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Monetary Exogeneity in Brazil
– A Restatement:
Brazil from 1966 to 1985

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ABSTRACT

Using monthly data spanning from 1966 to 1985, we examine the money supply exogeneity. The implemented tests investigated the plausibility of classical hypotheses. We employed GARCH processes and the bootstrap approach. The results are robust to the choice of Granger-causality tests. We showed that the real rate of interest did indeed 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 rate of inflation. Money creation and the rate of inflation were cointegrated and the causal relation between them was unidirectional from money to inflation. These empirical findings are sharply different from many previous results. Our main contribution is to demonstrate that the monetary supply was exogenous with respect to the rate of inflation and that the monetary authority had enough independence to execute an active monetary policy.

KEYWORDS: money exogeneity, noncausality tests, model selection procedures, non-normality correction, GARCH, bootstrap, Cagan's model.

PALAVRAS-CHAVE: moeda exógena, testes de não-causalidade, procedimento de seleção de modelos, correção de não-normalidade, GARCH, bootstrap, modelo de Cagan.

JEL: C32, C51, C52, E31, E51, E52

1. Introduction

From 1966 to 1985, the inflation fighting policy was largely centered on aggregate demand management and wage and price controls. The fact that during these years the economic policy was conducted in a somewhat “orthodox way” may induce one to ask about the monetary policy behavior. Are there reasons that support the claiming 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 the present paper is to confirm that from 1966 to 1985 the monetary supply was exogenously determined. We intend to verify this assumption by means of three arguments. First, the monetary authority conducted the open market policy increasing the real interest of rate in order to stimulate the demand for bonds. Second, even if during this period the seigniorage collection remained constant as a share of GDP, the government succeeded in keeping its portion of the revenues by reducing the base multiplier. Moreover, the seigniorage-GDP ratio was independent from the rate of inflation and followed a white noise process. 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 1966-1985 period is largely diffused among many Brazilian economists. One probable rationale that gives support to 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, ruling out the possibility of rational explosive bubbles. In this case, the money supply is endogenous.

An alternative argument is made based on Sargent and Wallace's (1973) scheme derived from Cagan's model under the hypotheses of adaptive expectations and a monetary rule which depends on inflation past rates. This is a model in which the adaptive mechanism is rational. In their model, the best way to forecast the subsequent rates of money creation is to extrapolate lagged rates of inflation. This in turn implies that inflation itself is best predicted by extrapolating lagged rates of inflation. So, both money creation and inflation are best predicted by extrapolating current and lagged rates of inflation. Lagged rates of money creation add nothing to the predictions formed in this way. In this model, past values of inflation influence money creation but the converse is not true; thus, money creation is passive.

An essential element in this argument is the hypothesized feedback that occurs from expected inflation to money creation. This feedback emerges because of the government's attempt to finance a roughly constant rate of real expenditures principally by money creation. In this sense, this is also a description of Blanchard and Fischer's (1989) version of Cagan's model, in which the monetary expansion is endogenously determined by expected inflation given a constant level of seigniorage.

If the seigniorage collection 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. The inflationary process is then totally inertial because the current inflation rate is a function of the lagged rates of money creation; but given lagged rates of money creation, past rates of inflation exert no influence on money creation. The system is one in which money creation causes inflation, in Granger's sense, while 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 gave rise to a somewhat efficient indexation system that protected agents from the effects of inflation. Even if the indexation rules did not contemplate the agents in a fair way, one can not deny that they prevented the high inflation to degenerate into a public panic and an open hyperinflation process. Furthermore, the indexation rules were developed little by little along the seventies and eighties amid an increasing inflation rate. Then the rigidity in the price system was introduced gradually, which increased the inflation inertia without destroying the inflationary memory. These arguments are supposed to explain why agents had adaptive expectations about inflation. Therefore, Brazil's experience over the period did not show evidence of expectations being formed in a rational way.

One can also argue that the monetary policy followed a rule, which was independent of the inflation rate. In Cerqueira (1993), this assertion was tested with quarterly data. An empirical investigation using monthly series is reported in the appendix.

The assumption of endogenous money growth may not be supported by the empirical evidence when tested. Surprisingly, some authors found a unidirectional relation between inflation and money. We can suppose that their results came about because of their use of lower frequency data (quarterly). The resulting loss of information may have distorted the results of the 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[1] (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 mode. It is therefore unwise to rely exclusively on this test in checking for model adequacy. However, it can be, valuable when used in conjunction with other checks. The Breusch-Godfrey Lagrange multiplier test is a common complement to the Q-test (Granger and Newbold, 1986).

As it will be shown in this paper, by carrying out both tests and using monthly data from 1966.01 to 1985.12[2], we achieved results that contrast sharply with previous findings of others Brazilian authors[3]. This might explain why in applying the same causality tests different conclusions emerged.

Why the monetary expansion followed an I(1) process between 1966 to 1985 remains an open question. Although the answer surpasses the objectives of the present paper, 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. If the rule governing monetary authority was to achieve debt sustainability then it was urged to support the deficit financing by increasing the money growth. This led to a pegging of the inflation rate. Thus, the monetary authority chose, or was compelled to choose, the deficit

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

[2] We chose this period because the economic policy was more homogenous than if we had included nearby additional years. As we pointed out, the aggregate demand and price controls were the main tools of the macroeconomic policy. The beginning of this period is marked by the end of a resolute and successful orthodox inflation stabilization plan. The end is set off by the beginning of an assortment of stabilization attempts, which were essentially characterized by price freeze and the break of existing private contracts of wages, borrowings, rents and so on.

[3] In this paper, we used monthly data to compute our results that span the whole period. Quarterly series or for that with missing data we interpolated or simulated the values. 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 the lag length of autoregressive process and rigidly observed the classical econometric hypotheses.

inflationary financing in order to sustain the debt[4]. In this sense, there was a choice of economic policy.

The paper is organized as follows. Section 2 presents a brief review on the relations among deficit, debt, and monetary expansion. Section 3 describes the Brazilian monetary regime and reports on the causality tests between the real interest rate and the public debt. Section 4 discusses theoretical issues on seigniorage collection and monetary exogeneity. Section 5 presents the causality test results between money creation and inflation rate, and in section 6, we present the conclusion. In the appendices A, B and C we provide, respectively, the statistical procedures for expanding the public debt series range, for interpolating the GDP series and for determining the seigniorage-GDP ratio stochastic process. Appendices D, E and F show, respectively, the statistical reports concerning the causality tests between the real interest rate and the public debt stock, the estimates of the demand for bonds, and the causality tests between money creation and inflation rate.

2. The dynamics of public debt and the Sargent and Wallace dilemma: A Review[5]

When does a public deficit financed with debt expansion lead to an insolvent path or when does it give rise to an inflationary process? In a monetary regime where the monetary authority (MA) is independent, the fiscal authority (FA) determines the public debt growth and the monetary authority settles the debt financing composition between bonds and money. In such regime, the MA controls the money creation by following

[4] Issler and Lima (1997) suggest that seigniorage revenues were critical to restore intertemporal budget equilibrium and keep debt sustainable.

[5] This section is based in the following references: Alesina (1988), Barro (1974, 1978, 1979, 1986), Pastore (1990), Sargent (1984), Sargent e Wallace (1973, 1975, 1976, 1982).

its own rule of monetary expansion that reflects trade-off between efficiency and political economy considerations. In this regime, the monetary control is realized through open market transactions.

Let us suppose an economy, which the government budget constraint is given by

$$f_t = (g_t - \tau_t) + ib_t\,, \tag{1}$$

where g_t represents the fiscal expenditures excluding debt-service payments, τ_t is the tax revenues and b_t is the government bonds stock. All variables are expressed as a proportion of GDP. The real interest rate is defined as [r=i-π], where i and π are respectively, the nominal interest rate and the inflation rate. The public operational deficit (f_t less the nominal debt-service πb_t) is then given by

$$d_t^o = (g_t - \tau_t) + rb_t\,. \tag{2}$$

The first term is the primary deficit and the second the financial deficit component. Consolidating the government budget constraint, (Treasury plus Central Bank), we get:

$$f_t = \overset{\circ}{M}_t / Y_t + \overset{\circ}{B}_t / Y_t\,, \tag{3}$$

where $\overset{\circ}{M}_t$ and $\overset{\circ}{B}_t$ are respectively the monetary base and the nominal public debt derivatives with respect to time. Differentiating $b_t = B_t / Y_t$ and using the fact that $Y_t = y_t p_t$, where Y_t is the real output, we get

$$\overset{\circ}{b}_t = \overset{\circ}{B}_t / Y_t - (\pi + y) b_t\,, \tag{4}$$

where y is the output rate of growth. Combining the above equations, we get:

$$\overset{\circ}{b}_t = [(g_t - \tau_t) - \overset{\circ}{M}_t / Y_t] + (r - y) b_t\,, \tag{5}$$

which describes the dynamics of the debt-GDP ratio.

This result shows that the public debt increases by the uneven between the primary deficit and the seigniorage, and by the difference between the real interest rate (r) and the output growth rate (y). However, this can lead to an ever-increasing debt growth, if the public deficit is permanent and the monetary authority refuses to give it liquidity. This can generate the perception of a weakened monetary authority that can lead to deficit inflationary financing.

Suppose r, y and $\overset{\circ}{M}/Y$ are constants and the primary deficit is d=(g-τ), we then get the following solution to equation (5):

$$b_t = b_0 e^{(r-y)t} + \frac{d - \overset{\circ}{M}/Y}{r-\rho}(e^{(r-y)t} - 1), \tag{6}$$

which describes the public debt path. Note that if $(\dot{M}/Y)$=y=d=0 and only the real interest rate increases, the debt will grow at this rate. A recession and a primary deficit increase the operational deficit financial component, which can become dominant relatively to the primary term. These conditions may lead the fiscal authority to cut expenditures and increase taxes by an amount that becomes greater, at each instant this adjustment is postponed. Alternatively, they may induce the deficit "monetization" in order to avoid a probable government's default.

If the fiscal authority follows an intertemporal budget constraint, under the condition of a no-Ponzi path, the relation between r and y determines, if y>r, the convergence to a stationary debt given a permanent primary deficit. Integrating the equation (5) in infinite horizon, and imposing the transversality condition that b_t increases at an inferior rate than (r-y), we get

$$b_t = \int_t^{\infty} (\tau_s - g_s + \mu_s m_s) e^{-(r-y)(s-t)} ds, \tag{7}$$

where $\mu = \overset{\circ}{M}/M$ e m=M/Y. This intertemporal budget restriction implies that τ_t and g_t paths are such that, after a while, the primary surplus has to be risen, if m=y=0, but it does not imply a bounded growth nor a finite stationary value for b_t. Furthermore, even if d=0, the budget restriction is not respected if r>0 and y≤0. In this case, the only way to restore it is by generating a fiscal surplus or through expansion of the monetary base.

However, if it is the case that y>r, independently of the size of the deficit and the proportion of the deficit financing with monetary expansion, b_t will converge to a steady state finite value. We verify this by taking the limit of (6), so we get:

$$b^* = \frac{d - \mu m}{y - r}. \qquad (8)$$

Furthermore, if keeping the operational deficit constant is feasible and y>0, we get an stationary equilibrium even if y<r. This will permit to achieve a Ricardian fiscal regime. Combining (2) and (5), we have:

$$\overset{\circ}{b}_t = (d^o - \mu m) - y b_t, \qquad (9)$$

where d^o is the operational deficit constant level. Solving this equation, b_t converges to

$$b_t^* = \frac{d^o - \mu m}{y}.$$

Suppose the real interest rate does not grow with b_t, and the fiscal authority does not succeed to keep constant the operational deficit through expenditures cuts, thus the FA must raise the tax burden. Substituting the solution of the differential equation (9) in (2), we get the path to τ_t:

$$\tau_t = (g^* - d^o) + r b_0 e^{-yt} - (r/y)(d^o - \mu m)(e^{-yt} - 1). \qquad (10)$$

Then taking the limit, we obtain the desired burden tax

$$\tau^* = g^* + (r - y) b^*. \qquad (11)$$

This kind of rule stabilizes the public debt-GDP ratio, since now it oscillates around the long-run path. It implies a fiscal regime where the current deficits are compensated by future surplus. Moreover, if consumers have Ricardian behavior[6] any temporary deficit cannot be inflationary.

[6] In a Ricardian world, the public deficits are always feasible and may be financed by public bonds without pressure on the monetary base, because bonds are not taken as wealth but rather as a future liability. Thus, their present values are deducted from the existing wealth. Therefore, permanent income and consumption remain constant with the debt growth, while saving grows by the same amount of the debt growth in such a

However, to follow an intertemporal budget constraint does not imply any limit to the debt growth and the problem of a perceived lack of fiscal control persists, as well as the risk of uncontrolled inflation. This situation is harder if agents anticipate higher real interest rate and lower economic growth. So, the confidence is only restored when information confirms that the government is meeting its budget constraint, showing explicitly that it produces fiscal surplus.

Nevertheless, as pointed above, budget constraints may have infinite horizons and they are compatible with unlimited debt-GDP ratio growth. Thus, skepticism about the government's ability to cut expenses may prompt agents to ignore bonds in an attempt to protect them from probable government default.

Let us assume that the real interest of rate is greater than the output growth (r>y). Suppose also that after time T the government fails to finance its primary deficit by issuing bonds and, up to T, the monetary authority has been following a moderate monetary expansion (μ) rule, such that $\mu=\pi$. Then from time T ahead, the MA resistance vanishes and the entire deficit begins to be financed with high-powered money creation. The implied inflation size is then as large as b_t[7] and gets larger the tighter the monetary policy was until time T.

Let us assume a monetarist hypothesis of a "super-orthodox" public deficit, such that its financing with monetary expansion is not inflationary. This restriction is imposed with a demand for money like $M^d=kPY$. Suppose also $M^s=M^d$, then the inflation rate at period T will be $\pi=\mu-y$. Thus, the seigniorage is given by

$$\overset{\circ}{M}/Y=(\pi+y)m+\overset{\circ}{m}, \qquad (12)$$

way that the real interest rate stays constant and the crowding-out effect does not happen.

[7] However empirical data show countries with increasing debt-GDP ratios without presenting any kind of inflationary disturbance; see Alesina (1988).

where the first term is the inflation tax increased by the output growth and the second is the base-GDP ratio variation. In steady state $\overset{\circ}{\pi} = \overset{\circ}{\mu} = 0$ and the seigniorage is S=$\mu m = (y + \pi)m$. Given that beyond time T, $\overset{\circ}{b}_t = 0$, i.e., the whole deficit is financed with monetary expansion, from (5) and (2) we get

$$b_T = \frac{\mu m - (g - \tau)}{r - y} \tag{13}$$

and

$$d^o = \mu m = (\pi + y)m\,. \tag{14}$$

Combining these two equations gives

$$(\pi + y)m = (g - \tau) + (r - y)b_T\,, \tag{15}$$

which is rearranged to yield

$$\pi = \frac{(g - \tau) + (r - y)b_T}{m} - y\,. \tag{16}$$

The fraction term represents the operational deficit deduced from the public debt expansion allowed by the economy rate of growth, divided by the money-GDP ratio. By (14), this last expression is identical to the monetary expansion rate. Equation (16) shows the compatible inflation rate, after time T, with the operational deficit. Note that π is an increasing function in the public debt.

Therefore, in a world where r>y, even in the presence of an orthodox and independent MA, an expansionist fiscal policy may have its inflationary effects postponed just by sometime, provided there is an upper bound to the debt size.

If r>y the inflation starts at the time the monetary authority changes its policy to finance a permanent deficit. Nevertheless, under rational expectations and if the money demand is sensitive to interest rate fluctuations, the inflation rate begins to grow before the monetary expansion acceleration. Then a permanent primary deficit leads to an ever-growing debt perceived by the agents, who become skeptical about the debt sustainability. Thus, they anticipate a fatal deficit "monetization" and an increase in the cost of holding money, which leads to the immediate reduction of the desired money stock and the increase of the money velocity and inflation rate.

