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Parametric Bootstrap for Unit Root Testing - Brazilian Evidence

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ABSTRACT

This paper develops a procedure for bootstrapping Dickey-Fuller and Ouliaris, Park, and Phillips unit root tests. The procedure allows for deterministic trend polynomials in the maintained hypothesis, seasonal dummies in the test equation, and error terms following ARMA processes. The approach may be used to verify the unit root nonstationarity or the stationarity around a deterministic polynomial trend indicated by these tests. It may also work as a tool to verify the size and power distortion presented by the sample tests. We illustrate the use of the procedure by applying it to Brazilian macroeconomic time series. In most cases we obtained different results from the sample tests. The results confirmed that both tests have material size or power distortions that can be corrected by the procedure. Our evidence show the bootstrap approach perform better than the traditional unit root tests concerning the size accuracy and the higher power in hypothesis testing unit root against the alternative of stationarity. Using this approach, we conclude some of the series can be modeled as stationary processes around a drift or a polynomial trend, in contrast to previous findings.

KEYWORDS: unit root tests, bootstrap, deterministic terms, serial correlation correction, size and power distortion.
JEL: C12, C15, C22, E00.

RESUMO

O artigo desenvolve um procedimento de Monte Carlo, conhecido como “bootstrap”, para ser aplicado aos testes de raízes unitárias de Dickey-Fuller e Ouliaris, Park, e Phillips. O procedimento permite a introdução de termos polinomiais de tendência na hipótese nula, dummies para tratamento de sazonalidade na equação de teste, e permite especificar o termo de erro como processos ARMA. O método pode ser utilizado para verificar a validade dos resultados apontados pelos testes acima de não estacionariedade ou de estacionariedade de uma série, em torno de um polinômio de tendências. Ele pode também ser empregado para verificar as distorções de tamanho e potência presentes nos testes amostrais. A operacionalidade do procedimento é mostrada com a sua aplicação à séries macroeconômicas brasileiras. Em muitas aplicações os resultados obtidos divergem daqueles obtidos dos resultados amostrais. Os resultados confirmam que os dois testes possuem importantes distorções de tamanho ou potência, que podem ser corrigidas pelo procedimento proposto. Nossas evidências mostram que o procedimento de bootstrap tem melhor desempenho que os testes tradicionais de raiz unitária, no que concerne a precisão do tamanho do teste, e a maior potência no teste da presença de uma raiz unitária contra a hipótese alternativa de estacionariedade. Com este método, concluímos que algumas das séries estudadas podem ser modeladas como processos estacionários em torno de um “drift” ou de um polinômio de tendências, o que contrasta com resultados obtidos por outros autores.

PALAVRAS CHAVES: testes de raízes unitárias, bootstrap, termos determinísticos, correção de correlação serial, distorção de tamanho e potência.

1 Introduction

The purpose of this paper is to examine once more the stationarity and the underlying data generation process of some macroeconomics series, in the periods 1966 to 1985 and 1966 to 1990. The analysis will be performed by way of the Augmented Dickey-Fuller test (ADF) and the Ouliaris, Park and Phillips (1989) unit root tests (OPP), complementarily with the Philips-Perron (PP) test. Considering the Brazilian distinct policy regimes can induce wrongly unit root detection, we combine these tests with a Monte Carlo procedure, the bootstrap approach, which is potentially more powerful[1] in rejecting a false hypothesis than the standard Dickey-Fuller tests and its extensions, such as the OPP; see Li and Maddala (1996). *Hence our final aim is to propose a new methodology for unit root testing.*

In the tests of Dickey and Fuller and Phillips and Perron, the maintained hypothesis is that the time series is integrated with drift but with no trend. On the other hand, the Ouliaris, Park and Phillips test extends these tests to explicitly allow for a deterministic polynomial time trend in the maintained hypothesis[2]. Since an important feature of this procedure is its invariance to the presence of drift and polynomial trend in the true data generation process, it should be helpful in discriminating unit root nonstationarity and processes that are stationary around a deterministic polynomial trend; see OPP (1989).

The bootstrap technique is a well-known tool since it often provides a way of improving on the approximations of asymptotic theory. In the stable first order autoregressions the bootstrap can improve estimation of p-values, power of tests, etc; see Vinod (1993). Rayner (1990) advises against using the traditional *t* tests in favor of the bootstrap. However, the results about the superiority of the t-bootstrap over the Dickey-Fuller test are not yet conclusive. In general both have the same power, although for some non-normal distributions, its performance was slightly better than that of the Dickey-Fuller test; see Li and Maddala (1996). Nevertheless, it provides a simple and accurate *test specially designed for use in the context of particularly empirical work*.

The basic idea of bootstrapping is to use the single available data set to design a sort of Monte Carlo experiment in which the data itself is used to approximate the distribution of some random quantity[3]. This idea is implemented by performing a Monte Carlo experiment in which the error terms are drawn from an assumed distribution, such as the normal, with variance taken from the sample residuals[4]; see Davidson and MacKinnon (DM), 1993, p. 763-766. Then, with a given initial value of the dependent variable and/or with the other regressors' observations, the dependent variable series is simulated. By repeating this procedure the bootstrap can approximate, for instance, the unknown true distribution of the estimator with the empirical 'bootstrap' distribution. It is in such cases that the bootstrap can be particularly useful.

In this paper, we propose the application of the bootstrap technique to the ADF and OPP unit root tests in order to obtain approximated critical values to the test statistics. The procedure allows for the presence of serial correlation in the error term and a deterministic polynomial time trend in the maintained hypothesis. Because there are repeated conflicts between the ADF and the OPP sample results, we will apply the technique to both tests and

[1] Considering unit root as null and the stationarity as the alternative.
[2] However, it does not imply that the bulk of shocks that took place in the Brazilian economy may be mimicked by a deterministic polynomial trend with a higher order than to one. Then, the OPP test would not be the best alternative either. The experiments performed with unit root tests that permit one structural break were not successful, since no choice of break points led to the rejection of the unit root hypothesis. Furthermore, the presence of many different breakpoints made the research for a test specification too hard to obtain.
[3] The approach is based on the idea that the sample we have is a good representation of the underlying population, which is all right if we have a large enough sample. Li and Maddala (1996) states that for bootstrapping models with AR(1) errors and trending regressors a large sample n>100 is required for the method to work well.
[4] This is the often called parametric bootstrap.

observe if in both cases the nature respects the null hypothesis. As we will show, the basic bootstrap structure is the same for both tests.

Following this strategy, we intend to clarify the existing contradictions between the test results. This is particularly important when one of them is near to reject H_0 and the other is not. *In this context, we suggest that the bootstrap approach have the function of rendering conclusive results. Therefore, we will take as a decision rule to reject H_0 if the bootstrap test provides this result.*

In this paper, we show with the bootstrap approach that some Brazilian series like the real output, the real balances, the money velocity and "black" dollar premium are stationary between 1966 and 1990. These results contrast with previous findings[5] that pointed out these series as being I(1). Moreover, since our Monte Carlo tests are similar to Monte Carlo experiments, in which artificial data generation processes are replaced by observed data (the Brazilian series), the results may be used to assess the size and power properties of the ADF and OPP statistics, when applied to Brazilian data. We show that the OPP statistics present bigger size distortion compared to the ADF test and both tests have important power distortion.

The organization of this paper is as follows. Section 2 presents the Monte Carlo test we apply in the paper. In the first subsection, we describe the bootstrap's outlines, its application to unit root hypothesis testing, and the bootstrap approach to the ADF test. In subsection 2.2, we show how to incorporate into the bootstrap unit root model (i) a polynomial time trend in the maintained hypothesis and (ii) the Dickey-Fuller standard parametric correction to the presence of serial correlation in the error terms. Subsection 2.3 shows the bootstrap procedure we will apply to unit root testing with trend terms and underlying serial correlation in the error term. In section 3 our procedure is then applied to macroeconomic time series to test whether the original data stands up to the null hypothesis of a unit root. We also provide an evaluation of the unit root tests performance and some economic implications of our results. In section 4 we summarize the main results and discuss the bootstrap performance, as well the results concerning the size and power distortion in the bootstrapped tests. In section 5, we present the conclusion. In the appendices A, B and C, respectively, we describe the methodology applied for performing the data unit root tests, we depict the data and the recursive estimates of the lagged dependent variable corresponding to the ADF test equation, and we display the first fourth moments of all simulated bootstrap statistics.

2. The Monte Carlo Procedure

2.1 General Guidelines

The bootstrap method introduced by Efron (1979) is a resampling method used to obtain critical values. In other words, the bootstrap can approximate the unknown true distribution of the estimator with the empirical 'bootstrap' distribution. Let $(y_1, y_2,\ldots,y_n)$ be a random sample from a distribution characterized by a parameter θ. Inference about θ will be based on a statistic T. The basic bootstrap approach consists of drawing repeated samples (with replacement) of size m (which may or may not be equal to n, although it usually is) from $(y_1, y_2,\ldots,y_n)$. Call this sample $(y_1^*, y_2^*,..., y_n^*)$. This is the bootstrap sample. We repeat step N times. N is the number of bootstrap replications. For each bootstrap sample, we compute the statistic T. Call this T*. The distribution of T* is known as the bootstrap distribution of T. This bootstrap distribution is used to make inference about θ. Under some circumstances, the bootstrap

[5] Like Pastore (1995), Nakane (1993), Novaes (1991), Rossi (1988), Giambiaggi and Vals Pereira (1989) and Vals Pereira (1988). By using the ADF or the Phillips and Perron (PP) test, they identified a unit root in the same series we studied.

distribution enables one to make more accurate inferences than the asymptotic distribution of T[6].

The bootstrap method described above is the simplest one and it is valid only for IID observations, otherwise the bootstrap resampling does not provide the correct approximation of the true distribution, and so the method needs to be modified. Furthermore, in the cases of bootstrapping unit root autoregressive processes the asymptotic theory does not hold. Thus, when ε_t are AR processes (the ADF test) one needs both a parametric specification for the errors and some extensions of the bootstrap method, as described later; see Jeong and Maddala (1993) and Li and Maddala (1996). Finally, since we are interested in applying this method in hypothesis testing, as pointed out in the literature, it is important to apply significance tests using (asymptotically) pivotal statistics. Otherwise, one cannot expect much of an improvement over the asymptotic results.

