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PUBLIC DEBT SUSTAINABILITY AND ENDOGENOUS SEIGNORAGE IN BRAZIL: TIME-SERIES EVIDENCE FROM 1947-92

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Abstract

Using national accounts data for the revenue-GDP and expenditure-GDP ratios from 1947 to 1992, we examine two central issues in public finance. First, was the path of public debt sustainable during this period? Second, if debt is sustainable, how has the government historically balanced the budget after shocks to either revenues or expenditures? The results show that (i) public deficit is stationary (bounded asymptotic variance), with the budget in Brazil being balanced almost entirely through changes in taxes, regardless of the cause of the initial imbalance. Expenditures

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are weakly exogenous, but tax revenues are not; (ii) a rational Brazilian consumer can have a behavior consistent with Ricardian Equivalence; (iii) *seignorage* revenues are critical to restore intertemporal budget equilibrium, since, when we exclude them from total revenues, debt is not sustainable in econometric tests.

1. Introduction

At least since the end of WWII the Brazilian economy has been plagued with chronic public deficits and relatively high inflation. However, there were only very few episodes when a sharp increase in public debt was observed. Using national accounts data for the revenue-GDP and the expenditure-GDP ratios from 1947 to 1992, this paper studies two central issues in public finance. First, was the path of public debt sustainable during this period? Second, if debt is sustainable, how has the government historically balanced the budget after shocks to either revenues or expenditures are observed? For example, given an unpredictable increase in expenditures, there are two polar forms to balance the budget. One is to increase the present value of taxes and the other is to decrease the present value of expenditures. From the point of view of a rational Brazilian taxpayer, it is important to learn to what extent these two forms of balancing the budget have occurred.

Following Hamilton and Flavin(1986) and Bohn(1991), these issues are investigated using unit-root tests, cointegration tests, and calculating unconventional impulse-response functions based on Vector Error-Correction Models (VECM's) where a budget balance restriction is imposed[1]. Moreover, this framework is used to test hypotheses implicit on the *tax-and-spend* and *spend-and-tax* models[2].

This paper has three main findings. First, public deficit is stationary (bounded asymptotic variance), with the budget in Brazil being balanced almost entirely through changes in taxes, regardless of whether the initial imbalance was due to a change in expenditures or in taxes. Expenditures are weakly exogenous in the sense of Engle, Hendry and Richard(1983). The evidence points towards the spend-and-tax rather than the tax-and-spend model.

[1]We use this label for the impulse-response function because the response is not calculated for a given series of the system but for its present value innovation.

[2]See Von Furstemberg(1986), and Miller and Russek(1990).

Second, a rational Brazilian consumer can have a behavior consistent with Ricardian Equivalence (Barro(1974)), since, for example, given a fixed expenditure sequence, a tax break today will be fully reverted by future tax increases, leaving consumption and welfare unchanged. Third, we show that *seignorage* revenues are critical to restore intertemporal budget equilibrium, since, when we exclude them from total revenues, debt is not sustainable in econometric tests. These results match the conventional wisdom about Brazilian public finance and are broadly consistent with the theoretical model of seignorage of Bruno and Fischer(1990) and the extension in Ruge-Murcia(1995), and with the empirical findings of Pastore(1995) and Ruge-Murcia.

The paper is organized as follows: in Section 2 the methodology is presented; in Section 3 the data set is discussed; in Section 4 the empirical results are presented and in Section 5 we present the conclusions.

2. Methodology

The econometric techniques used here to test whether or not debt is sustainable follow Hamilton and Flavin(1986), and Bohn(1991). The calculations of the unconventional impulse-response function follow Bohn(1991). We discuss a slight caveat to this approach. Exogeneity tests are performed following the typology in Engle, Hendry and Richard(1983); see also the test implementation in Johansen(1995). A brief discussion of these techniques is presented here for the sake of completeness.

The Government budget constraint can be written in the following form:

$$B_{t+1} = G_t - T_t + (1+r)B_t + \epsilon_{t+1}, \tag{2.1}$$

where T_t represents fiscal revenues including seignorage, G_t represents fiscal expenditures excluding debt-service payments, r is the real interest rate (assumed fixed), B_t is beginning of period public debt, and ϵ_{t+1} is a stationary measurement error inherited from assuming that $r_t = r$ for all t. All time series are measured as a proportion of GDP.

Without loss of generality, we work with the following version of equation (2.1):

$$B_{t+1} = G_t^* - T_t + B_t + \epsilon_{t+1}, \tag{2.2}$$

where $G_t^* = G_t + rB_t$. Disregarding measurement error, and rearranging (2.1) we get:

$$B_{t+1} - B_t = G_t^* - T_t = Def_t, \tag{2.3}$$

where Def_t is the public deficit in period t. Equation (2.3) is the basis for the debt sustainability test of Hamilton and Flavin(1986). It clearly shows that if B_t is difference-stationary, there is a long-run relationship between G_t^* and T_t, and Def_t will not be an integrated series. Thus, debt sustainability is tested via a unit-root test on ΔB_t.

A similar argument can be made from an intertemporal perspective. The government intertemporal budget constraint is:

$$B_t = \sum_{j=0}^{\infty} \rho^j \, E_t \left[T_{t+j} - G_{t+j}^* - \epsilon_{t+j+1} \right], \tag{2.4}$$

where $\rho = \frac{1}{(1+r)}$ is the one-period discount rate for future taxes and expenditures.

Trehan and Walsh(1988) showed that (2.4) holds whenever public debt is difference-stationary. From (2.3), this last condition implies that G_t^* and T_t cointegrate with coefficient $(1, -1)$. This is the test proposed in Bohn(1991) to check whether or not debt is sustainable.

