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

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Cointegration and Tests of a Classical Model of Inflation in Argentina, Bolivia, Brazil, Mexico, and Peru

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Abstract

Monetary models of inflation with rational expectations carry a number of testable implications. First, money growth and inflation should be cointegrated. Second, the equilibrium error may anticipate future monetary policy due to 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 this classical model of inflation is generally consistent with the inflation experiences of Argentina, Bolivia, Brazil, Mexico, and Peru despite their alleged heterogeneity. In all countries the data do not reject these three conditions all but one of the periods studied.

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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. Because an acceleration in inflation surprisingly accompanied 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. In the mid-1980s, the new-structuralist view came to dominate policy making circles in Latin America. But 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 and rational expectations.

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

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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).

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 the Mexican program proved a long term success. Four of the countries' 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, Bolivia and Mexico, had brought their inflation rates back to moderate levels by 1990. 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 a simple classical model successfully describes the inflationary process across these experiences.

Econometric studies conducted in the 1980s came to dismiss the relevance of monetary models because they detected instability in the demand for money.[3] Many of these studies, including Cardoso (1983), Gerlach and Simone (1985) and Rossi (1985), found that Brazilian money demand was unstable in the 1970s and early 1980s.[4] A notable exception is Calomiris and Domowitz (1989) who correctly model Brazilian money demand within an error correction framework and find that money demand was indeed stable in Brazil during a comparable period. As shown by Engle and Granger (1987), although not noted by the Calomiris and Domowitz, cointegrated series always have an error correction representation. The instability of the prior studies could have resulted from the misspecification involved with estimating cointegrated systems without regard to the order of integration of the variables and cointegrating relationship.

[3]One should note that these models did not directly hypothesize rational expectations,

[4]Darrat (1985) questions these conclusions on the basis of how lags are modeled, e.g. an Almon versus Cardoso's Koyck-adjustment mechanism.

Specifically, if the elements of money demand are cointegrated, then an autoregressive representation of money demand may display instability. This paper extends this research by systematically testing the validity of a simple classical monetary model with rational expectations that incorporates cointegration for a number of Latin American countries. Contrary to these studies, we find that this simple model is consistent with the inflationary experiences of Argentina, Bolivia, Brazil, Mexico, and Peru.

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 in models where the present price level (inflation rate) is a function of future expected price levels (inflation rates) (Diba and Grossman 1988a and 1988b). Theoretically, inflation can accelerate infinitely even though money growth remains stationary. Such bubbles, however, would have the growth rates of prices and money continuously diverging, ruling out 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 nonexistence 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 that 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 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).[5]

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, deriving the cross equation restrictions from the error correction form is relatively straightforward. The cross equation restrictions portend to test the rational expectations assertion that agents do not make systematic errors in predicting inflation. Unfortunately, rejection of these cross equation restrictions does not lead to any clear interpretation of the underlying inflation-money growth dynamic. But when we correctly account for regime changes, we find that the data satisfy these restrictions.[6]

Although we find that the data are consistent with this classical model, we cannot claim that this model explains inflation in these countries better than one that allows for "inertial inflation" stemming from widespread indexation and staggered wage contracts. But Novaes

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[5]The fact that money growth innovations follow innovations in inflation (Hanson 1980) has been treated as an anomaly as noted by Calomiris and Domowitz (1989). In our interpretation, such causality is consistent with the classical model.

[6]Another study by Phylaktis and Taylor (1993) rejected cross-equation restrictions in Cagan's (1956) money demand model. Our study differs from theirs, however, in three ways. First, we impose a specific stochastic behavior on real output. Second, our money demand specification is more general. And third, we carefully control for regime change. These differences may account for our more conclusive results.

(1993) offers evidence that purely inertial inflation cannot significantly explain Argentine and Brazilian inflation during the period we cover here. She rejects contractual inertia as an important source of inflation propagation in a modified version of Taylor's (1980) model. And her econometric results on the time series behavior are consistent with our own. We find her results encouraging and we feel they lend credence the appropriateness of the monetary model we present in the next section for Argentina, Bolivia, Brazil, Mexico, and Peru.

I. A Classical Model of Inflation

The model starts with a version of the Cagan's (1956) money demand specification.

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

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.[7] 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.[8] One can interpret real output here to represent permanent income as

[7]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).

[8]One can test this assumption using the techniques of Johansen and Juselius (1990). Data limitations force us to leave this to future research.

opposed to current income. And, as assumed below, this measure of income should display a unit root.

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-k+1} is the information set at time t-k+1. The model subsumes rational expectations, i.e. individuals use all information available to them to form expectations about future inflation rates.

Real output and real interest rates are assumed to follow random walks (real output also has a drift).[9]

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

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

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.

[9]Our interpretation that y_t represents permanent income is not necessary, however. Although recent evidence calls into question the assumption of a unit root in GDP in developed countries, this is not the case in developing countries. Basu and McLeod (1992) show that real output in developing countries is highly persistent and develop a model where temporary terms of trade shocks generate permanent changes in real output. The debate over persistence in output in industrialized nations is still not resolved as a recent study by Mocan (1994) finds a significant unit root in U.S. GNP.

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

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+1} and solving forward n periods into the future yields equation (9).

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

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

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

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

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

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

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

Solutions to (12) 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

$$E[\zeta_t|\Phi_{t-k+1}] = 0 \quad \forall\ k \geq 0 \tag{14}$$

The solution for inflation with a bubble is[10]

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

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

[10]To see this note that

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

Substituting this value into equation (9) yields the additive term B_t.

differencing of the inflation series. Taking first differences of the bubble in equation (13) using the lag operator L yields[11]

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

One could continue differencing this representation of the bubble. The ARMA representation of equation (16), 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 nonstationarity that cannot be differenced away.

The presence of bubbles also rules out cointegration between inflation and money growth. Reconsider equation (15) which states

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

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 (17) 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

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

growth and inflation rules out any nonstationarity 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, and that economic agents do not make systematic errors in forming these expectations. These relationships generate a set of easily testable restrictions on the inflationary process for rational expectations. 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

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

where C(L) is a 2 x 2 matrix in the lag operator and ν_t is a vector white noise process with ν_t

$= [\nu_{1,t}, \nu_{2,t}]$.

Engle and Granger (1987) show that the corresponding ARMA representation of the MA process of equation (19) will not be invertible and that an error correction form more appropriate. To see this, multiply both sides of equation (19) by the cointegrating vector [1, -1] to get

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

where Z_t equals the negative of real money growth π_t - μ_t.

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

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

where $\bar{0}$ 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 (19) will have an error correction representation[12]

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

where $A^*(0) = 0$, λ is a vector of constants, λ is a (2 x 1) vector of constants, $\det[C(L)] = [(1-L)b(L)]$, b(L) is a scalar lag polynomial, and λ is a vector of constants. As b(L) is invertible, premultiplying equation (20) by $b^{-1}(L)$ yields

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

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

where $D(L) = b^{-1}(L)[I-A^{*}(L)] = b^{-1}(L)A(L)$ and $g(L) = b^{-1}(L)$. Equation (23) can be rearranged in the following way[13]

$$H(L)\begin{bmatrix}(1 - L)\mu_t \\ Z_t\end{bmatrix} = \omega_t \qquad (24)$$

In order to generate optimal forecasts of money growth, we rewrite the VAR representation of equation (24) in the following way

$$Y_t = \Theta Y_{t-1} + e_t \qquad (25)$$

where

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

$$
Y_t = \begin{bmatrix} (1-L)\mu_t \\ (1-L)u_{t-1} \\ \cdot \\ \cdot \\ \cdot \\ \cdot \\ (1-L)\mu_{t-p-1} \\ (1-L)\mu_{t-p} \\ Z_t \\ Z_{t-1} \\ \cdot \\ \cdot \\ \cdot \\ Z_{t-p-1} \\ Z_{t-p} \end{bmatrix} \quad e_t = \begin{bmatrix} \omega_{1t} \\ 0 \\ \cdot \\ \cdot \\ \cdot \\ 0 \\ \omega_{2t} \\ 0 \\ \cdot \\ \cdot \\ \cdot \\ 0 \end{bmatrix} \tag{26}
$$

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

$$\Theta = \begin{bmatrix} \theta_{111} & \theta_{112} & \cdots & \theta_{11p-1} & \theta_{11p} & \theta_{121} & \theta_{122} & \cdots & \theta_{12p-1} & \theta_{12p} \\ 1 & 0 & \ldots & 0 & 0 & 0 & 0 & \ldots & 0 & 0 \\ 0 & 1 & \ldots & 0 & 0 & 0 & 0 & \ldots & 0 & 0 \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ 0 & 0 & \ldots & 1 & 0 & 0 & 0 & \ldots & 0 & 0 \\ \theta_{211} & \theta_{212} & \cdots & \theta_{21p-1} & \theta_{21p} & \theta_{221} & \theta_{222} & \cdots & \theta_{22p-1} & \theta_{22p} \\ 0 & 0 & \ldots & 0 & 0 & 1 & 0 & \ldots & 0 & 0 \\ 0 & 0 & \ldots & 0 & 0 & 0 & 1 & \ldots & 0 & 0 \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ \cdot & \cdot & \cdots & \cdot & \cdot & \cdot & \cdot & \cdots & \cdot & \cdot \\ 0 & 0 & \ldots & 0 & 0 & 0 & 0 & \ldots & 1 & 0 \end{bmatrix} \tag{27}$$

Optimal forecasts of the Y_t will thus be generated by

$$E\left[Y_{t+i} | \hat{\Phi}_{t-k+1}\right] = \Theta^{k+i} Y_{t-k} \tag{28}$$

where $\hat{\Phi}_{t-k+1} \equiv [Y_{t-j+1} | j \geq k] \subseteq \Phi_{t-k+1}$ is the information set available to the econometrician.

One important aspect of the VAR of equation (28) is that if the cointegrated present value model holds, Z_t either Granger-causes or is Granger-caused by 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, we test for such a causal relationship.

Equation (1) implies a set of restrictions on the optimal forecast equation (28).

Recall the money demand function

$$\mu_t - \pi_t = \bar{y} \\ - \alpha\left(E[\pi_{t+1} | \Phi_{t-k+1}] - E[\pi_t | \Phi_{t-k}]\right) + \xi_t \tag{29}$$

Rewriting equation (29) in the new notation yields

$$Z_t = - \bar{y} + \\ \alpha\left(E[Z_{t+1} | \Phi_{t-k+1}] + E[\mu_{t+1} | \Phi_{t-k+1}] - E[Z_t | \Phi_{t-k}] - E[\mu_t | \Phi_{t-k}]\right) \\ - \xi_t \tag{30}$$

Taking expectations conditional on Φ_{t-k+1} of equation (30) and rearranging yields

$$E\left[Z_t - \frac{\alpha}{1+\alpha} Z_{t+1} - \frac{\alpha}{1+\alpha} \Delta\mu_{t+1} - \frac{1}{1+\alpha} \bar{y} \;| \Phi_{t-k+1}\right] = 0 \tag{31}$$

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 (31) can be expressed as

$$H_1^o = R_1 \Theta^k - \frac{\alpha}{1+\alpha} R_1 \Theta^{k+1} - \frac{\alpha}{1+\alpha} R_2 \Theta^{k+1} = 0 \tag{32}$$

that are nonlinear in the parameter matrix Θ. The Wald statistic for this test is

$$T H_1^{o\prime} \left[\left(\frac{\partial H_1^o}{\partial \Theta} \right)' \hat{\Sigma}_\Theta \left(\frac{\partial H_1^o}{\partial \Theta} \right) \right]^{-1} H_1^o \sim \chi^2_{2p} \tag{33}$$

where T is the number of observations, Σ_Θ is the estimated covariance matrix of the estimated

Θ matrix and 2p are the number of restrictions and degrees of freedom.

