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An Approach for Testing Money Supply Exogeneity in Brazil Mixing Kalman Filter and Cointegration Procedures

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Abstract

Using monthly data spanning from 1964.04 to 1986.02, we examine the extent of money supply exogeneity. The implemented tests investigated the plausibility of classical hypotheses. We employed Kalman Filter procedures to estimate unobserved components and extracting outliers, Johansen & Engle-Granger's cointegration procedures and the bootstrap approach. The results are robust to the choice of Granger-causality tests. From past studies, we argued that the real rate of interest did cause, in the Granger sense, the bond stock and that the demand for bonds was very sensitive to interest rate variations. This implies that the monetary authority was able to perform indirect monetary control through the open market transactions. The results show that seigniorage collection was a white noise and econometrically independent from the inflation rate. Money creation and the inflation rate were cointegrated. We found that money growth was weakly exogenous for the parameter of interest in the conditional model of inflation, but the reverse is not true for the inflation. Moreover, Granger's causal relation between them was unidirectional from money to inflation. Therefore, money growth is strongly exogenous concerning the inflation rate. Thus, Cagan's adaptative model is not rational, and inflationary expectations were formed adaptatively. These empirical findings are sharply different from many previous results. Our main contribution is having demonstrated that the monetary supply was exogenous with respect to the inflation rate, that the monetary authority had enough independence to execute an active monetary policy, and that the inflation expectations followed the adaptative formation rule.

Resumo

Usando dados mensais de 1964.4 a 1986.2, examinamos o grau da exogeneidade da oferta de moeda. Os testes implementados investigaram a plausibilidade das hipóteses clássicas. Empregamos métodos de espaço de estado, os procedimentos de cointegração de Johansen e Engle-Granger, processos GARCH e a abordagem de bootstrap. Os resultados são robustos aos tipos de testes de causalidade de Granger. De estudos passados, argumentamos que a taxa real de juros causava a dívida pública no sentido de Granger e que a demanda de títulos era sensível às variações da taxa de juros. Isto implicava que a autoridade monetária era capaz de realizar indiretamente o controle monetário através das operações de mercado aberto. Os resultados mostraram que a coleta de senhoriagem se comportava como um ruído branco e era econometricamente independente da taxa de inflação. A expansão monetária e a taxa de inflação eram cointegradas. Os testes indicaram que a expansão monetária era fracamente exógena para os parâmetros de interesse no modelo condicional da inflação, mas o inverso não é valido para a taxa de inflação. Ademais, a relação de causalidade de Granger entre essas variáveis era unidirecional da expansão monetária para a inflação. Consequentemente, o crescimento monetário era fortemente exógeno no que concerne à taxa de inflação. Logo o modelo adaptativo de Cagan não é racional, e as expectativas inflacionárias eram formadas com a regra adaptativa. Esses resultados empíricos são inteiramente divergentes de estudos anteriores. Nossa principal contribuição é ter demonstrado que a oferta monetária era exógena com respeito à taxa de inflação, que a autoridade monetária tinha independência suficiente para executar uma política monetária ativa e que as expectativas inflacionárias seguiam a regra de formação adaptativa.

Keywords: Kalman filtering, cointegration procedures, bootstrap, econometric modeling, inflation, money supply, monetary policy.
JEL: C32, C52, E31, E51, E52.

1. Introduction

From 1964 to 1985, the inflation stabilization policy was centered on aggregate demand management and wage, exchange rate, public prices and concentrated sectors price controls. The fact that during these years the economic policy was conducted in a seemingly "orthodox way" may induce us to wonder about the monetary policy behavior. Are there reasons to support the claim that the money supply was exogenously determined concerning the inflation rate, that is, did the monetary authority follow a rule of money creation that did not accommodate the movements in the inflation rate?

The purpose of this paper is to confirm that the monetary supply was exogenously determined from 1964.04 to 1986.02. That is, we claim that the monetary policy was active, and within Cagan's (1956) and Sargent & Wallace's (1973) models show that the inflation expectation formation rule followed an adaptative scheme. We thus postulate that the monetary rule was executed independently and the rule that guided the monetary execution was exogenous with respect to both the 'model in question' and the inflation rate (for a detailed description of the monetary policy of the period, see Simonsen, 1985 and Cerqueira 2006b).

We intend to support this assumption with three arguments. First, the monetary authority conducted the open market policy increasing the real interest rate in order to stimulate the demand for bonds. Second, even if during part of this period the seigniorage collection remained constant as a share of GDP, the government succeeded in keeping its fraction of revenues by reducing the monetary base multiplier[1]. Moreover, the seigniorage-GDP ratio followed a white noise process and was therefore independent of the inflation rate. Third, the existing causality between money growth and inflation was unidirectional from the former to the latter.

Nevertheless, the belief that the money supply was passive during most of the 1964.04-198.06 period is largely spread among many Brazilian economists. One probable rationale supporting this belief is the hypothesis of rational expectations. If the demand function for real balances follows Cagan's form, the solution for the current inflation is a function of the rate of expected money creation, excluding the possibility of rational explosive bubbles. In this case, money supply is endogenous.

An alternative argument is based on Sargent & Wallace's (1973) scheme derived from Cagan's model under the hypotheses of adaptive expectations and a monetary rule, which depend on past inflation rates. This is a model in which the adaptive mechanism is rational. In Sargent &

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[1] Since M1=kB, where ***B*** is the monetary base and ***k*** the multiplier, the inflation tax-GDP ratio is given by IT=$\pi(B/PY)=\pi(M1/PY)(1/K)$, where ***P*** is the price index and π the inflation rate. This means that the government collects 1/k of the produced inflation tax by the real money balances. The difference (1-1/k) represents the inflationary transfer from the private sector to the bank system.

Wallace's model, the best way to forecast the subsequent rates of money creation is by extrapolating lagged rates of inflation. This in turn implies that inflation itself is best predicted by extrapolating past inflation rate. So both money creation and inflation are best forecast by extrapolating current and lagged rates of inflation. Lagged rates of money creation add nothing to predictions formed in this way. In this model, past values of inflation influence money creation but the converse is not true; thus, money supply is passive.

An essential element in this argument is the hypothesized feedback that occurs from expected inflation to money creation. This feedback emerges due to the government's attempt to finance a roughly constant rate of operational deficit by money creation. In this sense, this is also a description of Bruno & Fischer's (1990) version of Cagan's model; in this description monetary expansion is endogenously determined by expected inflation, given a constant level of seigniorage.

However, if the monetary expansion follows a purely autoregressive process or a white noise process, then under adaptive expectations the monetary expansion and the money supply are exogenous with respect to the inflation rate. In this system money creation influences current and subsequent inflation rates; but given lagged rates of money creation, past inflation rates exert no influence on money creation. The system is one in which money creation causes inflation, in Granger's sense, whereas inflation does not cause money creation. This is a model in which adaptive expectations are not rational. In the model, feedback occurs from the expected inflation to the current inflation rate, a feedback that emerges from an autonomous increase in the monetary expansion. Therefore, under adaptive expectations the money supply passivity is a consequence of the monetary rule followed by the monetary authority.

Brazil's long experience with high inflation rates gave rise to an efficient indexation system that protected agents from the effects of inflation. Even if the indexation rules did not fairly contemplate the agents, one cannot deny that such rules prevented the ever-rising inflation to degenerate into public panic, speculative run and open hyperinflation process. Furthermore, the indexation rules were developed little by little along the seventies and eighties simultaneously with an increasing inflation rate. The rigidity of the price system was then introduced gradually, which augmented the inflation inertia though did not destroy the inflationary memory. Thus, as the economy indexation degree rose, the inflation rate became inertial, i.e., its present values began to depend on its past ones. These arguments supposedly explain why the agents had adaptive expectations about the inflation in the period. Therefore, Brazil's experience over the period did not provide evidence of expectations being formed rationally. One can therefore argue that the monetary policy followed a rule independently of the inflation rate. This assertion is tested in section 3.

The assumption of endogenous money growth cannot be supported by empirical evidence. Surprisingly, some authors found a unidirectional relation between inflation and money creation. We can suppose that their results came about because of their use of lower frequency data (quarterly). The resulting information loss may have distorted the results of causality tests. They relied on the Ljung-Box test and the related correlogram for detecting serial correlation and setting the lag length in autoregressive models. It is well known that the portmanteau test may have very low power[2] (since the significant correlation at one lag may be diluted by insignificant correlation at other lags) in the detection of specific important departures from the assumed model. It is therefore unwise to rely exclusively on this test in checking model adequacy. However, it can be valuable when used with other tests. The Breusch-Godfrey Lagrange multiplier test is a common complement to the Q-test (Granger and Newbold, 1986). By carrying out both tests and using monthly data from 1964.04 to 1986.02[3], we achieved results that contrast sharply with previous findings of other Brazilian authors. This might explain why in applying the same causality tests different conclusions emerged.

Why in that period the monetary expansion increased uninterruptedly, with a behavior near to an I(1) process, remains an unresponded question. We conjecture that the answer may be found in the chronic public deficit that has been partially financed by issuing bonds. This produced an ever-increasing financial component[4]. If the rule governing the monetary authority was to achieve debt sustainability, then it was urgent to support the deficit financing by increasing the money creation. This led to a pegging of the inflation rate. Thus, the monetary authority chose, or was compelled to choose, the deficit inflationary

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[2] Reversibly, if one chooses too small a lag, the test may not detect serial correlation at high-order lags.