If r>y the fiscal policy imposes the behavior on the monetary authority. After time T, the MA will provide liquidity to the deficit with a sufficient increase in the seigniorage, to finance the treasury through the increase of the inflation rate (see equation 15). Nevertheless, if y>r, b_t will converge to a finite steady state value. From equation (16), and given the seigniorage and the operational deficit levels, we get the expression:

$$b^* = \frac{(g-\tau) - m(\pi + y)}{y - r}, \tag{17}$$

which represents the value of b_t given some steady state inflation rate. In this case, the monetary policy determines the public debt path. Setting the monetary expansion at the rate ($\mu=\pi+y$), also sets a stable inflation rate that produces the appropriate seigniorage level, which in turn will lead to the stabilization of the debt-income ratio.

In contrast, if the real interest rate is influenced by the debt growth even if, in time t_0, the rate of growth is greater than the real interest rate (y>r), this inequality may be inverted before the stabilization of b_t occurs. This will induce an increasing path for b_t and will bring back the inflationary instability associated with permanent primary deficits. Thus if the money and bonds demands are elastic to the real interest rate, a primary deficit may not be sustainable.

We can rewrite equation (15) in order to take into account generic function forms for bonds and money demands. Then we get

$$(g-\tau) = (\pi + y)m^d(r,\pi) + (y-r)b_t^d(r,\pi). \tag{18}$$

This expression is a steady state condition, when the money and bonds markets have reached equilibrium. The equation states that a primary deficit is feasible when it equals the inflation tax increased by the output growth, plus the revenue from selling bonds, permitted by the difference between the output growth rate and the real interest rate. If the economy's rate of growth is zero, in order to have a feasible deficit, it is necessary that $\pi>0$ and $r<0$.

Now let us return to the economy that has a Cambridgean money demand. We assume the MA has fixed the monetary expansion to a level μ (and hence the inflation rate), and we suppose that the output is not affected by real interest rate oscillations. Then, a permanent increase in the primary deficit must be completely financed, in steady state, by additional sales of bonds. We show this by taking the derivative of (18):

$$(\overset{\circ}{g-\tau}) = -\overset{\circ}{r}\, b^d(r,\pi) + (y-r)\frac{\partial b^d(r,\pi)}{\partial r}\overset{\circ}{r}\,. \tag{19}$$

It follows that to make this deficit increase feasible, without changing the seigniorage level, the real interest rate must be inferior to a maximum r*. Otherwise, the revenue from selling bonds will be reduced. But even when r<r*, the deficit increment may be superior to the marginal bonds' revenue, which will make this increment unfeasible. In this case, the pressure for financing will be felt by the monetary policy.

To sum up the precise paths followed by inflation and bonds depend on the assumptions made about deficit financing and the relation between the real interest rate and the output rate of growth. In this section, we assume either a constant growth rate of money with the remainder deficit financed by borrowing or constant shares of the deficit financed by borrowing and seigniorage. Another possibility is that the government collects a constant amount of seigniorage revenue and finances the remainder by borrowing. This will be discussed in the section 4.

3. The Brazilian monetary regime

In the monetary regime described in the previous section the monetary control may be conducted through open market operations, discount loans or reserve requirements. In this section, we turn our eyes to an economy where the monetary policy is conducted indirectly, through a "peculiar" form of open market transactions. In addition, this is an economy with a consistently increasing inflation. In this sense, we are relaxing some hypotheses of the model depicted above and taking distance from its main characteristics.

In Brazil, the monetary regime was different from what was described in the previous section. First, the Treasury financed itself directly through the Central Bank. Second, 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 the bonds quantity 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, they could resell their holdings of excess bonds to the Central Bank.

The repurchase agreements were necessary because free reserves were expensive to the banks. If the banks had to keep the government securities until maturity, and the Central Bank did not provide (inexpensive) liquidity to the system, the banks would have had to hold a much larger volume of free reserves (within an inflationary environment), or they would have resorted more often to the discount loans. This would have been unbearably costly to them.

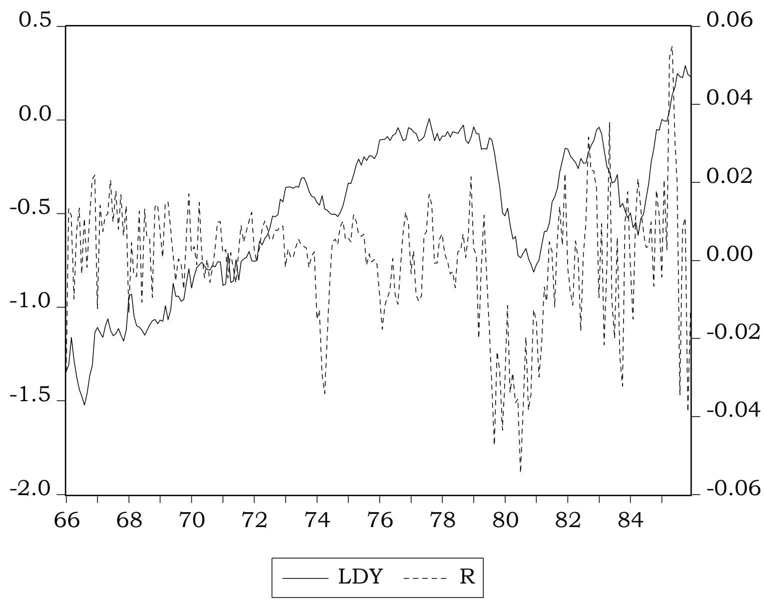
The main consequence of this procedure was the elimination of the open market operations as an instrument of monetary policy. The money supply was controlled indirectly by increasing the interest rate to expand the demand for bonds. This procedure was efficient from 1974 to 1985, as we will demonstrate later.

In the monetary regime described in section 2, when the Treasury sells bonds, the real interest rate and the public debt rise simultaneously. It is not possible to settle the causal direction in

Granger's sense. The Brazilian demand for bonds increased only when financial intermediaries bought bonds in primary auctions (to warrant their overnight deposits increased by the higher real interest rate); see Pastore (1990). Therefore, we may postulate that the increase in real interest rate came before the expansion of the debt stock. If this assertion is true, the real interest rate caused the public debt oscillations.

In figure 1, we illustrate the monthly debt-GDP ratio[8] (LDY) and the ex-post real interest rate (R), computed by discounting the inflation rate from the nominal interest rate on overnight deposits. We note that the real interest turning points came before the bond stock turning points.

FIGURE 1: DEBT-GDP (LDY) AND REAL INTEREST RATE (R)



For a more formal test of the above hypothesis, we begin by verifying the integration order of each series. The results are taken from Cerqueira (1998) and are shown in Table D.1. The ADF and the Phillips-Perron tests point out that the real interest is I(0) and the debt-GDP ratio

[8] The ratio is defined as the government bonds stock held by private agents over the nominal GDP; see appendices A and B for further details

is I(1) in the analyzed period. In this case, there is no meaning in looking for a cointegration relation.

Our strategy was therefore to begin the causality analysis by estimating a VAR with the debt-GDP in first difference and real interest rate in levels and by plotting impulse-response functions.

The first basic criterion in selecting the appropriate lag-length is the absence of serial correlation in the residuals. This can be diagnosed by the Ljung-Box and Breusch-Godfrey tests. These tests led us to conclude that only systems with 3 or more lags were potentially properly specified. We decided to start on sequential testing with a sixth order VAR. We then tested that the coefficients corresponding to the largest lag are zero, using the likelihood ratio statistics (LR) and information criteria.

The results are reported in Table D.2. For the Schwarz and Hannan-Quinn criteria, the optimal lag length is three. If we consider only orders between 4 and 6, all criteria will choose a lag length of four. It is well known that omitting relevant lagged values of the dependent variable can inflate the coefficients of the lagged “independent” variables (in a causality test); see Harvey (1990). In addition, recent findings by Giles and Mirza (1998) suggest that the pretesting procedures can result in severe overrejection of the noncausality null hypothesis while the overfitting method results in less distortion in the empirical power. We decided to work with four lags on the VAR[9].

Table D.3 shows the VAR diagnostics tests. The presence of heteroskedasticity and the lack of normality in the residuals system are striking. Since these problems are mainly caused by large outliers concentrated in some parts of the analyzed period, one way of correcting them is to model the error terms as a multivariate generalized autoregressive conditional heteroskedasticity process of order (r,m), VGARCH(r,m); see Harvey (1990) and Hamilton (1994). An alternative

[9] We advert this choice must be taken with careful since we performed the selection under the violation of the normality hypothesis; see diagnostics reports in appendix D.

procedure for handling the residual non-normality is to simulate the causality test statistics with the bootstrap approach; as in Li and Maddala (1996) and Giles and Mirza (1998). This alternative is implemented to validate the results obtained from the GARCH models or if the latter are explosive or fail to converge.

The causality tests are carried out using a LR ratio, Wald, and/or an F statistic, whose distributions depend on the assumption of Gaussian error terms; see Green (1990). So theoretically, the violation of the normality hypothesis may invalidate the causality tests. Therefore, our selection criterion was to choose the more parsimonious model that led to residuals NIID. As indicated by the diagnostics tests, a GARCH(1,1) process with a long run component seemed to be a suitable representation. The estimates and diagnostics tests are reported in Table D.4.

Before presenting the results of the causality tests we will discuss the impulse-response function computed for the VAR models without any treatment for the lack of normality. In figure 2 and 3 we show the responses of debt-GDP first difference (DLDY) and real interest rate (R) to a one standard deviation orthogonalized impulse in both endogenous variables (following the order R LDY[10]). Estimated two standard error bounds are depicted as dotted lines.

In the VAR(4), a one-time shock on real interest leads to a transitory increase in the public debt-GDP growth, with a peak of 1.1% in the fourth period. Around the 11th after-shock period, the series returns to its previous equilibrium value of zero. In the VAR(3), the effect on the debt-GDP series is weaker but the return to the equilibrium is not as quick. On the other hand, the apparent effect of an innovation in the

[10] In our case, the order of equations does not significantly change the impulse responses. In the fourth order VAR, the correlation coefficient between the two residuals is 0.189, with a p-value of 0.505, thus it is not significant at either 5 or 10%. We can also construct a 95% confidence interval around the residuals inner product (=0.028)

debt stock on real interest rate that appears in the 5th period is not significant (p-value of 0.386[11]).

FIGURE 2: Response to One S.D. Innovations ± 2 S.E. - VAR (3)

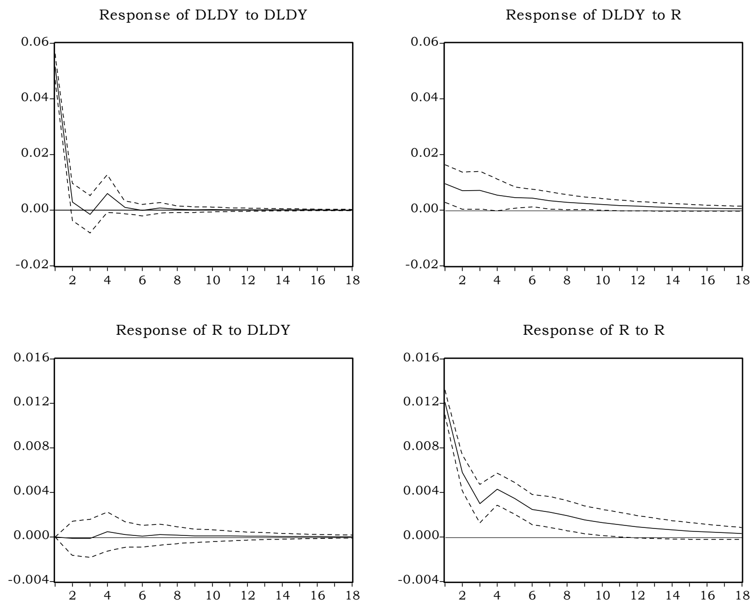
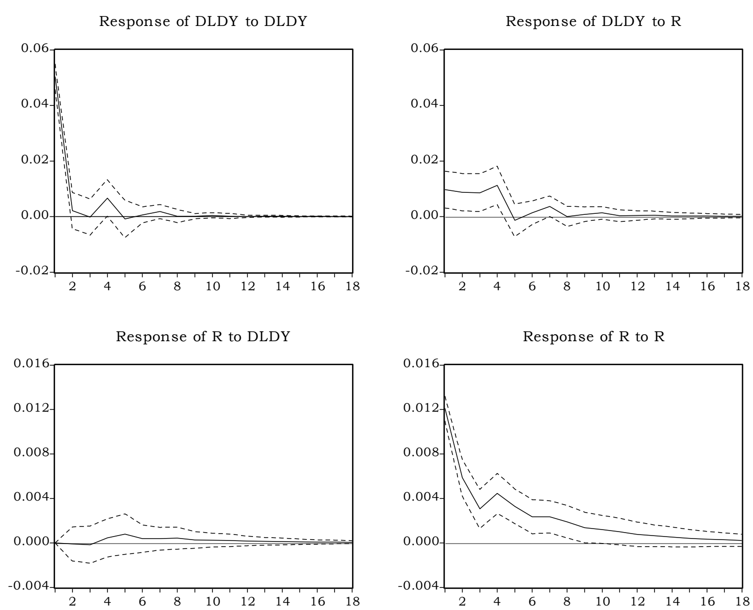


FIGURE 3: Response to One S.D. Innovations ± 2 S.E. - VAR(4)



with bounds [-0.278; 0.333]. This shows the consistency of the residual orthogonality hypothesis. Similar conclusions apply to the VAR(3).

[11] With higher order VAR (5 and 6) this effect increases, but it is not corroborated by the causality tests and it is marginally significant.

These results strongly suggest that the real interest rate causes the public debt and there is no feedback between the bonds stock and the real interest rate. Furthermore, a shock to the real interest rate leads to a permanent increase in the demand for bonds, which induces a decrease in the long run real interest rate.

In Tables 1 and 2, we present the causality tests results. In Table 1, the tests are reported in spite of the violation of the normality and homoskedasticity hypotheses. In Table 2, either the tests were performed observing the spherical conditions or empirical distributions were simulated to the test statistics. Each column designates the dependent variable in the respective test.

We implemented the Granger direct test with the VAR(4) specification described earlier. For this test, the lack of normality was corrected with an ARCH error specification[12], which is equivalent to a (nonlinear) restricted GARCH(2,2). Below Table 2, we also reported the Monte Carlo p-values simulated using the bootstrap method[13].

The results show the null hypothesis that R fails to cause LDY is rejected with a p-value of 0.1%, while LDY fails to cause R is accepted with a p-value near 60%. Thus, the tests demonstrate that the causality goes unidirectionally from the real interest rate to the bonds stock.

For comparison purposes, we also performed the Geweke, Meese and Dent (GMD) and the Sims causality tests; see Hamilton (1994). The diagnostics tests are reported in Tables D.5 to D.7.

[12] This model allows the (conditional variance) mean reversion to a varying level q_t. This component is interpreted as the time varying long run volatility (that converges to the mean of the conditional variance) and its coefficient as the associated persistence.

[13] We bootstrapped the causality test statistics by simulating each dependent variable with the VAR(4) coefficients. The error terms were drawn from a multivariate normal distribution with covariance matrix taken from the VAR estimates. The number of replications was set to 1000.

TABLE 1: CAUSALITY TESTS (P-VALUES) WITH NON-GAUSSIAN RESIDUALS*

PROCEDURE	GRANGER DIR. TEST		GMD TEST		SIMS TEST	
NULL HYPO.	LDY	R	LDY	R	LDY	R
$R \nRightarrow LDY$						
F	0.001		0.076	0.002	0.040	0.022
LR	0.009		0.062	0.001	0.030	0.016
$LDY \nRightarrow R$						
F		0.883	0.923	0.541	0.555	0.655
LR		0.876	0.915	0.501	0.520	0.623
$R \nLeftrightarrow LDY$						
F			0.013	0.013	0.004	0.004
LR			0.010	0.010	0.003	0.003

*The symbol $\nRightarrow$ means “does not Granger cause”.

