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Abstract

In this paper, a new procedure to test for stationarity is proposed. The new test has three features. First, it is a point optimal test because its derivation is based on the Neyman-Pearson principles. Second, it achieves what is called a near ideal asymptotic rejection (NIAR) profile in the sense of Müller (2005). And third, it provides information about the potential relationship between the empirical size and the previously adopted significance level. The simulations presented at the end of the paper corroborate the asymptotic results and show that its performance improves that of other tests.

Key words: Near integration, point optimal, Monte Carlo, unit root.

JEL classifications: C12, C15, C22

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1.- Introduction

When one considers the model $y_t = \phi y_{t-1} + u_t$ to test for stationarity, the null hypothesis is that $H_0 : \phi < 1$, so the null is a composite hypothesis and not a simple one. This composite character of the null hypothesis creates problems because there are as many sizes as there are values under the null hypothesis and, as a consequence, we do not know what to say about the size of the test.

To avoid this problem, Kwiatkowski et al. (1992) propose using a reparameterization in which the time series is decomposed into the sum of a deterministic term, a pure random walk and a stationary process. The null hypothesis of trend stationarity, which is simple, corresponds to the hypothesis that the variance of the random walk is zero. See also Leybourne and McCabe (1994, 1999) and Lanne and Saikkonen (2003).

In this paper, we follow an alternative route to solve the size indetermination problem. We propose a modified Dickey-Fuller (DF) test, Dickey-Fuller (1979), based on the Neyman-Pearson framework that permits us to inform about the size corresponding to different values of the parameter of interest under the null hypothesis. We show that it achieves what, following Müller (2005), can be called a near ideal asymptotic rejection (NIAR) profile. This NIAR profile refers to a situation in which the empirical size is equal to or smaller than the previously chosen significance level, for any value of the parameter under the null hypothesis, except for a region that depends on the sample size. As the sample size grows, the region becomes smaller. At the same time, the probability of rejecting the null hypotehsis when it is false tends to one.

The rest of the paper is organized as follows. Section 2 introduces the model and applies the Neyman-Pearson Lemma. The limiting behaviour of the test procedure is derived in Section 3. In Section 4, we assess the finite sample performance of the test using Monte-Carlo simulation methods. The main conclusions can be seen in Section 5.

2.- Models and Some Preliminary Results

Consider the following DGP for an observed time series, y_t

$$y_t = d_t + u_t \quad (1)$$

$$u_t = \phi_1 u_{t-1} + ... + \phi_p u_{t-p} + \varepsilon_t \quad (2)$$

where ε_t is i.i.d. N(0, σ^2). The deterministic term, d_t, is specified as $\mathrm{d_t} = 0$ for Model 1, $\mathrm{d_t} = \delta_0$ for Model 2 and $\mathrm{d_t} = \delta_0 + \delta_1 \mathrm{t}$ for Model 3.

For the case with no deterministic terms, (1)-(2) can be written as

$$\Delta y_t = (\phi^* - 1) y_{t-1} + \phi_1^* \Delta y_{t-1} + \ldots + \phi_{p-1}^* \Delta y_{t-p+1} + \varepsilon_t \tag{3}$$

where $\phi^* = \sum_{i=1}^{p} \phi_i$ and $\phi_j^* = -\sum_{i=j+1}^{p} \phi_i$.

The local-to-unity framework we assume in this paper specifies that $\phi^* = 1 - \frac{\mathrm{c}}{\sqrt{\mathrm{T}}}$, for $\mathrm{c} \geq 0$. The reason why we propose to change the traditional local-to-unity approach based on the $1 - \frac{\mathrm{c}}{\mathrm{T}}$ is because we show that, under the new local-to-unity framework based on $\frac{\mathrm{c}}{\sqrt{\mathrm{T}}}$, the test proposed in this paper achieves what, in the Introduction, we called the NIAR profile.

Let us now derive the log-likelihood function corresponding to the process generated by (3). Let $(\phi^* - 1)$ be denoted by $\overline{\phi}$. As can be seen in Hamilton (1994), conditional on *p* initial values, the log-likelihood function for the complete sample is:

$$l(\overline{\phi}) = k_1 - \frac{\sum_{p+1}^{T} (\Delta y_t - \overline{\phi} y_{t-1} - \phi_1^* \Delta y_{t-1} - \ldots.. - \phi_{p-1}^* \Delta y_{t-p+1})^2}{2\sigma^2} \tag{4}$$

where k_1 is a constant that depends on the parameters of the model and on *p* initial values.

Consider a particular value of $\overline{\phi}$ under the null hypothesis, $\overline{\phi}^* < 0$. According to the Neyman-Pearson Lemma, when we are testing $H_0 : \overline{\phi} = \overline{\phi}^*$ against the alternative, $H_1 : \overline{\phi} = 0$, the best testing procedure is the test that rejects the null hypothesis when:

$$l(\overline{\phi}^*) - l(0) < \log h = h^* \tag{5}$$

where h is a positive constant, smaller than one.

