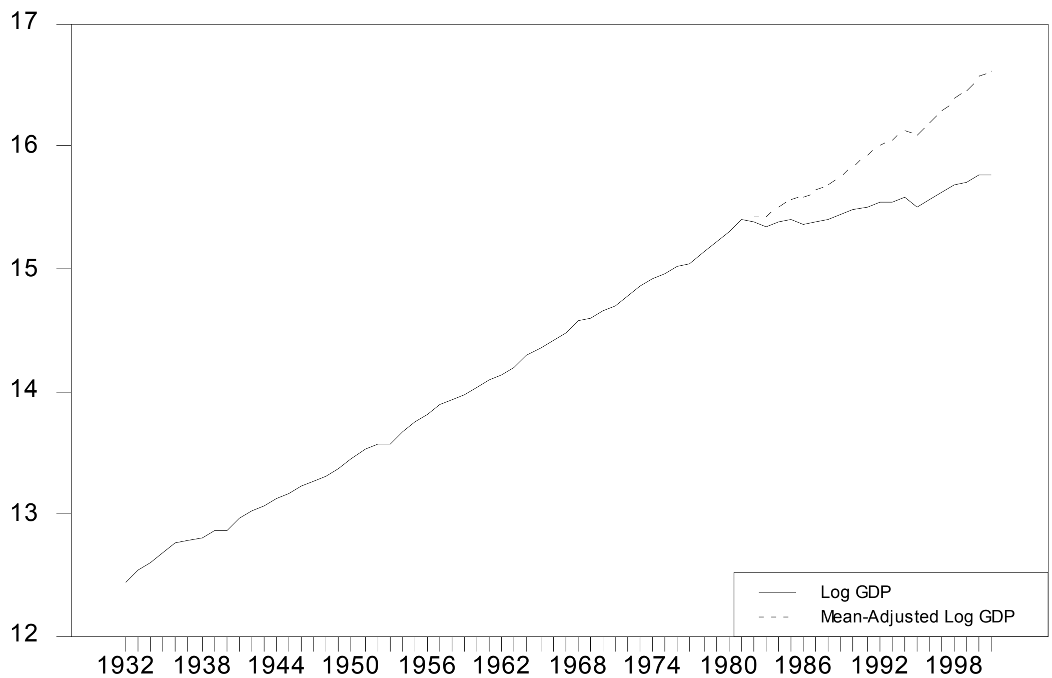
[18] We investigated two alternative explanations for the non-neutrality result. Specifically the severe 1995 recession known in Mexico as the 'Crisis' and the 1982-1990 period when banks were nationalized . For the restricted 1932-1994 sample, all $b_k$ coefficients are significant and negative. Thus rejection of LRN occurs due to events prior to the Crisis in 1995. Nor can the non-neutrality result be attributed to bank nationalization. The coefficients on differenced money are still negative and significant with the inclusion of a dummy variable for 1982-1990 in the FS test.

value of real GDP, denoted $\tilde{y}_t$, is equal to actual GDP adjusted for the change in mean growth.[19]

$$\tilde{y}_t = y_t + (t - 1981)(6.042973 - 1.881805) \qquad t > 1981 \tag{6}$$

In this equation the product of the number of years since the 1981 shift and the difference in the growth rates for the two sub-periods is added to the actual value of real GDP in order to obtain the hypothetical value. Actual and mean-adjusted values of (log) real GDP are shown in Figure 5.