TABLE 2: CAUSALITY TESTS (P-VALUES) PERFORMED WITH GAUSSIAN RESIDUALS OR MONTE CARLO SIMULATIONS

PROCEDURE	GRANGER DIR. TEST*		GMD TEST		SIMS TEST	
NULL HYPO.	LDY	R	LDY	R	LDY	R
$R \nRightarrow LDY$						
F	0.001		0.002	0.004	0.039	0.033
LR	0.001		0.007	0.010	0.039	0.033
$LDY \nRightarrow R$						
F		0.548	0.691	0.614	0.558	0.681
LR		0.629	0.628	0.520	0.558	0.681
$R \nLeftrightarrow LDY$						
F			0.018	0.243	0.005	0.006
LR			0.001	0.210	0.005	0.006

*MONTE CARLO P-VALUES: $R \nRightarrow LDY \begin{Bmatrix} F = 0.001 \\ LR = 0.001 \end{Bmatrix}$, $LDY \nRightarrow R \begin{Bmatrix} F = 0.866 \\ LR = 0.866 \end{Bmatrix}$.

The GMD test was specified using the series in the same form as in the VAR system. To the equation which has the debt-GDP first difference as dependent variable we began by specifying the regression with 6 lagged first-difference of the dependent variable and 6 leads and lags of the real interest. The Schwarz and Hannan-Quinn criteria chose the autoregressive model of fourth order. The lack of normality was solved with a nice ARCH(3) specification for the residuals. In the other equation, we selected the most parsimonious specification that supported the absence of serial correlation in the residuals. This turn out to be a regression with 5 leads and lags. Furthermore, we opted for accepting a somewhat overfitted model. The residuals were specified with a GARCH(1,1) process.

The results are similar to the Granger direct tests whether we take the debt-GDP or the real interest rate as the dependent variable.

However, it is worth mentioning some present differences. In the LDY equation, correction for the residuals' lack of normality strengthens the unidirectional causality and it confirms that the debt stock fails to cause the real interest rate. In the other equation, there are evidences of instantaneous causality[14] from the real interest rate to the debt stock. We conclude that the GMD tests indicate that the real interest rate causes the debt-GDP ratio and there is no feedback between them.

To perform the Sims' causality test we first pre-whitened both series using the autoregressive process found by the debt-GDP unit root test[15]; see Hamilton (1994). We employed an AR(4) process with a seasonal dummy corresponding to the fourth month of the year. For the test equations, we chose to take the same number of lags and leads (six) based on t-tests about the significance of the last lag or lead. Table D.7 shows the residuals diagnostics tests from the causality test regressions. As required, the accomplished tests with the pre-filtered series have no serial correlation. In the present case, the lack of normality and homoskedasticity is not very serious. However, for the sake of comparison we decided to perform Monte Carlo simulations[16] to derive the empirical distributions of the tests.

When the debt stock is regressed on the real interest of rate, the future values of R are clearly insignificant and its past values are significant at 4%. On the other hand, if we regress R on LDY only the debt stock future values are significant[17]. These empirical findings

[14] Nevertheless, the occurrence of instantaneous causality may be a function of the data and in general, it is not possible to differentiate between instantaneous causation in either direction or instantaneous feedback. So, the idea of instantaneous causality may be of little or no practical value; see Granger and Newbold (1994).
[15] If we had used the respective DGP for each of the series, we would have obtained the same results concerning the causal relation between them.
[16] Since we are handling regressions that have small F-statistics and heteroskedastic residuals, a GARCH specification leaves little "space" for performing LR or F tests and, so consequently, causality tests.
[17] The trials with more parsimonious specifications yielded stronger conclusions than those reported. We also remarked that the results above indicate evidences about instantaneous feedback, but they do not matter in our context.

demonstrate the hypothesis that causality was unidirectional from the real interest rate to debt-GDP ratio, during 1966 to 1985[18]. Technically speaking, the bond stock held by the public could be expressed as a distributed lag of current and past values of real interest (with no future values of R), with a disturbance process that was orthogonal to past, present, and future R's given that LDY did not Granger-cause R. In this sense we may assume the real interest rate was strictly exogenous with respect to the demand for bonds; see Sargent (1987).

Accordingly, we conclude that the monetary authority changed the real interest rate to induce variations on the public bond demand. Moreover, given that the public debt is a non-monetary liability, this mechanism worked as an instrument for controlling the monetary supply.

3.1 Public bonds demand

The final step is then to specify a public bond demand. The real interest rate and debt-GDP series have different integration orders and therefore can not be cointegrated. In addition, the real interest rate caused the debt-GDP ratio, but the latter failed to Granger-cause the real interest. Thus, an AD(p,q) equation may be used to model the existing short run relation between these variables.

We use as point of departure, the VAR specification described earlier. Thus, we specified an AD(6,6), without correction for normality. The debt-GDP is transformed in first difference and the real interest rate remains in levels. We added a linear trend term in order to consider

[18] We should comment that performing the causality tests under the verification of the spherical conditions (residuals NIID) or with a procedure that replaces the verification of these hypotheses did not lead to the change on the main conclusions. We conjecture this is due to the clearness of the causal relation between the variables, what can be deduced from the result robustness to the type of the employed noncausality test. Anyway, this suggests that in the present context one should not be concerned about

potential demand shifting. We proceeded with the information criteria and the last significant lag analysis. Thus, we reduced the model to an AD(4,4) whose residuals do not present serial correlation; see Table E.1. In order to obtain NIID residuals we applied an ARCH(3) specification, a choice that was supported by diagnostics tests (see Table E.2).

Since our focus is on the demand for bonds we reparametrized the estimated model to a fully AD(4,4)-ARCH(3) in levels. The estimation results are reported in the Table 3. Bollerslev-Wooldrige standard errors were used in order to be conservative, given that there is an integrated series among the regressors.

The hypothesis that the lagged dependent variable coefficients sum up to one is rejected with a p-value near 0%. The diagnostics tests indicate that the residuals are NIID and are approximately white noise. We have no reason to doubt the model's stability.

TABLE 3: BONDS DEMAND ESTIMATES
DEPENDENT VARIABLE: LDY*

VARIABLES	Coeff.	Prob.
C	-0.042745	0.0002
T	0.000209	0.0000
R	0.793968	0.0000
R(-1)	0.173784	0.2987
R(-2)	0.462111	0.0938
R(-3)	0.341945	0.1194
R(-4)	-0.675880	0.0002
LDY(-1)	0.984878	0.0000
LDY(-2)	-0.017434	0.8202
LDY(-3)	0.110946	0.1577
LDY(-4)	-0.119647	0.0246
VARIANCE EQUATION		
c	0.001528	0.0000
u_{t-1}^2	-0.078758	0.0000
u_{t-2}^2	0.239366	0.0084
u_{t-3}^2	0.215110	0.0147

R^2=0.984	F=0.000	SER=0.052	AIC=-3.142	SIC=-2.922	DW=1.904
Q(24)=0.285	Q(36)=0.747	ARCH(1)=0.649	ARCH(4)=0.868	Q^2(18)=0.757	BJ=0.157

*P-values calculated with Bollerslev-Wooldrige standard errors. KS=0.0541(10% critical value=0.0880).

heteroskedasticity or non-normality. Further, it points out the bootstrap approach did not provide a better improvement on the asymptotic theory.

The estimates show the bonds demand was very sensitive to real interest rate variations[19]. The hypothesis of a long run elasticity superior to one is accepted with a p-value superior to 69%. Over the period between 1966 and 1985, the real interest rate changed from a minimum of –5.4% to a maximum of 5.5%, with an almost zero mean of 0.01%. The debt-GDP ratio fluctuated between 21.8% and 133.6% with a mean of 65.5%[20]. An increase of 5 percent points in real interest rate would have led to an increase of 4% in the debt-GDP ratio during the same month. Taking the ratio's mean value it would have increased to 68.1%, this would have financed an operational deficit with respect to the GDP of 2.6%. A five-month growth of 1 percent point in the real interest would have induced an increase of 1.1% in the debt-GDP ratio which, being in its mean value, would have grown to 66.2% and would have financed an operational deficit-GDP of 0.7%.

This demonstrates that the public debt was elastic regarding the real interest rate and the overnight interest rate was an efficient instrument to stimulate public bond sales. They are also empirical facts that suggest an independent behavior of the monetary authority, rather than a passive monetary policy and an endogenous money supply.

4. Financial innovations, money demand and seigniorage collection

Financial innovations in general produce new assets warranted by public bonds. These assets have higher liquidity and lesser risk of capital loss than bonds. Hence, for a given level of deficit financed with seigniorage, they cause higher inflation, because they contract the

[19] This conclusion is an indirect way of stating that during 1966 to 1985, the assumption of Ricardian equivalence was not empirically verified.
[20] Note that with monthly frequency data this ratio is greater than otherwise. Because even if the bond stock had a constant quarterly value, the flow of monthly output is smaller compared to its quarterly aggregated value.

monetary demand and increase the money income velocity (given the output, the real interest rate, and the inflation expectations).

In Brazil, the Central Bank repurchase agreements, which began at 1976, increased the intensity of the innovation process. The repurchases almost reduced to zero the risk of capital loss. They also gave the financial institutions an almost instantaneous liquidity to any unbalance between their bond holdings and liabilities, at a price near the bonds' yield curve. This mechanism gave the government's securities a degree of liquidity close to primary money and then led the public debt to crowd-out the demand for money[21].

Even without the presence of financial innovations, the search by the monetary authority of a given level of seigniorage may lead to an ever-increasing inflation. An increase in the expected inflation rate reduces the desired money stock, which can increase or reduce the inflation tax depending on whether the economy is on the upward or on the downward segment of the Laffer curve. However, the seigniorage component $(\overset{\circ}{M/P})$[22] can be negative if the monetary authority does not increase the monetary expansion simultaneously with the increase in the expected inflation. If the economy is operating on the downward portion of the Laffer curve and if the inflation expectations are increasing, then a growing rate of monetary expansion is necessary to keep constant the level of seigniorage collection. This will produce an ever-increasing inflation rate.

Along the upward side of Laffer curve, a contraction in the money demand will lead to a higher inflation rate. The contractions of money demand may reach a point at which the maximum inflation tax falls

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[21] This claim is best explained with Baumol's money demand: $M^d = p\sqrt{cy/2i}$. Given the real income (y) and the nominal interest rate (i), a reduction in the transaction cost (c) in the conversion of a financial asset into money reduces the money demand.
[22] The seigniorage is the sum of the inflation tax with the growth of the real monetary base. Differentiating the real base (M/P) with respect time, we get: $\frac{\dot{M}}{P} = \frac{M}{P}\pi + (\overset{\circ}{M/P})$.

short of the public deficit financed with money creation. In this case, the only way to keep the seigniorage collection constant is by continuously accelerating the monetary expansion and hence the inflation. Consequently, inflation grows indefinitely as the monetary authority collects seigniorage by accelerating inflation[23].

Is there evidence that the demand for money contracted in such a way to limit the seigniorage collection and cause the money supply endogeneity?

Cerqueira (1993) estimated the demand for real balances for 1966-1985 period, using quarterly data. We showed that, owing to financial innovations, the money demand experienced a contraction[24] between 1976 and 1984, which reduced the monetary authority's ability to collect seigniorage. This is supported by our estimates of the maximum seigniorage that could be collected as a proportion of GDP which decreased from 4.5% per quarter in 1976 to 3.3% in 1985 (using the average base multiplier between 1976 and 1985[25]). Consequently, the money velocity increased from 1.8 to 7.1 between 1975 and 1984/85, while the inflation rate grew from an average monthly rate of 2.2% to 10.5% during the same period.

However, if we had considered the base multiplier from 1976/78 we would have estimated a maximum seigniorage collection of only 2.6%. This could have led the monetary authority to increase money growth. Thus, the relationship between financial innovations and inflation might be characterized by a potential feedback coming from the money growth

[23] These arguments are described more fully in Cagan (1956) and Blanchard and Fisher (1989) ch.4.

[24] The reduction in the money demand is represented by the decrease in its constant term; see equation (20) ahead.

[25] The maximum seigniorage is given by $S^*=c/\alpha ek$ (see development ahead), where k is the base multiplier. Since M1=kB, where B is the monetary base, the inflation tax-GDP ratio is given by $IT=\pi\frac{B}{PY}=\pi\frac{M1}{PY}\frac{1}{K}$. This means the government collects a fraction 1/k of the produced inflation tax by the real money balances. The difference (1-1/k) represents the inflationary transfer inside the private sector to the banking system.

increase. To provide a formal exposition of this argument we begin by supposing an economy described by the following version of Cagan's (1956) model:

$$m \equiv \frac{M}{PY} = ce^{-\alpha\pi_e}\text{, } c > 0 \text{ e } \alpha > 0 \text{ (money demand equation)} \qquad (20)$$

$$\overset{\circ}{\pi}_e = \beta(\pi - \pi_e)\text{, } \quad \beta > 0 \text{ (expectation rule equation),} \qquad (21)$$

where π_e is the expected inflation rate, α is the semi-elasticity of the money demand with respect the expected inflation and β is the inverse of inflationary memory (the bigger is β the smaller is the effect of past inflation on the present inflationary expectations). We assume a constant real output and a constant real interest rate. For a given level of exogenous money growth μ, the seigniorage flow is given by:

$$S = \frac{\overset{\circ}{M}}{PY} = \frac{\overset{\circ}{M}}{M}\frac{M}{PY} = \mu ce^{-\alpha\pi_e}\text{ .} \qquad (22)$$

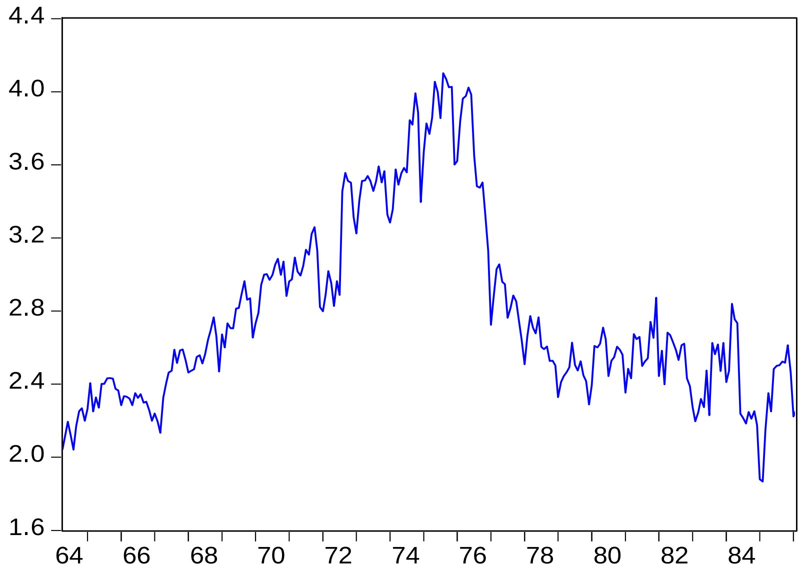
In steady state $\overset{\circ}{\pi}_e$=0 and $\mu = \pi_e = \pi$ and the inflation tax equals the seigniorage. The seigniorage is maximized (S*=c/αe) when $\pi = 1/\alpha$. With a constant operational deficit at level $S = \bar{S}$, the monetary authority will react according to:

$$\mu = \frac{\bar{S}}{ce^{-\alpha\pi_e}}\text{ .} \qquad (23)$$

The monetary expansion rate is increasing with the expected level of inflation and then is passive. Thus, a reduction in the constant term c, caused by financial innovations, shifts down the reaction curve that results in increases the monetary expansion and the inflation rate.

We conjecture that the reduction in the ability of collecting seigniorage was the reason that induced the monetary authority to reduce the base multiplier, in order to increase its fraction on the seigniorage collection. Actually, Cerqueira (1993b) showed that the reduction in the constant term of the money demand generated a reduction in the desired holdings of real balances of nearly 27.9%. Meanwhile the base multiplier was reduced by 27.4%. This implied an increase of 37.7% in the government-collected proportion of the inflation tax, which more than compensated the first effect.