From the *Granger Representation Theorem*, e.g., Engle and Granger(1987), we have that every cointegrating system of $X_t = (G_t^*, T_t)'$, has an error-correction representation[3]:

$$A(L)\Delta X_t = -\alpha \, \beta' X_{t-1} + \mu_t, \tag{2.5}$$

or, if (2.3) holds in I(0) space,

$$A(L)\Delta X_t = -\alpha \, Def_{t-1} + \mu_t \tag{2.6}$$

where β is the cointegrating vector, α is the adjustment coefficient vector of the error-correction term, and μ_t is a multivariate white-noise process. The matrix polynomial $A(L)$, of order k in the lag operator L, satisfies: $A(0) = I$, and $A(1) = \alpha\beta'$ is finite. Using these results we rewrite (2.6) as a first-order system

[3] Although we are basically using $X_t = (G_t^*, T_t)'$ as the vector of fiscal variables, Bohn suggests augmenting it to include series that may increase the information set of the econometrician. Examples are GNP, hours worked, etc.

of equations. This will be useful in understanding the unconventional impulse-response function. Start by considering the following identity:

$$Def_t = \beta' \Delta X_t + Def_{t-1}. \tag{2.7}$$

Since $A(L) = I - A_1 L - A_2 L^2 - ... - A_k L^k$, we can rewrite (2.6) as:

$$\Delta X_t = A_1 \Delta X_{t-1} + \cdots + A_k \Delta X_{t-k} - \alpha\, Def_{t-1} + \mu_t. \tag{2.8}$$

Using identity (2.7) lagged in equation (2.8) and rearranging:

$$\begin{aligned} \Delta X_t &= \left(A_1 - \alpha\beta'\right) \Delta X_{t-1} + \cdots + A_k \Delta X_{t-k} \\ &\quad -\alpha\, Def_{t-2} + \mu_t. \end{aligned} \tag{2.9}$$

After repeating this process $k-1$ times, we get:

$$\Delta X_t = A_1^* \Delta X_{t-1} + \cdots + A_k^* \Delta X_{t-k} - \alpha\, Def_{t-k-1} + \mu_t, \tag{2.10}$$

where $A_i^* = A_i - \alpha\beta'$. Using (2.10) and stacking below a few obvious identities, we finally get a first-order stationary representation of the system as follows:

$$\begin{aligned} \begin{pmatrix} \Delta X_t \\ \Delta X_{t-1} \\ \vdots \\ \Delta X_{t-k+1} \\ Def_{t-k} \end{pmatrix} &= \begin{pmatrix} A_1^* & A_2^* & \cdots & A_k^* & -\alpha \\ I & 0 & \cdots & 0 & 0 \\ \vdots & & & \vdots & \vdots \\ 0 & \cdots & I & 0 & 0 \\ 0 & \cdots & 0 & \beta' & 1 \end{pmatrix} \begin{pmatrix} \Delta X_{t-1} \\ \Delta X_{t-2} \\ \vdots \\ \Delta X_{t-k} \\ Def_{t-k-1} \end{pmatrix} \\ &\quad + \begin{pmatrix} \mu_t \\ 0 \\ \vdots \\ 0 \\ 0 \end{pmatrix}, \end{aligned} \tag{2.11}$$

or, compactly as:

$$X_t^* = A^* X_{t-1}^* + \mu_t^*. \tag{2.12}$$

where $X_t^* = \left(\Delta X_t', \Delta X_{t-1}', \cdots, \Delta X_{t-k+1}', Def_{t-k}\right)'$ and $\mu_t^* = \left(\mu_t', 0, \cdots, 0\right)'$ are $nk+1$ vectors, and the matrix A^* is $[nk+1]$ by $[nk+1]$. As we discuss below, the matrix A^* is critical in finding the response of the system to innovations in the variables. Indeed, the response j-periods ahead of an innovation in one variable in the system can be calculated by selecting the appropriate row of $(A^*)^j$.

The implications of the intertemporal budget constraint on the behavior of fiscal variables of the system can now be calculated using (2.12). To see how, consider again equation (2.4) in its two possible forms, using either G_{t+j} or G_{t+j}^*:

$$(1+r) \cdot B_t = \sum_{j=0}^{\infty} \rho^j \, E_t \left[T_{t+j} - G_{t+j} - \epsilon_{t+j+1}\right], \tag{2.13}$$

or,

$$B_t = \sum_{j=0}^{\infty} \rho^j \, E_t \left[T_{t+j} - G_{t+j}^* - \epsilon_{t+j+1}\right], \tag{2.14}$$

For an unchanged debt value, and disregarding the error term, any increase in expenditures (not accompanied by an increase in taxes) would, in the future, either require a decrease in expenditures or an increase in taxes. Moreover, in present value terms, all future changes either in taxes or expenditures should be exactly equal to the initial change in expenditures (a similar result applies for a change in taxes). To prove this result, we start by defining the present value of a generic variable z, labelled $PV(z)_t$, $PV(z)_t = \sum_{j \geq 1} \rho^j \, z_{t+j}$, an innovation in z, labelled $\widehat{z}_t$, $\widehat{z}_t = z_t - E_{t-1} z_t$, and an innovation on the present value of z, labelled $\widehat{PV}(z)_t$, $\widehat{PV}(z)_t = E_t PV(z)_t - E_{t-1} PV(z)_t$[4]. Note that we can rewrite (2.13) as:

$$\begin{aligned}(1+r)B_t &= T_t + E_t PV(T)_t \\ &\quad - \left[G_t + E_t PV(G)_t + E_t \left(\frac{PV(\epsilon)_t}{\rho}\right)\right],\end{aligned} \tag{2.15}$$

or, in terms of innovations as:

[4]There is no contradiction between the definition of $\widehat{z}_t = z_t - E_{t-1} z_t$, and $\widehat{PV}(z)_t = E_t PV(z)_t - E_{t-1} PV(z)_t$. The former can written as $\widehat{z}_t = E_t z_t - E_{t-1} z_t$, due to the fact that z_t is in the information set when the conditional expectation is taken.

$$\widehat{T_t} + \widehat{PV}\,(T)_t = \widehat{G_t} + \widehat{PV}\,(G)_t + (1+r)\,\widehat{B_t} + \frac{\Omega_t}{\rho}, \tag{2.16}$$

where $\Omega_t = E_t[PV(\epsilon)_t] - E_{t-1}[PV(\epsilon)_t]$ is the inherited measurement error term. If the series in (2.16) have a unit-root, it is convenient to rewrite it in first-difference format. Consider the identity[5] $(1-\rho)\,[z_t + PV(z)_t] = z_t + PV(\Delta z)_t$, and the fact that $\Delta\widehat{X_t} = \widehat{X_t}$, to obtain:

$$\Delta\widehat{T_t} + \widehat{PV}\,(\Delta T)_t = \Delta\widehat{G_t} + \widehat{PV}\,(\Delta G)_t + r\Delta\widehat{B_t} + r\Omega_t. \tag{2.17}$$

An equivalent result can be obtained if we start this process with (2.14):

$$\Delta\widehat{T_t} + \widehat{PV}\,(\Delta T)_t = \Delta\widehat{G_t^*} + \widehat{PV}\,(\Delta G^*)_t + r\Omega_t. \tag{2.18}$$

Equation (2.18) is critical for the discussion of intertemporal equilibrium, since it rewrites the intertemporal equilibrium condition in terms of innovations in taxes and expenditures ($\Delta\widehat{T_t}$ and $\Delta\widehat{G_t^*}$). Given an innovation in one of the fiscal variables, either $\Delta\widehat{T_t}$ or $\Delta\widehat{G_t^*}$, and disregarding measurement errors, there must be an innovation in the present value of either taxes, revenues, or both to restore equilibrium. Equation (2.18) also implies a restriction in terms of the responses from an initial innovation. Say, for instance, that there is a positive shock to expenditures. Then, the present values of taxes and expenditures will change in a way that their difference must equal the initial shock to the system, i.e., $\Delta\widehat{G_t^*} = \widehat{PV}\,(\Delta T)_t - \widehat{PV}\,(\Delta G^*)_t$ must hold. Notice that when we consider the measurement error term, this equation will not hold exactly. The absolute difference between the left- and the right-hand side is an increasing function of r. It should also be clear that the equation $\Delta\widehat{G_t^*} = \widehat{PV}\,(\Delta T)_t - \widehat{PV}\,(\Delta G^*)_t$ holds exactly when $r = 0$ (i.e., $\rho = 1$), since the error term vanishes in this case.

Calculating the marginal impact of current innovations on the innovations of the present values can be easily done using (2.12). Since $E_t X^*_{t+k} = A^{*k} X^*_t$, using the definition of present value, $PV(X^*)_t = \sum_{j\geq 1} \rho^j\, X^*_{t+j}$, and recalling that $\widehat{X_t^*} = E_t X^*_t - E_{t-1} X^*_t$, and $\widehat{PV}(X^*)_t = E_t PV(X^*)_t - E_{t-1} PV(X^*)_t$, we can then calculate $(E_t - E_{t-1})\,PV(X^*)_t$ and finally $\widehat{PV}(X^*)_t$:

$$E_t PV(X^*)_t = \sum_{k\geq 1} \rho^k E_t X^*_{t+k} = \sum_{k\geq 1} (\rho\, A^*)^k\, X^*_t. \tag{2.19}$$

[5]See Campbell(1987).

Using the *law of iterated expectations*, $E_{t-1}\left[E_t PV(X^*)_t\right] = \sum_{k\geq 1} (\rho\, A^*)^k\, E_{t-1} X_t^*$, and, $(E_t - E_{t-1})\, PV(X^*)_t = \sum_{k\geq 1} (\rho\, A^*)^k\, \mu_t^*$. Thus,

$$\widehat{PV}(X^*)_t = \sum_{k\geq 1} (\rho\, A^*)^k\, \mu_t^*. \tag{2.20}$$

Since $\mu_t^* = \left(\mu_t', 0, \cdots, 0\right)'$, equation (2.20) illustrates that we can write the innovation of the present value of any variable in the system as a linear combination of the VECM innovations μ_t. The loadings on these shocks are derived from the matrix A^*, whose coefficients come from the VECM estimation and ρ, the discount factor. Thus, the only parameter that cannot be estimated from the system (2.5) is ρ[6]. The partial derivative of $\widehat{PV}(X^*)_t$ with respect to μ_t^* is simply:

$$\frac{\partial \widehat{PV}(X^*)_t}{\partial \mu_t^*} = \sum_{k\geq 1} (\rho\, A^*)^k = \rho\, A^* \left(I - \rho\, A^*\right)^{-1}. \tag{2.21}$$

Equation (2.21) makes it easy to calculate the innovation present-value response function for either taxes or expenditures. All that is needed is a selection vector h_i, which selects the i-th element of a given vector. Recalling that $\Delta \widehat{X_t^{*i}} = \widehat{X_t^{*i}}$, where X^{*i} is the i-th variable in X^*, we obtain:

$$\frac{\partial\, \widehat{PV}(X^{*i})_t}{\partial \mu_t^*} = h_i \left\{\rho\, A^* \left(I - \rho\, A^*\right)^{-1}\right\}. \tag{2.22}$$