**Figure 5**
**Actual and Mean Adjusted (Log) Mexican Real GDP (1932-2001)**

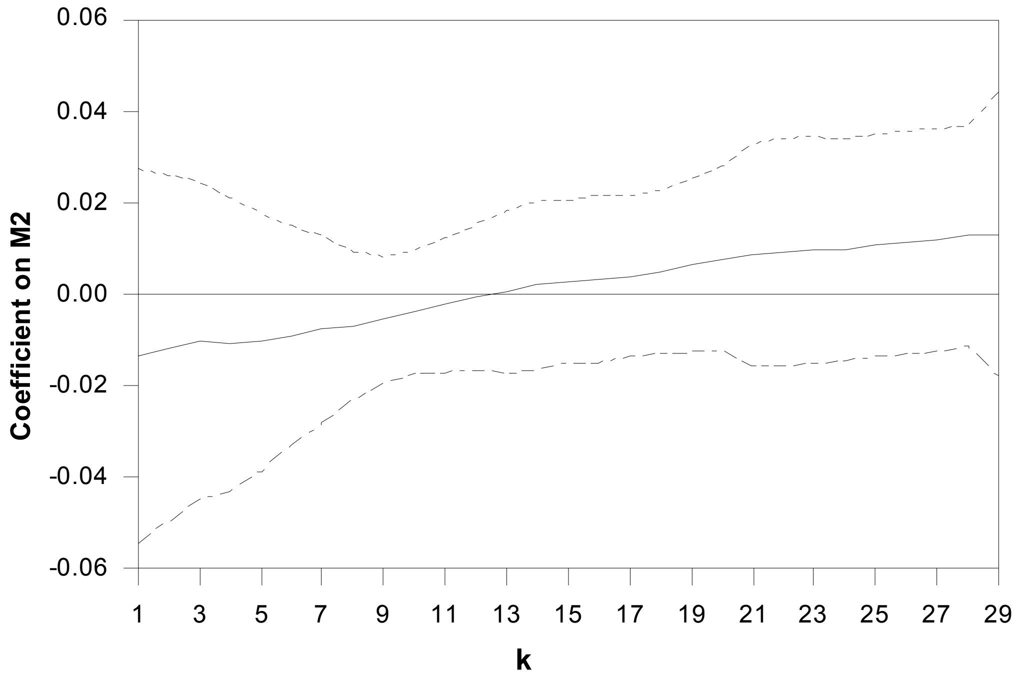



Results of the FS test using M2 and hypothetical, mean-adjusted real GDP are shown in Figure 6. Zero lies within the confidence intervals for all values of k; therefore,

[19] Unit-root tests indicate that the growth rate of mean adjusted log-GDP is stationary and does not contain a trend.

one would not reject long run neutrality over the full sample if the mean of Mexican real GDP growth had not changed after 1981. This supports the conjecture that the 1981-1982 event explains the non-neutrality result.

**Figure 6**
**FS Regression Results**
**Mean-Adjusted Output and M2 (1932-2001)**




We conducted an additional experiment by estimating equation (5) for 1932-1982 then sequentially including an additional year of data until we have estimated the neutrality regressions for samples from 1932-1982 to 1932-1987. This allows us to investigate the amount of time required for the impact of the 1982 shift in mean real GDP growth to reverse pre-1982 findings in favor of neutrality. To conserve space only the plots for 1932-84, 1932-1986 and 1932-1987 are shown in Figures 7a-c.[20]

[20] The omitted figures can be obtained from the authors.