III Empirical Results

a. Cointegration Tests

Before moving to the tests for cointegration, tests on the order of integration are in order.[14] Tables 1 through 5 show the Dickey-Fuller (1979) tests[15] for stationarity for money and prices in Argentina, Bolivia, Brazil, Mexico, and Peru. In all countries, 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 inflation is I(1) in all of them.

Generally, cointegration means that (nonstationary) 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 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,

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[14]Inflation in all countries is measured by the wholesale price index and either M_1 or M_2 were used as monetary aggregates. 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 1964 to December 1990 from the Banco Central de Peru.

[15]Phillips-Perron (1988) tests confirm these results.

there exists an error correction form

$$\Delta X_t = A_1 \Delta X_{t-1} + \ldots + A_{k-1} \Delta X_{t-k+1} + \Pi X_{t-1} + \varepsilon_t \tag{34}$$

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 (34).[16]

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + \Phi D_t + \varepsilon_t \tag{35}$$

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

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

[16]The Π matrix is the same in equation (42) and equation (43). One can show 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.

problem.[17]

$$|\lambda S_{kk} - S_{ko}S_{oo}^{-1}S_{ok}| = 0 \tag{36}$$

where S_{ij} are the residual moment matrices from the OLS regressions of ΔX_t and X_{t-k} on ΔX_{t-j}, $j = 1,...,k-1$. 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 Rank(Π)=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 Rank(Π)=r<p, then the data are I(1) and we have r cointegrating vectors. If Rank(Π)=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).

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 in tables 6 through 10. 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. The significant cointegrating relationships rule out rational inflationary bubbles in each case.

Tests on the cointegrating vector accord well with the model's specification. Specifically,

[17]The equations are estimated using RATS 3.10 software and Micro-TSP.

we test for long run money neutrality that takes the form of testing whether $\beta' = [1, -1]$. In all countries but Mexico, one cannot reject the neutrality of money at a 5% significance level. However, augmented Dickey-Fuller stationarity tests - confirmed in Figures 6 through 10 - for the growth in real balances show that in all cases real balances strongly reject the null hypothesis of nonstationarity. Growth in real balances in Mexico rejects the null hypothesis when no trend is included but does not when a trend is included. This result is somewhat unexpected but indicates that growth in real balances in Mexico has no trend.

b. Cross equation Restrictions

Tests on the cross equation restrictions are sensitive to regime changes because the VAR in equation (24) is not in reduced form. In other words, if the central bank changes its policy rule, the change in inflation expectations will necessarily change the parameter estimates of the VAR even when parameters of the money demand equation do not change.[18] Therefore, not only do we look at the full period under consideration but also subperiods.

Tables 11 through 15 present the remaining tests of the model. As mentioned above, if the model holds, equilibrium errors, Z_t, either anticipate changes in money growth or are anticipated by money growth. In all countries for the whole time period, Z_t significantly Granger-causes $\Delta\mu_t$ while significant causality in the other direction appears in Bolivia, Brazil, Mexico, and Peru.

In tests covering the period as a whole, the Argentine (Table 11), Mexican (Table 14),

[18]CUSUM tests for structural stability appear in appendix A. Even though structural breaks were not found, we divided the periods studied for each country.

and Peruvian (Table 15) data generally accord with the hypothesized model. For reasonable values of the semi-elasticity of money demand with respect to interest rates, α, the cross equation restrictions are not violated for information lags of (k-1=) 1 quarter in Argentina and 1 month in Mexico and Peru.[19] The model performs well also for the sub-periods 1972:1 to 1982:12 and 1983:1 to 1989:9 in Mexico and 1964:1 to 1985:6 and 1985:7 to 1990:12 in Peru. In other words, the model survives a number of external and policy shocks: the liberalization starting in 1977 and financial problems of 1982 in Argentina, the devaluation crises of 1976 and 1982 and the fall of oil prices in the 1980s in Mexico, Alan Garcia's heterodox experiment in the mid-1980s in Peru, and the international financial turbulence in all three countries after the onset of the debt crisis.

On the other hand, the model does not fare as well for Bolivia and Brazil. We then tested the model during different sub-periods for these two countries. For Bolivia, we tested the model for the period before the hyperinflation (1980:6 to 1984:12) and after the hyperinflation (1986:1 to 1990:9). In each of the sub-periods, the model performs well. Equilibrium errors anticipate changes in monetary policy in both periods while there is feedback in the first sub-period.

The Brazilian case is more complicated. After experimentation with different sub-periods, we settled on 1974:1 to 1980:12, 1981:1-1982:12, all of 1983, 1984:1 to 1985:12. We decided on the first sub-period because studies of Brazilian money demand usually find that the year 1980 is a watershed in that the indexation of financial assets was fixed at 50% while inflation approached 100% following on the heels of a maxi-devaluation in 1979 (Rossi 1989, Calomiris

[19]Dividing the Argentine data into sub-periods does not alter the results substantially. We do not present these results here but are available on request.

and Domowitz 1989, and Welch 1993). The next sub-period extends to the end of 1982 as another maxidevaluation was staged at the beginning of 1983 initiating a year of significant financial turmoil. The final period extends to December 1985, a few months before the implementation of the Cruzado Plan in March 1986.

The cross equation restrictions are not rejected in Brazil for any of these sub-periods except for the year 1983. The strong rejection of the restrictions for 1983 may reflect the uncertainty generated by the maxi-devaluation. Otherwise, the model performs well for the Brazilian case.

IV Conclusions

The inflation processes in Argentina, Bolivia, Brazil, Mexico, and Peru generally conform to the implications of the new classical model despite its simple form. We find that the data cannot reject the joint hypothesis that the classical model is valid and that agents form and act on rational expectations. Inflation and money growth are cointegrated in all countries ruling out speculative inflationary bubbles. Agents apparently anticipate future changes in money growth (and, by implication, inflation) in line with the rational expectations monetary model. Further, the cross equation restrictions implied by the model are not rejected for Argentina when the information lag is one quarter and for Bolivia, Brazil, Mexico, and Peru when the information lag is one month. These restrictions, however, are only rejected for Brazil in the period extending the year 1983.

The results show that forward looking expectations do play a part in the inflation process of all countries. Further, purely "speculative" sources play an insignificant role in all countries.

Certainly, the model is too simple to explain many other important aspects of inflation in these Latin American countries. Our findings certainly do not dismiss results pertaining to more structural explanations of inflation in Latin America because we do not explicitly incorporate them in the alternative hypothesis. But recent work by Novaes (1993) shows that "inertial" sources of inflation in Argentina and Brazil are not important in explaining inflation persistence. And the stochastic properties of the series she investigated, although consistent with our findings, are not consistent with inertial explanations of inflation. In light of her findings, we feel more comfortable concluding that our model reasonably explains the inflation process in these countries. And our results cast doubt on the assertion that classical models of inflation are irrelevant for the Latin American experience. The simple model tested in this paper performs surprisingly well.

Further research on inflation and money demand should improve our understanding of inflation in Latin America that incorporates new techniques on evaluating cointegrated time series.[20] A full analysis of all the variables in the model using the Johansen and Juselius (1990) technique in addition to incorporating the possibility of I(2) processes should yield more informative results.

[20]Examples consistent with the results found here for Argentina are found in Ahumada (1992 and 1993).

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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
Inflation[b]	-2.19
ΔInflation[b]	-6.81***
Money Growth (M_2)[b]	-1.77
ΔMoney Growth (M_2)[b]	-5.76***

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

Variable	Augmented Dickey-Fuller
Inflation[b]	-2.67
ΔInflation[b]	-6.75***
Money Growth (M_2)[b]	-2.19
ΔMoney Growth (M_2)[b]	-5.70***

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

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

The order of the autoregressive terms, 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 nonnormality 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 2
Bolivia: Unit Roots Tests[a]

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

Variable	Augmented Dickey-Fuller
Inflation[b]	-3.28**
ΔInflation[b]	-12.03***
Money Growth (M_2)[b]	-2.54
ΔMoney Growth (M_2)[b]	-4.05***

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

Variable	Augmented Dickey-Fuller
Inflation[b]	-3.38*
ΔInflation[b]	-12.00***
Money Growth (M_2)[b]	-2.60
ΔMoney Growth (M_2)[b]	-4.12***

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

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

The order of the autoregressive terms, 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 nonnormality 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 3
Brazil: Unit Roots Tests[a]

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

Variable	Augmented Dickey-Fuller
Inflation[b]	-1.72
ΔInflation[b]	-14.88***
Money Growth (M_1)[b]	1.813
ΔMoney Growth (M_1)[b]	-7.37***

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

Variable	Augmented Dickey-Fuller
Inflation[b]	-5.41***
ΔInflation[b]	-14.93***
Money Growth (M_1)[b]	0.180
ΔMoney Growth (M_1)[b]	-7.73***

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

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

The order of the autoregressive terms, 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 6 lags.

(b) Series showed significant nonnormality 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 4
Mexico: Unit Roots Tests[a]

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

Variable	Augmented Dickey-Fuller
Inflation[b]	-2.48
ΔInflation[b]	-10.57***
Money Growth (M_1)[b]	0.932
ΔMoney Growth (M_1)[b]	-4.12***

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

Variable	Augmented Dickey-Fuller
Inflation[b]	-3.58**
ΔInflation[b]	-10.56***
Money Growth (M_1)[b]	-1.31
ΔMoney Growth (M_1)[b]	-4.12***

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

$$y_t = \mu + \tau t + \phi' y_{t-1} + \sum_{i=1}^{q} \phi_i \Delta y_{t-i} \tag{i}$$

The order of the autoregressive terms, 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 4 lags.

(b) Series showed significant nonnormality 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 5
Peru: Unit Roots Tests[a]

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

Variable	Augmented Dickey-Fuller
Inflation[b]	-1.50
ΔInflation[b]	-8.985***
Money Growth (M_2)[b]	-2.02
ΔMoney Growth (M_2)[b]	-12.63***

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

Variable	Augmented Dickey-Fuller
Inflation[b]	-2.715
ΔInflation[b]	-9.02***
Money Growth (M_1)[b]	-3.27
ΔMoney Growth (M_2)[b]	-6.91***

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

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

The order of the autoregressive terms, q, was chosen to render the residuals of the regression white noise according to the Box-Pierce Q(22) statistic. The regressions used 4 lags.

(b) Series showed significant nonnormality 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 6
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
Tests on β' 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 nonstationary	Without Trend	With Trend
Augmented Dickey-Fuller	-3.73***	-3.716**

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 nonautocorrelated residuals.

