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COINTEGRATION AND TESTS OF A CLASSICAL MODEL OFINFLATION IN ARGENTINA, BOLIVIA, BRAZIL, MEXICO, AND PERU

by

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Abstract

The failure of the "heterodox" inflation stabilization attempts of the 1980s in Argentina, Brazil and Peru along with the apparent success of the more orthodox Bolivian and Mexican stabilizations have left researchers again looking at the dynamics of inflation. This paper seeks to add to recent findings by testing whether the recent seemingly different inflation experiences in Argentina, Bolivia, Brazil, Mexico and Peru are consistent with the new classical model of inflation. Monetary models of inflation with rational expectations carry a number of testable implications. First, money growth and inflation should be cointegrated while the short term dynamics display temporary and stochastic dislocations from this equilibrium relationship. Second, the equilibrium error anticipates future monetary policy due the fact that agents have superior information to that of the econometrician. Third, cointegration between money growth and inflation implies, as Campbell and Shiller (1987 and 1988) show, that cross equation restrictions can be readily generated from the error correction form. Our results show that the new classical model of inflation is generally consistent with the inflation experiences of Argentina, Bolivia, Brazil, Mexico, and Peru in spite of their supposed heterogeneity.

Introduction

The inflationary process in Latin America has received a large amount of attention over the last decade especially after the onset of the debt crisis in 1982. In the aftermath of the initial orthodox adjustment and acceleration of inflation, a strong revisionist version of the old monetarist-structuralist debate emerged.[1] The older monetarist inflation theories have been supplanted by rational expectations models of inflation while structuralist theories grew into new-structuralist or "inertial" inflation theories. The failure of the "heterodox" inflation stabilization attempts of the 1980s in Argentina, Brazil, and Peru along with the apparent success of the more orthodox Bolivian and Mexican stabilizations have left researchers again looking at the dynamics of inflation. This paper seeks to add to recent findings by testing whether the recent inflation experiences in Argentina, Bolivia, Brazil, Mexico and Peru are consistent with the new classical model of inflation.

The inflation experiences of these countries differed in the 1980s (Figures 1 - 5).[2] Four of the countries attempted "heterodox" stabilization policies combining incomes policies with monetary and fiscal austerity: Argentina in 1985-1987, 1988, and 1989, Brazil in 1986, 1987, 1989, and 1990, Mexico in the period 1988 to the present, and Peru in 1986-1987. Only

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[1]See the articles in Baer and Welch (1987) for discussions of the initial adjustment which ignited the revised monetarist-structuralist debate. Sargent (1986) succinctly lays out the new classical view of inflation. The works contained in Baer and Kerstenetzky (1964) present a good summary of the old debate.

[2]Good reviews on the recent inflation experiences in each of these countries include Bruno, *et al*, (1991), Pastor (1989), Sachs and Morales (1988), Paredes and Sachs (1991), and Dornbusch and Edwards (1991).

the Mexican program proved a long term success. Four of the countries's inflation rates reached hyperinflationary levels: Argentina in June and July 1989 and December 1989, Bolivia in early 1985, Brazil in 1990, and Peru in 1990-1991. Only two of these countries had brought their inflation rates back to moderate levels by 1990, Bolivia and Mexico. The sample of countries used in this study offers a rich diversity of high inflation experiences in 1980s. The aim of this paper is to see if the classical model successfully describes the inflationary process across these experiences.

Monetary models of inflation with rational expectations carry a number of testable implications. First, the main tenet of monetary models is that inflation is (ultimately) a monetary phenomenon. This precludes the existence of speculative sources of inflation. In the context of rational expectations, speculative bubbles can theoretically emerge due the inconsistency inherent in models where the present price level is a function of future expected price levels [Diba and Grossman 1988a and 1988b]. Theoretically, inflation can accelerate infinitely even though money grow remains stationary. Such bubbles, however, would have the growth rates of prices and money continuously diverging which precludes cointegration between inflation and money growth. Hence, one can empirically rule out inflationary bubbles if money growth and inflation are cointegrated. We will interpret the non-existence of rational inflation bubbles to mean that the inflationary process is consistent with monetary models in general.

Second, forward looking or rational expectations imply structural restrictions on the monetary model which can be interpreted best in the context of cointegrated models. The solution for the inflation rate in these models resembles the general form for the present value models of Campbell and Shiller (1987 and 1988). Specifically, the models imply that money growth and inflation are cointegrated in the long run while the short term dynamics display temporary and stochastic dislocations from this equilibrium relationship. These "disequilibria", however, in turn do not imply that the present value model is not valid. On the contrary, the equilibrium error can be seen to anticipate future monetary policy due the fact that agents have superior information than does the econometrician. Causality running from the equilibrium error to money growth and inflation does not imply that the error causes changes in the variables of the model but instead anticipates them [Campbell and Shiller 1988: 506-507].

Cointegration between money growth and inflation, given the above interpretation, suggests that the appropriate framework should be an error correction representation for the two joint processes. As Campbell and Shiller (1987 and 1988) show, the cross equation restrictions can be readily generated from the error correction form. Unfortunately, rejection of these cross equation restrictions, however, does not lead to any clear interpretation of the underlying inflation-money growth dynamic. The ultimate goal of the analysis is to gauge the ability of the model to describe the inflationary process in each of these countries.

The paper is organized as follows. Section I presents a model of inflation in line the with classical theory and discusses the stationarity properties of money growth and inflation. Cross equation restrictions are developed in section II. The empirical results of the model are shown in section III while section IV summarizes and concludes the paper.

I. A Classical Model of Inflation

The model starts with the money demand specification of Cagan (1956).

$$m_t - p_t = y_t - \alpha i_t + \epsilon_t \tag{1}$$

where m_t is the natural logarithm of the money stock, p_t is the natural logarithm of the price level, y_t is the natural logarithm of real output, i_t is the nominal interest rate, and ϵ_t is a zero mean random error term all evaluated at time t.[3] The standard assumption describes ϵ_t as a random walk of the form

$$\epsilon_t = \epsilon_{t-1} + \eta_t \tag{2}$$

where η_t is white noise.

The classical model assumes a Fisher relationship for the nominal interest rate.

$$i_t = r_t + E[\pi_{t+1} | \Phi_{t-k+1}] \tag{3}$$

where r_t is the real interest rate, E[·] is the expectations operator, $\pi_{t+1} = p_{t+1} - p_t$ is the logarithmic inflation rate, and Φ_t is the information set at time t. The model subsumes rational expectations, i.e. individuals use all information available to them to form expectations about future inflation rates.

[3] This error term can be viewed as one which is either viewed by market participants or constructed by them. ϵ_t, however, is not observed by the researcher. See Diba and Grossman (1988a) and Campbell and Shiller (1987 and 1988).

Real output and real interest rates are assumed to follow random walks (real output also has a drift).

$$y_t - y_{t-1} = \bar{y} + \omega_{1t} \tag{4}$$

$$r_t - r_{t-1} = \omega_{2t} \tag{5}$$

where ω_{1t} and ω_{2t} are white noise.

Taking first differences on equation (1) and combined with equations (2) to (5) yields the following expression.

$$\mu_t - \pi_t = \bar{y} - \alpha\left(E[\pi_{t+1}|\Phi_{t-k+1}] - E[\pi_t|\Phi_{t-k}]\right) + \xi_t \tag{6}$$

where μ_t is the logarithmic growth of money and

$$\xi_t = \eta_t + \omega_{1t} - \alpha\omega_{2t} \tag{7}$$

is white noise.

Rearranging equation (6) yields

$$\pi_t = \mu - \bar{y} + \alpha\left(E[\pi_{t+1}|\Phi_{t-k+1}] - E[\pi_t|\Phi_{t-k}]\right) - \xi_t \tag{8}$$

Taking expectations on equation (8) conditional on Φ_{t-k} and solving forward n periods into the future into equation (9) yields

$$E[\pi_t|\Phi_{t-k}] = \frac{1}{1+\alpha}\left[-\sum_{i=0}^{n-1}\left(\frac{\alpha}{1+\alpha}\right)^i \tilde{y} + \sum_{i=0}^{n-1}\left(\frac{\alpha}{1+\alpha}\right)^i \left(E[\mu_{t+i}|\Phi_{t-k}]\right)\right] + \left(\frac{\alpha}{1+\alpha}\right)^n E[\pi_{t+n}|\Phi_{t-k}] \tag{9}$$

For the evolution of inflation expectations (and thus inflation) to be stable (no bubbles), they must satisfy the following transversality condition

$$\lim_{n\to\infty}\left(\frac{\alpha}{1+\alpha}\right)^n E[\pi_{t+n}|\Phi_{t-k}] = 0 \tag{10}$$

If equation (10) is satisfied, the no bubbles solution to the inflation rate is

$$\pi_t = \mu_t - \tilde{y} + \frac{\alpha}{1+\alpha}\sum_{i=0}^{\infty}\left(\frac{\alpha}{1+\alpha}\right)^i \left(E[\mu_{t+i+1}|\Phi_{t-k+1}] - E[\mu_{t+i}|\Phi_{t-k}]\right) - \xi_t \tag{11}$$

On the other hand, if the transversality condition is violated, a rational bubble can exist. For the bubble to be consistent with expectations, it must evolve in the following way

$$E[B_{t+1}|\Phi_{t-k}] - \left(\frac{1+\alpha}{\alpha}\right)B_t = 0 \tag{12}$$

Solutions to (14) satisfy the stochastic difference equation

$$B_{t+1} - \left(\frac{1+\alpha}{\alpha}\right)B_t = \zeta_{t+1} \tag{13}$$

where the random variable ζ_t satisfies

$$E[\zeta_t | \Phi_{t-k}] = 0 \quad \forall\, k \geq 0 \tag{14}$$

The solution of inflation rate with a bubble is[4]

$$\pi_t = \mu_t - \bar{y} + \frac{\alpha}{1+\alpha}\sum_{i=0}^{\infty}\left(\frac{\alpha}{1+\alpha}\right)^i \left(E[\mu_{t+i+1} | \Phi_{t-k+1}] - E[\mu_{t+i} | \Phi_{t-k}]\right) + B_t - \xi_t \tag{15}$$

The presence of bubbles carries a number of implications [Diba and Grossman 1988a: 522-523]. The first is that the presence of bubbles precludes the stationarity of any degree of differencing of the inflation series. Taking first differences of the bubble in equation (16) using the lag operator L yields[5]

$$\left[1 - \left(\frac{1+\alpha}{\alpha}\right)L\right](1 - L)B_t = (1 - L)\zeta_t \tag{16}$$

One could continue differencing this representation of the bubble. The ARMA representation of equation (20), however, will never be stationary (as the root of [1-$((1+\alpha)/\alpha)$z] = 0 lies inside the unit circle) nor invertible. The bubble introduces a non-

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[4]To see this note that

$$E[B_{t+1} | \Phi_{t-k}] - B_t = \frac{1}{\alpha}B_t$$

Substituting this value into equation (8) yields the additive term B_t.

[5]The following discussion follows Diba and Grossman's (1988a and 1988b).

stationarity which cannot be differenced away.

The presence of bubbles would also rule out cointegration between inflation and money growth. Reconsider equation (19) which /αates

$$\pi_t = \mu_t - \tilde{y} + \frac{\alpha}{1+\alpha}\sum_{i=0}^{\infty}\left(\frac{\alpha}{1+\alpha}\right)^i \left(E[\mu_{t+i+1}|\Phi_{t-k+1}] - E[\mu_{t+i}|\Phi_{t-k}]\right) + B_t - \xi_t \quad (17)$$

Suppose both inflation and money growth are stationary after first differencing (i.e. integrated of order 1 or I(1)) and recall that the growth rate of real output is assumed to be constant. In this classical representation, the left hand side of equation (21) is an equilibrium relationship of inflation and money growth with cointegrating vector $\alpha' = [1, -1]$ and an intercept while the right hand represents the residuals Z_t. If there are no bubbles, the residuals are stationary and inflation and money growth are cointegrated of order (1,1). In the presence of bubbles, however, the residuals of the cointegrating regression are not stationary. Hence, if inflation and money growth are cointegrated, no bubbles exist. Further, cointegration of money growth and inflation rules out any non-stationarity of the unobserved variables [Diba and Grossman 1988a: 525-526]. Hamilton and Whiteman (1985) come to similar conclusions by showing that if money growth is stationary after d differences and inflation is stationary after differencing d times, then speculative inflationary bubbles cannot exist.