**Figure 7a**
**FS Regression Results**
**Output and M2 (1932-1984)**

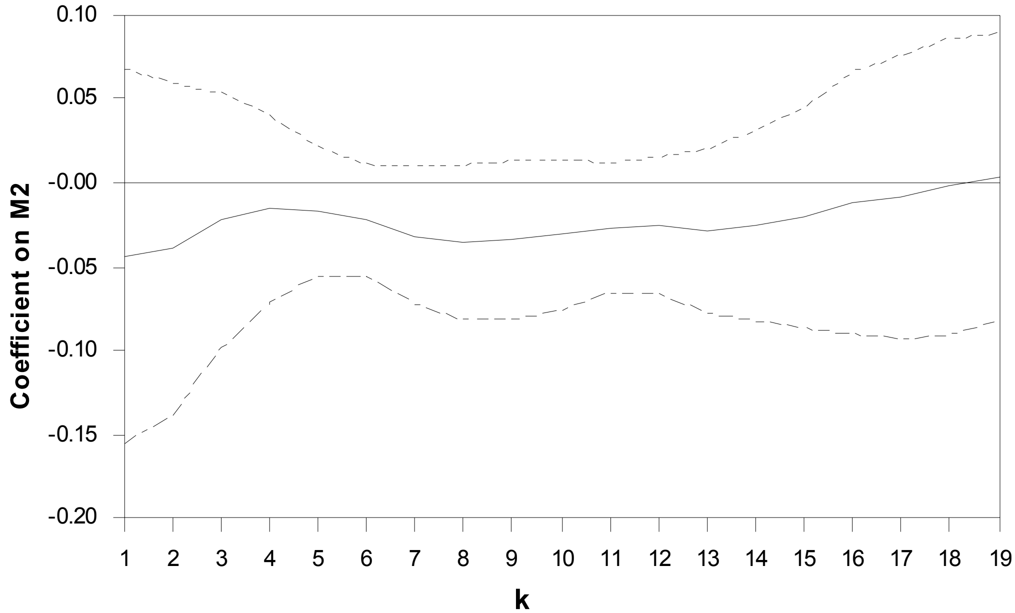



**Figure 7b**
**FS Regression Results**
**Output and M2 (1932-1986)**

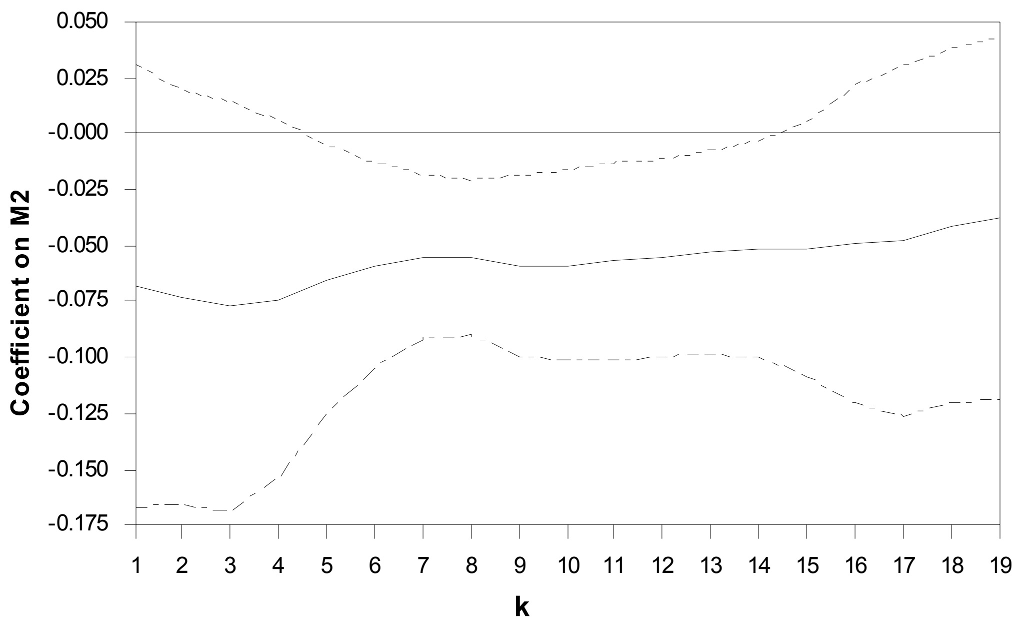
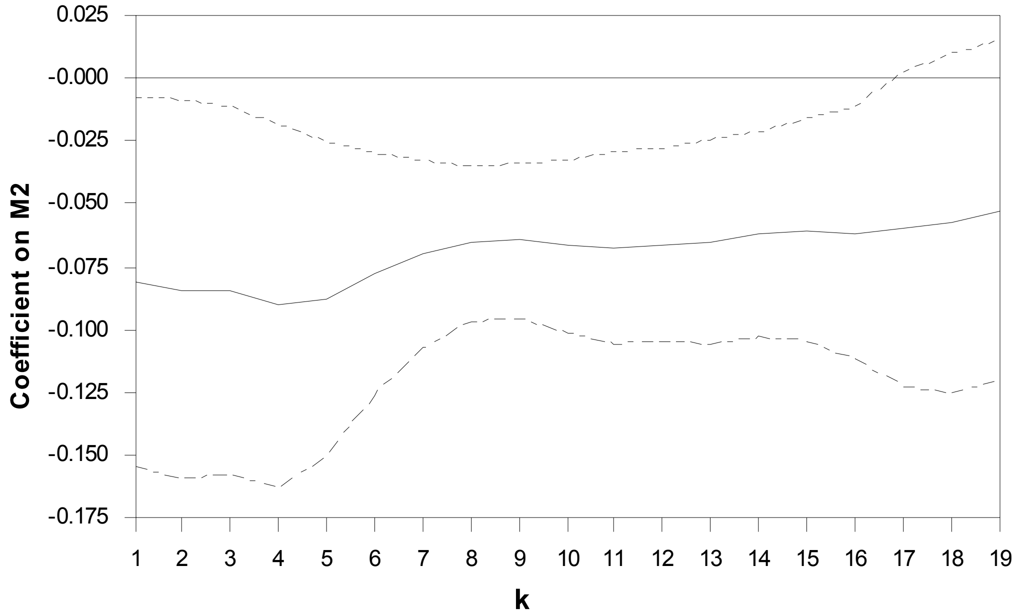