(b) Series showed significant nonnormality 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 7
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
Tests on β' 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: $\beta'X_t$ is nonstationary	Without Trend	With Trend
Augmented Dickey-Fuller	-7.84***	-3.716**

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 nonautocorrelated residuals.

(b) Series showed significant nonnormality 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 8
Brazil: Tests for number (r) of Cointegrating Vectors
for $X_t = [\pi_t, \mu_t]$ with M_1[(a)]

TRACE TESTS	$H_o{:}r=0$ $H_1{:}r=2$	$H_o{:}r=1$ $H_1{:}r=2$
test statistic	25.33^{***}	0.025
MAXIMUM EIGENVALUE	$H_o{:}r=0$ $H_1{:}r=1$	$H_o{:}r=1$ $H_1{:}r=2$
test statistic	25.30^{***}	0.025
UNRESTRICTED ESTIMATES	β_μ	β_π
	1.000	-0.866
	α_μ	α_π
	-0.616	0.186
Tests on ß' for inflation and M_1		
	$H_o{:}\beta_\pi=1,\ \beta_\mu=-1$ $\chi^2_{(1)}=1.094$	
Dickey-Fuller and Phillips-Perron Tests on the Final $\beta'X_t$[(a)]		
H_o: $\beta'X_t$ is nonstationary	Without Trend	With Trend
Augmented Dickey-Fuller	-4.68^{***}	-4.66^{***}

Notes: (a) Two 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 nonautocorrelated residuals.

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

Table 9
Mexico: Tests for number (r) of Cointegrating Vectors
for $X_t = [\pi_t, \mu_t]$ with M_1[(a)]

TRACE TESTS	H_o:r=0 H_1:r=2	H_o:r=1 H_1:r=2
test statistic	54.75***	5.91
MAXIMUM EIGENVALUE	H_o:r=0 H_1:r=1	H_o: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
Tests on ß' for inflation and M_1		
	H_o:β_π=1, β_μ=-1	
	$\chi^2_{(1)}$=15.972***	
Dickey-Fuller and Phillips-Perron Tests on the Final $\beta'X_t$[(a)]		
H_o: $\beta'X_t$ is nonstationary	Without Trend	With Trend
Augmented Dickey-Fuller	-4.97***	-2.78

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 nonautocorrelated residuals.

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

Table 10
Peru: Tests for number (r) of Cointegrating Vectors
for $X_t = [\pi_t, \mu_t]$ with M_2[(a)]
(with trend)

TRACE TESTS	H_o:r=0 H_1:r=2	H_o:r=1 H_1:r=2
test statistic	41.84***	0.71
MAXIMUM EIGENVALUE	H_o:r=0 H_1:r=1	H_o:r=1 H_1:r=2
test statistic	41.13***	0.71
RESTRICTED ESTIMATES	$ß_\mu$	$ß_\pi$
	1.000	-0.903
	α_μ	α_π
	-0.949***	-0.106
	Tests on ß' for inflation and M_2	
	H_o:$ß_\pi$=1, $ß_\mu$=-1	
	$\chi^2_{(1)}$ = 3.02	
	Dickey-Fuller and Phillips-Perron Tests on the Final $ß'X_t$[(a)]	
H_o: $ß'X_t$ is nonstationary	Without Trend	With Trend
Augmented Dickey-Fuller	-4.73***	-4.80***

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 nonautocorrelated residuals.

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

Table 11

Argentina: Tests of the Present Value Model

Causality Tests[a]	$\Delta\mu_t \rightarrow Z_t$	3.78	R^2_μ	0.479
	$Z_t \rightarrow \Delta\mu_t$	41.18***	R^2_Z	0.195
Cross Equation Restrictions (24)[b]	α=1.45	(k=2) 17.17***	$Q_\mu(12)$[c]	12.7
	α=3.624	(k=2) 2.67	$Q_Z(12)$[c]	14.9
	α=5.145	(k=2) 2.17		

Notes: (a) Test statistics are Wald statistics distributed as a $\chi^2_{(2)}$.

(b) Test statistics are Wald statistics distributed as a $\chi^2_{(4)}$.

(c) Test statistics are Box-Pierce Q statistics distributed as a χ^2 with two lags in the VAR and degrees of freedom equal to the order of autocorrelation.

Table 12a
Bolivia: Tests of the Present Value Model (1980:6-1990:9)

Causality Tests[(a)]	$\Delta\mu_t \rightarrow Z_t$	42.07***	R^2_μ	0.34
	$Z_t \rightarrow \Delta\mu_t$	16.58***	R^2_Z	0.33
Cross Equation Restrictions (24)[(b)]	α=2.23	(k=2) 18.5***	$Q_\mu(33)^{(c)}$	28.6
	α=3.36	(k=2) 17.32***	$Q_Z(33)^{(c)}$	24.3
	α=5.86	(k=2) 16.1**		

Table 12b
Bolivia: Tests of the Present Value Model (1980:6-1984:12)

Causality Tests[(a)]	$\Delta\mu_t \rightarrow Z_t$	13.07***	R^2_μ	0.37
	$Z_t \rightarrow \Delta\mu_t$	12.62***	R^2_Z	0.23
Cross Equation Restrictions (24)[(b)]	α=2.23	(k=2) 5.87	$Q_\mu(21)^{(c)}$	13.3
	α=3.36	(k=2) 5.05	$Q_Z(21)^{(c)}$	11.7
	α=5.86	(k=2) 5.49		

Table 12c
Bolivia: Tests of the Present Value Model (1986:1-1990:9)

Causality Tests[(a)]	$\Delta\mu_t \rightarrow Z_t$	4.51	R^2_μ	0.46
	$Z_t \rightarrow \Delta\mu_t$	12.27***	R^2_Z	0.21
Cross Equation Restrictions (24)[(b)]	α=2.23	(k=2) 5.98	$Q_\mu(21)^{(c)}$	11.4
	α=3.36	(k=2) 4.93	$Q_Z(21)^{(c)}$	16.4
	α=5.86	(k=2) 5.56		

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 χ^2 with three lags in the VAR and degrees of freedom equal to the order of autocorrelation.

Table 13a
Brazil: Tests of the Present Value Model 1974:1-1985:12

Causality Tests[a]	$\Delta\mu_t \rightarrow Z_t$	7.02*	R^2_μ	0.532
	$Z_t \rightarrow \Delta\mu_t$	22.0***	R^2_Z	0.179
Cross Equation Restrictions (40)[b]	α=2.956	(k=2) 42.26*** (k=3) 14.56**	$Q_\mu(19)$[c]	20.6
	α=5.231	(k=2) 40.42*** (k=3) 21.06**	$Q_Z(19)$[c]	16.8
	α=11.249	(k=2) 40.04*** (k=3) 10.78*		

Table 13b
Brazil: Tests of the Present Value Model 1974:1-1980:12

Causality Tests[a]	$\Delta\mu_t \rightarrow Z_t$	8.70*	R^2_μ	0.502
	$Z_t \rightarrow \Delta\mu_t$	9.03**	R^2_Z	0.111
Cross Equation Restrictions (24)[b]	α=2.956	(k=2) 4.15	$Q_\mu(19)$[c]	24.7
	α=5.231	(k=2) 4.19	$Q_Z(19)$[c]	20.5
	α=11.249	(k=2) 4.48		

Table 13c
Brazil: Tests of the Present Value Model 1981:1-1982:12

Causality Tests[a]	$\Delta\mu_t \rightarrow Z_t$	1.52	R^2_μ	0.599
	$Z_t \rightarrow \Delta\mu_t$	2.95	R^2_Z	0.151
Cross Equation Restrictions (40)[b]	α=2.956	(k=2) 6.58	$Q_\mu(19)$[c]	16.7
	α=5.231	(k=2) 6.64	$Q_Z(19)$[c]	10.1
	α=11.249	(k=2) 7.83		

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 χ^2 with three lags in the VAR and degrees of freedom equal to the order of autocorrelation.

Table 13d
Brazil: Tests of the Present Value Model 1983:1-1983:12

Causality	$\Delta\mu_t \rightarrow Z_t$	6.39^{*}	R^2_{μ}	0.566
Tests[(a)]	$Z_t \rightarrow \Delta\mu_t$	1.11	R^2_{Z}	0.474
Cross	α=2.956	(k=2) 50.5^{***}	$Q_{\mu}(6)$[(c)]	4.02
Equation Restrictions	α=5.231	(k=2) 54.3^{***}	$Q_Z(6)$[(c)]	3.49
(24)[(b)]	α=11.249	(k=2) 96.38^{***}		

Table 13e
Brazil: Tests of the Present Value Model 1984:1-1985:12

Causality	$\Delta\mu_t \rightarrow Z_t$	4.81	R^2_{μ}	0.618
Tests[(a)]	$Z_t \rightarrow \Delta\mu_t$	2.71	R^2_{Z}	0.277
Cross	α=2.956	(k=2) 7.51	$Q_{\mu}(19)$[(c)]	23.3
Equation Restrictions	α=5.231	(k=2) 8.35	$Q_Z(19)$[(c)]	19.5
(24)[(b)]	α=11.249	(k=2) 8.645		

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 χ^2 with three lags in the VAR and degrees of freedom equal to the order of autocorrelation.

Table 14a

Mexico: Tests of the Present Value Model (1972:1-1989:9)

Causality Tests[(a)]	$\Delta\mu_t \rightarrow Z_t$	15.57***	R^2_μ	0.46
	$Z_t \rightarrow \Delta\mu_t$	23.67***	R^2_Z	0.07
Cross	α=2.27	(k=2) 9.29	$Q_\mu(41)$[(c)]	45.7
Equation Restrictions	α=3.24	(k=2) 9.07	$Q_Z(41)$[(c)]	43.0
(40)[(b)]	α=4.78	(k=2) 8.34		

Table 14b

Mexico: Tests of the Present Value Model (1972:1-1982:12)

Causality Tests[(a)]	$\Delta\mu_t \rightarrow Z_t$	6.67*	R^2_μ	0.43
	$Z_t \rightarrow \Delta\mu_t$	19.21***	R^2_Z	0.01
Cross	α=2.27	(k=2) 3.43	$Q_\mu(33)$[(c)]	45.7
Equation Restrictions	α=3.24	(k=2) 3.15	$Q_Z(33)$[(c)]	43.0
(40)[(b)]	α=4.78	(k=2) 2.91		

Table 14c

Mexico: Tests of the Present Value Model (1973:1-1989:9)

Causality Tests[(a)]	$\Delta\mu_t \rightarrow Z_t$	9.45**	R^2_μ	0.48
	$Z_t \rightarrow \Delta\mu_t$	8.61**	R^2_Z	0.04
Cross	α=2.27	(k=2) 6.62	$Q_\mu(27)$[(c)]	21.2
Equation Restrictions	α=3.24	(k=2) 6.86	$Q_Z(27)$[(c)]	9.7
(24)[(b)]	α=4.78	(k=2) 6.96		

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 χ^2 with three lags in the VAR and degrees of freedom equal to the order of autocorrelation.