II Cross Equation Restrictions

The new classical view of inflation posits that inflation rates are functions of current and expected future money growth rates. The form of these relationships generate a set of easily testable restrictions on the inflationary process. The inflation generation process of the classical model without bubbles followed

$$\pi_t = \mu_t - \bar{y} + \frac{\alpha}{1 + \alpha}\sum_{i=0}^{\infty}\left(\frac{\alpha}{1 + \alpha}\right)^i\left(E[\mu_{t+i+1}|\Phi_{t-k+1}] - E[\mu_{t+i}|\Phi_{t-k}]\right) - \xi_t \tag{18}$$

The task now is to derive an error correction form of the monetary growth process in order to generate forecasts of μ_{t+i} and then test the restrictions implied by equation (1).

Suppose inflation and money growth are both I(1) and cointegrated CI(1,1). The trick now is to generate an error correction representation of the inflationary process. Let the time series vector $X_t = [\pi_t, \mu_t]$. By the Wold decomposition theorem, X_t can be represented

$$(1 - L)X_t = C(L)v_t \tag{19}$$

where C(L) is a 2 x 2 matrix in the lag operator and v_t is a vector white noise process with $v_t = [v_{1,t}, v_{2,t}]$.

Engle and Granger (1987) show that the corresponding ARMA representation of the MA process of equation (2) will not be invertible and that an error correction form would be the appropriate model choice. To see this, multiply both sides of equation (2) by the cointegrating vector [1, -1] to get

$$(1-L)Z_t = \alpha'(1 - L)X_t = \alpha' C(L)\nu_t \tag{20}$$

For Z_t to be stationary, i.e. I(0),

$$\alpha' C(1) = \vec{0} \tag{21}$$

where 0 is a 1 x 2 vector of zeros. Hence, $C(L) = C(1) + (1-L)C^*(L)$ cannot be simply inverted to form an AR representation of X_t. Granger and Engle (1987) show that the CI(1,1) process of equation (2) will have an error correction representation[6]

$$(1 - L)X_t = A^*(L)(1 - L)X_t - \lambda Z_{t-1} + b(L)\nu_t \tag{22}$$

where $A^*(0) = 0$, λ is a vector of constants, λ is a (2 x 1) vector of constants, $\det[C(L)] = (1-L)b(L)]$, and b(L) is a scaler lag polynomial. As b(L) is invertible, premultiplying equation (5) by $b^{-1}(L)$ yields

$$D(L)(1 - L)X_t = -g(L)\lambda Z_{t-1} + \nu_t \tag{23}$$

where $D(L) = b^{-1}(L)[I-A^*(L)] = b^{-1}(L)A(L)$ and $g(L) = b^{-1}(L)$. Define the matrix M as

[6]The Granger Representation Theorem [Engle and Granger 1987: 255-256]. These results follow from factoring the adjoint matrix of C(L).

$$M = \begin{bmatrix} 0 & 1 \\ 1 & -1 \end{bmatrix} \tag{24}$$

Now

$$MX_t = \begin{bmatrix} 0 & 1 \\ 1 & -1 \end{bmatrix} \begin{bmatrix} \pi_t \\ \mu_t \end{bmatrix} = \begin{bmatrix} \mu_t \\ Z_t \end{bmatrix} \tag{25}$$

and

$$X_t = M^{-1} \begin{bmatrix} \mu_t \\ Z_t \end{bmatrix} = \begin{bmatrix} 1 & 1 \\ 1 & 0 \end{bmatrix} \begin{bmatrix} \mu_t \\ Z_t \end{bmatrix} \tag{26}$$

Substituting equation (30) into equation (27) yields

$$D(L)(1 - L)M^{-1} \begin{bmatrix} \mu_t \\ Z_t \end{bmatrix} = -g(L)\lambda Z_{t-1} + \nu_t \tag{27}$$

Since (1-L) is a scaler in L, equation (10) can be rearranged in the following way[7]

[7]See Campbell and Shiller (1988), p. 510-511. The intuition behind this reformulation lies in the fact that Z_t is stationary.

$$H(L)\begin{bmatrix}(1 - L)\mu_t \\ Z_t\end{bmatrix} = \omega_t \tag{28}$$

In order to generate optimal forecasts of money growth, we rewrite the VAR representation of equation (11) in the following way

$$Y_t = \Theta Y_{t-1} + e_t \tag{29}$$

where

$$Y_t = \begin{bmatrix}(1-L)\mu_t \\ (1-L)u_{t-1} \\ . \\ . \\ . \\ . \\ (1-L)\mu_{t-p-1} \\ (1-L)\mu_{t-p} \\ Z_t \\ Z_{t-1} \\ . \\ . \\ . \\ Z_{t-p-1} \\ Z_{t-p}\end{bmatrix} \quad e_t = \begin{bmatrix}\omega_{1t} \\ 0 \\ . \\ . \\ . \\ 0 \\ \omega_{2t} \\ 0 \\ . \\ . \\ . \\ 0\end{bmatrix} \tag{30}$$

and Θ is the companion matrix of the VAR of the form

$$\Theta = \begin{bmatrix} \theta_{111} & \theta_{112} & \cdots & \theta_{11p-2} & \theta_{11p} & \theta_{121} & \theta_{122} & \cdots & \theta_{12p-1} & \theta_{12p} \\ 1 & 0 & \ldots & 0 & 0 & 0 & 0 & \ldots & 0 & 0 \\ 0 & 1 & \ldots & 0 & 0 & 0 & 0 & \ldots & 0 & 0 \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ 0 & 0 & \ldots & 1 & 0 & 0 & 0 & \ldots & 0 & 0 \\ \theta_{211} & \theta_{212} & \cdots & \theta_{21p-1} & \theta_{21p} & \theta_{221} & \theta_{222} & \cdots & \theta_{22p-1} & \theta_{22p} \\ 0 & 0 & \ldots & 0 & 0 & 1 & 0 & \ldots & 0 & 0 \\ 0 & 0 & \ldots & 0 & 0 & 0 & 1 & \ldots & 0 & 0 \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ 0 & 0 & \ldots & 0 & 0 & 0 & 0 & \ldots & 1 & 0 \end{bmatrix} \tag{31}$$

Optimal forecasts of the Y_t will thus be generated by

$$E\left[Y_{t+i} \mid \Phi_{t-k}\right] = \Theta^{k+1+i} Y_{t-k} \tag{32}$$

One important aspect of the VAR of equation (36) is that if the cointegrated present value model holds, Z_t will "Granger-cause" changes in money growth and changes in inflation [Campbell and Shiller 1988: 513]. If economic agents have superior information to that of the econometrician, one would find that the equilibrium error anticipates the changes in inflation and money growth. Hence, below we test for such a causal relationship.

Equation (1) implies a set of restrictions on the optimal forecast equation (15).

Recall the money demand function

$$\pi_t = \mu - \tilde{y} + \alpha\left(E[\pi_{t+1}|\Phi_{t-k+1}] - E[\pi_t|\Phi_{t-k}]\right) - \xi_t \tag{33}$$

Rewriting equation (37) in the new notation yields

$$Z_t = -\tilde{y} + \alpha\left(E[Z_{t+1}|\Phi_{t-k+1}] + E[\mu_{t+1}|\Phi_{t-k+1}] - E[Z_t|\Phi_{t-k}] - E[\mu_t|\Phi_{t-k}]\right) - \xi_t \tag{34}$$

Taking expectations conditional on Φ_{t-k} of equation (38) and rearranging yields

$$E\left[Z_t - \frac{\alpha}{1+\alpha}Z_{t+1} - \frac{\alpha}{1+\alpha}\Delta\mu_{t+1} - \frac{1}{1+\alpha}\tilde{y}\ |\Phi_{t-k}\right] = 0 \tag{35}$$

Let R_1 = [0,0,...,0,1,0,0,...,0,0] and R_2 = [1,0,...,0,0,0,0,...,0,0]. The classical restrictions in equation (18) can be expressed as

$$H_1^o = R_1\Theta^k - \frac{\alpha}{1+\alpha}R_1\Theta^{k+1} - \frac{\alpha}{1+\alpha}R_2\Theta^{k+1} = 0 \tag{36}$$

which are non-linear in the parameter matrix Θ. The Wald statistic for this test is

$$TH_1^{o\prime}\left[\left(\frac{\partial H_1^o}{\partial\Theta}\right)^{\prime}\hat{\Sigma}_{\hat{\Theta}}\left(\frac{\partial H_1^o}{\partial\Theta}\right)\right]^{-1}H_1^o \sim \chi_p^2 \tag{37}$$

where T is the number of observations and Σ_Θ is the estimated covariance matrix of the estimated Θ matrix.

III Empirical Results

a. Cointegration Tests

Before moving to the tests for cointegration, tests on the order of integration are in order.[8] Tables 1 through 10 show the Dickey-Fuller (1979) and Phillips-Perron (1988) tests for stationarity for money and prices in Argentina, Bolivia, Brazil, Mexico, and Peru. The Phillips and Perron (1988) tests correct for non-normality of variables (tested for using the Jarque and Bera (1980) tests). In Argentina, Bolivia, Brazil, Mexico, and Peru, M_2 money growth and inflation are strongly stationary after differencing. In all countries, inflation is not unambiguously I(1) as opposed to I(0). The cointegration tests below, however, indicate that it is I(1) as is money growth.[9]

Generally, cointegration means that (non-stationary) time series variables tend to move together such that a linear combination of them is stationary. As in the analysis above, some have interpreted cointegration as representing a long run equilibrium

[8]Inflation in both countries is measured by the wholesale price index and M_2 was picked as the monetary aggregate based upon cointegration tests on the money demand equation (1). The conclusions of the tests, however, do not depend on the choice of money aggregate. The Argentine data are quarterly observations from 1970 to 1984 and come from INDEC. The Bolivian data are monthly observations from June 1980 to September 1990 from the Banco Central de Bolivia. The Brazilian data are monthly observations from 1974 to 1985 and come from the Fundação Getúlio Vargas. The Mexican data are monthly observations from January 1972 to September 1989 from the Banco de Mexico Indicadores Economicos and from the data bank of Sie-Mexico. Finally, Peru's data are monthly observations from January 1983 to June 1990 from the Banco Central de Peru.

[9]M_2 was used as the money aggregate in all countries except Mexico. We used M_1 for Mexico but the results hold for M_2 as well.

relationship. Differencing X_t d times to generate a stationary time series and then estimating a VAR based upon the differenced series is inappropriate in the presence of cointegration. Recall that if a (px1) vector time series X_t (p=2 in this case) is first difference stationary, i.e. I(1), and cointegrated, i.e. b=1, there exists an error correction form

$$\Delta X_t = A_1 \Delta X_{t-1} + \dots + A_{k-1} \Delta X_{t-k+1} + \Pi X_{t-1} + \varepsilon_t \tag{38}$$

where $\Pi = \alpha \beta'$, $\beta' = [\beta_\pi, \beta_\mu]$ is the cointegrating vector, $\alpha' = [\alpha_\pi, \alpha_\mu]$ is the error correction. coefficient (or speed of adjustment). An important aspect of this theorem is that the VAR should incorporate the long run equilibrium relationship between the levels. A VAR based purely upon differences would exclude this relevant information in addition to displaying infinite variance.

In general, there can exist (p-1) independent cointegrating vectors. A weakness in the Engle and Granger (1987) approach is that it offers no clear criterion for choosing the number of cointegrating vectors. Johansen and Juselius (1990) take a general maximum likelihood approach to choosing the number of independent cointegrating vectors, estimating Π, α, β', and testing restrictions on α and β. Their technique is based upon the following general version of equation (1).[10]

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[10]The Π matrix is the same in equation (42) and equation (43). It can be shown that the level variable can take on any lag from 1 to k without affecting Π. The coefficients on the lagged differenced variables, of course, change.

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + \Phi D_t + \varepsilon_t \qquad (39)$$

where D_t is a set of seasonal dummies which sum to zero.

The analysis of the negative of the growth in real money balances looks at the behavior of $ß' = [1, -1]$ of the vector time series $X_t = [\pi_t, \mu_t]$. The maximum likelihood estimates for the cointegrating vector ß' can be obtained from the following eigenvalue problem.[11]

$$|\lambda S_{kk} - S_{ko} S_{oo}^{-1} S_{ok}| = 0 \qquad (40)$$

where S_{ij} are the residual moment matrices from the OLS regressions of ΔX_t and X_{t-j}, ΔX_{t-j}, $j = 1,\ldots,k-1$. The estimates of ß' are just the corresponding eigenvectors while the (maximum) eigenvalues along with the trace (computed from the eigenvalues) are used as test statistics for the rank of Π. Notice that if $\mathrm{Rank}(\Pi) = r = p$, any vector is a cointegrating vector and hence the original vector times series X_t is stationary. Hence, if inflation and money growth are I(0), then we should find two cointegrating vectors. If $\mathrm{Rank}(\Pi) = r < p$, then the data are I(1) and we have r cointegrating vectors. If $\mathrm{Rank}(\Pi) = r = 0$, then we find no cointegrating vectors and a VAR based purely on the first difference of X_t is appropriate. The critical values and sizes of the test statistics appear in the appendix of Johansen and Juselius (1990).