**Figure 7c**
**FS Regression Results**
**Output and M2 (1932-1987)**

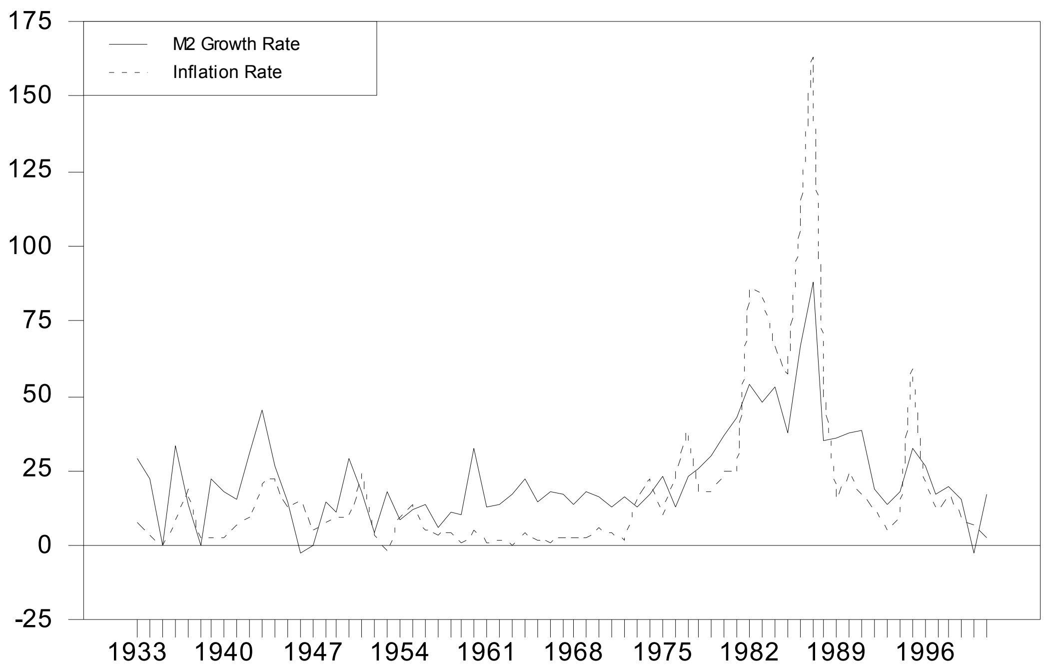



In the sequence of graphs, the null hypothesis of neutrality cannot be rejected for samples through 1984, but once data for 1985 are included LRN is violated. This conclusion is strengthened with the inclusion of data from 1986 and 1987. Sequential examination of the graphs reveals a downward shift of the coefficient plot with each additional year of data. For the 1932-82 period, the coefficients are either positive or (virtually) zero and insignificant. With the addition of data from 1983, the second year of relatively high inflation (greater than 84%) many of the coefficients are negative although none are significant. By the time data from 1984 and 1985 are included, all coefficients for k = 1…19 are negative and two of the $b_k$ are significant. By 1987 all values of $b_k$ for k < 17 are significant and negative. This is consistent with our conjecture, discussed in the next section, that the failure of LRN to hold for the full sample appears attributable to the high money growth, high inflation regime of 1982-1987. Further, although the FS methodology has been criticized by Coe and Nason for lack of power, it does quickly identify the non-neutrality occurring in Mexico.

**6. Causes of the real growth slowdown and long run non-neutrality.**

Long run neutrality is rejected for the 1932-2001 sample due to a significant slowdown in real GDP growth after 1981. What accounts for this slowdown? Mexico has been politically stable since the 1920's and has not been involved in any major military conflicts. Instead, the source of economic instability appears to lie in the time paths of money growth and inflation.

**Figure 8**
**Mexican M2 Growth and Inflation**



A plot of M2 growth and the inflation rate is provided in Figure 8, while Panel A of Table 2 shows average annual growth rates of money and the price level for the full sample as well as for both the 1933-1981 and 1982-2001 sub-periods. As can be seen in the graph and the table, the behavior of both money growth and the price level are very different for the two sub-periods. Average annual rates of money growth are below 18% for all three measures from 1933-1981. However, average annual money growth is above 33% for all three measures from 1982-2001. The price level increased at an average annual rate of less than 10% from 1933-1981, yet increased by an average of more than 41% annually from 1982-2001. The annual inflation rate in Mexico went from nearly 25% in 1981 to almost 86% in 1982. During the next six years annual changes in the GDP deflator ranged from 50% to 163% annually. In 1989 inflation decreased to an annual rate of 16% and, except for a brief jump to 59% during the 'Crisis,' has since stayed below 25% annually.[21]

**Table 2**

**Panel A-Annual Average Percentage Growth Rates of Output and Money**

| Variable | 1933-2001 | 1933-1981 | 1982-2001 |
|---|---|---|---|
| **GDP** | 4.84 | 6.04 | 1.88 |
| **M1** | 21.07 | 15.69 | 34.24 |
| **M2** | 22.42 | 17.87 | 33.56 |
| **M3** | 22.87 | 17.94 | 34.94 |
| **Price Level** | 18.56 | 9.32 | 41.22 |

**Panel B-Standard Deviation of Growth Rates of Output, Money, and Prices**

| Variable | 1933-2001 | 1933-1981 | 1982-2001 |
|---|---|---|---|
| **GDP** | 3.40 | 2.47 | 3.61 |
| **M1** | 16.63 | 9.95 | 21.97 |
| **M2** | 15.82 | 10.34 | 21.01 |
| **M3** | 16.56 | 10.52 | 22.05 |
| **Price Level** | 27.83 | 8.78 | 42.61 |

[21] The annual changes are December to December. Inflation data prior to 1982 are based on a price index which includes both consumer and producer prices. A complete GDP deflator series is not available for the full sample period.