Equation (2.22) is what we have labelled the unconventional impulse-response function. Rather than calculating the response over time of a given variable in the system to shocks, it calculates the innovation present-value response to shocks to the system. It is tempting to associate the l-th element of the vector $h_i \left\{\rho\, A^* \left(I - \rho\, A^*\right)^{-1}\right\}$ with a unit impulse of the l-th element of μ_t. Although Bohn(1991) claims that the results from the use of the unconventional impulse-response function require no orthogonalization of the errors in the VECM, the interpretation above embeds the assumption that no other element of μ_t changed when its l-th element did. Of course, this can only be true if the covariance matrix of the innovations is diagonal, or if the elements of μ_t are orthogonalized.

[6]What we do below is to calculate several response functions varying ρ. We also discuss which values of ρ best fit the data.

The exogeneity tests conducted here follow the typology in Engle, Hendry and Richard(1983); see Maddala(1992) for an introduction and Ericsson and Irons(1994) for a complete review. There are three definitions of exogeneity: *weak*, *strong*, and *super exogeneity*. They are relevant respectively to conduct conditional inference, conditional forecasts and to obtain parameter-invariant models, i.e., models robust to the Lucas Critique.

To keep the discussion here at a basic level, we consider the bi-variate case. Suppose that we are interested in conducting inference on a given set of parameters β, and have two series z_t, and y_t. Then, z_t is said to be weakly exogenous for β if we can factor the joint density of $(z_t, y_t)'$, $f(z_t, y_t)$, into the conditional and marginal distributions $g(y_t|z_t)$and $h(z_t)$ respectively, where the parameters of the joint density are separable in such a way that, to learn β, one does not need to conduct inference on the parameters of the marginal density $h(z_t)$. Thus, it is valid to do conditional inference on β using z_t.

Strong exogeneity for β requires, in addition to weak exogeneity, that z_t is not *Granger-caused* by y_t. If that is not the case, it is impossible to do conditional forecasting for y_t, since one needs to learn y_t in order to learn the future of z_t. Super exogeneity requires, in addition to weak exogeneity, that the parameters β of the conditional density $g(y_t|z_t)$remain invariant to changes in the parameters of the marginal density $h(z_t)$. In this case, the Lucas Critique does not apply.

Since the implementation of Granger-causality tests is well known, we do not discuss it here at all. In the presence of unit roots and cointegration, weak exogeneity tests are simple to perform, since the parameters of the marginal distribution enter the conditional model through the error-correction term. Thus, testing for weak exogeneity is simply a test of exclusion of the error-correction term; see Johansen(1995, pp. 122-123). If it is not significant in a given equation, the regressand is weakly exogenous for the parameters of interest in the conditional model of the other variable in the system. However, if the error-correction term is significant weak exogeneity is rejected, and the parameters of the conditional and marginal distributions are not separable.

Testing for super exogeneity requires identifying factors that lead to changes in the parameters of the marginal model and then testing whether the parameters of the conditional model have changed due to these same factors; see Engle and Hendry(1993). This is usually done using dummy variables for potential regime shifts, although the size and power of the tests have to be taken into account.

3. The Database

Although the focus of this paper is public-debt sustainability, Brazil does not have long-span time-series data on debt. The data on expenditures, which include payments of *nominal* interest on public debt, and the data on revenues (excluding seignorage) were extracted from the national accounts calculated by FIBGE. They cover the period from 1947 to 1992. Since G_t^* includes real interest expenditures, not nominal, there is a mismatch between the data and the theoretical framework. Unfortunately, FIBGE does not provide disaggregate data on the nominal debt service, which made it impossible to match the data to the model.

Tax revenues from national accounts do not include seignorage, whose series were extracted from Cysne(1995), and then added to them. Both total tax revenues and expenditures, as described above, were then divided by GDP, and labelled T_t and G_t^* respectively. This is equivalent to using the same deflator for all series, which seems justified based on the broad spectrum of economic activity covered by tax collection and expenditures.

As pointed out by Ahmed and Rogers(1995), who had the same type of problem in their data set, the fact that we use nominal interest payments in place of rB_t may bias toward rejection the restricted version of the cointegration test between G_t^*, and T_t. Indeed, we may get a cointegrating vector different from $(1, -1)'$, possibly with the coefficient of G_t^* being greater than unity in absolute value. This, however, is not the only problem. Since the nominal interest rate can be thought of as the sum of the real interest rate and the *ex-ante* rate of inflation, we may not get cointegration at all if inflation is an integrated process. Needless to say, this is a real possibility for Brazil. In order to account for that problem, we increased the significance level of the cointegration test. Instead of the usual 5% or 10% levels (Johansen and Juselius(1990)), we used the 20% level.

4. Empirical Results

4.1. Unit-Root and Cointegration Tests

Before presenting the results of unit-root tests it should be mentioned that the fact that our series are ratios, therefore lying in the real interval $[0, 1]$, does not rule-out them being integrated processes; see Ahmed and Yoo(1989). Indeed, unit roots are about the persistency properties of the series, and not about the range

of random variables. The limited range is not really a problem, since one can think of data transformations (e.g. logit) that would map the initial ratio into the real line.

Figures 1 and 2 show the data set used in this paper. It is clear that all series are smooth, indicating a high level of persistence. The ratios G_t^* and T_t show a steady increase since the beginning of the sample, almost doubling in value. Despite this behavior, $G_t^* - T_t$ shows mean reversion, especially after the expenditure and tax increases of the mid-eighties.