Figure 3: Base Multiplier



Therefore, the fact that the seigniorage collection was constant during the period did not mean that the monetary policy was passive. It also explains why the estimated steady state inflation rates[26] remained

[26] If we combine equation (20) and (21), we get the equation about the dynamics of the expected inflation rate: (24) $\overset{\circ}{\pi}_e = \frac{\beta}{1-\alpha\beta}(\mu - \pi_e)$, whether money creation is exogenous or not. If S is less than S* there are two equilibria: π_e^1 which is the low-inflation equilibrium and, π_e^2, the high inflation equilibrium. If the Cagan stability condition $\alpha\beta < 1$ is respected, for a initial expectations $\pi_e^0 < \pi_e^1$, π_e converges to π_e^1, a stable target, since $\overset{\circ}{\pi}_e < 0$. If $\pi_e^0 > \pi_e^2$ then π_e increases indefinitely ($\overset{\circ}{\pi}_e > 0$). For this reason, π_e^2 is called the unstable equilibrium point. If S=S* then there is only one steady state point $\pi_e = 1/\alpha$.
Under rational expectations hypothesis the inflationary memory coefficient goes to zero and so β goes to infinity. From the difference equation above, the expected inflation dynamics is now given by: (25) $\overset{\circ}{\pi}_e = \frac{\pi_e - \mu}{\alpha}$. Now π_e^1 is the unstable equilibrium and the economy converges along the slippery side of Laffer's curve to the high inflation rate π_e^2.

approximately constant from 1976 to 1985. In the end, the policy of reducing the base multiplier has a similar effect to cutting the public deficit financed with issuing money.

From 1976 to 1985, the host of financial innovations restricted the government ability to collect seigniorage. However, the policy of reducing the base multiplier counterbalanced this effect. Moreover, even if the monetary authority had increased the money growth to make up for this effect, the observed inflation and monetary rates, while increasing, were very distant from the estimated unstable levels, and thus far from the path of a hyperinflationary disequilibrium[27]; see Cerqueira (1993).

5. Money supply exogeneity

The money demand contractions restrict the seigniorage collection and increase the observed and the expected rates of inflation (in Cagan's model). If the monetary authority wants to maintain the previous level of seigniorage, it will have to increase monetary expansion, which will render money creation endogenous. Moreover if the expected inflation is predicted by extrapolating lagged rates of inflation, then money creation is caused in Granger's sense by the rate of inflation. Lagged rates of money creation add nothing to the predictions of inflation and their own

This is the same solution as if the Cagan's stability condition were being violated. Consider now the Bruno and Fischer (1990) model with variable parameters. Suppose β is a positive function of the expected inflation, thus the inflationary memory shrinks as the inflation increases. The dynamics of the expected inflation is given by (24) if the expectations are adaptive and by (25), if the agents have perfect foresight. However, under adaptive expectations we cannot say a priori anything about what will happen to the expectations. Then suppose that there is π_e^* such that $\alpha\beta(\pi_e^*)=1$. So, given some level of seigniorage the convergence of π_e depends on the expectations position in the initial time (i.e., π_e^0). Hence if $\pi_e^0<\pi_e^*$ the stable inflation rate will be π_e^1, but if $\pi_e^0>\pi_e^*$ the economy will converge toward the high inflation trap π_e^2.

[27] In 1984/85, the observed average values of the base monetary growth and rate of inflation were respectively, 35.6%% and 33.9%. Meanwhile the estimated stable and unstable inflationary levels were respectively, 12.6% and 124.6%.

expansion rates. This is a model in which Cagan's adaptive scheme is rational (Sargent and Wallace, 1973).

On the other hand, it may be possible that the monetary expansion is exogenous with respect to inflation and uncorrelated with the random terms in the demand function for money. In this case, money creation influences current and future rates of inflation; but given past lagged rates of money creation, past values of inflation exert no influence on money creation. The model is one in which money creation causes inflation, while inflation does not cause money creation. In such a model, adaptive expectation schemes like Cagan's are not rational.

In the last section we postulated a reaction curve (equation (23)) in which the money supply was passive. However, there is no hard evidence that proves that the monetary authority followed this kind of rule.

In fact, the Central Bank repurchase agreements facilitated, but did not necessarily implied an endogenous money creation. If the inflation rate increased due to a negative supply shock and the Central Bank target was to keep the real interest rate constant, it would have to buy bonds through the repurchase agreements. Hence, it would accommodate the price increase and the money supply would be endogenously determined (or caused) by the inflation rate. This outcome comes about because of the interest rate policy and not because of the repurchase agreements per se. Money endogeneity emerges regardless of the monetary regime, as long as the goal is to keep the real interest rate constant. If the target were to control the money supply with a fluctuating real interest rate, this causal relation would not necessarily take place.

Figure 1 shows large real interest rate fluctuations[28] between 1966 to 1985. It was argued in section 3 that these fluctuations caused the public debt stock. Since the debt stock held by the Central Bank is a

[28] During the period the real interest rate followed an AR(2) process meaning its past values had information about its contemporaneous behavior.

non-monetary liability, it is an instrument of monetary control. This finding is in contradiction with other authors' empirical studies who found that money creation was caused by inflation; see Marques (1983), Triches (1992), Pastore (1994), and Pastore (1997).

In appendix C we show that the seigniorage collection as a GDP proportion could be taken as constant with a mean value of 0.61% per month or 1.82% per quarter from 1974 to 1988. This series followed a white noise process, meaning that even if it had a constant mean, its behavior could not be predicted due to the series lack of memory. Thus, the seigniorage behaved as a shock. It exhibited no relation with either the money demand contractions or the increases in expected inflation. It was shown that the inflation rate had no impact on the seigniorage–GDP ratio; see Table C.3.

Even if the monetary authority was attempting to finance a roughly constant rate of real expenditure by money creation, the seigniorage collection did not follow a path consistent with endogenous money creation. The policy of reducing the base multiplier in order to maintain the effectiveness of the inflation tax contributes to the plausibility of an exogenous monetary policy.

Accordingly, we assume that there was a public deficit permanently financed with money creation that assured the debt sustainability while preserving the exogeneity of the money supply. This means that there was a steady state level of public deficit financed with money creation, as in Cagan's adaptive scheme[29]. To complete the proof of this hypothesis it is "necessary" to show that money creation caused inflation, and that inflation did not cause money creation (Sargent and Wallace, 1973).

The first step in looking for a causal relation is to determine the integration order of the series. Table F.1 reproduces the unit root tests

[29] That is the seigniorage collection equaled the inflationary tax-GDP ratio.

for both variables, taken from Cerqueira (1998). The Phillips-Perron test rejects the null of a unit root in the series levels. On the other hand, the ADF test indicates both series are difference stationary. In Cerqueira (1998) using the bootstrap approach on the ADF statistics, it was concluded that the unit root hypothesis could not be rejected for both series. The computed p-values were 0.999 and 0.443, respectively for the rate of inflation and the money growth. In spite of the mixed evidence, we will treat both series as I(1).

The next step is to perform cointegration tests. We used the likelihood-based cointegration tests of Johansen (1991). It is well known that the results of cointegration tests using this technique depend on the deterministic components included in the VAR and on the chosen lag length. Therefore, some pre-testing was done to insure a proper specification.

One striking characteristic that appeared in our first trials was the presence of large outliers, which induced leptokurtic distributions and caused the VAR residuals to violate the normality assumption. Since the Johansen's procedure is based upon the Gaussian likelihood, large deviations from Gaussian white noise can invalidate the test conclusions. This problem may cause spurious cointegration findings (Lee and Tse, 1996). On the other hand, Cheung and Lai (1993) argue that Johansen tests are reasonably robust to excess kurtosis. In our initial trials, the post-estimation weak exogeneity tests were very sensitive to outliers, which warrants caution in interpreting the results of the weak exogeneity tests.

For this reason, we chose to introduce a battery of policy dummies[30] acting as exogenous variables and excluded from the long-run relation in the Johansen procedure. This changes the asymptotic

[30] The dummies are defined as follows: D661=1 if t=1966.01; D8112=1 if t=1981.12; D836=1 if t=1983.06; D842=1 if t=1984.02; D8485=1 if t=1984.12 and –1 if t=1985.01;

distributions but it is a common procedure to account for important short-run effects that have to be controlled to lessen violations of the Gaussian assumption about the stochastic part of the process (Johansen and Juselius, 1992; Hansen and Juselius 1995).

The pretesting also identified intrinsic explosiveness in the data by detecting roots lying outside the unit circle. This kind of nonstationarity can not be removed by differencing and is an indication of the inadequacy of the chosen model (Johansen, 1995). One alternative for handling this problem is to use dummy variables like those mentioned above. Another possible method is to employ a dummy-type variable like a linear trend term. Pretesting suggested the inclusion of a linear trend in the cointegration space and of exogenous policy dummies to address this nonstationarity issue.

To choose the lag length we searched for parsimonious models without autocorrelation. The number of lags necessary to respect this condition is in the interval between 13 and 17[31]. Then the lag length was selected using two types of information criteria (Schwarz and Hannan-Quinn)[32]. The pretest results are shown in Table F.2.

Both criteria chose sequentially the more parsimonious VAR models. Nevertheless, there are some arguments against this choice. Monte Carlo results in Gonzalo (1995) show that the efficiency loss is small for overfitting, while consistency is lost if the lag length chosen is too small. Dolado and Lütkepohl (1996) point out that the loss in power in the Wald test caused by extra unnecessary lags, is likely to be relatively small if the true order of the VAR, k, is large and the dimension of the system, n, is small. The intuition is that one or two extra lags on all of the variables are not likely to drastically reduce the estimation

D85=1 if t=1985.08, -1 if t=1985.09, and 1 if t=1985.11; Dum84=1 if t≥1984.01; 0 otherwise.

[31] Below 12 and between 18 and 21 lags all estimated models showed problems in multivariate and/or univariate Ljung-Box tests.

precision. Recent findings from Giles and Mirza (1998) suggest that the pretesting procedures can result in severe overrejection of the null of noncausality while overfitting methods cause less distortion with often little or no loss of power. Their suggestion is to abandon pretesting for cointegration in favor of a more straightforward overfitting method when unidirectional causality is suspected. Finally, the system normality tests (considering the unrestricted VAR and the restricted VECM models) are more problematic for the specification with 13 lags than for higher ordered VARs. We then chose to work with 15 lags.

The results of the cointegration test are presented in Table F.3. The inclusion of dummy-type variables changes the asymptotic distributions, which may lead to a problem of under rejection of the null of no cointegration. One way to deal with this is to increase the significant level from the usual 5% or 10% levels to the 15% level.

Thus at 15% we reject the null of no cointegration vector and can not reject the null of one cointegrating vector. Although we used the 15% level, it is worth mentioning that the λ_{max} statistics rejects the null of zero cointegration vectors at the 5% level[33].

The inspection of the roots of the companion matrix (see figure F.1) indicates there are two roots very close to 1 (-0.9653 and 0.9128), but none of the others are close to other points on the unit circle. Thus, the non-stationarity can be removed by differencing and we can proceed the analysis of the VAR with 15 lags.

The presence of cointegration corroborates the hypothesized absence of rational bubbles between 1966 and 1985[34]. Furthermore, it

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[32] The strategy of determining the lag length with sequential a likelihood ratio statistic led to a number of lags inferior to 13. Therefore, we did not report it.

[33] Without the dummy variables, the trace statistics would reject the null at the 5% level. Nonetheless, testing the battery of policy dummies when the number of cointegrating vector is one, we get the MLR statistics $\chi^2(14)=153.564$, which implies a p-value very near zero. Then we reject the restricted model without dummies.

[34] We note that cointegration relation is robust to changes on the sample data. Either we extend the sample until 1964.01 or until 1990.03, the presence of a long run relationship remains.

rules out any non-stationarity in the unobserved variables since it can be eliminated by differencing (Diba and Grossman, 1988; Welsh, 1991). Testing for plausibility of the (1,-1,#,#) cointegrating vector confirmed that the long run relation between money growth and inflation is characterized by homogeneity. This is the "classical" representation, in which the existing equilibrium relationship of inflation and money growth has cointegrating vector β=(1,-1) and a moving stationary drift term[35]. This could represent the output rate of growth an I(0) variable; see Cerqueira (1998).

Diagnostics tests reported in Tables F.4 and F.5 show that the restriction on the cointegrating space (by setting the cointegration rank equals to 1) approximates the residuals of Gaussian innovation. The cause for the lack of normality can be traced to the residuals of the money growth equation. Fortunately, this violation is not very strong and it turns out the money growth variable is weakly exogenous.

The diagnostics tests of the error correction model[36] show that our choice of 15 lags for the VAR was appropriate. First, because the residuals have no serial correlation and second, choosing the thirteen-lag VAR would have led to the estimation of a misspecified VEC model with autocorrelated residuals.

Table F.6 shows the results of two weak exogeneity tests conditioned on the existence of one cointegrating vector: the first uses the estimated vector and the second the theoretical vector (1,-1,#,#). At usual significance levels we found that money creation is weakly exogenous for the parameter of interest in the conditional model of inflation, but the reverse is not true for the inflation.

These results corroborate the implicit idea in equation (24). A monetary shock causes the acceleration of expected inflation, which by

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

this turn increases the inflation. The inflation rate drifts above its steady state path, and since the adjustment coefficient α is negative[37], the acceleration in expectation is reduced, thereby forcing the inflation rate down towards its long-run path.

Tables 4 and 5 show three causality tests[38]. Below each table, we also report the strong exogeneity tests[39]. When the residuals were not Gaussian, we relied on GARCH models, or on the simulation of the test statistics' distribution following the bootstrap approach. The Granger direct test specification is taken from the VECM estimates[40]. The methodology for performing the other tests followed the same paths described in section 3, thus all technical details are explained in the appendix.

TABLE 4: CAUSALITY TESTS (P-VALUES) WITH NON-GAUSSIAN RESIDUALS*

PROCEDURE	GRANGER DIR. TEST		GMD TEST		SIMS TEST	
NULL HIP0.	MI	PI	MI	PI	MI	PI
PI $\not\Rightarrow$ MI						
F	0.746		0.206	0.000	0.678	0.521
LR	0.574		0.149	0.000	0.648	0.477
MI $\not\Rightarrow$ PI						
F		0.001	0.002	0.001	0.005	0.005
LR		0.000	0.001	0.000	0.004	0.003
MI $\not\Leftrightarrow$ PI						
F			0.379	0.353	0.894	0.830
LR			0.348	0.333	0.891	0.823

*The symbol $\not\Rightarrow$ means "does not Granger cause". Strong exogeneity: $\pi \not\Rightarrow \mu \begin{Bmatrix} F = 0.787 \\ LR = 0.625 \end{Bmatrix}$, $\mu \not\Rightarrow \pi \begin{Bmatrix} F = 0.000 \\ LR = 0.000 \end{Bmatrix}$.

TABLE 5: CAUSALITY TESTS (P-VALUES) PERFORMED WITH GAUSSIAN RESIDUALS OR MONTE CARLO SIMULATIONS*

PROCEDURE	GRANGER DIR. TEST		GMD TEST		SIMS TEST	
NULL HIP.	MI	PI	MI	PI	MI	PI
PI $\not\Rightarrow$ MI	0.750		0.611	0.371	0.691	0.543

36 For a matter of space, we did not report the VECM estimates, which can be obtained from the author upon request.

37 The value of α was found to be -0.170.

38 Since the causality between money growth and inflation is a controversial matter, we decided to implement three different tests.

39 Strong exogeneity is the conjunction of weak exogeneity and Granger noncausality, it insures valid conditional forecasting; see Ericsson and Irons (1994).

40 For the VECM as whole, the residuals are not Gaussian, so we performed Monte Carlo simulations for the distributions of the test statistics. Given that the system's residuals are orthogonal, the error terms were drawn from a normal distribution with a diagonal covariance matrix.

F LR	0.750		0.697	0.260	0.691	0.543
MI $\nRightarrow$ PI						
F		0.002	0.050	0.000	0.007	0.009
LR		0.002	0.051	0.001	0.007	0.009
MI $\nLeftrightarrow$ PI						
F			0.131	0.026	0.896	0.828
LR			0.236	0.029	0.896	0.828

*Monte Carlo p-values of the strong exogeneity tests: $\pi \nRightarrow \mu \begin{Bmatrix} F = 0.792 \\ LR = 0.792 \end{Bmatrix}$, $\mu \nRightarrow \pi \begin{Bmatrix} F = 0.001 \\ LR = 0.001 \end{Bmatrix}$.