Not only did average annual rates of growth of money, the price level, and output change after 1981 but the volatility of the variables increased as well. Panel B of Table 2 displays the standard deviation of the annual changes for the five variables over the two periods. For 1982-2001 the standard deviations of the three money measures are more than double those from the earlier period. The standard deviation of annual price level changes increased more than fourfold after 1981. Despite substantially lower average growth rates post-1981, output volatility was also greater.

The picture that emerges from these data is one of high inflation driven by excessive money growth. This is consistent with the conclusion of Feliz and Welch (1997) that actual and anticipated money growth drive inflation in Mexico and four other Latin American countries.[22] Accompanying the inflation in Mexico was an increase in economic uncertainty as measured by the standard deviations of the money, output, and price level variables. These results are consistent with the finding of Bruno and Easterly (1998) that inflation rates above 40% annually have a negative effect on economic growth. The *average* annual inflation rate in Mexico was above the 40% for 1982-2002, and exceeded 100% in 1986 and 1987. The findings are also supportive of theoretical models suggesting that the welfare losses from inflation may be sizeable. See Wu and Zhang (1998) and Ball and Romer (2003) for examples.

A full explanation for the reason the Mexican government selected a high money growth/inflation regime is beyond the scope of this paper. Lustig (2001) provides an account of events during this period. However a brief sketch is appropriate to assist in the understanding of our empirical findings. During the late 1970s Mexico discovered large reserves of oil. These discoveries along with rising petroleum prices caused sharp

[22] For Mexico they use monthly data from January 1972 to September 1989.

increases in government revenues since oil production is a government-owned monopoly in Mexico. Furthermore, most forecasters predicted that oil prices would continue to rise. Based on the expectation of increasing revenue streams, the Mexican government expanded public expenditures and borrowed heavily in domestic and international credit markets. When petroleum prices began to decline in the early 1980s, expenditure cuts were deferred in the hope that the price decrease was temporary. As the oil prices continued to decline, Mexico incurred larger deficits and confronted borrowing constraints in international markets. Mexico chose to finance its deficits with money creation and inflation accompanied by welfare losses followed.

**7. Conclusions**

A wide range of diagnostic tests for unit-roots, Granger causality, and structural breaks leads us to conclude that the test for long run neutrality of money developed by Fisher and Seater can be applied to data for Mexico for the 1932-2001 period. For this time period, we find that money is not neutral regardless of the money measure used. However, money is long run neutral if we consider only the 1932-1981 period; permanent shocks to money had no long run effects on real output through 1981. Further, we find that money is long run neutral if we adjust for a shift in mean GDP growth occurring in 1982. This indicates that non-neutrality is a result of the 1982 shift in real growth. The evidence is consistent with our conjecture that this event was caused by excessive money growth causing inflation to rise to levels unprecedented in Mexico at least since 1925.[23]

[23] Economic data are generally unavailable for Mexico from 1910 to 1924 or so, the period of the Mexican revolution and subsequent turmoil as various factions struggled for control.

Fisher and Seater originally concluded that M2 was not neutral in the United States during 1968-1975. Subsequently, Boschen and Otrok showed that the inclusion of the Great Depression period was responsible for the FS non-neutrality result. Haug and Lucas also find that the Great Depression affects the test of LRN for Canada. Our results for Mexico are in much the same spirit and provide further evidence that money neutrality tends to be a long run characteristic of economies except during periods of aberrant economic performance.

## Appendix A

Some adjustments must be made to develop continuous money and output series for the 1932-2001 period. Real output as measured by real GDP in 1980 prices for 1932-1995 is from INEGI (2003).[24] INEGI also reports real GDP for 1995-2001 in 1993 prices. Annual growth rates for 1995-2001 are calculated from the GDP series in 1993 prices and applied to the 1995 real GDP figure in 1980 prices until a complete series is constructed in 1980 prices. For example real GDP in 1993 prices grew 5.15% in 1996 and 6.77% in 1997. These percentages are applied to 1995 real GDP of 5.45 billion pesos to obtain estimates of 1996 real GDP (5.73 billion pesos) and 1997 real GDP (6.12 billion pesos).