Thus, consider the simple case of an AR(1) unit root model:

$$y_t = \beta y_{t-1} + \varepsilon_t, \quad \beta = 1, \quad t=1,2,\ldots,n \tag{1}$$

where $y_0 = 0$, $\varepsilon_t \sim iid(0,\sigma^2)$. In this case the OLS estimator is a function of the standard Wiener process W(r) and has a nonnormal limiting distribution

$$\left(\sum_1^n y_{t-1}^2\right)^{1/2}(\hat{\beta}-1) \xrightarrow{d} \frac{\sigma}{2}\left((W(1))^2-1\right)\left(\int_0^1 (W(r))^2 dr\right)^{-1/2}. \tag{2}$$

Thus, conventional tests based on normal asymptotic theories are not valid. The associated tests in this case are the Dickey-Fuller coefficient test $\hat{\rho} = n(\hat{\beta}-1)$ and t-test, $\hat{t} = (\hat{\beta}-1)/SE(\hat{\beta})$ which have the following limiting distributions

$$\hat{\rho} = n(\hat{\beta}-1) \Rightarrow \frac{1}{2}\left((W(1))^2-1\right)\left(\int_0^1 (W(r))^2 dr\right)^{-1} \tag{3}$$

$$\hat{t} \Rightarrow \frac{1}{2}\left((W(1))^2-1\right)\left(\int_0^1 (W(r))^2 dr\right)^{-1/2}. \tag{4}$$

Now consider the following modification of model (1) with trend terms t^k

$$y_t = \Sigma_0^p \beta_k t^k + \alpha y_{t-1} + \varepsilon_t, \quad \alpha = 1, \tag{5}$$

where y_0 is a random variable with a distribution that is independent of n, the sample size, and ε_t is a weakly stationary, zero mean innovation sequence. The hypothesis we want to test is α=1, and β_p=0. The limiting distributions of $\hat{\rho}$ and t_p are nonstandard. They depend on nuisance parameters, which require transformations of the statistics that eliminate the nuisance parameters asymptotically. The asymptotic distributions of these statistics, after the pertinent transformations are presented below[7]

$$\hat{\rho} = n(\hat{\alpha}-1) \xrightarrow{d} \left(\int_0^1 W_p(r)dW(r)\right)\left(\int_0^1 (W_p(r))^2 dr\right)^{-1} \tag{6}$$

[6] When, for instance, the asymptotic theory is tractable but not very accurate in samples of the size used in applications. In such cases the bootstrap often provides a way of improving on the approximations of asymptotic theory

[7] Following OPP (1989) $W_p(r)$ is defined to be the stochastic process on [0,1] such that $W_p(r)$ is the projection residual of a Brownian motion W(r) on the subspace generated by the polynomial functions 1,r,…, r^k in $L^2[0,1]$. Here, denotes the Hilbert space of square integrable functions on [0,1] with the inner product (f,g)= $\int_0^1 fg$ for f,g$\in L^2[0,1]$.

$$t(\hat{\alpha}) \xrightarrow{d} \left(\int_0^1 W_p(r) dW(r) \right) \left(\int_0^1 (W_p(r))^2 dr \right)^{-1/2}. \tag{7}$$

Under the null hypothesis of a unit root, the left-tailed critical values of the asymptotic distribution of both statistics increase in absolute value with the number of included deterministic regressors. This is due to the fact that the p^{th} trend term increases more rapidly with t^{p-1} than either the constant or the linear trend; see DM (1993).

To construct the bootstrap test corresponding to (2), we start with the OLS estimation of (1), compute $\hat{\varepsilon}_t$, and its standard error ($\hat{\sigma}$). We then get the bootstrap sample $\hat{\varepsilon}_t^*$ by drawing pseudo-random numbers from a normal distribution $N(0, \hat{\sigma})$. We now generate y_t^* using $\hat{\varepsilon}_t^*$ and $\hat{\beta}$ using the coefficients of the estimated equation. In this case, the bootstrap estimates of $\hat{\rho}$ and $\hat{t}$ must converge in distribution to the same expressions as (3) and (4); see Ferreti and Romo (1994), and Li and Maddala (1996).

One final point concerns the use of pivotal statistics. There are two Dickey-Fuller tests: the coefficient test (3) and the t-test (4). When it comes to the bootstrap approach there is again the question of whether to consider the coefficient test or the t-test. Li and Maddala (1996) looked into this issue and found the t-test only marginally better. The case for considering pivotal statistics may not be as strong for the unit root model (as in the stationary models).

2.2 Unit Root Tests with Deterministic Terms

As suggested in DM (1993), we can get a combined model that nests the underlined models of the ADF and OPP unit root tests. The model will enable us to use only one general basic structure to construct all the bootstrap tests. Moreover, it will permit to follow the same procedure described in the last section concerning the Dickey-Fuller tests.

The bootstrap we suggest in this paper assumes that the error of the underlying data-generation process follows an ARMA process. So we will use the Dickey-Fuller augmented version, which employs a parametric correction to the simple model by adding sufficient terms in Δy_{t-i}, to whiten the residuals. It is a well-known fact that this correction does not change the asymptotic Dickey-Fuller distribution; see Hamilton (1994) and DM (1993). For the same reason, the Ouliaris, Park and Phillips test will also be extended for correcting serial correlation. This procedure also produces an asymptotically valid test in the presence of serial correlation; see DM (1993).

We can then determine a generalized model that nests the trend-stationary model with a maintained polynomial trend of order p, and the random walk with drift added by a linear time trend (thus allowing the drift term to change over time). One plausible way that includes both models as special cases is presented below.

Suppose the DGP of a time series y_t can be written as

$$y_t = \gamma_0 + \gamma_1 t + \gamma_2 t^2 + \gamma_3 t^3 + \ldots + \gamma_p t^p + v_t; \quad v_t = \alpha v_{t-1} + \varepsilon_t, \tag{8}$$

where ε_t follows a stationary ARMA process. We then get

$$y_t = \gamma_0 + \gamma_1 t + \gamma_2 t^2 + \gamma_3 t^3 + .. + \gamma_p t^p + \alpha(y_{t-1} - \gamma_0 - \gamma_1(t-1) - \gamma_2(t-1)^2 - .. - \gamma_p(t-1)^p) + \varepsilon_t.$$

Developing this expression, we obtain

$$y_t = (1-\alpha)\gamma_0 + \alpha(\gamma_1 - \gamma_2 + \gamma_3 - \gamma_4 + \ldots + \gamma_p) + [(1-\alpha)\gamma_1 + \alpha(2\gamma_2 - 3\gamma_3 + 4\gamma_4 - \ldots - k_1\gamma_p)]t + [(1-\alpha)\gamma_2 + \alpha(3\gamma_3 - 6\gamma_4 + 10\gamma_5 - \ldots + {}_2\gamma_p)]t^2 + [(1-\alpha)\gamma_3 + \alpha(4\gamma_4 - 10\gamma_5 + \ldots - k_3\gamma_p)]t^3 + [(1-\alpha)\gamma_5 + \alpha(5\gamma_5 - \ldots + k_4\gamma_p)]t^4 + [(1-\alpha)\gamma_5 + \alpha(\ldots - k_5\gamma_p)]t^5 + \ldots + (1-\alpha)\gamma_p t^p + \alpha y_{t-1} + \varepsilon_t, \tag{9}$$

where the set $(k_1, k_2, \ldots, k_p)$ is in N. For an even p, it has a positive signal and a negative if p is odd. This set is taken from the p^{th} row of Pascal's triangle. Because (9) is nonlinear in the parameters, it is convenient to reparametrize it as

$$y_t = \beta_0 + \beta_1 t + \beta_2 t^2 + \beta_3 t^3 + \ldots + \beta_p t^p + \alpha y_{t-1} + \varepsilon_t. \tag{10}$$

The above equation may be extended to include other nonstochastic regressors, such as dummy variables. Because the seasonal dummies are of the same order as the constant term, which is already included, their inclusion does not change the asymptotic distributions of the test statistics. We shall consider in this paper centered seasonal dummies which are constructed so that they add up to zero for each t, i.e., $\sum_{i=1}^{s}\tilde{D}_{it}=0$.

Since our interest is to assume that the error term follows an ARMA process we need unit root tests that are (asymptotically) valid in the presence of serial correlation. The procedure is to add to equation (10) a polynomial of lagged first differences in y_t with length sufficiently large to whiten the residuals[8]. In the choice of the truncation lag parameter one may employ the sequential test for the significance of the last lagged first difference ("kmax" procedure), suggested by Campbell and Perron (1991). This procedure may be combined with the information based model selection rules (such as the Akaike and the Schwartz criteria) described in Ng and Perron (1993). Therefore, the general combined model has the following expression

$$y_t = \sum_{o}^{p}\beta_j t^j + \sum_{1}^{s}\delta_r D_r + \sum_{1}^{k}\theta_i \Delta y_{t-i} + (\alpha-1)\Delta y_{t-1} + \varepsilon_t\,. \tag{11}$$

This model stands as the general specification for performing the ADF and the OPP unit root tests. Under the null hypothesis $H_0{:}\alpha=1$ we obtain the nonreparametrized version:

$$y_t=(\gamma_1 - \gamma_2 + \gamma_3 - \gamma_4 + \ldots + \gamma_p)+ (2\gamma_2 - 3\gamma_3 + 4\gamma_4 - \ldots - k_1\gamma_p)t+(3\gamma_3\text{-}6\gamma_4+10\gamma_5 -\ldots+ k_2\gamma_p)t^2+(4\gamma_4\text{-}10\gamma_5+\ldots\text{-}\ k_3\gamma_p)t^3+(5\gamma_5 -\ldots+ k_4\gamma_p)t^4+(6\gamma_6 -\ldots\text{-}\ k_5\gamma_p]t^5+\ldots+k_{p-1}\gamma_p t^{p-1}+ \sum_{1}^{s}\delta_l D_l + \sum_{1}^{k}\theta_i\Delta y_{t-i} +y_{t-1} +\varepsilon_t. \tag{12}$$

Thus, if we are testing if y_t is a random walk with a polynomial trend of fourth order in the maintained process, under $H_0{:}\alpha=1$ equation (12) takes the expression

$$y_t=(\gamma_1\text{-}\gamma_2+\gamma_3 - \gamma_4 +\gamma_5) +(2\gamma_2 - 3\gamma_3 + 4\gamma_4\text{-}5\gamma_5)t + (3\gamma_3 \text{-}6\gamma_4+10\gamma_5)t^2+(4\gamma_4 - 10\gamma_5)t^3+5\gamma_5 t^4 + \sum_{1}^{s}\delta_l D_l + \sum_{1}^{k}\theta_i\Delta y_{t-i} +y_{t-1} + \varepsilon_t. \tag{13}$$

If y_t is just a random with drift, under the null (12) becomes

$$y_t = \gamma_1 + \sum_{1}^{s}\delta_l D_l + \sum_{1}^{k}\theta_i\Delta y_{t-i} + y_{t-1} + \varepsilon_t. \tag{14}$$

It is the model (12), or its special cases like equation (13) and (14), that we will apply to simulate the bootstrap sample $(y_1^*, y_2^*,\ldots,y_n^*)$. This procedure of simulation follows Hamilton (1994), who states that the goal of unit root tests is to find a parsimonious representation that gives a reasonable approximation of the true data process. So if $H_0{:}\alpha=1$ is accepted for a given unit root test specification, we can conclude that the series has an ARIMA(k+1,1,0) representation. Otherwise if H_0 is rejected the true process is approximated by a stationary ARMA(k+1,0) process. It is direct that each representation may include the test specified seasonal dummies and the pertinent deterministic terms.

