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Is there Really a Unit Root in the Inflation Rate? More Evidence from Panel Data Models*

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Abstract

Time series unit root evidence suggests that inflation is nonstationary. By contrast, when using more powerful panel unit root tests, Culver and Papell (1997) find that inflation is stationary. In this paper, we test the robustness of this result by applying a battery of recent panel unit root tests. The results suggest that the stationarity of inflation holds even after controlling for cross-sectional dependence and structural change.

JEL Classification: C32; C33; E31.
Keywords: Unit Root; Inflation; Cross-Sectional Dependence; Structural Change.

1 Introduction

In recent years, there has been a great deal of research focusing on the persistence of inflation. Some of the most recent advances within this field include Cook (2005), Charemza *et al.* (2005), Österholm (2004) and Holmes (2002), to mention a few. This is an important and relevant topic because inflation is typically regarded as a key variable in many economic models, whose validity hinges critically on whether inflation is I(0) or not. There is also a large body of empirical work based cointegration that relies on inflation being I(1).

The empirical evidence of the stationarity of inflation is quite mixed. For example, Johansen (1992) finds that log prices are I(2) and that inflation therefore is I(1), while Rose (1988) finds US inflation to be I(0). Recently, Culver and Papell (1997) apply time series unit root and stationarity tests to 13 OECD countries and find overwhelming evidence in favor of inflation being I(0). However, when accounting for the low power of the individual time series tests by using the Levin and Lin (1992) panel data unit root test, the authors strongly reject the null hypothesis of unit root for the panel as a whole.

Unfortunately, the Levin and Lin (1992) test suffers from several shortcomings that make the conclusions of Culver and Papell (1997) somewhat questionable. First,

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it is based on the assumption that the cross-sectional units are independent of each other, which is of course highly unlikely to hold because of the strong comovement of prices across countries. Second, even though long span inflation rates are prone to structural change, the test does not allow for the possibility of such change. Third, the test requires that the autoregressive behavior of the individual cross-sections are exactly equal under both the null and alternative hypotheses, which is clearly an overly restrictive assumption. Fourth, the test is not applicable in the presence of heteroskedasticty.

The potential consequences of violating any of these assumptions are well known. In particular, it has been shown that a violation can cause the test to be biased towards stationarity. Therefore, although Culver and Papell (1997) account for the fourth critique by using bootstrapped critical values, it is still perfectly possible that their results do not reflect the actual process generating inflation but rather a small-sample bias on behalf of the test employed.

In this paper, we use the same data as Culver and Papell (1997) to test the robustness of their conclusions. In so doing, we employ several recently developed panel data unit root tests, some of which are very general and permit both cross-sectional dependency and structural breaks, as well as disparate autoregressive behavior and heteroskedasticity, which are likely to be highly relevant features when testing the inflation data. Our results indicate that the stationarity of inflation holds even after controlling for cross-sectional dependence and structural change.

The rest of the paper is organized as follows. Section 2 provides a brief description of the tests used in this paper, while Section 3 presents the empirical results. Some concluding remarks appear in Section 4.

2 Panel Unit Root Tests

To test the results of Culver and Papell (1997), we employ a battery panel unit root tests, which are all based on the assumption that inflation, denoted π_{it}, can be given the following autoregressive representation

$$\Delta\pi_{it} = \alpha_i + \phi_i \pi_{it-1} + \sum_{j=1}^{p} \gamma_{ij} \Delta\pi_{it-j} + u_{it}. \tag{1}$$

where $t = 1, ..., T$ and $i = 1, ..., N$ indexes the time series and cross-sectional units, respectively. The tests that we use can be classified into three groups depending on whether they allow for structural breaks or cross-sectional correlation. The first group consists of the Z_{tbar} and $Z_{\tilde{t}bar}$ tests of Im *et al.* (2003), the t_δ^* test of Levin *et al.* (2002), the $\hat{\varphi}_{LSDV}$ test of Harris and Tzavalis (1999) and the $\Gamma_{LM}(p)$ test of Im *et al.* (2005). These tests are based on the assumptions of no cross-sectional dependence and no structural breaks.