[11]The Johansen and Juselius (1990) procedure assumes normality. The equations are estimated using RATS 3.10 software.

The estimated ß' can then be substituted into equation (43) to derive estimates of α. One can also impose restrictions on Π in the form of individual vectors ß' and α. In this case, we are interested in testing whether ß' = [1, -1]. The likelihood ratio test is distributed as a $\chi^2_{(1)}$.

The results of the rank tests appear for Argentina in tables 11a, for Bolivia in 12a, for Brazil in tables 13a, for Mexico in 14a, and for Peru in 15a. In all cases, the trace and maximum eigenvalue tests indicate that the Π matrix is rank=1, i.e. r=1, at the 5% significance level. In other words, there is only one cointegrating vector for inflation and money growth. This also indicates that the original time series are not stationary as the Π is not full rank, i.e. equal to 2. The significant cointegrating relationships rule out rational inflationary bubbles in each case.

Tests on the cointegrating vector appear for Argentina in table 11b, for Bolivia in 12b, for Brazil in 13b, for Mexico in 14b, and for Peru in 15b. Specifically, we test for long run money neutrality which takes the form of testing whether ß' = [1, -1]. In all countries but Mexico, one cannot reject the neutrality of money. This anomoly probably results from non-normalities in the series. To confirm this, a Phillips-Perron stationarity test on the growth in real balances appears in tables 11b through 15b. In all cases, real balances strongly reject the null hypothesis of non-stationarity.

Tables 16 through 21 present the remaining tests of the present value model. As mentioned above,[12] if the present value model holds, equilibrium errors, Z_t, should

[12]CUSUM tests for structural stability appear in appendix A. Even though structural breaks were not found, dividing the periods studied for each country did not significantly alter the results of the paper. These tests are available upon request.

anticipate changes in money growth. In all countries, Z_t significantly Granger-causes $\Delta\mu_t$ while significant causality in the other direction occurs only in Bolivia, Mexico, and Peru.

The cross equation restrictions, on the other hand, do not conform as readily to the present value model. For all reasonable values of the semi-elasticity of money with respect to the interest rate, α, in Argentina, Mexico and Peru, one does not reject the present value model for an information lag of (k-1=1) 1 quarter. In the Bolivian and Brazilian cases, the model is rejected for all reasonable values of α for information lags of (k-1=) 1 month at less than the 1% significance level while rejected (k-1) = 2 months at the 5% level for Bolivia and at the 10% level for Brazil.

IV Conclusions

The inflation processes in Argentina, Bolivia, Brazil, Mexico, and Peru generally conform to the implications of the new classical model in spite of its simple form. Inflation and money growth are cointegrated in all countries ruling out speculative inflationary bubbles. Agents apparently anticipate future changes in money growth (and, by cointegration, inflation) in line with the rational expectations monetary model. Further, the cross equation restrictions implied by the model are not rejected in Argentina when the information lag is one quarter and Mexico and Peru when the information lag is one month. These restrictions, however, are rejected in Bolivia and Brazil with a one month information lag.

The results show that forward looking expectations do play a part in the inflation process of all countries. Further, "speculative" sources play an insignificant role at least in the medium to long term. Certainly, the model is too simple to explain many other important aspects of inflation in these Latin American countries especially in the Brazilian case. A more detailed structural specification incorporating forward looking variables may improve the performance of the model.

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Table 1
Argentina: Unit Roots Tests[a]

a. Null Hypothesis: Variable has a Unit Root (No Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
Inflation[b]	-2.19	-3.37**
Money Growth (M_2)[b]	-1.77	-2.77

b. Null Hypothesis: Variable has a Unit Root (Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
Inflation[b]	-2.67	-4.05***
Money Growth (M_2)[b]	-2.19	-3.47**

Table 2
Argentina: Unit Roots Tests[a]

a. Null Hypothesis: Variable has a Unit Root (No Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
ΔInflation[b]	-6.81***	-11.55***
ΔMoney Growth (M_2)[b]	-5.76***	-11.91***

b. Null Hypothesis: Variable has a Unit Root (Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
ΔInflation[b]	-6.75***	-11.45***
ΔMoney Growth (M_2)[b]	-5.70***	-11.79***

Notes: (a) Unit root tests on the time series variable y_t are based upon the following regression

$$y_t = \mu + \tau t + \phi^* y_{t-1} + \sum_{i=1}^{q} \phi_i \Delta y_{t-i} \qquad \text{(i)}$$

Dickey-Fuller tests assume normality while Phillips-Perron test make a correction for non-normal time series. The order of the autoreggressive termes, q, was chosen to render the residuals of the regression white noise according to the Box-Pierce Q(22) statistic. The inflation regression used 1 lag while the money growth equation used 1 lag.

(b) Series showed significant non-normality either because of skewness or kurtosis by the Jarque-Bera test.

Table 3
Bolivia: Unit Roots Tests[a]

a. Null Hypothesis: Variable has a Unit Root (No Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
Inflation[b]	-3.28**	-4.57***
Money Growth (M_2)[b]	-2.54	-5.15***

b. Null Hypothesis: Variable has a Unit Root (Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
Inflation[b]	-3.38*	-4.65***
Money Growth (M_2)[b]	-2.60	-5.16**

Table 4
Bolivia: Unit Roots Tests[a]

a. Null Hypothesis: Variable has a Unit Root (No Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
ΔInflation[b]	-12.03***	-14.06***
ΔMoney Growth (M_2)[b]	-4.05***	-23.15***

b. Null Hypothesis: Variable has a Unit Root (Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
ΔInflation[b]	-12.00***	-14.01***
ΔMoney Growth (M_2)[b]	-4.12***	-23.23***

Notes: (a) Unit root tests on the time series variable y_t are based upon the following regression

$$y_t = \mu + \tau t + \phi^* y_{t-1} + \sum_{i=1}^{q} \phi_i \Delta y_{t-i} \qquad \text{(i)}$$

Dickey-Fuller tests assume normality while Phillips-Perron test make a correction for non-normal time series. The order of the autoreggressive termes, q, was chosen to render the residuals of the regression white noise according to the Box-Pierce Q(22) statistic. The inflation regression used 2 lag while the money growth equation used 10 lag.

(b) Series showed significant non-normality either because of skewness or kurtosis by the Jarque-Bera test.

Table 5
Brazil: Unit Roots Tests[a]

a. Null Hypothesis: Variable has a Unit Root (No Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
Inflation[b]	-1.72	-2.84*
Money Growth (M_2)[b]	0.932	-7.20***

b. Null Hypothesis: Variable has a Unit Root (Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
Inflation[b]	-5.41***	-7.72***
Money Growth (M_2)[b]	-1.31	-11.98***

Table 6
Brazil: Unit Roots Tests[a]

a. Null Hypothesis: Variable has a Unit Root (No Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
ΔInflation[b]	-14.88***	-19.97***
ΔMoney Growth (M_2)[b]	-3.82***	-38.81***

b. Null Hypothesis: Variable has a Unit Root (Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
ΔInflation[b]	-14.93***	-19.98***
ΔMoney Growth (M_2)[b]	-4.12***	-40.80***

Notes: (a) Unit root tests on the time series variable y_t are based upon the following regression

$$y_t = \mu + \tau t + \phi^* y_{t-1} + \sum_{i=1}^{q} \phi_i \Delta y_{t-i} \quad \text{(i)}$$

Dickey-Fuller tests assume normality while Phillips-Perron test make a correction for non-normal time series. The order of the autoreggressive termes, q, was chosen to render the residuals of the regression white noise according to the Box-Pierce Q(22) statistic. The inflation regression used 1 lag while the money growth equation used 10 lags.

(b) Series showed significant non-normality either because of skewness or kurtosis by the Jarque-Bera test.

Table 7
Mexico: Unit Roots Tests[a]

a. Null Hypothesis: Variable has a Unit Root (No Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
Inflation[b]	-2.48	-3.87***
Money Growth (M_1)[b]	0.932	-7.20***

b. Null Hypothesis: Variable has a Unit Root (Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
Inflation[b]	-3.58**	-5.90***
Money Growth (M_1)[b]	-1.31	-11.98***

Table 8
Mexico: Unit Roots Tests[a]

a. Null Hypothesis: Variable has a Unit Root (No Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
ΔInflation[b]	-10.57***	-24.57***
ΔMoney Growth (M_1)[b]	-3.82***	-38.81***

b. Null Hypothesis: Variable has a Unit Root (Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
ΔInflation[b]	-10.56***	-24.53***
ΔMoney Growth (M_1)[b]	-4.12***	-40.80***

Notes: (a) Unit root tests on the time series variable y_t are based upon the following regression

$$y_t = \mu + \tau t + \phi^* y_{t-1} + \sum_{i=1}^{q} \phi_i \Delta y_{t-i} \qquad \text{(i)}$$

Dickey-Fuller tests assume normality while Phillips-Perron test make a correction for non-normal time series. The order of the autoreggressive termes, q, was chosen to render the residuals of the regression white noise according to the Box-Pierce Q(22) statistic. The inflation regression used 4 lag while the money growth equation used lags.

(b) Series showed significant non-normality either because of skewness or kurtosis by the Jarque-Bera test.

Table 9
Peru: Unit Roots Tests[a]

a. Null Hypothesis: Variable has a Unit Root (No Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
Inflation[b]	-1.91	-3.88***
Money Growth (M_2)[b]	0.085	-3.44***

b. Null Hypothesis: Variable has a Unit Root (Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
Inflation[b]	-3.10	-5.79***
Money Growth (M_2)[b]	-1.48	-5.96***

Table 10
Peru: Unit Roots Tests[a]

a. Null Hypothesis: Variable has a Unit Root (No Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
ΔInflation[b]	-8.73***	-17.45***
ΔMoney Growth (M_2)[b]	-3.97***	-20.04***

b. Null Hypothesis: Variable has a Unit Root (Time Trend)

Variable	Augmented Dickey-Fuller	Phillips-Perron
ΔInflation[b]	-8.71***	-17.39***
ΔMoney Growth (M_2)[b]	-4.20***	-20.71***

Notes: (a) Unit root tests on the time series variable y_t are based upon the following regression

$$y_t = \mu + \tau t + \phi^* y_{t-1} + \sum_{i=1}^{q} \phi_i \Delta y_{t-i} \qquad \text{(i)}$$

Dickey-Fuller tests assume normality while Phillips-Perron test make a correction for non-normal time series. The order of the autoreggressive termes, q, was chosen to render the residuals of the regression white noise according to the Box-Pierce Q(22) statistic. The inflation regression used 1 lag while the money growth equation used 10 lags.

(b) Series showed significant non-normality either because of skewness or kurtosis by the Jarque-Bera test.

Table 11a

Argentina: Tests for number (r) of Cointegrating Vectors for $X_t = [\mu_t, \pi_t]$ with M_2[a]

TRACE TESTS	H_o:r=0 H_1:r=2	H_o:r=1 H_1:r=2
test statistic	22.61***	3.20
MAXIMUM EIGENVALUE	H_o:r=0 H_1:r=1	H_o:r=1 H_1:r=2
test statistic	25.81***	3.20
UNRESTRICTED ESTIMATES	β_μ	β_π
	1.000	-0.953
	α_μ	α_π
	-0.469	0.271

Table 11b

Argentina: Tests on β' and α for inflation and M_1

H_o:β_π=1, β_μ=-1,
$\chi^2_{(2)}$=0.063

Dickey-Fuller and Phillips-Perron Tests on Growth in Real Balances[a]

H_o: $\beta'X_t$ is non-stationary	Without Trend	With Trend
Augmented Dickey-Fuller	-3.73***	-3.716**
Phillips-Perron[b]	-6.578***	-6.53***

Notes: (a) One lag was used in these tests of stationarity. The lag structure was chosen by adding lags until the Q(12) statistic did not reject the null hypothesis of non-autocorrelated residuals.

(b) Series showed significant non-normality either because of skewness or kurtosis by the Jarque-Bera test.

