

# **The Long-Run Relationship between Money, Nominal GDP, and the Price Level in Venezuela: 1950 to 1996**

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## **Abstract**

This paper explores whether a significant long-run relationship exists between money and nominal GDP and between money and the price level in the Venezuelan economy. We apply time-series econometric techniques to annual data for the Venezuelan economy for 1950 to 1996. An important feature of our analysis is the use of tests for unit roots and cointegration with structural breaks. Certain characteristics of the Venezuelan experience suggest that structural breaks may be important. Since the economy depends heavily on oil revenue, oil price shocks have had important influences on most macroeconomic variables. Also since the economy possesses large foreign debt, the world debt crisis that exploded in 1982 had pervasive effects on the Venezuelan economy. Radical changes in economic policy and political instability may have also significantly affected the movement of the macroeconomy. We find that a long-run relationship exists between narrow money (M1) and nominal GDP, the GDP deflator, and the CPI when one makes allowances for one or two structural breaks. We do not find such long-run relationships when broad money (M2) is used.

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## **1. Introduction**

Whether a long-run relationship exists between money and income or between money, income, and the interest rate crucially determines the role of money in the design and implementation of monetary policy. To use monetary aggregates as intermediate targets requires that they exhibit a long-run relationship with the final variable(s) that the monetary authority wants to affect. Though the topic has received extensive analysis in the U.S. and other countries, the issues are far from settled.

Studies apply cointegration techniques to U.S. data for different periods, different data frequencies, and different definitions of money. Miller (1991) uses the Engle-Granger technique to test for cointegration between different nominal monetary aggregates, real GNP, and the implicit price deflator - measured in natural logarithms - for the period 1959:I to 1987:IV. He also includes the four-to-six-month commercial-paper rate and the dividend-price ratio separately to test for cointegration. He finds a cointegrating relation only between M2, the real GNP, the implicit price deflator, and the four-to-six-month commercial-paper rate. Splitting the data in 1973:IV, Miller (1991) finds evidence of instability in the cointegrating vector. He suggests that structural changes may explain the lack of cointegration between some definitions of money and real GNP, the implicit price deflator, and the interest rate.

Friedman and Kuttner (1992) use Johansen technique to search for cointegration between real money balances and real income, and between real money balances, real income, and the commercial-paper rate for the period 1960:II to 1990:IV. They conclude that a long-run relationship exists between real M2 and real income, and between real M2, real income, and the commercial-paper rate for the whole period, but that extending the data to include the 1980's sharply weakens the evidence. They also show that tests using data from 1970 onward possess no evidence of cointegration.

Stock and Watson (1993) use annual data from 1900 to 1989 and find a stable long-run relationship between real M1, real net national product, and the commercial-

paper rate. They point out, however, that differences between the prewar and postwar estimates of the cointegrating vectors “raise the possibility that there has been a shift in the long-run money demand relation” (Stock and Watson, 1993, p. 804).

A common thread woven into these papers is that structural shifts may play an important role. None explicitly include, however, structural breaks in the estimated cointegrating vectors. During the 1990’s, unit-root and cointegration tests that explicitly allow for endogenously determined structural breaks have been developed. Boucher and Flynn (1997) examine whether a long-run relationship exists between real money, real income, and the interest rate using cointegration tests that allow for one structural break. Applying the Gregory-Hansen (1996) cointegration tests to quarterly data from 1959 to 1990, they find evidence of a stable long-run relationship between credit, income, and the Aaa bond rate with a breakpoint at 1980:I. Using an error-correction model derived from the Gregory-Hansen (1996) regime shift model, they find tentative evidence that M1 cointegrates with income and the six-month commercial-paper rate with a structural break at 1975:II.

Our paper explores whether a significant long-run relationship exists between money and nominal GDP, and between money and the price level for the Venezuelan economy. We apply time-series econometric techniques to annual data of the Venezuelan economy for the period 1950 to 1996. An important feature of our analysis is the extensive use of unit-root and cointegration tests that include structural breaks. Certain characteristics of the Venezuelan economy suggest that structural breaks may be important. Since the economy depends heavily on oil exports, oil shocks importantly affect most macroeconomic variables. Since the country possesses a large foreign debt, the world debt crisis that exploded in 1982 has pervasive effects on the Venezuelan economy. Radical changes in economic policy and political instability may also significantly affect the behavior of the economy.

Our paper contains five sections. Section 1 develops a simple partial-equilibrium model of the money market and describes how this market adjusts to changes in the money supply under different exchange rate regimes. We compare the results of the model to the actual movement of money, prices, and the exchange rate in the Venezuelan economy. Section 2 presents the results of applying unit-root tests to the logarithms of the M1 and M2 velocities of circulation. After testing for the orders of integration of the variables considered, section 3 applies residual-based cointegration tests to the pairwise relations between money - M1 and M2 - and nominal GDP, and between money and the price level - the CPI and the GDP deflator. Section 4 examines Granger causality tests and innovation accounting using vector error-correction models for the pairwise-cointegrated relationships between money and the nominal GDP/price level found. Finally, section 5 concludes with some policy considerations.

## 2. Money Supply Changes and Equilibrium in the Money Market<sup>1</sup>

This section considers how the money market adjusts when the money supply changes. Such adjustment depends on the exchange rate regime.

Consider the following money demand function:

$$M^d/P = f(y, i), \quad (1)$$

where

$M^d$  = nominal money demand:

P = price level:

Y = real income; and

i = nominal interest rate.

This money demand function is expressed in nominal terms as follows:

$$M^d = Pf(y, i). \quad (2)$$

The money supply conforms to the monetary base - multiplier model as follows:

$$M^s = kB, \quad (3)$$

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<sup>1</sup> The basic model in this section is developed in Sachs and Larrain (1993, chapter 9, section, 9.5).

where

$M^s$  = nominal money supply;

$k$  = money multiplier; and

$B$  = monetary base.

Equilibrium in the money market requires that money demand equals money supply:

$$M^d = Pf(y, i) = kB = M^s. \quad (4)$$

Figure 1 shows equilibrium in the money market in terms of the price level. The supply of money is a vertical line as it is independent of the price level. The demand for money is shown as a straight line starting at the origin as the real value of desired money balances does not change with variations in the price level. In this graph, along the demand for money function, the level of income ( $y$ ) and the rate of interest ( $i$ ) are constant. Any change in these variables shifts the demand for money schedule. At point A, the demand for money equals to the money supply, and the price level is  $P_0$ .

Consider now an open market operation that expands the monetary base and the money supply. At the initial price level, interest rate, and level of income, an excess supply of money exists (AB in Figure 2). Equilibrium reestablishes itself in the money market in different ways depending crucially on the exchange rate regime.

Figure 2 depicts the adjustment mechanism under a floating exchange rate. Starting at point A, the expansionary monetary policy shifts the money supply schedule from  $M_0^s$  to  $M_1^s$ . If the economy possesses some degree of price stickiness, then the money market clears in the short-run by a fall in the interest rate (point B). The lower interest rate reduces the velocity of circulation ( $V$ ), and given that the slope of the money demand schedule is  $V/y$ ,  $M^d$  rotates downward to  $M^{d'}$ . In short-run equilibrium, the price level is unchanged. In the long-run, equilibrium emerges through an increase in the price level from  $P_0$  to  $P_1$  that equates the nominal money demand with the higher money supply (point C), while the interest rate returns to its initial value.

An expansion of the money supply has different effects under a fixed exchange rate. As shown in Figure 3, the excess of money generated by the shift from  $M_0^s$  to  $M_1^s$  reduces the stock of international reserves as economic agents purchase foreign assets, goods, and services in an attempt to reestablish their desired stock of real money balances. This, in turn, shifts the money supply schedule back to its starting position ( $M_0^s$ ). In the final equilibrium, the stock of money and the price level remain unaltered. Under a fixed exchange rate regime, the monetary authority loses control over the money supply, which becomes an endogenous variable.

Note, however, that the previous results depend crucially on several assumptions: perfect capital mobility, low barriers to trade, and the absence of exchange controls. In particular, the combination of a fixed exchange rate with exchange controls moves the adjustment to an expansion in the money supply toward that of an economy with a flexible exchange rate (Figure 2). Exchange controls make it harder, though not impossible, to reduce the excess supply of money caused by an expansionary monetary policy through transactions that deplete international reserves. Instead, equilibrium in the money market is partially achieved in the long run by a rise in the price level that reestablishes the original stock of real money balances.

This combination of a fixed exchange rate with exchange controls mirrors the situation frequently used by Venezuelan policymakers. Exchange controls were implemented from 1960 to 1964, 1983 to 1988, and 1994 to 1996. The last two periods of exchange controls were accompanied by expansive monetary policies (Figure 4 compares the growth rates of M1 and the inflation rates in Venezuela and the United States from 1951 to 1996).

The hypothesis examined in this paper is as follows: Venezuela possesses stable long-run relationships between money, nominal income, and the price level for the period 1950 - 1996. Periods of fixed exchange rates with no controls correspond, in general, to periods of relatively low monetary expansion, while periods of expansive monetary

policies have been accompanied by fixed exchange rates combined with exchange controls, or more flexible exchange rates (i.e. 1989 - 1993). An exception is the period 1973 - 1982, during which a relatively high rate of monetary growth is observed, and the fixed exchange rate was maintained without exchange controls. This can be explained by the large net inflows of capital that the Venezuelan economy experienced during most of this period.