Table 1 shows the results of the unit-root tests performed: the Augmented Dickey-Fuller (ADF) and the Phillips-Perron test. The results of the latter are especially relevant, since the data clearly show signs of heteroskedasticity. Regardless of the test type, the results support one unit root for expenditures and taxes, whereas the deficit $(G^* - T)$ is stationary even at 1%[7].

The next step is to perform cointegration tests. We used the likelihood-based cointegration test 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 is done in order to choose these. Lag length was selected using two types of information criteria (Schwarz and Hannan-Quinn). To choose the deterministic components we used the likelihood ratio test discussed in Johansen. The pre-test results are presented in Tables 2 and 3.

The results in Table 2a are based on a VAR estimated with an unrestricted constant term and the results in Table 2b on a VAR estimated without a constant. In both cases, using the Schwarz criterium, the optimal lag length is two. Using the Hannan-Quinn criterium, the optimal lag length is three. Given the Monte-Carlo results in Gonzalo(1995), showing that the efficiency loss is small for overfitting, while consistency is lost if the lag length chosen is too small, we chose to work with three lags on the VAR. It is also worth mentioning that the VAR with two lags showed problems in diagnostic tests.

Testing the deterministic components in the VAR is more subtle, since it requires conditioning on the number of cointegrating relationships to be performed; see Johansen(1991). Table 3 shows test results when the number of cointegrating vectors is one. When the model with an unrestricted constant is tested against the one with a restricted constant the test statistic is 1.64. Since it is distributed as a χ_1^2, we cannot reject the restricted model. The same result is obtained when we

[7]Tests for two unit roots reject this hypothesis for all three series.

tested the restricted constant against the no-constant model, or the unrestricted constant versus the no-constant model. Thus, we used a VAR without a constant.

The results of the cointegration test are presented in Table 4. As discussed above, we used the 20% significance level for the critical values of the Trace and the $\lambda_{\max}$ statistics. At 20% we reject the null of no cointegration vectors and cannot reject the null of one cointegrating vector. Although we used the 20% level, it is worth mentioning that the test statistics for the null of zero cointegrating vectors are very close to the critical value at 10%. The point estimate of the cointegrating vector is $(1, -0.94)$ (normalization on expenditures), which matches the prior that using the nominal-interest payments rather than the real-interest payments would bias upwards the absolute value of the expenditures coefficient. It is natural at this point to test the theoretical value of $(1, -1)$ for the cointegrating vector. At usual significance levels this theoretical vector could not be rejected. Thus, cointegration tests corroborate our prior findings from unit-root tests that the deficit is stationary, and that debt is sustainable.

4.2. Exogeneity Tests

Table 5 shows the results of the error-correction model estimates. First, it seems that our choice of 3 lags for the VAR was appropriate. For the system as a whole, ΔG^*_{t-2} is significant, while ΔT_{t-2} is marginally significant. Choosing only two lags in the VAR would probably have led to the estimation of a misspecified model with autocorrelated residuals.

Table 6 shows the results of the weak exogeneity tests under the existence of one cointegration vector. At usual significance levels we found that expenditures are weakly exogenous for the parameters of interest in the conditional model of revenues, but the converse is not true for revenues. The fact that the coefficient of Def_{t-1} on the ΔG^*_t equation is not significant led us to test the joint hypotheses that the cointegrating vector and the adjustment coefficient are respectively $(1, -1)$ and $(0, \alpha)$. The χ^2_2 statistic of this test is 1.78, which is not rejected at the 41% significance level. The results on Table 7 clearly indicate that expenditures Granger-cause taxes and vice-versa at the 5% significance level. At 1%, however, taxes do not precede expenditures.

4.3. Unconventional Impulse-Response Function

Before presenting the results of the unconventional impulse-response function we investigated whether or not the two innovations in the VECM are correlated. As discussed above, only in the case that they are uncorrelated do impulse-response results have the usual interpretation. The correlation coefficient between the two innovations is 0.244, with a *p-value* of 0.19, thus not significant at either 5 or 10%. A likelihood ratio test using the VECM's covariance matrix also confirms that the two shocks are orthogonal at usual significance levels. The log-likelihood ratio test statistic equals 2.67. Since we are imposing only one zero constraint (the covariance between innovations), this statistic has a χ_1^2 distribution, with critical values being 3.84 and 2.71 at 5% and 10% respectively. Thus, at these significance levels we do not reject the null of orthogonal innovations. Therefore, impulse-response results have the interpretation given above and no orthogonalization technique is used[8].

Unconventional impulse-response functions were calculated using $(1, -1)$ as the cointegrating vector and the adjustment factor in the form $(0, \alpha)$[9]. The VECM used two lags, which is equivalent to a VAR of order 3. Insignificant regressors were deleted based on their *t-tests*. Based on these restricted VECM estimates, and using (2.22), we calculated the marginal impact of innovations on G_t^* and T_t on $\widehat{PV}(\Delta G^*)_t$ and $\widehat{PV}(\Delta T)_t$. Since the results depend on the discount factor used, we chose to include on Table 8 the following four cases: $\rho = 0.97$, $\rho = 0.98$, $\rho = 0.99$, and $\rho = 1$. To be able to test hypotheses on the parameters being estimated, we performed a Monte Carlo experiment with 1,000 replications for each value of ρ. The sample size used for each replication is the same of the original data and pre-sample observations values were also drawn. Standard errors of point-estimates were then computed based on the Monte Carlo results and are included in Table 8.