The results from the causality tests indicate that money growth causes inflation while the rate of inflation fails to cause the monetary expansion. The surprising exception is the GMD test that indicates a significant feedback from inflation to money when the dependent variable is the rate of inflation. Nevertheless, if the residuals are specified as a GARCH process, we obtain the same results as with the other tests[41]. This is an indication that one must be careful in respecting the spherical conditions when handling GMD causality tests.

The existence of some feedback from inflation to money is not a blow to the notion that the money supply was exogenous. This feedback may be interpreted as adjustments in monetary policy following a monetary shock. This is illustrated by the impulse-response functions[42] plotted in figure 4. The one standard deviation shock on money growth causes a period of fluctuations in this series until it achieves its new steady state point. Indeed, beyond period 27 the impulse functions are statistically different from zero, with p-values near 0%, which back this claiming.

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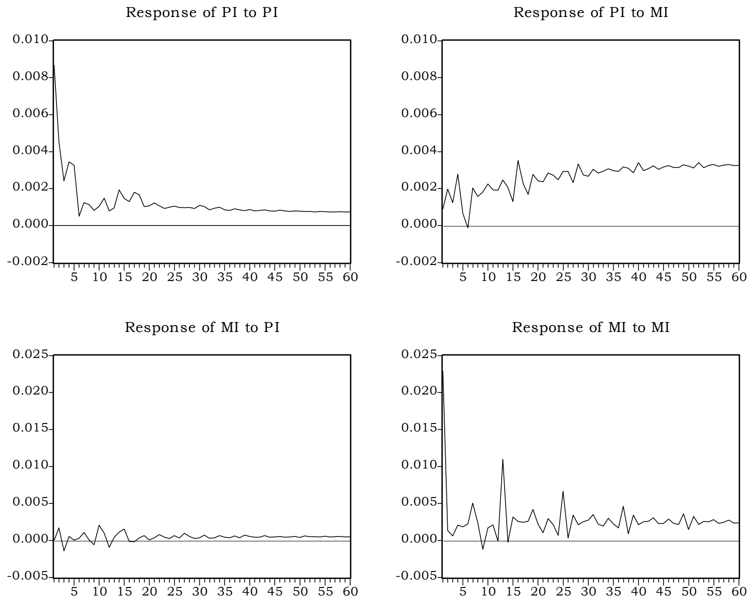
[41] In addition, from Tables 4 and 5 we note that the observation of the Gaussian assumption weakened, in the equation with the money growth as dependent variable, the causality from money to inflation.
[42] As in Lütkepohl and Reimers (1992) the impulse-response functions were calculated using (1,-1,#,#) as the cointegrating vector and the adjustment factor in the form (-0.172,0). The other VECM components are the same as those mentioned in the text. We remark the order of equations is not important in these impulse responses, due to the lack of correlation between the residuals in the VECM. The correlation coefficient between them is 0.102 and is not significant with a p-value above 40%. Furthermore, they are orthogonal since we can not reject the hypothesis of a zero-inner value (=0.005) with a p-value of 0.179.

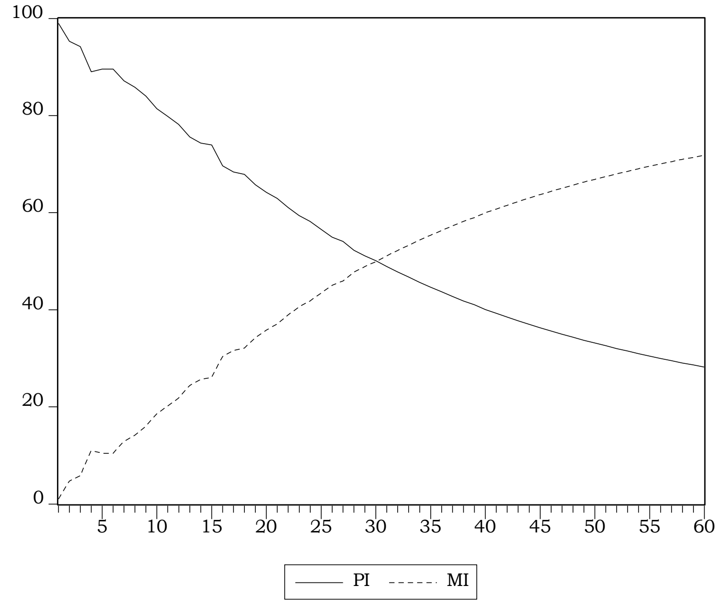
FIGURE 4: Response to One S.D. Innovations



The impulse in the monetary expansion leads to a permanent increase in the inflation rate but conversely an inflationary shock has a transitory and not significant impact on the money growth[43]. The effect of an inflationary innovation on money growth disappears after 16 months, and during this period, it has an average impact on money growth of 0.044 percent points. By this turn, a monetary shock has a permanent impact on the inflation rate near 0.33 points. When we consider innovations in either inflation or money growth, 76% of the inflation final variation is caused by a monetary shock. This last result shows a feedback effect (around 24%) on the inflation due to the expected inflation. This is illustrated by the variance decomposition of forecast error of the inflation, depicted in figure 5. Thus, we conclude that the existing “persistence” in the inflation was due mainly to monetary causes, rather than being caused by disturbances that came about in the “real side” of the economy.

[43] The persistence is measured by taking the value to which converged the inflation over its initial impulse value. The persistence equals to 8.2%.

FIGURE 5: VARIANCE DECOMPOSITION OF PI



We can conclude that there is enough evidence to validate the hypothesis that money growth is strongly exogenous[44] in what concerns the inflation rate, between 1966 and 1985[45]. Evidence shows that the causality is unidirectional and moves from money expansion to inflation. This means that the money supply was not passive and it was econometrically exogenous with respect to the price determination.

[44] Because it is weakly exogenous for the inflation parameters of interest in its conditional model, and MI is not Granger-caused by PI.

[45] If we use the amplified monetary aggregate M2 (M1 plus government bonds held by the public) as money definition, we will conclude that the money growth G-caused the rate of inflation and the inflation G-caused the money growth. Both series are I(1) and they cointegrate, but the theoretical vector (1,-1), as expected, is marginally accepted (p-value of 0.082). The hypothesis of weak exogeneity is rejected in both directions. Thus the causality relation it was only in the Granger sense, and a clear feedback occurred in both directions. These results are not surprising, since the public bonds return (the overnight deposits remuneration), during the period, where tightly associated with the rate of inflation. Moreover, due to the ever-growing inflation and the host of financial innovations, the M2 fraction corresponding to the bond stock increased steadily (from 13% in 1966.01 to 70% in the end of the period). However, these results are sharply different from the previous findings of Pastore (1997). Moreover, this conclusion has only theoretical interest and no practical meaning in explaining the money exogeneity. The reasons are that the overnight deposits were not accessible by the greatest part of the agents, during almost the whole period, and their liquidity, even being near the primary money, were far from being instantaneous. Thus, the M2 aggregate can not be taken, to this period, as the money representative concept.

Moreover, it is a strong indication that the monetary authority did not follow a reaction curve like equation (23) as a monetary rule[46].

6. Conclusions

This paper presents tests on the exogeneity of Brazilian monetary supply for the military period from 1966.01 to 1985.12, using monthly data. We chose this period because the macroeconomic policy was more homogenous regarding the inflation stabilization, than if we had included nearby years. The results show the monetary authority worked towards increasing the real interest rate to induce the public bond demand. They also show that, even if between 1974 and 1985 the seigniorage collection was predetermined in a constant level, this policy did not lead to the endogeneity of the money supply since the causality was unidirectional from money growth to inflation rate. All results show that the monetary expansion caused in Granger's sense the inflation rate[47]. This was probably possible because the MA 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 rate of inflation. Therefore, Brazilian inflation followed an ever-increasing path without exploding a hyperinflationary process.

[46] This conclusion contradicts Marques (1983), Pastore (1994) and especially Pastore (1997), but goes in the same direction of Pastore (1990). We suppose that differences in the methodologies and in data frequency can explain the divergence between our results and what could be regarded as the "accepted wisdom" among Brazilian economists.

[47] The fact that all tests achieved the same results demonstrates they are robust to the applied noncausality-type test. Moreover, our empirical trials reveal the violation of homoskedasticity and normality assumptions, in general do not affect very seriously the causality tests performance. Nevertheless, some carefulness must be taken when performing the Geweke, Meese, and Dent test.

We may conclude that money creation influences current and future rates of inflation but, given lagged rates of money creation, past rates of inflation exert no influence on money creation. This is an indication that the rational Cagan's adaptive schemes are not well fitted to the Brazilian economy. It also contrasts sharply with an existing tradition among Brazilian economists that assumes the monetary policy was completely passive during the seventies and eighties.

Indeed, our result reveal the monetary policy was executed in an independent way, i.e., the rule that guided the monetary execution was taken exogenously respecting the considered model and the rate of inflation. Therefore, we postulate that the monetary authority chose to finance a rough fraction[48] of the public deficit by issuing money, which explains the intermittent growth of the monetary expansion and the rate of inflation. This policy conduction generated a vicious cycle because, by exacerbating the already volatile 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[49].

To reject the causality from prices to money does not mean to propose there was rigid monetary control. This depended on the monetary regime. In the regime described in section 2, the monetary authority is able to fulfill almost any target of money stock. Compelled by the public deficits, it 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

[48] We estimate that the deficit proportion financed with issuing money oscillated between 50.8% and 13.7% (these values correspond respectively to 1966 and 1985), with an average value of 39.1%. We note that this decrease is due to the money demand contraction that occurred during the period, reduced the ability to collect seigniorage, and so augmented the burden of the financing with public debt.

[49] 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%.

responsibility was the burden of the Central Bank. It was enough that the real interest was restricted to fluctuate between a given range, to determine the deficit monetization through the repurchase agreements. In this regime, the monetary control was indirect and the instruments less efficient, but they were enough to manage an independent monetary policy from the inflation variations. Perhaps these are somewhat old monetarist ideas, but we can not deny they stamped their mark on the data.

References

Alesina, A. 1989. "The End of Large Public Debts", in Giavazzi, F. and Spaventa, L. (Eds.), High Public Debt: The Italian Experience". Cambridge: Cambridge University Press, 1989.

Ardeo, V. & Giambiagi, F. 1988. "PIB Trimestral a Preços Correntes para o Período 1980/1988: Resultados Aproximados". Boletim Conjuntural, IPEA/INPES no.5, out..

Banerjee, A., Dolado, J.J., Galbraith, J.W. & Hendry, D.F. 1993. “Cointegration, Error Correction, and the Econometric Analysis of Non-Stationary Data”. Oxford: Oxford University Press.

Barro, R. 1974. "Are Government Bonds Net Wealth?", in Barro, R. "Money, Expectations, and Business Cycles". New York: Academic Press, 1981.

Barro, R. 1978. "Public Debt and Taxes", in Barro, R. op. cit..

Barro, R. 1979. "On the Determination of the Public Debt", in Barro, R. op. cit..

Barro, R. 1986. "U.S. Deficits since World War I", in Barro, R. "Macroeconomic Policy". Cambridge: Harvard University Press, 1990.

Blanchard, O. & Fisher, S. 1989. “Lectures on Macroeconomics". Cambridge: The MIT Press.

Bodin, P.M. 1988. “Keynes, Sargent e o Papel da Politica Monetária em um Plano de Estabilização”. Pesquisa e Planejamento Econômico 18: 145-160.

Bruno, M. & Fischer, S. 1990. "Seigniorage, Operating Rules, and the High Inflation Trap". The Quarterly Journal of Economics: May 353-374.

Cagan, P. (1956) "The Monetary Dynamics of Hyperinflation", in Friedman, M. (Ed.), “Studies in the Quantity Theory of Money". Chicago: The University of Chicago Press.

Campbell, J.Y. & Perron, P. 1991. “Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots”. Cambridge: NBER Macroeconomics Annual

Carneiro, D. D. & Werneck, R.L.F. 1992. "Public Savings and Private Investment: Requirements for Growth Resumption in the Brazilian Economy". Working Paper 283: Departamento de economia, PUC, RJ.

Carneiro, D. D., Werneck, R.L.F. & Garcia, M.G.P 1993. "Strengthening the Financial Sector in the Brazilian Economy". Working Paper 307: Departamento de Economia, PUC, RJ.

Cerqueira, L.F. 1993. "Demanda por Moeda, Regressões Espúrias e Níveis de Instabilidade Hiperinflacionarias". Departamento de Economia, PUC, RJ, unpublished.

Cerqueira, L.F. 1998. "Bootstrapping ADF and Ouliaris, Park, and Phillips Unit Root Tests: An Analysis of the DGP's of Brazilian Macroeconomic Time Series". Département d'économique, Université Laval, unpublished.

Cheung, Y.W & Lai, K.S. 1993. "Finite Sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration". Oxford Bulletin of Economics and Statistics 55: 313-328

Contador, C.R. 1978. "A Exogeneidade da Oferta de Moeda no Brasil". Pesquisa e Planejamento Econômico, agosto.

Dagenais, M.G. & Dufour, J-M. 1991 "Invariance, Nonlinear Models, and Asymptotic Tests". Econometrica: 59: 1601-1615.

Diba, B.T. & Grossman, H.I. 1988. "Rational Inflationary Bubbles?". Journal of Monetary Economics 21:35-46.

Dolado, J.J. & Lütkepohl, H. 1996. "Making Wald Tests Work for Cointegrated VAR Systems". Econometric Reviews 15: 369-386.

Dufour, J-M. & Khalaf, L. 1994. "Simulation Based Finite and Large Sample Inference Methods in Seemingly Unrelated Regressions". C.R.D.E, Université de Montréal, unpublished.

Ensgsted, T. 1994. "The Classic European Hyperinflations Revisited: Testing the Cagan Model Using a Cointegrated VAR Approach". Economica 61:331-343.

Ericsson, N.R. & Irons, J.S. 1994. "Testing Exogeneity". Oxford: Oxford University Press.

Giles, J.A & Mirza, S. 1998. "Some Pretesting Issues on Testing for Granger Noncausality". University of Victoria, unpublished.

Gonzalo, J. 1994. “Five Alternative Methods of Estimating Long Run Equilibrium Relationships”. Journal of Econometrics 60: 203-233.

Granger, C.W.J. 1986. “Developments in the Study of Cointegrated Economic Variables”. In Engle, R. & Granger, C. (Eds.), Long-Run Economic Relationships: Readings in Cointegration. New York: Oxford Univ. Press, 1991.

Granger, C.W.J. & Newbold, P. 1986. “Forecasting Economic Time Series”. 2nd edition, San Diego: Academic Press.

Green, W. 1990. “Econometric Analysis”. New York: Macmillan.

Hamilton, J.D. 1994. “Time Series Analysis”. Princeton: Princeton University Press.

Hansen, H. & Juselius, K. 1995. “Cats in Rats – Cointegration Analysis of Time Series”. Evanston: Estima.

Harvey, A.C. 1990a. “The Econometric Analysis of Time Series”. Exeter: Philip Alan.

Harvey, A.C. 1990b. “Forecasting, Structural Time Series and the Kalman Filter”. Cambridge: Cambridge University Press.

Issler, J.V. & Lima, L.R. 1997. “Public Debt Sustainability and Endogenous Seigniorage in Brazil: Time-Series Evidence from 1947-92”. Working Paper: Graduate School of Economics – EPGE, Fundacao Getulio Vargas.

Jeong, J. & Maddala, G.S. 1993. “A Perspective on Application of Bootstrap Methods in Econometrics”. In Maddala, G.S., Rao, C.R., Vinod, H.D. (Eds.) «Handbook of Statistics», vol. 11, Elsevier Science Publishers.

Johansen, S. 1991. “Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models". Econometrica 59: 1551-1580.

Johansen, S. 1995. “Likelihood-Based Inference in Cointegrated Vector Auto-Regressive Models”. New York: Oxford University Press.

Johansen, S. & Juselius, K. 1990. "Maximum Likelihood Estimation and Inference on Cointegration - with Applications to Demand for Money". Oxford Bulletin of Economics and Statistics 52: 169-210.

