Exogeneity of Money Supply in Brazil from 1966 to 1985

Luiz Fernando Cerqueira*

Abstract – Using monthly data spanning from 1966 to 1985, we examine the extent of 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. It was shown that the real rate of interest did cause, in the Granger sense, the bond stock and that the demand for bonds was very sensitive to interest rate variations. This implies that the monetary authority was able to perform indirect monetary control through the open market transactions. The results show that seigniorage collection was a white noise and econometrically independent from the inflation rate. Money creation and the inflation rate were cointegrated. We found that money growth was weakly exogenous for the parameter of interest in the conditional model of inflation, but the reverse is not true for the inflation. Moreover, Granger’s causal relation between them was unidirectional from money to inflation. Therefore, money growth is strongly exogenous concerning the inflation rate. 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 inflation rate and that the monetary authority had enough independence to execute an active monetary policy.

Keywords – Time series models, econometric modeling, bootstrap, inflation, money supply, monetary policy.

JEL – C32, C51, C52, E31, E51, E52

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1. Introduction

From 1966 to 1985, the inflation stabilization policy was centered on aggregate demand management and wage and price controls. The fact that during these years the economic policy was conducted in a seemingly “orthodox way” may induce us to ask about the monetary policy behavior. Are there reasons to support the claim that the money supply was exogenously determined concerning the inflation rate, that is, did the monetary authority follow a rule of money creation that did not accommodate the movements in the inflation rate?

The purpose of the present paper is to confirm that the monetary supply was exogenously determined from 1966 to 1985. That is, we claim that the monetary rule was executed in an independent way and the rule that guided the monetary execution exogenously respected the considered ‘model in question’ as well as the inflation rate. We intend to support 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 fraction of the revenues by reducing the base multiplier. Moreover, the seigniorage-GDP ratio was independent from the inflation rate 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 supporting this belief is the hypothesis of rational expectations. If the demand function for real balances follows Cagan’s form, the solution for the current inflation is a function of the rate of expected money creation, excluding the possibility of rational explosive bubbles. In this case, the money supply is endogenous.

An alternative argument is 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 past rates of inflation. This is a model in which the adaptive mechanism is rational.
Sargent and Wallace’s 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 past rate of inflation. So both money creation and inflation are best forecast by extrapolating current and lagged rates of inflation. Lagged rates of money creation add nothing to predictions formed in this way. In this model, past values of inflation influence money creation but the converse is not true; thus, money 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 Bruno and Fischer’s (1990) 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 rate of inflation 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 rate of inflation, 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 rates of inflation gave rise to an efficient indexation system that protected agents from the effects of inflation. Even if the indexation rules did not fairly contemplate the agents, one cannot deny that such rules prevented the high inflation to degenerate into public panic and open hyperinflation processes. Furthermore, the indexation rules were developed little by little along the seventies.
and eighties simultaneously with an increasing rate of inflation. Then the rigidity of the price system was introduced gradually, which augmented 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 argue that the monetary policy followed a rule 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 A.

The assumption of endogenous money growth cannot be supported by empirical evidence, when tested. Surprisingly, some authors found a unidirectional relation between inflation and money increase. 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 (since the significant correlation at one lag may be diluted by insignificant correlation at other lags) in the detection of specific important departures from the assumed model. It is therefore unwise to rely exclusively on this test when checking for model adequacy. However, it can be valuable when used with other tests. The Breusch-Godfrey Lagrange multiplier test is a common complement to the Q-test (Granger and Newbold, 1986).

By carrying out both tests and using monthly data\(^1\) from 1966.01 to 1985.12\(^2\), we achieved results that contrast sharply with previous findings of others Brazilian authors. 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\(^3\). 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 inflationary financing in order to sustain the debt. There was, then, a choice of economic policy.

The paper is organized as follows. Section 2 describes the Brazilian monetary regime and reports on the causality tests between the real interest rate and the public debt. Section 3 discusses theoretical and empirical issues on seigniorage collection and monetary exogeneity. Section 4 presents causality and exogeneity test results between money creation and inflation rate, and in section 5, we present the conclusion. In Appendix A we provide the statistical procedure for determining the seigniorage-GDP ratio stochastic process. Appendices B and C show, respectively, the statistical reports concerning the causality tests between the real interest rate and the public debt stock, and the cointegration tests between money creation and inflation rate. To save space we do not report all relevant tests and diagnostics but one can access them in working paper 210 in the site www.uff.br/econ/publicações/textosdiscussão.

2. The Brazilian monetary regime

In a monetary regime with an independent monetary authority, the fiscal authority determines the public debt growth, and the monetary authority settles the debt financing composition between bonds and money. In such regime, the Central Bank controls the money creation by following its own rule of monetary expansion, which reflects a trade-off between efficiency and political economy considerations. So, the monetary control is realized through open market transactions, discount loans or reserve requirements.

In this section, we turn our eyes to an economy in which the monetary policy is conducted indirectly, through a particular form of open market transactions. Besides this is an economy with a consistently increasing inflation.

In the Brazilian monetary regime, firstly the Treasury financed itself directly through the Central Bank. Secondly, public bonds were not
sold to the final takers, but rather to financial institutions, which financed themselves through overnight deposits from the private sector. At the same time, the Central Bank informally gave liquidity to the excess of bonds 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, then they could resell their holdings of excess bonds to the Central Bank.