Using (4), the left hand side of (5) can be written as:

$$l(\overline{\phi}^*) - l(0) = \tfrac{1}{2\sigma^2}\left[\sum_{p+1}^{T} (\Delta y_t - \phi_1^* \Delta y_{t-1} - \ldots - \phi_{p-1}^* \Delta y_{t-p+1})^2 - \right.$$
$$\left. - \sum_{p+1}^{T} (\Delta y_t - \overline{\phi}^* y_{t-1} - \phi_1^* \Delta y_{t-1} - \ldots - \phi_{p-1}^* \Delta y_{t-p+1})^2 \right] \tag{6}$$

Now, since

$$\begin{aligned} &(\Delta y_t - \overline{\phi}^* y_{t-1} - \phi_1^* \Delta y_{t-1} - \ldots - \phi_{p-1}^* \Delta y_{t-p+1})^2 = \\ &(\Delta y_t - \phi_1^* \Delta y_{t-1} - \ldots - \phi_{p-1}^* \Delta y_{t-p+1})^2 + \overline{\phi}^{*2} y_{t-1}^2 - \\ &-2\overline{\phi}^* y_{t-1} (\Delta y_t - \phi_1^* \Delta y_{t-1} - \ldots - \phi_{p-1}^* \Delta y_{t-p+1}) \end{aligned} \tag{7}$$

we can write (5) as

$$l(\overline{\phi}^*) - l(0) = \frac{1}{2\sigma^2}\left[-\overline{\phi}^{*2} \sum_{p+1}^{T} y_{t-1}^2 + 2\overline{\phi}^* \sum_{p+1}^{T} y_{t-1} (\Delta y_t - \phi_1^* \Delta y_{t-1} - \ldots - \phi_{p-1}^* \Delta y_{t-p+1}) \right] \tag{8}$$

Hence, since $\overline{\phi}^*$ is negative, the inequality in (5) is equivalent to

$$\frac{\sum_{p+1}^{T} y_{t-1} (\Delta y_t - \phi_1^* \Delta y_{t-1} - \ldots\ldots - \phi_{p-1}^* \Delta y_{t-p+1})}{\sigma\, (\sum_{p+1}^{T} y_{t-1}^2)^{\frac{1}{2}}} - \frac{\overline{\phi}^* (\sum_{p+1}^{T} y_{t-1}^2)^{\frac{1}{2}}}{\sigma} > \frac{h^* \sigma}{\overline{\phi}^* (\sum_{p+1}^{T} y_{t-1}^2)^{\frac{1}{2}}} - \frac{\overline{\phi}^* (\sum_{p+1}^{T} y_{t-1}^2)^{\frac{1}{2}}}{2\sigma} \tag{9}$$

Note that the right hand side term of (9) is positive.

3.-Test procedure and its limiting behaviour

Write the matrix form of (3) as:

$$\Delta \mathrm{y} = \mathrm{X}\beta + \varepsilon \tag{10}$$

where $\Delta \mathrm{y}$ is a $\mathrm{Tx1}$ vector of observations of $\Delta \mathrm{y}_\mathrm{t}$, $\mathrm{X} = (\mathrm{x}_1, \mathrm{X}_2)$ where x_1 is a $\mathrm{Tx1}$ vector of observations of $\mathrm{y}_{\mathrm{t}-1}$ and X_2 is the $\mathrm{T} \times (\mathrm{p}-1)$ matrix of observations of $\Delta \mathrm{y}_{\mathrm{t}-1}, \ldots, \Delta \mathrm{y}_{\mathrm{t}-\mathrm{p}+1}$; $\beta' = (\beta_1, \beta_2') = (\overline{\phi}, \phi_1^*, \ldots, \phi_{\mathrm{p}-1}^*)$.

The Maximum-Likelihood (ML) estimators of β and σ^2 are:

$$\hat{\beta} = (\mathrm{X}'\mathrm{X})^{-1}\mathrm{X}'\Delta \mathrm{y} \tag{11}$$

$$\hat{\sigma}^2 = \frac{\hat{\varepsilon}'\hat{\varepsilon}}{\mathrm{T}} \tag{12}$$

where $\hat{\varepsilon} = \Delta \mathrm{y} - \mathrm{X}\hat{\beta}$.

The test-statistic we consider is the following:

$$PONI = \frac{\sum_{p+1}^{T} y_{t-1} (\Delta y_t - \hat{\phi}_1^* \Delta y_{t-1} - \ldots\ldots - \hat{\phi}_{p-1}^* \Delta y_{t-p+1})}{\hat{\sigma}\, (\sum_{p+1}^{T} y_{t-1}^2)^{\frac{1}{2}}} + \frac{c(\sum_{p+1}^{T} y_{t-1}^2 / T)^{\frac{1}{2}}}{\hat{\sigma}} \tag{13}$$

where $\hat{\phi}_1^*,...,\hat{\phi}_{p-1}^*$ are the last *p-1* elements of the vector defined in (11) and *c* is a constant whose value will be fixed in the next section. PONI is the acronym of Point Optimal Near Integration.

The critical region we propose is:

$$\mathrm{PONI} > \mathrm{N}_\alpha$$

where N_α is the critical point of a Standard Normal distribution for a chosen nominal size, α .

In the two next theorems we establish the limiting behaviour of the probability of rejecting the null hypothesis assuming, first, that the data have been generated by the null hypothesis (Theorem 1) and then, that the data have been generated by the alternative hypothesis (Theorem 2).