Koop, G., Pesaran, M.H. & Potter, S. M. 1996. “Impulse Response Analysis in Nonlinear Multivariate Models”. Journal of Econometrics 74: 119-147.

Lee, T-W. & Tse, Y. 1996. “Cointegration Tests with Conditional Heteroskedasticity”. Journal of Econometrics 73: 401-410.

Li, H. & Maddala, G.S. 1996. “Bootstrapping Time Series Models”. Econometric Review 15: 115-158.

Lütkepohl, H. & Reimers, H-E. 1992. “Impulse Response Analysis of Cointegrated Systems”. Journal of Economic Dynamics and Control 16: 53-78.

Marques, M.S. 1983. “Moeda e Inflação: A Questão da Causalidade". Revista Brasileira de Economia 37: 13-38.

Marques, M.S. & Werlang, S.R. 1989. “Moratória Interna, Dívida Pública e Juros Reais". Pesquisa e Planejament0 Econômico 19: 19-44.

McCallum, B.T. 1989. “Monetary Economics: Theory and Practice”. New York: Macmillian.

Novaes, A.D. 1991. “Um Teste da Hipótese de Inflação Inercial no Brasil". Pesquisa e Planejament0 Econômico 21: 377-396.

Pastore, A.C. 1991. "Déficit, Dívida Pública, Moeda e Inflação: uma Resenha", mimeo.

Pastore, A.C. 1995. “Déficit Publico, a Sustentabilidade do Crescimento das Dividas Interna e Externa, Senhoriagem e Inflação: Uma Analise do Regime Monetário Brasileiro”. R. de Econometria 14(2): 177-234.

Pastore, A.C. 1997. “Passividade Monetária e Inercia”. Revista Brasileira de Economia 51: 1-51.

Pindick, R.S. & Rubinfeld, D.L. 1981. “Econometric Models and Economic Forecasts". 2nd Edition. Singapore: McGraw Hill.

Rossana, R.J. & Seater, J.J. 1995. “Temporal Aggregation and Economic Time Series”. Journal of Business and Economic and Statistics 13(4): 441-451.

Sargent, T. 1984. "Stopping Moderate Inflations: The Methods of Poincaré and Thatcher", in Dornbusch, R. e Simonsen, M.H. (Eds.) "Inflation, Debt, and Indexation". Cambridge: The MIT Press.

Sargent, T. 1987. "Macroeconomic Theory". San Diego: Academic Press.

Sargent, T. & Wallace, N. 1973. "Rational Expectations and the Dynamics of Hyperinflation", in Lucas, R. e Sargent, T. (Eds.) "Rational Expectations and Econometric Practice". Minneapolis: The University of Minnesota Press, 1981.

Sargent, T. & Wallace, N. 1975. "Rational Expectations, the Optimal Monetary Instrument, and the Optimal Money Supply Rule", in Lucas, R. e Sargent, T. (eds.) op. cit..

Sargent, T. & Wallace, N. 1976. "Rational Expectations and the Theory of Economic Policy", in Lucas, R. e Sargent, T. (eds.) op. cit.

Sargent, T. & Wallace, N. 1982. “Some Unpleasant Monetarist Arithmetic”, in Sargent, T. “Rational Expectations and Inflation”. New York: Harper & Row, 1986.

Spanos, A. 1990. “Statistical Foundations of Econometric Modeling". Cambridge: Cambridge University Press.

Triches, D. 1992. “Demanda por Moeda no Brasil e a Causalidade entre as Variáveis Monetárias e a Taxa de Inflação: 1972/87”. BNDES.

Vinod, H.D. 1993. “Bootstrap Methods: Applications in Econometrics”. In Maddala, G.S., Rao, C.R., & Vinod, H.D., (Eds.) op. cit.

Welsh, J.H., 1991. “Rational Inflation and Real Internal Debt Bubbles in Argentina and Brasil?". L.A.M. of the Econometric Society, unpublished.

Working, H., 1960. “Note on the Correlation of Differences of Average in a Random Chain”. Econometrica 28: 916-918.

Appendix A: Simulating missing values in the public debt series

The series of public debt between 1964 and 1969 is only available in the last month of the year. In this sense, the alternative, in order to extend the analysis until 1966, is to perform an interpolation procedure to replace the missing data.

We chose to interpolate the series in logarithms (denoted as LD), since it made the estimations easier, and later to invert the simulations to obtain the nominal public debt. We used, as the simulation period, the months from 64.12 (which is the month with the first available data) to 69.12. We chose a longer period than needed to avoid an eventual learning-time inside our sample of interest, where the simulations would take a while until they achieved the long run path.

We first determined if the series had a unit root[50] between 1970.01 and 1973.12[51]. Tables A.1 and A.2 show the tests results and the respective diagnostics tests[52] to the series in levels.

TABLE A.1: UNIT ROOT TESTS REPORT PERIOD: 1970.01 TO 1973.12

SERIES	TEST	LAGS	$t\hat{\alpha}$
LD	ADF	3	-2.358
	PP	3	-2.874
ΔLD	ADF	2	-6.594**
	PP	3	-8.622**

Notes: ADF and PP tests specified with trend term. The symbol (**) represents rejection of the null of a unit root at the 1% significance level.

TABLE A.2_ADF TEST: RESIDUALS DIAGNOSTICS

DW=1.929	Q(12)=0.792	LM(1)=0.647	LM(4)=0.384	BJ=0.470
AIC=-5.754	SIC=-5.513	SER=0.053		

Since LD is integrated of order one, then for any k LD_t and LD_{t-k} will be cointegrated; see Granger (1991). Therefore, for matter of

[50] The ADF lag truncation was made based on the t-test of significance of the last lagged first difference combined with the information criteria. To the Phillips-Perron test, the choice was done with the Bartlett kernel.
[51] This period was chosen since it has resembling features to the former concerning the economic behavior.
[52] The notation follows the tradition used in the economic literature. The Durbin-Watson test as the information criteria (Akaike and Schwarz) and standard error

simulation we can take the static regression of the Engle-Granger two steps procedure, since it represents the long run relation between the contemporaneous and the series past values[53]. We decided to add a policy dummy, corresponding to 1971.01 to 1971.07, for controlling the residual normality and improving the simulations on the period 1970/1973[54]. The results are in Table A.3.

TABLE A.3: ENGLE-GRANGER COINTEGRATION TEST
DEPENDENT VARIABLE: LD

Variable	Coeff.	P-value
C	-3.141539	0.0004
T	0.014701	0.0005
D717	-0.068662	0.0054
LD(-1)	0.577799	0.0000

R^2=0.991	DW=1.721	SER=0.049	$t_{\hat{\alpha}}$ =-6.507**	lags=0
Q(12)=0.440	LM(1)=0.887	LM(4)=0.207	BJ=0.081	SIC=-5.854

Note: The symbol (**) represents rejection of the null of a unit root at the 1% significance level.

Given the second step unit root test statistic is well above the rejection value of 1% (-4.665), we do not have to worry about the introduction of the policy dummy in the long run equation.

We then drew from a normal distribution an error series with similar characteristics (maximum, minimum, mean, variance and kurtosis), in the range from 1964.12 to 1973.12, to the residuals of the first step procedure. We then performed the simulation for the period 1965.01/73.12 using the coefficient estimation from the equation above and with start value taken from 1964.12. The forecast evaluation to the sample 1970.01/73.12 is in the table below.

regression are reported with the test statistics. The Ljung-Box, the Breusch-Godfrey LM version, and the Bera-Jarque are reported with the respective p-values.

[53] A more natural way of performing the simulation would be to employ the ARIMA process implied by the series unit root test. However, the simulations with autoregressive processes presented periods of learning with striking volatility that extended until 1967. For this reason, we decided to abandon this alternative. We could also have interpolated the series by running an OLS regression of each month against December and then simulating the respective missing value with the estimated equation. Nevertheless, this procedure can only work with flow variables. With stock variables this method produced at the beginning of each simulated year a cycle not supported by the behavior of the remaining years.

[54] We tried several statistical specifications and this was that led to the best goodness of fit.

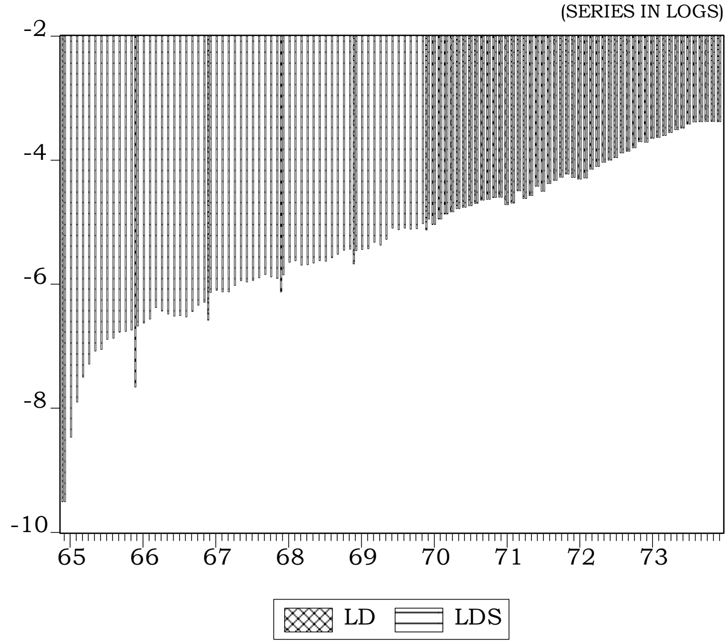
We notice that the forecast looks good, with no systematic bias and forecast error coming from random factors. In figure A.1, we show the observed and the simulated series during 1964.12/1973.12. The simulations computed for the period 1965.12/1969.12 was then matched with the original public debt series.

TABLE A.4: GOODNESS OF FIT

Root Mean Squared Error	0.006
Root Mean Squared Percentage Error	1.471
Mean Absolute Error	0.051
Mean Absolute Percentage Error	1.222
Theil Inequality Coefficient	0.007
Bias Proportion	0.028
Variance Proportion	0.030
Covariance Proportion	0.967

FIGURE A.1: PUBLIC DEBT OBSERVED AND SIMULATED

(SERIES IN LOGS)



Appendix B: Disaggregating the output index series

Since the pioneer paper of Working (1960), it is a well-known fact that temporal aggregation has statistical implications on the time series properties. As pointed out in the work of Rossana and Seater (1995) the temporal aggregation causes substantial losses of information about the underlying data processes. They argue that non-aggregated data is governed by complex time series processes with much low frequency cyclical variation, whereas aggregated data is 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 is aggregated. Moreover, the aggregated data shows more long-run persistence than the underlying disaggregated data.

For all the above reasons, we decided to interpolate the Brazilian GDP index quarterly series (YQ). This series is constructed by matching the IBGE index product with the index computed by Rossi (1988); see Cerqueira (1993).

Our empirical strategy was to determine the series AR(1) representation, between 1966.1 and 1991.4[55], using the equation provided by the unit root level test reported below.

TABLE B.1: UNIT ROOT TESTS REPORT PERIOD: 1966.1 TO 1991.4

SERIES	TEST	LAGS	$t\,\hat{\alpha}$
YQ	ADF	8	-1.683
	PP	3	-1.147
ΔYQ	ADF	2	-13.464**
	PP	3	-12.982**

Notes: ADF and PP tests specified without trend term. The symbol (**) represents rejection of the null of a unit root at the 1% significance level.

TABLE B.2_ADF TEST: RESIDUALS DIAGNOSTICS

DW=1.989	Q(24)=0.681	Q(36)=0.880	LM(1)=0.993	LM(3)=0.920	LM(6)=0.978
LM(9)=0.180	LM(12)=0.137	BJ=0.144	AIC=2.196	SIC=2.492	SER=2.840

Solving the difference equation, given by the ADF test equation, in terms of the variable first difference we got an AR(1) representation with serial correlation coefficient ≈ 0.9824[56]. We then distributed the quarterly series on a monthly basis using the Kalman filtering procedure, given the series followed the above AR(1) process; see Hamilton (1994) and Harvey (1990).

[55] We chose a longer period than we needed because the Kalman filtering procedure employs backwards and forwards estimations, which demands some space for accommodations at each extreme point.

[56] This coefficient is computed with the expression $\hat{\rho} = 1 + 1 \big/ (1 - \sum_{i=1}^{8} \Delta y_{t-i})$

In figure B.1 the interpolated series[57] (YQI) plot is overlapped on the original data. To illustrate the disaggregation impact we report in the next table the monthly GDP series ADF test. The number of first difference lags necessary to correct the serial correlation, and the series stationarity are striking. This result is coherent with our previous findings using the bootstrap technique; see Cerqueira (1998).

FIGURE F.1: GDP INDEXES - INTERPOLATED AND OBSERVED DATA

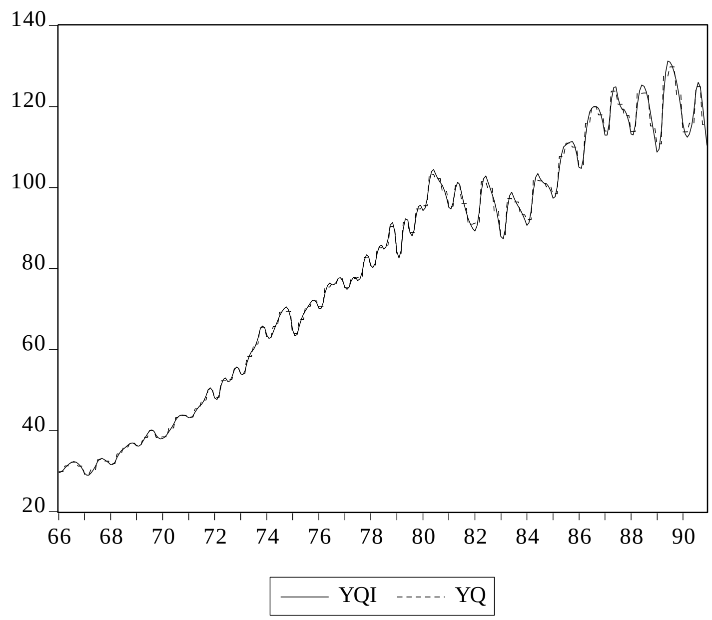


TABLE B.3: UNIT ROOT TESTS REPORT PERIOD: 1966.01 TO 1990.03

SERIES	TEST	LAGS	$t\hat{\alpha}$		
YQI	ADF	17	-3.013*		
DW=1.970	Q(24)=0.990	Q(36)=0.821	LM(1)=0.425	LM(3)=0.441	LM(6)=0.312
LM(9)=0.316	LM(12)=0.466	BJ=0.000	AIC=-11.120	SIC=-10.790	SER=0.004

Notes: ADF test specified without trend term. The symbol (*) represents rejection of the null of a unit root at the 5% significance level.

The final step is to obtain the monthly real GDP series, in values of 1980, and the nominal GDP. We computed these series in the following manner:

$$\text{i) Real GDP}_{j.i} = (GDP_{80}/12)*(YQI_{j.i}/100)$$

$$\text{ii) Nominal GDP}_{j.i} = (\text{Real GDP}_{j.i})*(P_{j.i})$$

[57] Because we have an indexed series, the output series was multiplied by the ratio between periodicities, e.g. 3 for quarterly to monthly, in order to do the mean values of

where, $GDP_{j.i}$=GDP at prices of 1980, month j, year i; GDP_{80}=GDP of 1980; $YQI_{j.i}$=monthly GDP index, month j, year i, 1980=100; $P_{j.i}$=general price index, 1980=100.

Appendix C: Estimating the seigniorage steady-state level

In this appendix, we show the methodology employed to estimate the seigniorage-GDP ratio from 1974.01 to 1988.06. We define seigniorage as the first difference of the monetary base (high-powered money) over the price index. Since we are mainly concerned with studying monthly series, the present estimation is made using this data frequency; see appendix B. The nominal GDP series is taken from the interpolation made possible on the latest appendix.