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

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

In a monetary regime with an independent Central Bank, the real interest rate and the public debt rise simultaneously, when the Treasury sells bonds. It is not possible to settle the causal direction, in Granger’s sense. In Brazil, the 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 is true, the real interest rate caused, in Granger’s sense, the public debt oscillations.

In figure 1, we illustrate the logarithm of the monthly debt-GDP ratio (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.

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...
(1999) and are shown in Table B.1. The ADF and the Phillips-Perron tests reveal that the real interest is $I(0)$ and the debt-GDP ratio is $I(1)$ in the analyzed period. In this case, there is no point 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, as well as 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 three or more lags were potentially properly specified. We decided to start on sequential testing with a sixth order VAR. We then tested if the coefficients corresponding to the largest lag are zero, using the likelihood ratio statistics (LR) and information criteria.

The results are reported in Table B.2. For the Schwarz and Hannan-Quinn criteria, the optimal lag length is 3. If we consider only orders between 4 and 6, all criteria will choose a lag length of 4. 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, findings by Giles and Mirza (1998) suggest that the pretesting procedures can result in severe overrejection of the noncausality null hypothesis, whereas the overfitting method results in less distortion in the empirical power. We decided to work with four lags on the VAR.

The VAR diagnostics tests points out the presence of heteroskedasticity and the lack of normality in the residuals system. Since such problems are caused mainly by large outliers concentrated in some parts of the analyzed period, a 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 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 statistics, 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.

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 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). Estimated two standard error bounds are depicted as dotted lines.

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) impulse-response functions (not presented), the effect on the debt-GDP series is weaker, but the return to
the equilibrium is not as quick. There is also an apparent feedback from the debt stock to the real interest rate, that becomes visible in the 5th period, but that is not significant (p-value of 0.386).

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.  

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

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 were the tests performed observing the spherical conditions or empirical distributions were simulated to the test statistics. Each column represents 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 Component GARCH error specification, 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. The results show that R fails to G-cause LDY is rejected with a p-value of 0.1%, while LDY fails to G-cause R is accepted with a p-value near 60%. Thus, the tests demonstrate that the causality in the Granger sense goes unidirectionally from the real interest rate to the bonds stock.

For comparison, we also performed the Geweke, Meese and Dent (GMD) and the Sims causality tests (see footnote 7). The GMD test was specified using the series in the same form as in the VAR system. To the equation that 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. That turned 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 existing 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 G-cause the real interest rate. In the R equation, there are evidences of instantaneous causality from the real interest rate to the debt stock. We con-
include that the GMD tests indicate that the real interest rate $G$-causes the debt-GDP ratio and there is no feedback between them.

**Table 1 – Causality tests (p-values) with non-gaussian residuals***

<table>
<thead>
<tr>
<th>PROCEDURE</th>
<th>GRANGER DIR. TEST</th>
<th>GMD TEST</th>
<th>SIMS TEST</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>LDY</td>
<td>R</td>
<td>LDY</td>
</tr>
<tr>
<td>$R \Rightarrow LDY$</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>F</td>
<td>0.001</td>
<td></td>
<td>0.076</td>
</tr>
<tr>
<td>LR</td>
<td>0.009</td>
<td></td>
<td>0.062</td>
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<tr>
<td>$LDY \Rightarrow R$</td>
<td>0.883</td>
<td>0.923</td>
<td>0.915</td>
</tr>
<tr>
<td>F</td>
<td></td>
<td>0.876</td>
<td>0.915</td>
</tr>
<tr>
<td>LR</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$R \leftrightarrow LDY$</td>
<td>0.013</td>
<td>0.013</td>
<td>0.004</td>
</tr>
<tr>
<td>F</td>
<td></td>
<td>0.010</td>
<td>0.010</td>
</tr>
<tr>
<td>LR</td>
<td></td>
<td></td>
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</tbody>
</table>

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

**Table 2 – Causality tests (p-values) performed with gaussian residuals or Monte Carlo simulations***

<table>
<thead>
<tr>
<th>PROCEDURE</th>
<th>GRANGER DIR. TEST*</th>
<th>GMD TEST</th>
<th>SIMS TEST</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>LDY</td>
<td>R</td>
<td>LDY</td>
</tr>
<tr>
<td>$R \Rightarrow LDY$</td>
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</tr>
<tr>
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<td></td>
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</tr>
<tr>
<td>LR</td>
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<td></td>
<td>0.007</td>
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<td>$LDY \Rightarrow R$</td>
<td>0.548</td>
<td>0.691</td>
<td>0.614</td>
</tr>
<tr>
<td>F</td>
<td></td>
<td>0.629</td>
<td>0.628</td>
</tr>
<tr>
<td>LR</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$R \leftrightarrow LDY$</td>
<td>0.018</td>
<td>0.243</td>
<td>0.005</td>
</tr>
<tr>
<td>F</td>
<td></td>
<td>0.001</td>
<td>0.210</td>
</tr>
<tr>
<td>LR</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

* Monte Carlo p-values: $R \not\Rightarrow LDY \begin{cases} F = 0.001 \end{cases}$, $LDY \not\Rightarrow R \begin{cases} F = 0.866 \end{cases}$

To perform the Sims’ causality test, we first pre-whitened both series using the autoregressive process found by the debt-GDP unit root test; 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. 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 to derive test empirical distributions.