Nearly complete series for M1, M2, and M3 compiled using an older methodology are also from INEGI (1994, 1999, 2003) for 1932-2000. Money measures constructed using the current methodology are available from the Bank of Mexico (2003) only from 1985 to the present. In a manner analogous to that employed for GDP, the growth rate in the 'new method' money variables for 2000-2001 is applied to the 2000 figures from the old methodology to obtain figures for 2001 that are consistent with the older approach.

[24] Our real output data are different from those used in the paper by Wallace. In that study real GDP was constructed from several different INEGI sources.

## Appendix B
## Unit-root Test Results

### Table B1
### Log-Series, 1932-2001

| | GDP | M1 | M2 | M3 | 10% Critical Value | 1% CV Critical Value |
|---|---|---|---|---|---|---|
| **ADF** (LM) | .2300 (0-lags) | -1.0371 (1-lag) | -1.4030 (1-lag) | -1.4955 (1-lag) | -3.15 | -4.04 |
| **ADF** (AIC) | .2300 (0-lags) | -1.0371 (1-lag) | -1.9806 (3-lags) | -1.4955 (1-lag) | -3.15 | -4.04 |
| **ADF** (BIC) | .2300 (0-lags) | -1.0371 (1-lag) | -1.4030 (1-lag) | -1.4955 (1-lag) | -3.15 | -4.04 |
| **ADF** (GS) | .2300 (0-lags) | -1.0371 (1-lag) | -1.9806 (3-lags) | -1.8765 (3-lags) | -3.15 | -4.04 |
| **PP** (3-lags) | -.0420 | -.5778 | -.9530 | -.9349 | -3.15 | -4.04 |
| **DFGLS** (4-lags) | -.605 | -1.687 | -1.977 | -1.642 | -2.57 | -3.48 |
| **KPSS ($\mu$)** (8-lags) | .8780 | .8376 | .8507 | .8485 | .347 | .739 |
| **KPSS ($\tau$)** (8-lags) | .1831 | .1999 | .2007 | .2008 | .119 | .216 |

### Table B2
### OLS Tests for Trend in Log-Differenced Series
### 1933-2001

| | Newey-West Correction | OLS with Lagged Dependent Variables (GS Lag Selection) |
|---|---|---|
| **GDP** | t-stat = -2.7801<br>msl = .0054 | Lags = 1<br>t-stat = -2.1549<br>msl = .0349 |
| **M1** | t-stat = 2.0699<br>msl = .0385 | Lags = 1<br>t-stat = 1.8302<br>msl = .0718 |
| **M2** | t-stat = 1.6799<br>msl = . 0930 | Lags = 3<br>t-stat = .3635<br>msl = .7175 |
| **M3** | t-stat = 1.7717<br>msl = .0765 | Lags = 1<br>t-stat = 1.3817<br>msl = .1718 |

**Table B3**
**Log-Differenced Series**
**1933-2001**

| | **GDP** | **M1** | **M2** | **M3** | **10% Critical Value*** | **5% Critical Value*** | **1% Critical Value*** |
|---|---|---|---|---|---|---|---|
| **ADF** (LM) | -6.7411 (0-lags) | -4.0367 (0-lags) | -2.1071 (2-lags) | -3.5719 (0-lags) | -2.58 (-3.15) | -2.89 (-3.45) | -3.51 (-4.04) |
| **ADF** (AIC) | -6.7411 (0-lags) | -4.0367 (0-lags) | -3.8842 (0-lags) | -3.5719 (0-lags) | -2.58 (-3.15) | -2.89 (-3.45) | -3.51 (-4.04) |
| **ADF** (BIC) | -6.7411 (0-lags) | -4.0367 (0-lags) | -3.8842 (0-lags) | -3.5719 (0-lags) | -2.58 (-3.15) | -2.89 (-3.45) | -3.51 (-4.04) |
| **ADF** (GS) | -6.7411 (0-lags) | -4.0367 (0-lags) | -2.1071 (2-lags) | -3.5719 (0-lags) | -2.58 (-3.15) | -2.89 (-3.45) | -3.51 (-4.04) |
| **PP** (3-lags) | -6.9072 | -4.5732 | -4.2426 | -3.9442 | -2.58 (-3.15) | -2.89 (-3.45) | -3.51 (-4.04) |
| **KPSS** **($\mu$)** (8-lags) | .4595 | .3409 | .2808 | .2903 | .347 | .463 | .739 |

* Critical values are given for a specification including a constant but no trend; critical values for a specification including both a constant and a trend are given in parenthesis.