Table 15a
Peru: Tests of the Present Value Model (1964:1-1990:12)

Causality	$\Delta\mu_t \rightarrow Z_t$	7.91^{**}	R^2_μ	0.695
Tests[(a)]	$Z_t \rightarrow \Delta\mu_t$	25.17^{***}	R^2_Z	0.225
Cross	α=1.08	(k=2) 9.04	$Q_\mu(19)^{(c)}$	19.8
Equation Restrictions	α=3.66	(k=2) 3.85	$Q_Z(19)^{(c)}$	8.55
(24)[(b)]	α=15.56	(k=2) 3.41		

Table 15b
Peru: Tests of the Present Value Model (1964:1-1985:6)

Causality	$\Delta\mu_t \rightarrow Z_t^{(a)}$	1.78	R^2_μ	0.771
Tests[(a)]	$Z_t \rightarrow \Delta\mu_t^{(a)}$	129.31^{***}	R^2_Z	0.331
Cross	α=1.08[(b)]	(k=2) 10.28	$Q_\mu(19)^{(c)}$	8.69
Equation Restrictions	α=3.66[(b)]	(k=2) 9.80	$Q_Z(19)^{(c)}$	8.52
(24)[(b)]	α=15.56[(a)]	(k=2) 8.34		

Table 15c
Peru: Tests of the Present Value Model (1985:7-1990:12)

Causality	$\Delta\mu_t \rightarrow Z_t$	1.83	R^2_μ	0.339
Tests[(a)]	$Z_t \rightarrow \Delta\mu_t$	12.45^{***}	R^2_Z	0.052
Cross	α=1.08	(k=2) 1.97	$Q_\mu(19)^{(c)}$	10.10
Equation Restrictions	α=3.66	(k=2) 1.33	$Q_Z(19)^{(c)}$	9.44
(40)[(b)]	α=15.56	(k=2) 1.18		

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 χ^2 with three lags in the VAR and degrees of freedom equal to the order of autocorrelation.