### 3. Is The Velocity of Circulation Stationary?

Typical approaches to examining the relationship, if any, between money and predetermined final variables (i.e., real GDP, nominal GDP, or the price level) estimate money demand functions or “St. Louis” type equations. In this section, we apply unit-root tests to the velocity of circulation, a different approach. If the velocity of circulation contains a unit root, random shocks have permanent effects on this variable, suggesting an unstable connection between money and nominal income.

Two measures of velocity -- M1 and M2-- are analyzed using annual data from Venezuela for the period 1951 to 1996. Velocity in logs is defined as follows:

$$\text{Velocity} = V = \ln[\text{GDP}_t / ((M_t + M_{t-1})/2)], \quad (5)$$

where

$\text{GDP}_t$  = nominal GDP; and

$M_t$  = nominal money supply.

Given that the nominal GDP is a flow variable, and that monetary aggregates are stocks, we used an average of the latter [i.e.,  $(M_t + M_{t-1})/2$ ] in measuring the velocity of circulation.

Figure 5 shows the two measures of velocity, LVM1 and LVM2. The graph suggests a relatively close relation between these two measures of velocity. A simple regression between LVM1 and LVM2 yields a  $R^2$  of 0.46, significantly different from zero ( $F = 36.77$ ), but small enough to indicate a different reaction of these variables to changes in those common factors affecting them.

The results of the standard augmented Dickey-Fuller tests (ADF) are presented in Table 1. Using the MacKinnon critical values, the  $t_{\alpha}$  statistics for both series do not reject the null hypothesis of a unit root. In choosing  $k$  in the different tests, we have followed Perron's (1989) strategy. Given the following model:

$$y_t = \mu + \alpha y_{t-1} + \sum_{j=1}^k c_j \Delta y_{t-j} + e_t \quad (6)$$

a value of  $k = k^*$  was selected, if the  $t$ -statistic on  $c_j$  was greater than 1.60 in absolute value, and the  $t$ -statistic on  $c_l$  for  $l > k^*$  was less than 1.60, starting with a maximum value of  $k = 4$ . A deterministic trend is also included in the Dickey-Fuller tests, if it presents a  $t$ -statistic greater than two in absolute value.<sup>2</sup>

It is evident from Figure 5 that both series present large swings, mainly around 1973 to 1976 and 1986 to 1989. Perron (1989) has shown that when a series contains a significant shift, it is difficult to reject the unit-root hypothesis, even if the series is stationary before and after the shift. He proposes three models to test for a unit root while allowing for a one-time change in the structure occurring at time  $T_B$  ( $1 < T_B < T$ ):<sup>3</sup>

Model A (Change in the intercept) (7)

$$\text{Null hypothesis: } y_t = \mu + dD(TB)_t + y_{t-1} + e_t$$

$$\text{Alternative hypothesis: } y_t = \mu_1 + \beta t + (\mu_2 - \mu_1)DU_t + e_t$$

Model B (Change in the slope) (8)

$$\text{Null hypothesis: } y_t = \mu_1 + y_{t-1} + (\mu_2 - \mu_1)DU_t + e_t$$

$$\text{Alternative hypothesis: } y_t = \mu + \beta_1 t + (\beta_2 - \beta_1)DT_t + e_t$$

Model C (Change in intercept and slope) (9)

$$\text{Null hypothesis: } y_t = \mu_1 + y_{t-1} + dD(TB)_t + (\mu_2 - \mu_1)DU_t + e_t$$

$$\text{Alternative hypothesis: } y_t = \mu_1 + \beta_1 t + (\mu_2 - \mu_1)DU_t + (\beta_2 - \beta_1)DT_t + e_t$$

<sup>2</sup> Enders (1995) describes a more elaborate procedure to choose the deterministic regressors included in the Dickey-Fuller tests.

<sup>3</sup> Perron (1989) unit-root tests can be implemented in two steps. In the first step, the raw series is detrended using the null hypothesis equation for either model A, B, or C. In the second step, an equation like equation (6), but without the constant, is estimated using the detrended series. The  $t$ -statistic for the null hypothesis  $\alpha=1$  can be compared to the critical values derived by Perron (1989).

where

$$D(TB)_t = 1 \text{ if } t = T_B + 1, 0 \text{ otherwise;}$$

$$DU_t = 1 \text{ if } t > T_B, 0 \text{ otherwise;}$$

$$DT_t^* = t - T_B \text{ if } t > T_B, 0 \text{ otherwise; and}$$

$$DT_t = t \text{ if } t > T_B, 0 \text{ otherwise.}$$

We apply these three models to the series LVM1 and LVM2 using 1973 as breakpoint. The choice of 1973, the year of the first oil shock, as the time of the structural break reflects the paramount importance of oil production for the Venezuelan economy. The results of applying Perron's (1989) approach appear in Table 2. Using the critical values for  $t_\alpha$  derived by Perron with  $\lambda = 0.5$ , we cannot reject the null hypothesis of a unit root for any of the series even at the 10-% level.

Zivot and Andrews (1992) criticize Perron's method of setting the breakpoint exogenously. They propose instead a data-dependent procedure to determine the time of the structural break. Their strategy chooses as the breakpoint the one that gives the smallest one-sided t-statistic for testing  $\alpha = 1$ . As in Perron's (1989) paper, Zivot and Andrews (1992) present three models in order to test for a unit root with an endogenously determined structural break. The null hypothesis for the three models is:

$$y_t = \mu + y_{t-1} + e_t \quad (10)$$

The alternative hypotheses for the different models are:

$$\text{Model A: } y_t = \mu + \theta DU_t(\lambda) + \beta t + \alpha y_{t-1} + \sum_{j=1}^k c_j \Delta y_{t-j} + e_t; \quad (11)$$

$$\text{Model B: } y_t = \mu + \gamma DT_t^*(\lambda) + \beta t + \alpha y_{t-1} + \sum_{j=1}^k c_j \Delta y_{t-j} + e_t; \text{ and} \quad (12)$$

$$\text{Model C: } y_t = \mu + \theta DU_t(\lambda) + \beta t + \gamma DT_t^*(\lambda) + \alpha y_{t-1} + \sum_{j=1}^k c_j \Delta y_{t-j} + e_t, \quad (13)$$

where

$$DU_t(\lambda) = 1 \text{ if } t > T\lambda, 0 \text{ otherwise, and}$$

$$DT_t^*(\lambda) = t - T\lambda \text{ if } t > T\lambda, 0 \text{ otherwise.}$$

Table 3 presents the results of applying Zivot and Andrews's method to the series LVM1 and LVM2. Using the critical values of  $t_{\alpha}(\lambda)$  derived by Zivot and Andrews, we still cannot reject the unit-root hypothesis for any of the series even at a 10-% level.

Lumsdaine and Papell (1996) extend the endogenous break methodology to allow for two-breaks. Given that both LVM1 and LVM2 show at least two large swings over the period of analysis, this approach is particularly relevant. They propose the following three models for the alternative hypothesis:

$$\text{Model AA: } y_t = \mu + \theta DU1_t + \omega DU2_t + \beta t + \alpha y_{t-1} + \sum_{j=1}^k c_j \Delta y_{t-j} + e_t; \quad (14)$$

$$\text{Model CC: } y_t = \mu + \theta DU1_t + \gamma DT1_t + \omega DU2_t + \psi DT2_t + \beta t + \alpha y_{t-1} + \sum_{j=1}^k c_j \Delta y_{t-j} + e_t; \text{ and} \quad (15)$$

$$\text{Model CA: } y_t = \mu + \theta DU1_t + \gamma DT1_t + \omega DU2_t + \beta t + \alpha y_{t-1} + \sum_{j=1}^k c_j \Delta y_{t-j} + e_t, \quad (16)$$

where

$$DU1_t = 1 \text{ for } t > T_B 1=1, 0 \text{ otherwise;}$$

$$DU2_t = 1 \text{ for } t > T_B 2=1, 0 \text{ otherwise;}$$

$$DT1_t = t - T_B 1 \text{ for } t > T_B 1=1, 0 \text{ otherwise; and}$$

$$DT2_t = t - T_B 2 \text{ for } t > T_B 2=1, 0 \text{ otherwise.}$$

The results of applying Lumsdaine and Papell's approach to the two measures of velocity appears in Table 4. Using model AA that allows for two breaks in the intercept, it is possible to reject the unit-root hypothesis for the log of the velocity of M1 at the 5-% level of significance. With model CA in which the first break is a change in intercept and slope and the second break is a change in intercept only, the unit-root hypothesis can also be rejected at a 5-% level of significance. Therefore, allowing for two breakpoints, we find evidence that the logarithm of the velocity of circulation of M1 in Venezuela during

the period 1950 to 1996 is stationary. This suggests a stable relationship between M1 and the nominal GDP during the period analyzed.

In contrast, the logarithm of the velocity of circulation of M2 does not seem to behave as a stationary variable, even when the possibility of two structural breaks is considered.