[3] Since the pioneer paper of Working (1960), it has been a well-known fact that temporal aggregation has statistical implications on the time series properties. Rossana and Seater (1995) point out that the temporal aggregation causes substantial loss of information about the underlying data processes. They argue that non-aggregated data are governed by complex time series processes with much low-frequency cyclical variation, whereas aggregated data are governed by simpler processes without the same rich cyclical pattern. Cycles of much more than a year's duration in the monthly data are reduced when the data are aggregated. Moreover, the aggregated data show more long-run persistence than the underlying disaggregated data. For all the above reasons, we decided to use monthly data to compute our results, which span the whole period. Thus, for quarterly series or for those series with missing data we interpolated or simulated them (the applied procedures are reported in the appendices). So we preserved or extended the available information of each series with the employment of higher frequency data. In general, such a procedure leads to more accurate estimations (Granger and Newbold, 1986). Furthermore, we followed a judicious criterion for setting the lag length of autoregressive process and rigidly observed the classical econometric hypotheses.

[4] We estimate that between 1966 and 1985 the proportion of the public deficit due to the nominal debt-service increased from 34.8% to 83.1%, while the real service decreased from 7.0% to 1.7%. At the same time, the nominal deficit as a GDP fraction rose from 2.9% to 17.2%, and the operational deficit from 2.1% to 3.2%.

financing in order to sustain the debt. There was, then, a choice of economic policy.

This paper is organized as follows. In section 2 we describe in general lines the results presented in Cerqueira (2006a) on Brazilian monetary regime and the relation between real interest rate and the public debt. In section 3 we present the paper's theoretical framework, and provide statistical procedure for determining the seigniorage collection stochastic process. Section 4 offers an econometric study about the long-run relationship between money creation and inflation rate. And in section 5 we present our conclusions. In appendix A we provide an analysis of the integration order of the two series. In appendices B and C we respectively report the results of the cointegration tests and estimate the series of adaptatively formed inflationary expectations.

2. The Brazilian monetary regime

In Brazil, the monetary regime was different from the usual standards, in which the monetary control is realized through open market transactions, discount loans or reserve requirements. Here the monetary policy was conducted indirectly, through a particular form of open market transactions. Besides this was an economy with a consistently increasing inflation.

In the Brazilian monetary regime, firstly the Treasury directly financed itself through the Central Bank. Secondly, public bonds were not sold to the final takers, but rather to financial institutions, which financed themselves through overnight deposits from the private sector. At the same time, the Central Bank informally gave liquidity to the excess of bonds over these deposits by means of repurchase agreements. If in a primary auction the financial intermediaries did not succeed in getting a permanent and equal increase in their funding, then they could resell their holdings of excess bonds to the Central Bank. The repurchase agreements were necessary because free reserves were costly to the banks. If the banks had to wait for government securities to grow mature and the Central Bank did not provide (inexpensive) liquidity to the system, banks would either have had to hold a much larger volume of free reserves (within an inflationary environment), or have resorted more often to the discount loans, which would have been unbearably costly to them.

The main consequence of this procedure was the elimination of the open market operations as an instrument of monetary policy. The money supply was controlled indirectly by increasing the interest rate to expand the demand for bonds. This procedure was efficient from 1966 to 1985, as demonstrated in Cerqueira (2006a). The author's results show that the causality in Granger's sense was unidirectional from the real interest rates to the ratio public debt/GDP from 1974 to 1985. Thus, one could assume that the real interest rates were strictly exogenous as to the demand for bonds; see Sargent (1987). One can thus conclude that the monetary authority altered the real interest rates to cause changes in the demand for bonds. Besides, once the public debt is a non-monetary liability, this mechanism operated as a money supply control instrument. Finally, the demand for bonds was elastic to the real interest rates, and the overnight interest rates were enough to encourage bonds sales. These empirical facts suggest an active behavior of the monetary authority instead of a passive monetary policy and an endogenous money supply.

3. Seigniorage and Inflation

Let us consider the following continuous time version of Cagan's model:

$$m \equiv \frac{M}{PY} = ce^{-\alpha\pi_e}\text{, } c > 0 \text{ e } \alpha > 0 \text{ (money demand equation)} \quad (1)$$

$$\overset{\circ}{\pi}_e = \beta(\pi - \pi_e)\text{, } \beta > 0 \text{ (expectation rule equation),} \quad (2)$$

where M is nominal money demand, P is index price, Y the GDP, c is a constant which captures nominal shocks and financial innovations changes, π_e is the expected inflation rate, α is the semi-elasticity of the money demand with regard to the expected inflation, and β is the inverse of inflationary memory (the bigger β, the smaller the effect of past inflation on the present inflation expectations). We assume a constant growth rate and a constant real interest rate. For a given level of exogenous money growth μ, the seigniorage flow is given by:

$$S = \frac{\overset{\circ}{M}}{PY} = \frac{\overset{\circ}{M}}{M}\frac{M}{PY} = \mu ce^{-\alpha\pi_e} \quad (3)$$

By making some operations we get to the expected inflation rate dynamics equation:

$$\dot{\pi}^e = \frac{b}{1-ab}(\mu - \pi^e) \quad (4)$$

In steady state $\dot{\pi}^e$=0 and $\mu = \pi^e = \pi$ and the inflation tax equals the seigniorage. The seigniorage is maximized (S*=c/αe) when $\pi = 1/\alpha$. With a constant operational deficit at level $S = \overline{S}$, the monetary authority will react according to:

$$\mu = \frac{\overline{S}}{ce^{-\alpha\pi_e}}\text{ .} \quad (5)$$

The monetary expansion rate is increasing with the expected level of inflation and thus is passive. Then, a reduction in the constant term c, caused by a financial innovation, shifts down the reaction curve that increases the monetary expansion and the inflation rate – in the (π^e,μ) mapping –, which implies increases in monetary expansion and inflation expectations.

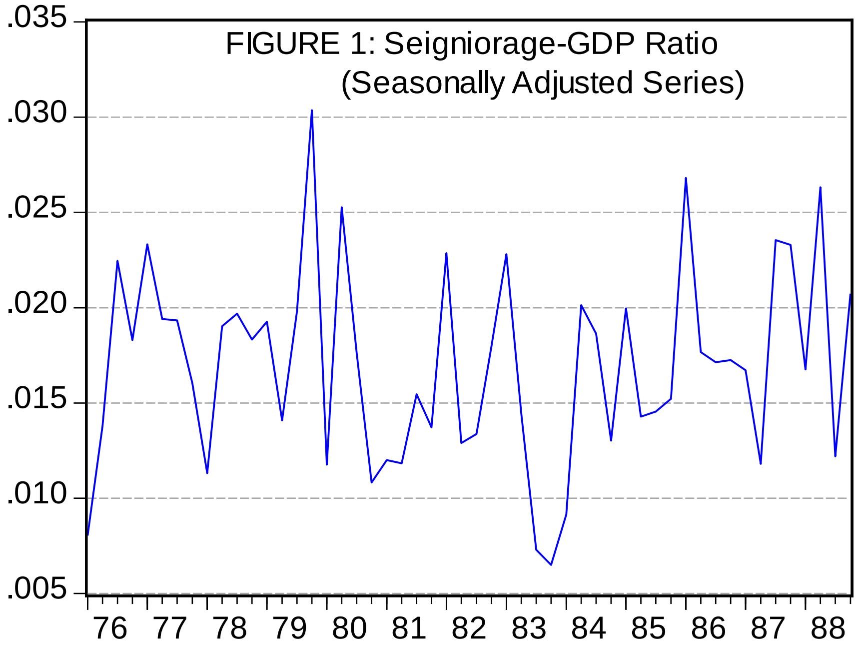
The postulation of a reaction curve as (5) and the existence of empirical evidence that the monetary authority has been following it are two different things. Such behaviour implies a passive money supply or caused in Granger's sense by lagged inflation rates, a popular hypothesis among Brazilian economists. This claim tends to be accepted due to the empirical evidence in Pastore (1994/95 and 1997) – an influential Brazilian economist –, who shows that the money growth is caused by the inflation, but the reverse is not true. Therefore the author concludes that the money supply was predominantly passive.

Intuition suggests that the Central Bank’s repurchase agreements tend to make the money supply endogenous. However, some reflection shows that the former facilitates the latter, though does not determine it. If the inflation rates grow due to a negative supply shock and the Central Bank’s goal is to keep the real interest rate constant, the Bank would automatically buy bonds through the repurchase agreements, sanctioning the price increase with a larger monetary base stock. The money supply was made endogenous or caused by the Central Bank’s rule to fix the real interest rate, and was but facilitated – or accelerated – by the repurchase agreements, not due to them per se. Money endogeneity runs independently of the monetary regime, provided that the Central Bank keeps the real interest rate constant.

Notwithstanding, if the Central Bank lets the real interest rate float to sell bonds, such causality will be eliminated or reduced. In the Brazilian monetary regime selling bonds would cause a simultaneous increase of the nominal and real interest rates, whereas in a conventional monetary regime the increase of interest rates would cause a bond demand raise, implying a more instantaneous and precise control of the monetary stock than in our case.