THEOREM 1: Suppose that the DGP is given by $\phi^*_{\mathrm{TRUE}} = 1 - \frac{\mathrm{c}_{\mathrm{TRUE}}}{\sqrt{\mathrm{T}}}$; then, asymptotically we have:

$$\begin{aligned} &\Pr\mathrm{ob}(\mathrm{PONI} > \mathrm{N}_\alpha) < \alpha \quad \text{if } \mathrm{c}_{\mathrm{TRUE}} > \mathrm{c} \\ &\Pr\mathrm{ob}(\mathrm{PONI} > \mathrm{N}_\alpha) = \alpha \quad \text{if } \mathrm{c}_{\mathrm{TRUE}} = \mathrm{c} \\ &\Pr\mathrm{ob}(\mathrm{PONI} > \mathrm{N}_\alpha) > \alpha \quad \text{if } \mathrm{c}_{\mathrm{TRUE}} < \mathrm{c} \end{aligned} \tag{14}$$

where *c* is the value assumed to define the PONI statistic in (13).

Proof: See Appendix.

THEOREM 2: If the series has a unit root, that is, if $\phi^* = 1$, then:

$$\lim \Pr\mathrm{ob}(\mathrm{PONI} > \mathrm{N}_\varepsilon) = 1 \tag{15}$$

Proof: See Appendix.

REMARK 1: In the introduction, we mentioned two reasons to justify the proposal of the new test. The first one was that the new test achieves the NIAR profile. This has been shown in the two theorems just presented. The second point was that this new method provides information about the size of the test, corresponding to the different values under the null hypothesis. This can be seen examining the results of Theorem 1. If the true value of the parameter of interest equals the value chosen to define the PONI statistic, that is, if $c_{TRUE} = c$, then the test has a size equal to the chosen theoretical size.

If the true value of the parameter of interest is smaller than the chosen value of this parameter, that is, if $c_{TRUE} > c$, then the test has a size smaller than the theoretical chosen size. Finally, if the true value of the parameter of interest is greater than the chosen value, that is, if $c_{TRUE} < c$, then the test has a size greater than the chosen theoretical size.

To conclude, we can say that, for values of the parameter of interest closer to one than the value chosen to define the test, the size will be superior to the selected theoretical size. If the value coincides, the size is this selected value. And, if the true value of the parameter is smaller than the value chosen to define the test, the size will be smaller than the theoretical size. Besides, it is clear that, when the parameters coincide, the PONI test provides the best trade-off between size and power because of the Neyman-Pearson Lemma.

When one considers models with deterministic terms, an appropriate detrending procedure of the data is needed.

The two best known approaches proposed in the literature to extract the deterministic terms are the OLS and the GLS detrending procedure proposed in Elliot et al. (1996) and Ng and Perron (2001). Some Monte Carlo simulations show the scarce capacity of the testing procedures to discriminate between the null and the alternative hypotheses when the model has a constant and a linear trend.

Consider the model with a constant in which we use the OLS detrending procedure:

$$\hat{u}_t = y_t - \hat{\delta}_0 = u_t - \frac{\sum_{t=1}^{T} u_t}{T} \text{ because } \hat{\delta}_0 = \frac{\sum_{t=1}^{T} y_t}{T} = \delta_0 + \frac{\sum_{t=1}^{T} u_t}{T}$$

Assume that the series has a unit root so that $\sum_{t=1}^{T} u_t$ is $O_p(T^{3/2})$. Thus, $\hat{\delta}_0 - \delta_0$ is $O_p(T^{1/2})$. We have:

$$\sum_{t=1}^{T} \hat{u}_t^2 = \sum_{t=1}^{T} u_t^2 - T\overline{u}^2$$

Both terms on the right-hand side are $O_p(T^2)$ and it is clear that $\sum_{t=1}^{T} \hat{u}_t^2$ is very different from $\sum_{t=1}^{T} u_t^2$. The consequence is that the performance of the test based on the OLS detrending procedure will be different to that of the case where the true values of the parameters are known, even asymptotically.

Consider, instead, the following detrending procedure:

$$\hat{u}_t^* = y_t - 2\hat{\delta}_0 = -\delta_0 + u_t - 2\overline{u}$$

In this case, we have:

$$\sum_{t=1}^{T} \hat{u}_t^{*2} = T\delta_0^2 + \sum_{t=1}^{T} u_t^2 + 4T\overline{u}^2 + 2\delta_0 \sum_{t=1}^{T} u_t + 4\delta_0 T\overline{u} - 4\overline{u}\sum_{t=1}^{T} u_t$$

Since all terms with δ_0 are of an order smaller than $O_p(T^2)$ and $4T\overline{u}^2 = 4\overline{u}\sum_{t=1}^{T} u_t$ we obtain that, asymptotically,

$$\sum_{t=1}^{T} \hat{u}_t^{*2} \cong \sum_{t=1}^{T} u_t{}^2$$

We have carried out simulation experiments to compare the performance of the PONI test considering the three detrending procedures we have just mentioned. The results can be seen in Table 1 (MODEL 2), where the results for ϕ = 1 are the power of the test and for values of ϕ smaller than 1 are the size of the test.