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

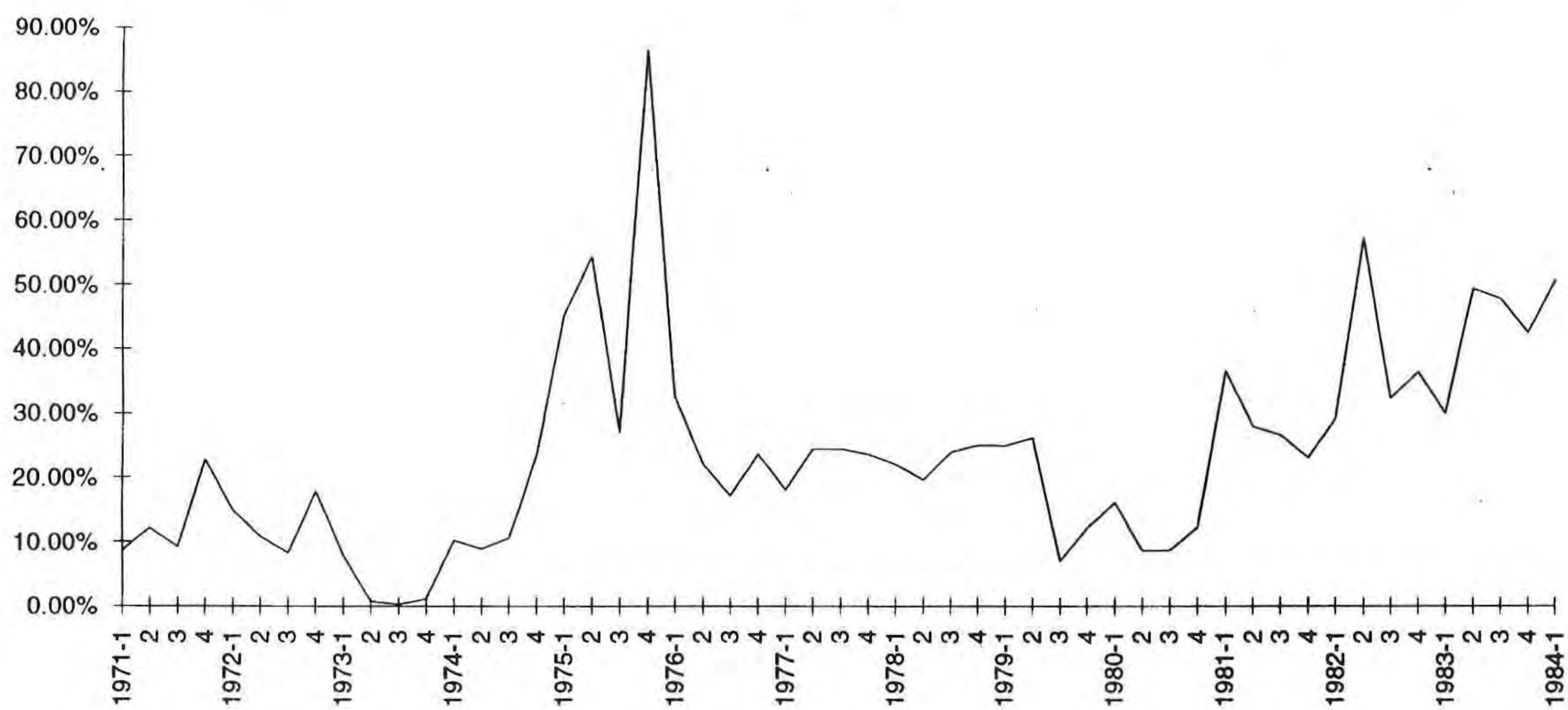


Figure 2
Bolivia: Monthly Inflation 1980:1-1990:9

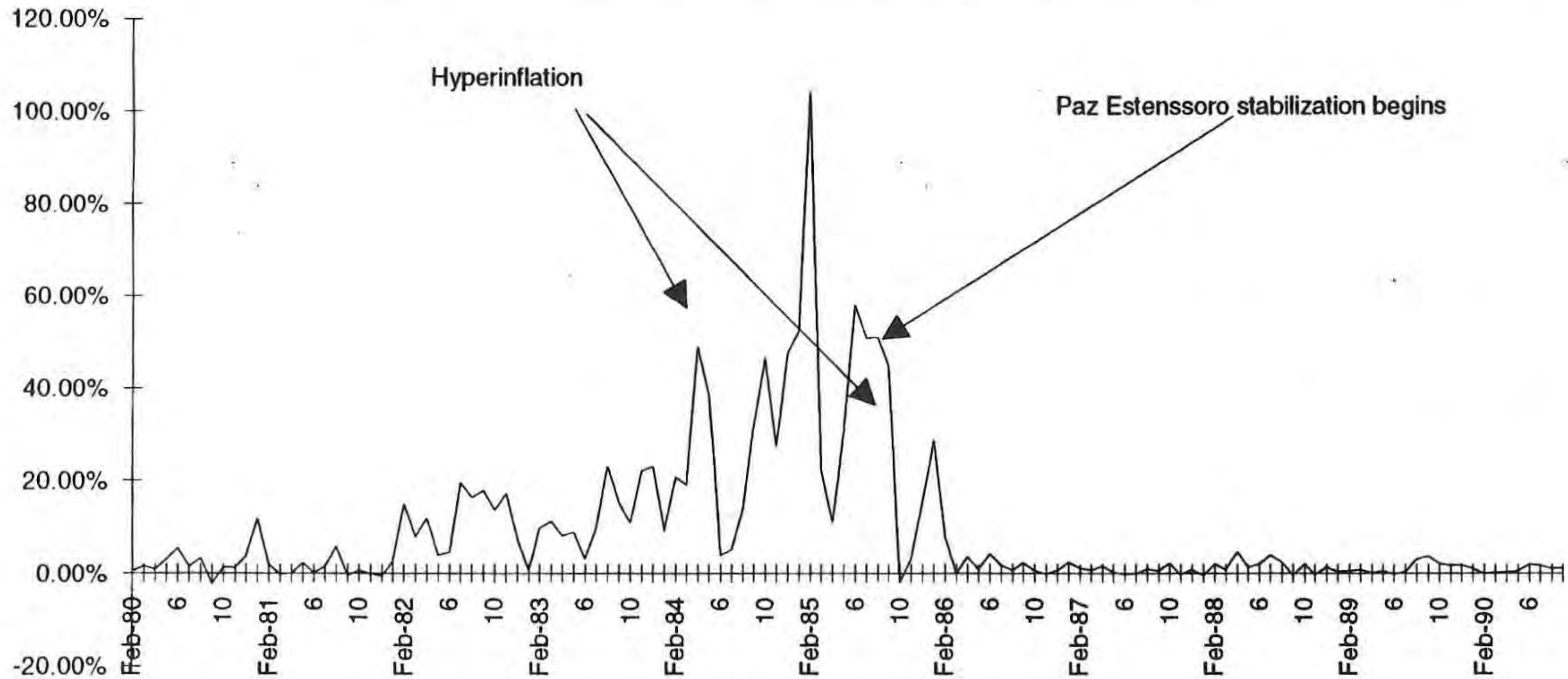


Figure 3
Brazil: Monthly Inflation 1974:2-1985:12

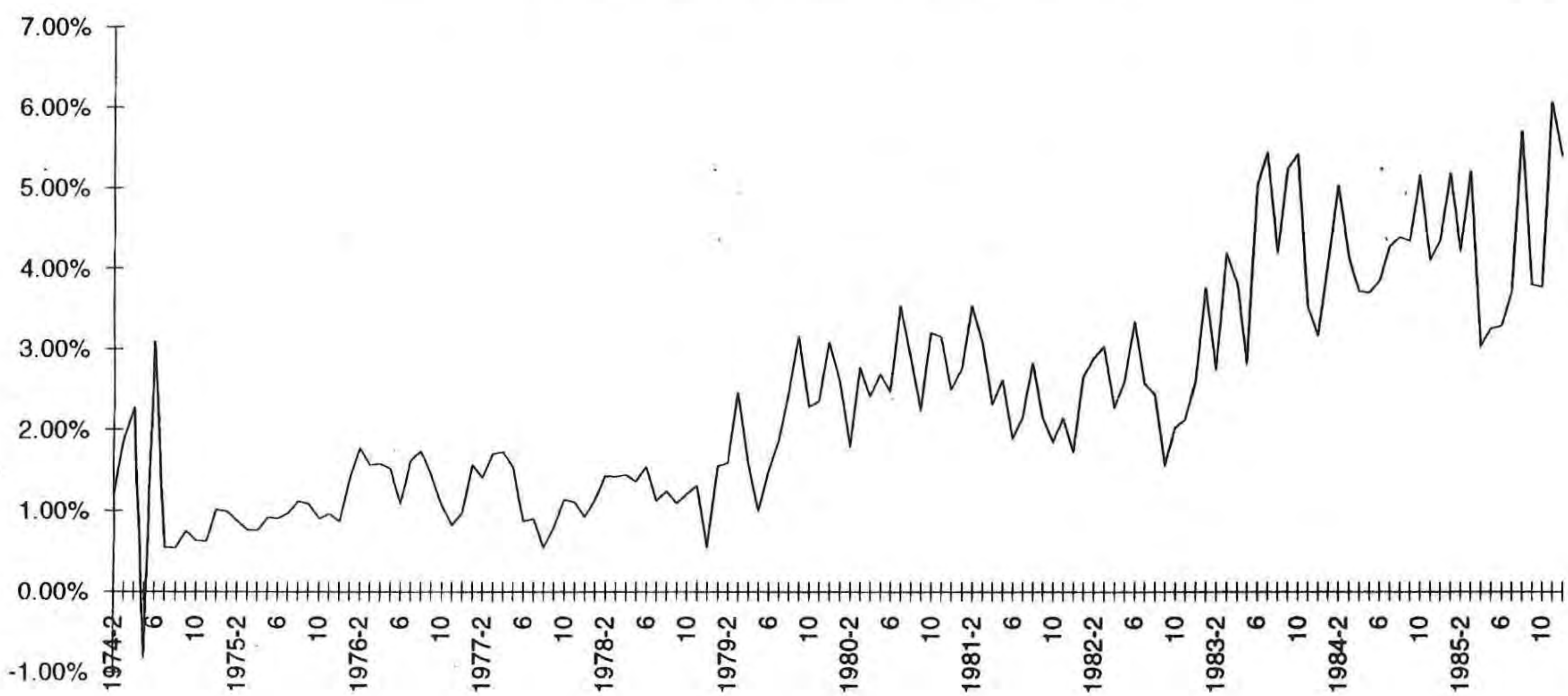


Figure 4
Mexico: Monthly Inflation 1960:2-1989:9

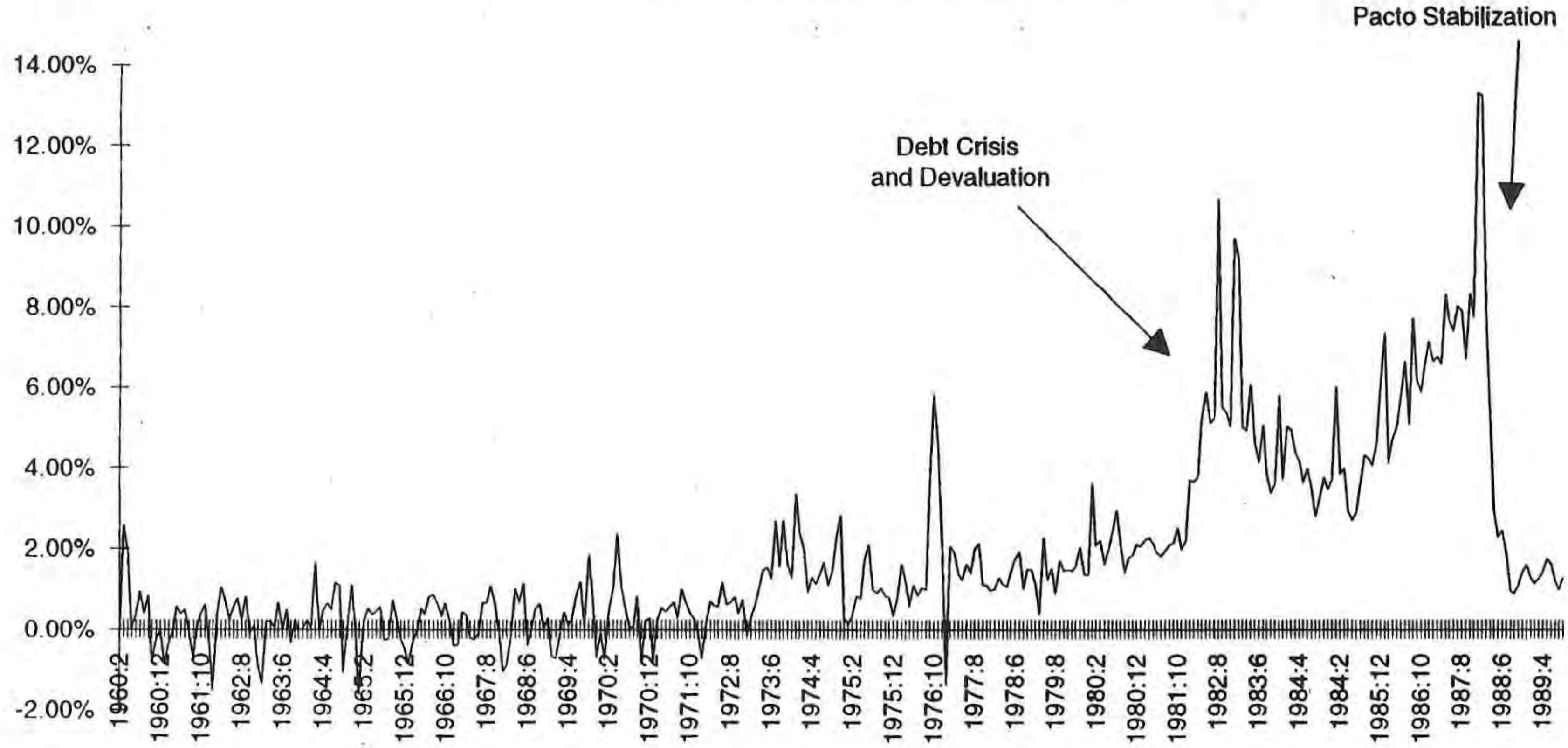


Figure 5:
Peru: Monthly Inflation 1982:2-1990:6

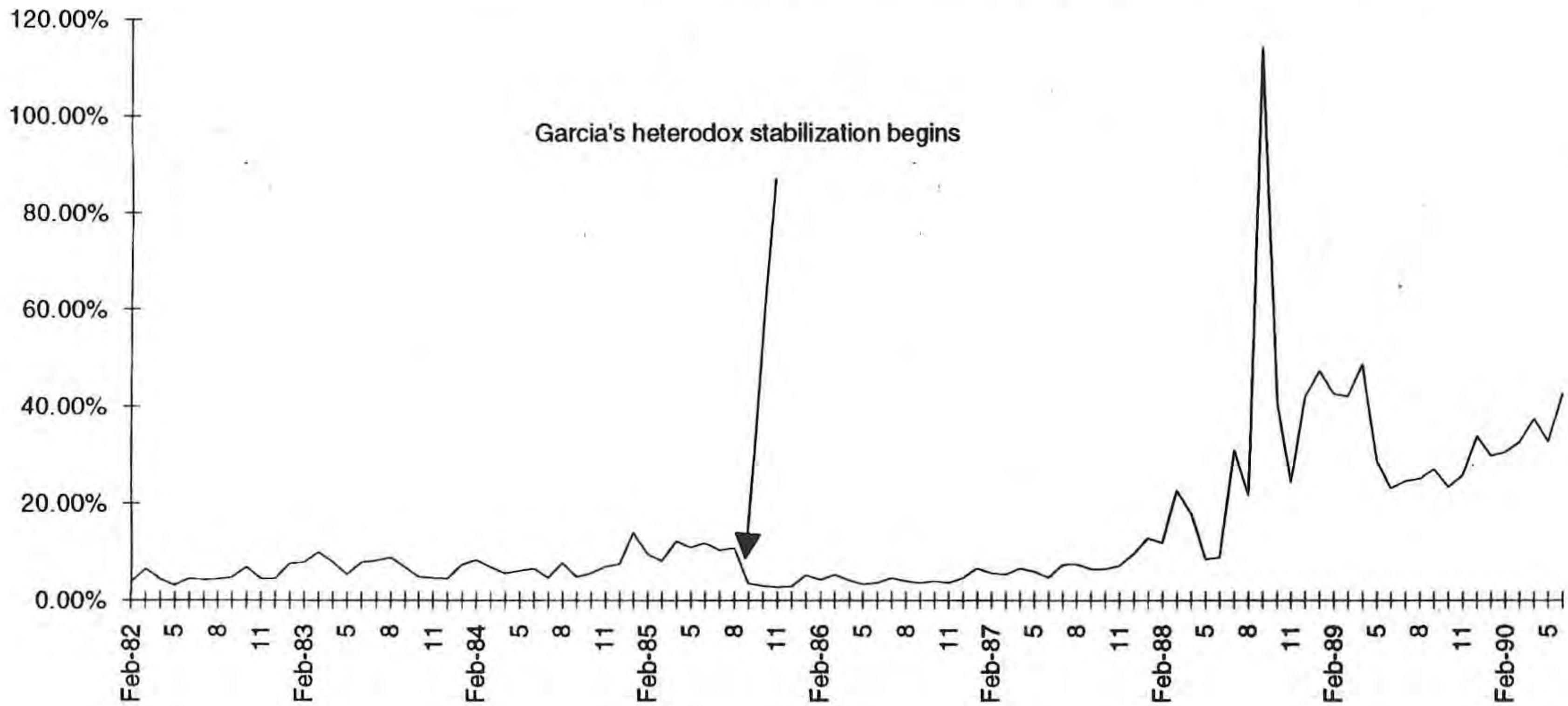


Figure 6
Argentina: Monthly Real M2 Growth 1971:1-1984:2

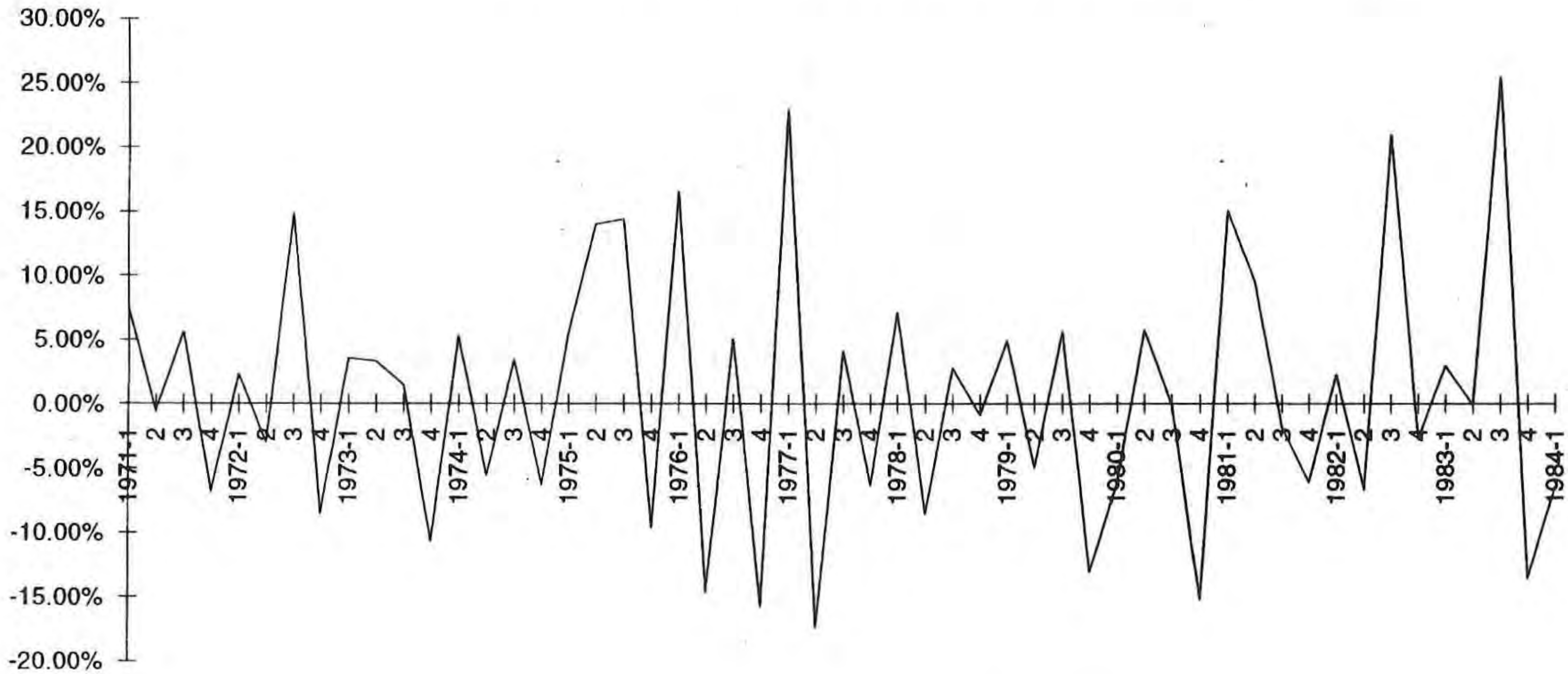


Figure 7
Bolivia: Real M2 Growth 1980:2-1990:9

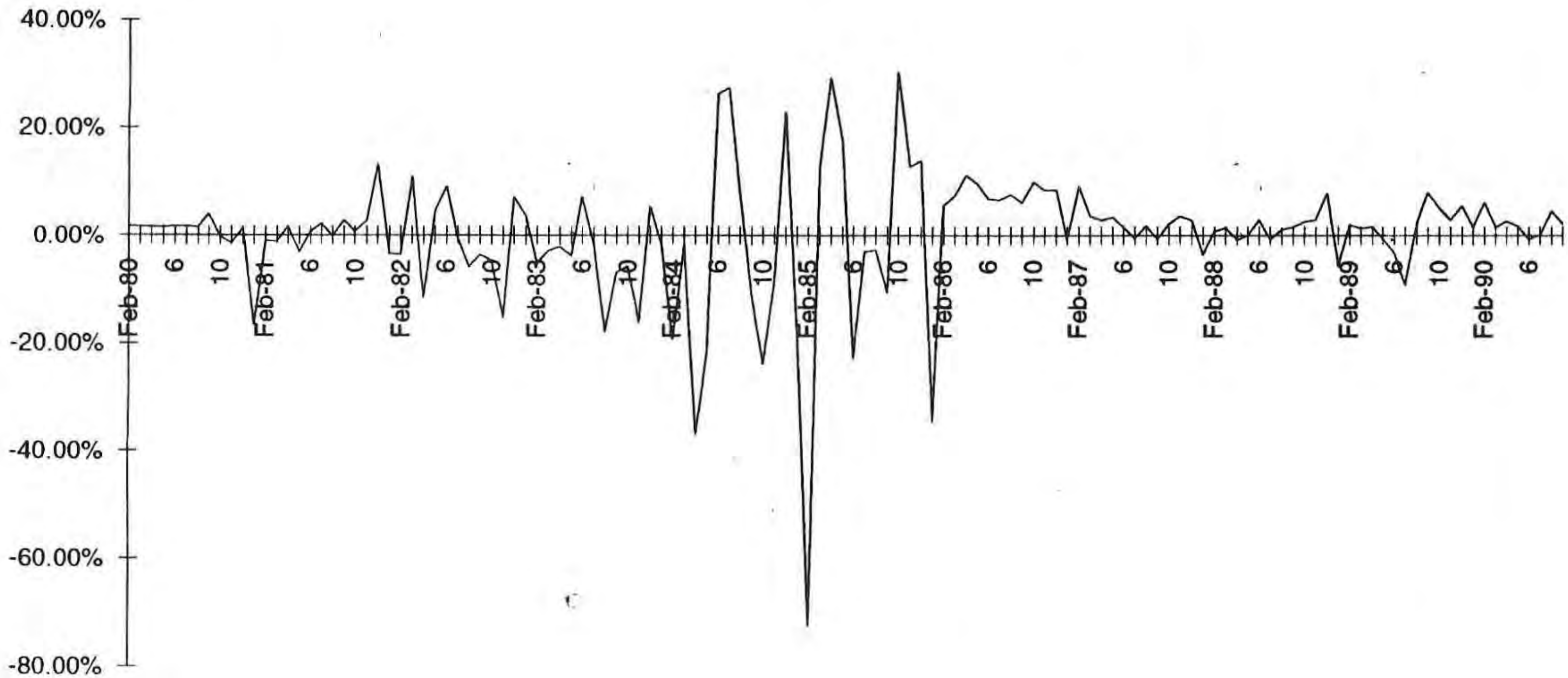


Figure 8
Brazil: Monthly Real M1 Growth 1974:2-1985:12

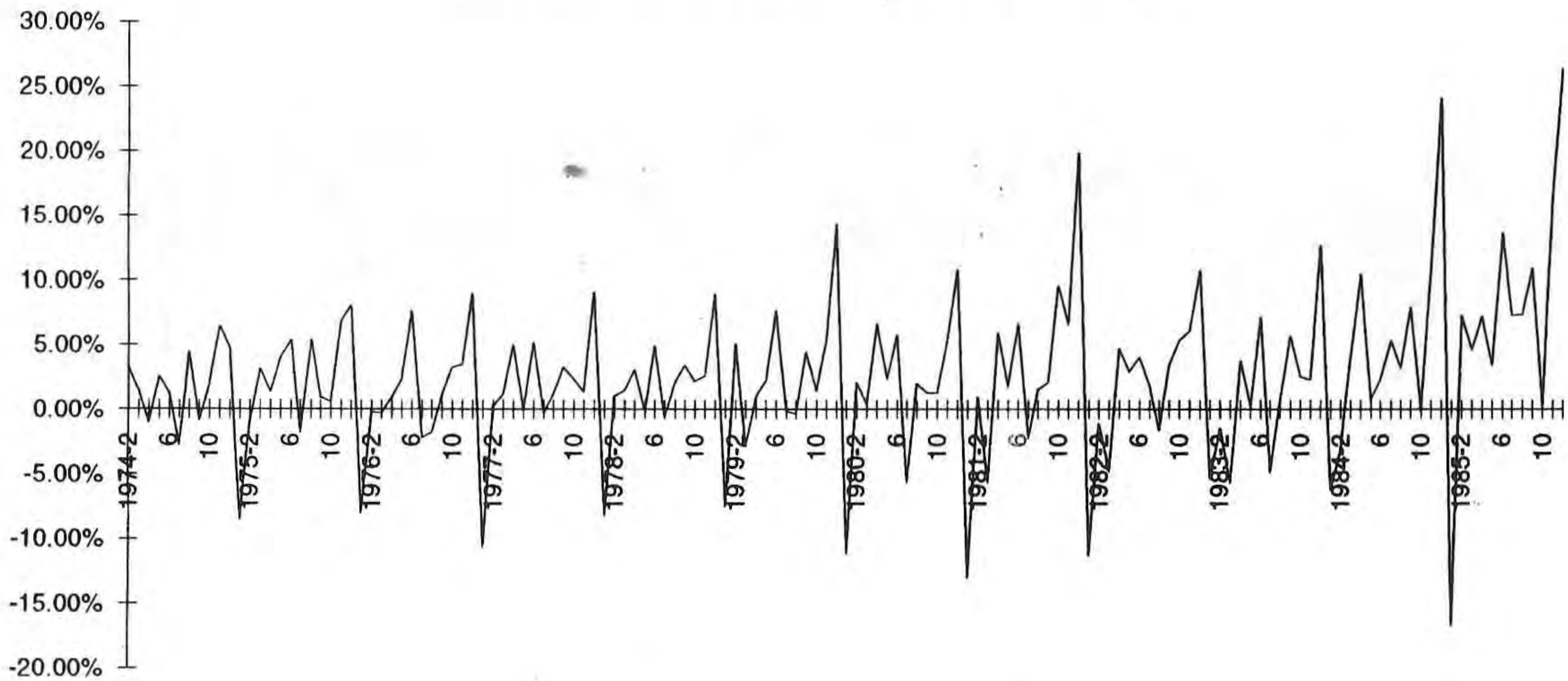


Figure 9
Mexico: Real Monthly M1 Growth 1960:2 to 1989:9

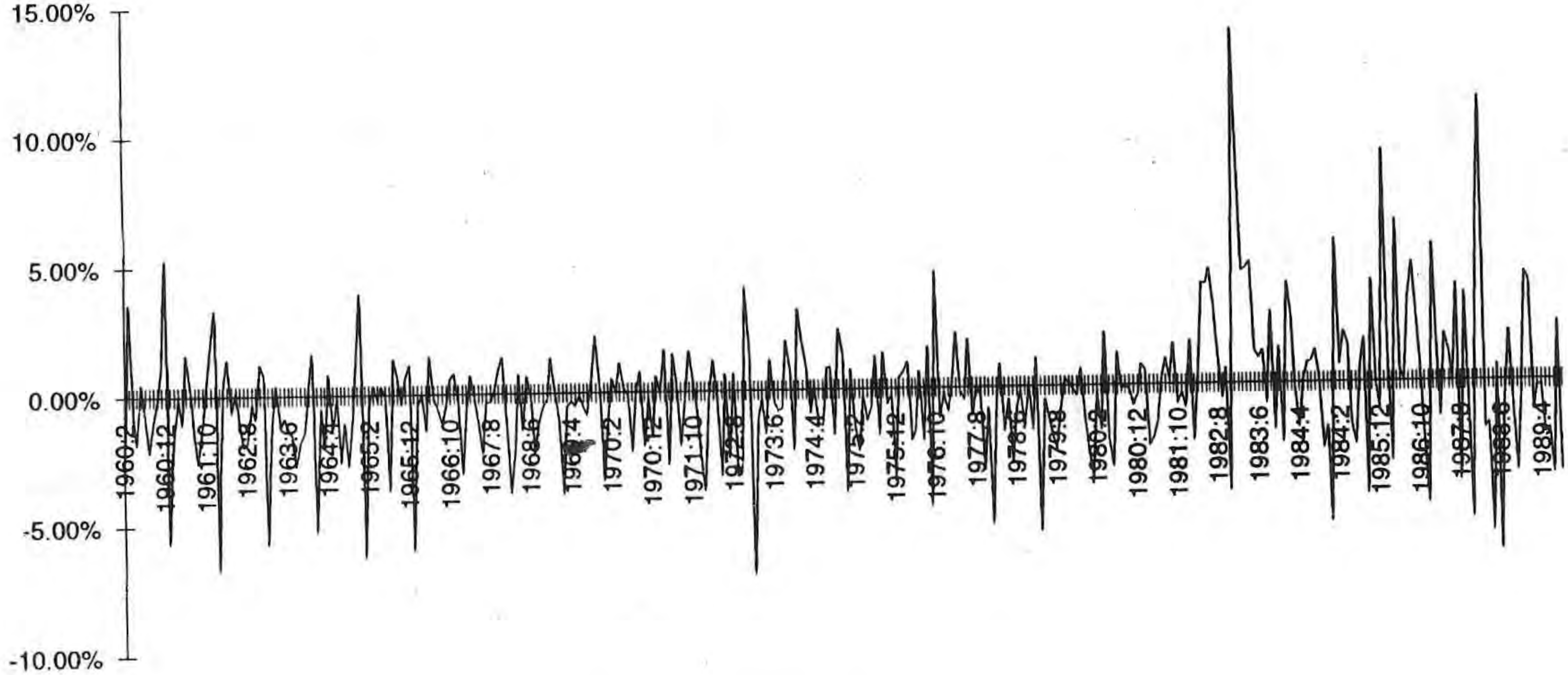


Figure 10
Peru: Real Monthly M2 Growth 1982:2-1990:6

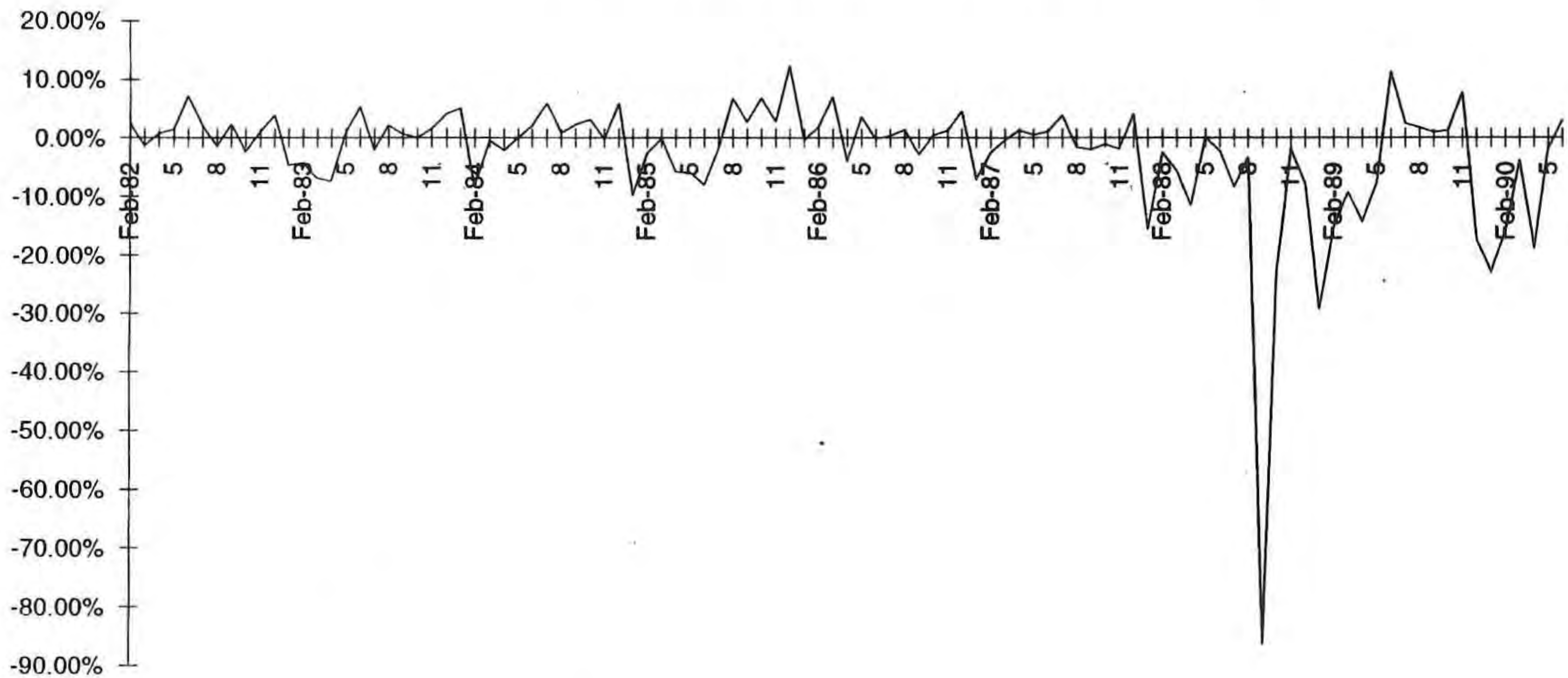


Figure A1
Argentina: CUSUM 1973:1-1984:2

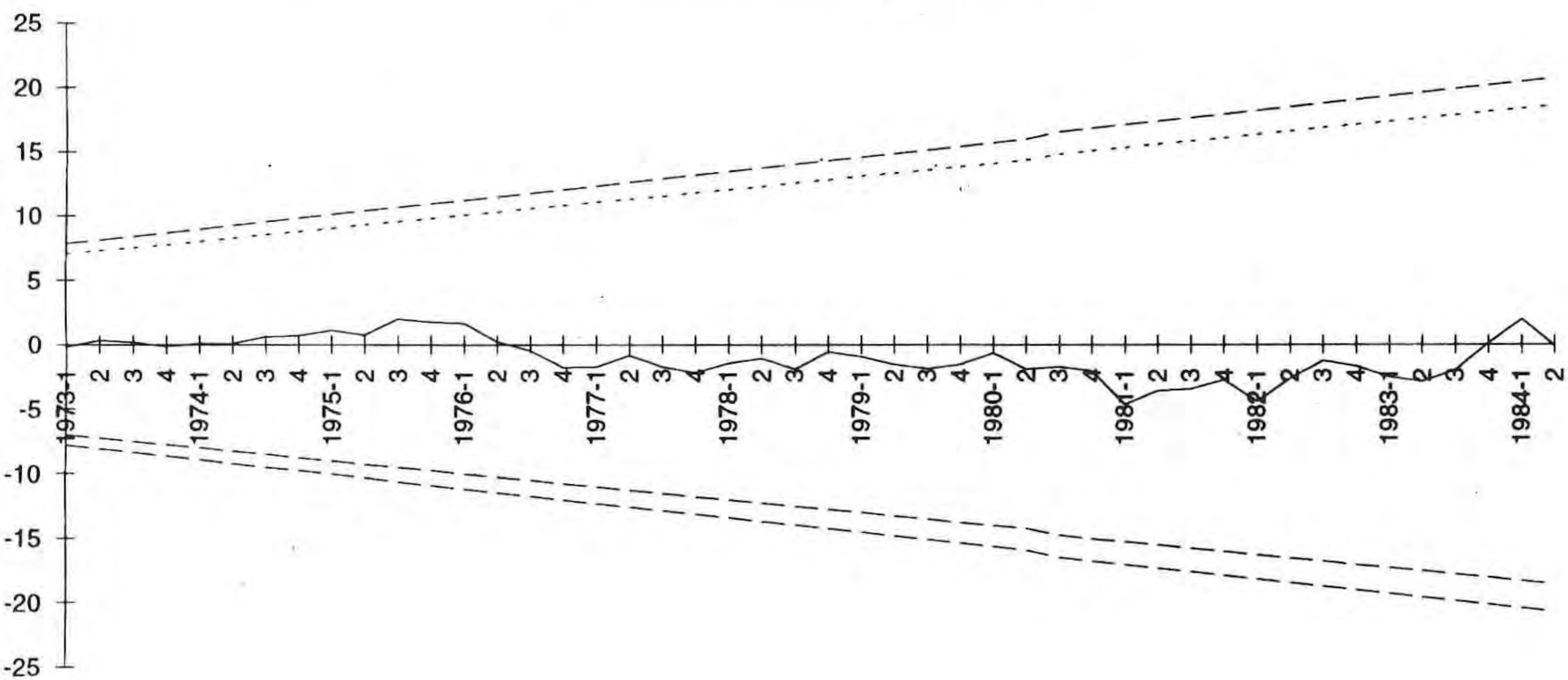


Figure A3
Brazil: CUSUM 1974:8-1985:12

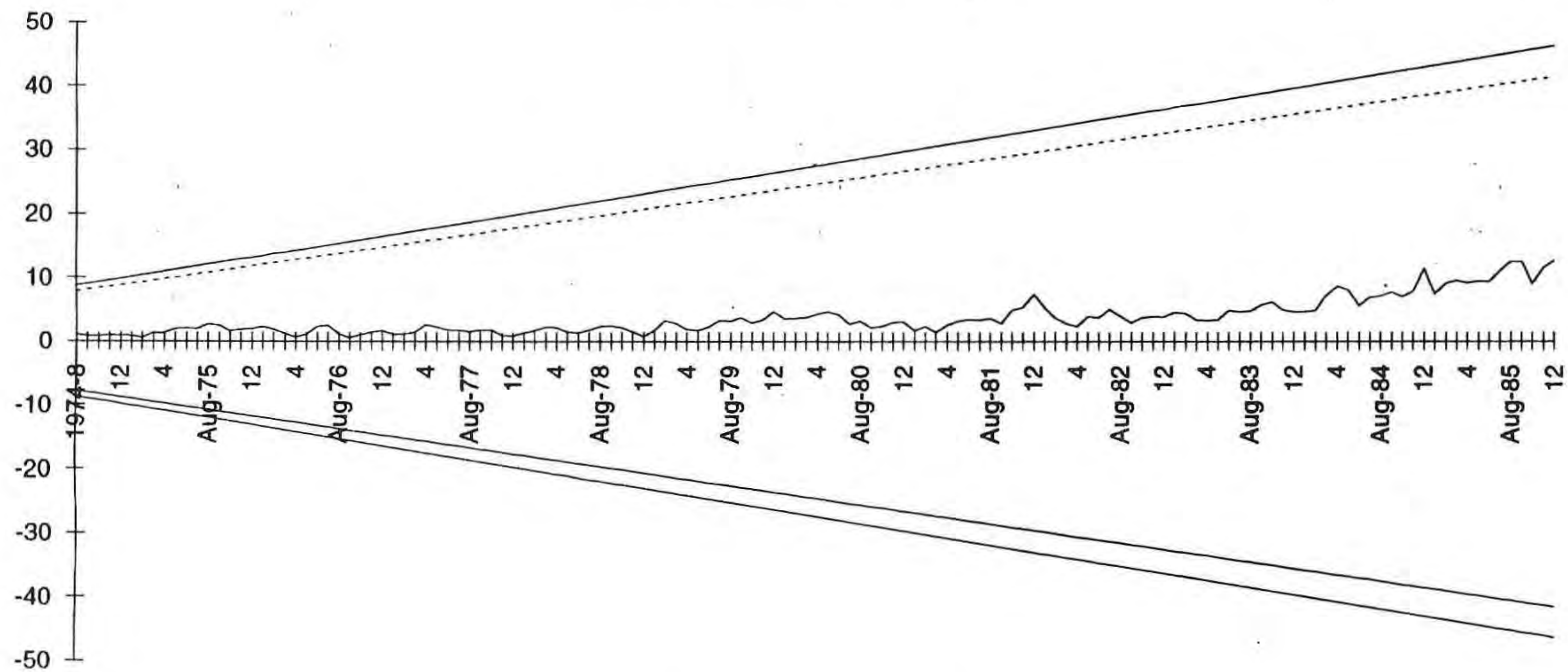
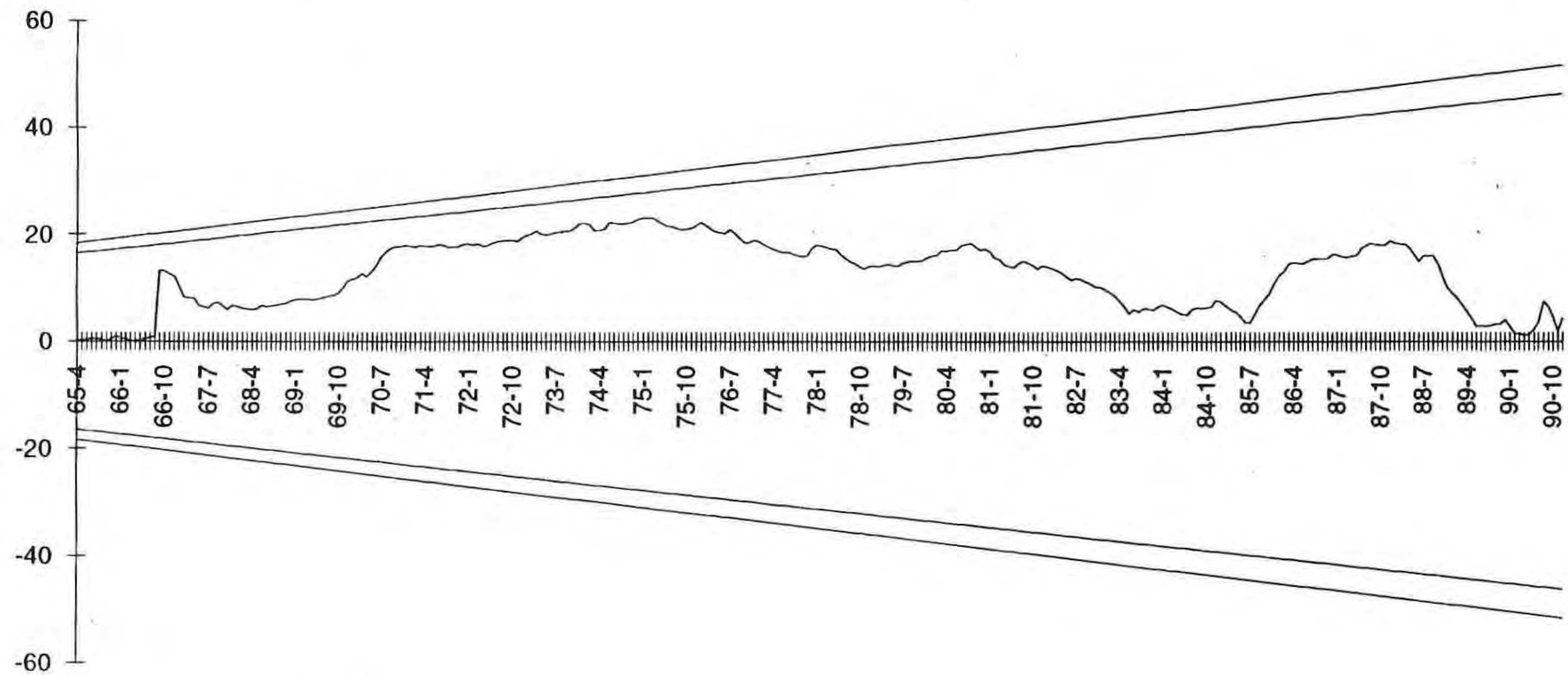


Figure A5
Peru: CUSUM 1964:1-1990:12



NOTES ON REVISIONS TO COMMENTS ON FELIZ AND WELCH

REFEREE #1

This referee noted the limitations of our results after succinctly outlining the main results of the paper. These limitations have to do with the fact that we did not have a well-defined alternative hypothesis, e.g., inertial inflation. We did not go back and address this issue directly by presenting test of the inertial proposition because Ana Novaes recently did so in a recent **JDE** article. Because many of our statistical tests provided results similar to hers, we deferred to her study in the introduction (pp.4-5) and in the conclusion (p.22). Although we fully agree with the referee's interpretation in paragraph three, we felt our results buttressed by Novaes rejecting inertial sources of inflation for Brazil.

Further, we went back and compared our results with earlier ones that found money demand instability (Cardoso 1983, Gerlach and Simone 1985, and Rossi 1985) and with others that found money demand stability but either ignored tests of rational expectations (Calomiris and Domowitz 1989) or rejected them (Phylaktis and Taylor 1993). This helped us articulate a clearer interpretation of our findings as the referee states in the third paragraph.