Cerqueira (2006a) showed the real interest rates floating not only appeared to be significant in the public bonds demand, but also caused the public debt. These empirical results contradict Pastore’s evidence, making us review his empirical analysis on the causality between the inflation rate and money growth. We shall start by analysing the behaviour of the seigniorage/GDP ratio series.

In 1976, the monthly inflation rate goes above 3% and from 1976.1 to 1988.4 the monetary base growth as GDP fraction did no present any significant change. We admit then that from this level the inflation acceleration would not have any impact over the government’s revenue with issuance base. Then the primary money growth can be taken as relatively constant and independent from the inflation rate. This leads us to presume that the seigniorage can be described as an erratic process similar to a white noise; see figure 1. This hypothesis is based on the increasing cost of holding money and on the process of fiat money substitution by other financial assets, which takes place from 1976 on; see Marques and Werlang (1989).



The seigniorage/GDP series has quarterly frequency as does the GDP. The real GDP is the one estimated in Cerqueira (2006c), by Kalman filter with benchmarking adjustment; see Durbin Koopman (2004). Once the seigniorage/GDP shows seasonality and some outliers, we decided firstly to extract these components. We estimated a stationary structural model using Kalman filter procedures with an AR(1) component, fixed seasonal dummies, intervention variables such as impulse dummies – 1986.2, 1986.3 and 1987.1 – and an irregular term. As required the smoothed disturbances model are approximately $N(0,\sigma^2)$[5]. The adjusted series is the sum of the AR(1) plus the irregular components (figure 1).

The statistical approach is that of submitting the adjusted seigniorage/GDP series to independence tests. Table 1 reports the unit root tests of Augmented Dickey-Fuller, Phillips-Perron and their modifications for the series level; see Maddala and Kim (2002). The ADF and DF-GLS tests were carried out with no lags, and the others with quadratic spectral kernel and size two window. All tests largely reject the null hypothesis of a unit root.

Table 1: Unit Root Tests Report

Test / Series	ADF	DF-GLS	PP	ERS-PO	Ng-Perron			
					MZ_{α}^{d}	MZ_{T}^{d}	MSB^{d}	MP_{T}^{d}
SY	-6.6891 (0.0000)	-4.5181	-6.6875 (0.0000)	2.7432*	-14.9169	2.6692	0.1789	1.8762

Note: The symbol (*) means rejection of the null at 5% significance level; the absence of symbols indicates rejection at 1%.

Figure 2 shows the autocorrelation functions, the spectrum – with window 6 –, the cumulative periodogram, the CUSUMSQ, the histogram and the QQ of the Seigniorage/GDP series. Ljung-Box test statistics takes the hypothesis accepts the null of no autocorrelation up to order 12 and 24 with p-values

[5] The estimated autoregressive coefficient is ϕ = 0.1838. The results are available upon request.

0.9883 and 0.2698, respectively. Spectral density is close to that of a stationary series like a white noise. The cumulative periodogram behaves likewise, and the associated Kolmogorov-Smirnov statistics accepts the null hypothesis that the series is close to a white noise at 5%[6]. Evidence shows that the money growth series over the GDP is an ergodic process close to a strong white noise[7]; see Hendry (1995).

The seigniorage/GDP series during the period is described by the following statistics: mean=0.017, std.dev.=0.005, skew.=0.218, kurt.=2.765. The normality tests statistics of Bera-Jarque and Hansen-Doornik have p-values equal to 0.766 and 0.776, respectively, pointing that the series is a Gaussian white noise. Assuming that the above results are correct, Seigniorage/GDP (SY) regression against a constant and the inflation rate (PI) must reach the following results: (i) constant significant at the level of the sample mean (0.017); (ii) non-significant inflation rate coefficient[8]; (iii) R^2 close to zero; (iv) residuals approximately NIID. These results are evidenced in table 2. With p-value 0.9879, we cannot reject the hypothesis that the equation constant term approximately equals the series' sample mean. Nor can we reject the hypothesis that the inflation rate coefficient is non-significant; and that R^2, likewise residuals behave according to expectations.

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[6] The computed statistics equals 0.1513, and the bilateral critical value at 5% is 0.15495.

[7] Indeed, considering Goldfeld-Quandt heteroskedasticity test statistics, h(17) = 0.90578 (*p-value* = 0.5796), one cannot reject the hypothesis that the series has constant non-conditional variance along the period; the LM ARCH(1) and ARCH(2) tests, with *p-values* respectively 0.7337 and 0.6412, and McLeod-Li(4) test with *p-value* = 0.5641, lead to the hypothesis that the series does not have autoregressive conditional variance.

[8] We get the quarterly inflation rate by accumulating geometrically each trimester's monthly data. Evidence that the quarterly inflation rate be I(1) is inconclusive, *i.e.*, the unit root tests do not allow agreement as for the diagnosis; ACFs and PACFs as well as the spectral density point at a strong serial correlation, with a first order autoregressive root around 0.65. For these reasons we used the série in levels. Besides, the regression with the inflation in first difference did not show any material difference in the results and in the reported diagnostic statistics.

FIGURE 2: Seigniorage-GDP Ratio Statistics

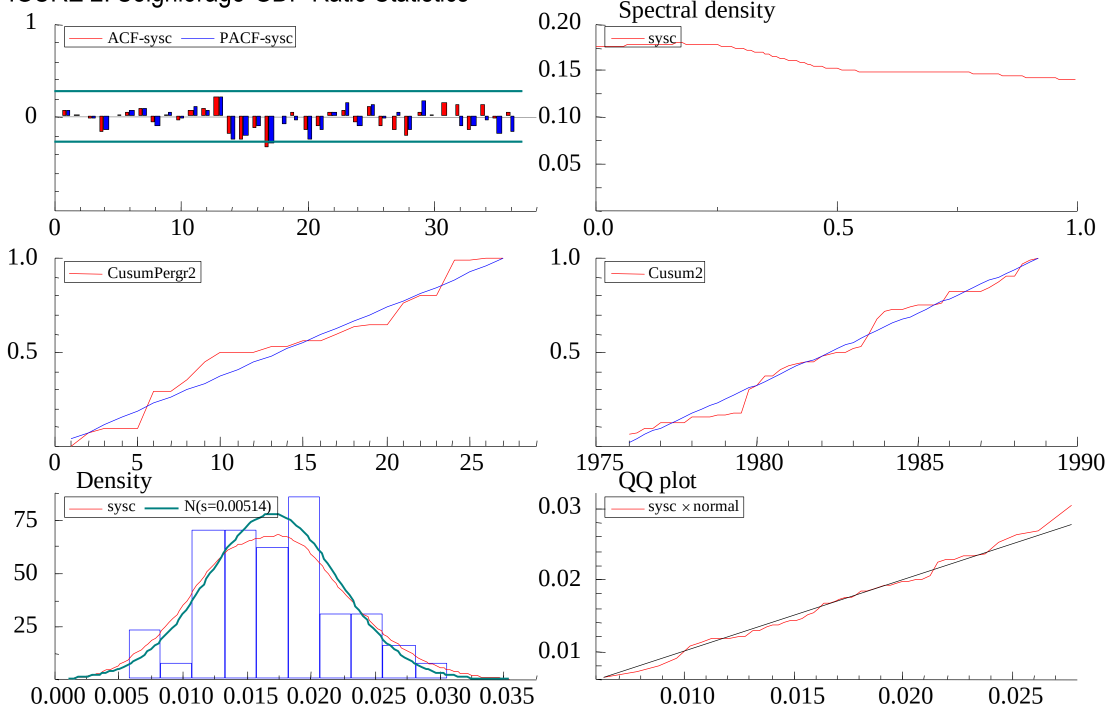


Table 2: OLS Regression SY Against PI

Variable	Coefficients	P-value			
C	0.0170	0.0000			
PI	-0.0002	0.9852			
N=52	R^2=0.000001	F=0.0004	SER=0.0052	DW=1.7914*	Q(12)=0.9883
Q(24)=0.2362	LM(1)=0.6246	LM(2)=0.8849	LM(3)=0.9674	LM(4)=0.8284	LM(6)=0.9484
WHITE=0.1230	ARCH(1)=0.7382	ARCH(4)=0.6427	$Q^2(4)$=0.5255	Sk=0.2198	Ek=0.2345
BJ=0.7641	BDS=0.9426				

* The hypothesis test in which DW = 2 is taken with p-value = 0.4520.

The above results confirm the hypothesis that from 1976.1 to 1988.4 the ratio seigniorage/GDP followed a white noise statistic process ~ (0.0170; 0.0052). <u>This means that, even having a finite average, its behavior could not be predicted due to the series' lack of memory</u>. Thus, seigniorage collection behaved as a random shock. It did not show any relation with money holdings contraction or with inflation rate rise. It can, therefore, be taken as independent from the inflation rate, which did not at all alter this government-financing source. Such conclusion provides empirical support to the hypothesis that the monetary policy was not passive in the period, once the inflation rate did not have any impact on causing seigniorage collection to rise of fall.