We have extended these results to the case where the model has a linear trend. In this case, we have considered the following detrending procedure:

$$\hat{u}_t^* = y_t - \lambda\hat{\delta}_0 - \hat{\delta}_1 t$$

where λ is a constant between 1 and 2. From the Monte Carlo experiments we have carried out, we conclude that the best performance of the test corresponds to a value of λ between 1.3 and 1.35 depending on the sample size. When this sample size is 200 or smaller we propose to use $\lambda = 1.3$ while, if the sample size is larger than 200, then we propose to use $\lambda = 1.35$. The results can be seen in Table 1 (MODEL 3).

It is clear that the test based on the new detrending procedure outperforms the other two procedures.

Table 1: Size and power of the PONI test for different detrending procedures for model 2 and 3

		MODEL 2			MODEL 3		
TM	ϕ	**OLS**	**GLS**	**NEW**	**OLS**	**GLS**	**NEW**
50	0.8	0.00	0.00	0.00	0.00	0.00	0.01
50	0.9	0.00	0.00	0.02	0.00	0.00	0.05
50	0.95	0.01	0.03	0.14	0.00	0.00	0.11
50	0.97	0.03	0.05	0.26	0.00	0.00	0.16
50	0.98	0.03	0.07	0.34	0.00	0.00	0.19
50	**1**	**0.06**	**0.13**	**0.55**	**0.00**	**0.00**	**0.25**
100	0.8	0.00	0.00	0.00	0.00	0.00	0.00
100	0.9	0.00	0.00	0.00	0.00	0.00	0.00
100	0.95	0.00	0.06	0.06	0.00	0.00	0.05
100	0.97	0.03	0.14	0.19	0.00	0.00	0.10
100	0.98	0.05	0.20	0.31	0.00	0.00	0.14
100	**1**	**0.16**	**0.41**	**0.63**	**0.00**	**0.01**	**0.23**
500	0.8	0.00	0.00	0.00	0.00	0.00	0.00
500	0.9	0.00	0.00	0.00	0.00	0.00	0.00
500	0.95	0.00	0.05	0.00	0.00	0.00	0.00
500	0.97	0.00	0.14	0.00	0.00	0.00	0.02
500	0.98	0.03	0.28	0.09	0.01	0.05	0.12
500	**1**	**0.62**	**0.88**	**0.89**	**0.21**	**0.44**	**0.53**

4.- Monte Carlo experiments

In this section, we provide Monte Carlo simulation results to illustrate the finite sample performance of the PONI test in relation to the modified KPSS test proposed by Sul et al. (2005).

We start by considering the following DGP:

$$y_t = \delta_0 + \delta_1 t + u_t, \quad t = 1, 2, ..., T,$$
$$u_t = \phi u_{t-1} + \varepsilon_t, \quad t = 2, ..., T, \quad \varepsilon_t \sim iidN(0,1)$$

We have results for three different models depending on the form adopted by the deterministic terms: Model 1: when $\delta_0 = 0$ and $\delta_1 = 0$; Model 2: when $\delta_0 \neq 0$ and Model 3: when $\delta_0 \neq 0$ and $\delta_1 \neq 0$. The values considered for ϕ are $\phi = 0.8, 0.9, 0.95, 0.97, 0.98, 1$

For the particular case of Model 1 without lags, the PONI statistic we use is:

$$PONI = \frac{\sum_{p+1}^{T} y_{t-1}\Delta y_t}{\tilde{\sigma}\left(\sum_{p+1}^{T} y_{t-1}^2\right)^{1/2}} + \frac{c\left(\sum_{p+1}^{T} y_{t-1}^2 / T\right)^{1/2}}{\tilde{\sigma}}$$

where $\tilde{\sigma}^2$ is the maximum-likelihood estimator of the variance of the disturbance in the following regression $\Delta y_t = \beta_1 y_{t-1} + u_t$; c is set as 0.5 so that a theoretical size of 5% is guaranteed for a value of the null parameter, ϕ, equal to $1 - \frac{0.5}{\sqrt{T}}$ given T. For example, a 5% significance level corresponds to $\phi = 0.95$ when T=100.

For Model 1, the modified KPSS test of Sul et al. (2005) is the standard KPSS statistic constructed as:

$$KPSS = \frac{T^{-2}\sum_{t=2}^{T}\left(\sum_{i=2}^{t}\hat{u}_i\right)^2}{\hat{\omega}^2}$$

where $\hat{\omega}^2$ is any standard long-run variance estimator of the form:

$$\hat{\omega}^2 = \hat{\gamma}_0 + 2\sum_{j=1}^{T-1}\lambda(j/l)\hat{\gamma}_j, \quad \hat{\gamma}_j = T^{-1}\sum_{t=j+1}^{T}\hat{u}_t\hat{u}_{t-j},$$

where $\lambda(.)$ is a kernel function. We have used the Quadratic spectral window (as depicted in Sul et al. (2005)) and the automatic method depicted in Andrews (1991) to estimate the bandwidth. And, as Carrion-i-Silvestre and Sansó (2006) propose, we use the modification of Sul et al. (2005) to control the estimated bandwidth, and to avoid the inconsistency of the KPSS detected by Choi and Ahn (1995,1999) and Kurozumi (2002). 5% critical values of KPSS are known to be 1.656 for Model 1, 0.463 for Model 2 and 0.146 for Model 3. These asymptotic critical values can be found in Kwiatkowski et al. (1992) for Models 2 and 3, and Hobijn et al. (1998) for Model 1.