2.3 Performing the Bootstrap Approach

In this section, we shall implement the parametric specification for the error terms described in section 2.2. Concerning the question of which test statistics to choose, we consider it is more prudent to perform the experiment for both of them. Thus, we apply the bootstrap approach to the coefficient test $\hat{\rho}$ [9] and the t-test statistics.

We are concerned with two periods of the Brazilian economic history: 1966 to 1985 and 1966 to 1990. For a quarterly series, in the second period there are 93 observations (we only consider the first quarter of 1990), implying the bootstrap method may present size distortions. To correct this inconvenience we will extend the number of replications up to 5000. So even

[8] If the series has moving average components the procedure remains valid, provided one lets the number of lagged difference terms tend to infinity at a rate no faster than $n^{1/3}$.

[9] Since we are correcting for serial correlation, the test statistics must be redefined to $\hat{\rho}=n(\hat{\beta}-1)\big/(1-\sum_{1}^{k}\Delta y_{t-i})$.

having a number of extractions smaller than 100, we will draw a sufficiently large number of samples. For the monthly series, we will do no more than 1000 iterations.

As the bootstrap sample initial value we take the first sample observation, which however may be an outlier. Then, to solve this problem we will not consider the first 100 simulated values and will take the 101st simulation as the effective start value.

In what follows we describe the procedure we perform:

(i) the first stage is to specify the ADF and the OPP unit root tests, and estimate the respective regression coefficients, residuals and standard error ($\hat{\sigma}$); then we follow for each test the steps below;

(ii) with the SE residuals, $\hat{\sigma}$, we draw N samples of size n to the error terms ε_t^* , assuming that they have a normal distribution $N(0, \hat{\sigma}^2)$ [10]; in order to be able to reproduce the same sample we assigned a particular random seed;

(iii) as the series initial value y_0, we will always pick up the first data sample value (that is the value originated by the nature); with this initial value, we simulate the bootstrap sample $(y_1^*, y_2^*, ..., y_n^*)$, under H_0:α=1, by using the estimated coefficients of model (12), and the drawn error terms ε_t^* ; we leave out the first 100 simulations; then we repeat this step N times;

(iv) using the specification achieved in step (i) we run N regressions and compute from each one the bootstrap statistics ρ^* and t^*;

(v) we compute the bootstrap distribution to the ρ^* and t^* statistics by sorting them in descending order; thus we identify the critical value corresponding to the data observed statistics (i.e., given by the nature) and calculate the p-value bootstrap[11] (or p-value Monte Carlo);

(vi) we follow the decision rule: reject H_0 if the p-value Monte Carlo is smaller than a specified nominal size test (that will be 5%).

3 Empirical Applications

Our sample covers the period from 1966 to 1990. As we mentioned in the introduction, there were significant changes in the conduction of the political economy, especially after 1986. The trials of stabilization inflation by means of demand control gave place to supply shocks via application of price freeze and changes in the indexation rules. These changes significantly affected the short run behavior of most of the series, which this paper is concerned. In particular the real balances, the money velocity, the rate of inflation, and the interest rate series, exhibit after 1986 great outliers that disturb the econometric work. Thus, we chose to make separate reports for the periods 1966.01/1985.12 and 1966.01/1990.03, and to investigate what happened with the series DGPs during each of the periods.

The accomplishment of a bootstrap test was directly connected with the results of the sample tests. If the null hypothesis of a unit root was far from being rejected by the ADF and OPP tests, we did not perform the bootstrap. If in the period the tests displayed contradictory results, then we performed the bootstrap test.

[10] In the next section, we present the bootstrap results for the different number of replications in an increasing order. We note the drawing of the error term was done (to each size of replication) in a sequential way, in order to get coherent simulations. Thus for a given random seed we are always able to obtain the same simulated sample.

[11] Let $\hat{T}$ be the statistics observed from the data and T* the bootstrap statistics. The p-value Monte Carlo is defined as: $P[T^* \geq \hat{T}] = (1 - \text{rank}(\hat{T})/N) + 1/N$. Alternatively, we can write $P[T^* \leq \hat{T}] = 1 - ((1 - \text{rank}(\hat{T})/N) + 1/N)$.

As we have stated, we adopt the following decision rule: if the sample tests have different outcomes about the presence of a unit root and the bootstrap results present one same conclusion for both tests, then we accept it.

With the exception of the real output series, which has quarterly frequency, all the remaining series are available in monthly frequency. It is well known that the data aggregation may lead to the elimination of cycles, generation of nonexistent serial correlation structure and the accentuation of the persistence; see Working (1960), Harvey (1990b) and Rossana and Seater (1995). So if we keep the disaggregated data we will preserve all the available information included in each series. As a rule, the more information one uses, the more accurate the estimations will be; see Granger and Newbold (1986) and Harvey (1990b). For these reasons, only the real output series is studied with a quarterly frequency.

In this section, we report the sample t-statistics for the ADF, OPP, and Phillips-Perron (PP) unit root tests. For a matter of space, we report neither the coefficient test $\hat{\rho}$ nor the diagnostics tests of the regression tests[12].

In appendix A, we give further details about the accomplishment of the diagnostics tests and tabulate the asymptotic distributions for the relevant significance levels. In appendix B we depict the studied series and provide the pictures of the recursive estimates of the lagged dependent variable (coefficient ($\hat{\rho}$-1)) of the ADF test equation. In appendix C, we tabulate for the simulated bootstrap statistics, the moments of the respective frequency distribution.

In the cases where the bootstrap provides different results from the sample data test, we present a Monte Carlo experiment to investigate the bootstrap size distortions.

3.1 Real Output

This is a quarterly series formed by linking the IBGE index product with the index computed by Rossi (1988). We followed this methodology in order to have a series length that covered all the studied period. This series is seasonally unadjusted and we proceeded the tests with the series in logs.

The unit root test results for the series in level (LY) and for its first difference (ΔLY) are presented in the Table below. In the shorter period, all tests accepted the unit root hypothesis with sufficiently large p-values. In the longer period, the ADF and PP test rejected the null with a significance level of 10%, a result that is not confirmed by the OPP t-statistics. We suppose this conflict is an indication that specifying the unit root test with high order trend terms is not the best procedure.

TABLE 1: UNIT ROOT TESTS — SERIES: OUTPUT

SERIES	TEST	1966.1 TO 1985.4		1966.1 TO 1990.1	
		LAGS	$t\,\hat{\alpha}$	LAGS	$t\,\hat{\alpha}$
LY	ADF	4	-2.207	4	-2.753†
	PP	3	-1.814	3	-2.610†
	OPP	4	-3.359	4	-3.536
ΔLY	ADF			3	-3.569**
	PP			2	-12.461**

Notes: (1) ADF and PP tests specified without trend term. (2) OPP tests performed with a polynomial trend of fourth and fifth order respectively to each period. The symbols (†) and (**) represent rejection of the null of a unit root at the 10 and 1% significance levels respectively.

Figure B.2 shows the recursive estimate of the y_{t-1} coefficient in the ADF specification. After 1975 this coefficient presents a remarkable stability and exhibits a behavior that does not point out the presence of a unit root in the series. Therefore, a bootstrap test on the ADF and OPP tests is more appropriate for the whole period. The next step in our procedure is to verify if

[12] The estimated test equations for the series in levels and in first differences, as well as the respective diagnostics residual tests may be obtained from the author on request.

the Monte Carlo tests corroborate the results obtained from the sample data. The approach was performed taking the specifications of the ADF and the OPP tests and imposing the unit root hypothesis α=1. The results are in Table 2.

TABLE 2: MONTE CARLO P-VALUES PERIOD: 1966.1 TO 1990.3

REPLICATIONS	ADF				OPP			
	t-STATISTICS		$\hat{\rho}$ -STATISTICS		t-STATISTICS		$\hat{\rho}$ -STATISTICS	
	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE
100	4	0.030	1	0.000	44	0.430	30	0.290
500	11	0.020	1	0.000	220	0.438	163	0.324
1000	20	0.019	1	0.000	468	0.467	346	0.395
5000	69	0.014	1	0.000	2364	0.473	1752	0.350

The first conclusion is that both bootstrapped ADF statistics strongly reject the presence of a unit root, which supports the judgement that the real output is stationary around a level (drift) term. The results show both ADF statistics have low power (in the case of the output series DGP) and that the Monte Carlo test behaves in order to correct this power distortion.

Nevertheless, the bootstrap approach applied to the OPP test produced results that contradict our initial beliefs, i.e., the bootstrap approach only yields the exact size of the sample test and did not lead to the H_0 rejection. With a test specification that includes a fifth order time polynomial in the fitted regression, the unit root hypothesis is accepted with p-values superior to 35% (case of 5000 replications) in both bootstrap tests.

The results suggest that the OPP test does not have any important size distortion concerning the data generation process of this particular series[13]. In addition, they lead to the conclusion that the ADF test specification is better fitted to the series than the OPP equation with high order trend terms. For the output series, the inclusion of deterministic terms, other than the constant, reduces the test power in capturing a non-existent unit root.

In the appendix C, the features of each bootstrapped statistic distribution, computed using 5000 replications, reveal that we can assume the distributions are close to being a functional standard Wiener process. This is an important issue since, in the sample test, the coefficient (α) that is being tested is approximately 1 (see figure B.2). Hence, at first sight, we do not have reason to doubt the validity of our Monte Carlo tests.