There was a public deficit permanently financed partly by money growth, which ensured the public debt sustainability (see Cerqueira, 2006a), not making the money supply endogenous. This means that there was a steady state level of public deficit financed by money issuing[9], which is coherent with Cagan's

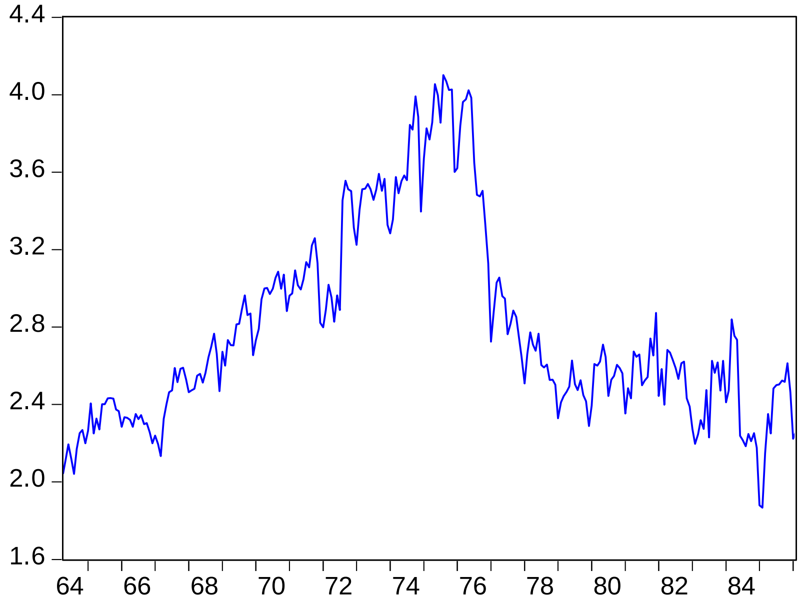
[9] In the period the inflation tax/GDP ratio mean lies around 1.7898%. This is similar to the unadjusted seigniorage/GDP series (2.1608%). The equality between them is marginally accepted with *p-value* 0.0588. However, the average difference between them, representing the real base variation over GDP, equals approximately zero, with *p-value* 0.7150. This is a strong

non-rational adaptative model. To complete the exogeneity hypothesis money supply proof, it is necessary to show that the money growth caused in Granger's sense the inflation rate, but inflation did not cause money growth (see Sargent and Wallace, 1993). This will be seen in the next section.

4. Money supply exogeneity

From 1976 to 1985 the host of financial innovations contracted the monetary demand that restricted the seigniorage collection. However, the policy to reduce the monetary base multiplier counterbalanced this effect;[10] see Cerqueira (1993, 2006ª and figure 3 below). Anyhow, money demand contraction increases observed and expected inflation rates according to Cagan's model. If the monetary authority is to keep the previous level of seigniorage collection, it will have to increase money creation, which will turn out to be endogenous. Besides, if the expected inflation is predicted by extrapolating lagged inflation rates, money issuing is, in Granger's sense, caused by the inflation rate. In this system, monetary expansion past values add nothing either to inflation predictions or to money growth expected rates. In this model, Cagan's adaptative scheme is rational; see Sargent and Wallace (1973).

Figure 3: Base Multiplier



On the other hand, money supply might be exogenous as for the inflation rate and error term of money demand function. In such case, money supply issuances influence current and future inflation rates, but past inflation values do not influence money supply. The system is such that money creation causes

hint that there is a steady state equilibrium.

[10] Whereas the reduction in the money demand constant term implied the decreasing of the required money stock by 43.9166%, from 1975 to 1984/85, the money multiplier fell by 39.3997%, which represented an increase of 65.0160% in the rate of inflation tax collected on the M1 stock. This more than compensated for the first effect; see Cerqueira (2006a).

inflation, though not the other way round. In this system, Cagan's adaptative scheme is not rational.

In section 2 we argue that monetary policy let the real interest rate float in order to stimulate demand for bonds. Since the debt stock kept by the Central Bank is a non-monetary liability, it was an instrument for monetary control. In section 3 we show that, even if the monetary authority was attempting to finance a roughly constant rate of public deficit by money creation, the seigniorage collection did not follow a path consistent with the endogenous money supply. Much on the contrary, seigniorage collection appeared to be a random variable and independent from inflation rate, which is likely to be related to the policy of reducing the base multiplier. This policy contributed to keep the effectiveness of inflationary tax collection, reinforcing the claim of an exogenous monetary policy.

We thus assert that a public deficit fixed rate permanently funded by money creation supported the sustainability of the public debt, as the exogeneity of the money supply was preserved. This points at the existence of a steady state public deficit financed by seigniorage, as in Cagan's adaptative model. It is necessary to prove that money supply caused inflation, whereas the other way round is not true.

The inflation rate and money supply series[11] are shown in figure 3, in monthly frequency. The study comprises the 1964.04-1986.02 span. The former marks the end of a period of monetary imbalance and the beginning of a vigorous and successful plan to stabilize inflation, based on severe fiscal and monetary policies. The period truncation is in 1986.02 due to radical change in the inflation fighting policy – which took place in 1986.3 – and a permanent change in the inflation dynamics, which are supposed to have changed inflation expectation formation rules.

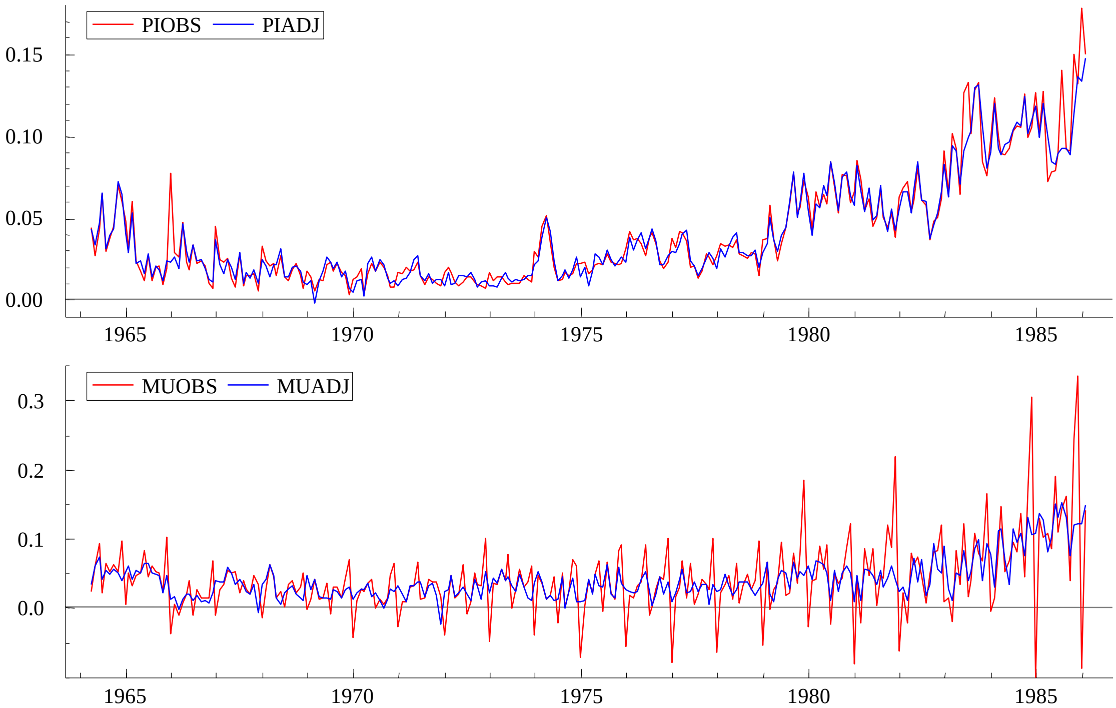
When looking for a causal relation, the first step is determining the series integration order, and then testing the existence of a cointegration relation between them. If there is such, it is tested if any of the variables can be regarded as weakly exogenous with a given interest parameter. Ultimately, causality tests in Granger's sense are carried out.

As the series present many outliers and their presence causes difficulty to the implementation of cointegration tests (see Cerqueira, 2006a), we chose to have the series go through previous treatment by using univariate structural models[12] with Kalman filtering procedures. We also decided to seasonally adjust the series by putting aside the respective stochastic seasonal component identified by same approach. Both conducts imply causing the VAR to be parsimonious, especially when working with monthly data. The adjusted and observed series are plotted in figure 3.

Unit root tests are reported in Appendix A. Inflation series shows no ambiguities I(1), the money supply series is also taken as I(1).

[11] By money growth rate we mean the percent variation of the M1 monetary aggregate; by inflation rate we mean the percent variation of IGP-DI of FGV, the Brazilian general price index.
[12] Experiments were also carried out with SUTSE bivariate models for outliers' extraction, which virtually pointed at the same dates as the univariate models.

FIGURE 3: Rate of inflation (PI) and Money Growth (MU) - Observed and Adjusted



The next step is finding a cointegration relation between the variables[13]. We chose Johansen's (1991) co-integration procedure test. Choosing the Var lag length and the deterministic components is crucial for test results. We decided to specify the VAR according to recommendation by Gonzalo (1994), Dolado & Lütkepohl (1996), Giles & Mirza (1998) and Lütkepohl & Saikkonen (1999). Thus we searched for a stable VAR – which led to the introduction of a linear trend term – whose residuals assessed by multivariate portmanteau (with minimal lag adjusted by the degrees of freedom and T/4) and Breusch-Godfrey type tests (1 to 12 lags) did not carry any serial co-relation. This made us reach 10 as the minimal lag number.