## Appendix C
## Testing for a Unit-Root vs. a Trend with Structural Break

The tests permit two ways of modeling potential break. In either case, at most one break is allowed at some true break-date $T_b$. The first is called an additive outlier (AO) model. This approach assumes that a break occurs suddenly and is unaffected by the dynamics of the series. In order to test for a unit-root versus a trend-break using the AO mixed model the series of interest, $x_t$ , is first detrended using the regression

$$x_t = \mu + \beta t + \theta DU_t + \phi DT_t + \tilde{x}_t , \tag{C1}$$

where $DU_t = 1$ for t >$T_b$ and $DT_t$ = t - $T_b$ for t > $T_b$. A ~ denotes a detrended variable. The parameter $\theta$ measures the size of the shift in the intercept, while $\phi$ measures the size of the shift in the slope of the trend function. Second, the unit-root hypothesis is tested with the t-statistic for the null hypothesis of $\lambda = 1$ from the regression:

$$\tilde{x}_t = \sum_{i=0}^{j} \omega_i D(T_b)_{t-i} + \lambda \tilde{x}_{t-1} + \sum_{i=1}^{j} c_i \Delta \tilde{x}_{t-i} + u_t , \tag{C2}$$

where $D(T_b)_t = 1$ for t = $T_b$+1. Inclusion of the additional dummy terms $D(T_b)_{t-i}$ (i=0,…,j) allows a break in the constant term under the null hypothesis of a unit-root and ensures the proper limiting distributions of the t-statistics. The choice of the number of lags, j, is described below.

The second approach, referred to as an innovative outlier (IO) model, allows the break to occur gradually. Typically an IO model is estimated assuming that the variable of interest reacts to trend function shocks in the same fashion that it responds to shocks in the innovation process. The mixed version of the IO model specifies the potential trend-break as

$$x_t = \mu + \beta t + \psi(L)(\theta DU_t + \phi DT_t + e_t) , \tag{C3}$$

where $\psi(L)$ is a polynomial in the lag operator. Note that in the AO model the long run impact of a shift in slope is given by $\gamma$, while in the IO model, the long run impact is given by $\psi(1)\phi$. The unit-root test is conducted using the t-statistic for testing whether $\alpha$ =1 in the regression:

$$(4)\ \ y_t = \mu + \beta t + dD(T_b)_t + \theta DU_t + \gamma DT_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + u_t\ . \qquad \text{(C4)}$$

Again the inclusion of the additional dummy, $D(T_b)_t$ allows for the possibility of a break in the intercept under the null hypothesis.[25]

For both the AO and the IO models, the truncation parameter j is chosen by Perron's data-dependent method denoted by j(t-sig). The parameter j is allowed to take on a pre-specified maximum. A maximum lag of 6 years is used in this work. As is standard practice, for each value of $T_b$ the equations are initially estimated with the maximum number of lags and the final lag is eliminated if it is not significant at the 10% level. This sequential process is repeated until the coefficient on the last included lag term is significant but the coefficient on one additional lag term is not.

Test results are reported in Table B1 and the critical values at different significance levels are shown in Table B2. As can be seen from the tables none of the test statistics are sufficient to reject the null hypothesis of a unit-root at a 10% marginal significance level. Thus we are led to conclude that Mexican real GDP and money are not trend-break stationary over this sample period.

[25] Conclusions were unaffected by omitting this dummy variable.

**Table C1**
**Unit-root vs. Trend-Break Test Results**
**Mexican Aggregate Series**
**1932-2001**

| Model | Log Real GDP | Log M1 | Log M2 | Log M3 |
|---|---|---|---|---|
| IO Model | Test Stat = -4.56<br>$T_B$ = 1976; j = 1 | Test Stat = -4.20<br>$T_B$ = 1975; j = 6 | Test Stat = -4.71<br>$T_B$ = 1974; j = 8 | Test Stat = -4.60<br>$T_B$ = 1971; j = 7 |
| AO Model | Test Stat = -4.29<br>$T_B$ = 1975; j = 1 | Test Stat = -4.32<br>$T_B$ = 1969; j = 6 | Test Stat = -4.69<br>$T_B$ = 1966; j = 8 | Test Stat = -4.79<br>$T_B$ = 1967; j = 7 |

**Table C2**
**Approximate Finite Sample Critical Values**

| Model | 1.0% | 2.5% | 5.0% | 10% |
|---|---|---|---|---|
| IO[a] | -6.32 | -5.90 | -5.59 | -5.29 |
| AO[b] | -6.03 | -5.73 | -5.39 | -5.07 |

[a]From Perron (1997) Table 1.
[b]From Vogelsang & Perron (1998) Table 2.

## References


Andrews, D.W.K., 1989. Power in econometric applications. Econometrica 57, 1059-1090.

Bae, S. and R. Ratti, 2000, Long-run neutrality, high inflation, and bank insolvencies in Argentina and Brazil, Journal of Monetary Economics 46, 581-604.

Bai, B. and P. Perron, 1998, Estimating and testing linear models with multiple structural changes, Econometrica 66, 47-78.