** signifies rejection of H_o at a 5% significance level, *** signifies rejection of H_o at a 1% significance level.

Table 12a
Bolivia: Test for number (r) of Cointegrating Vectors for $X_t = [\mu_t, \pi_t]$ with M_2[(a)]

TRACE TESTS	H_o:r=0 H_1:r=2	H_o:r=1 H_1:r=2
test statistic	31.39***	2.85
MAXIMUM EIGENVALUE	H_o:r=0 H_1:r=1	H_o:r=1 H_1:r=2
test statistic	28.54***	2.85
UNRESTRICTED ESTIMATES	β_μ	β_π
	1.000	-0.881
	α_μ	α_π
	-0.532	0.598

Table 12b
Bolivia: Tests on ß' and α for inflation and M_2

H_o:β_π=1, β_μ=-1,
$\chi^2_{(2)}$=3.239*

Dickey-Fuller and Phillips-Perron Tests on Growth in Real Balances[(a)]

H_o: ß'X_t is non-stationary	Without Trend	With Trend
Augmented Dickey-Fuller	-7.84***	-3.716**
Phillips-Perron[(b)]	-9.660***	-6.53***

Notes: (a) Two lags were used in these tests of stationarity. The lag structure was chosen by adding lags until the Q(12) statistic did not reject the null hypothesis of non-autocorrelated residuals.

(b) Series showed significant non-normality either because of skewness or kurtosis by the Jarque-Bera test.

** signifies rejection of H_o at a 5% significance level, *** signifies rejection of H_o at a 1% significance level.

Table 13a
Brazil: Tests for number (r) of Cointegrating Vectors for $X_t = [\pi_t, \mu_t]$ with M_2[(a)]

TRACE TESTS	H_0:r=0 H_1:r=2	H_0:r=1 H_1:r=2
test statistic	27.66***	0.42
MAXIMUM EIGENVALUE	H_0:r=0 H_1:r=1	H_0:r=1 H_1:r=2
test statistic	28.07***	0.42
UNRESTRICTED ESTIMATES	β_μ	β_π
	1.000	-0.907
	α_μ	α_π
	-0.800	0.078

Table 13b
Brazil: Tests on β' and α for inflation and M_2

H_0:β_π=1, β_μ=-1,
$\chi^2_{(1)}$=1.405

Dickey-Fuller and Phillips-Perron Tests on the Final $\beta'X_t$[(a)]

H_0: $\beta'X_t$ is non-stationary	Without Trend	With Trend
Augmented Dickey-Fuller	-2.77*	-2.78
Phillips-Perron	-12.28***	-12.23***

Notes: (a) One lag was used in these tests of stationarity. The lag structure was chosen by adding lags until the Q(22) statistic did not reject the null hypothesis of non-autocorrelated residuals.

* signifies rejection of H_0 at a 10% significance level, ** signifies rejection of H_0 at a 5% significance level, *** signifies rejection of H_0 at a 1% significance level.

Table 14a
Mexico: Tests for number (r) of Cointegrating Vectors for $X_t = [\pi_t, \mu_t]$ with M_1[(a)]

TRACE TESTS	H_0:r=0 H_1:r=2	H_0:r=1 H_1:r=2
test statistic	54.75^{***}	5.91
MAXIMUM EIGENVALUE	H_0:r=0 H_1:r=1	H_0:r=1 H_1:r=2
test statistic	48.84^{***}	5.91
UNRESTRICTED ESTIMATES	β_μ	β_π
	1.000	-0.698
	α_μ	α_π
	-0.924	0.154

Table 14b
Mexico: Tests on ß' and α for inflation and M_1

H_0:$\beta_\pi = 1$, $\beta_\mu = -1$,
$\chi^2_{(1)} = 15.972^{***}$

Dickey-Fuller and Phillips-Perron Tests on the Final $\beta' X_t$[(a)]

H_0: $\beta' X_t$ is non-stationary	Without Trend	With Trend
Augmented Dickey-Fuller	-4.97^{***}	-2.78
Phillips-Perron	-18.04^{***}	-12.23^{***}

Notes: (a) Six lags were used in these tests of stationarity. The lag structure was chosen by adding lags until the Q(22) statistic did not reject the null hypothesis of non-autocorrelated residuals.

* signifies rejection of H_0 at a 10% significance level, ** signifies rejection of H_0 at a 5% significance level, *** signifies rejection of H_0 at a 1% significance level.

Table 15a
Peru: Tests for number (r) of Cointegrating Vectors for $X_t = [\pi_t, \mu_t]$ with M_2[(a)]

TRACE TESTS	$H_0{:}r=0$ $H_1{:}r=2$	$H_0{:}r=1$ $H_1{:}r=2$
test statistic	32.38^{***}	1.56
MAXIMUM EIGENVALUE	$H_0{:}r=0$ $H_1{:}r=1$	$H_0{:}r=1$ $H_1{:}r=2$
test statistic	30.82^{***}	1.56
RESTRICTED ESTIMATES	β_μ	β_π
	1.000	-0.007
	α_μ	α_π
	-0.281	74.39

Table 15b
Peru: Tests on β' and α for inflation and M_2

$H_0{:}\beta_\pi=1,\ \beta_\mu=-1,$
$\chi^2_{(1)}=22.651^{***}$

Dickey-Fuller and Phillips-Perron Tests on the Final $\beta' X_t$[(a)]

H_0: $\beta' X_t$ is non-stationary	Without Trend	With Trend
Augmented Dickey-Fuller	-1.922	-2.78
Phillips-Perron	-3.91^{***}	-12.23^{***}

Notes: (a) One lag was used in these tests of stationarity. The lag structure was chosen by adding lags until the Q(22) statistic did not reject the null hypothesis of non-autocorrelated residuals.

* signifies rejection of H_0 at a 10% significance level, ** signifies rejection of H_0 at a 5% significance level, *** signifies rejection of H_0 at a 1% significance level.

Table 16
Argentina: Tests of the Present Value Model

Causality	$\Delta\mu_t \rightarrow Z_t$	3.78	R^2_μ	0.479
Tests[(a)]	$Z_t \rightarrow \Delta\mu_t$	41.18***	R^2_Z	0.195
Cross	$\alpha=1.45$	(k=2) 17.17***	$Q_\mu(12)$[(c)]	12.7
Equation Restrictions	$\alpha=3.624$	(k=2) 2.67	$Q_Z(12)$[(c)]	14.9
(40)[(b)]	$\alpha=5.145$	(k=2) 2.17		

Table 17
Bolivia: Tests of the Present Value Model

Causality	$\Delta\mu_t \rightarrow Z_t$	42.07***	R^2_μ	0.34
Tests[(a)]	$Z_t \rightarrow \Delta\mu_t$	16.58***	R^2_Z	0.33
Cross	$\alpha=2.23$	(k=2) 18.5***	$Q_\mu(21)$[(c)]	28.6
Equation Restrictions	$\alpha=3.36$	(k=2) 17.32***	$Q_Z(21)$[(c)]	24.3
(40)[(b)]	$\alpha=5.86$	(k=2) 16.1**		

Table 18
Brazil: Tests of the Present Value Model

Causality	$\Delta\mu_t \rightarrow Z_t$	2.155	R^2_μ	0.521
Tests[(a)]	$Z_t \rightarrow \Delta\mu_t$	21.05***	R^2_Z	0.0975
Cross	$\alpha=2.956$	(k=2) 17.80*** (k=3) 12.60**	$Q_\mu(21)$[(c)]	23.8
Equation Restrictions	$\alpha=5.231$	(k=2) 15.76*** (k=3) 12.90**	$Q_Z(21)$[(c)]	18.3
(40)[(b)]	$\alpha=11.249$	(k=2) 19.37*** (k=3) 12.54*		

Notes: (a) Test statistics are Wald statistics distributed as a $\chi^2_{(3)}$.
(b) Test statistics are Wald statistics distributed as a $\chi^2_{(6)}$.
(c) Test statistics are Box-Pierce Q statistics distributed as a $\chi^2_{(21)}$ with three lags in the VAR.

Table 19
Mexico: Tests of the Present Value Model

Causality	$\Delta\mu_t \rightarrow Z_t$	15.57***	R^2_μ	0.43
Tests(a)	$Z_t \rightarrow \Delta\mu_t$	23.67***	R^2_Z	0.01
Cross	$\alpha=2.27$	(k=2) 9.29	$Q_\mu(21)$(c)	45.7
Equation Restrictions	$\alpha=3.24$	(k=2) 9.07	$Q_Z(21)$(c)	43.0
(40)(b)	$\alpha=4.78$	(k=2) 8.34		

Table 20
Peru: Tests of the Present Value Model

Causality	$\Delta\mu_t \rightarrow Z_t$	7.91**	R^2_μ	0.43
Tests(a)	$Z_t \rightarrow \Delta\mu_t$	25.17***	R^2_Z	0.01
Cross	$\alpha=1.08$	(k=2) 9.07	$Q_\mu(21)$(c)	21.4
Equation Restrictions	$\alpha=3.66$	(k=2) 7.07	$Q_Z(21)$(c)	14.6
(40)(b)	$\alpha=15.56$	(k=2) 19.38***		

Notes: (a) Test statistics are Wald statistics distributed as a $\chi^2_{(3)}$.
(b) Test statistics are Wald statistics distributed as a $\chi^2_{(6)}$.
(c) Test statistics are Box-Pierce Q statistics distributed as a $\chi^2_{(21)}$ with three lags in the VAR.