### **3. Cointegration Analysis**

A complementary approach to the application of unit root tests to the velocity of circulation checks for cointegration between the monetary aggregates, nominal GDP, and price level. In this section, we apply residual-based tests for cointegration between these variables using Venezuelan annual data for the period 1950 to 1996.

The first step in this method determines the order of integration of the variables under study. Table 5 shows the results of the standard ADF tests for the logs of M1 (LM1), M2 (LM2), nominal GDP (LGDP), GDP deflator (LDGDP), and the CPI (LCPI). It is quite evident from the positive  $t_{\alpha}$  values that the null hypothesis of a unit root cannot be rejected for any of the variables.

Following the analysis of the velocity of circulation, we apply unit-root tests considering structural breaks to these series. As shown in Tables 6, 7, and 8, we cannot reject the unit-root hypothesis for the variables in levels using either the Perron (1989), Zivot-Andrews (1992), or Lumsdaine-Papell (1996) unit-root testing strategies.

We also tested the first difference of these variables for unit roots. The standard ADF tests (Table 9) indicate that the first difference of the log of all variables, with the exception of the first differences of the log of GDP (DLGDP) and CPI (DLCPI), are stationary at the 1-% or 5-% levels of significance.

Given these results, we apply unit root tests that account for structural breaks to the series DLGDP and DLCPI. Perron (1989) unit-root tests with one exogenous breakpoint ( $T_B=1973$ ) allow the rejection of the unit-root hypothesis for DLGDP at the 5-

% level with models B and C (Table 10). For DLCPI, we can only reject the unit-root hypothesis at the 10-% level with model C (Table 10).

Zivot and Andrews (1992) unit-root tests with one endogenous breakpoint allow the rejection of the unit-root hypothesis for DLGDP at the 1-% level with either model B or model C, and at the 5-% level with model A (Table 11). In the case of DLCPI, we can reject the unit root hypothesis at the 1-% level with model C, and at the 5-% level with model B (Table 11).

Given that the unit-root tests support the position that the logarithms of the variables in levels are I(1), we tested whether the following pair of variables are cointegrated: LGDP/LM1, LDGDP/LM1, LCPI/LM1, LGDP/LM2, LDGDP/LM2, and LCPI/LM2

Table 12 reports the results from applying the Engle-Granger cointegration tests with a constant (C), and constant and trend (T) in the cointegrating vectors. Using the MacKinnon critical values, we cannot reject the null hypothesis of no cointegration for any of the pair of variables in Table 12 even at the 10-% level.

Similar to velocity of circulation, we may want to consider structural breaks in the relations between money, income, and the price level. Gregory and Hansen (1996) proposed residual-based cointegration tests that allow the cointegrating vectors to change at a single unknown time during the sample period. Their tests, as they note, are multivariate extensions of the univariate tests of Perron (1989), and Zivot and Andrews (1992).

Gregory and Hansen (1996) proposed models to test for cointegration in the presence of a single structural break-assuming only two variables to simplify-can be expressed as follows:

$$\text{Level Shift (C) Model: } y_t = \mu_1 + \mu_2 \phi_t(\lambda) + \alpha x_t + e_t; \quad (17)$$

$$\text{Level Shift with Trend (C/T) Model: } y_t = \mu_1 + \mu_2 \phi_t(\lambda) + \beta t + \alpha x_t + e_t; \text{ and } (18)$$

$$\text{Regime Shift (C/S) Model: } y_t = \mu_1 + \mu_2 \phi_\tau(\lambda) + \alpha x_t + \alpha x_\tau \phi_\tau(\lambda) + e_t, (19)$$

where

$$\varphi_t(\lambda) = 1 \text{ if } t > T\lambda, 0 \text{ otherwise.}$$

The results of applying Gregory and Hansen (1996) cointegration tests to the pairwise relations previously mentioned, are shown in Table 13. Accounting for structural breaks, the ADF<sup>\*</sup> (m=1) critical values derived by Gregory and Hansen (1996) indicate that the null hypothesis of no cointegration between LGDP and LM1 is rejected at the 5-% level for the level shift model (C), and at the 10-% level for the level shift with trend (C/T) and regime shift (C/S) models. LM1 cointegrates with LDGDP and LCPI at the 10-% level only when considering the regime shift model (C/S). In contrast, we still cannot reject the hypothesis of no cointegration between LM2 and LGDP, LDGDP, and LCPI by including a single breakpoint in the cointegrating vector. The residuals from the cointegrating vectors for the variables that cointegrate are shown in Figures 6, 7, 8, 9, and 10.

These results mirror those obtained in section 2, showing that structural breaks provide support for the hypothesis that a long-run equilibrium relationship exists between money (defined as M1), income, and prices in Venezuela. Note, however, that CUSUM of squares tests applied to the residuals from the cointegrating vectors from the Gregory-Hansen (1996) method indicate some instability for the LGDP/LM1 and LCPI/LM1 relationships around 1985 to 1989 and 1986 to 1988, respectively. Only for the LDGDP/LM1 relation does no indication of instability according to that test exist. Those findings suggest the possibility of one additional structural break in the cointegrating vectors for the LGDP/LM1 and LCPI/LM1 relationships. The problem is that, as far as we know, the critical values to test the null hypothesis of no cointegration in the presence of two or more structural breaks are not yet derived.

#### **4. Vector Error-Correction Models**

This section contains vector error-correction models (VEC) for the pairwise-cointegrated relationships. Using the VEC, we assess the direction of Granger causality and the effects

of innovations on the variables under analysis. In estimating the VEC models, we started with a maximum of four lags, and used likelihood ratio tests (Enders 1995), and Akaike (AIC) and Schwarz (SC) criteria to determine the appropriate numbers of lags.

For the relationship between DLGDP and DLM1, both the Akaike and Schwarz criteria favor the VEC model with zero lags, while the likelihood ratio test rejects the null hypothesis of zero lags in favor of one lag. Given that the residuals of the model with one lag better approximate a white-noise process, we use the former. The results are as follows (t-statistics in parenthesis):

$$\text{DLM1} = 0.06 + 0.045\text{DLM1}(-1) + 0.641\text{DLGDP}(-1) + 0.245\text{CRGDPM1}(-1); \quad (20)$$

(1.87) (0.27)                      (3.20)                      (1.74)

and

$$\text{DLGDP} = 0.0276 + 0.0113\text{DLM1}(-1) + 0.891\text{DLGDP}(-1) - 0.295\text{CRGDPM1}(-1), \quad (21)$$

(1.01)      (0.08)                      (5.22)                      (-2.448)

where

$$\text{CRGDPM1}(-1) = \text{LGDP}(-1) - 1.119*\text{LM1}(-1) + 0.659*\text{DU73}(-1) - 1.231$$

is the one-period lagged value of the residuals from the cointegrating vector with a level shift. The signs of the speed-of-adjustment coefficients are as expected for long-run equilibrium.

These results indicate that there is Granger causality in both directions. In equation (20), the coefficient of DLGDP(-1) is significantly different from zero at the 5-%, and the coefficient of CRGDPM1(-1) at the 10-%, level. In equation (21), only the coefficient of CRGDPM1(-1) is significantly different from zero the 5-% level .

Given that the residuals from the equations in the VEC possess a high correlation ( $r = 0.58$ ), we entertain innovation accounting for two possible orderings: DLM1-

DLGDP, and DLGDP-DLM1. With the ordering DLM1-DLGDP, the impulse response functions of DLM1 to its own shock and a shock to DLGDP show a slow but continuous convergence toward zero (top graph in the left side of Figure 11). A similar pattern appears for the impulse response functions of DLGDP (bottom graph in the left side of Figure 4.1). This slow convergence can be attributed to the relatively large coefficients that DLGDP(-1) exhibits in both equations (20) and (21). Using the ordering DLGDP-DLM1, the impulse response functions look noticeably different (graphs in the right side of Figure 11). In particular, shocks to DLM1 have no impact on DLGDP.

With the ordering DLM1-DLGDP, the variance decomposition of DLM1 (top graph in the left side of Figure 12) indicates that the percentage of the forecast error explained by its own shocks decreases significantly in the first 10 periods, and then stabilizes around 57 %. The percentage of the forecast error of DLGDP (lower graph in the left side of Figure 12) explained by its own shocks is fairly stable around 70% during a 20 period horizon. Reversing the ordering, the most noticeable change is that the variance decomposition of DLGDP indicates that its forecast error is completely explained by its own shocks (lower graph in the right side of Figure 12). In sum, the position of a variable in the ordering affects its explanatory power; first in the ordering implies higher explanatory power than second in the ordering, not a surprising result.