Finally, we hopefully provided a more intuitive interpretation of the cross equation restrictions and Granger causality implied by these models. The cross equation restrictions test the rational expectation's tenet that individuals do not make systematic errors in forecasting inflation (pp.4 and 10). The Granger causality tests follow from the assumption under rational expectations that agents use all available information - perhaps more than the econometrician. These are complementary to the cross equation restrictions.

REFEREE #2

Since this referee numbered his comments, I will address them accordingly.

1. This comment is similar to some of referee #1's comments about how we interpret of our results. Not only did we find cointegration between money growth and inflation but we also found that the model does not reject restrictions based upon rational expectations. Also, like referee #1, referee #2 worries about the alternative hypothesis. Hopefully, the reworked introduction and conclusion and the comparisons with other studies satisfy these concerns.

2. The referee has concerns about the unit root assumed in output. To get this, on p.5, we think of output as permanent income that should follow a unit root with drift. The justification for this comes from optimizing models. But we do not need this assumption as noted in footnote 9 on page 6. Evidence from developing countries reported by Basu and McLeod (1992) shows that real GDP does in fact have a unit root and that temporary shocks such as terms of trade changes generate highly persistent effects on real GDP. They justify their findings with an endogenous growth Solow-Swan model. Further, the

study by Mocan (1994) casts some doubt on the results that find no unit root in U.S. GNP.

If GDP does not have a unit root, we could have concentrated only on money demand. But we still feel comfortable with the more general approach. Subsequent to our finishing the paper, we read the Phylaktis and Taylor study. We appreciate the referee pointing this study out. Our study is more general in that we use a more fully specified model, e.g., the production side, while their study mainly looks at the demand for money. Further, our findings are not only consistent with a Cagan money demand function, but also with rational expectations -contrary to their findings - and a new classical view of the production side of the economy - not included in their study. We outline these points in footnote 6, p. 4.