We assume that since the beginning of the inflationary acceleration, which started in 1974, even for low levels of inflation (around 1.0% per month), the inflation acceleration had no impact on the government revenue with monetary issuing as GDP proportion. Thus, from this level of inflation the primary monetary expansion could be taken as constant and independent from the rate of inflation, so we may presume it might be described as a white noise process; see figure C.1. This hypothesis is based on the increasing cost of holding money and on the process of money substitution by other financial assets; see Marques and Werlang (1989).

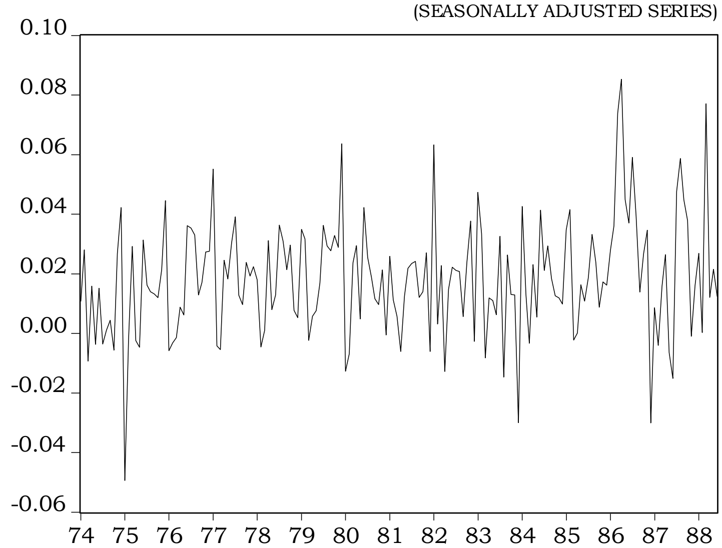
The choice of period was based primarily on the time series memory, that is, we tried to identify the longest period where the series behavior was close to an independent sequence of random variables; see correlogram series and periodogram below. We also note that 1974 marks the beginning of the inflation acceleration and the introduction of new financial assets (both increased the cost of holding money and started up a process of monetary demand contraction). The truncation

1980 equals to 100.

month is associated with the increasing tendency showed by the seigniorage collection.

Since the seigniorage-GDP series presents a deterministic seasonal pattern, we first chose to perform a seasonal adjustment, by running the observed data against 11 seasonal centered dummies[58]. This procedure brought into the series SY plotted in figure C.1.

FIGURE C.1: SEIGNIORAGE-GDP RATIO
(SEASONALLY ADJUSTED SERIES)



Our statistical approach was then to proceed using independence tests on the adjusted series. We first performed the unit root tests reported in Tables C.1 and C.2. The tests indicate the seigniorage-GDP ratio is stationary.

TABLE C.1: UNIT ROOT TESTS REPORT PERIOD: 1974.01 TO 1988.06

SERIES	TEST	LAGS	$t\hat{\alpha}$
SY	ADF	0	-11.817**
	PP	4	-11.895**

Notes: ADF and PP tests specified without trend term. The symbol (**) represents rejection of the null of a unit root at the 1% significance level.

TABLE C.2_ADF TEST: RESIDUALS DIAGNOSTICS

DW=2.004	Q(24)=0.840	Q(36)=0.785	LM(1)=0.713	LM(3)=0.576	LM(6)=0.683
LM(9)=0.708	LM(12)=0.323	BJ=0.001	AIC=-5.034	SIC=-4.998	SER=0.019

[58] A centered dummy is a variable, which assumes the value (1-1/12) in a specific month and (-1/12) otherwise.

In figure C.2, we show the series correlogram. It is immediate to see that the autocorrelation dies off very quickly with increasing lags, which shows the series is asymptotically uncorrelated. The Ljung-Box test statistics accepts the null of no autocorrelation up to order 12, 24 and 48 with p-values 0.226, 0.654 and 0.659, respectively. This result is corroborated by the Kolmogorov-Smirnov statistics, which accepts the null of a non-serially correlated series with p-value well superior to 10%[59]; see figure C.3. The spectral shape also suggests that the SY series is white noise; see figure C.4.

The seigniorage-GDP series during the period is described by the following statistics: mean=0.606%, std. dev.=0.649%, skew.=0.270, kurt.=4.361. However, it was not normally distributed considering the Bera-Jarque statistics is rejected with a p-value near of 0.004%.

If the above results are corrected, we will expect the regression of SY against a constant and the rate of inflation (pi) must give the following outcomes: (i) constant significant at the level of the sample mean (0.606%); (ii) coefficient of pi non-significant; (iii) R^2 near zero; (iv) residuals approximately IID. These results are testified by the report below.

TABLE C.3: OLS REGRESSION[60] SY AGAINST PI[61]
(WHITE HETEROSKEDASTICITY S.E. CONSIDERED)

Variable	Coeff.	P-value
C	0.006739	0.0000
PI	-0.010295	0.3340

R^2=0.007	F=0.285	SER=0.006	DW=1.793	Q(24)=0.443	Q(36)=0.382
LM(1)=0.174	LM(3)=0.214	LM(6)=0.327	LM(9)=0.443	LM(12)=0.167	BJ=0.001
WHITE=0.129	ARCH(1)=0.013	ARCH(4)=0.124	Q^2(18)=0.823		

[59] The test statistics is equal to 0.0898 and the critical value of a two-sided test of size 10% is 0.125; see Harvey (1990).
[60] We employed the White heteroskedasticity correction due to the presence of some ARCH processes in the residuals. The notation is the same as in previous appendixes. We added to the report the White, ARCH, and Q^2 tests of heteroskedasticity.
[61] Since the residuals are not normal, we computed the likelihood and F Monte Carlo distributions. The p-values Monte Carlo associated with the constant tem are equal to 0.001, while for the inflation coefficient the p-values are both equal to 0.292.

With p-value equals to 0.449, we can not reject the hypothesis that the equation constant term is approximately equals to the series sample mean. We can not also reject the assumption of a non-significant inflation coefficient and that the R^2 and the residuals are as expected.

All the above results corroborate the hypothesis that during 1974.01 to 1988.06 the seigniorage-GDP ratio followed a white noise process ~ (0.606%,0.649%). In addition, it could be taken as independent from the rate of inflation that had no effect in changing this government revenue. This conclusion gives empirical support to the assumption that the monetary policy was not passive during 1974/1988, in the sense the rate of inflation had no impact in increasing or reducing the seigniorage collection.

There was a public deficit permanently financed with monetary expansion which assured the debt sustainability (see Cerqueira (1998)), without made the money supply endogenous. This means there was a steady state level of public deficit financed with issuing money[62], which is coherent with the non-rational Cagan's adaptive model. To complete the prove of the hypothesis of a money supply exogeneity, it is "necessary" to show the money creation caused inflation, in the Granger's sense, while inflation did not cause money creation (Sargent and Wallace, 1973). A task performed in section 5 of this paper.

[62] We notice that along the studied period the inflation tax-GDP ratio mean value was around 0.668%. Then with a p-value of 0.275 we can not reject the mean equality hypothesis between this series and the seigniorage-GDP ratio.

FIGURE C.2

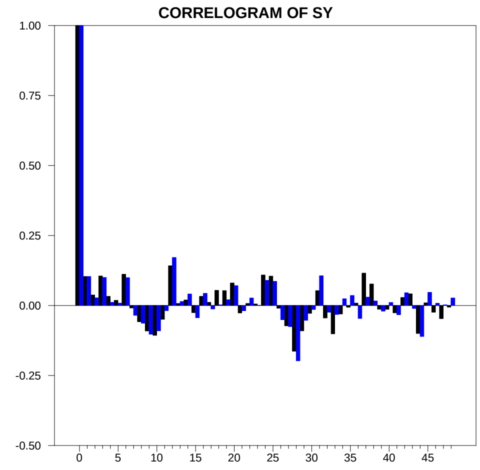


FIGURE C.3

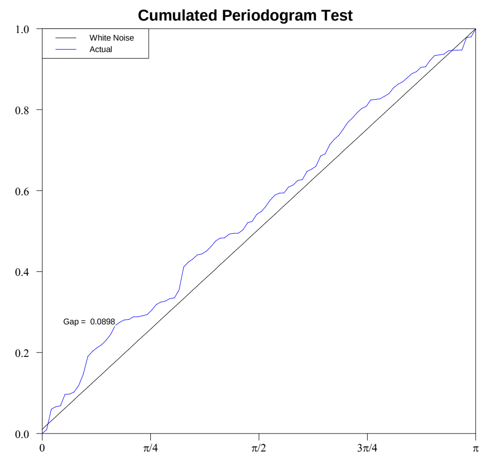
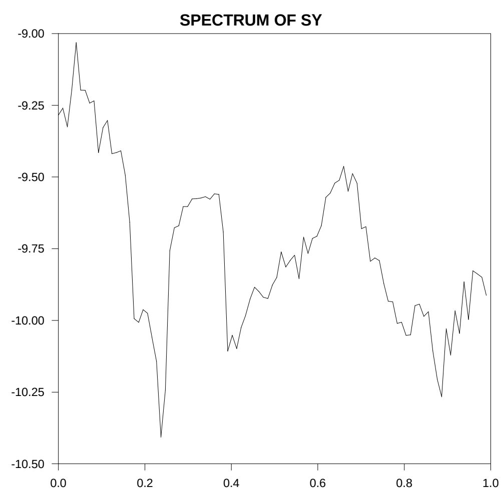


FIGURE C.4



Appendix D: Causality tests between the debt-GDP ratio and the real interest rate

The debt-GDP series is denoted by LDY (series in logarithms) and the real interest rate by R. The remaining notation is the same as used in the previous appendixes. With the exception of the R^2, SER, AIC, SIC, DW and KS statistics, the others are reported by their respective p-values. Unit root tests performed as earlier.

TABLE D.1: UNIT ROOT TESTS PERIOD: 1966.01 TO 1985.12

SERIES	ADF		Phillips-Perron	
	LAGS	$t\hat{\alpha}$	LAGS	$t\hat{\alpha}$
Real Interest Rate(R)	2	-4.370**	4	-8.260**
Debt-GDP Ratio(LDY)	3	-2.326	4	-2.047
ΔLDY	2	-7.067**	4	-13.980**

Notes: (1) Real interest of rate tests specified without trend term. Debt-GDP ratio level tests specified with trend term and first difference without trend. The symbol (**) represents rejection of the null of a unit root at the 1% significance level.

TABLE D.2: VAR LAG TRUNCATION

VAR ORDER	LR RATIO* (P-VALUES)	AKAIKE CRITERION	SCHWARZ CRITERION	HANNAN-QUINN CRITERION
6	---	-8.938	-8.553	-8.783
5	0.104	-8.754	-8.428	-8.844
4	0.025	-8.979	-8.714	-8.872
3	0.002	-8.966	-8.761	-8.883

*LR statistics has a $\chi^2(4)$ distribution.

Table D.3 reports the diagnostics tests to the VAR model without correction for the residual non-normality. The tests point out the presence of strong heteroskedasticity in the system residuals. The normality violation is due mainly to the existence of large outliers concentrated in specific periods. One procedure for correcting this problem is to model the error terms with a GARCH process. The maximum likelihood estimates (i.e., under the assumption that the errors are conditionally normally distributed and without pre-specifying any upper bound to the number of iterations) are reported in Table D.4.

TABLE D.3: VAR MODEL (4 LAGS) RESIDUALS DIAGNOSTICS
EQUATION 1 (ΔLDY)

R^2=0.115	F=0.001	SER=0.052	AIC=-3.026	SIC=-2.894	DW=2.019
Q(24)=0.577	Q(36)=0.899	LM(1)=0.254	LM(3)=0.235	LM(6)=0.441	LM(9)=0.664
LM(12)=0.667	WHITE=0.015	ARCH(1)=0.262	ARCH(4)=0.001	Q^2(18)=0.003	BJ=0.066

EQUATION 2 (R)

R^2=0.394	F=0.000	SER=0.012	AIC=-5.916	SIC=-5.783	DW=2.012
Q(24)=0.991	Q(36)=0.338	LM(1)=0.186	LM(3)=0.151	LM(6)=0.378	LM(9)=0.291
LM(12)=0.484	WHITE=0.000	ARCH(1)=0.144	ARCH(4)=0.000	Q^2(18)=0.000	BJ=0.000

TABLE D.4: VGARCH MODEL ESTIMATES*

Variable	EQUATION 1 (ΔLDY)		EQUATION 2 (R)	
	Coeff.	P-value	Coeff.	P-value
C	0.005503	0.0974	0.000414	0.5396
R(-1)	0.689496	0.0191	0.555874	0.0000
R(-2)	0.227801	0.4190	0.018095	0.8321
R(-3)	0.534059	0.1302	0.062956	0.3828
R(-4)	-0.795382	0.0026	0.041472	0.5547
ΔLDY(-1)	0.036518	0.6212	0.000435	0.9688
ΔLDY(-2)	0.011702	0.8802	0.015996	0.1354
ΔLDY(-3)	0.104474	0.1626	0.001440	0.9026
ΔLDY(-4)	-0.016253	0.8137	0.004926	0.6907
VARIANCE EQUATION				
c	0.002584	0.0000	0.000292	0.6119
q_{t-1}-c	0.634255	0.0030	0.975887	0.0000
$u^2_{t-1}-\sigma^2_{t-1}$	0.273876	0.1123	0.340961	0.0000
$u^2_{t-1}-q_{t-1}$	-0.224628	0.2793	-0.209700	0.0024
$\sigma^2_{t-1}-q_{t-1}$	0.034716	0.9635	0.335714	0.4283

Note: u^2_{t-1} represents the ARCH term, σ^2_{t-1} the GARCH term, and q_{t-1} the long run component.

EQUATION 1: RESIDUALS DIAGNOSTICS

R^2=0.113	F=0.012	SER=0.053	AIC=-3.052	SIC=-2.846	DW=2.014
Q(24)=0.360	Q(36)=0.722	ARCH(1)=0.692	ARCH(4)=0.526	Q^2(18)=0.397	BJ=0.097

EQUATION 2: RESIDUALS DIAGNOSTICS

R^2=0.370	F=0.000	SER=0.013	AIC=-6.248	SIC=-6.041	DW=2.158
Q(24)=0.927	Q(36)=0.516	ARCH(1)=0.960	ARCH(4)=0.994	Q^2(18)=0.881	BJ=0.281

Tables D.5 and D.6 show the diagnostics tests for the Geweke, Meese, and Dent (GMD) causality test. Test residuals of Table D.5 point out the presence of heteroskedasticity and the violation of the normality assumption. These problems vanished with the filters ARCH(3) and GARCH(1,1) applied respectively for each residual equation; see Table D.6. Note that the information criteria chose the GARCH specifications.

TABLE D.5: GMD TEST RESIDUALS DIAGNOSTICS WITHOUT CORRECTION FOR LACK OF NORMALITY
DEPENDENT VARIABLE: ΔLDY*

R^2=0.150	F=0.001	SER=0.052	AIC=-3.024	SIC=-2.818	DW=2.030
Q(24)=0.361	Q(36)=0.742	LM(1)=0.118	LM(3)=0.127	LM(6)=0.349	LM(9)=0.508
LM(12)=0.286	WHITE=0.065	ARCH(1)=0.212	ARCH(4)=0.001	Q^2(18)=0.001	BJ=0.006

* REGRESSION SPECIFIED WITH 4 LAGS TO ΔLDY AND R AND 4 LEADS TO R.
DEPENDENT VARIABLE: R*

R^2=0.475	F=0.000	SER=0.012	AIC=-5.989	SIC=-5.738	DW=1.993
Q(24)=0.942	Q(36)=0.352	LM(1)=0.912	LM(3)=0.770	LM(6)=0.295	LM(9)=0.290
LM(12)=0.318	WHITE=0.001	ARCH(1)=0.180	ARCH(4)=0.000	Q^2(18)=0.000	BJ=0.000

* REGRESSION SPECIFIED WITH 5 LAGS TO ΔLDY AND R AND 5 LEADS TO ΔLDY.