Like many other macroeconomic variables, inflation usually exhibit strong comovements across countries. The second group of tests allows for such cross-sectional dependence by assuming that (1) admits to a common factor representation. It includes the G_{ols}^{++}, P_m and Z tests of Phillip and Sul (2003), the P_e^c test of Bai and Ng (2004), the t_a and t_b tests of Moon and Perron (2004) and the Pesaran (2003) IPS^* test.

A number of papers find that inflation is subject to structural change. For example, Garcia and Perron (1996) find evidence of structural breaks in both the mean and variance of US inflation between 1961 and 1986. The third group of tests allows for such breaks and consists of the $\Gamma_{LM}^{B}(p)$ test of Im *et al.* (2005) and the $LM(\lambda)$ test of Carrion-i-Silvestre *et al.* (2005). The $\Gamma_{LM}^{B}(p)$ test allows for a single break, whereas $LM(\lambda)$ allows for an unknown number of breaks in the level of each series.

All tests except $LM(\lambda)$ take nonstationarity as the null hypothesis, and all tests except t_{δ}^{*} and $\hat{\varphi}_{LSDV}$ permit ϕ_i to differ across i, which make them more general than the Levin and Lin (1992) test used by Culver and Papell (1997). Except for IPS^{*}, each statistic is normally distributed under the null hypothesis. The direction of the divergence under the alternative hypothesis determines whether we should use the right or left tail of the normal distribution to reject the null. The Z_{tbar}, $Z_{\tilde{t}bar}$, t_{δ}^{*}, $\hat{\varphi}_{LSDV}$, G_{ols}^{++}, P_m, $\Gamma_{LM}(p)$ and $\Gamma_{LM}^{B}(p)$ statistics diverge to negative infinity and are compared to the left tail, whereas the Z, P_e^c and $LM(\lambda)$ statistics diverge to positive infinity and are thus compared to the right tail.

3 Empirical Results

The monthly data that we use is taken directly from Culver and Papell (1997), and covers 13 OECD countries between February 1957 and September 1994.[1] For the implementation of the tests, we use the Bartlett kernel, and all bandwidths and lag lengths are chosen according to $4(T/100)^{2/9}$. The maximum number of common factors and structural breaks are set equal equal to five, which is a common choice in the literature. Also, to allow for at least some form of cross-sectional correlation in the first and third group of tests, we use data that has been demeaned with respect to common time effects.

The test results are reported in Table 1. Looking first at the first group of tests, we see that the I(1) null is strongly rejected for all tests and all panels, which corroborate the Culver and Papell (1997) findings. However, as pointed out in the introduction, the restrictive nature of these tests does not allow one to discriminate between stationarity and nonstationarity with cross-sectional dependence and structural breaks.

To get at feeling of the size of the cross-sectional dependence problem in the inflation data, we computed the long-run cross-sectional correlation matrix of the OLS residuals obtained from (1).[2] The results show that all correlations lie between 0.26 and 0.84, with an overall average of 0.55, suggesting that the independence assumption is clearly violated.

Therefore, we now proceed by considering the results of the second and third group of tests, which are reported in Table 1. It is seen that all tests except $LM(\lambda)$ lead to a clear rejection of the I(1) null at the 5% level of significance, which we take as a strong evidence in favor of the stationarity of inflation. There is no difference depending on whether ϕ_i is restricted to be equal for all i or not. For the $LM(\lambda)$ test, we end up with five rejections at the 5% level and one rejection at the 1% level. However, since the rejections are only marginal, and since the alternative hypothesis for this particular test only requires a single series to be nonstationary, we choose to interpret these results as evidence in favor of stationarity for the panel as a whole.

[1] The data is downloadable from Journal of Applied Econometrics data archive available online at http://qed.econ.queensu.ca/jae/.

[2] Results are not shown here to save space, but are available from the corresponding author upon request.