On the other hand, an overfitting model has little efficiency loss, whereas consistency is lost if the lag length is too small. Wald test's power loss is small when extra lags are added, in case the true VAR order is large and the system's dimension, small. Moreover, when unidirectional causality is suspected, overfitting methods cause less distortion with often little or no power loss if compared to the pre-testing procedures. For those reasons, we chose to work with an overfitted VAR(11).

Although the series are treated as outliers, the experiments' residuals show strong asymmetry[14], leading to rejection of the normality hypothesis according to Bera-Jarque tests in Lütkepohl and Urzua versions[15]. According to

[13] If we had considered the money growth series as I(0), we could have estimated a VAR in levels form, following Sims, Stock & Watson's (1990) recommendations; or a VAR with inflation in difference and money growth in levels. The experiments carried out, however, show that, with few lags ($\ell \leq 5$), there is a strong serial correlation, and with a lot of lags the VAR is unstable. We decided to leave this route and impose the cointegration restriction; see Lütkepohl (1993). Besides, our intuition points at the existence of a cointegration relation between inflation and money growth.

[14] This feature appears in the inflation rate equation residuals.

[15] Cheung and Lai (1993) argue that skewness in innovations has a statistically significant effect

Doornik-Hansen version, rejection is caused by the presence of kurtosis. Though Johansen's procedure is derived from a Gaussian likelihood function, its asymptotic properties exclusively depend on IID error hypothesis. The normality hypothesis is thus not vital to conclusions – although the ARCH effect might be; see Johansen (1995).

The cointegration tests' results are displayed in Appendix B[16]. Table B1 shows non-restrict VAR residuals diagnostic tests without imposition of cointegration relation. The tests point at the presence of ARCH elements and non-conditional heteroskedasticity; besides, residuals are not Gaussian. The causes of such violations derive fundamentally from inflation rate, which does not hamper a cointegration relation. We therefore rejected the null hypothesis that there is no cointegration at 1% significance level according to the trace statistics and at 1.8% with the maximum eigenvalue statistics.

The hypothesis that there is at most one cointegrating vector[17] cannot be rejected. The inspection of the residuals after restricting the cointegrating space to one vector enhances the residuals' stochastic properties, closer to being Gaussian; see table B3. As in the non-restricted model, some hypotheses violations can be traced to the residuals of the inflation equation. There are also four roots in the companion matrix around 0.90. This shows that the system's stability is far from ideal, although it is not explosive. On the other hand, the multivariate diagnostics tests show that our choice of 11 lags for the VAR was appropriate. Firstly because the residuals have no serial correlation and secondly for choosing a more parsimonious VAR would have led to the estimation of misspecified VECM model with autocorrelated residuals.

Cointegration corroborates the hypothetical absence of no rational bubbles from 1964.04 to 1986.02. This implies excluding the hypothesis that inflation acceleration in the late-1985 to early-1986 period was causing speculative hyperinflation. Testing for the plausibility of the (1,-1,#,#) cointegrating vector confirmed a homogeneous long-run relation between money growth and inflation. Such is the classical representation: balance between inflation and money growth has cointegrating vector $\beta=(1,-1)$ and a

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on the test sizes of the Johansen tests, but less so far for the trace test. On the other hand, the trace test appears to be more robust to excess kurtosis (in innovations) than the maximum eigenvalue test. Overall, the trace test is found to be more robust to nonmormality than the maximum eigenvalue test.

[16] It is interesting to observe that ours results are close to Cerqueira (2006a). The author's Var was specified with centred seasonal dummies and a series of intervention dummies acting as exogenous variables and excluded from the long-term relation in Johansen's procedure. As it is well known these variables changes the cointegration tests asymptotic distributions, but is a common procedure to account for important short run effects that must be controlled to lessen violations of the Gaussian assumption about the stochastic part of the process; see Johansen and Juselius (1990), and Hansen and Juselius (1995).

[17] These results must be taken cautiously, once one of the series – money growth – has long memory and it is not a pure I(1) process. As discussed in the appendix, it is hard to differentiate the series from a I(1) process with unit root tests. Another situation concerns when the I(1) variable VAR representation has a near singular covariance matrix, in such case Johansen's LR test tends to find spurious cointegration, with probability asymptotically equals to 1; Engle-Granger test is thus more robust to avoid deceitful results, it is recommended to use both tests; see Gonzalo & Lee (1995) and Maddala & Kim (2002). We carried out a Engle-Granger test with constant and trend – with residuals of the second step not showing serial correlation – and concluded that at 5% we reject the hypothesis of zero cointegration between inflation rate and money growth. Thus, we can continue our analysis with certain relieve.

moving stationary drift term[18]. This could be represented by the real interest rate, which is a I(0) series; see Cerqueira (2000).

Table B4 shows the results of two weak-exogeneity tests conditioned on the existence of one cointegrating vector: the first uses a theoretical vector (1,-1,#,#); the second, the estimated vector shown in table B2. At usual significance levels, we found that money creation is weakly exogenous for the parameters of interest in the conditional model of inflation. However, the reverse is not true for the inflation.

Such results corroborate the implicit idea in equation (4). Monetary shock causes expected inflation to accelerate, which in turn increases inflation. The inflation rate drift above is steady state. Once the adjustment coefficient is negative[19], expectation increase is reduced, thereby forcing inflation rate down towards a long run path[20].

Table B5 shows the model's residuals evaluation with all above restrictions. Multivariate tests reject conditional and non-conditional homoskedasticity hypotheses precluding the validity of the normality tests, which reject the null. Yet, the companion matrix roots indicate that the system's stability was not changed by the imposed restrictions.

Two non-causality tests are shown in tables 3 and 4, and below each, the strong exogeneity tests are reported[21]. Granger's direct test specification is taken from the VECM estimates – with 11 lags – and the restriction given by (1,-1,#,#). The money creation equation residuals are approximately NIID, whereas inflation equation residuals are not. We thus used a parametrical bootstrap[22] to access the specific distribution of Wald (F) and likelihood ratio (LR) test statistics. As on can observe, conclusions are maintained.

Geweke, Meese & Dent's (1983) test is specified with the same number of lags and leads – 11. As the money growth equation residuals are NIID, results in tables 3 and 4 are the same. For the inflation rate equation dummies were introduced to take care of outliers, and the conditional variance was specified as a GARCH(1,1) process. With NIID residuals, results are altered, and all tests produce the same conclusions. Maybe, due to the excess of estimated parameters, the GMD test results were very sensitive to the GARCH specification used, and many experiments resulted in F statistics with negative signs. Thus, we considered that the results must be taken only as an illustration[23].

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[18] That is, in the cointegration relation the drift term is allowed to move over time due to the presence of a linear time trend.

[19] The value of α was found to be -0.2054.

[20] While not reported the results lead to the same conclusions presented in Cerqueira (2006a). We conclude that Johansen's procedure, in this exercise, is robust to the presence of exogenous terms in VAR.

[21] Strong exogeneity is and the conjunct hypothesis of weak exogeneity and Granger's noncausality.

[22] Because VECM residuals are orthogonal, the error terms could be extracted from a normal distribution with diagonal covariance matrix.

[23] Anyway, the fact that the tests present the same results demonstrates that they are robust to the non-causality test. Besides, our experiments reveal that violation of homoskedasticity and normality hypotheses in general do not have material effect on the tests' performance.

TABLE 3: Granger Causality Tests (P-Values) with non-Gaussian Residuals*

TESTS	GRANGER DIR. TEST		GMD TEST	
NULL HIP.	MI	PI	MI	PI
PI $\nRightarrow$ MI				
F	0.5785		0.5148	0.4388
LR	0.5097		0.3929	0.3409
MI $\nRightarrow$ PI				
F		0.0014	0.0030	0.1293
LR		0.0076	0.0115	0.0690

* The symbol $\nRightarrow$ means "does not cause in Granger's sense". Strong exogeneity: PI $\nRightarrow$ MI $\begin{Bmatrix} F = 0.6568 \\ LR = 0.5916 \end{Bmatrix}$, MI $\nRightarrow$ PI $\begin{Bmatrix} F = 0.0003 \\ LR = 0.0001 \end{Bmatrix}$. Mi represents the money growth and PI the inflation rate.

