Testing for Group-Wise Convergence with an Application to Euro Area Inflation

Lopez, Claude and Papell, David

University of Cincinnati

2010

Testing for Group-Wise Convergence with an Application to Euro Area Inflation

Claude Lopez [*]
University of Cincinnati

David H. Papell [**]
University of Houston

Abstract

We propose a new procedure to increase the power of panel unit root tests when used to study group-wise convergence. When testing for stationarity of the differential between a group of series and their cross-sectional means, although each differential has non-zero mean, the group of differentials has a cross-sectional average of zero for each time period by construction. We incorporate this constraint for estimation and generating finite sample critical values. Applying this new procedure to Euro Area inflation, we find strong evidence of convergence among the inflation rates soon after the implementation of the Maastricht treaty and a dramatic decrease in the persistence of the differential after the occurrence of the single currency.

We thank Paul Evans, Chris Murray, and participants at the Midwest Econometric Group, Purdue University, Sam Houston State University, Southern Economic Association and Society for Nonlinear Dynamics and Econometrics meetings for helpful comments and discussions. Lopez would like to acknowledge the financial support of the Taft Research Center.

[*] Department of Economics, University of Cincinnati, Cincinnati, OH 45221-0371 Tel: +1 (513) 556-2346. Email: Claude.Lopez@uc.edu
[**] Department of Economics, University of Houston, Houston, TX 77204-5882 Tel: +1 (713) 743-3807. Email: dpapell@uh.edu

1. Introduction

Time series investigation of the convergence hypothesis often relies on unit root tests. The rejection of the null hypothesis is commonly interpreted as evidence that the series have converged to their equilibrium state, since any shock that causes deviations from the equilibrium eventually dies out. The extension of these tests to the panel framework has significantly influenced the literature on how to measure convergence of macroeconomic variables. The combination of time-series and cross-sectional information leads to tests with improved performance, especially for the data lengths usually encountered in macroeconomic analysis.

Initial work on panel unit root tests adapted univariate unit root tests to the panel setting. Levin, Lin and Chu (2002) consider a homogenous speed of mean reversion across the series while others, such as Im, Pesaran and Shin (2003) and Maddala and Wu (1999) allow for heterogeneous speeds of reversion. These tests, and improved versions of them, are widely used to investigate the stationary behavior of macroeconomic time series such as real GDP (Fleissig and Strauss (1999)), inflation (Culver and Papell (1997)), interest rates (Wu and Zang (1996)) and real exchange rates (Papell (1997)).

While the first generation of panel unit root tests relied on cross-sectional independence of the series, more recent work relaxes this assumption by proposing different specifications for cross-sectional dependency. O'Connell (1998), Papell and Theodoridis (2001), Chang (2004), and Lopez (2009) account for cross-sectional dependence by estimating the residual covariance matrix.[1] Among the studies that apply these tests, Rapach (2002) and Hegwood and Papell (2007) examine real GDP, Afonso and Rault (2007) investigate government expenditure and revenue, Kappler (2006) focuses on worked hours, and Papell (2006) and Lopez (2008) study real exchange rates.

Panel unit root tests for convergence among series, or group-wise convergence, utilize Bernard and Durlauf's (1995, 1996) definition of time series convergence for long-run output movements, where two (or more) countries converge when long-run forecasts of per capita output differences tend to zero as the forecasting horizon tends to infinity. In the bivariate context, tests for time series convergence require cross-country per capita output differences

[1] An alternative is to use a factor structure approach as in Bai and Ng (2004). Breitung and Pesaran (2005) provide a survey of the existing literature.

to be stationary. In the multivariate, or panel, context, a group of countries converge if the null hypothesis that the difference between each country's output and the cross-sectional mean has a unit root can be rejected in favor of the alternative hypothesis that each difference is stationary. Panel methods have been used to investigate output convergence by Ben-David (1993, 1996), Islam (1995), Evans and Karras (1996), Evans (1998), and Fleissig and Strauss (2001), among others.

Convergence of inflation rates, especially within the European Union, is another topic for which panel unit root tests have been fruitfully employed. Whether the variable of interest is the inflation rate, as Lee and Wu (2001), or the inflation differential with respect to the cross-sectional mean, as in Kočenda and Papell (1997) and Weber and Beck (2005), these studies report conclusions on the mean reverting behavior of the inflation rates.

While the first and second generations of panel unit root tests have significantly enhanced finite sample performance, these tests can still have low power to reject the unit root null in a panel of stationary series if the panels consist of highly persistent series, contain a small number of series, and/or have series with a limited length. This paper proposes a new procedure that improves the power of panel unit root tests when testing for group-wise convergence.