4 Conclusion

In this paper, we examine the robustness of the somewhat controversial finding of Culver and Papell (1997) that the rate of inflation in I(0). It is argued that the panel unit root test employed by Culver and Papell (1997) is based on unrealistic assumptions and that there is a need to reevaluate the results while allowing for more general data generating processes. Results obtained from a large battery of recent panel data unit root tests suggest that the stationarity of inflation holds even after allowing for general forms of cross-sectional dependence and multiple structural breaks in each cross-section.

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Table 1: Panel unit root tests of inflation

		Culver and Papell (1997) defined panels[a]											
Study	Test	1	2	3	4	5	6	7	8	9	10	11	5% val
Common time effects and no breaks													
Im *et al.* (2003)	Z_{tbar}	−27.25	−26.54	−26.32	−25.79	−25.59	−25.04	−24.81	−24.03	−17.77	−20.92	−19.99	−1.65
	$Z_{\tilde{t}bar}$	−24.97	−24.29	−24.10	−23.68	−23.39	−22.95	−22.75	−22.00	−16.50	−18.93	−18.29	−1.65
Levin *et al.* (2002)	t_{δ}^{*}	−9.92	−9.16	−9.77	−10.22	−9.00	−9.48	−10.08	−9.33	−8.65	−4.98	−7.94	−1.65
Harris and Tzavalis (1999)	$\hat{\varphi}_{LSDV}$	−546.03	−532.08	−533.72	−522.13	−519.39	−507.80	−509.40	−494.70	−296.52	−419.26	−434.93	−1.65
Im *et al.* (2005)	$\Gamma_{LM}(p)$	−17.56	−16.64	−17.51	−16.04	−16.58	−15.05	−15.96	−14.94	−10.57	−14.42	−12.80	−1.65
Common factors and no breaks													
Pesaran (2003)	IPS^{*}	−6.17	−6.16	−6.18	−6.18	−6.18	−6.17	−6.19	−6.19	−6.19	−6.19	−6.19	−2.20
Bai and Ng (2004)[b]	P_{e}^{c}	29.60	28.44	28.44	28.44	27.23	27.23	27.23	25.96	21.72	20.11	21.72	1.65
Phillips and Sul (2003)	G_{ols}^{++}	−63.96	−61.86	−61.28	−59.47	−61.55	−59.07	−56.76	−55.76	−28.67	−46.92	−44.07	−1.65
	P_{m}	−12.88	−12.33	−12.33	−12.33	−11.76	−11.76	−11.76	−11.16	−7.94	−8.32	−9.11	−1.65
	Z	28.44	27.23	27.23	27.23	25.96	25.96	25.96	24.63	16.26	18.36	20.11	1.65
Moon and Perron (2004)[b]	t_{a}	−631.45	−626.09	−576.60	−579.08	−566.26	−575.46	−499.03	−501.63	−259.21	−300.85	−503.36	−1.65
	t_{c}	−48.50	−48.33	−43.00	−43.80	−42.27	−43.65	−38.98	−39.90	−23.63	−24.99	−36.98	−1.65
Common time effects and structural breaks													
Im *et al.* (2005)	$\Gamma_{LM}^{B}(p)$	−23.18	−22.42	−22.73	−21.02	−21.95	−20.17	−20.49	−19.61	−14.78	−18.16	−17.09	−1.65
Carrion-i-Silvestre *et al.* (2005)[c]	$LM(\lambda)$	2.18	2.26	1.66	1.55	1.73	1.63	1.01	1.08	3.12	1.15	0.75	1.65

Notes: The tests are computed using the Bartlett kernel. All bandwidths and lag lengths are chosen according to $4(T/100)^{2/9}$. For t_{δ}^{*}, the bandwidth is chosen according to the rule $3.21T^{1/3}$. All tests except $LM(\lambda)$ take nonstationarity as the null hypothesis. For $\Gamma_{LM}^{B}(p)$ and $LM(\lambda)$, we use a trimming parameter of $0.1T$

[a]The panels are numbered in the order they appear from top to bottom in Table III of Culver and Papell (1997).

[b]The test is based on a maximum of five common factors. The IC_1 criterion of Bai and Ng (2004) is used to select the appropriate number of breaks.

[c]The test is based on a maximum of five breaks.