If data are generated using Model 2 or 3, then the PONI test is the same after eliminating the deterministic terms from y_t using the new procedure mentioned in Section 3.

In the experiments, we have considered $\delta_0 = 1$ for Model 2, and $\delta_0 = 1,\ \delta_1 = 1$ for Model 3. We have studied the behaviour of the tests for three different sample sizes, *T = 50, 100, 500,* and all the experiments are based on *10,000* replications.

Since the distribution of the PONI test depends on the initial value of the autoregressive process, in order to assess the robustness of the test procedure, in the experiments we use different values of $\varepsilon_1 = (0,1,3,5)$. As the results are quite robust to the initial values considered, in Table 3 we present the results corresponding to $\varepsilon_1 = 0$. Those corresponding to other values are available from the authors upon request.

The simulation results are presented in Table 2. It gives the rejection rate of the *PONI* and the modified *KPSS* tests corresponding to each sample size and for each value of the autoregressive parameter for a 5% nominal significance level, for each model. Note that when $\phi = 1$, the rejection rate is the power of the corresponding test.

With respect to the performance of the *PONI* test, the main conclusion that can be extracted from the results presented in Table 2 for Model 1 is that, the empirical size corresponding to values of ϕ equal to or smaller than $1-\frac{0.5}{\sqrt{T}}$ is smaller than the theoretical 5%. This implies that the PONI procedure guarantees a 5% empirical level for values of ϕ closer to 1, as the sample size grows. Note, in this respect, that when *T = 500* the empirical size corresponding to $\phi = 0.98$ is *0.06*.

When we compare the performance of the *PONI* test with that of the modified *KPS*S test, we find that, for moderate size samples, say *T = 100*, the trade-off between size and power of the two tests is very similar. However, when the sample size grows, say *T = 500*, then the *PONI* test clearly outperforms the modified *KPSS* test. For values close to the alternative hypothesis, the rejection rate corresponding to the *PONI* test is much smaller than that of the modified *KPSS* test and the power is similar for both tests.

The results corresponding to Model 2, are very similar to those for Model 1. But, even for a sample size of 100, the PONI test achieves a higher power and a lower size than the KPSS test.

For Model 3, the trade-off between size and power of the PONI test is better than that for the modified KPSS test.

In order to analyze the robustness of the PONI test to different distributions of the perturbance term of the DGP, ε_t , we present, in Table 3, the results of a Monte Carlo experiment where we have considered the same DGP as in the previous experiment, for Model 1, but now the disturbance term follows alternative distributions to the Standard Normal. In particular, we have considered a Normal distribution with variance equal to 4, N(0,4), a t-Student of 2 degrees of freedom, t(2,) and a Uniform distribution, U(-2,2). It can be seen for Model 1 that the PONI test is quite robust to different assumptions about the probability distribution of ε_t; for models 2 and 3, the test is also robust, and the results are available upon request.

Table 2: Size and power of the tests

		MODEL 1		MODEL 2		MODEL 3	
TM	ϕ	**PONI**	**KPSS**	**PONI**	**KPSS**	**PONI**	**KPSS**
50	0.8	0.00	0.02	0.00	0.02	0.01	0.02
50	0.9	0.01	0.16	0.03	0.07	0.05	0.02
50	0.95	0.11	0.40	0.14	0.17	0.11	0.03
50	0.97	0.26	0.54	0.26	0.23	0.16	0.03
50	0.98	0.36	0.62	0.34	0.25	0.19	0.03
50	**1**	**0.61**	**0.76**	**0.54**	**0.34**	**0.25**	**0.03**
100	0.8	0.00	0.02	0.00	0.03	0.00	0.03
100	0.9	0.00	0.06	0.00	0.04	0.01	0.02
100	0.95	0.04	0.30	0.06	0.19	0.05	0.07
100	0.97	0.18	0.49	0.19	0.31	0.10	0.12
100	0.98	0.32	0.59	0.31	0.37	0.15	0.14
100	**1**	**0.68**	**0.80**	**0.63**	**0.54**	**0.23**	**0.18**
500	0.8	0.00	0.04	0.00	0.04	0.00	0.04
500	0.9	0.00	0.04	0.00	0.04	0.00	0.03
500	0.95	0.00	0.04	0.00	0.04	0.00	0.03
500	0.97	0.00	0.18	0.00	0.17	0.02	0.14
500	0.98	0.06	0.38	0.09	0.36	0.12	0.32
500	**1**	**0.90**	**0.92**	**0.89**	**0.88**	**0.53**	**0.72**

Table 3: Size and power of the PONI test for different distributions of ε_i. Model 1.