For the relationship between DLDGDP and DLM1, the Akaike and Schwarz criteria choose a model with zero lags, while the likelihood ratio test favors a model with one lag. As in the case of DLGDP and DLM1, the residuals of the VEC with one lag better approximate a white-noise process. Therefore, we chose to estimate the model with one lag. The results are as follows:

$$\begin{aligned}
 \text{DLM1} &= 0.093 + 0.096\text{DLM1}(-1) + 0.524\text{DLDGDP}(-1) \\
 &\quad (3.23) \quad (0.59) \quad (3.10) \\
 &\quad + 0.373\text{CRDGDPM1}(-1); \quad (22) \\
 &\quad (2.10)
 \end{aligned}$$

and

$$\begin{aligned}
 \text{DLGDP} = & 0.034 - 0.026\text{DLM1}(-1) + 0.86\text{DLDGDP}(-1) \\
 & (1.26) \quad (-0.17) \qquad \qquad (5.34) \\
 & - 0.468\text{CRDGDPM1}(-1), \qquad \qquad \qquad (23) \\
 & (-2.76)
 \end{aligned}$$

where

$$\begin{aligned}
 \text{CRDGDPM1}(-1) = & \text{LDGDP}(-1) - 0.278\text{LM1}(-1) + 8.597\text{DU76}(-1) \\
 & - 0.839\text{SLM176}(-1) - 0.7
 \end{aligned}$$

is the one-period lagged value of the residuals from the cointegrating vector obtained from the regime-shift model. The signs of the speed-of-adjustment coefficients match prior expectations. Moreover, the speeds of adjustment toward long-run equilibrium exceed those for the VEC of DLM1 and DLGDP.

These results indicate two-way Granger causality. In the DLM1 equation, the coefficients of DLDGDP(-1) and CRDGDPM1(-1) are both significantly different from zero at the 5-% level. In the DLDGDP equation, the coefficient of CRDGDPM1(-1) is significantly different from zero at the 5-% level. Also the speed-of-adjustment coefficient in the equation for DLDGDP exceeds that in the equation for DLM1.

As in the previous model, the residuals from the equations in the VEC possess high correlation ( $r = 0.55$ ). Hence, we report innovation accounting using the orderings DLM1-DLDGDP and DLDGDP-DLM1. With the ordering DLM1-DLDGDP, the impulse-response functions of DLM1 and DLDGDP (graphs in the left side of Figure 13) exhibit a slow but continuous convergence towards zero. Given that the coefficients of DLDGDP(-1) in equations (22) and (23) are smaller than in equations (20) and (21), convergence is faster in this model.

The variance decomposition of DLM1 (top graph in the left side of Figure 14) indicates that after decreasing in the first 10 periods, the percentage of the forecast error explained by its own shocks stabilizes around 66% for the rest of the 20 period horizon.

The percentage of the forecast error of DLDGDP (lower graph in the left side of Figure 14) explained by its own shocks is fairly stable around 69% during a 20 period horizon. Using the ordering DLDGDP-DLM1, the impulse response functions and the variance decompositions indicate that shocks to DLM1 have little impact on DLDGDP (graphs on the right sides of Figures 13 and 14). Once again, the position in the ordering matters; first in the ordering implies more importance than second.

For DLM1 and DLCPI, we cannot specify a standard VEC model that satisfies all the diagnostic checks. The Akaike and Schwarz criteria favor a VEC with no lags. This specification yields residuals for DLCPI that, however, do not approximate a white-noise process. The likelihood ratio test favors a VEC with one lag, but the impulse-response functions from this model are explosive. Moreover, extended models with two to six lags also exhibit explosive impulse-response functions. For this reason, we estimated a near VAR error-correction model (Enders (1995), p. 313). In this model, the DLM1 equation has no lags, while the DLCPI equation has one lag. This model yields residuals for DLCPI that better approximate a white-noise process than a VEC with zero lags, and has convergent impulse response functions. The estimation using seemingly unrelated regressions (SUR) yielded the following results:

$$\begin{aligned} \text{DLM1} &= 0.159 + 0.591\text{CRCPI}1(-1); \text{ and} & (24) \\ & (7.11) \quad (3.16) \end{aligned}$$

$$\begin{aligned} \text{DLCPI} &= 0.032 - 0.057\text{DLM1}(-1) + 0.922\text{DLCPI}(-1) \\ & (1.82) \quad (-0.63) \quad (9.81) \\ & - 0.216\text{CRCPI}1(-1), & (25) \\ & (-1.978) \end{aligned}$$

where

$$\begin{aligned} \text{CRCPI}1(-1) &= \text{LCPI}(-1) - 0.207\text{LM1}(-1) + 10.431\text{DU74}(-1) \\ & - 1.013\text{SLM174}(-1) - 1.577 \end{aligned}$$

is the lagged value of the residuals of the cointegrating vector from the regime shift model. The speed-of-adjustment coefficients possess the signs necessary to achieve long-run equilibrium. Moreover, the speed-of-adjustment coefficient for DLCPI is lower than for DLDGDP.

The results suggest Granger causality in both directions. The coefficient of CRCPIM1(-1) is significantly different from zero at the 5-% level in the equation for DLM1, and at the 10-% (nearly 5-%) level in the equation for DLCPI.

Given that the DLM1 equation contains no lags, convergence to long-run equilibrium occurs in one period. As in the previous cases, the residuals from the equations in this model possess high correlation ( $r = 0.51$ ). Therefore, we report innovation accounting for DLCPI using the orderings DLM1-DLCPI and DLCPI-DLM1. The graphs in the left side of Figures 15 and 16 show that with the ordering DLM1-DLCPI, shocks to DLM1 affect the behavior of DLCPI. When the ordering is reversed, however, shocks to DLM1 have no impact on DLCPI (graphs in the right side of Figures 15 and 16).

We do not have a clear-cut explanation for the substantial differences observed in innovation accounting when the ordering of the variables is altered. We suspect, however, that the frequent changes in exchange rate regimes observed in the Venezuelan economy during the period analyzed may explain these results.

## **5. Conclusions and Some Policy Considerations**

This paper examines whether a long-run relationship exists between money and nominal income, and money and the price level in the Venezuelan economy during the period 1950 to 1996. Using time-series econometric techniques, we find evidence that such relationships exist for the narrow aggregate M1 when due allowance for structural breaks is made. Unit-root tests with two endogenously determined breakpoints indicate that the log of the velocity of M1 is stationary. A complementary approach, applying residual based cointegration tests that allow for one structural break in the cointegrating vector,

indicates that M1 cointegrates with the nominal GDP, the GDP deflator, and the CPI. The breakpoints determined endogenously can, in general, be related to relevant economic events of the Venezuelan economy. In contrast, the results obtained using M2 do not suggest a significant long-run relationship between this broader aggregate and nominal income, or the price level.

The vector error-correction models estimated using the residuals from the cointegrating vectors between LM1/LGDP, LM1/LDGD, and LM1/LCPI yield speed-of-adjustment coefficients with the signs necessary for long-run equilibrium. Moreover, in each model, those coefficients are significantly different from zero, at least at the 10-% level. That implies two-way Granger causality between the natural logarithm of M1 and the natural logarithm of GDP, DGDP, and CPI. Innovation accounting shows that the variables converge toward equilibrium after a shock, although we had to use a non-standard vector error-correction specification for the M1-LCPI model to eliminate instability.

Even though a detailed discussion of the selection of an intermediate target for monetary policy is beyond the scope of this paper, we will offer some remarks about the policy implications of our results. Friedman and Kuttner (1992) note that “the quantity of money or its growth rate, can play a useful role in the monetary process only to the extent that fluctuations in money over time regularly and reliably correspond to fluctuations in income, prices, or whatever other aspects of economic activity the central bank seeks to influence”. (p. 472). Our empirical results support the contention that a regular and reliable link between money, nominal income, and prices exists in the Venezuelan economy. Therefore, money narrowly defined contains relevant information, if the final objective of monetary policy is to influence the rate of growth of nominal GDP, or directly the rate of inflation.

This is a notable finding, since most economists and policymakers in Venezuela consistently reject the view that the acceleration of inflation, especially since the 80s,

relates to monetary factors. For about 20 years, policymakers have used the nominal exchange rate as the nominal anchor without adopting the necessary discipline in fiscal and monetary policies. That policymakers err does not indict the fixed exchange rate system. Hence, we can still argue that a fixed exchange rate regime, appropriately implemented, provides a good strategy for monetary policy in Venezuela. The problem is that other factors in the Venezuelan economy weaken the credibility of a fixed exchange rate. As de Grauwe (1996) explains, the collapse of fixed exchange regimes frequently reflects adjustment problems. A country with a fixed exchange rate that encounters a negative shock producing a balance of payment deficit can only eliminate this deficit by reducing aggregate demand. If policymakers adjust the exchange rate to avoid the undesirable costs from adjusting aggregate demand, the commitment to a fixed exchange rate weakens. Those adjustment problems are particularly relevant for the Venezuelan economy whose exports depend fundamentally on one product (oil)<sup>4</sup>. Figure (17) compares the percentage rate of change of the terms of trade of Venezuela and the United States, showing that fluctuations in the terms of trade of Venezuela substantially exceed those of the United States.

The evidence in favor of a long-run relationship between money and nominal income, and between money and the price level, together with the adjustment problems of the fixed exchange rate system, should lead naturally to a reconsideration of the monetary policy strategy in Venezuela.