Finally, I have amended the paragraph on bubbles according to the referee's comments.

SERIE DOCUMENTOS DE TRABAJO

The following working papers from recent year are still available upon request from:

Rocío Contreras,
Centro de Documentación, Centro de Estudios Económicos, El Colegio de México A.C., Camino al Ajusco # 20 C.P. 01000 México, D.F.

90/I Ize, Alain. "Trade liberalization, stabilization, and growth: some notes on the mexican experience."

90/II Sandoval Musi, Alfredo. "Construction of new monetary aggregates: the case of Mexico."

90/III Fernández, Oscar. "Algunas notas sobre los modelos de Kalecki del ciclo económico."

90/IV Sobarzo, Horacio E. "A consolidated social accounting matrix for input-output analysis."

90/V Urzúa, Carlos M. "El déficit del sector público y la política fiscal en México, 1980 - 1989."

90/VI Romero, José. "Desarrollos recientes en la teoría económica de la unión aduanera."

90/VII García Rocha, Adalberto. "Note on mexican economic development and income distribution."

90/VIII García Rocha, Adalberto. "Distributive effects of financial policies in Mexico."

90/IX Mercado, Alfonso and Taeko Taniura "The mexican automotive export growth: favorable factors, obstacles and policy requirements."

91/I Urzúa, Carlos M. "Resuelve: a Gauss program to solve applied equilibrium and disequilibrium models."

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91/III Young, Leslie and José Romero. "A dynamic dual model of the North American free trade agreement."

91/IV Yúnez-Naude, Antonio. "Hacia un tratado de libre comercio norteamericano; efectos en los sectores agropecuarios y alimenticios de México."

91/V Esquivel, Hernández Gerardo. "Comercio intraindustrial México-Estados Unidos."

91/VI Márquez, Colín Graciela. "Concentración y estrategias de crecimiento industrial."

92/I Twomey, J. Michael. "Macroeconomic effects of trade liberalization in Canada and Mexico."

92/II Twomey, J. Michael. "Multinational corporations in North America: Free trade intersections."

92/III Izaguirre Navarro, Felipe A. "Un estudio empírico sobre solvencia del sector público: El caso de México."

92/IV Gollás, Manuel y Oscar Fernández. "El subempleo sectorial en México."

92/V Calderón Madrid, Angel. "The dynamics of real exchange rate and financial assets of privately financed current account deficits"