TM	ϕ	**N(0,1)**	**N(0,4)**	**t(2)**	**U(-2,2)**
50	**0.8**	0.00	0.00	0.00	0.00
50	**0.9**	0.01	0.03	0.02	0.03
50	**0.95**	0.11	0.09	0.09	0.09
50	**0.97**	0.26	0.24	0.26	0.24
50	**0.98**	0.36	0.34	0.37	0.34
50	**1**	**0.61**	**0.59**	**0.62**	**0.58**
100	**0.8**	0.00	0.00	0.00	0.00
100	**0.9**	0.00	0.00	0.00	0.00
100	**0.95**	0.04	0.03	0.02	0.03
100	**0.97**	0.18	0.16	0.16	0.16
100	**0.98**	0.32	0.31	0.32	0.32
100	**1**	**0.68**	**0.67**	**0.70**	**0.68**
500	**0.8**	0.00	0.00	0.00	0.00
500	**0.9**	0.00	0.00	0.00	0.00
500	**0.95**	0.00	0.00	0.00	0.00
500	**0.97**	0.00	0.00	0.00	0.00
500	**0.98**	0.06	0.06	0.06	0.06
500	**1**	**0.90**	**0.90**	**0.91**	**0.91**

5.- Conclusions

In this paper, we have derived a new procedure to test the null hypothesis of stationarity against the alternative that the series has a unit root. Since the null we have considered is that of stationarity, we have adopted a local-to-unity setting based on $1-\frac{c}{\sqrt{T}}$ instead of $1-\frac{c}{T}$.

The new procedure has been derived using the Neyman-Pearson approach, after having chosen a particular value under the null hypothesis. Thus, the test we have proposed is a point optimal test.

Then, we have derived the limiting behaviour of the test under both hypotheses, stationarity and unit root. We have shown that the new test achieves what we have called the NIAR profile. This means that the size of the new test is equal to or smaller than the size chosen to define the critical region except for the region corresponding to the values of the parameter of interest which are closer to one than the value chosen to define the test. Using the notation of Section 3, the latter region is determined by $c > c_{TRUE}$. We have shown that this region is a decreasing function of the sample size. We have also recommended the use of a new procedure to demean and detrend the series and we have shown the improvement achieved with this new procedure.

The simulation results we have presented in the last section confirm the results derived in previous sections and allow us to conclude that the performance of the new test compares favorably in terms of the trade-off between size and power to that of others tests available in the literature.

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APPENDIX

Before presenting the proofs of the two theorems, let us consider the following proposition:

Proposition A.1: Let y_t be generated by $y_t = \phi y_{t-1} + u_t$ with $\phi = 1 - cT^{-\frac{1}{2}}$.

Assume that $u_t = \Psi(L)\varepsilon_t$ where $\varepsilon_t \sim i.i.d.(0,\sigma_\varepsilon^2)$ and $\Psi(L) = \sum_{i=0}^{\infty}\psi_i L^i$. Let $y_t^* = b_T^{\frac{1}{2}} y_t$ with $b_T = T^{-\frac{1}{2}}$ and let ω^2 be the long-run variance of u_t, $\omega^2 = \sigma_\varepsilon^2\left(\sum_{i=1}^{\infty}\psi_i\right)^2$. Then, as $T \to \infty$:

(i) $T^{-\frac{1}{4}} y_t = y_t^* \longrightarrow N\left(0, \frac{\omega^2}{2c}\right)$.

(ii) $T^{-\frac{3}{2}}\sum_{t=1}^{T} y_t^2 = T^{-1}\sum_{t=1}^{T} y_t^{*2} \to \frac{\omega^2}{2c}$

(iii) $T^{-1}\sum_{t=1}^{T} y_{t-1}u_t = T^{-3/4}\sum_{t=1}^{T} y_{t-1}^* u_t \to \frac{\omega^2}{2} - \frac{\sigma^2}{2}$

(iv) $T^{-\frac{3}{4}}\sum_{t=1}^{T} y_{t-1}\varepsilon_t = T^{-\frac{1}{2}}\sum_{t=1}^{T} y_{t-1}^*\varepsilon_t \longrightarrow N\left(0, \frac{\sigma_\varepsilon^2\omega^2}{2c}\right)$

$\to$ denotes convergence in probability.

Proof:

(i) We can write y_t^* as

$$y_t^* = b_T^{\frac{1}{2}}\left\{u_t + \phi u_{t-1} + \phi^2 u_{t-2} + ... + \phi^h u_{t-h}\right\} =$$
$$= b_T^{\frac{1}{2}}\left\{\psi(L)\varepsilon_t + \phi\psi(L)\varepsilon_{t-1} + \phi^2\psi(L)\varepsilon_{t-2} + \cdots + \phi^h\psi(L)\varepsilon_{t-h}\right\} =$$
$$= b_T^{\frac{1}{2}}\left\{\psi_0\varepsilon_t + (\psi_1 + \phi\psi_0)\varepsilon_{t-1} + (\psi_2 + \phi\psi_1 + \phi^2\psi_0)\varepsilon_{t-2} + ...\right\} = \psi^*(L)\varepsilon_t$$

For $h \to \infty$, where $\psi^*(L) = \left\{\psi_0^* + \psi_1^* L + \psi_2^* L^2 + ...\right\}$ and $\psi_i^* = b_T^{\frac{1}{2}}\left(\psi_i + \phi\psi_{i-1} + ... + \phi^i\psi_0\right)$.