TABLE 4: Granger Causality Tests (P-Values) With Gaussian Residuals or Monte Carlo Simulations*

TESTS	GRANGER DIR. TEST		GMD TEST	
NULL HIP.	MI	PI	MI	PI
PI $\nRightarrow$ MI				
F	0.5749		0.5148	0.2665
LR	0.5749		0.3929	0.1530
MI $\nRightarrow$ PI				
F		0.0161	0.0030	0.0136
LR		0.0161	0.0115	0.0036

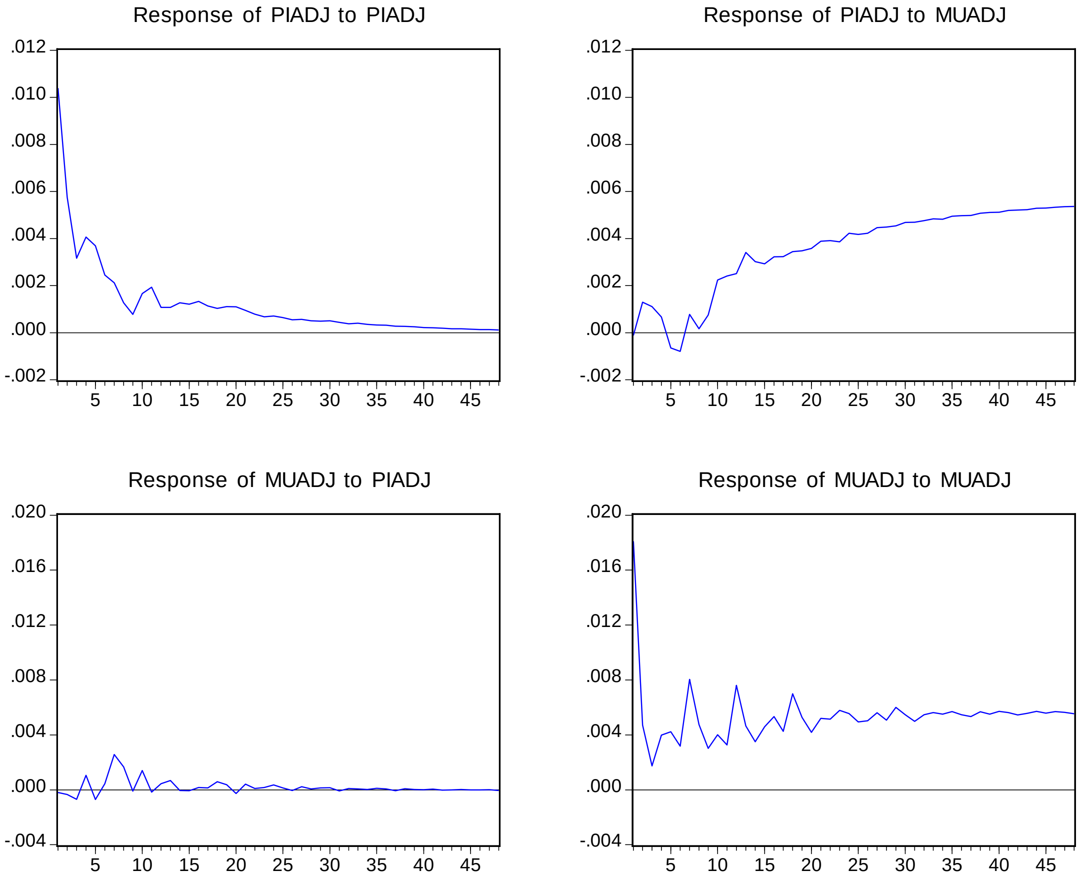
* Monte Carlo p-values for the strong exogeneity tests: PI $\nRightarrow$ MI $\begin{Bmatrix} F = 0.6510 \\ LR = 0.6510 \end{Bmatrix}$, MI $\nRightarrow$ PI $\begin{Bmatrix} F = 0.0002 \\ LR = 0.0002 \end{Bmatrix}$.

Causality tests' results show that money growth causes inflation, whereas inflation rate fails to cause money growth. There is not any type of feedback from inflation to the currency. This result contrasts with Cerqueira (2006a). That can be seen in the impulse-response functions [24] (figure 4) estimated with generalized impulse. This procedure constructs an orthogonal set of innovations that does not depend on the VAR ordering[25]; see Pesaran and Shin (1998).

[24] As in Lütkepohl and Reimers (1992), the impulse-response functions were calculated using (1,-1,#.#0) as the cointegrating vector and the adjustment factor in the form (-0.2054,0).

[25] However the variables order is, in our context, irrelevant, once the correlation coefficient between restrict VECM residuals is -0.0106, which is approximately zero with p-value 0.6533 (Pearson's test), and the hypothesis that the inner product (0.0005) is zero is accepted with p-value = 0.8794 (by using conventional *t* test).

FIGURE 4: Response to Generalized One S.D. Innovations



The one standard shock in money growth causes a period of fluctuations in this series, until it achieves its new steady state level, around 0.56 percent (hypothesis accepted with p-value = 0.8379) above its initial level. Indeed, beyond period 27, the impulse functions are statistically different from zero, with p-value around 0%, which back this claim.

The impulse in monetary expansion leads to a permanent increase of the inflation rate, but an inflationary shock does not have any meaningful effect over money growth. The peak shown in the seventh period is not significant with p-value 0.8544, and the average impact is close to zero all over the period. On the other hand, a monetary shock has permanent impact over inflation rate, next to 0.55 points. When innovations are considered, either in inflation rate or in money growth, 84% of final variation in inflation rate is caused by monetary shock. This result – attained with the inflation rate prediction error variance decomposition – shows a feedback effect (around 16%) over the inflation rate due to inflationary expectations. We can conclude that the existing persistence in inflation was due to monetary causes mainly, not to disturbances on the real side of economy or to the price indexation system.

As for the inflation rate from 1964.04 to 1986.02, there was enough evidence to validate the hypothesis that money growth is strongly exogenous[26]. Evidence supports the claim that causality is unidirectional and moves from

[26] Because it is weakly exogenous for the parameters of interest in the conditional model of inflation and MI is not caused in Granger's sense by PI.

money expansion to inflation[27]. This means that money supply was not passive and was econometrically exogenous stricto sensu (Sims, 1972) with respect to determining prices[28]. Besides, this strongly indicates that monetary authority did not, in the period, follow as monetary rule a reaction curve as the one described in equation (5)[29].

Finally, once the money supply was exogenous, Sargent and Wallace's (1973) model reveal that the inflation expectation from 1964.04 to 1986.02 followed Cagan's adaptative scheme. This implies that the rule must be applied in future research. Appendix C shows an estimated series of expected inflation formed adaptatively.

5. Conclusions

This paper presents tests on the exogeneity of Brazilian monetary supply for the military period comprising 1964.01 to 1986.02, using monthly data. We chose this period because the macroeconomic policy was more homogenous regarding the inflation stabilization policy than that of nearby years.

The results show that even if there was nearly stable average seigniorage between 1976.1 and 1988.4, this did not lead to the endogeneity of the money supply, since money growth was strongly exogenous with respect to the inflation rate. Results show that the monetary expansion caused the inflation rate in Granger's sense, and the former was weakly exogenous compared with

[27] By means of recursive estimation we investigate our model parameter constancy. The recursive cointegrating coefficients are, according to PcGive procedure, within the confidence bands, whereas Cats in Rats procedure suggests some breaks after 1982. On the other hand, the inflation equation adjustment coefficient is stable and does not present relevant fluctuations. From another viewpoint, the trace tests statistics, despite a dramatic change in the early 80s, describe an upward trajectory, as required (see Hansen and Juselius, 1995). Whereas the log-likelihood function is monotonously descending, but it is not within the 95% confidence interval. That does not happen to the eigenvalue, which is in the whole sample within the 95% asymptotic bands, showing there is not non-constancy of $\hat{\beta}_i$ and $\hat{\alpha}_i$ parameters in the partial model. The 1-step ahead and N-steps ahead prediction tests point at the existence of many breaks all over the period after 1974, which confirms with CUSUMSQ; these point at problems in the short-term parameters. The last result should be expected due to the turbulences the Brazilian economy went through in the period in question; see Cerqueira (2006a). It seems that in long run there are not any material signs making us reject $\hat{\alpha}$ and $\hat{\beta}$ non-constancy. Further, there is no indication that overidentification of imposed restriction is taking place. Reason shows that the estimated model has acceptable constancy properties. To save space, the tests carried out were not reported, but they are available upon request.

[28] Using the same series and extending the period to 1960.01 the results change, despite the presence of cointegration relation and the accepted homogeneity hypothesis. The hypothesis of weak exogeneity in any of the equations is not possible, and noncausality is rejected in both directions, which means that the inflation rate and money growth are endogenous, which is a similar result to Pastore's (1997). However, such result is not a surprise, once the period prior to 1964.04 is marked by complete lack of monetary control, with public deficits being financed basically by money issuing. In this period, inflation tax collection was essential for maintaining public expenditures. This imbalance turned out to modify the relation between money growth and inflation. Thus, we underline that the results in this article refer only to the period comprised between 1964.04 and 1986.02.

[29] This conclusion contradicts Marques (1983), Pastore (1994/95) and especially Pastore (1997). We conjecture that difference in the methodologies and data frequency can explain the divergence between our results and the so-called "common wisdom" among Brazilian economists.

the latter. This was possible because the monetary authority chose to reduce the base multiplier in order to keep its proportion of seigniorage collection. Therefore, even with (i) a permanent deficit with the seigniorage playing a crucial role in balancing the public accounts, and (ii) a host of financial innovations that led to the money demand contraction, the money supply remained exogenous respecting the inflation rate. Therefore, Brazilian inflation followed an ever-increasing path without setting off a hyperinflationary process.

We may conclude that money creation influences current and future inflation rates, but, given lagged rates of money creation, past inflation rates exert no influence on money creation. This is an indication that Cagan's rational adaptive schemes are not adequate to the Brazilian economy and that the rule followed by the economic agents to form expectations about inflation was adaptative. This contrasts sharply with an existing tradition among Brazilian economists, who assume that the monetary policy was completely passive during the 70s and 80s.