Table 8 has two interesting results. First, it is clear that expenditures change very little when we consider innovations in either expenditures or revenues. For $\rho = 1$, and thus with no measurement error, 89% of expenditure-generated deficits are eliminated *via* an increase in taxes, versus only 11% of decrease in future expenditures. Notice that the former is not statistically different from 100% and

[8]Bohn(1991) claims that there is no need for orthogonalization of shocks in his technique. As we discussed above, there is a potential interpretation problem when innovations are not orthogonal.

[9]The value of α was found to be 0.25.

the latter not statistically different from zero. Second, impulse-response results are consistent with the existence of Ricardian-Equivalence consumers in Brazil (Barro(1974, 1989)). This is true since a change in taxes, not accompanied by a change in expenditures, will be offset 100% by a change in future taxes. Thus, Ricardian-Equivalence consumers will not change their optimal consumption allocation given a change in taxes today, since they expect this to be offset exclusively by a change in future taxes. This result is a direct consequence of the weak exogeneity of expenditures.

These two basic results are also obtained for $\rho \neq 1$. For example, for $\rho = 0.99$, point estimates of $\frac{\partial \widehat{PV}(\Delta T)_t}{\partial \widehat{G_t^*}}$ and $\frac{\partial \widehat{PV}(\Delta G^*)_t}{\partial \widehat{G_t^*}}$ are 93% and -11% respectively, and are statistically equal to 100% and zero respectively. Moreover, $\frac{\partial \widehat{PV}(\Delta T)_t}{\partial \widehat{G_t^*}}$ is still statistically equal to -100%. Thus, Ricardian Equivalence holds. For other values of $\rho \neq 1$, impulse-response estimates are not very different from the results above, although these parameters are not precisely estimated and hypothesis-testing is not very informative. This is a consequence of the fact that $\Delta \widehat{T_t} + \widehat{PV}\,(\Delta T)_t \neq \Delta \widehat{G_t^*} + \widehat{PV}\,(\Delta G^*)_t$, and that there is an increasing measurement error term.

It is always appropriate to ask what is the proper value of ρ to use for the sample period 1947-1992. Due to the lack of long-span data on interest rates on debt, we can only conjecture here. Given the high and ascending level of Brazilian inflation during this period, it is hard to imagine a positive value for the real interest rate paid on debt. An observed phenomenon under high inflation is the erosion of the real value of nominal contracts. Even if debt is indexed to lagged inflation, as was the case in Brazil and other high-inflation countries, real interest rates may still be reduced by an increase in the level of inflation. Therefore, we propose using $\rho \simeq 1$.

Our last empirical test checks whether or not seignorage was important for long-term fiscal balance. We perform a cointegration test with the same revenue series as before with seignorage revenues excluded. For the sake of comparison with previous results, we use the same significance level as before (20%). The results are presented in Table 9. We find no cointegration between the expenditure-GDP ratio and the modified tax-GDP ratio. Notice that the Trace and the $\lambda_{\max}$ statistic are much smaller than their critical values at 20%.

The evidence for Brazil suggests that debt is sustainable and that expenditures are exogenous. Thus, the country followed much closer a spend-and-tax than a tax-and-spend policy. Indeed, our impulse-response results suggests that, given

a shock to expenditures, the present value of taxes will fully accommodate this imbalance restoring long-run equilibrium. Given that seignorage is critical for the existence of equilibrium, it was probably through seignorage revenues that most expenditure increases were financed. In this case, Brazilian inflation is simply a consequence of the type of spend-and-tax policy followed in the post-war period.

It is interesting to compare our results to those of others who investigated public-finance issues in Brazil. Pastore(1995) tests debt sustainability using the techniques of Hamilton and Flavin(1986). From unit-root test results, he finds that the first difference of public debt is stationary. Since this series is equal to Def_t (see (2.3)), his results and ours are identical. Since both are achieved using different techniques, different series, and different samples, they complement each other in confirming the conventional wisdom about Brazilian public finance.

One of the contributions of this paper is to test for exogeneity of expenditures. Several authors have used the exogeneity of expenditures in theoretical models, some of which were later applied to the data, e.g. Bruno and Fischer(1990), the extension in Ruge-Murcia(1995) and the references therein. Since expenditures were found to be weakly exogenous under proper econometric tests, we validate this prior modelling. When we consider the fact that debt was only sustainable when seignorage was included as a government revenue, it becomes clear that the fundamental character of Brazilian public finance was that of endogenous seignorage used to accommodate expenditure increases. Indeed, this conclusion is a common feature of the work of Pastore(1995), Ruge-Murcia(1995), and ours.

5. Conclusions

This paper presents tests on the sustainability of Brazilian public debt for the post-war period - 1947-92. They show that debt is sustainable only if seignorage is included as a government revenue. In this case, exogeneity tests suggest that expenditures are exogenous. Unconventional impulse-response results show that, regardless of how an initial fiscal imbalance is originated (shocks to expenditures or revenues), it is eliminated through a change in taxes. This last result is consistent with the existence of Ricardian-Equivalence consumers in Brazil. Based on these evidence, we find that Brazil has followed closely a spend-and-tax policy in the post-war era, with seignorage having a crucial role in balancing the budget. As a consequence, Brazilian inflation was consistently high during this period.

It is interesting to ask under which circumstances Brazil can achieve fiscal

balance under low inflation. Buchanan and Wagner(1977) point out that institutional issues related to the management of the government budget are critical. An alternative would be to increase Central Bank independence, so that it would not be at all responsible for accommodating fiscal expansions.