Note that:

$$\sum_{i=0}^{\infty}\psi_i^{*2} = b_T\left(1+\phi^2+\phi^4+\cdots+\phi^{2h}\right)\left[\sum_{i=0}^{\infty}\psi_i^2 + 2\sum_{i=0}^{\infty}\sum_{j=1}^{\infty}\phi^j\psi_i\psi_{i+j}\right]$$

Now, since:

$$b_T\left(1+\phi^2+....+\phi^{2h}\right) = b_T\,\frac{1-\left(1-cb_T\right)^{2h}}{1-\left(1-cb_T\right)^2} =$$

$$= b_T\,\frac{1-\exp\left\{-2c\left[Tr\right]b_T\right\}}{2cb_T - c^2b_T^2} \to \frac{1}{2c}$$

and $\phi^j = 1+o(1)$ we have:

$$\sum_{i=0}^{\infty}\psi_i^{*2} \to \frac{1}{2c}\left(\sum_{i=0}^{\infty}\psi_i\right)^2$$

To study the asymptotic distribution of y_t^*, let us consider the limiting behaviour of its moment generating function:

$$E\left[e^{sy_t^*}\right] = \prod_{i=0}^{\infty}\left[1+\frac{\sigma_\varepsilon{}^2\psi_i^{*2}}{2}s^2 + o\left(\psi_i^{*2}s^2\right)\right] =$$

$$= \exp\left\{\frac{\sigma_\varepsilon{}^2s^2}{2}\sum_{i=0}^{\infty}\psi_i^{*2}\right\} \to \exp\left\{\frac{\sigma_\varepsilon{}^2s^2}{2}\,\frac{\left(\sum_{i=1}^{\infty}\psi_i\right)^2}{2c}\right\}$$

Thus, y_t^* behaves asymptotically as a normal variable with zero mean and a variance equal to $\frac{\omega^2}{2c}$.

(ii) Using the Central Mapping Theorem we have:

$$\frac{1}{T}\sum_{t=1}^{T}y_t^{*2} = \frac{1}{T^{3/2}}\sum_{t=1}^{T}y_t^2 \to \frac{\omega^2}{2c}$$

(iii) We can write:

$$y_t^* = \phi y_{t-1}^* + b_T^{1/2}u_t\,,$$

so that:

$$\sum_{t=1}^{T}y_t^{*2} = \phi^2\sum_{t=1}^{T}y_{t-1}^{*2} + b_T\sum_{t=1}^{T}u_t^2 + 2\phi b_T^{1/2}\sum_{t=1}^{T}y_{t-1}^*u_t$$

Now, since $\phi^2 = 1+\frac{c^2}{T}-\frac{2c}{\sqrt{T}}$ and $\phi = 1+o(1)$, we obtain:

$$\sum_{t=1}^{T} y_{t-1}^{*}u_t = \frac{y_T^{*2}}{2b_T^{1/2}} - \frac{c^2}{2Tb_T^{1/2}}\sum_{t=1}^{T} y_{t-1}^{*2} + \frac{2c}{2\sqrt{T}b_T^{1/2}}\sum_{t=1}^{T} y_{t-1}^{*2} -$$

$$-\frac{b_T}{2b_T^{1/2}}\sum_{t=1}^{T} u_t^2$$

The first term on the right hand side is $O_P(T^{1/4})$, the second is $O(T^{1/4})$, and the last two terms are $O(T^{3/4})$. The result is obtained using (ii) and $\frac{1}{T}\sum_{t=1}^{T} u_t^2 \to \sigma^2$.

(iv) Define $x_t = y_{t-1}^{*}\varepsilon_t$. Note that x_t is martingale difference sequence. Next, $\frac{1}{\sqrt{T}}\sum_{t=1}^{T} y_{t-1}^{*}\varepsilon_t$ can be written as:

$$\frac{1}{\sqrt{T}}\sum_{t=1}^{T} y_{t-1}^{*}\varepsilon_t = \frac{1}{\sqrt{T}}\left(\psi_0^{*}\varepsilon_0 + \psi_1^{*}\varepsilon_1 + ...\right)\varepsilon_1 +$$

$$+\frac{1}{\sqrt{T}}\left(\psi_0^{*}\varepsilon_1 + \psi_1^{*}\varepsilon_0 + ...\right)\varepsilon_2 + ... + \frac{1}{\sqrt{T}}\left(\psi_0^{*}\varepsilon_{T-1} + \psi_1^{*}\varepsilon_{T-2} + ...\right)\varepsilon_T =$$

$$= \frac{1}{\sqrt{T}}\sum_{i=0}^{\infty} \psi_i^{*}\varepsilon_{T-i}^{*}$$

where $\varepsilon_{T-i}^{*} = \sum_{t=1}^{T} \varepsilon_t \varepsilon_{t-i-1}$

Note that $E\varepsilon_{T-i}^{*} = 0, Var(\varepsilon_{T-i}^{*}) = T\sigma_\varepsilon^4$ and they are independent from each other. Since $\sum_{i=0}^{\infty} \psi_i^{*2} \to \frac{1}{2c}$, we obtain:

$$E\left[e^{s\frac{1}{\sqrt{T}}\sum_{t=1}^{T} y_{t-1}^{*}\varepsilon_t}\right] = \prod_{i=0}^{T}\left[1 + \frac{s^2}{2}\sigma_\varepsilon^4 \psi_i^{*2} + o\left(\psi_i^{*2}s^2\right)\right] =$$

$$= \exp\left\{\frac{s^2\sigma_\varepsilon^4}{2}\sum_{i=0}^{\infty} \psi_i^{*2}\right\} \to \exp\left\{\frac{s^2\sigma_\varepsilon^4}{2}\frac{\left(\sum_{i=0}^{\infty} \psi_i\right)^2}{2c}\right\}$$

and the result follows.