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, when we regress R on LDY, only the debt stock future values are significant. These empirical findings demonstrate the hypothesis that causality in Granger’s sense was unidirectional from the real interest rate to debt-GDP ratio, from 1966 to 1985.

For that reason, 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, in the Sims’ sense, with respect to the demand for bonds; Sargent (1987) and Sims (1972). Accordingly, we claim 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. This establishes the first evidence about the monetary policy independence during the analyzed period.

2.1 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, thus cannot be cointegrated. In addition, the real interest rate caused the debt-GDP ratio, but the latter failed to Granger-cause the real interest. In consequence, an AD(6,6) 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 quartered 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 potential demand shifting. We proceeded with the information criteria and the last significant lag analysis. We therefore reduced the model to an AD(4,4) whose residuals do not present serial correlation; see footnote 7. In order to obtain NIID residuals, we applied an ARCH(3)
specification, a choice supported by diagnostics tests. Since we focus on the demand for bonds we re-parameterized the estimated model to a fully AD(4,4)-ARCH(3) in levels. Bollerslev-Wooldrige standard errors were used in order to remain conservative, given the presence of an integrated series among the regressors.

The hypothesis that the lagged dependent variable coefficients sum up to 1 is rejected with a p-value of 0.01%. The diagnostics tests indicate that the residuals are NIID and are approximately white noise. We have no reason to doubt the stability of the model.

The estimates show the bonds demand was very sensitive to real interest rate variations. The hypothesis of a long run elasticity superior to one is accepted with a p-value superior to 75%. Over the 1966-1985 period, 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 quartered debt-GDP ratio fluctuated between 7.3% and 44.5% with a mean of 21.8%. An increase of 5 percent points in real interest rate would have led to an increase of 4.4% in the debt-GDP ratio during the same quarter. Taking the ratio’s mean value, it would have increased to 22.8%; this would have financed an operational deficit with respect to the GDP of 0.96%. A five month growth of 1 percent point in real interest would have induced an increase of 1.1% in the debt-GDP ratio which, in its mean value, would have grown to 22.1% and would have financed an operational deficit-GDP of 0.24%.

That 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.

3. Financial innovations, money demand and seigniorage collection

Financial innovations in general produce new assets warranted by public bonds. These assets have higher liquidity and less risk of capital
loss than bonds. Hence, for a given level of deficit financed with seigniorage, they cause higher inflation, because they contract the 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 that began at 1976 intensified the innovation process and all but reduced to zero the risk of capital loss. They also gave the financial institutions an almost instantaneous liquidity to any unbalance between their bonds 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.

Even without financial innovations, the search of a given level of seigniorage by the monetary authority may lead to ever-increasing inflation. An increase in the expected rate of inflation 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. The seigniorage component \((\frac{M}{P})^{11}\) may be negative if the monetary authority does not increase the monetary expansion simultaneously with the rise in the expected inflation. However, if the economy is operating on the downward region 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 rate of inflation.

Along the upward side of Laffer curve, a contraction in the money demand will lead to a higher rate of inflation. Furthermore, contractions of money demand may be of such amount that the maximum inflation tax falls below 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.

Is there evidence that the demand for money contractions limited the seigniorage collection and caused the money supply endogeneity? Cerqueira (1993) estimated the demand for real balances for the 1966-1985 period by using quarterly data. It was shown that, owing to financial
innovations, the money demand experienced a contraction 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 that decreased from 4.5% per quarter in 1976 to 3.3% in 1985 (using the average base multiplier between 1976 and 1985). 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 increase. To provide a formal argument, suppose an economy described by the following version of Cagan’s (1956) model:

\[ m \equiv [M/(PY)] = ce^{-\alpha \pi_c} \quad c > 0 \quad \alpha > 0 \]  
(1)

\[ \pi_c = \beta(\pi - \pi_c) \quad \beta > 0 \]  
(2)

where \( \pi_c \) is the expected rate of inflation, \( \alpha \) is the semi-elasticity of the money demand with regard to the expected inflation, and \( \beta \) is the inverse of inflationary memory (the bigger \( \beta \), the smaller the effect of past inflation on the present inflation expectations). We assume a constant rate of growth and a constant real interest rate. For a given level of exogenous money growth \( \mu \), the seigniorage flow is given by:

\[ S = \dot{M}/(PY) = (\dot{M}/M)[M/(PY)] = \mu ce^{-\alpha \pi_c}. \]  
(3)

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

\[ \mu = (\bar{S} / ce^{-\alpha \pi_c}). \]  
(4)
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 a financial innovation, shifts down the reaction curve that 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 (see figure 3), in order to increase its fraction on the seigniorage collection. Actually, Cerqueira (1993) 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.

Thus, the fact that the seigniorage collection was constant during the period (see Appendix A) did not mean that the monetary policy was passive regarding the inflation rate. Furthermore, it helps to explain why the estimated steady state rates of inflation remained approximately constant from 1976 to 1985. Indeed, the policy of reducing the base

**FIGURE 3: BASE MULTIPLIER**
multiplier had 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 such effect. Moreover, even if the monetary authority had increased the money growth to compensate for this effect, the observed inflation and monetary rates, while increasing, were very distant from the estimated unstable levels, thus far from the path of a hyperinflationary disequilibrium15; see Cerqueira (1993).

4. 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 inflation rate. Lagged rates of money creation add nothing to the predictions of inflation and their own 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, whereas 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 4) 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 imply 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 takes place 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 from 1966 to 1985. It was argued in section 2 that these fluctuations caused the public debt stock. Since the debt stock held by the Central Bank is a non-monetary liability, it is an instrument of monetary control.