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<sup>4</sup>From 1993 to 1995, oil exports represented an average of 73.4% of total exports

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**TABLES:**

**Table 1: Standard Augmented Dickey-Fuller Tests**

Series	t	k	Constant	Trend
LVM1	-1.506	3	Yes	No
LVM2	-1.53	1	Yes	No

- a Significant at the 1-% level  
 b Significant at the 5-% level  
 c Significant at the 10-% level

**Table 2: Perron (1989) Unit-Root Tests with One Structural Break ( $T_B=1973$ )**

Series	Model A		Model B		Model C	
	t	k	t	k	t	k
LVM1	-2.519	0	-1.96	3	-3.225	0
LVM2	-1.465	0	-0.915	1	-2.03	0

- a Significant at the 1-% level  
 b Significant at the 5-% level  
 c Significant at the 10-% level

**Table 3: Zivot & Andrews (1992) Unit Root-Tests with One Structural Break**

Series	Model A			Model B			Model C		
	TB	t	k	TB	t	k	TB	t	k
LVM1	1988	-3.652	3	1986	-3.357	3	1974	-4.139	0
LVM2	1991	-1.535	1	1988	-2.415	1	1981	-2.345	0

- a Significant at the 1-% level  
 b Significant at the 5-% level  
 c Significant at the 10-% level

**Table 4: Lumsdaine & Papell (1996) Unit-Root Tests with Two Structural Breaks**

Series	Model AA			Model CC			Model CA		
	TB1/2	t	k	TB1/2	t	k	TB1/2	t	k
LVM1	74/88	-6.93b	3	74/88	-6.38	3	74/88	-6.73b	3
LVM2	74/91	-2.85	1	74/81	-2.99	0	74/81	-2.74	0

- a Significant at the 1-% level  
b Significant at the 5-% level  
c Significant at the 10-% level

**Table 5: Standard Augmented Dickey-Fuller Tests**

Series	t	k	Constant	Trend
LM1	4.175	0	Yes	No
LM2	2.392	1	Yes	No
LGDP	4.309	0	Yes	Yes
LDGDP	4.506	4	Yes	No
LCPI	5.071	4	Yes	No

- a Significant at the 1-% level  
b Significant at the 5-% level  
c Significant at the 10-% level

**Table 6: Perron (1989) Unit-Root Tests with One Structural Break ( $T_B=1973$ )**

Series	Model A		Model B		Model C	
	t	k	t	k	t	k
LM1	-0.114	3	-2.299	3	-2.439	4
LM2	-0.703	4	-2.942	4	-2.715	3
LGDP	0.156	2	-1.02	2	-1.289	2
LDGDP	0.318	2	-0.903	2	-0.764	1
LCPI	0.076	2	-1.16	2	-1.362	2

- a Significant at the 1-% level  
b Significant at the 5-% level  
c Significant at the 10-% level

**Table 7: Zivot & Andrews (1992) Unit-Root Tests with One Structural Break**

Series	Model A			Model B			Model C		
	TB	t	k	TB	t	k	TB	t	k
LM1	1993	-1.968	3	1988	-1.699	3	1989	-1.715	3
LM2	1988	-1.525	1	1984	-1.781	1	1984	-1.763	1
LGDP	1992	1.02	4	1984	-1.294	0	1982	-1.317	0
LDGDP	1986	1.618	3	1983	-0.626	0	1986	-0.363	4
LCPI	1991	2.53	4	1985	0.511	4	1984	0.433	4

- 
- a Significant at the 1-% level  
b Significant at the 5-% level  
c Significant at the 10-% level

**Table 8: Lumsdaine & Papell's (1996) Unit-Root Tests with Two Structural Breaks**

Series	Model AA			Model CC			Model CA		
	TB1/2	t	k	TB1/2	t	k	TB1/2	t	k
LM1	73/93	-2.831	4	73/89	-3.448	3	73/89	-2.992	0
LM2	73/88	-2.307	1	72/84	-3.47	1	73/84	-3.688	1
LGDP	86/92	0.077	4	82/93	-2.108	0	82/95	-2.12	0
LDGDP	86/93	-0.941	0	70/86	-2.494	4	73/86	-1.973	4
LCPI	82/91	1.258	4	71/84	-3.693	4	78/84	-1.594	4

- 
- a Significant at the 1-% level  
b Significant at the 5-% level  
c Significant at the 10-% level

**Table 9: Standard Augmented Dickey-Fuller Tests**

Series	t	k	Constant	Trend
DLM1	-5.608a	0	Yes	Yes
DLM2	-3.855b	0	Yes	Yes
DLGDP	-3.205c	0	Yes	Yes
DLDGDP	-4.033b	0	Yes	Yes
DLCPI	1.976	4	Yes	No

- 
- a Significant at the 1-% level  
b Significant at the 5-% level  
c Significant at the 10-% level

**Table 10: Perron (1989) Unit-Root Tests with One Structural Break ( $T_B=1973$ )**

Series	Model A		Model B		Model C	
	t	k	t	k	t	k
<b>DLGDP</b>	-3.543c	0	-4.309b	0	-4.48b	0
<b>DLCPI</b>	-0.996	4	-2.181	4	-3.991c	2

- a Significant at the 1-% level  
b Significant at the 5-% level  
c Significant at the 10-% level

**Table 11: Zivot & Andrews (1992) Unit-Root Tests with One Structural Break**

Series	Model A			Model B			Model C		
	TB	t	k	TB	t	k	TB	t	k
<b>DLGDP</b>	1986	-5.214b	0	1985	-5.474a	0	1981	-5.825a	0
<b>DLCPI</b>	1986	-3.304	4	1984	-4.925b	2	1986	-6.354a	3

- a Significant at the 1-% level  
b Significant at the 5-% level  
c Significant at the 10-% level

**Table 12: Engle-Granger Cointegration Tests**

Series	(C)		(T)	
	t	k	t	k
<b>LGDP/LM1</b>	-1.864	3	-2.403	3
<b>LDGDP/LM1</b>	-0.706	3	-2.539	3
<b>LCPI/LM1</b>	-1.047	3	-2.504	3
<b>LGDP/LM2</b>	0.509	0	-1.043	0
<b>LDGDP/LM2</b>	0.109	2	-1.546	0
<b>LCPI/LM2</b>	-0.081	1	-1.902	1

- a Significant at the 1-% level  
b Significant at the 5-% level  
c Significant at the 10-% level

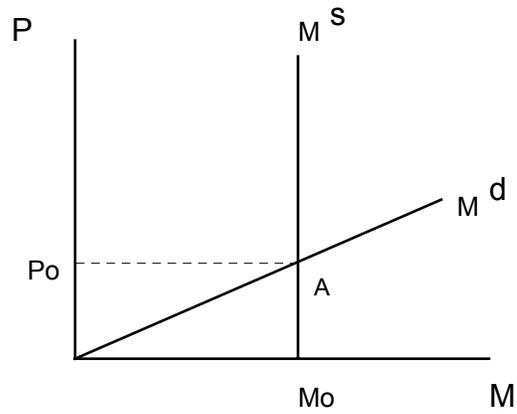
**Table 13: Gregory & Hansen (1996) Cointegration Tests with One Structural Break**

Series	Model (C)			Model (C/T)			Model (C/S)		
	TB	t	k	TB	t	k	TB	t	k
<b>LGDP/LM1</b>	1973	-4.636b	0	1973	-4.904c	0	1973	-4.806c	0
<b>LDGDP/LM1</b>	1990	-2.984	0	1973	-4.375	0	1976	-4.751c	0
<b>LCPI/LM1</b>	1972	-2.824	0	1973	-4.221	0	1974	-4.832c	0
<b>LGDP/LM2</b>	1990	-3.499	1	1975	-3.621	0	1981	-4.471	0
<b>LDGDP/LM2</b>	1990	-2.589	0	1974	-3.959	1	1979	-2.884	0
<b>LCPI/LM2</b>	1993	-2.981	4	1972	-3.762	1	1979	-4.28	1

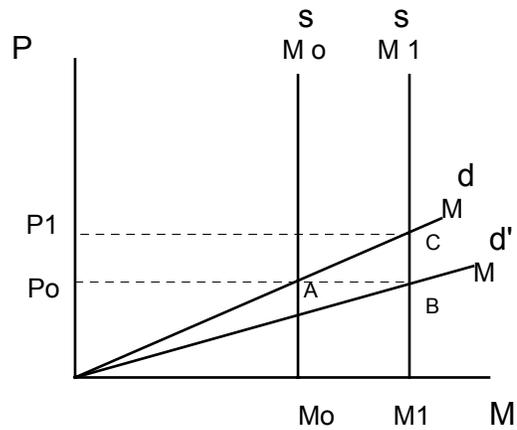
- 
- a Significant at the 1-% level
  - b Significant at the 5-% level
  - c Significant at the 10-% level