92/VI Esquivel Hernández, Gerardo. "Política comercial bajo competencia imperfecta: Ejercicio de simulación para la industria cervecera mexicana."

93/I Fernández, Jorge. "Debt and incentives in a dynamic context."

93/II Fernández, Jorge. "Voluntary debt reduction under asymmetric information."

93/III Castañeda, Alejandro. "Capital accumulation games."

93/IV Castañeda, Alejandro. "Market structure and innovation a survey of patent races."

93/V Sempere, Jaime. "Limits to the third theorem of welfare economics."

93/VI Sempere, Jaime. "Potential gains from market integration with individual non-convexities."

93/VII Castañeda, Alejandro. "Dynamic price competition in inflationary environments with fixed costs of adjustment."

93/VIII Sempere, Jaime. "On the limits to income redistribution with poll subsidies and commodity taxation."

93/IX Sempere, Jaime. "Potential gains from integration of incomplete markets."

93/X Urzúa, Carlos M. "Tax reform and macroeconomic policy in Mexico."

93/XI Calderón, Angel. "A stock-flow dynamic analysis of the response of current account deficits and GDP to fiscal shocks."

93/XII Calderón, Angel. "Ahorro privado y riqueza financiera neta de los particulares y de las empresas en México."

93/XIII Calderón, Angel. "Política fiscal en México."

93/XIV Calderón, Angel. "Long-run effects of fiscal policy on the real levels of exchange rate and GDP."

93/XV Castañeda, Alejandro. "On the invariance of market innovation to the number of firms. The role of the timing of innovation."

93/XVI Romero, José y Antonio Yúnez. "Cambios en la política de subsidios: sus efectos sobre el sector agropecuario."

94/I Székely, Miguel. "Cambios en la pobreza y la desigualdad en México durante el proceso de ajuste y estabilización".

94/II Calderón, Angel. "Fiscal policy, private savings and current account deficits in Mexico".

94/III Sobarzo, Horacio. "Interactions between trade and tax reform in Mexico: Some general equilibrium results".

94/IV Urzúa, Carlos. "An appraisal of recent tax reforms in Mexico". (Corrected and enlarged version of DT. Núm. X-1993)

94/V Urzúa, Carlos. "Privatization and fiscal reforms in Eastern Europe: Some lessons from Latin America".

94/VI Feliz, Raúl. "Terms of trade and labour supply: A revision of the Laursen-Metzler effect".

94/VII Feliz, Raúl and John H. Welch. "Cointegration and tests of a classical model of inflation in Argentina, Bolivia, Brazil, Mexico, and Peru".