In Appendix A we show that the seigniorage collection as a GDP proportion could be taken as constant with a mean value of 1.82% 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 A.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 a constant seigniorage collection that assured the debt sustainability while preserving the exogeneity of the money supply. This means there was a steady state level of public deficit financed with money creation, as in Cagan’s adaptive scheme. To reach the proof of such hypothesis, it is “necessary” to show that money creation was exogenous respecting inflation, while inflation was determined by money creation (Sargent and Wallace, 1973).
The first step in searching for a causal relation is to determine the integration order of the series. Table C.1 reproduces the unit root tests for both variables, taken from Cerqueira (1999). 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. Cerqueira (1999), by using the bootstrap approach on the ADF statistics, 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 inflation rate 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). 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 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 dummies19 acting as exogenous variables and excluded from the long-run relation in the Johansen procedure. This changes the asymptotic distributions but 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 pre-testing also identified intrinsic explosiveness in the data by detecting roots lying outside the unit circle. This kind of non-stationarity cannot be removed by differencing and indicates the inadequacy of the
chosen model (Johansen, 1996). 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. Pre-testing suggested the inclusion of a linear trend in the cointegration space and of exogenous policy dummies to address this non-stationarity 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. Then the lag length was selected using two types of information criteria (Schwarz and Hannan-Quinn). The pre-test results are shown in Table C.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, whereas 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 precision. Findings from Giles and Mirza (1998) suggest that the pre-testing procedures can result in severe overrejection of the null of non-causality while overfitting methods cause less distortion with often little or no loss of power. Their suggestion is to abandon pre-testing 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 C.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 15%. Thus, at 15% we reject the null of no cointegration vector and cannot reject the null of one cointegrating vector. Although we used the 15% level, it is worth mentioning that the $\lambda_{\text{max}}$ statistics rejects the null of zero cointegration vector at the 5% level.
The inspection of the roots of the companion matrix (not reported) indicates that there are two roots very close to 1 (-0.9653 and 0.9128), but none of the others is close to other points in the unit circle. Consequently, 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 from 1966 to 1985. Moreover, it 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 between inflation and money growth has cointegrating vector \(\beta=(1,-1)\) and a moving stationary drift term\(^{21}\). This could represent the output rate of growth, an I(0) variable; see Cerqueira (1999).

Diagnostics tests reported in Tables C.4 and C.5 show that the restriction on the cointegration space (by setting the cointegration rank equal 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 vanishes in the estimated system, with money growth variable taken as weakly exogenous. The diagnostics tests of the error-correction model show that our choice of 15 lags for the VAR was appropriate. Firstly, because the residuals have no serial correlation, and secondly, for choosing the thirteen-lag VAR would have led to the estimation of a misspecified VEC model with autocorrelated residuals.

Table C.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. Those results corroborate the implicit idea in equation (5). A monetary shock leads the acceleration of expected inflation, which by its turn increases the inflation.
rate drifts above its steady state path, and since the adjustment coefficient \( a \) is negative\(^{22} \), 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\(^{23} \). Below each table, we also report the strong-exogeneity tests\(^{24} \). When the residuals were not Gaussian, we relied on GARCH models, or on the simulation of the test statistic distribution following the bootstrap approach. The Granger direct test specification is taken from the VECM estimates\(^{25} \). The methodology for performing the other tests followed the same paths described in section 3; see footnote 7.

**Table 4 – Causality tests (p–values) with non-gaussian residuals**

<table>
<thead>
<tr>
<th>PROCEDURE</th>
<th>GRANGER DIR. TEST</th>
<th>GMD TEST</th>
<th>SIMS TEST</th>
</tr>
</thead>
<tbody>
<tr>
<td>PI (\Rightarrow) MI</td>
<td>MI: 0.746, PI: 0.574</td>
<td>MI: 0.206, PI: 0.000</td>
<td>MI: 0.678, PI: 0.521</td>
</tr>
<tr>
<td>MI (\Rightarrow) PI</td>
<td>F: 0.001, LR: 0.000</td>
<td>F: 0.002, PI: 0.001</td>
<td>F: 0.005, PI: 0.005</td>
</tr>
<tr>
<td>MI (\Leftrightarrow) PI</td>
<td>F: 0.379, PI: 0.353</td>
<td>F: 0.894, PI: 0.850</td>
<td></td>
</tr>
</tbody>
</table>

\(^*\) The symbol \( \not\Rightarrow \) means “does not Granger–cause”. Strong exogeneity: \( \pi \not\Rightarrow \mu \) \( \{ F = 0.787 \} \), \( \mu \not\Rightarrow \pi \) \( \{ F = 0.000 \} \) \( \{ LR = 0.625 \} \).