Figure 1
Argentina: Monthly Inflation 1971:1-1984:2

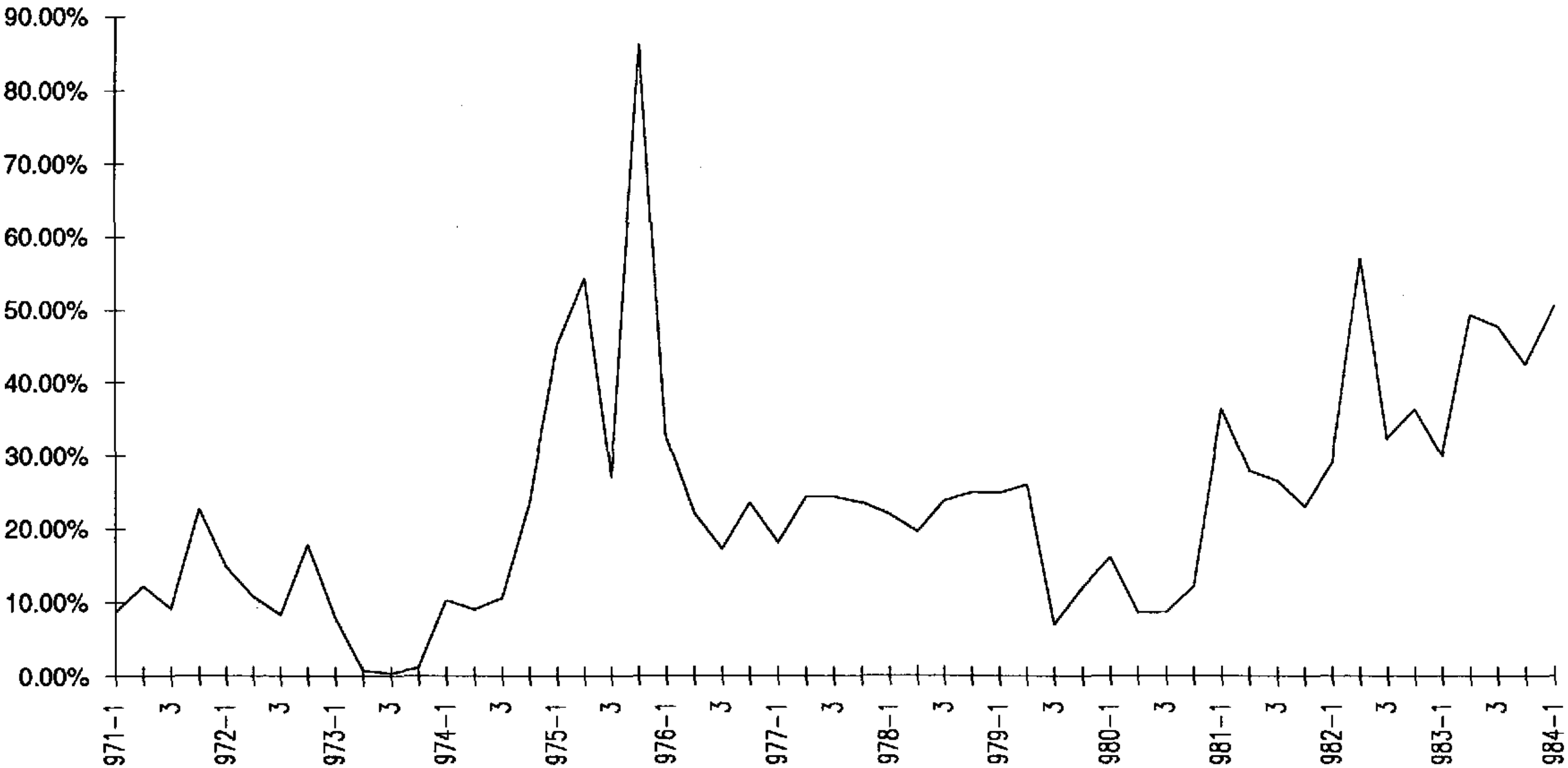


Figure 2
Bolivia: Monthly Inflation 1980:1-1990:9

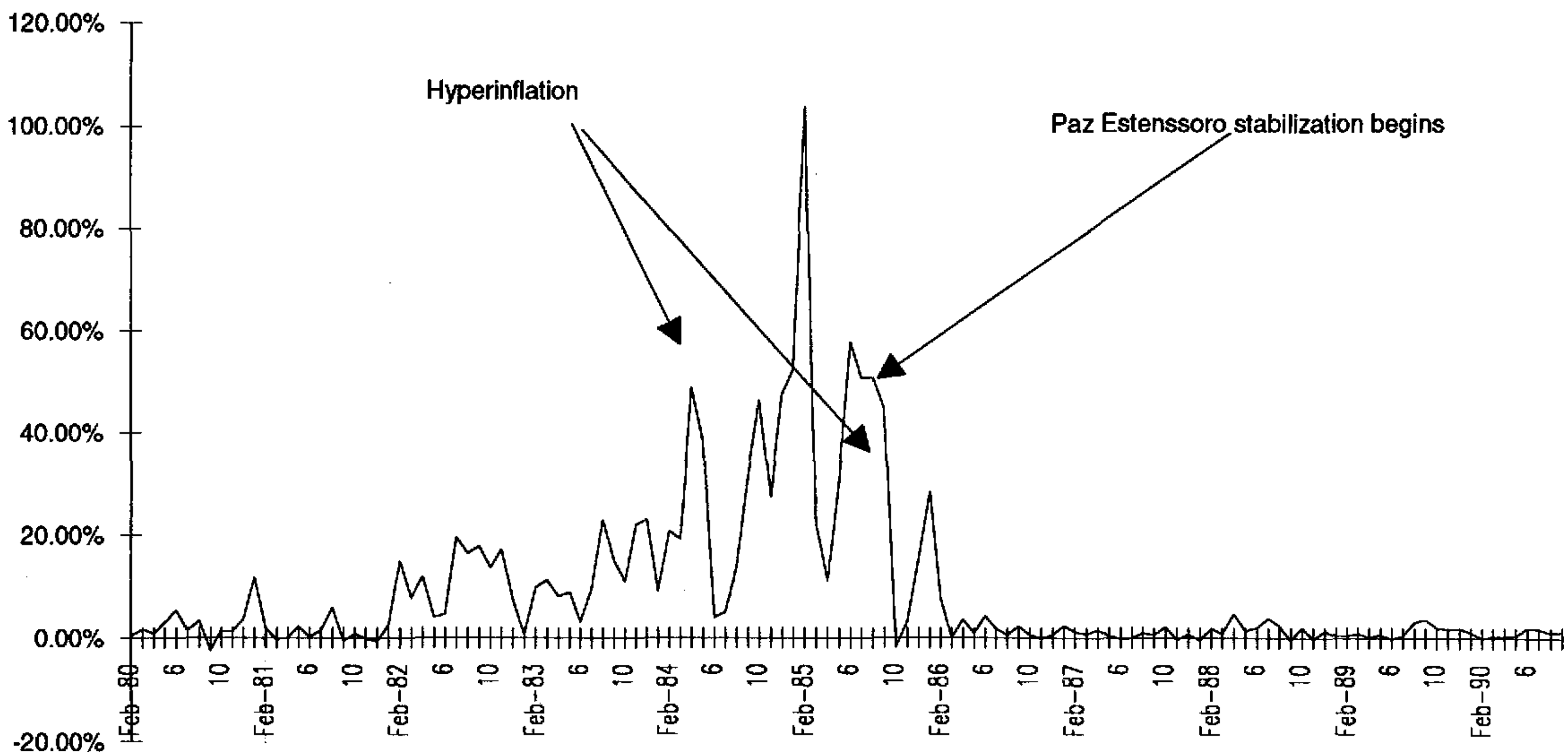


Figure 3
Brazil: Monthly Inflation 1974:2-1985:12

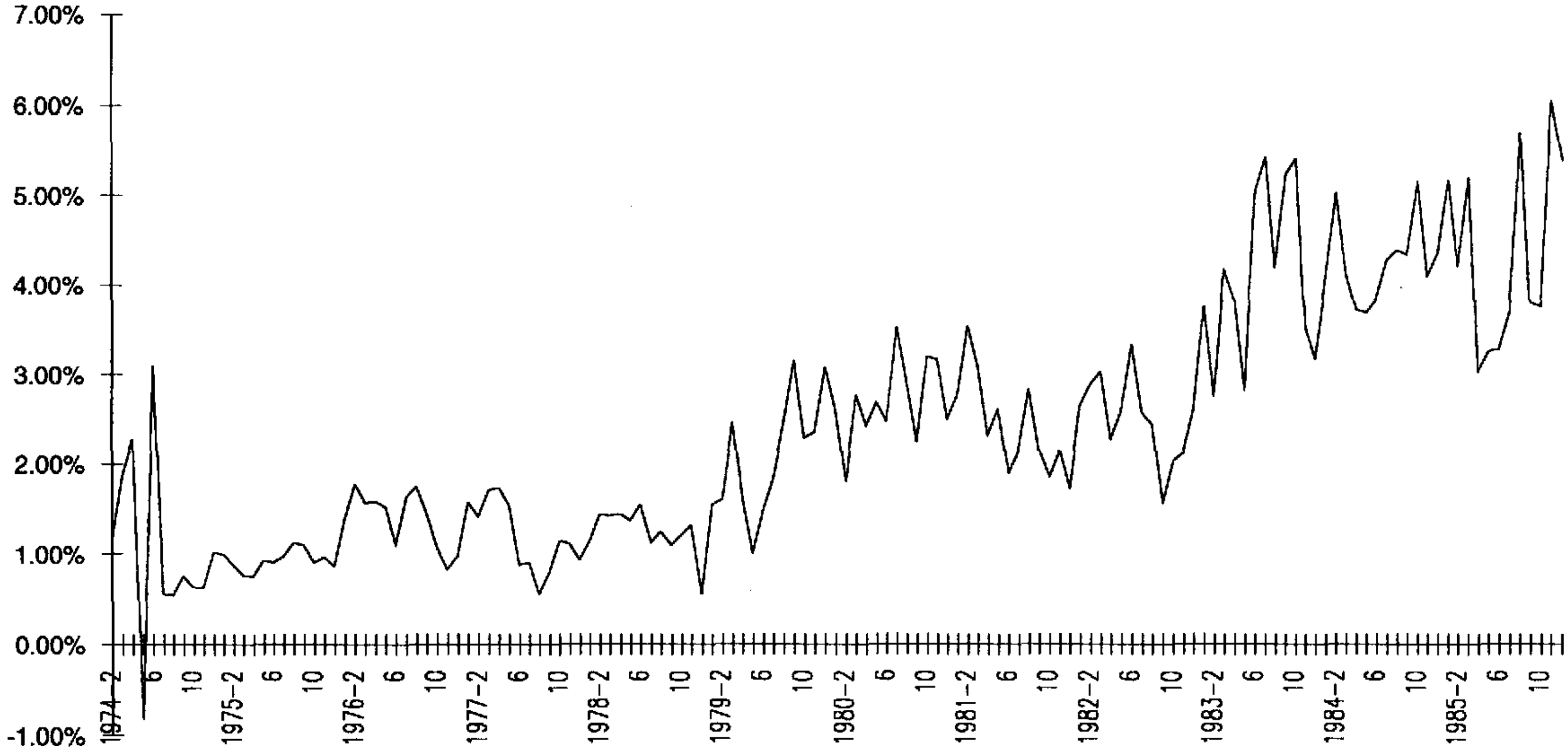


Figure 4
Mexico: Monthly Inflation 1960:2-1989:9

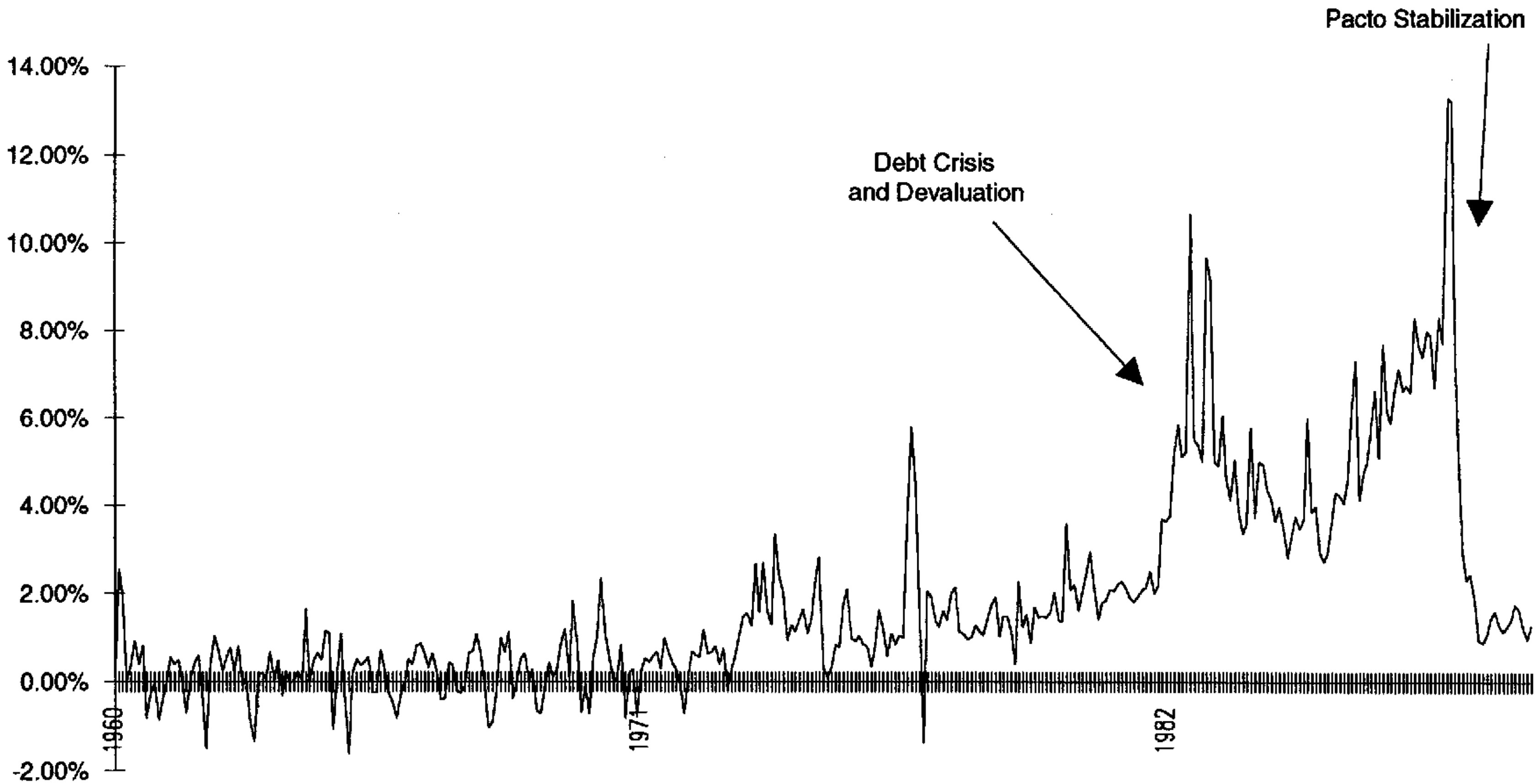


Figure 5:
Peru: Monthly Inflation 1982:2-1990:6

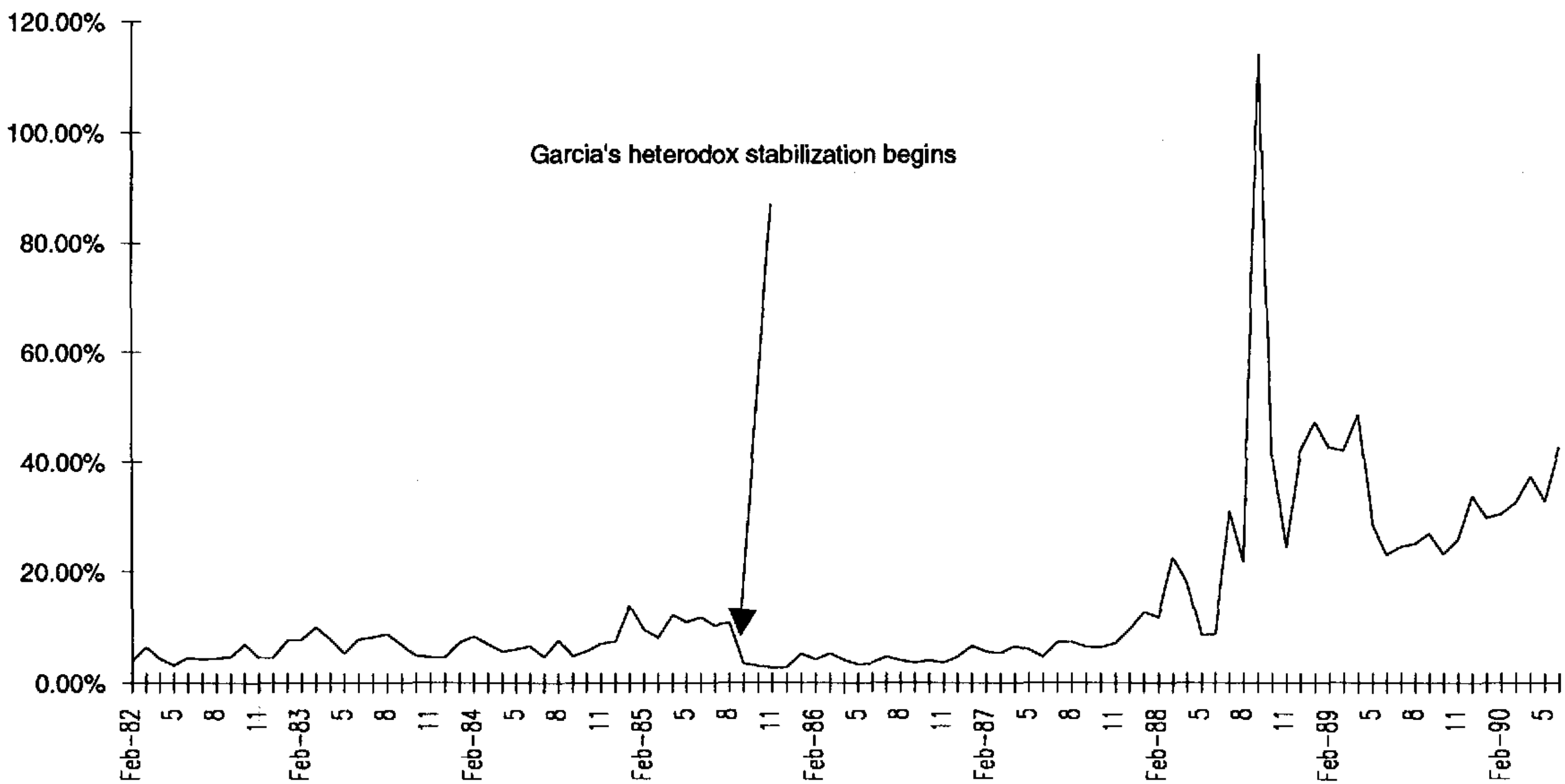


Figure 6
Argentina: Monthly Real M2 Growth 1971:1-1984:2

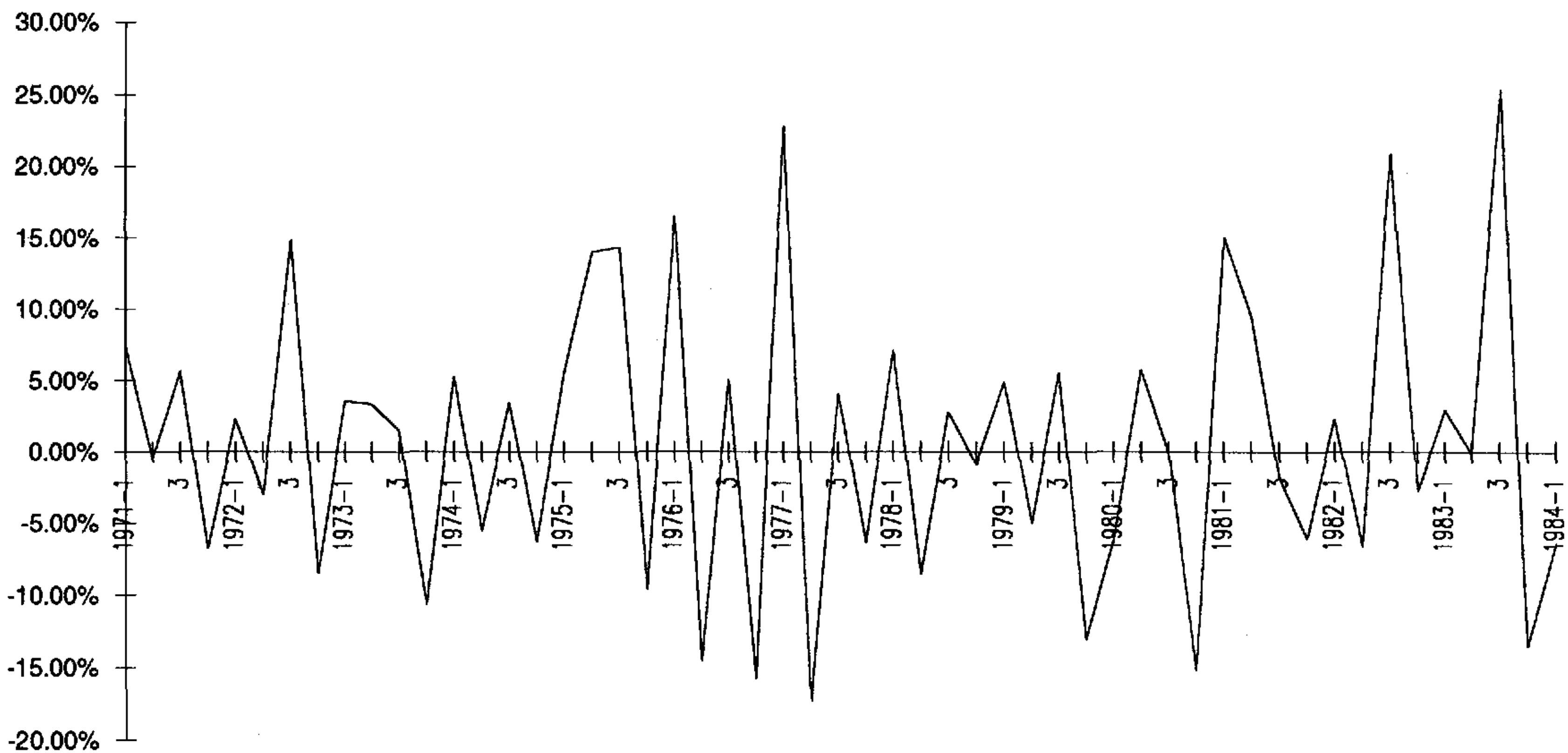


Figure 7
Bolivia: Real M2 Growth 1980:2-1990:9

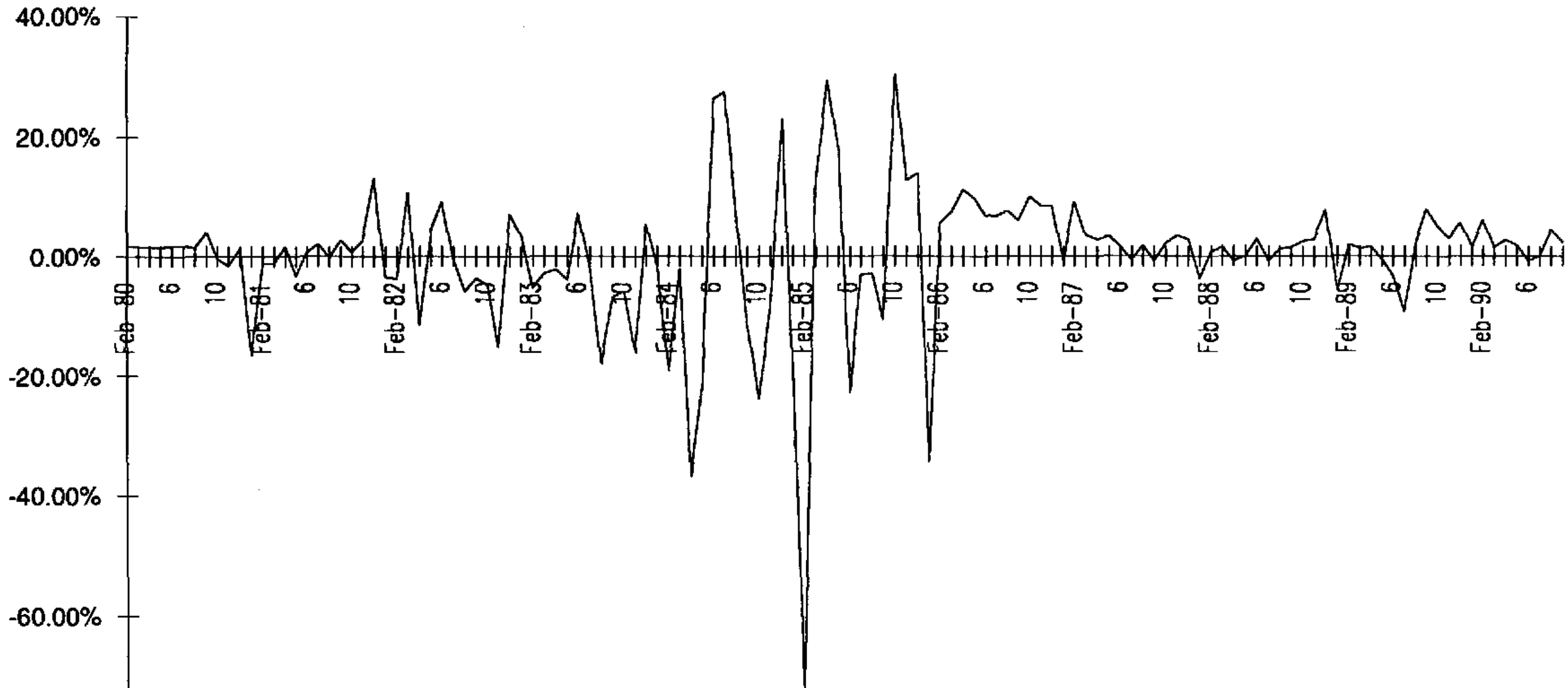