Proof of Theorem 1:

The numerator of the first term of the PONI statistic can be written as:

$$\sum_{p+1}^{T} y_{t-1}\left[\overline{\phi}^{*} y_{t-1} + (\phi_1^{*} - \hat{\phi}_1^{*})\Delta y_{t-1} + ... + (\phi_{p-1}^{*} - \hat{\phi}_{p-1}^{*})\Delta y_{t-p+1} + \varepsilon_t\right]$$

since each element $(\phi_i^{*} - \hat{\phi}_i^{*})$ is $O_p(T^{-1/2})$, the statistic in (13) asymptotically is equivalent to:

$$\frac{\bar{\phi}^* \sum_{p+1}^{T} y_{t-1}^2}{\sigma\left(\sum_{p+1}^{T} y_{t-1}^2\right)^{1/2}} + \frac{\sum_{p+1}^{T} y_{t-1}\varepsilon_t}{\sigma\left(\sum_{p+1}^{T} y_{t-1}^2\right)^{1/2}} + \frac{c\left(\sum_{p+1}^{T} y_{t-1}^2 / T\right)^{1/2}}{\sigma}$$

or, equivalently

$$\frac{\sum_{p+1}^{T} y_{t-1}\varepsilon_t / T^{3/4}}{\sigma\left(\sum_{p+1}^{T} y_{t-1}^2 / T^{3/2}\right)^{1/2}} + \frac{(c - c_{TRUE})T^{1/4}\left(\sum_{p+1}^{T} y_{t-1}^2 / T^{3/2}\right)^{1/2}}{\sigma}$$

Note that, as it is shown in Proposition A.1 (iv), $\mathrm{T}^{-3/4}\sum_{p+1}^{T} y_{t-1}\varepsilon_t$ has a well defined limiting distribution and $\sum_{p+1}^{T} y_{t-1}^2$ is $O_p(T^{3/2})$. The result follows because the first term converges to a well-defined limit.

Proof of Theorem 2:

The OLS of β in (10) can be written as:

$$\hat{\beta} = \begin{pmatrix} 0 \\ \beta_2 \end{pmatrix} + (X'X)^{-1}X'\varepsilon$$

so that,

$$\hat{\beta} - \begin{pmatrix} 0 \\ \beta_2 \end{pmatrix} = \gamma_T^{-1}(\gamma_T^{-1}X'X\gamma_T^{-1})^{-1}\gamma_T^{-1}X'\varepsilon$$

where γ_T is the following normalization matrix:

$$\gamma_T = \begin{pmatrix} T & 0 \\ 0 & \gamma_T^* \end{pmatrix} = \begin{pmatrix} T & 0 & \dots & 0 \\ 0 & T^{1/2} & \dots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & T^{1/2} \end{pmatrix}$$

It is seen that each element $(\hat{\phi}_i^* - \phi_i^*)$ is $O_p(T^{-1/2})$. Under the alternative hypothesis, the first term of the numerator of the PONI statistic can be written as:

$$\sum_{p+1}^{T} y_{t-1}\left[(\phi_1^* - \hat{\phi}_1^*)\Delta y_{t-1} + \ldots + (\phi_{p-1}^* - \hat{\phi}_{p-1}^*)\Delta y_{t-p+1} + \varepsilon_t\right]$$

This expression, asymptotically, is dominated by $\sum_{p+1}^{T} y_{t-1}\varepsilon_t$ that is $O_p(T)$ because, as we commented before, the probability order of each element of $(\phi_1^* - \hat{\phi}_1^*),...,(\phi_{p-1}^* - \hat{\phi}_{p-1}^*)$ is $O_p(T^{-1/2})$ and because $\sum_{p+1}^{T} y_{t-1}\Delta y_{t-j}$ needs to be divided by T in order to converge.

Using this result, we can say that the PONI test, asymptotically, follows the same limiting behaviour as:

$$\frac{\sum_{p+1}^{T} y_{t-1}\varepsilon_t / T}{\sigma(\sum_{p+1}^{T} y_{t-1}^2 / T^2)^{\frac{1}{2}}} + \frac{\sqrt{T}c(\sum_{p+1}^{T} y_{t-1}^2 / T^2)^{\frac{1}{2}}}{\sigma}$$

The first term of this expression is $O_p(1)$ and it is the t-ratio proposed by Dickey-Fuller (1979) to test for the presence of a unit root. The result follows because when the model has a unit root, $\sum_{p+1}^{T} y_{t-1}^2$ is $O_p(T^2)$ and thus the second term tends to infinity as $T \to \infty$. Note that Theorem 2 can be stated by saying that if the alternative hypothesis holds, the PONI statistic is $O_p(T^{1/2})$.

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