**Table 5 – Causality tests (p–values) performed with gaussian residuals or Monte Carlo simulations**

<table>
<thead>
<tr>
<th>PROCEDURE</th>
<th>GRANGER DIR. TEST</th>
<th>GMD TEST</th>
<th>SIMS TEST</th>
</tr>
</thead>
<tbody>
<tr>
<td>PI (\Rightarrow) MI</td>
<td>MI: 0.750, PI: 0.574</td>
<td>MI: 0.611, PI: 0.371</td>
<td>MI: 0.691, PI: 0.548</td>
</tr>
<tr>
<td>MI (\Rightarrow) PI</td>
<td>F: 0.002, LR: 0.002</td>
<td>F: 0.050, PI: 0.000</td>
<td>F: 0.007, PI: 0.009</td>
</tr>
<tr>
<td>MI (\Leftrightarrow) PI</td>
<td>F: 0.151, PI: 0.029</td>
<td>F: 0.896, PI: 0.828</td>
<td></td>
</tr>
</tbody>
</table>

\(^*\) Monte Carlo p–values of the strong exogeneity tests: \( \pi \not\Rightarrow \mu \) \( \{ F = 0.792 \} \), \( \mu \not\Rightarrow \pi \) \( \{ F = 0.001 \} \) \( \{ LR = 0.792 \} \).

The results from the causality tests indicate that money growth G-causes inflation while the inflation rate fails to G-cause the monetary

expansion. The surprising exception is the GMD test, which indicates a significant feedback from inflation to money when the dependent variable is the inflation rate. Nevertheless, if the residuals are specified as a GARCH process, we obtain the same results as we get with the other tests. 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. It may be interpreted as adjustments in monetary policy following a monetary shock. This is illustrated by the impulse-response functions 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 claim.

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. 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, near 0.33 points impact on the inflation rate. When innovations are considered 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 computed by the variance decomposition of forecast error of inflation; see footnote 7. Therefore, we conclude that the existing “persistence” in the inflation was due to mainly monetary causes, rather than being triggered by disturbances arisen in the “real side” of the economy.

We can conclude that there is enough evidence to validate the hypothesis that money growth is strongly exogenous concerning the inflation rate, from 1966 to 1985. Evidence supports the claim 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. Moreover, it is a strong indication that the monetary authority did not follow a reaction curve like equation (4) as a monetary rule.

5. Conclusions

This paper presents tests on the exogeneity of Brazilian monetary supply for the military period comprising 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 that the monetary authority worked towards increasing the real interest rate to induce the public bond demand. They also show that, though from 1974 to 1985 the seigniorage collection was predetermined in a constant level, this policy did not lead to the endogeneity of the money supply, since money growth
was strongly exogenous with respect to the inflation rate. The results show that the monetary expansion caused in Granger’s sense the inflation rate, and the former was weakly exogenous with respect to the latter. This was possible probably because the monetary authority chose to reduce the base multiplier in order to keep its proportion of seigniorage collection. Therefore, even with (i) a permanent deficit with the seigniorage playing a crucial role in balancing the public accounts, and (ii) a host of financial innovations that led to the money demand contraction, the money supply remained exogenous respecting the inflation rate. Therefore, Brazilian inflation followed an ever-increasing path without ‘detonating’ a hyperinflationary process.

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 assuming that the monetary policy was completely passive during the seventies and eighties.

Indeed, our results reveal that the monetary policy was executed in an independent way, that is, the rule guiding the monetary execution was taken exogenously with regard to the considered model and the inflation rate. Therefore, we postulate that the monetary authority chose to finance a rough fraction of the public deficit by issuing money, which explains the intermittent monetary expansion and the inflation rate. This policy 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.

To reject the causality from prices to money does not mean to propose there was rigid monetary control. This depends on the monetary regime. In a regime whose monetary authority is independent, it is able to fulfill almost any target of money stock. Compelled by the public deficits, the monetary authority may refuse to buy public bonds in the open market and then impose upon the fiscal authority the cost of increasing
the real interest rate through the primary auctions. In the Brazilian regime, this responsibility was a burden to the Central Bank. It was enough that the real interest rate 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 enough to manage an exogenous money supply from the inflation variations. Perhaps these are somewhat old monetarist ideas, but we cannot deny they stamped their mark on the data.

Notes

I thank Bruno Larue, Benoît Carmichael and two anonymous referees for comments and suggestions. Financial support from CNPq and FAPERJ-Brazil is gratefully acknowledged. All errors are mine.

1 With the exception of the seigniorage/GDP series, all remaining data were not previously adjusted.

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

3 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%.

4 The ratio is defined as the government bonds stock held by private agents (published in Brazilian Central Bank bulletin) over the nominal GDP; see Cerqueira (2006) for the applied interpolation procedure.
The nominal interest rate is the overnight rate yielded by the three months treasury bonds most negotiated in the monetary market. This series is published in the Brazilian Central Bank bulletin; see Cerqueira (1993) for further details. We converted the series to the equivalent monthly rate. The rate of inflation is calculated with the general price index (IGP) computed by Fundação Getúlio Vargas.

As mentioned in the introduction these diagnostics reports are presented in the paper full version – TD 210 – that can be accessed in the site www.uff.br/econ/publicações/textosdiscussão.

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, it is not significant at either 5 or 10%. We can also construct a 95% confidence interval around the residuals inner product (=0.028) with bounds [-0.278; 0.333]. This shows the consistency of the residual orthogonality hypothesis. Similar conclusions apply to the VAR(3).

Saying that the debt-GDP increased is the same as saying that the stock of public bonds held by the public increased, which means the demand for bonds increased.

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.

The concepts of predeterminedness and strict exogeneity are neither sufficient nor necessary to obtain the parameters of interest of the equation between LDY and R. None of the concepts is sufficient for efficient (without loss of information) statistical inference or at all necessary (see Engle, Hendry & Richard, 1983 and Ericsson, 1994). Only if R were weakly exogenous for the parameters of interest in the conditional model of LDY, efficient estimation and testing could be conducted by analyzing only this model, ignoring the information of R marginal process. Unfortunately, the series cannot cointegrate, which makes it hard to perform weak exogeneity testing. Being aware of this drawback, we decided to propose an AD process in order to study the dynamic relation between those variables.