**FIGURES:**

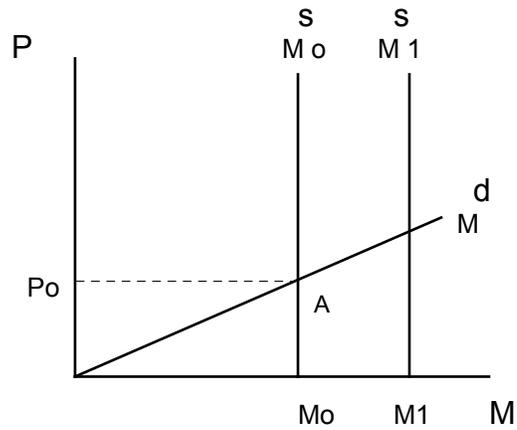
**Figure 1: Money Market Equilibrium**



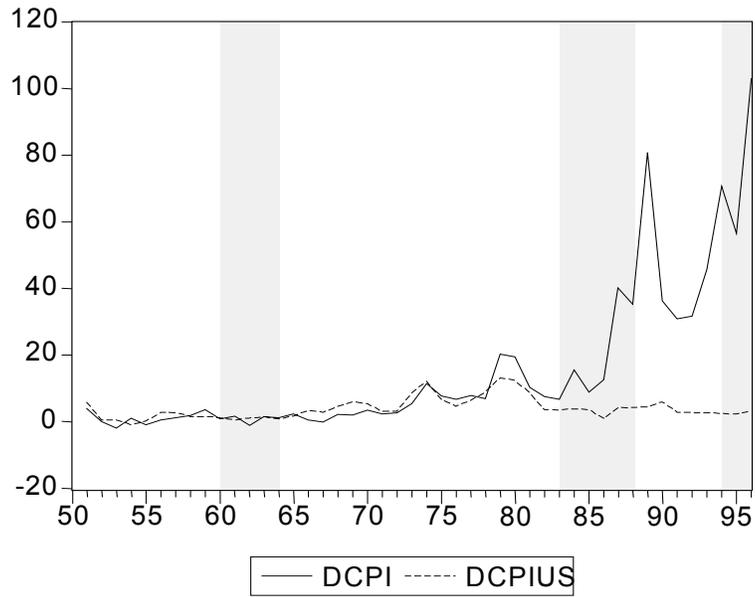
**Figure 2: Increase in Money Supply: Floating Exchange Rates**



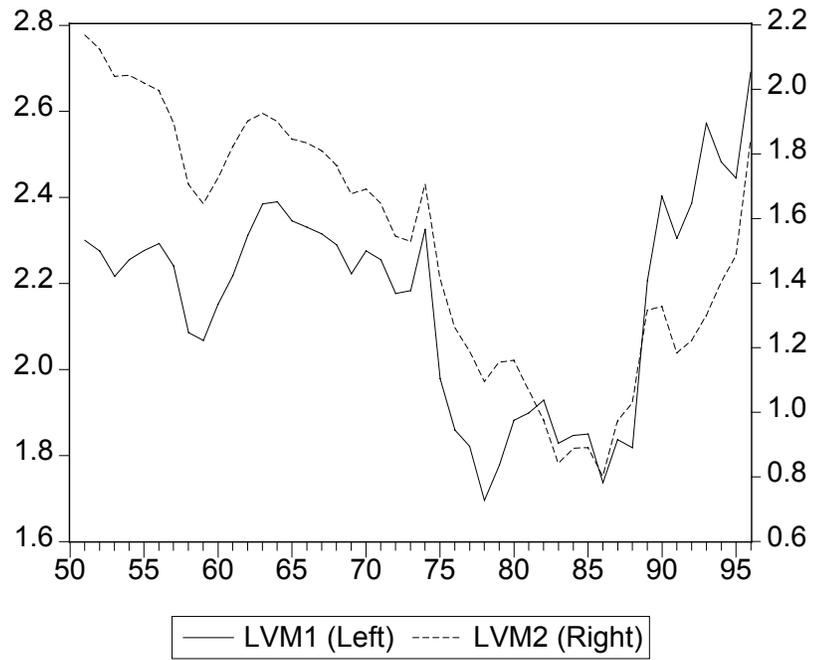
**Figure 3: Increase in Money Supply: Fixed Exchange Rates**



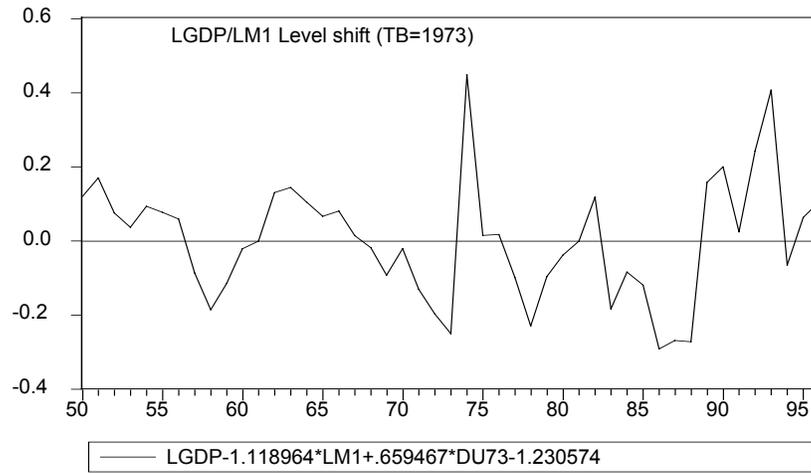
**Figure 4: Inflation and Money Growth: United States and Venezuela**



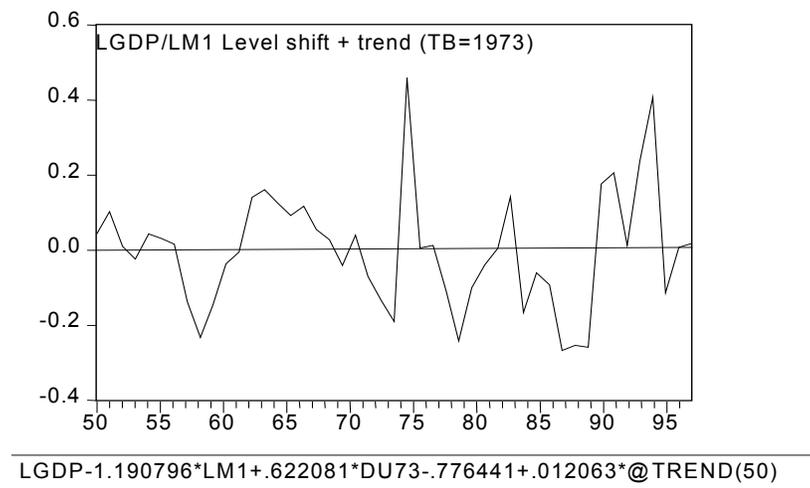
**Figure 5: M1 and M2 Velocities of Circulation for Venezuela**



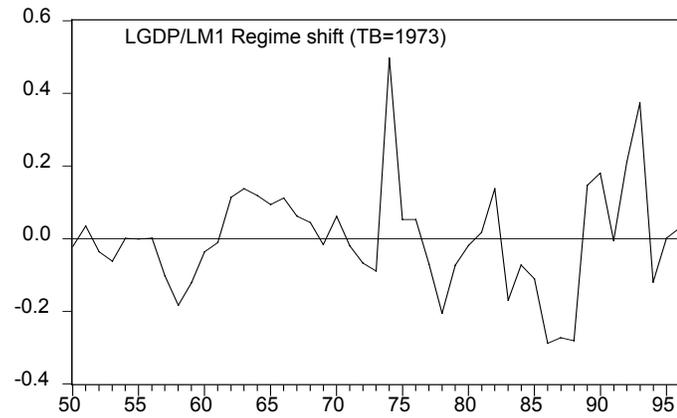
**Figure 6: Residuals from GDP and M1 Cointegration with Level Shift**



**Figure 7: Residuals from GDP and M1 Cointegration with Level Shift and Trend**

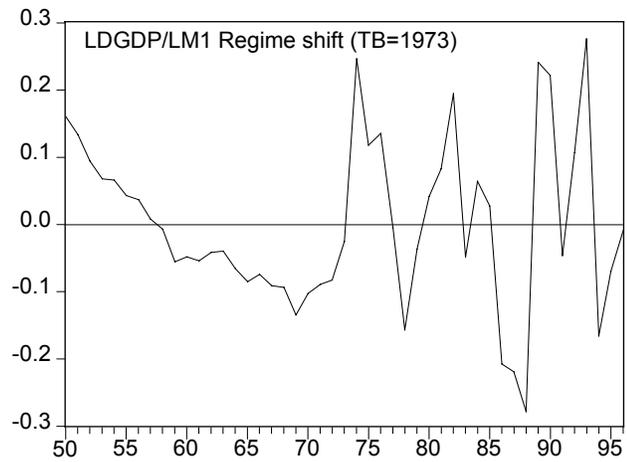


**Figure 8: Residuals from GDP and M1 Cointegration with Regime Shift**



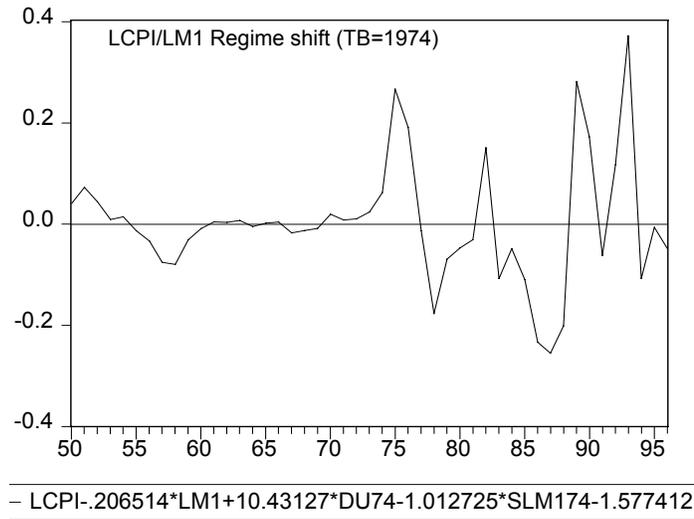
$$- LGDP-.963904*LM1+2.228525*DU73-.179829*SLM173-2.510345$$