Indeed, our results reveal that the monetary policy was executed in an independent way, that is, the rule guiding the monetary execution was taken exogenously with regard to the considered model and the inflation rate. Therefore, we postulate that the monetary authority chose to finance a roughly fixed rate of the public deficit by issuing money. This explains the intermittent monetary expansion and the inflation rate. Such policy generated a vicious cycle: by exacerbating the inflation expectations, it introduced a feedback in the inflation growth. In addition, it provoked the uninterrupted growth of the debt-service payments and, consequently, the continuous increase of the nominal public deficit.

Rejecting the causality from prices to money does not mean to propose there was rigid monetary control. This depends on the monetary regime. In a regime whose monetary authority is independent, it is possible to fulfill **almost** any target of money stock. Compelled by the public deficits, the monetary authority may refuse to buy public bonds in the open market and then impose upon the fiscal authority the cost of increasing the real interest rate through the primary auctions. In the Brazilian regime, this was a task of the Central Bank. It was enough that the real interest rate was restricted to float between a given range to determine the deficit monetization through repurchase agreements. In this regime, the monetary control was indirect, and the instruments less efficient but enough to manage an exogenous money supply from the inflation variations. Perhaps these are somewhat old monetarist ideas, but we cannot deny they stamped their mark on the data.

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Appendix A: Integration Order of the Inflation Rate (PI) and Money Growth (MI)

Table A.1 reports the results of unit root tests to the series in levels and in first difference with monthly frequency. Besides the traditional ADF (Augmented Dickey-Fuller) and PP (Phillips-Perron) tests we also report their well-known modifications, to wit: DF-GLS (Dickey-Fuller test with GLS Detrending), ERS-PO (Elliot, Rothenberg & Stock point optimal test) e Ng-Perron (NG and Perron test); see Maddala and Kim (2002). Inspecting the sample autocorrelation functions and analyzing the results sensibility to the ℓ variations chose the bandwidths ℓ. These procedures lead us to employ for all tests the number of lagged difference terms specified to carry out the ADF test[30], which turned out to be the tests conservative. Tests to inflation rate level were specified with ℓ = 2; for its first difference ℓ = 1; and to the monetary growth, respectively, ℓ = 11 and ℓ = 10. To estimate zero frequency spectra we chose quadratic spectral kernel[31]. Computed p-values in brackets were also reported to ADF and PP tests statistics.

The inflation rate is I(1), but tests are inconclusive about money growth since the reported tests do not show the same results if the series is stationary or integrated. The class of tests that uses spectral corrections for serial correlation control rejects with size 1% the unit root hypothesis. While the tests that make use of lagged difference of the dependent variable, with the same intention, accept H_0 to the series in levels but differ on the stationarity of its first difference[32].

Monte Carlo studies[33] by Elliott et alii (1996), Stock (1994 and 1999) point that the DF-GLS statistic appears to represent a compromise, in the sense that its power is high – based on results in Elliott et alii (1996), typically as high as the asymptotic point optimal test P_T – but its size distortions are low, although not as low as the ADF-t statistic. Indeed, in our study both tests indicate the monetary expansion is at least I(1). Moreover, the KPSS test that H_0 is the stationarity, rejects the null at 1% level for the series in levels and accepts this hypothesis for its first difference. Additionally, we computed Bayesian test for unit root (Sims, 1988) – which places the stationarity and alternative hypotheses on the same footing –, specified with constant, 11 lags, stationarity a priori probability equal to or less than 0.70, lower limit for stationary prior equal or less than 0.50. The test presents strong evidence that the series has a unit root with marginal a posteriori probability equal to or greater than 0.699[34]. Furthermore, the structural model estimated with Kalman filter procedures indicates the presence of a stochastic trend among the series components. Finally, we implemented a parametric bootstrap to the ADF $\hat{t}$ e $\hat{\rho}$ statistics with above specified parameters; with p-values 0.777 and 0.887,

[30] We employed enough lagged difference to get residuals with no serial correlation tested by the Ljung-Box statistics with lags from to 1 to ¼ of the sample (near 105), and by the Breusch-Godfrey test in all lags from 1 to 12.

[31] See Andrews (1991) for an extensive analysis of kernel function properties.

[32] These confused results are the same to the observed series, that is it, not seasonally adjusted and with outliers extracted.

[33] These studies do not include the MP_T^d test.

[34] Using HEGY test we did not find evidence of a seasonal unit root neither in series levels nor in its first difference.

respectively, there is no reason to reject that the series has a unit root. Considering the previous arguments we decided to take the monetary growth as being I(1)[35].

TABLE A1: Unit Root Tests

Test / Series	ADF	DF-GLS	PP	ERS-PO	Ng-Perron			
					MZ_{α}^{d}	MZ_{T}^{d}	MSB^{d}	MP_{T}^{d}
PI	-0.3250 (0.9175)	-0.6547	-1.3833 (0.5905)	5.4947	-5.9332†	-1.2513	0.2109‡	5.4723
ΔPI	-17.3389* (0.0000)	-13.5691*	-20.7605* (0.0000)	1.1901*	-29.8104*	-3.7964*	0.1274*	1.0253*
MI	0.3954 (0.9825)	0.8579	-8.3199* (0.0000)	0.5988*	-117.137*	-7.4376*	0.0064*	0.5846*
ΔMI	-6.1662* (0.0000)	-0.3041	-58.8791* (0.0001)	7.0571	-4.0041	-1.2602	0.3147	6.2954

Note: (†) represents rejection of a unit root at the 10% significance level; (‡) at 5% significance; (*) rejection at 1%; no symbol means acceptation of the null hypothesis at 10%.

Appendix B: Cointegration Tests

In this appendix we report the diagnostic tests to the estimated VARs. We report the p-values to the Q (Ljung-Box), LM (Breusch-Godfrey), White, ARCH-LM, Q^2 (McLeod-Li), BJ (Bera-Jarque), LUTK (Lütkepohl), DH (Doornik-Hansen), Urzua[36]. The statistics of model evaluation R^2, SER (std. error reg.), LogL (log likelihood), CorTr (trace correlation) and the information criteria AIC e SIC are presented with the estimated values. We also report the four biggest roots of the VAR companion matrix and the Johansen's procedures.

Table B1 shows the statistics of the unrestricted VAR without imposition of cointegration relation; table B2, the cointegration tests with critical values at 1% level and the respective p-values; see Mackinnon-Haug-Michelis (1999); and table B3, the diagnostic analysis of the VECM with the restriction the cointegrating vector reported in table B2.

TABLE B1: Unrestrict VAR Diagnostics Tests

MULTIVARIATE TESTS					
N=263	LogL=1539.179	AIC=-11.3398	SIC=-10.6878	Q(12)=0.0513	Q(24)=0.2619
Q(36)=0.1983	Q(65)=0.5742	LM(1)=0.1275	LM(3)=0.4668	LM(6)=0.5947	LM(9)=0.9751
LM(12)=0.2093	WHITE=0.0000	LUTK=0.0033	DH=0.0000	URZUA=0.0336	ρ1=0.9923
ρ2=0.9177	ρ3=0.9007	ρ4=0.8614			
UNIVARIATE TESTS – EQUATION 1: DPI					
R^2=0.9015	SER=0.0104	AIC=-6.2085	SIC=-5.8826	Q(12)=0.9950	Q(24)=0.9635
Q(36)= 0.8668	Q(65)=0.9458	LM(1)=0.9653	LM(3)=0.9961	LM(6)=0.9999	LM(9)=0.9999
LM(12)=0.9953	ARCH(1)=0.2510	ARCH(4)=0.0745	ARCH(12)=0.0000	Q^2(4)=0.0405	Q^2(12)=0.0000
Sk=0.5929	Ek=1.0402	BJ=0.0000	DH=0.0007		
UNIVARIATE TESTS – EQUATION 2: DMI					
R^2=0.6824	SER=0.0178	AIC=-5.1311	SIC=-4.8051	Q(12)=0.9759	Q(24)=0.6992
Q(36)=0.4740	Q(65)=0.3971	LM(1)=0.6008	LM(3)=0.9396	LM(6)=0.9941	LM(9)=0.9994
LM(12)=0.9744	ARCH(1)=0.4982	ARCH(4)=0.4865	ARCH(12)=0.3269	Q^2(4)=0.5078	Q^2(12)=0.2126
Sk=-0.1756	Ek=0.1181	BJ=0.4715	DH=0.4461		

[35] As showed in table A1 the ADF and DF-GLS have the same results for the series levels; however, they disagree for the first difference series. This is a common fact when one works with unit root tests. We conjecture this ambiguity is due to the presence of outliers along the whole series. In general outliers disturb the tests performance by bringing them about less conservative.

[36] The last three are multivariate extensions of the Bera-Jarque residual normality test, which distinguish among themselves by the factorization method of the residuals covariance matrix.

TABLE B2: Johansen Cointegration Test

TEST STATÍISTICAS (P-VALUES)				COINTEGRATING VECTOR
Trace		λ-Max		(MONEY, INFLATION, TREND, CONSTANT)
r=0	r≤1	r=0	r≤1	
35.8742	13.5674	22.3039	13.5673	(1.0000, -0.9645, -0.0002, 0.0903)
(31.1539)	(16.5539)	(23.9753)	(16.5539)	
(0.0020)	(0.0333)	(0.0183)	(0.0333)	
COINTEGRATION RESTRICTION TEST				
RESTRICTION: (1,-1,#,#) $\chi^2(1)$ = 0.0459; P-VALUE = 0.8309				

Note: The symbol # means the parameter is unrestricted. The estimated eigenvalues are 0.0813 and 0.0503.