Figure 8
Brazil: Monthly Real M2 Growth 1974:2-1985:12

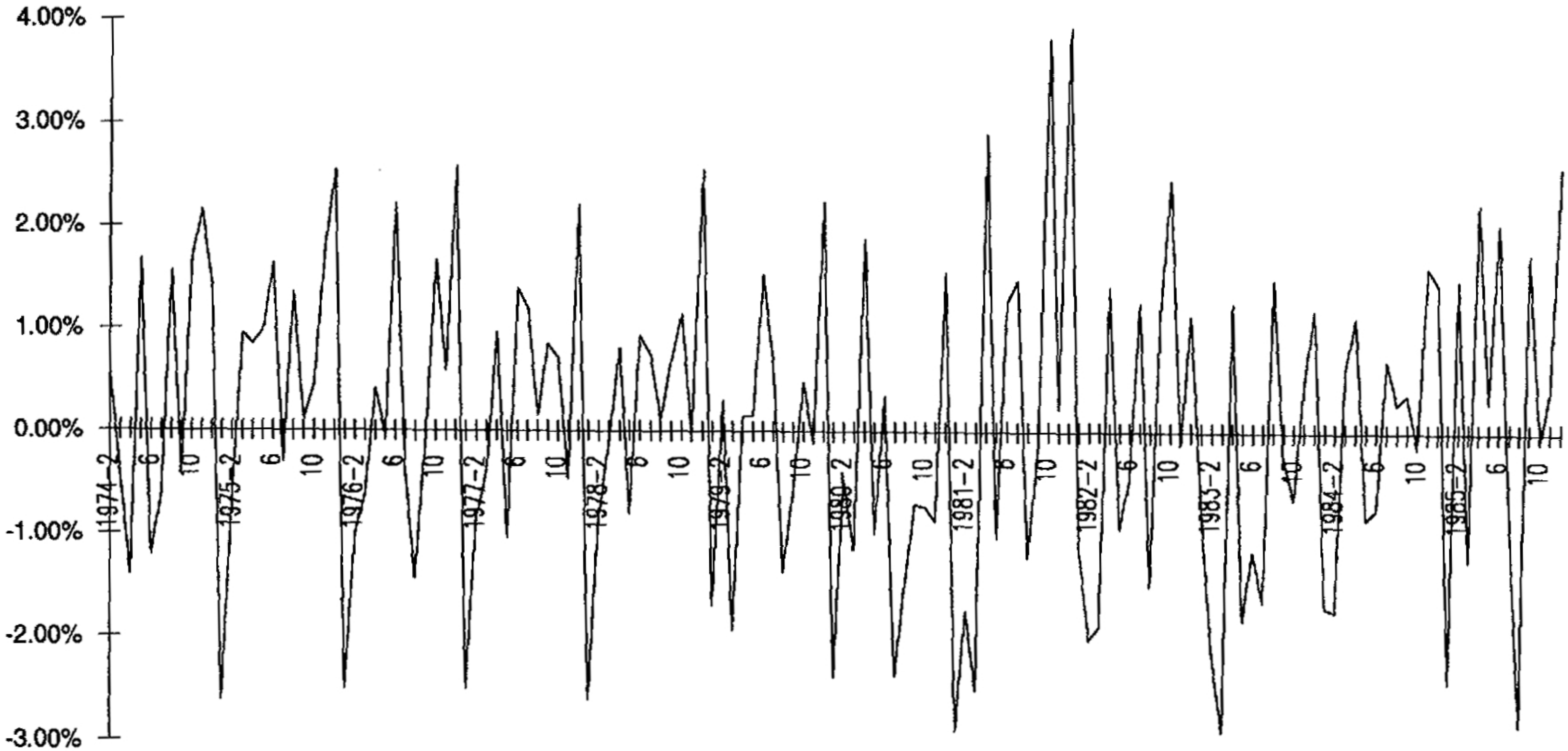


Figure 9
Mexico: Real Monthly M1 Growth 1960:2 to 1989:9

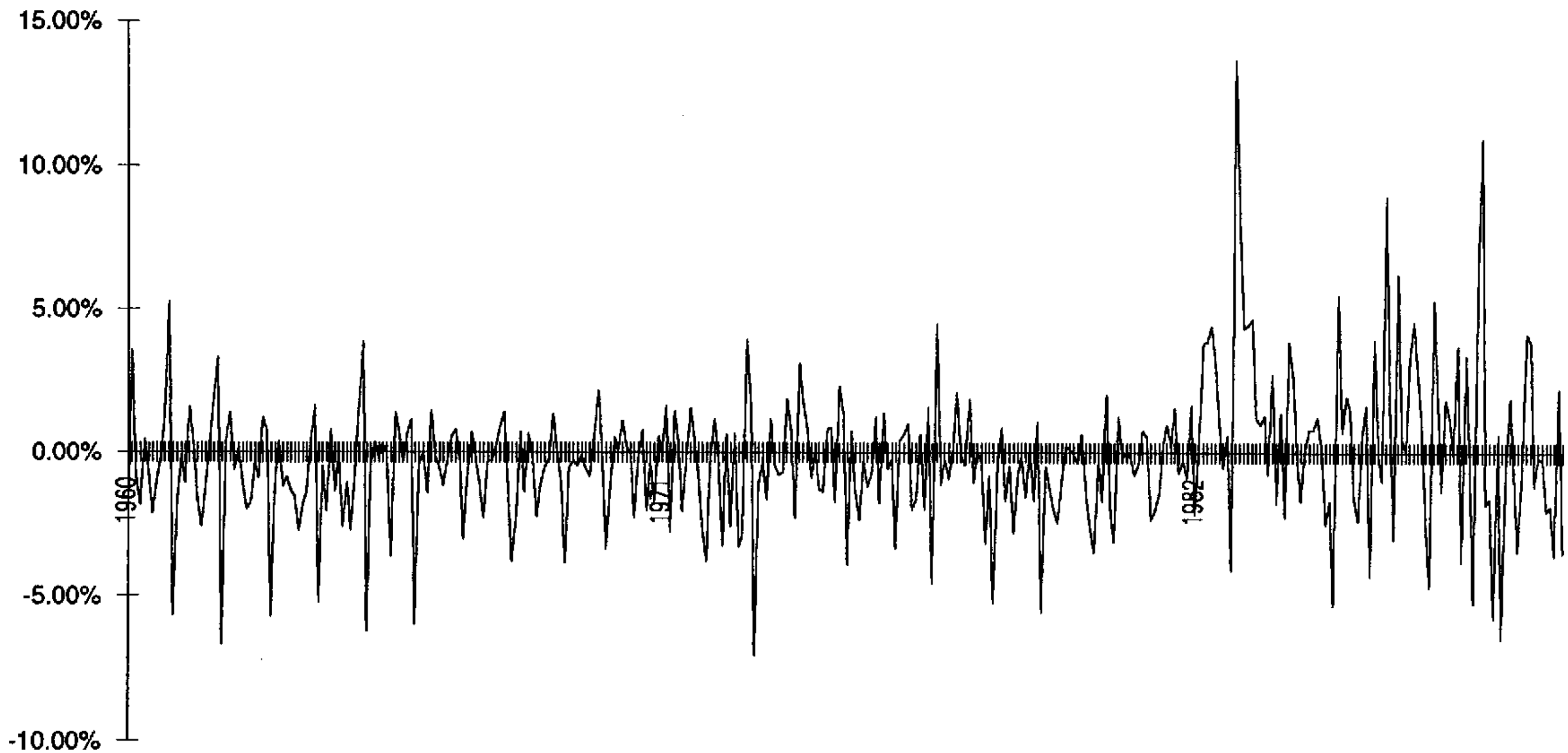
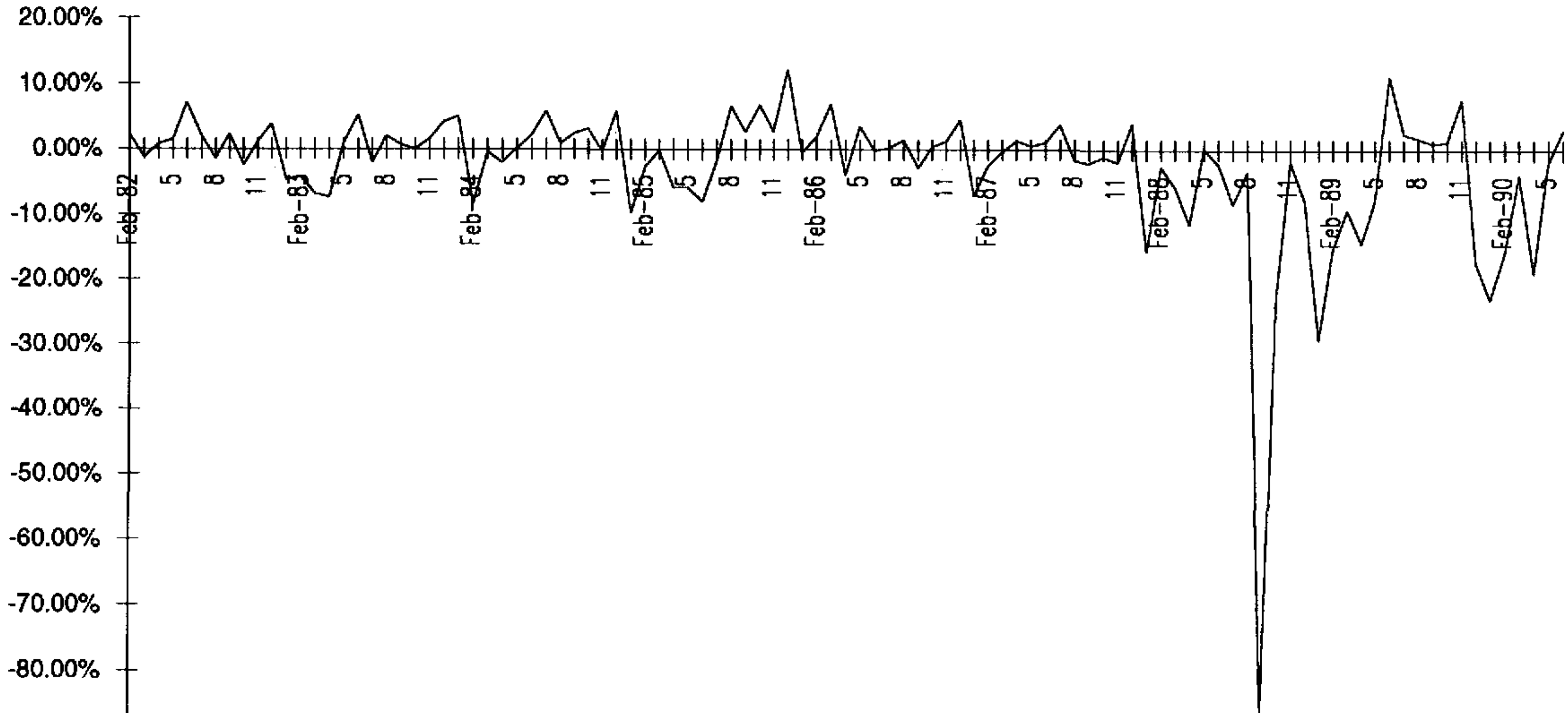


Figure 10
Peru: Real Monthly M2 Growth 1982:2-1990:6



Appendix: Tests of Stability

This appendix shows the results of CUSUM calculations for the inflation model in Argentina, Bolivia, Brazil, Mexico, and Peru.[1] The CUSUMs are calculated based upon the recursive residuals from the first equation of the VAR in equation (28) in the text of the following form

$$\Delta\mu_t = \theta_{11}(L)\Delta\mu_{t-1} + \theta_{12}(L)Z_{t-1} \tag{A1}$$

The CUSUM series are presented in Figures A1 through A5 along with 5% and 10% confidence bands. None of the series cross the significance bounds indicating no significant structural change. Sharp movements in the series, however, might indicate structural break. Argentina, (1980-81, 1983), Bolivia (1985), Mexico (1982-83, and Peru (1988 and 1990) show some moderate movements in the CUSUM series. Still, permanent regime changes and structural instability do not seem to be indicated by the data.

[1]For a good discussion of recursive residuals and CUSUM, see Harvey (1990).