A brief reflection on the current status of the Real Plan is appropriate. Since the beginning of the plan in July 1994, seignorage revenues have decreased sharply. Expenditures have increased considerably, generating a persistent deficit. These and other factors have led to an almost three-fold increase in the stock of public debt since then. If expenditures are still exogenous, as verified for the sample 1947-92, there are only two polar forms to restore long-run fiscal balance: (i) increase taxes, excluding seignorage, or (ii) increase seignorage revenues. In the first case, Brazilians will be the most heavily taxed citizens in Latin America. The problem is that there are very few services as a counterpart of these taxes - education, infra-structure, legal system, etc. In the second case, inflation will increase again, a price Brazilians may not be willing to pay for fiscal balance. Let us just hope that expenditures have ceased to be exogenous.

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APPENDIX

Figure 1

Expenditures (Including Nominal Interest Paid on Debt) and Revenues (Including *Seignorage*) as a Proportion of GDP

$G_t^* = G_t + rB_t$ and T_t

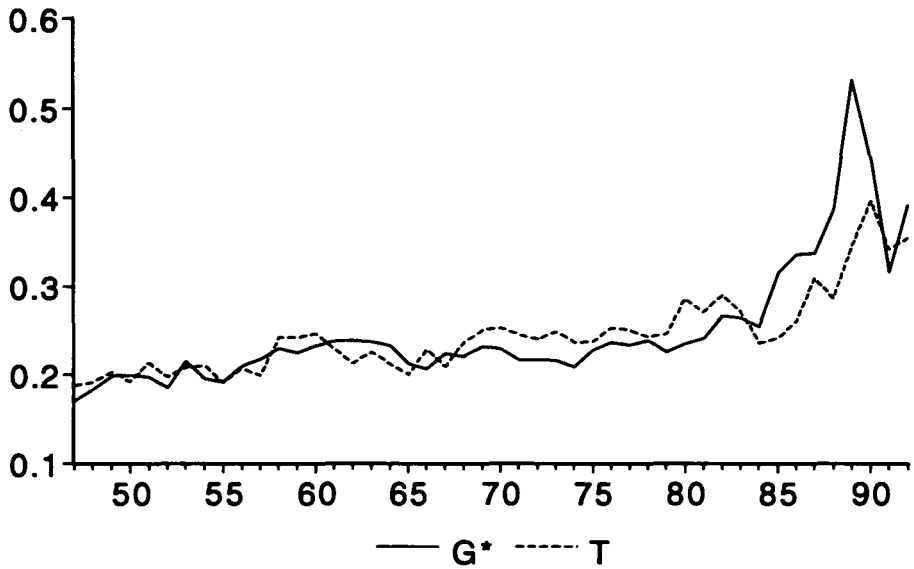

Figure 2

Public Deficit as a Proportion of GDP

$G_t^* - T_t$

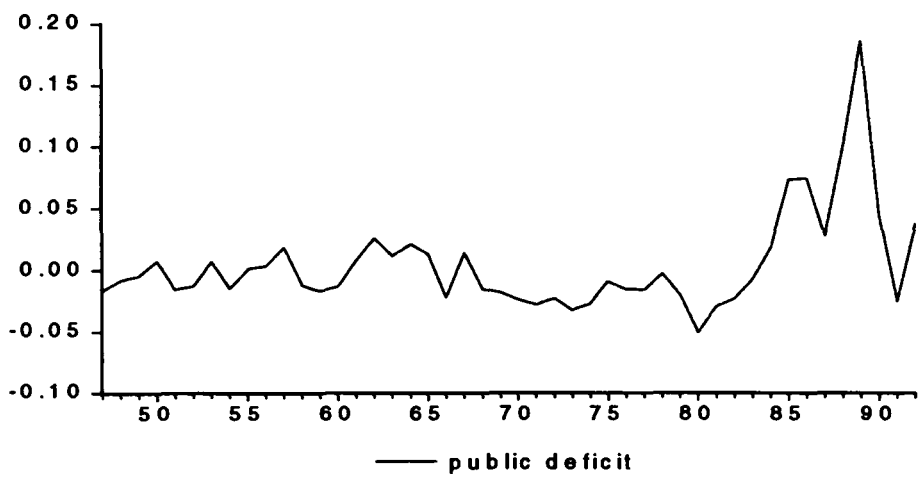

Table 1

Unit-Root Tests

Series	Lags[1]	ADF Test	Phillips-Perron Test[2]
Tax Revenues (T)	3	-3.03	-3.37
Expenditures (G^*)	4	-2.14	-2.54
Deficit ($G^* - T$)	4	-2.98**	-3.27**

Notes: (1) The number of lags applies only to the ADF test. The final choice was made based on the t-test of significance of the last lagged first-difference. (2) The lag truncation chosen for the Bartlett kernel was 3. For (G^*) and (T) a constant and a time trend were used. Critical values are -3.53 and -4.20 for the 5 and 1% significance levels respectively. For ($G^* - T$) the no-constant specification was used. Critical values are -1.95 e -2.62 for the 5 and 1% significance levels respectively. The symbols (*) and (**) represent rejection of the null of a unit root at the 5 and 1% significance levels respectively.