Table 3 shows the results of a simple Monte Carlo experiment designed to assess the size properties of the simulated statistics. We restrict our attention to the statistics associated with the ADF test since they rejected the null. The fundamental innovations are normally distributed with mean zero and variance taken from the ADF test residuals. Varying the parameters of the polynomial lag structure one may assess the size of the statistics. Several essays indicated that the relevant parameters are those associated with the third and fourth lagged first difference of the dependent variable. Since the statistics are invariant to the true parameter values under the null hypothesis, size distortion (if any) may be evaluated by setting α=1 in the data generation function and varying the value of the interest parameters (θ_3 and θ_4); see OPP (1989). We chose to use as the range of variation, the extreme values of the 99% confidence intervals of the mentioned coefficients[14], in order to get a wide range of variation. We note that this procedure is technically valid, since usual t-tests associated with hypotheses about any individual coefficient of a unit root specification can be compared with standard *t* or N(0,1) tables; see Hamilton (1994).

[13] This observation concerns the t-statistics, because for the sample data the OPP ρ-test rejected H_0 at the 1% significance level. In this case, the ρ-statistics exhibits important size distortions, which were corrected by the bootstrap approach.

[14] In fact being constructed simultaneously the intervals provide a confidence region with probability at least of 97%.

TABLE 3: MONTE CARLO EXPERIMENT

PARAMETERS θ_3/θ_4	ADF			
	t-STATISTICS		$\hat{\rho}$ -STATISTICS	
	RANK	P-VALUE	RANK	P-VALUE
-0.42/0.18	24	0.023	1	0.000
-0.04/0.66	120	0.119	1	0.000

Notes: (1) Number of observations = 97. (2) Number of replications = 1000. (3) The true model is the same of the bootstrap procedure with the additional hypotheses on the coefficients of ΔLY_{t-3} and ΔLY_{t-4}. (4) Rejections based on a nominal size of 5%.

Concerning the t-statistics, the simulation points out that, when the parameters exceed their average values the bootstrap approach possesses some size distortion[15]. However, the reported size test is an extreme value and it is only 2 percent points above the conventional significance level of 10%. Consequently, we can deduce that the bootstrap tests applied to the output series do not have a material size distortion in the range of values chosen to the third and fourth difference lagged parameters. Thus, our bootstrap procedure performed well even if, only a little better than the Dickey-Fuller test.

We then infer that during the period 1966.1 to 1990.1, the output series is I(0) and it is a stationary AR(5) process around a drift. This conclusion can be perceived by inspecting the series plot (figure B.1).

The fact that the real output is an I(0) series, or near integrated, is not very surprising. In the beginning of the analyzed period, the Brazilian economy passed through a phase of accelerated growth. During this period, known as the "economic miracle", the GDP growth achieved the rate of 14% in 1973. This picture is well represented by the behavior of the recursive estimates, which points out that the series possessed explosive features until 1975. However, the presence of a unit root in the early seventies was dominated in the following period. The end of the business cycle expansion, associated with external shocks and the implementation of a systematical inflation control by means of aggregated demand control, caused the reduction of the GDP rate of growth. This process went deeper during the eighties. Hence, the real GDP annual growth slowed from an average rate of 8.4% between 1970 and 1980 to only 1.5% from 1980 to 1990. One striking feature that contributed to the country's poor economic performance in the 1980's was the fall in the ratio investment to GDP, from a peak of 25.8% in 1975 to 15.5% in 1990; see Carneiro (1997). This issue was not only related with a new wave of external shocks but also with the reduction in the public investment, in conjunction with the internal inflationary instability and dramatic changes of macroeconomic policies.

In resume, during most of the period from 1966 to 1990 the GDP rate of growth was well below its trend path of 7% per year. This probably induced the variance stationarity of the GDP series, even if it passed through an explosive stage in the period's initial years. Once more, we call attention to the recursive ($\hat{\rho}$-1) estimates plot, which clearly shows the coefficient near to the unit root bounds, but without ever crossing this limit after 1976.

3.2 Rate of Inflation and Monetary Expansion

The rate of inflation is calculated with the general price index (IGP) computed by Fundacao Getulio Vargas. The monetary expansion is the percent variation of the M1 (currency plus personal checking accounts) monetary aggregate computed by the Brazilian Central Bank.

[15] Regarding the $\hat{\rho}$ -statistics there is no apparent size distortion.

Tables 4 and 5 show the results of the ADF, PP, and OPP unit root tests, respectively for each series.

TABLE 4: UNIT ROOT TESTS SERIES: RATE OF INFLATION

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.03	
		LAGS	$t\,\hat{\alpha}$	LAGS	$t\,\hat{\alpha}$
PI	ADF	2	-2.382	19	+3.345
	PP	4	-5.761**	5	+3.287
	OPP	2	-5.288**	19	-1.115
ΔPI	ADF			12	-3.777*
	PP			5	-14.807**

Notes: (1) ADF and PP tests specified with and without trend term, respectively, to 1966/85 and 1966/90. (2) OPP tests performed with a polynomial trend of third and fifth order respectively to each period. (3) To the series first difference in period 1966.01/90.**02,** the ADF test rejected the null at the level of 1%. The symbols (*) and (**) represent rejection of the null of a unit root at the 5 and 1% significance levels respectively.

TABLE 5: UNIT ROOT TESTS SERIES: MONETARY EXPANSION

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.03	
		LAGS	$t\,\hat{\alpha}$	LAGS	$t\,\hat{\alpha}$
MI	ADF	11	+0.507	15	+4.562
	PP	4	-15.943**	5	-5.115**
	OPP	12	-4.921*	15	+2.287
ΔMI	ADF			13	-2.799/ -4.390**
	PP			5	-24.856**

Notes: (1) ADF and PP tests specified with and without trend term, respectively, to 1966/85 and 1966/90. (2) OPP tests performed with a polynomial trend of fifth and fourth order respectively to each period. (3) ADF test to series first difference is reported to periods 66/90.03 and 66/90.02. The symbols (*) and (**) represent rejection of the null of a unit root at the 5 and 1% significance levels respectively.

Concerning the rate of inflation, there is a discord among the periods' results. To the shorter, the PP[16] and OPP test statistics reject the null with a p-value inferior to 1%, while the three statistics accept the null of the presence of a unit root in 1966/90. The main reason that seems to support this distinction is the number of lags necessary in the longer period, to make the test residuals independent, which caused the ADF and PP test statistics positive signals. The long polynomial of lags is due to the existence of different regimes in the Brazilian economy during the whole period. This changing regime environment is very well illustrated by the coefficient instability of the lagged dependent variable, in the ADF equation (figure B.4). The above results suggest the rate of inflation followed a random walk[17] process from 1966 to 1990.

The test results for the money growth are similar to the inflation, in spite of some differences. We would like to point out the stronger persistence and seasonality of this series compared to the rate of inflation. We also stress that the PP test rejects H_0 in the period 1966/90, while the ADF test only rejects the null in the series first difference, if we truncate the period on month 1990.02. This is caused by the presence of a great outlier that marks the end of the Cruzado era plans. The intrinsic instability of the Brazilian monetary policy is evident in figure B.5 (recursive estimate coefficient).

For the two series, we have two distinct periods concerning the unit root tests. In the shorter, there is some probability of rejecting the null, while in the greater the evidence of a unit root is too strong. So following these indications, we performed the bootstrap approach only for

[16] It is a known fact that the finite-sample properties of the unit root tests have poor results for at least some specifications of the error process. Pertaining to the Phillips-Perron test it is likely to reject the null hypothesis of a unit root although it is true, when the error term has a MA(1) close to –1. However several experiments showed that if the reported tests have poor properties (if any), the reason is not due to a MA(1) representation in the error term.

[17] In the ADF test of the inflation first difference, we identified a strong autocorrelation structure in the residuals. If we had used all the lagged first difference necessary to make the residuals independent, we would have accepted the unit root hypothesis and would have concluded that the rate of inflation was I(2). To withdraw this misfortune we truncated the polynomial lagged difference at the 13th lag, which was significant, and accepted the presence of some autocorrelation in the residuals. We then performed the Kolmogorov-Smirnov (K-S) test, which, with maximum gap of 0.0392 at frequency 1.9635 accepted at the level of 10% the residual serial independence. The bayesian unit root test also corroborates the rejection of the unit root hypothesis in the inflation first difference.

the years 1966/85. For the rate of inflation, we bootstrapped the ADF and OPP tests, and for the money growth only the OPP test.

The bootstrap tests applied to the inflation rate strongly support the conclusion of the ADF sample test (see Table 6). Moreover, the computed Monte Carlo p-values are bigger than those of the sample data tests are. Since the statistics distributions are in accordance with the theory, we accept these results.

TABLE 6: MONTE CARLO P-VALUES SERIES: RATE OF INFLATION PERIOD: 1966.01 TO 1985.12

REPLICA TIONS	ADF				OPP			
	t-STATISTICS		$\hat{\rho}$ -STATISTICS		t-STATISTICS		$\hat{\rho}$ -STATISTICS	
	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE
100	100	0.990	100	0.990	96	0.950	100	0.990
500	500	0.998	500	0.998	452	0.902	475	0.948
1000	1000	0.999	1000	0.999	900	0.899	938	0.937

Bootstrapping the OPP test, we attained different results from those computed with the observed data. The null hypothesis is accepted with a significance level near 90%, and the test statistics distribution has the expected negative skewness. This outcome leads to the conclusion that the OPP tests have important size distortions in this case, that is, the test is likely to reject a null hypothesis of a unit root in finite samples although it is true. This technique seems to be poor for the series underlying DGP and our approach is at work toward correcting this distortion.

Table 7 shows the Monte Carlo test results for the monetary expansion. The bootstrap OPP t-statistics has different results from the sample test and suggests the acceptation of the null. Notwithstanding, the $\hat{\rho}$ -statistics induces to the null rejection, but it has a degenerated distribution[18] (meanwhile the t-bootstrap approximates to a normal distribution). Thus, we will not take into account this result, considering that, in this case, the approach did not perform as it should have.

In resume, we have enough evidence to conclude that both series were integrated of order one during the whole period. In regarding the existing doubts about the period from 1966 to 1985, the bootstrap tests upheld the ADF test and denied the OPP sample results. These results are coherent with the studies of Vals Pereira (1988), Novaes (1991), Cerqueira (1993), and Pastore (1995 and 1997).

TABLE 7: MONTE CARLO P-VALUES
SERIES: MONEY GROWTH PERIOD: 1966.01 TO 1985.12

REPLICA TIONS	OPP			
	t-STATISTICS		$\hat{\rho}$ -STATISTICS	
	RANK	P-VALUE	RANK	P-VALUE
100	45	0.440	2	0.010
500	212	0.422	24	0.046
1000	444	0.443	50	0.049

The remaining question is to determine if the maintained processes have or have not a polynomial trend term. Following Ouliaris, Park and Phillips (1989) paper we apply a likelihood ratio sequential test for the hypothesis, $\mathbf{H_0:\alpha=1,}$ **and** $\boldsymbol{\beta_p=0}$, on the OPP test specification. We

[18] We conjecture this behavior is caused by the large number of lags demanded by the parametric correction for serial correlation in the residuals. This fact seems to have disturbed the bootstrap simulations, which generated values (in absolute terms) much greater than the sample statistics and caused the degeneration of the bootstrap-ρ (OPP) distribution.

compare the obtained statistics with Table III presented in the paper above and Table B.7 in Hamilton (1994).