Figure A1
Argentina: CUSUM 1973:1-1984:2

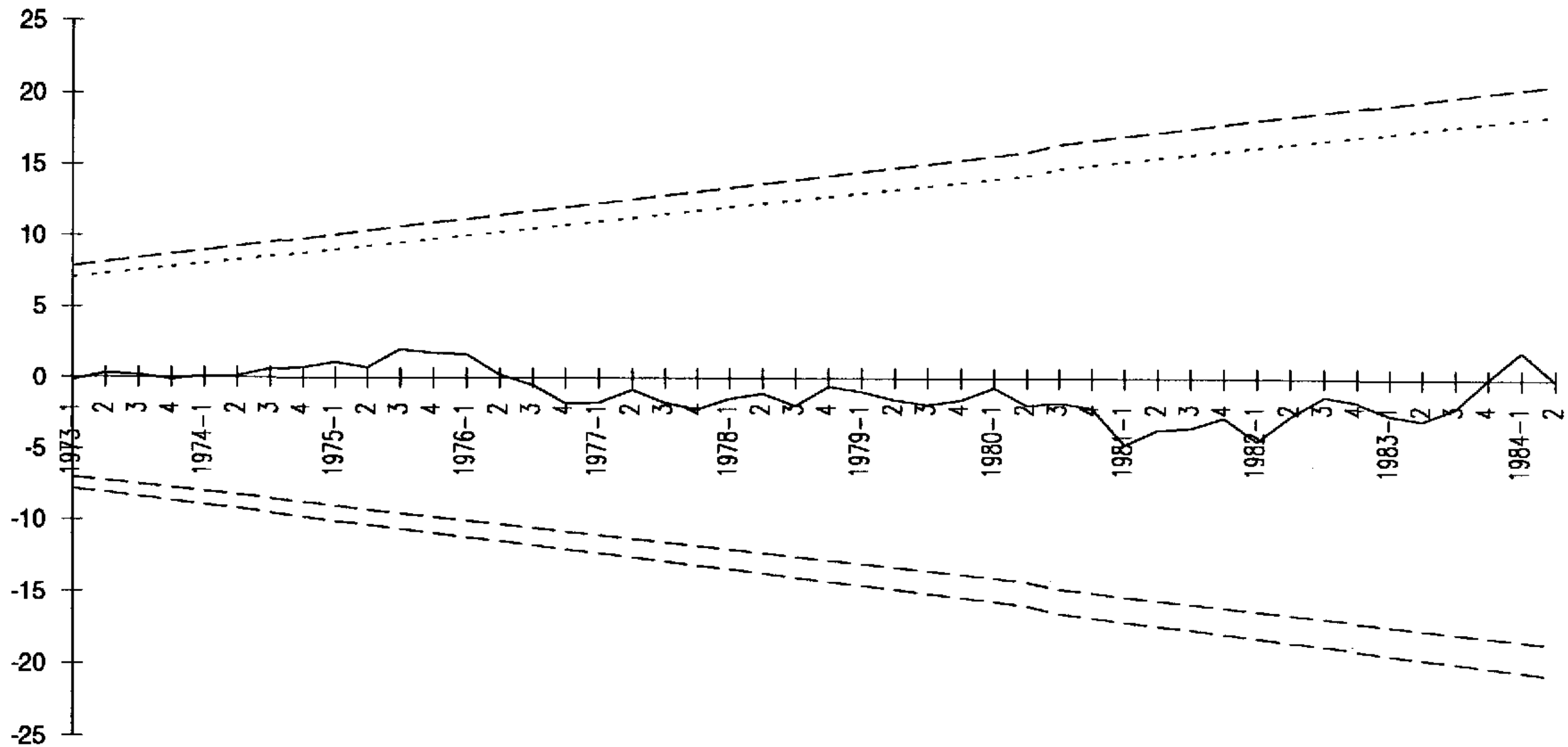


Figure A2
Bolivia: CUSUM 1981:1 to 1990:6

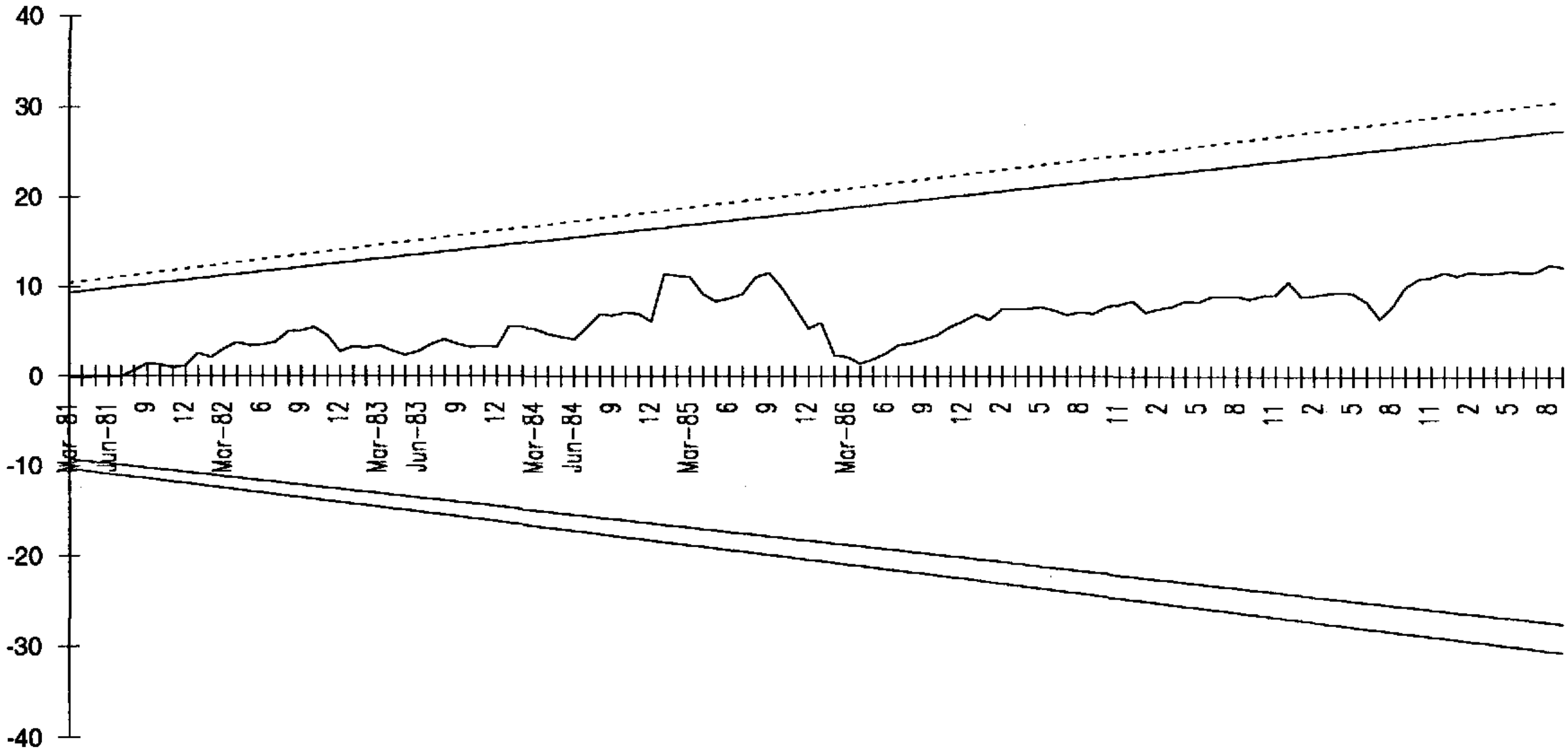


Figure A3
Brazil: CUSUM 1974:11-1985:12

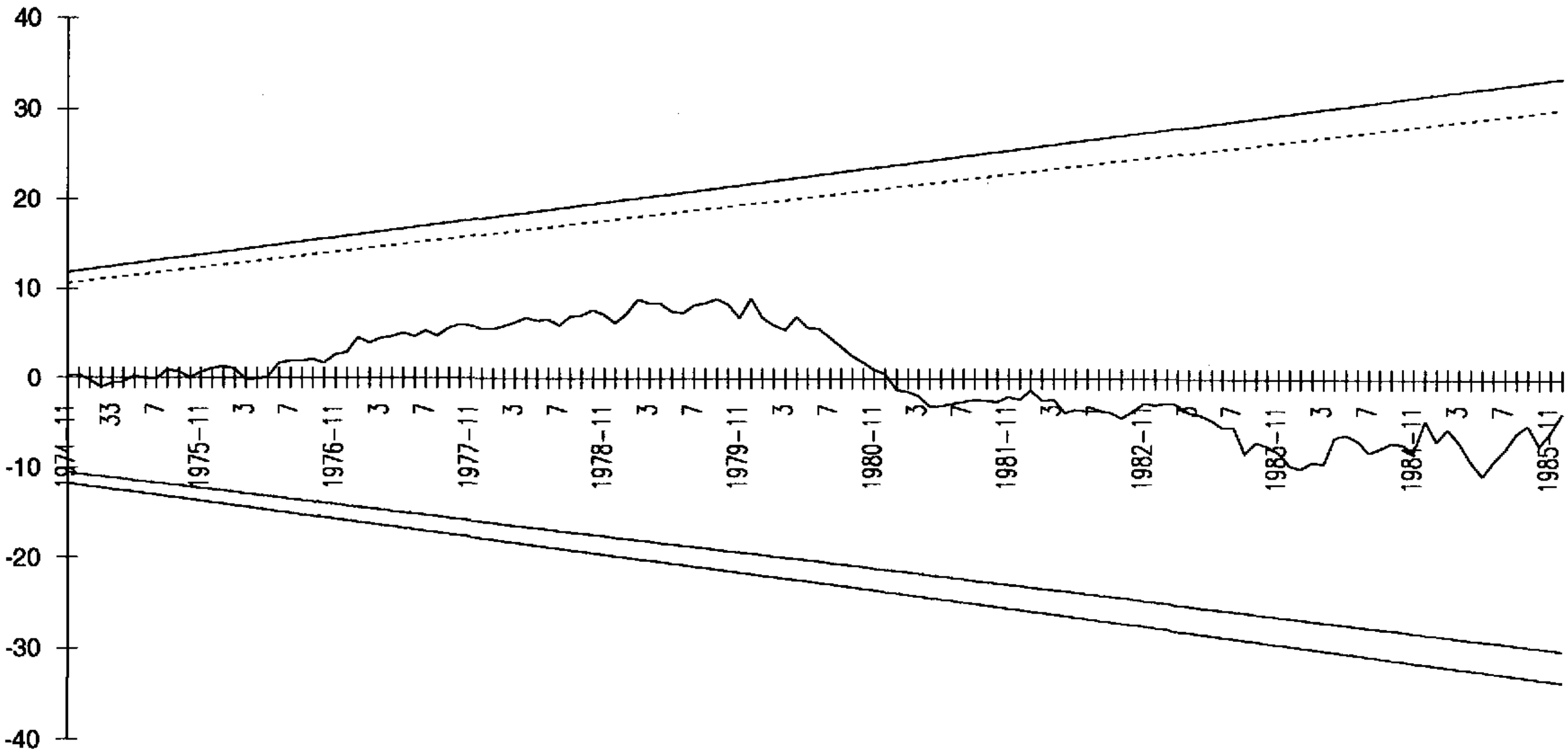


Figure A4
Mexico: CUSUM 1971:7 to 1989:9

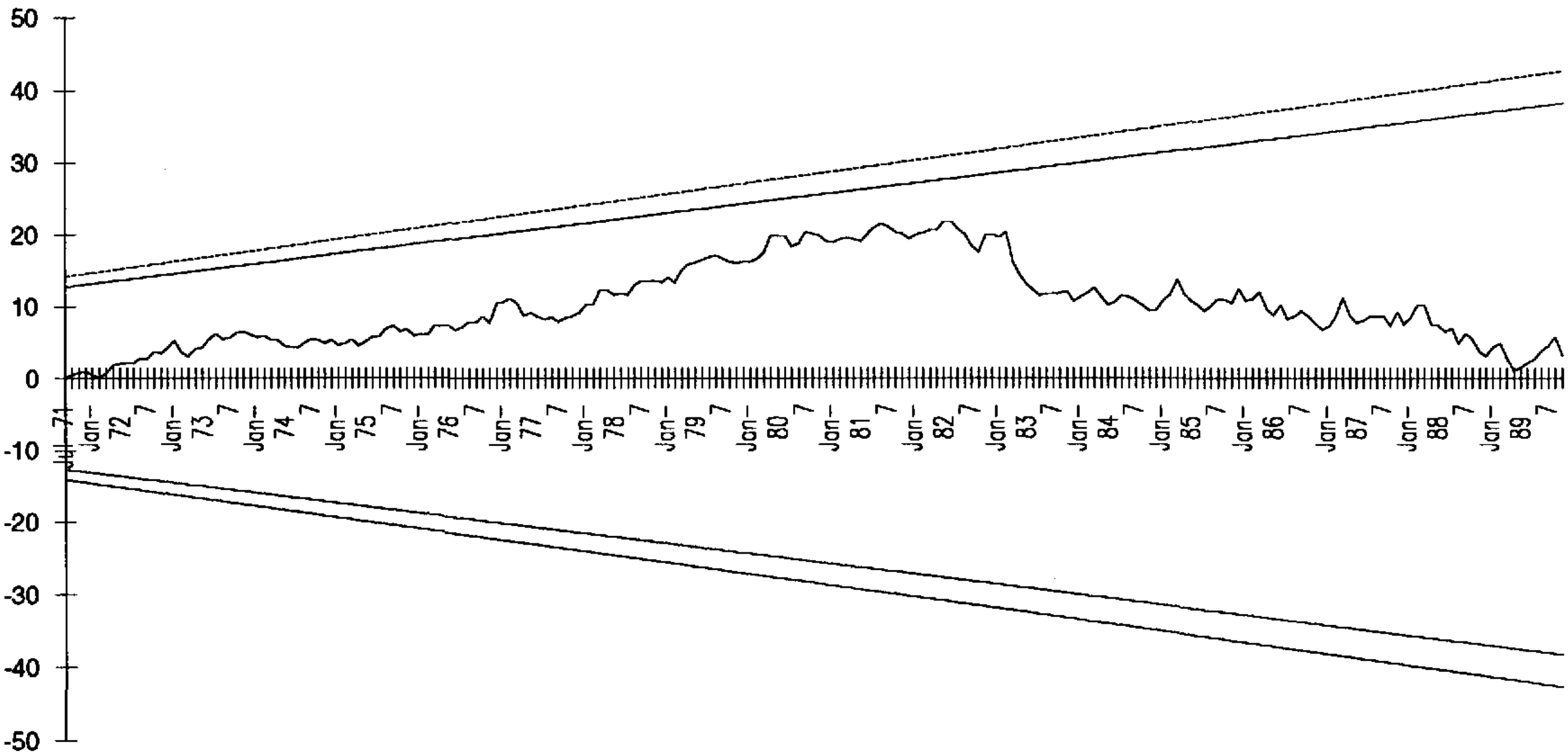


Figure A5
Peru: CUSUM 1983:1-1990:6