TABLE D.6: GMD TEST RESIDUALS DIAGNOSTICS WITH CORRECTION FOR LACK OF NORMALITY
DEPENDENT VARIABLE: ΔLDY

R^2=0.144	F=0.007	SER=0.053	AIC=-3.093	SIC=-2.828	DW=2.015
Q(24)=0.191	Q(36)=0.500	ARCH(1)=0.510	ARCH(4)=0.688	Q^2(18)=0.479	BJ=0.301

*RESIDUALS SPECIFIED WITH AN ARCH(3) PROCESS.
DEPENDENT VARIABLE: R

R^2=0.423	F=0.000	SER=0.012	AIC=-6.254	SIC=-5.958	DW=2.286
Q(24)=0.952	Q(36)=0.389	ARCH(1)=0.359	ARCH(4)=0.724	Q^2(18)=0.772	BJ=0.414

*RESIDUALS SPECIFIED WITH A GARCH(1,1) PROCESS.

To perform Sims' causality test we first pre-whitened the series with an AR(4) process that included a treatment for the seasonality presented in the year's fourth month. The pre-filtering process originated the series LDYF and RF. The diagnostics tests reported in Table D.7 show that the employed filters conducted the residuals of the Sims' procedure to be approximately innovation processes.

TABLE D.8: SIMS' TEST RESIDUALS DIAGNOSTICS WITHOUT CORRECTION FOR LACK OF NORMALITY (REGRESSIONS SPECIFIED WITH 6 LAGS AND LEADS).
DEPENDENT VARIABLE: LDYF*

R^2=0.124	F=0.008	SER=0.050	AIC=-3.085	SIC=-2.872	DW=2.191
Q(24)=0.561	Q(36)=0.882	LM(1)=0.153	LM(3)=0.192	LM(6)=0.428	LM(9)=0.455
LM(12)=0.479	WHITE=0.651	ARCH(1)=0.540	ARCH(4)=0.002	Q^2(18)=0.026	BJ=0.030

*KS=0.0958 (5% CRITICAL VALUE=0.0981)
DEPENDENT VARIABLE: RF*

R^2=0.125	F=0.007	SER=0.011	AIC=-6.147	SIC=-5.934	DW=2.090
Q(24)=0.809	Q(36)=0.786	LM(1)=0.465	LM(3)=0.796	LM(6)=0.926	LM(9)=0.990
LM(12)=0.914	WHITE=0.753	ARCH(1)=0.040	ARCH(4)=0.167	Q^2(18)=0.000	BJ=0.238

*KS=0.048 (10% CRITICAL VALUE=0.0880)

Appendix E: Demand for bonds estimates

Table E.1 shows the diagnostics tests for the bonds demand equation, specified without correction for normality. In Table E.2, the same model is re-estimated using an ARCH(3) structure in the residuals.

TABLE E.1: AD(4,4) MODEL RESIDUALS DIAGNOSTICS DEP. VAR.: ΔLDY

R^2=0.147	F=0.000	SER=0.052	AIC=-3.046	SIC=-2.884	DW=2.027
Q(24)=0.434	Q(36)=0.797	LM(1)=0.141	LM(3)=0.224	LM(6)=0.476	LM(9)=0.664
LM(12)=0.361	WHITE=0.007	ARCH(1)=0.231	ARCH(4)=0.001	Q^2(18)=0.004	BJ=0.003

TABLE E.2: AD-ARCH MODEL ESTIMATES
DEPENDENT VARIABLE: ΔLDY

VARIABLES	Coeff.	P-value
C	0.008738	0.2097
T	-1.59E-05	0.7496
R	0.822129	0.0026
R(-1)	0.155766	0.6238
R(-2)	0.356950	0.1713
R(-3)	0.290088	0.3675
R(-4)	-0.718307	0.0046
ΔLDY(-1)	0.023257	0.6465
ΔLDY(-2)	0.006518	0.9266
ΔLDY(-3)	0.109573	0.1427
ΔLDY(-4)	-0.040650	0.5152
VARIANCE EQUATION		
c	0.001584	0.0000
u_{t-1}^2	-0.072964	0.0001
u_{t-2}^2	0.223443	0.0177
u_{t-3}^2	0.215645	0.0081

R^2=0.144	F=0.002	SER=0.052	AIC=-3.113	SIC=-2.892	DW=1.987
Q(24)=0.242	Q(36)=0.574	ARCH(1)=0.489	ARCH(4)=0.797	Q^2(18)=0.483	BJ=0.131

Appendix F: Causality tests between monetary expansion and inflation rate

Table F.1 shows the unit root tests. The tests are reported as in the previous appendixes. Table F.2 shows the computed statistics for selecting the unrestricted VAR lag truncation. Table F.3 reports the Johansen's procedure in the conventional way. In the reports of the diagnostics tests (tables F.4 and F.5), we also present the Bowman & Shenton (BS) normality test described in Hansen and Juselius (1995). Figure F.1 shows the roots of companion matrix associated with the unrestricted estimated VAR.

TABLE F.1: UNIT ROOT TESTS PERIOD: 1966.01 TO 1985.12

SERIES	ADF		Phillips-Perron	
	LAGS	$t\,\hat{\alpha}$	LAGS	$t\,\hat{\alpha}$
Inflation Rate (PI)	2	-2.382	4	-5.761**
ΔPI	1	-9.549**	4	-26.651**
Money Growth (MI)	11	+0.507	4	-15.943**
ΔMI	10	-8.849**	4	-44.272**

Notes: (1) Inflation rate level tests specified with trend term and first difference tests without trend. Money creation tests specified with trend term. The symbol (**) represents rejection of the null of a unit root at the 1% significance level.

TABLE F.2: VAR LAG TRUNCATION

VAR ORDER	CONSTANT	LINEAR TREND	SCHWARZ CRITERION	HANNAN-QUINN CRITERON
13	unrestricted	coint. space	-14.936	-15.733
14	unrestricted	coint. space	-14.871	-15.702
15	unrestricted	coint. space	-14.796	-15.662
16	unrestricted	coint. space	-14.723	-15.623
17	unrestricted	coint. space	-14.636	-15.571

TABLE F.3: JOHANSEN'S COINTEGRATION TEST

TEST STATISTICS (CRITICAL VALUES AT THE 15% LEVEL)				**COINTEGRATING VECTOR**
L-max		Trace		(MONEY, INFLATION, TREND, CONSTANT)
r=0	r≤1	r=0	r≤1	
13.535*	7.942	21.477	7.942	(1.0000, -1.3400, -0.0002, 0.0432)
(11.977)	(9.427)	(21.404)	(9.427)	

COINTEGRATION RESTRICTION TEST

RESTRICTION: (1,-1,#,#) $\chi^2(1) = 0.33$; P-VALUE = 0.566

Notes: The symbol (*) indicates rejection of the null at the 5% significance level. The symbol # means the parameter is unrestricted.

FIGURE F.1

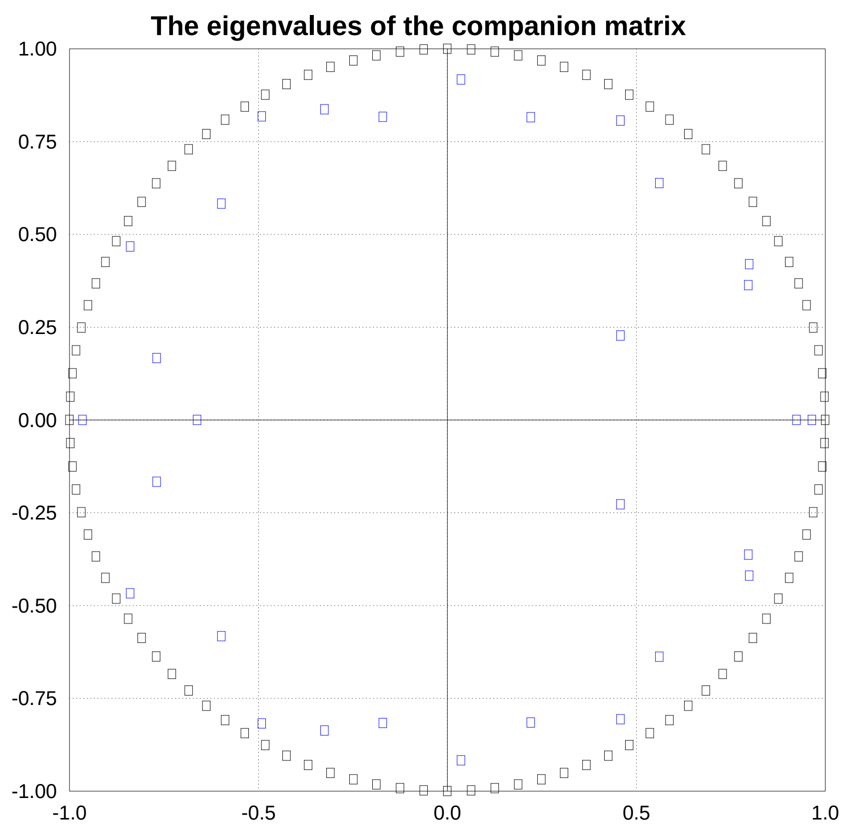


TABLE F.4: UNRESTRICTED VAR DIAGNOSTICS TESTS

MULTIVARIATE TESTS				
	Q(60)=0.078	LM(1)=0.339	LM(4)=0.795	BS=0.001
UNIVARIATE TESTS				
EQ.1(ΔPI):	Q(60)=0.727	BS=0.338	BJ=0.071	ARCH(15)=0.262
EQ.2:(ΔMI):	Q(60)=0.935	BS=0.010	BJ=0.002	ARCH(15)=0.218

TABLE F.5: RESTRICTED VAR (VECM) DIAGNOSTICS TESTS*

MULTIVARIATE TESTS				
	Q(60)=0.109	LM(1)=0.290	LM(4)=0.802	BS=0.014
UNIVARIATE TESTS				
EQ.1(ΔPI):	Q(60)=0.682	BS=0.325	BJ=0.107	ARCH(15)=0.247
EQ.2:(ΔMI):	Q(60)=0.910	BS=0.101	BJ=0.026	ARCH(15)=0.364

*Tests computed using (1,-1,#,#) as cointegrating vector.

In the next table, 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 F.3 and these two adjustment coefficients. Below the table we report the corresponding Monte Carlo p-values to the test which used the cointegrating vector (1,-1,#,#).

TABLE F.6: 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	0.34 (0.23)	0.844 (0.632)
PI is weakly exogenous for the parameter of interest of the MI conditional model	10.91 (5.12)	0.004** (0.024)

Note: The symbol (**) represents rejection of the null at the 1% significance level. If the VAR had 13 lags, the p-values, respectively for each hypothesis would be 0.477 and 0.004. MI "IS W.E. TO" PI $\begin{Bmatrix} F = 0.935 \\ LR = 0.935 \end{Bmatrix}$; PI "IS W.E. TO MI $\begin{Bmatrix} F = 0.001 \\ LR = 0.001 \end{Bmatrix}$.

Table F.7 shows the diagnostics tests of the Geweke, Meese, and Dent causality tests. For the equation which dependent variable is the inflation first difference, we began with 12 lags. Since we could get uncorrelated residuals with a smaller lag length, we reduced the model until an AR(5) using the Schwarz and Hannan-Quinn (HQ) criteria. The violation of homoskedasticity and normality assumptions was solved using a Component-GARCH specification in the residuals. The diagnostics tests are reported in Table F.8. In the equation which dependent variable is the money growth first difference, we could choose between a model with twelve lags and leads, and another with 5 lags and leads with the residuals specified with an AR(6) process. The Schwarz and Hannan-Quinn criteria chose the more parsimonious model. We got NIID residuals by specifying them with an ARCH(2) process; see Table F.8. We remark that for both equations all information criteria chose the models with Gaussian residuals. Furthermore, the conclusion about the causality direction is not affected by the model selection.

TABLE F.7: GMD TEST RESIDUALS DIAGNOSTICS WITHOUT CORRECTION FOR LACK OF NORMALITY

DEPENDENT VARIABLE: ΔPI*

R^2=0.413	F=0.000	SER=0.012	AIC=-5.882	SIC=-5.621	DW=1.972
Q(24)=0.492	Q(36)=0.202	LM(1)=0.236	LM(3)=0.267	LM(6)=0.445	LM(9)=0.089
LM(12)=0.136	WHITE=0.000	ARCH(1)=0.010	ARCH(4)=0.000	Q^2(18)=0.000	BJ=0.000

* REGRESSION SPECIFIED WITH 5 LAGS TO ΔPI AND ΔMI AND 5 LEADS TO ΔMI. HQ=-5.777.

DEPENDENT VARIABLE: ΔMI*

R^2=0.859	F=0.000	SER=0.029	AIC=-4.113	SIC=-3.692	DW=1.929
Q(24)=0.659	Q(36)=0.570	LM(1)=0.486	LM(3)=0.121	LM(6)=0.083	LM(9)=0.047
LM(12)=0.037	WHITE=0.132	ARCH(1)=0.000	ARCH(4)=0.000	Q^2(18)=0.000	BJ=0.000

* REGRESSION SPECIFIED WITH 5 LAGS TO ΔPI AND ΔMI AND 5 LEADS TO ΔPI. HQ=-3.943.

TABLE F.8: GMD TEST RESIDUALS DIAGNOSTICS WITH CORRECTION FOR LACK OF NORMALITY

DEPENDENT VARIABLE: ΔPI*

R^2=0.307	F=0.000	SER=0.014	AIC=-6.263	SIC=-5.944	DW=2.076
Q(24)=0.863	Q(36)=0.844	ARCH(1)=0.924	ARCH(4)=0.853	Q^2(18)=0.751	BJ=0.349

* RESIDUALS SPECIFIED WITH A GARCH(1,1) PROCESS. HQ=-6.134.

DEPENDENT VARIABLE: ΔMI*

R^2=0.844	F=0.000	SER=0.031	AIC=-4.315	SIC=-3.851	DW=1.997
Q(24)=0.500	Q(36)=0.245	ARCH(1)=0.475	ARCH(4)=0.742	Q^2(18)=0.146	BJ=0.355

* RESIDUALS SPECIFIED WITH AN ARCH(2) PROCESS. HQ=-4.128.

To perform the Sims' causality test we first filtered the series in order to obtain serially uncorrelated residuals from the test equations. The procedure originated the filtered series PIF (inflation) and MIF (money growth). Since as a must we have to use for both series the same filter, we chose to employ the autoregressive process obtained from the money growth ADF test. Then we estimated two AR(12) processes including 11 seasonal dummies. The filters are not reported, as also the test specifications, but may be obtained from the author on request. The same remark holds for the GMD tests.

The first requirement for the model selection was the absence of serial correlation in the test residuals. For both test equations the lag length choice was made based either on the Schwarz and Hannan-Quinn criteria or on the sequential test to the significance of the largest lag (we began with a lag length of 8). Table F.9 reports the respective residual diagnostics to each estimated equation. It can be observed that the estimated residuals filled the requirement of being independent innovations. However, they do not have neither homoskedasticity nor normality property. Since the equations have small F-statistics significance levels, we chose to perform Monte Carlo tests for simulating the causality test empirical distributions.

TABLE F.9: SIMS' TEST RESIDUALS DIAGNOSTICS WITHOUT CORRECTION FOR LACK OF NORMALITY (REGRESSIONS SPECIFIED WITH 7 LAGS AND LEADS).
DEPENDENT VARIABLE: PIF*

R^2=0.116	F=0.031	SER=0.011	AIC=-6.158	SIC=-5.916	DW=1.950
Q(24)=0.378	Q(36)=0.104	LM(1)=0.790	LM(3)=0.617	LM(6)=0.784	LM(9)=0.804
LM(12)=0.692	WHITE=0.609	ARCH(1)=0.030	ARCH(4)=0.000	Q^2(18)=0.000	BJ=0.000

*HQ=-6.060; KS=0.0510[CV(10%)=0.0880].
DEPENDENT VARIABLE: MIF*

R^2=0.097	F=0.048	SER=0.027	AIC=-4.327	SIC=-4.117	DW=1.942
Q(24)=0.555	Q(36)=0.600	LM(1)=0.752	LM(3)=0.435	LM(6)=0.426	LM(9)=0.475
LM(12)=0.542	WHITE=0.000	ARCH(1)=0.000	ARCH(4)=0.000	Q^2(18)=0.000	BJ=0.000

*HQ=-4.243; KS=0.0674 [CV(10%) = 0.0880].