Table 2

VAR Lag Truncation

(a)

VAR Order	Constant	Linear Trend	Schwarz Criterium	Hannan-Quinn Criterium
1	unrestricted	no trend	-14.21	-14.36
2	unrestricted	no trend	-14.25	-14.52
3	unrestricted	no trend	-14.24	-14.61
4	unrestricted	no trend	-13.98	-14.45

(b)

VAR Order	Constant	Linear Trend	Schwarz Criterium	Hannan-Quinn Criterium
1	no constant	no trend	-14.21	-14.31
2	no constant	no trend	-14.39	-14.60
3	no constant	no trend	-14.36	-14.68
4	no constant	no trend	-14.12	-14.54

Table 3

Deterministic Components Test (VAR(3))

Model	VAR Order	Constant	Linear Trend	Restricted Vs. Unrestricted Model	Test Statistic
1	3	unrestricted	no trend	2 versus 1	1.64
2	3	restricted	no trend	3 versus 2	2.02
3	3	no constant	no trend	3 versus 1	3.66

Notes: Results conditioned on the existence of one cointegrating vector. The critical value for the $\chi^2(1)$ statistic is 3.84 and 6.63 for the 5 and 1% significance levels respectively.

Table 4

Johansen's Cointegration Test (*Seignorage* Included as Revenues)

Test Statistic (Critical Value at the 20% Level)				Cointegrating Vector
λ_{max}		Trace		(Expenditures, Taxes)
$K = 0$	$K \leq 1$	$K = 0$	$K \leq 1$	
9.185*	1.638	10.82*	1.638	(1.0, -0.94)
(7.58)	(1.82)	(8.45)	(1.82)	

Cointegration Restriction Test

Restriction $(1, -1)\chi^2(1) = 1.50$; *p-value*: 22.14%

Notes: The symbol (*) indicates rejection of the null at the 20% significance level.

Table 5

Vector Error Correction Model Estimates

	EQ. 1 (ΔG_t^*)			EQ. 2 (ΔT_t)		
Regressor	Est. Coeff.	*t*-stat	*t*-prob	Est. Coeff.	*t*-stat	*t*-prob
ΔG_{t-1}^*	0.17964 (0.18197)	0.987	0.3298	0.23177 (0.1131)	2.049	0.0474
ΔG_{t-2}^*	-0.73603 (0.20688)	-3.558	0.0010	-0.1790 (0.12861)	-1.392	0.1720
ΔT_{t-1}	-0.30270 (0.26027)	-1.163	0.2521	-0.3066 (0.1618)	-1.895	0.0657
ΔT_{t-2}	0.61572 (0.25036)	2.459	0.0186	0.15161 (0.15565)	0.974	0.3362
Def_{t-1}	0.07624 (0.1516)	0.503	0.6179	0.25661 (0.09425)	2.723	0.0097

Notes: Standard errors in parentheses.

Table 6

Adjustment-Coefficient Weak Exogeneity Test

Null Hypothesis	Test Statistic	P-Value
T_t is weakly exogenous for the parameter of interest of the G_t^* conditional model	7.36	0.0067**
G_t^* is weakly exogenous for the parameter of interest of the T_t conditional model	1.11	0.29

Note: The symbol (**) indicates rejection of the null at the 1% significance level.

Table 7

Granger-Causality Test

VAR(3)	
Null Hypothesis	P-Value
T_t does not Granger-cause G_t^*	0.023*
G_t^* does not Granger-cause T_t	0.0000**

Notes: The symbols (*) and (**) represent rejection of the null at the 5 and 1% significance levels respectively.

Table 8

Unconventional Impulse-Response Function

		(ρ=0.97)		
Innovation on:	T	T	G*	G*
Response of:	$P\hat{V}(\Delta T)$	$P\hat{V}(\Delta G^*)$	$P\hat{V}(\Delta T)$	$P\hat{V}(\Delta G^*)$
Point-Estimate	-1.22	-0.07	1.12	0.05
Standard Error	(2.13)	(0.47)	(2.13)	(0.51)
		(ρ=0.98)		
Innovation on:	T	T	G*	G*
Response of:	$P\hat{V}(\Delta T)$	$P\hat{V}(\Delta G^*)$	$P\hat{V}(\Delta T)$	$P\hat{V}(\Delta G^*)$
Point-Estimate	-1.13	-0.04	1.01	-0.09
Standard Error	(1.05)	(0.39)	(0.99)	(0.42)
		(ρ=0.99)		
Innovation on:	T	T	G*	G*
Response of:	$P\hat{V}(\Delta T)$	$P\hat{V}(\Delta G^*)$	$P\hat{V}(\Delta T)$	$P\hat{V}(\Delta G^*)$
Point-Estimate	-1.06	-0.02	0.93	-0.11
Standard Error	(0.09)	(0.03)	(0.20)	(0.18)
		(ρ=1.00)		
Innovation on:	T	T	G*	G*
Response of:	$P\hat{V}(\Delta T)$	$P\hat{V}(\Delta G^*)$	$P\hat{V}(\Delta T)$	$P\hat{V}(\Delta G^*)$
Point-Estimate	-1.00	0	0.89	-0.11
Standard Error	(-)	(-)	(0.15)	(0.15)

Notes: Standard errors calculated using a Monte Carlo, where the Data Generating Process (DGP) was assumed to be the Restricted (cointegration) Vector Error Correction Model. The estimated model in the Monte Carlo corresponds to the DGP. Pre-sample data were also drawn, and not assumed fixed. The number of replications was set to 1,000. When ρ=1, the responses to innovations to T_t are not stochastic.

Table 9

Johansen's Cointegration Test (*Seignorage* Excluded as Revenues)

Test Statistic (Critical Value at the 20% Level)				Cointegrating Vector
λ_{max}		Trace		
K = 0	K ≤ 1	K = 0	K ≤ 1	
7.305	0.0454	7.351	0.0454	-
(10.04)	(1.66)	(11.07)	(1.66)	

Notes: The symbol (*) indicates rejection of the null at the 20% significance level. The VAR is modeled with an unrestricted constant and two lags, according to the Schwarz Criterium and to likelihood ratio tests of the deterministic components respectively.

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