The seigniorage is the sum of the inflation tax with the growth of the real monetary base. This can be seen by differentiating the real base (M/P) respecting time, then we get: $\frac{M}{P} = (\frac{M}{P})\pi + (\frac{M}{P})$.

The reduction in the money demand is represented by the decrease in its constant term; see equation (1).

The maximum seigniorage is given by $S^* = \frac{c}{\alpha ek}$ (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 $IT = \pi\frac{B}{(PY)} = \pi[M1/(PY)](1/k)$. 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.

If we combine equations (1) and (2), we get the equation of the dynamics of the expected rate of inflation: $\hat{\pi}_c = [\beta/(1 - \alpha\beta)](\mu - \pi_e)$, whether money creation is
exogenous or not. If $S$ is less than $S^*$, there are two equilibria: $\pi^e_1$, which is the low-inflation equilibrium, and $\pi^e_2$, the high inflation equilibrium. If the Cagan stability condition $ab < 1$ is respected, for initial expectations $\pi^0 < \pi^e_1$, $p_e$ converges to $\pi^e_1$, a stable target, since $\dot{\pi}_e < 0$. If $\pi^0 > \pi^e_2$, then $p_e$ increases indefinitely ($\dot{\pi}_e > 0$). If $S = S^*$, then there is only one steady state point $\pi^e = 1/a$.

15 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%.

16 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%. This decrease was due to the money demand contraction that took place during the period, reducing the ability to collect seigniorage, and so augmenting the burden of the financing with public debt.

17 That is, the seigniorage collection equaled the inflationary tax-GDP ratio; see appendix A.

18 By money creation we mean the percent variation of M1 (currency plus checking accounts) monetary aggregate computed by Brazilian Central Bank.

19 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 $t=1985.01$; $D85 = 1$ if $t=1985.08$, -1 if $t=1985.09$, and 1 if $t=1985.11$; $Dum84 = 1$ if $t \geq 1984.01$; 0 otherwise. The set of chosen dummies includes indicator variables that have the effect of deleting the observation from the least squares computation. This is an admitted approach if the deleted outliers are data generated by the same mechanism as the others but were drawn from the tails of the distribution. Since the data we employed in the paper have a homogeneous definition during the whole sample, we may assume that they were drawn from the same distribution (see Cerqueira 1993 for further details on the data treatment).

20 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 1, we get the MLR statistics $c^2(14) = 153.564$, which implies a p-value very near zero. Then we reject the restricted model without dummies.

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

22 The value of $a$ was found to be $-0.170$.

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

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

25 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.
The fact that all tests achieved the same results demonstrates they are robust to the applied noncausality-type test. Moreover, our empirical trials reveal that the violation of homoskedasticity and normality assumptions, in general, does not affect very seriously the causality tests performance.

As stated 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 (-0.170,0). The other VECM components are the same as those mentioned in the text. We remark that 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 0.093, with a p-value of 0.15, thus not significant at either 5 or 10%. A likelihood test using the VECM's covariance matrix also confirms that the two residuals are independent at 0.744 probability value. Since we cannot reject the hypothesis of a zero inner value (=0.004) with a p-value of 0.223, we still have an additional evidence of orthogonality.

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

This conclusion contradicts Marques (1983), Pastore (1994), and specially Pastore (1997), but meets 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 “received wisdom” among Brazilian economists.

Data were taken from the Central Bank bulletin.

To the ADF test the lag truncation was made based on the t-test of significance of the last lagged first difference whereas to the Phillips-Perron the lag selection was done with the Bartlett kernel.

The test statistics equals 0.0898, and the critical value of a two-sided test of size 10% is 0.125; see Harvey (1990).

In this period the inflation rate is I(0). The ADF test performed with 18 lags, and a trend term reject the null with significance level inferior to 1%.

The notation follows the tradition used in the economy literature. We reported the computed values to the statistics R², SER, AIC, SIC, and DW. The F-statistics, the Ljung-Box, the Breusch-Godfrey LM version, and the Bera-Jarque are reported with the respective p-values. We also reported the p-values to the White, ARCH, and Q² tests of heteroskedasticity. We employed the White heteroskedasticity correction due to the presence of some ARCH processes in the residuals.

Since the residuals are not normal, we computed the likelihood and F Monte Carlo distributions. The Monte Carlo p-values associated with the constant term are equal to 0.001, while for the inflation coefficient both p-values equal 0.292. If the errors are specified as an ARCH (1) process the residuals are normal (BJ=0.135) but the remaining reported results are the same.

Using ADF or Phillips-Perron unit root tests with significance level inferior to 1%, we reject the hypothesis that the series first difference has a unit root. The debt-GDP
series is therefore difference stationary, which implies the public debt was sustainable and the government intertemporal budget constraint was being satisfied.

During the studied period, the inflation tax-GDP ratio mean and median values were respectively around 2.0% and 1.75%. Then, with p-values of 0.275 and 0.265, we cannot reject the mean and the median equality hypotheses between this series and the seigniorage-GDP ratio.