TABLE B3: Restricted VECM Diagnostics Tests

MULTIVARIATE TESTS					
N=263	LogL=1539.089	CorTR=0.3859	AIC=-11.2859	SIC=-10.5932	Q(12)=0.1848
Q(24)=0.3617	Q(36)=0.3418	Q(65)=0.7651	LM(1)=0.4243	LM(3)=0.4483	LM(6)=0.8758
LM(9)=0.8327	LM(12)=0.3725	WHITE=0.0000	LUTK=0.0062	DH=0.0000	URZUA=0.0624
ρ1=0.9307	ρ2=0.9186	ρ3=0.9095	ρ4=0.9023		
UNIVARIATE TESTS – EQUATION 1: DPI					
R^2=0. 2969	SER=0. 0104	AIC=-6. 2090	SIC=-5.8830	Q(12)=0.9990	Q(24)=0.9765
Q(36)= 0.9020	Q(65)=0.9649	LM(1)=0.9117	LM(3)=0.9964	LM(6)=0.9998	LM(9)=0.9999
LM(12)=0.9990	ARCH(1)=0.2373	ARCH(4)=0.0626	ARCH(12)=0.0000	Q^2(4)=0.0319	Q^2(12)=0.0000
Sk=0.5821	Ek=1.0011	BJ=0.0000	DH=0.0009		
UNIVARIATE TESTS – EQUATION 2: DMI					
R^2=0.4758	SER=0.0181	AIC=-5.0996	SIC=-4.7736	Q(12)=0.9999	Q(24)=0.7185
Q(36)=0.5828	Q(65)=0.4222	LM(1)=0.9548	LM(3)=0.9834	LM(6)=0.9996	LM(9)=0.9999
LM(12)=0.9999	ARCH(1)=0.6335	ARCH(4)=0.4090	ARCH(12)=0.7291	Q^2(4)=0.4733	Q^2(12)=0.6027
Sk=-0.0982	Ek=0.1264	BJ=0.7418	DH=0.6326		

In table B4, we present the weak-exogeneity test using two different statistics. The first tests the joint hypothesis that the cointegrating vector is (1,-1,#,#) and the adjustment coefficients are respectively (0,α) and (α,0). The second, in brackets, uses the estimated cointegrated vector reported in Table B2 and these two adjustment coefficients. Table B5 shows the diagnostic tests to the restrict model with vector (1,-1,#, #) and adjustment coefficient equal to zero to the monetary expansion equation.

TABLE B4: Adjustment-Coefficient Weak Exogeneity Test

NULL HYPOTHESIS	TEST STATISTIC	P-VALUE
MI is weakly exogenous for the parameter of interest of the PI conditional model	0.0823 (0.0265)	0.9597 (0. 8707)
PI is weakly exogenous for the parameter of interest of the MI conditional model	8.8580 (8.7229)	0.0119 (0.0031)

Note: Monte Carlo p-values H_0: MI is weakly exogenous to PI $\begin{Bmatrix} F = 0.9090 \\ LR = 0.9090 \end{Bmatrix}$ **and H_0: PI is weakly exogenous to MI** $\begin{Bmatrix} F = 0.0003 \\ LR = 0.0003 \end{Bmatrix}$**.**

TABLE B5: Diagnostics Tests: VECM with Weakly Exogeneity Restriction

MULTIVARIATE TESTS					
N=263	LogL=1535.048	CorTR=0.3859	AIC=-11.2855	SIC=-10.5958	Q(12)=0.1602
Q(24)=0.3278	Q(36)=0.3117	Q(65)=0.7319	LM(1)=0.4258	LM(3)=0.4307	LM(6)=0.8687
LM(9)=0.8264	LM(12)=0.3300	WHITE=0.0000	LUTK=0.0107	DH=0.0001	URZUA=0.0945
ρ1=0.9410	ρ2=0.9178	ρ3=0.9100	ρ4=0.9015		
UNIVARIATE TESTS – EQUATION 1: DPI					
R^2=0.2968	SER=0.0104	AIC=-6.2089	SIC=-5.8829	Q(12)=0.9986	Q(24)=0.9672
Q(36)= 0.8778	Q(65)=0.9531	LM(1)=0.9153	LM(3)=0.9964	LM(6)=0.9998	LM(9)=0.9999
LM(12)=0.9987	ARCH(1)=0.2410	ARCH(4)=0.0545	ARCH(12)=0.0000	Q^2(4)=0.0265	Q^2(12)=0.0000
Sk=0.5551	Ek=0.9829	BJ=0.0000	DH=0.0014		
UNIVARIATE TESTS – EQUATION 2: DMI					
R^2=0.4757	SER=0.0181	AIC=-5.0994	SIC=-4.7734	Q(12)=1.0000	Q(24)=0.7014
Q(36)=0.5688	Q(65)=0.4036	LM(1)=0.8463	LM(3)=0.9783	LM(6)=0.9995	LM(9)=1.0000
LM(12)=0.9999	ARCH(1)=0.6715	ARCH(4)=0.4573	ARCH(12)=0.7137	Q^2(4)=0.4910	Q^2(12)=0.5970
Sk=-0.1023	Ek=0.1454	BJ=0.7082	DH=0.5963		

Appendix C: Estimating Inflationary Expectations Following Cagan's Adaptive Scheme (1956)

As established in section 5, the monetary expansion from 1964.04 to 1986.02 is strongly exogenous with respect to the inflation rate. Evidence supports the claim that the causality is unidirectional and moves from money growth to inflation. Thus, in accordance with Sargent and Wallace (1973) Cagan's adaptive scheme is not rational. Therefore, agents are assumed to compute current inflation expectations with lagged rates of inflation. As the inflation rate (π) is integrated of order one, for any k, π_t e π_{t-k} will be cointegrated; see Granger (1986). Then, for estimation matters we can take the static regression of Engle-Granger two-step procedure, since it represents the long-run relation between contemporaneous and past inflation values; in steady state the observed and the expected inflations are equal. With this equation we can predict the current expectation inflation series. The results are in table C1 and the forecast evaluation in the table C2. The computed and observed series are plotted in figures C1. We suggest employing these series in further studies.

TABLE C1: **Engle-Granger Cointegration Test**

Dependent Variable: PI

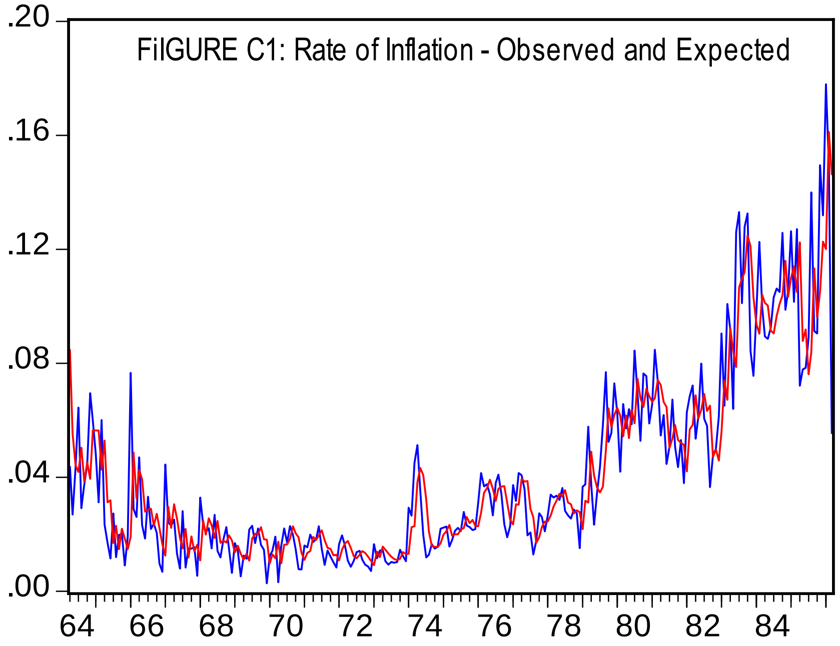
Variable	Coefficient	P-Value
C	0.0013	0.3613
PI(-1)	0.5502	0.0000
PI(-2)	0.1246	0.0725
PI(-3)	0.3045	0.0000

N=263	R^2=0.8219	SER=0.0143	AIC=-5.6443	SIC=-5.5900
$t_{\hat{\alpha}}$ = -8.5270*	Lags = 11	DW=1.9778	Q(12)=0.9945	Q(24)=0.9546
LM(1)=0.7876	LM(3)=0.6913	LM(6)=0.7980	LM(9)=0.7835	LM(12)=0.5786

Note: Critical value with is -5.0558; associated p-value = 0.0000.

TABLE C2: Goodness of Fit

Root Mean Squared Error	0.0152
Root Mean Squared Percentage Error	0.0100
Mean Absolute Percentage Error	0.3554
Theil Inequality Coefficient	0.1454
Bias Proportion	0.0005
Variance Proportion	0.0262
Covariance Proportion	0.9733
Pseudo R^2	0.7955



FIlGURE C1: Rate of Inflation - Observed and Expected