Panel unit root tests for group-wise convergence involve stationarity between a group of series and their cross-sectional means. As the series may not be characterized by absolute convergence toward the cross-sectional average, each differential can have a non-zero mean. By construction, however, the group of differentials has a cross-sectional average of zero for each time period. In order to improve the panel unit root test's performance, we exploit this extra information in the data by incorporating the appropriate restriction when estimating the model and generating finite sample critical values. To our knowledge, this constraint has not been utilized for previous tests of convergence using panel unit root tests. It should be emphasized that our proposed method is only applicable for tests of group-wise convergence. The power of panel unit root tests that examine the Purchasing Power Parity hypothesis by investigating the stationarity of real exchange rates, for example, cannot be improved by our method as, in this case, the series are individually converging to their own mean but not to a common target.

Monte Carlo simulations confirm the enhanced size and size-adjusted power of the test when using the constraint. Since imposing a valid constraint will increase power, and the constraint is valid by construction, this result is not surprising. What is more interesting is that the increase in power is generally larger for more persistent data, lower numbers of series, and smaller data spans, that is for the cases commonly encountered in testing for convergence among macroeconomic variables such as output and inflation.

We then investigate the performance of the test when dealing with mixed panels of stationary and unit root series. In contrast to the previous simulations, the frequency of rejection is not always higher when the constraint is imposed. The presence of unit roots leads to a significant decrease in the restricted test's ability to reject the unit null hypothesis. This is a desirable feature as it shows that the new restricted test leads to more reliable evidence of stationarity of the entire panel than the unrestricted test.

The enhanced performance of the testing procedure enables a reduction of the data length while preserving good power of the test, allowing us to analyze data sets for the post-Euro, 1999-2006 period. More specifically, we focus on inflation convergence among the Euro countries. While our main concern is to investigate whether the rates respect the Maastricht criterion after the Euro, we are also interested in any potential impact of the Euro on the rate of convergence among the series. As a result, the study analyzes the 1979:1 to 2006:12 period using a rolling window of eight years, starting with 1979:1-1986:12 and ending with 1999:1-2006:12. Note that the last window solely accounts for the post-Euro period, hence isolates its impact. Furthermore, this rolling window approach deals with any potential time break in the data due to events such as German reunification.

We first apply the new testing procedure. Our results show that the inflation rates have converged toward a common target as early as just after the implementation of Maastricht treaty and that this convergence remains strong until after the advent of the Euro. We then focus on the rate of persistence of the differentials, as it is directly linked to the degree of convergence among the inflations rate. The results highlight three phases: (i) periods ending between 1986:12-2002:1, or pre-Maastricht period where the persistence is quite high but stable; (ii) periods ending between 2002:2-2004:12, or the pre-Euro period where the persistence varies a lot observing, first, a drastic decrease that is later partially compensated; and finally (iii) periods ending between 2005:1 to 2006:12, or Euro period, where the persistence is, once again, stable, yet at a lower level than in the initial period. Note that the described behavior follows closely the European Monetary Union time table.

Finally, we generate median unbiased estimates of the rate of persistence of the differential, their 95% confidence intervals and the corresponding half-lives. These allow us to rigorously compare the rates across the different periods highlighted previously. The results confirm a dramatic decrease in the persistence of the differentials after the occurrence of the single currency. Based on the half-lives, the persistence of the differentials has decreased by more than 40 percent between the pre-Maastricht and Euro periods and by more than 50 percent between the pre-Euro and Euro periods.

The next section develops the new testing procedure, and conduct detailed size and power experiments while Section 3 present the empirical application and Section 4 concludes the paper.

2. Panel Unit Root Tests for Convergence

2.1 Group-wise stochastic convergence

In the panel framework, testing for (stochastic) convergence of a group of N time series requires studying the dynamic properties of the series differential with respect to the cross sectional mean. Group-wise (stochastic) convergence implies that:

$$\lim_{k \to \infty} E\left(y_{i,t+k} - \sum_{j=1}^{N} \frac{y_{j,t+k}}{N} \middle| I_t \right) = \eta_i \;\; for\; i = 1, \dots, N \tag{1}$$

Where I_t represents the information set available at time t. If η_i =0, then the convergence follows Bernard and Durlauf (1996)'s definition of absolute convergence. If $\eta_i \neq 0$ then the convergence is said conditional or relative as defined by Durlauf and Quah (1999), which implies that the series converge toward a time-invariant equilibrium differential. [2]

[2] The differentials will be stationary if either the series and the cross-sectional mean are both I(0) or if they are both I(1) and cointegrated.

2.2 Panel unit root test

We modify standard panel unit root tests to account for the restriction on the intercepts when testing for group-wise convergence. More specifically, we focus on the second generation of panel unit root tests that account for contemporaneous correlation by estimating the residual covariance matrix. The test considered is an extension of the Levin, Lin, and Chu (2002) application of the ADF test to the panel framework that investigates a homogeneous rate of convergence across the series. Let consider the following system of ADF regressions:

$$\begin{gathered}\Delta y_{it} = c_i + \rho y_{i,t-1} + \sum_{j=1}^{k_i} \phi_{ij} \Delta y_{i,t-j} + \varepsilon_{it} \\ with \quad i = 1,...,N \quad t = 1,...,T \quad and \quad \varepsilon_{\mathrm{it}} \sim WN(0;\Sigma)\end{gathered} \tag{2}$$

where ρ is the homogeneous rate of convergence, k_j the lagged first differences that account for serial correlation and $\sum$ the non-diagonal covariance matrix. The null and alternative hypotheses tested are $\rho = 0$ and $\rho < 0$.