For the rate of inflation, we sequentially accept the null up to the squared trend term. We reject the hypothesis of a non-significant linear trend term, at 5% significance level. We conclude this series is an ARIMA(20,1,0) process with trend. We follow the same course for the monetary growth and conclude that it follows an ARIMA(16,1,0) process with drift.

The conclusion that the rate of inflation and the monetary growth are both I(1) is coherent in an economy with a permanent operational public deficit not completely financed by issuing bonds. Moreover, the explosive feature of the money growth is in accordance with an ever-increasing public debt, which has a stationary rate of growth. In fact, the ratio of operational deficit to GDP increased between 1975/81 and 1982/89 from an annual average of 2.7% to 4.4%. Meanwhile the collected seigniorage as a GDP proportion grew from 1.9% to 2.9%. In the same way, an ever-increasing monetary expansion must be followed by a megainflation. From 1975 to 1989, the inflation rate increased from an annual rate of 29.4% to 1748%.

Although being explosive processes both series are difference stationary and so one necessary condition to the nonexistence of a rational inflationary bubble is already satisfied, i.e., the inflation series is stationary of a finite order of differentiation. It is well known that the presence of bubbles precludes the stationarity of any degree of differencing of the inflation series. Therefore, we have space to perform a test to verify if there was a rational bubble between 1986 and 1990; see Diba and Grossman (1988) and Welch (1991).

The second condition is that the inflation and the money growth must cointegrate. In Cerqueira (1998) we demonstrated the series were cointegrated with vector (1,-1) and a time-varying drift term, from 1966 to 1985, ruling out the possibility a rational inflationary bubble. Preliminary experiments also showed the same conclusion holds for the period from 1986.01 to 1990.03, but this issue will remain to be confirmed in further research.

3.3 Public bonds interest rate

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 real interest rate is the above series discounted by the monthly rate of inflation. As with the rate of inflation and the money growth, we worked with these series without logarithms.

Table 8 shows the unit root test outcomes for the nominal interest rate. It is easy to see that all tests point out the series as difference stationary[19]. Even the OPP t-statistics accepts the null hypothesis with a significance level well above 20%[20], for both periods. The recursive estimates on figure B.7 also indicate the presence of a unit root in the years after 1982 and the strong instability of this coefficient.

[19] In the ADF test of the first difference series we had problems in dealing with the residuals serial correlation. Therefore, we also performed the K-S test for serial independence, which accepted the null with a level well above 10%. Furthermore, the bayesian test strongly rejected the unit root hypothesis in the series first difference.
[20] The critical value corresponding to a 20% significance level is equal to -4.216.

TABLE 8: UNIT ROOT TESTS SERIES: NOMINAL INTEREST RATE

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.03	
		LAGS	t $\hat{\alpha}$	LAGS	t $\hat{\alpha}$
IM	ADF	13	-0.660	9	+2.046
	PP	4	-1.536	5	+1.311
	OPP	13	-2.371	9	-2.550
ΔIM	ADF			8	-9.376**
	PP			5	-8.114**

Notes: (1) ADF and PP tests specified with trend term. (2) OPP tests performed with a polynomial trend of fifth order to both periods. The symbol (**) represents rejection of the null of a unit root at the 1% significance level.

In the present case, we have no reason to carry out a bootstrap test. We then conclude that the nominal interest rate approximately follows an ARIMA(10,1,0) process with a linear trend[21].

Table 9 shows the results for the real interest rate. We firmly reject the unit root hypothesis and conclude that the series is well approximated by an AR(2) process with drift.

TABLE 9:UNIT ROOT TESTS SERIES: REAL INTEREST RATE

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.03	
		LAGS	t $\hat{\alpha}$	LAGS	t $\hat{\alpha}$
R	ADF	2	-4.370**	1	-9.530**
	PP	4	-8.260**	4	-8.684**

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

The fact that the nominal interest rate is I(1) is consistent with an ever-growing inflationary process, meaning that the cost of holding money follows the inflation path. This is one reason that supports the existence of a Fisher effect during the period. Another reason is the real interest rate stationarity, which corresponds to the hypothesis that the nominal interest rate and the inflation are cointegrated with a known cointegrating vector of (1,-1)[22], and the real rate is interpreted as the long-run equilibrium error of this relation. If we assume the expectations are rational and suppose the absence of any kind of noise in the error term, the expected inflation is well represented by the current inflation rate; see Garcia (1991). It may therefore be the case that there was a Fisher effect during the period 1966/90, and so, in the long-run, the changes in the nominal interest rate reflected the same percent variations of the expected rate of inflation.

As showed above, the real interest rate follows an AR(2) process, which means that its past values had information about its current behavior. This implies that the hypothesis of a constant expected value for the real interest rate has no empirical support. However, this is not a contradiction with hypothesizing the Fisher effect, since according to this hypothesis the only assumption is that the nominal interest rate and the inflation have the same long run trajectory.

[21] We achieved this conclusion by doing the same sequential procedure described in the former section. With usual significance levels, we accept the absence of trend terms of order superior to one.

[22] Our first experiments using Johansen procedure pointed out a cointegration relation between these two variables. However, further studies are necessary because we did not succeed in getting NIID residuals. Testing with the Engle-Granger two step test we got a cointegration relation with the inflation coefficient around 0.90. Since this procedure does not require Gaussian residuals, we can conclude that there is a stable long-run relation between the interest rate and inflation, which confers consistency to the Fisher effect. Albeit we can not perform hypothesis testing either on the coefficients equation or on the causality relation between the two variables.

3.4 Public debt

The nominal public debt is the stock of treasury bonds held by the private agents. We investigated the real debt and the debt-GDP[23] ratio series. Both series are in logarithms and we truncated the period at 1990.02[24].

Table 10 shows the unit root tests results for the real debt stock. In contrast with the others tests, the OPP test rejected H_0 for the shorter and the longer period, respectively, at 5% and 20% significance levels. Nevertheless, the recursive estimates point out a strong and evident unit root during the whole period. Thus, we decided to bootstrap these test statistics, reported in Table 11, in order to clarify the apparent contradiction among the results.

TABLE 10: UNIT ROOT TESTS SERIES: REAL PUBLIC DEBT

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.02	
		LAGS	$t\hat{\alpha}$	LAGS	$t\hat{\alpha}$
LDR	ADF	3	-1.233	3	-1.361
	PP	4	-1.214	5	-1.376
	OPP	9	-4.949*	9	-4.182
ΔDR	ADF			2	-8.736**
	PP			5	-16.452**

Notes: (1) ADF and PP tests specified without trend term. (2) OPP tests performed with a polynomial trend of fourth and fifth order respectively to each period. The symbols (*) and (**) represent rejection of the null of a unit root at the 5 and 1% significance levels respectively.

TABLE 11: MONTECARLO P-VALUES SERIES:REALDEBT TEST: OPP

REPLICATIONS	1966.01 TO 1985.12				1966.01 TO 1990.02			
	t-STATISTICS		$\hat{\rho}$ -STATISTICS		t-STATISTICS		$\hat{\rho}$ -STATISTICS	
	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE
100	85	0.840	56	0.550	79	0.780	74	0.730
500	413	0.824	229	0.456	390	0.778	352	0.702
1000	837	0.836	453	0.452	825	0.824	719	0.718

The bootstrap tests accept the null hypothesis with p-values far from those of the sample tests. This indicates the OPP test presents important size distortion and suggests that the OPP is not a performing unit root test for the real debt series. Hence, we accept the null of the presence of a unit root in the real debt series and conclude the series follows an ARIMA(4,1,0) process without trend but with drift[25].

By studying figure B.11, the presence of a unit root in the debt-GDP series is remarkable after 1980, but prior to this year this conclusion is not evident. However, none of the applied tests rejected the null for the period 1966/85; see Table 12. As showed in Table 13, the ADF bootstrap test statistics corroborate these results and help settle the correct size of the sample test.

[23] To compute this series we first interpolated the GDP series using the Kalman filter procedure; see Cerqueira 1998.
[24] This is a necessary course, since at 1990.03 the first Collor (stabilization) plan blocked 90% of private savings. This caused a big outlier in the debt series that can only be smoothed over with a sample expansion, which is out of the objective of the present paper.
[25] We inferred this conclusion from the sequential test earlier described. By inspecting figure B.10, one can observe that the series is better represented if a non-zero drift is supposed.

TABLE 12:UNIT ROOT TESTS SERIES: DEBT-GDP RATIO

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.02	
		LAGS	t $\hat{\alpha}$	LAGS	t $\hat{\alpha}$
LDY	ADF	3	-2.326	0	-2.068
	PP	4	-2.047	5	-2.337
	OPP	3	-3.697	3	-3.553
ΔLDY	ADF			0	-16.057**
	PP			5	-16.111**

Notes: (1) ADF and PP tests specified with trend term. (2) OPP tests performed with a polynomial trend of fifth order to both periods. The symbol (**) represents rejection of the null of a unit root at 1% significance level.

TABLE 13:MONTE CARLO P-VALUES
SERIES: DEBT-GDP PERIOD: 1966.01 TO 1985.12

REPLICA TIONS	ADF			
	t-STATISTICS		$\hat{\rho}$ -STATISTICS	
	RANK	P-VALUE	RANK	P-VALUE
100	53	0.520	73	0.720
500	278	0.554	383	0.764
1000	546	0.545	741	0.740

The convincing unit root in the period 1966/90 caused stronger bootstrap results (Table 14) compared to those achieved in the shorter period. We presume this outcome came about because we are working with a "pure" random walk process that does not need parametric corrections in the test specifications. Surprisingly the ADF tests displayed important size distortion, which seems somewhat odd as the ADF was originally designed to work with this type of stochastic process.

TABLE 14:MONTE CARLO P-VALUES SERIES: DEBT-GDP PERIOD: 1966.01 TO 1990.02

REPLICA TIONS	ADF				OPP			
	t-STATISTICS		$\hat{\rho}$ -STATISTICS		t-STATISTICS		$\hat{\rho}$ -STATISTICS	
	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE
100	100	0.990	100	0.990	85	0.840	90	0.890
500	500	0.998	500	0.998	439	0.876	450	0.898
1000	1000	0.999	1000	0.999	872	0.871	888	0.887

The same as for the others series, the OPP bootstrap results demonstrate that this is a test with acute size distortion. The only distinct feature is the non-degeneration of the $\hat{\rho}$-statistics distribution, probably due to the parsimonious number of lagged first difference included in the simulation of the dependent variable.