**Figure 9: Residuals from GDP Deflator and M1 with Regime Shift**



$$- LDGDP-.278077*LM1+8.596996*DU76-.838627*SLM176-.700059$$

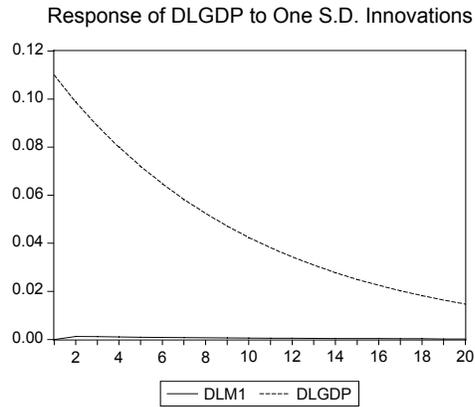
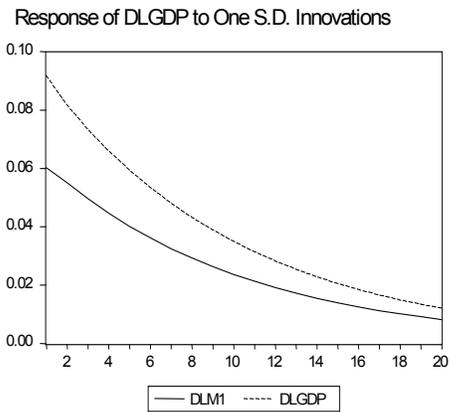
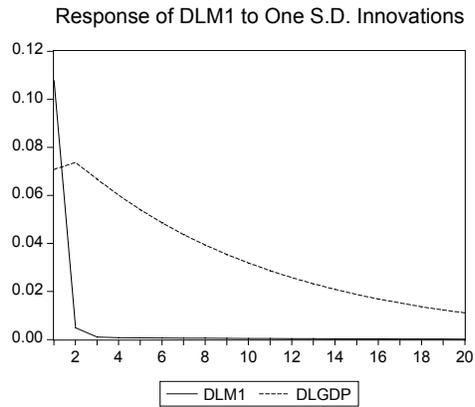
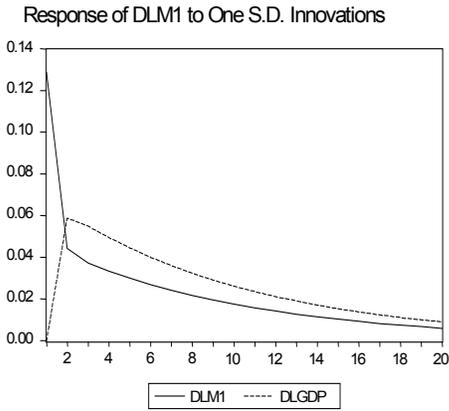
**Figure 10: Residuals from CPI and M1 Cointegration with Regime Shift**



**Figure 11: Impulse-Response Functions for GDP and M1**

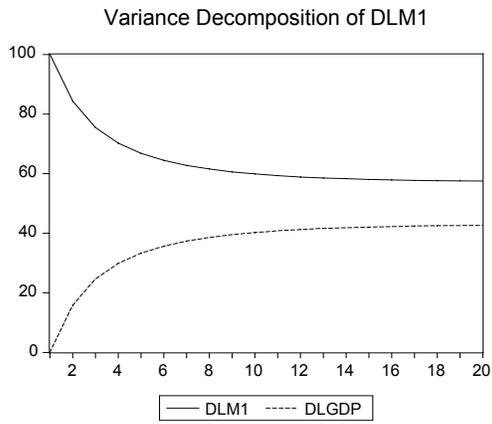
**Ordering: DLM1- DLGDP**

**Ordering: DLGDP - DLM1**

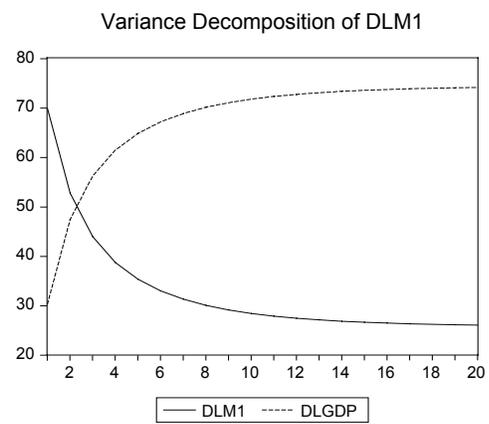


**Figure 12: Variance Decompositions for GPD and M1**

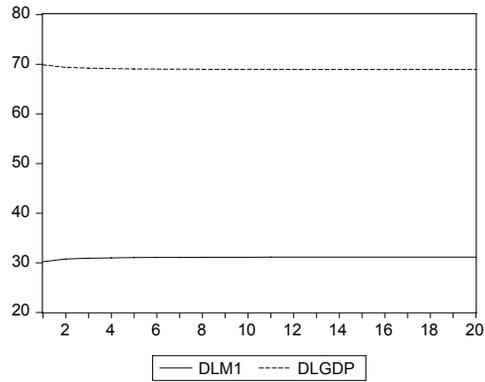
**Ordering: DLM1- DLGDP**



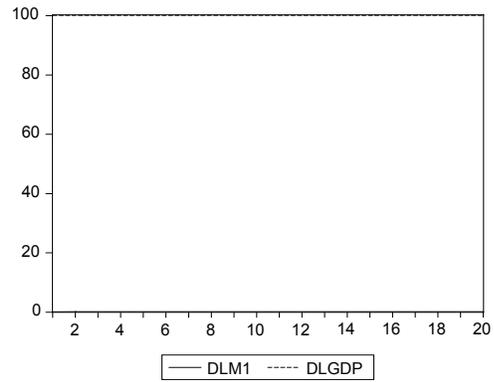
**Ordering: DLGDP - DLM1**



Variance Decomposition of DLGDP

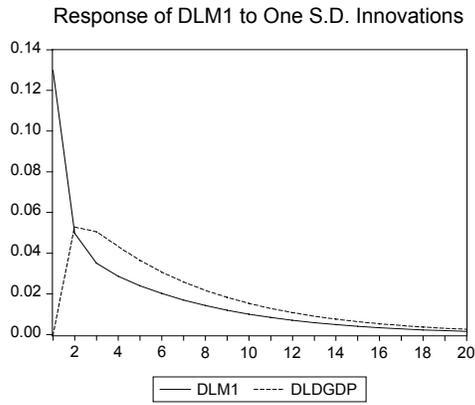


Variance Decomposition of DLGDP

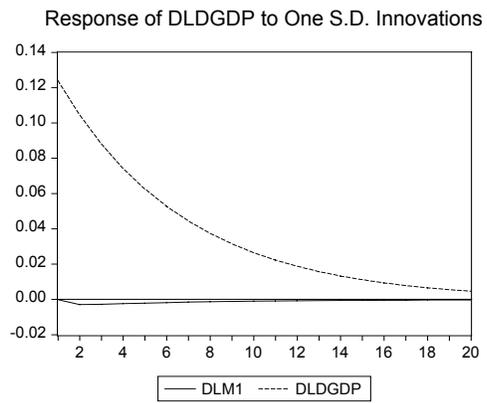
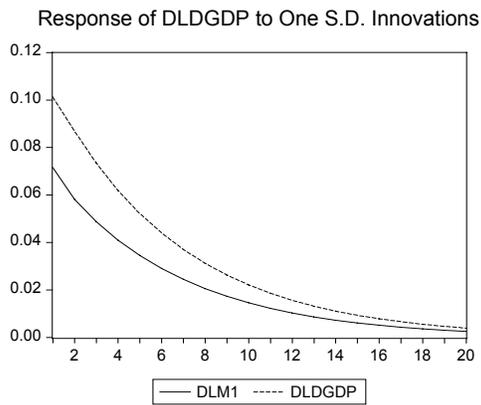
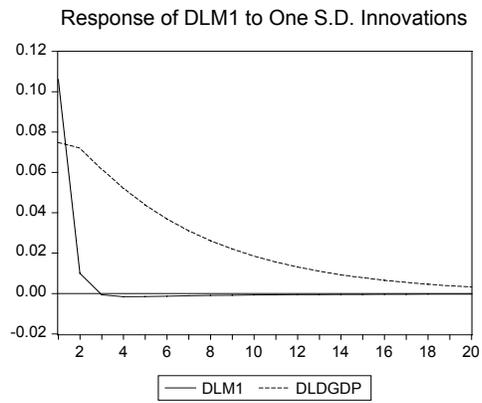