Bai, B. and P. Perron, 2003, Computation and analysis of multiple structural change models, Journal of Applied Econometrics 18, 1-22.

Ball, L. and D. Romer, 2003, Inflation and the informativeness of prices, Journal of Money, Credit, and Banking 35, 177-196.

Banarjee, A., R.L. Lumsdaine, and J.H. Stock, 1992, Recursive and sequential tests of the unit-root and trend-break hypothesis: theory and international evidence, Journal of Business and Economic Statistics 10, 271-287.

Bank of Mexico, 2003, website http://www.banxico.org.mx/.

Boschen, J.F. and C.M. Otrok, 1994, Long run neutrality and superneutrality in an ARIMA framework: comment, American Economic Review 84, 1470-1473.

Bruno, M. and W. Easterly, 1998, Inflation crises and long run growth, Journal of Monetary Economics 41, 3-26.

Bullard, J., 1999, Testing long run monetary neutrality propositions: Lessons from recent research, Federal Reserve Bank of St.Louis Review, 57-78.

Christiano, L.J., 1992, Searching for a break in GNP, Journal of Business and Economic Statistics 10, 237-250.

Coe, P. J. and J.M. Nason, 2002, The long-horizon regression approach to monetary neutrality: how should the evidence be interpreted? Economics Letters 78, 351-356.

Elliott, G., T. Rothenberg, and J.H. Stock, 1996, Efficient tests for an autoregressive unit-root, Econometrica 64, 813-836.

Feliz, R. and J.H. Welch 1997, Cointegration and tests of a classical model of inflation in Argentina, Bolivia, Brazil, Mexico, and Peru, Journal of Development Economics 52, 189-219.

Fisher, M.E. and J.J. Seater, 1993, Long run neutrality and superneutrality in an ARIMA framework, American Economic Review 83, 402-415.

Haug, A.A. and R.F. Lucas, 1997, Long run neutrality and superneutrality in an ARIMA framework: comment, American Economic Review 87, 756-759.

Insituto Nacional de Estadística, Geografía, e Informática (INEGI), 1994, Estadísticas Históricas de México (CD-ROM), Mexico City, Mexico.

INEGI, 1999, Estadísticas Históricas de México. Volumes I and II. Mexico City, Mexico.

INEGI, 2003. Website "Banco de Información Económica." Available at http://www.inegi.gob.mx/difusion/espanol/fiecons.html.

Kwiatkowski, D., P.C.B. Phillips, P. Schmidt, and Y. Shin, 1992, Testing the null hypothesis of stationarity against the alternative of a unit-root, Journal of Econometrics 54, 159-178.

Lucas, R.E., 1972, Expectations and the neutrality of money, Journal of Economic Theory 4, 103-124.

Lustig, N., 2001, Life is not easy: Mexico's quest for stability and growth, Journal of Economic Perspectives 15, 85-106.

Murray, C.J., and C.R. Nelson, 2000, The uncertain trend in US GDP, Journal of Monetary Economics 46, 79-95.

Newey, W.K., and K.D. West, 1994a, A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix, Econometrica 55, 703-708.

Newey, W.K., and K.D. West, 1994b, Automatic lag selection in covariance matrix estimation, Review of Economic Studies 61, 631-653.

Olekalns, N., 1996, Some further evidence on the long run neutrality of money, Economics Letters 50, 393-398.

Perron, P., 1989, The great crash, the oil price shock and the unit-root hypothesis, Econometrica 57, 1361-1401; Erratum 1993, 61, 248-249.

Perron, P., 1997, Further evidence on breaking trend functions in macroeconomic Variables, Journal of Econometrics 80, 355-385.

Perron, P., and T.J. Vogelsang, 1992, Nonstationarity and level shifts with an application to purchasing power parity, Journal of Business and Economic Statistics 10, 301-320.

Rappoport, P., and L. Reichlin, 1989, Segmented Trends and Non-Stationary Time Series, The Economic Journal 99, 168-177.

Sen, A., 2003, On unit-root tests when the alternative is a trend-break stationary Process, Journal of Business and Economic Statistics 21, 174-184.

Vogelsang, T.J., and P. Perron, 1998. Additional tests for a unit-root allowing for a break in the trend function at an unknown time, International Economic Review. 39, 1073-1100.

Wallace, F.H., 1999, Long run neutrality of money in the Mexican economy, Applied Economics Letters 6, 637--40.

Wu, Y. and J. Zhang, 1998, Endogenous growth and the welfare costs of inflation: A reconsideration, Journal of Economic Dynamics and Control 22, 465-482.

Zivot, E., and D.W.K. Andrews, 1992, Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis, Journal of Business and Economic Statistics 10, 251-270.