The debt-GDP ratio series is therefore a random walk with drift, i.e., an ARIMA(0,1,0) process.

The assumption that the real debt and the debt-GDP are difference stationary implies the public debt was sustainable and the government intertemporal budget constraint was being satisfied. Because the stock of public bonds increases by the difference between the operational deficit and the seigniorage, the real debt difference stationarity hypothesis is equivalent to assuming a cointegration relation between these two variables; see Welsh (1991) and Pastore (1995).

Thus, the result shows that the Brazilian public debt did not have an explosive growth[26], which is in accordance with the argument that during the eighties the government generated the necessary seigniorage, in order to avoid a debt explosive growth, by continuously increasing the monetary expansion. Hence, if some risk of debt default was perceived by the private agents, it could not have come about due to the government's lack of ability to pay, or to the violation of its intertemporal budget constraint.

3.5 Real balances and monetization ratio

The M1 aggregate is defined as the total currency plus the balances held in checking accounts. It is computed monthly by the Central Bank. The real money (or real balances) series is the M1 aggregate over the general price index. The monetization ratio is the real money divided by the real output. This last series is the inverse of the money velocity.

Table 15 shows that the OPP test rejected the null of a unit root for the period 1966/90 for the real money series. Comparing the recursive estimates of the lagged dependent variable of the ADF and OPP equations, one may note the stability of the coefficient in the OPP specification (note that the scale are different), and the large distance the coefficient kept from the bounds of a unit root during the whole period. This suggests that for this series the OPP test seems to provide a suitable specification. Moreover, the ADF test[27] is likely to provide a false result to the sample data, since it does not have an accurate specification regarding the deterministic terms.

TABLE 15: UNIT ROOT TESTS SERIES: REAL MONEY

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.03	
		LAGS	$t\hat{\alpha}$	LAGS	$t\hat{\alpha}$
LMR	ADF	12	-1.487	13	-1.697
	PP	4	-1.022	5	-1.887
	OPP	12	-4.297	16	-4.657*
ΔLMR	ADF			12	-5.054**
	PP			5	-15.772**

Notes: (1) ADF and PP tests specified with trend term. (2) OPP tests performed with a polynomial trend of fifth and fourth order respectively to each period. . The symbols (*) and (**) represent rejection of the null of a unit root at the 5 and 1% significance levels respectively.

The above remarks are firmly confirmed by the bootstrap approach, whose results are in Table 16. With the exception of the OPP $\hat{\rho}$-statistics[28], the others strongly reject the unit root hypothesis. In this case, the bootstrap approach corrected the power distortion held by the ADF test statistics. In addition, the OPP t-bootstrap produced the correct sample size test. We then performed a Monte Carlo experiment to investigate the bootstrap size distortion. The guidelines of this procedure are described in section 3.1[29] and the results are reported in Table 17. The Monte Carlo evidence supports that the realized bootstrap tests do not possess any material size distortion, in the range of values chosen to the first and fourth lagged differences parameters. This shows that our approach corrected the sample tests power distortions.

[26] Our results are complementary to the works of Pastore (1995) and Issler and Lima (1997), since we test sustainability using different period and technique from those employed by these authors.
[27] The residuals of this test equation presented serial correlation of order 11th and 12th. However, the K-S convincingly (well above the 10% level) accepted the residual series independence.
[28] Even in the present case, the statistics frequency distribution degenerated. We notice that in the sample test it has a positive value around 115.0753. Therefore, we do not consider this outcome.
[29] We chose to make variations on the first and third lagged difference parameters, since these are the regressors, which stand the t-statistics on its highest values. The observation applies for both test equations. We also performed experiments by changing the 16th lagged difference, which is the regressor that supports the non-serial correlation hypothesis. The results are quite similar to those reported in Table 17.

TABLE 16: MONTE CARLO P-VALUES SERIES: REAL MONEY PERIOD: 1966.01 TO 1990.03

REPLICATIONS	ADF				OPP			
	t-STATISTICS		$\hat{\rho}$ -STATISTICS		t-STATISTICS		$\hat{\rho}$ -STATISTICS	
	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE
100	1	0.000	1	0.000	1	0.000	66	0.650
500	1	0.000	1	0.000	1	0.000	342	0.682
1000	1	0.000	1	0.000	1	0.000	686	0.685

TABLE 17: MONTE CARLO EXPERIMENT P-VALUES SERIES: REAL MONEY PERIOD: 1966.01 T0 1990.03

PARAMETERS θ_1/θ_3	ADF		PARAMETERS θ_1/θ_3	OPP	
	t-STATISTICS	$\hat{\rho}$ -STATISTICS		t-STATISTICS	$\hat{\rho}$ -STATISTICS
0.082/0.019	0.000	0.000	0.109/0.077	0.000	0.750
0.403/0.362	0.000	0.009	0.433/0.428	0.000	0.842

Notes: (1) Number of observations = 291. (2) Number of replications=1000. (3) The true model is the same of the bootstrap procedure with the additional hypotheses on the coefficients of ΔLMR_{t-3} and ΔLMR_{t-3} . (4) Rejections based on a nominal size of 5%.

Thus, we achieved a different conclusion from the ADF sample test, given that by the bootstrap tests the real money series is I(0). We conclude the series is an AR(17) with a polynomial trend of fourth order in the deterministic terms.

The monetization ratio or the inverse of money velocity is pointed out as a stationary variable by the ADF and OPP tests[30], during the period from 1966 to 1990 (Table 18). Figure B.16 ratifies this conclusion. The bootstrap results in Table 19 endorse these results and provide accurate size tests[31]. However, concerning the OPP test the approach seems to have some size distortion (Table 20), when the values of the coefficients of ΔLMY_{t-1} and ΔLMY_{t-3} approach the interval higher bound. From a practical standpoint of view, some size distortion is not surprising in finite samples, when the parameters undertake extreme values. If we reduce the upper bound by tightening the confidence intervals for 95, 90 and 80%, the simulated t–statistics p-values fall to 0.146, 0.135, and 0.112, respectively. Thus, we do not need to be afraid of detecting some size distortion.

Therefore, we have evidence of the monetization ratio and the money velocity stationarity. We can conclude that these series follow a stationary AR(4) process around a second order polynomial trend.

TABLE 18: UNIT ROOT TESTS SERIES: MONETIZATION RATIO

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.03	
		LAGS	$t\,\hat{\alpha}$	LAGS	$t\,\hat{\alpha}$
LMY	ADF	12	-2.115	3	-3.196†
	PP	4	-1.669	5	-2.733
	OPP	15	-5.183**	3	-3.866*
ΔLMY	ADF			12	-5.139**
	PP			5	-15.167**

Notes: (1) ADF and PP tests specified with trend term. (2) OPP tests performed with a polynomial trend of fifth and second order respectively to each period. The symbols (†), (*) and (**) represent rejection of the null of a unit root at the 10, 5 and 1% significance levels respectively.

TABLE 19: MONTE CARLO P-VALUES SERIES: MONETIZATI0N RATIO PERIOD: 1966.01 TO 1990.03

REPLICATIONS	ADF				OPP			
	t-STATISTICS		$\hat{\rho}$ -STATISTICS		t-STATISTICS		$\hat{\rho}$ -STATISTICS	
	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE
100	1	0.000	2	0.010	5	0.040	7	0.060
500	1	0.000	4	0.006	22	0.042	43	0.084
1000	1	0.000	8	0.007	50	0.049	92	0.091

[30] The $\hat{\rho}$ -statistics rejected the null at a significance level near to 1%.

[31] Particularly to this series all frequency distributions corresponding to the bootstrap $\hat{\rho}$ -statistics are well behaved and have the required negative asymmetry.

TABLE 20: MONTE CARLO EXPERIMENT P-VALUES SERIES: MONETIZATION RATIO PERIOD: 1966.01 T0 1990.03

PARAMETERS θ_1/θ_3	ADF		PARAMETERS θ_1/θ_3	OPP	
	t-STATISTICS	$\hat{\rho}$ -STATISTICS		t-STATISTICS	$\hat{\rho}$ -STATISTICS
0.072/0.063	0.000	0.000	0.080/0.076	0.000	0.000
0.400/0.407	0.000	0.049	0.406/0.419	0.173	0.284

Notes: (1) Number of observations = 291. (2) Number of replications=1000. (3) The true model is the same of the bootstrap procedure with the additional hypotheses on the coefficients of ΔLMR_{t-3} and ΔLMR_{t-3}. (4) Rejections based on a nominal size of 5%.

The assumption that the real balances and the monetization ratio (money velocity) are I(0) series has important implications. The first one is that the money demand contraction that took place during the seventies and eighties did not have an explosive feature. In this sense, one may not conclude that the high inflationary levels observed at the end of the eighties was caused by a money demand reduction that led the economy to the Cagan's explosive region. Thus, the financial innovation process that occurred during the period and induced the monetary demand contraction was not strong enough to cause an explosive money velocity. The recursive estimates suggest that if these series were near to having integrated features, this tendency was reverted during the eighties for a reason that remains to be explained. One possible explanation is that the Cruzado Plans had succeeded in reducing the speed of this movement.

Engsted (1994) argues that if the real balances are I(1) and the velocity shock is stationary, then the Cagan's model, under rational expectations and with no bubbles, has the testable implication that the real money cointegrate with the growth rate of money. Nevertheless, we have showed in this study that the real money and the monetary growth have a different integration order. Thus, one cannot postulate a cointegration between them. This implies that this version of Cagan's model is not well fitted to the features of the Brazilian series.

3.6 Exchange rates

In this section, we focus our attention on the premium in the black market for American dollars. The premium is defined as the percentage excess of the black market price of dollars over the official exchange rate. We collected the series from the Central Bank bulletin. All calculations were made with the series in logarithms. In Tables 21 to 23, we report the respective unit roots tests for each of these three series.

TABLE 21: UNIT ROOT TESTS REPORT SERIES: OFFICIAL EXCHANGE RATE

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.03	
		LAGS	$t\hat{\alpha}$	LAGS	$t\hat{\alpha}$
LUSO	ADF	1	+9.163	1	+6.305
	PP	4	+13.898	5	+12.154
	OPP	1	-4.200*	1	-0.478
ΔLUSO	ADF			5	-5.496**
	PP			1	-5.300**

Notes: (1) ADF and PP tests specified without trend term. (2) OPP tests performed with a polynomial trend of third and fourth order respectively for each period. The symbols (*) and (**) represent rejection of the null of a unit root at the 5 and 1% significance levels respectively.