**Figure 13: Impulse-Response Functions for GDP Deflator and M1**

**Ordering: DLM1-DLDGDP**

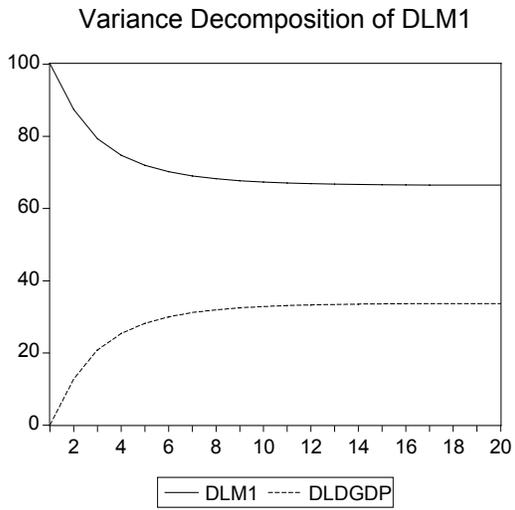


**Ordering: DLDGDP-DLM1**

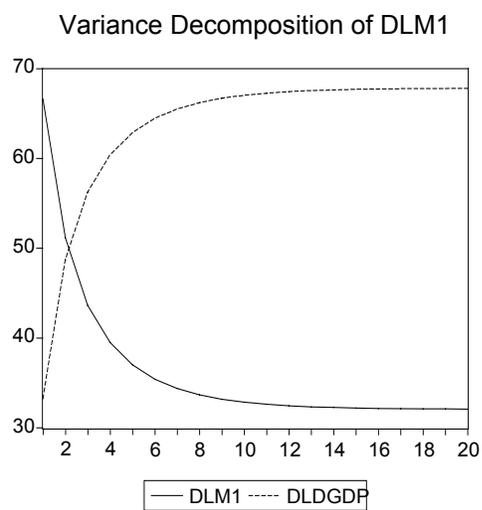


**Figure 14: Variance Decompositions for GDP Deflator and M1**

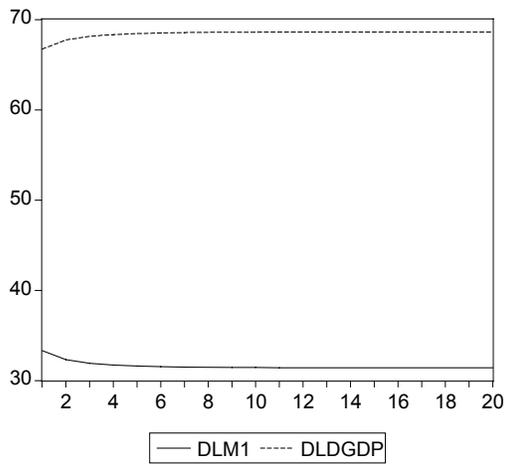
**Ordering: DLM1-DLDGDP**



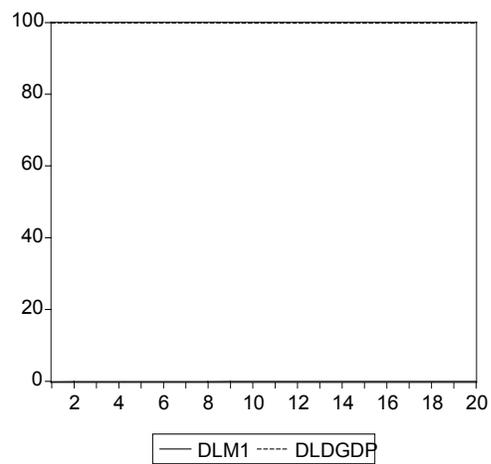
**Ordering: DLDGDP-DLM1**



**Variance Decomposition of DLDGDP**

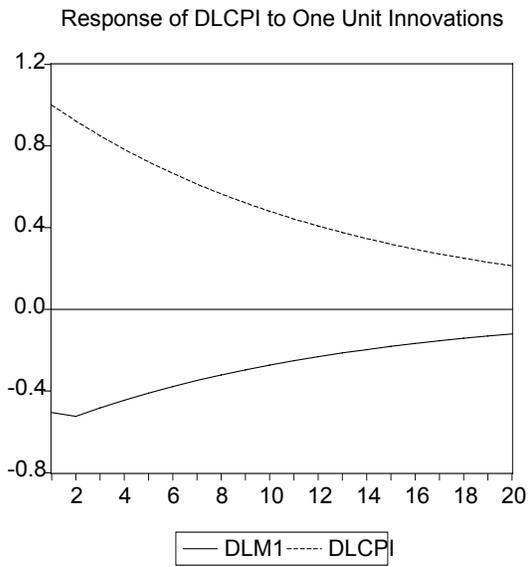


**Variance Decomposition of DLDGDP**

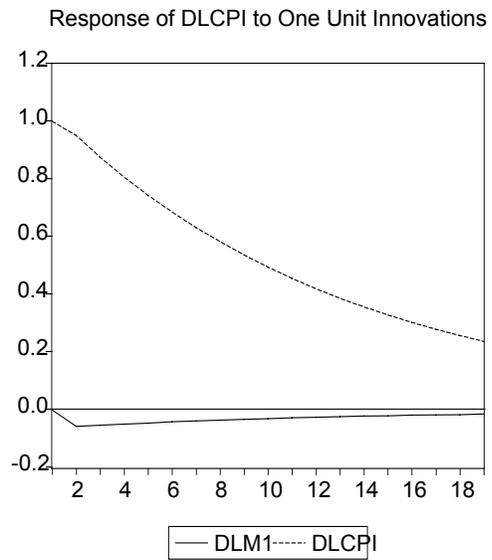


**Figure 15: Impulse-Response Functions for CPI**

**Ordering: DLM1- DLCPI**

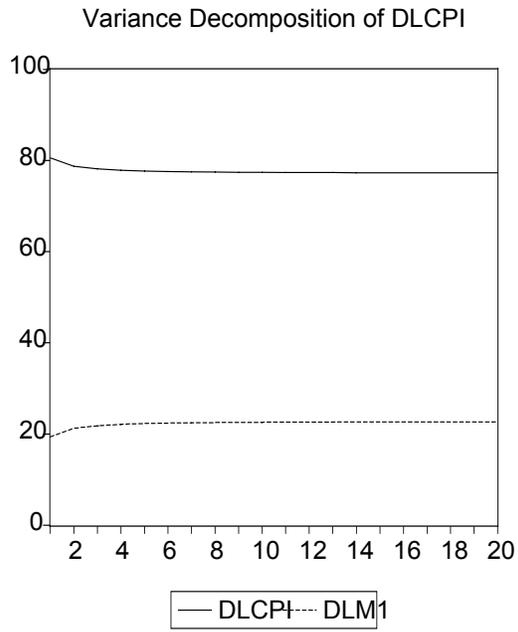


**Ordering: DLCPI - DLM1**

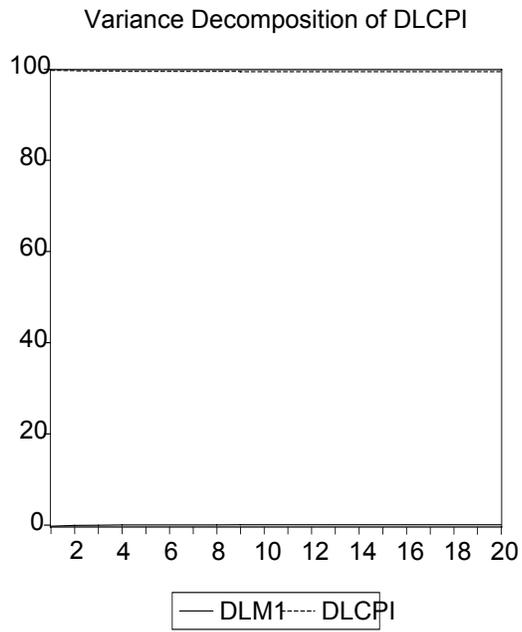


**Figure 16: Variance Decompositions for CPI**

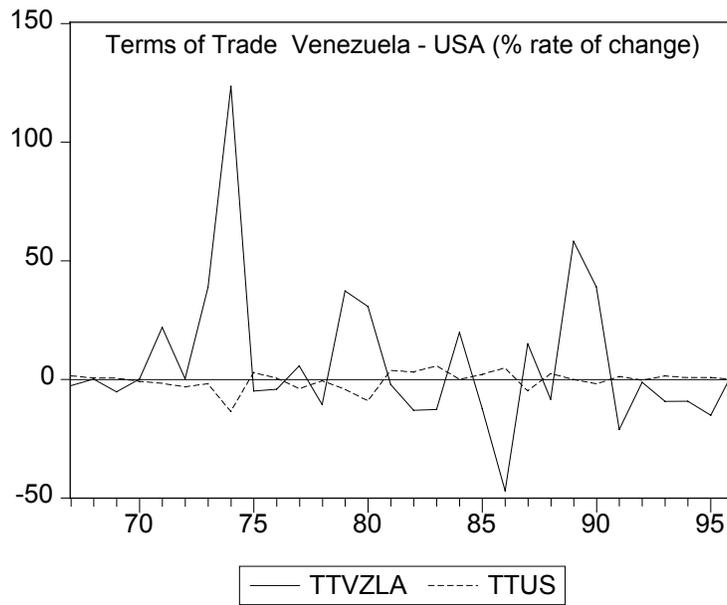
**Ordering: DLM1- DLCPI**



**Ordering: DLCPI - DLM1**



**Figure 17: Terms of Trade in US and Venezuela**



Source: IMF International Financial Statistics Yearbook 1997