Exogeneidade da oferta de moeda no Brasil entre 1966 e 1985

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

Palavras-chave – Modelos de séries temporais, modelagem econométrica, bootstrap, inflação, oferta de moeda, política monetária.

References


Appendix A: 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\(^{30}\) (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. The real GDP series is taken from the interpolation made possible by the employment of Kalman filtering; see footnote 7.

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 inflation rate, so we may presume it might be described as a white noise process; see figure A.1. This hypothesis is based on the increasing cost of holding money and on the process of money replacement by other financial assets; 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 in which 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 month is associated with the increasing tendency shown 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. This procedure brought into the series SY depicted in figure A.1.

Our statistical approach was then to proceed using independence tests on the adjusted series. We first performed the unit root tests reported in Tables A.1\(^{31}\) and A.2. The tests indicate the seigniorage-GDP ratio is stationary. In figure A.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%32. The spectral shape also suggests that the SY series is white noise; see footnote 7.

The seigniorage-GDP series during the period is described by the following statistics: mean=1.818%, std.dev.=1.947%, skew.=0.270, kurt.=4.361. However, it is not normally distributed considering that the Bera-Jarque statistics is rejected with a p-value near 0.004%. If the above results are corrected we will expect the regression of SY against a constant and the inflation rate33 must present the following outcomes: (i) constant significant at the level of the sample mean (1.818%); (ii) coefficient of pi non-significant; (iii) R² near zero; (iv) residuals approximately IID. Those results are testified by the report34 below.
Table A.1 - Unit root tests report

<table>
<thead>
<tr>
<th>SERIES</th>
<th>TEST</th>
<th>LAGS</th>
<th>t-α</th>
</tr>
</thead>
<tbody>
<tr>
<td>SY</td>
<td>ADF</td>
<td>0</td>
<td>-11.817**</td>
</tr>
<tr>
<td></td>
<td>PP</td>
<td>4</td>
<td>-11.895**</td>
</tr>
</tbody>
</table>

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 A.2 - ADF test: residuals diagnostics

<table>
<thead>
<tr>
<th></th>
<th>DW=2.004</th>
<th>Q(24)=0.840</th>
<th>Q(56)=0.785</th>
<th>LM(1)=0.715</th>
<th>LM(3)=0.576</th>
<th>LM(6)=0.685</th>
</tr>
</thead>
<tbody>
<tr>
<td>LM(9)=0.708</td>
<td>LM(12)=0.523</td>
<td>Bj=0.001</td>
<td>AIC=5.054</td>
<td>SIC=4.998</td>
<td>SER=0.019</td>
<td></td>
</tr>
</tbody>
</table>

With p-value equal to 0.448, we cannot reject the hypothesis that the equation constant term is approximately equal to the series sample mean. We cannot also reject the assumption of a non-significant inflation coefficient and that the $R^2$ and the residuals are as expected. Thus, the seigniorage collection could be taken as constant and uncorrelated with the inflation rate, which had no effect in changing this government revenue.

All the above results corroborate the hypothesis that from 1974.01 to 1988.06 the seigniorage-GDP ratio followed a white noise process ~ (1.818%;1.947%).

Table A.3 - Ols regression SY against PI

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coeff.</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>C</td>
<td>0.020218</td>
<td>0.0000</td>
</tr>
<tr>
<td>PI</td>
<td>-0.030884</td>
<td>0.3550</td>
</tr>
</tbody>
</table>

(White Heteroskedasticity S.E. considered)

This conclusion gives empirical support to the assumption that the monetary policy was not passive during the 1974-1988 period, in the sense that the inflation rate 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, 1999), without making the money supply endogenous. This means that there was a steady state level of public deficit financed with...
issuing money\textsuperscript{37}. To complete proving the hypothesis of a money supply exogeneity, it is “necessary” to show the money creation was \textit{exogenous} respecting inflation, while the reverse was not true for the inflation, a task that is realized in section 4 of the present paper.

Appendix B: 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.

<table>
<thead>
<tr>
<th>Table B.1 – Unit root tests</th>
<th>Period: 1966.01 to 1985.12</th>
</tr>
</thead>
<tbody>
<tr>
<td>SERIES</td>
<td>ADF</td>
</tr>
<tr>
<td></td>
<td>LAGS</td>
</tr>
<tr>
<td>Real Interest Rate (R)</td>
<td>2</td>
</tr>
<tr>
<td>Debt-GDP Ratio (LDY)</td>
<td>3</td>
</tr>
<tr>
<td>ΔLDY</td>
<td>2</td>
</tr>
</tbody>
</table>

Note: (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>
<thead>
<tr>
<th>Table B.2: VAR LAG truncation</th>
</tr>
</thead>
<tbody>
<tr>
<td>ORDER</td>
</tr>
<tr>
<td>-------</td>
</tr>
<tr>
<td>6</td>
</tr>
<tr>
<td>5</td>
</tr>
<tr>
<td>4</td>
</tr>
<tr>
<td>3</td>
</tr>
</tbody>
</table>

\*LR statistics has a c$'$(4) distribution.
Appendix C: Cointegration relation test between monetary expansion and inflation rate

Table C.1 shows the unit root tests, and Table C.2 the computed statistics for selecting the unrestricted VAR lag truncation. Table C.3 reports the Johansen’s procedure in the conventional way. In the diagnostics tests reports (tables C.4 and C.5), we also present the Bowman & Shenton (BS) normality test described in Hansen and Juselius (1995).