TABLE 22: UNIT ROOT TESTS REPORT SERIES: BLACK MARKET RATE

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.03	
		LAGS	t $\hat{\alpha}$	LAGS	t $\hat{\alpha}$
LUSB	ADF	3	+7.277	2	+7.429
	PP	4	+4.930	5	+12.320
	OPP	1	-4.630*	0	-1.052
ΔLUSB	ADF			2	-4.720**
	PP			4	-10.325**

Notes: (1) ADF and PP tests specified without trend term. (2) OPP tests performed with a polynomial trend of third and fifth order respectively for each period. The symbols (*) and (**) represent rejection of the null of a unit root at the 5 and 1% significance levels respectively.

TABLE 23: UNIT ROOT TESTS REPORT SERIES: DOLLAR PREMIUM

SERIES	TEST	1966.01 TO 1985.12		1966.01 TO 1990.03	
		LAGS	t $\hat{\alpha}$	LAGS	t $\hat{\alpha}$
LAG	ADF	6	-3.611*	3	-3.455*
	PP	4	-4.044**	5	-4.153**
	OPP	6	-3.707†	3	-3.966*

Notes: 1) ADF and PP tests specified with trend term. 2) OPP tests performed with a polynomial trend of second order for both periods. The symbols (†), (*) and (**) represent rejection of the null of a unit root at the 10, 5 and 1% significance levels respectively.

Regarding the dollar price series of both markets, one striking issue is the rejection of the null by the OPP test during the period 1966/85. This is a distortion held by the test when it handles ever-increasing nominal series (see figure B.17), that in general has a nearly exponential shape (when expressed in logarithms). This feature, in most of the cases, led the test to wrongly reject the unit root hypothesis. Therefore, we do not take into account these results and conclude that the official and the black market rate were difference stationary series between 1966 and 1985, as during 1966 to 1990.

On the other hand, the dollar premium is, by the three tests, a trend stationary series. The results are soundly confirmed by the bootstrap tests (table 24). We have enough evidence to assume that the dollar premium is an AR(4) stationary process around a second order polynomial trend.

TABLE 24: MONTE CARLO P-VALUES SERIES: DOLLAR PREMIUM PERIOD: 1966.01 TO 1990.03

REPLICATIONS	ADF				OPP			
	t-STATISTICS		$\hat{\rho}$ -STATISTICS		t-STATISTICS		$\hat{\rho}$ -STATISTICS	
	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE	RANK	P-VALUE
100	1	0.000	1	0.001	1	0.000	1	0.000
500	1	0.000	2	0.002	1	0.000	1	0.000
1000	1	0.000	8	0.007	1	0.000	1	0.000

The dollar premium stationarity implies the cointegration between the official rate and black dollar price with a known cointegrating vector of (1,-1). This means that even being a price from a speculative market the black dollar had a stable, long run relationship with the official market price. Moreover, this relationship implied that one price could be used to help forecast the other. At the same time, there might be Granger causality in at least one direction (a matter for further research)[32]. Therefore, from this point of view, the policy makers' apprehension of an explosion in the parallel market that occurred in the eighties did not have clear empirical support. Figure B.18 illustrates that the dollar premium was too distant from a non-stationary process.

[32] Preliminary experiments pointed out the series were cointegrated with vector (1,-1). The black dollar price is weakly exogenous for the parameters of the official rate but the inverse does not seem to be true. The causal relation is bi-directional in the Granger sense. However, all tests were performed with the violation of the Gaussian hypothesis.

One possible reason that may justify why both prices move closely in the long run was the way in which the crawling peg policy was conducted during the seventies and eighties. This policy prevented any speculative run in the black market even if it was not able to forestall an overvaluation throughout the entire period.

4. Monte Carlo Main Results

In section 3, we applied our procedure to some Brazilian macroeconomic time series. Then using the unit root test specification as start point we found the series representation. The approximations to the series true processes are shown in the table below.

TABLE 25: DATA GENERATION PROCESSES

SERIES	INTEGRATION ORDER	DGP	DETERMINISTIC TERMS	
			TREND	SEASONALS
GDP	I(0)	ARIMA(5,0,0)	drift	S2
INFLATION	I(1)	ARIMA(20,1,0)	drift, t	S2
MONEY GROWTH	I(1)	ARIMA(16,1,0)	drift	S2 TO S12
NOMIN. INTEREST	I(1)	ARIMA(10,1,0)	drift, t	S3
REAL INTEREST	I(0)	ARIMA(2,0,0)	drift	--
REAL DEBT	I(1)	ARIMA(4,1,0)	drift	--
DEBT-GDP	I(1)	ARIMA(0,1,0)	drift	S4
REAL MONEY	I(0)	ARIMA(14,0,0)	drift, t^2, t^3, t^4	S2 TO S12
MONETIZ. RATIO	I(0)	ARIMA(4,0,0)	drift, t^2	S2 TO S12
OFFICIAL DOLLAR	I(1)	ARIMA(2,1,0)	drift	--
BLACK DOLLAR	I(1)	ARIMA(3,1,0)	drift	--
DOLLAR PREMIUM	I(0)	ARIMA(4,0,0)	drift, t^2	--

In the empirical applications we obtained two types of results. The first type comprise the set of tests, which the bootstrap approach led to the rejection of the unit root hypothesis, by at least one of the bootstrapped tests.

For this type of result, when we bootstrapped the ADF test, the bootstrap coefficient and the t-test had similar behaviors. In three cases, the bootstrap approach changed the sample test results, suggesting that both statistics possess substantial power distortion. In one case (money velocity), the sample coefficient statistics rejected the null while the t-test marginally rejected it. The bootstrap approach strongly rejected the null. In general, the bootstrap tests proved to have more power than the ADF tests, especially concerning the t-test, suggesting the sample $\hat{\rho}$-statistics have better performance than the t-test. The results demonstrated that both bootstrap statistics had similar performances and they did not present any material size distortion.

The bootstrap statistics corresponding to the OPP test had a different pattern from the ADF test. In the case (the GDP series) in which the deterministic regressors were misspecified, the bootstrap approach corrected the size distortion presented in the sample test statistics and ratified the H_0 acceptation. In one case (real balances) the lengthy polynomial of lagged first difference caused the degeneration of the bootstrap $\hat{\rho}$-statistics distribution and induced the null acceptation. Meanwhile, the null was rejected by the t-test. In the cases where the sample tests had a small number of lags, the bootstrap-$\hat{\rho}$ did not degenerate and both bootstrap statistics rejected the null. The bootstrap-t performed better than the coefficient test regarding the distribution behavior and the power test. The bootstrap procedure helped in setting the accurate significance level for both statistics (when the distribution did not degenerate), without presenting important size distortion.

With the exception of the coefficient test associated with the OPP tests, the remaining three bootstrap statistics revealed great ability in detecting an “evident” false unit root hypothesis. It is interesting to note that when H_0 was rejected by the bootstrap-t OPP, it was simultaneously rejected by both bootstrap unit root type tests. Bearing in mind that we are handling a set of particular sample data, this may suggest the OPP test is just giving information about the deterministic trend terms the series process may have. Thus, it is playing no role in determining the series integration order. Therefore, for unit root testing we should only perform the ADF test.

The second type of results includes the cases in which the four bootstrap statistics accepted the null hypothesis. For the series in which the sample OPP tests rejected the null, the bootstrap gave opposite results and got the same conclusion as the sample ADF tests[33]. This demonstrated that the OPP tests have, in our context, significant size distortions. As mentioned above, in one case (monetary expansion) the bootstrap coefficient statistics degenerated due to a large number of lagged difference, which suggests that the bootstrap-t OPP had better performance, when dealing with series with strong persistence.

In the typical case of a random walk (debt-GDP ratio) the ADF statistics failed to give a correct real size test. Their values were not too far from the rejection region, as should be expected if we take as benchmark the bootstrap p-values. This shows that one must be careful in working with this test as it may have significant size distortions and may lead to the null rejection even if it is a true hypothesis. In our experiments both bootstrap test statistics had similar performances regarding the size tests.

In the cases of the null acceptation, a likelihood sequential procedure was performed in order to verify the significance of the high order trend terms. In all cases their non-significance was accepted. This is in accordance with the above remark about the usefulness of employing the OPP tests. Moreover, in our particular context the sample ADF tests presented lower size distortion than the OPP type tests.

Broadly speaking, the bootstrap technique performed better when applied to the ADF tests than when applied to the OPP tests. This conclusion came forth the non-degeneration of the bootstrap-$\hat{\rho}$ ADF statistics, when a large number of lags were necessary to control residual serial correlation, and from the absence of any striking size distortion.

Furthermore, the bootstrap approach performed quite better than the ADF and the OPP tests regarding the size accuracy and the higher power in hypothesis testing unit root against the alternative of stationarity. *Thus, it seems to be wise to perform the sample unit root tests and proceed the analysis by bootstrapping them.*

From the reported testing is not completely clear the OPP test usefulness, since if for some cases it showed greater power than the ADF test, for others it disclosed meaningful size distortions. Thus, its performance is not unequivocal when compared to the traditional and easier applying ADF. Anyway, we believe it remains as an ADF complementary test since it is useful in specifying the deterministic trend terms to be included in a series (trend stationary) process. However, it has not a clear function if one only wants to determine a series integration order.

One last striking feature that came about from our trials, concerns the number of replications to be used in the Monte Carlo test. As can be deduced from the reported p-values in the previous section, the increasing number of simulations only increased the p-values precision and had no effect on the final decision. Thus, if not too costly, one may simulate with 1000 replications, otherwise 500 replications will be sufficient.

5. Conclusion

This paper has developed a procedure for bootstrapping the Dickey-Fuller and Ouliaris, Park and, Phillips unit root tests. The procedure allows explicitly for polynomial trends, drift, seasonal dummies and an ARMA error term. Our aim was to develop a procedure with the

[33] These cases are all concentrated in the period from 1966 to 1985.

practical objective of studying the Brazilian economy. We illustrated the procedure using several macroeconomic time series. Because they were designed to be applied in a particular empirical context, the accomplished bootstrap tests provided accurate p-values for the sample tests. In this case, the bootstrap afforded a way of improving on the approximations of asymptotic theory.

The empirical results may also be interpreted as a set of Monte Carlo experiments, because the bootstrap simulations mimicked artificial series designed to access the size and power properties of the ADF and OPP test statistics. Thus, the reported results constitute a study on the size and power distortions of these tests, when applied to the Brazilian series. Our results showed the bootstrap approach performed a lot better than the ADF and the OPP tests regarding the size accuracy. As also has higher power in hypothesis testing unit root against the alternative of stationarity. The results also showed the bootstrap approach performed better when applied to the ADF tests than when applied to the OPP tests, because of its reduced size distortion. Moreover, the Monte Carlo results showed the ADF sample tests have lower size distortion compared to the OPP tests, while the OPP-t has better power properties (than the ADF tests) when a polynomial trend is presented in a series (stationary) DGPs.

From the Monte Carlo results we concluded that the OPP test is not very useful for unit root testing, but it may be used complementarily for the purpose of specifying the deterministic trend terms to be included in a series process.

Our main finding was to evidence that one optimal way for unit root testing is to perform the ADF traditional tests combined with the bootstrap technique.

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Appendix A: Guidelines used for performing the unit root tests to the sample data

Regarding the ADF and the OPP tests, for all series we chose the truncation lag of the polynomial of lagged first difference by combining the kmax criterion proposed by Campbell and Perron (1991), with the information based selection rules (such as the Akaike and Schwartz criteria) suggested by Ng and Perron (1993). For the quarterly series we took the kmax=8 and to the monthly series we chose kmax=18. To the PP test, the lag truncation is chosen by the Bartlett kernel.

As a rule, we looked for specifications that maximized the p-values of the hypotheses of nonexistence of serial correlation of any order.

As diagnostics tests we employed three different type tests for checking the presence of serial correlation in the residuals: the first is the traditional Durbin-Watson test. The second is the portmanteau Ljung-Box Q-statistics based on the estimated auto and crosscorrelations of the first 24 and 36 lags. For the monthly data, we made use of the Q-statistics with the two lags and for the quarterly data, we used only Q(24). The third is the LM-version of the Breusch-Godfrey autocorrelation test of order p, p=1 to 4 for quarterly data, and p=1, 3, 6, 9 and 12 for monthly data.

We also looked into the residual normality property with the Bera-Jarque test and investigated the choice criteria like the Akaike and Schwarz statistics and the equation standard error regression (SER).

For most of the series we got for the unit root tests independent residuals. Nevertheless, in the cases where some serial correlation was marginally present we followed Harvey (1990b) suggestion and complemented the analysis by computing the Kolmogorov-Smirnov statistic (this is a frequency domain test that compares the residuals cumulative periodogram with the theoretical spectral distribution for a white noise). The tests for the dollar premium series (for the period 1966/85) were the only striking exceptions which presented serial correlation (of 12th order) pointed by the LM test (p-values around 2%). The tests for the real balances (ADF for 1966/90) and the monetization ratio (ADF and OPP for 1966/90) series had the hypotheses of no serial correlation of 12th order, marginally accepted with p-values a little bit superior to 5%. In all mentioned cases the Kolmogorov-Smirnov statistic accepted the residuals independence with p-values superior to 10%.

In the table below, we tabulated the critical values for the different specifications of the ADF/PP and OPP t-tests.

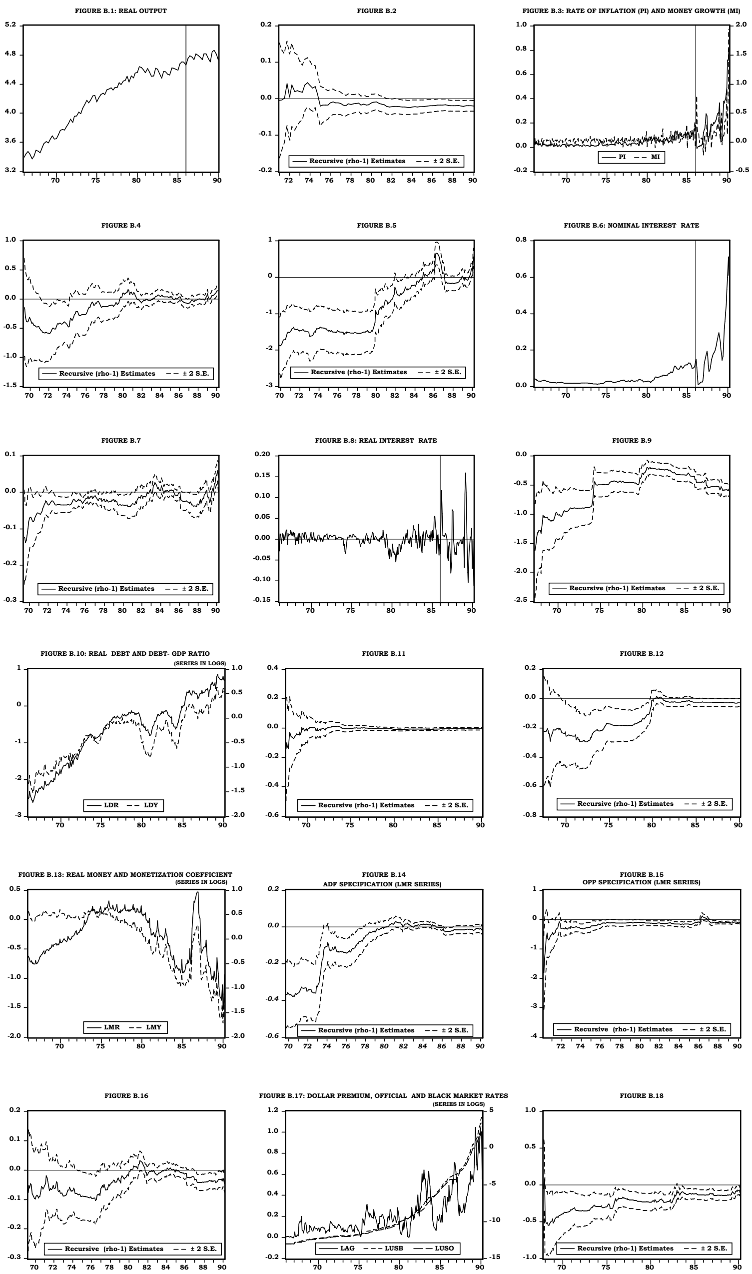
CRITICAL VALUES* FOR THE DICKEY-FULLER, PHLLIPS-PERRON AND OULIARIS, PARK, AND PHILLIPS UNIT ROOT t-TESTS

CRITICAL VALUES	ADF/PP*		OPP			
	t_{ct}	t_{tt}	t_{t2}	t_{t3}	t_{t4}	t_{t5}
0.10	-2.574	-3.139	-3.560	-3.923	-4.252	-4.553
0.05	-2.874	-3.431	-3.828	-4.207	-4.513	-4.825
0.01	-3.461	-4.000	-4.377	-4.740	-5.063	-5.389

SOURCE: MACKINNON (1991) AND OULIARIS, PARK, AND PHILLIPS (1989). * CRITICAL VALUES FOR A SAMPLE OF SIZE N≈230.

Appendix B: Data and recursive estimates

For all series we depicted the data set used in the paper. Aside each series we plotted the recursive estimates of the lagged dependent variable corresponding to the ADF test equation, in order to verify either the mentioned coefficient was stable during the whole period and if its values were near the bounds of a unit root. For the real balances series we also show the recursive coefficient estimates associated with the OPP test. To the dollar series only the premium dollar recursive estimates are presented.



Appendix C: Bootstrap statistics frequency distributions

The table below shows the first fourth moments of all simulated bootstrap statistics and the Bera-Jarque statistic. In the statistics name, the first letter designates the estimated statistics (t or $\hat{\rho}$-statistics) and the next three represent the test name (ADF or OPP).

TABLE C.1: BOOTSTRAP STATISTICS FREQUENCY DISTRIBUTIONS

Series	Period	Statist.	Mean	S.E.	Skew.	Kurt.	BJ*
GDP	66-90	t-ADF	-2.243	0.221	-0.200	3.030	33.511
	66-90	ρ-ADF	-1.391	0.101	-0.174	3.178	31.810
	66-90	t-OPP	-3.518	0.355	-0.181	3.185	34.351
	66-90	ρ-OPP	-50.368	20.618	-2.462	15.164	35880.9
π	66-85	t-ADF	-3.332	0.264	0.050	2.963	0.462
	66-85	ρ-ADF	-80.166	14.674	-0.717	3.517	17.054
	66-85	t-OPP	-5.806	0.405	-0.227	2.859	9.376
	66-85	ρ-OPP	-80.166	14.674	-0.717	3.517	96.780
μ	66-85	t-OPP	-4.844	0.573	-0.103	3.140	2.592
	66-85	ρ-OPP	-168.16	9237.9	-23.197	619.41	15.92E6
LDR	66-85	t-OPP	-5.204	0.268	-0.114	2.997	2.168
	66-85	ρ-OPP	-389.42	6731.0	8.663	307.34	3.87E6
	66-90	t-OPP	-4.377	0.214	-0.147	2.907	3.948
	66-90	ρ-OPP	-84.455	17.153	-0.981	4.817	297.96
LDY	66-85	t-ADF	-2.342	0.118	-0.019	2.675	4.460
	66-85	ρ-ADF	-12.099	1.202	-0.179	2.789	7.196
	66-90	t-ADF	-3.436	0.154	0.557	8.929	1516.3
	66-90	ρ-ADF	-22.817	1.923	0.176	6.074	398.78
	66-90	t-OPP	-3.741	0.164	-0.039	3.126	0.914
	66-90	ρ-OPP	-32.686	3.208	-0.241	3.259	12.449
LMR	66-90	t-ADF	-1.021	0.151	-0.074	3.282	4.240
	66-90	ρ-ADF	-4.254	0.857	-0.353	3.217	22.692
	66-90	t-OPP	-3.673	0.270	-0.135	3.011	3.023
	66-90	ρ-OPP	-204.15	5194.1	2.766	120.23	57.4E4
LMY	66-90	t-ADF	-2.478	0.201	-0.081	3.174	2.348
	66-90	ρ-ADF	-16.337	2.160	-0.286	3.199	15.316
	66-90	t-OPP	-3.538	0.206	-0.067	3.023	0.775
	66-90	ρ-OPP	-31.266	3.624	-0.262	3.103	11.842
LAG	66-90	t-ADF	-2.476	0.281	-0.057	3.354	5.750
	66-90	ρ-ADF	-19.868	3.335	-0.356	3.421	28.430
	66-90	t-OPP	-2.509	0.284	-0.140	3.761	27.415
	66-90	ρ-OPP	-20.307	3.420	-0.438	3.817	59.805

*The Bera-Jarque statistic has $\chi^2(2)$ distribution, the C.V.(0.05) ≈ 5.99.

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