### Table C.1 – Unit root tests

<table>
<thead>
<tr>
<th>SERIES</th>
<th>ADF</th>
<th>Phillips-Perron</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>LAGS</td>
<td>t-α</td>
</tr>
<tr>
<td>Inflation Rate (PI)</td>
<td>2</td>
<td>-2.582</td>
</tr>
<tr>
<td>ΔPI</td>
<td>1</td>
<td>-9.549**</td>
</tr>
<tr>
<td>Money Growth (MI)</td>
<td>11</td>
<td>0.507</td>
</tr>
<tr>
<td>ΔMI</td>
<td>10</td>
<td>-8.849**</td>
</tr>
</tbody>
</table>

Note: (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 C.2 – VAR LAG truncation

<table>
<thead>
<tr>
<th>VAR ORDER</th>
<th>CONSTANT</th>
<th>LINEAR TREND</th>
<th>SCHWARZ CRITERION</th>
<th>HANNAN-QUINN CRITERON</th>
</tr>
</thead>
<tbody>
<tr>
<td>15</td>
<td>unrestricted</td>
<td>cointegration space</td>
<td>-14.956</td>
<td>-15.735</td>
</tr>
<tr>
<td>14</td>
<td>unrestricted</td>
<td>cointegration space</td>
<td>-14.871</td>
<td>-15.702</td>
</tr>
<tr>
<td>15</td>
<td>unrestricted</td>
<td>cointegration space</td>
<td>-14.796</td>
<td>-15.662</td>
</tr>
<tr>
<td>16</td>
<td>unrestricted</td>
<td>cointegration space</td>
<td>-14.723</td>
<td>-15.623</td>
</tr>
<tr>
<td>17</td>
<td>unrestricted</td>
<td>cointegration space</td>
<td>-14.656</td>
<td>-15.571</td>
</tr>
</tbody>
</table>

### Table C.3 – Johansen’s cointegration test

<table>
<thead>
<tr>
<th>TEST STATISTICS</th>
<th>COINTEGRATING VECTOR</th>
</tr>
</thead>
<tbody>
<tr>
<td>λ-max Trace</td>
<td>(MONEY, INFLATION, TREND, CONSTANT)</td>
</tr>
<tr>
<td>r=0 r≤1</td>
<td>r=0 r≤1</td>
</tr>
<tr>
<td>15.555* 7.942</td>
<td>21.477 7.942</td>
</tr>
<tr>
<td>(11.977) (9.427)</td>
<td>(21.404) (9.427)</td>
</tr>
</tbody>
</table>

COINTEGRATION RESTRICTION TEST

RESTRICTION: (1,-1,#,#) $\chi^2(1) = 0.38; \text{P-VALUE} = 0.566$

Note: The symbol (*) indicates rejection of the null at the 5% significance level. The symbol # means the parameter is unrestricted.
Table C.4 – Unrestricted VAR diagnostics tests

<table>
<thead>
<tr>
<th>MULTIVARIATE TESTS</th>
<th>LM(1)</th>
<th>LM(4)</th>
<th>BS</th>
</tr>
</thead>
<tbody>
<tr>
<td>Q(60) = 0.078</td>
<td>0.539</td>
<td>0.795</td>
<td>0.001</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>UNIVARIATE TESTS</th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>EQ.1 (ΔPI): Q(60)</td>
<td>0.727</td>
<td>BS = 0.538</td>
<td>Bj = 0.071</td>
</tr>
<tr>
<td>EQ.2 (ΔM): Q(60)</td>
<td>0.955</td>
<td>BS = 0.010</td>
<td>Bj = 0.002</td>
</tr>
</tbody>
</table>

Table C.5 – Restricted VAR (VECM) diagnostics tests*

<table>
<thead>
<tr>
<th>MULTIVARIATE TESTS</th>
<th>LM(1)</th>
<th>LM(4)</th>
<th>BS</th>
</tr>
</thead>
<tbody>
<tr>
<td>Q(60) = 0.109</td>
<td>0.290</td>
<td>0.802</td>
<td>0.014</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>UNIVARIATE TESTS</th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>EQ.1 (ΔPI): Q(60)</td>
<td>0.682</td>
<td>BS = 0.525</td>
<td>Bj = 0.107</td>
</tr>
<tr>
<td>EQ.2 (ΔM): Q(60)</td>
<td>0.910</td>
<td>BS = 0.101</td>
<td>Bj = 0.026</td>
</tr>
</tbody>
</table>

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

In table C.6, 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 C.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 C.6 – Adjustment – Coefficient weak exogeneity test

<table>
<thead>
<tr>
<th>NULL HYPOTHESIS</th>
<th>TEST STATISTIC</th>
<th>P-VALUE</th>
</tr>
</thead>
<tbody>
<tr>
<td>MI is weakly exogenous for the parameter of interest of the 1 conditional model</td>
<td>0.54</td>
<td>0.844</td>
</tr>
<tr>
<td></td>
<td>(0.23)</td>
<td>(0.632)</td>
</tr>
<tr>
<td>PI is weakly exogenous for the parameter of interest of the 1 conditional model</td>
<td>10.91</td>
<td>0.004**</td>
</tr>
<tr>
<td></td>
<td>(5.12)</td>
<td>(0.024)</td>
</tr>
</tbody>
</table>

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 {F = 0.935}; PI "IS W.E. TO MI" {F = 0.001}. {LR = 0.935}; {LR = 0.001}