The pooled ADF test relies on feasible generalized least squares (SUR) method, hence the name ADF-SUR test. It is performed in two steps. First, for each series, k_j is selected with the recursive lag-selection procedure of Hall (1994) and Ng and Perron (1995). Then, the residuals covariance matrix is deduced and used to estimate (2) with the SUR method, constraining the values of ρ to be identical across equations and using the preselected k_j. Finally, the estimated ρ and its corresponding standard deviation allow us to calculate the t-statistics corresponding to the null ρ = 0. Since the focus of the paper is on a panel of macroeconomic variables where the time series dimension is large compared to the cross-section dimension, it is assumed that T>N.

While it would be desirable to allow for heterogeneous rates of convergence, the choices are problematic. Following Im, Pesaran, and Shin (2003), several tests that average t-statistics across the members of the panel have been developed. The alternative hypothesis for these tests, however, is that $\rho_i < 0$ for at least one *i*, which is not economically relevant for investigating convergence among a group of countries. The tests developed by Breuer, McNowan, and Wallace (2002), which allow ρ_i to be heterogeneous across countries in a framework similar to (2), provides (at best) modest increases in power over univariate tests.

2.3 The new testing procedure

Our testing procedure benefits from extra knowledge available about the data and is directly used to design a model that accounts for all information available prior to the estimation. More specifically, this non-sample information is included as a restriction in the estimation and when generating the finite sample critical values. The restriction being true by construction, the final estimator ends up with a smaller variance than the unrestricted one. Greene (2008, p89) suggests that "one way to interpret this reduction in variance is as the value of the information contained in the restriction". [3]

The procedure relies on the knowledge that, once transformed, the data may have a non-zero mean for each differential *i* but a cross-sectional mean equal to 0 at every period. If y_{it}^{diff} is the differential for country *i* at time *t* with respect to the cross-sectional mean such that $y_{it}^{diff} = y_{it} - \sum_{i=1}^{N} y_{it} / N$; then by construction, for each period of time *t*, the sum of the differentials is equal to 0, that is $\sum_{i=1}^{N} y_{it}^{diff} = \sum_{i=1}^{N} \left(y_{it} - \frac{\sum_{i=1}^{N} y_{it}}{N} \right)$=0. Let replace y_{it} by y_{it}^{diff} in (2), then the intercepts c_i are on average equal to 0. Hence, the estimation uses the restriction $\sum_{i=1}^{N} c_i = 0$. Note that, since each regression allows for an intercept, we are not testing for absolute convergence. The resulting system of equations is:

$$\begin{gathered} \Delta y_{it}^{diff} = c_i + \rho y_{it-1}^{diff} + \sum_{j=1}^{k_i} \phi_{ij} \Delta y_{it-j}^{diff} + \varepsilon_{\mathrm{it}} \\ with \quad i = 1,...,N \quad t = 1,...,T \quad and \quad \varepsilon_{\mathrm{it}} \sim WN \end{gathered} \tag{3}$$

where ρ is the homogenous rate convergence, and y_{it}^{diff} is the data differential with respect to the cross-sectional mean. The error terms ($\varepsilon_{1t},...,\varepsilon_{Nt}$) are stationary with a non-diagonal covariance matrix $\sum$. The standard hypotheses, H_0: $\rho = 0$ versus H_1: $\rho < 0$, are tested.

The estimation procedure follows three steps:

---

[3]Judge et al. (1988, p812) explains that" if nonsample information is correct, then using it in conjunction with the sample information will lead to an unbiased estimator that has a precision matrix superior to the unrestricted least squares estimator".

1. *Data transformation*: the differentials with respect to the cross-sectional mean are calculated for all series
2. *Lag selection*: the number of lagged first difference terms allowing for serial correlation, k_i in (3), is selected using the recursive procedure for each series
3. *Estimation*: The residual covariance matrix $\sum$ is estimated. The resulting $\hat{\Sigma}$, along with the pre-selected $k_{j,}$ is then used in the estimation of (3) with the SUR method while two restrictions are imposed:
    a. $\sum_{i=1}^{N} c_i = 0$, that is the non-sample information
    b. $\rho_i = \rho$, that is a homogeneous rate of convergence

The estimated ρ and its corresponding standard deviation are obtained, and the t-statistic is calculated for $\mathrm{H}_0 : \rho = 0$.

The interpretation of the two restrictions is very different. (a) is true by construction, and therefore there is no question whether or not it is correct. (b) is almost surely false, as there is no reason why each country should have the same rate of convergence. There are two ways, however, for the restriction of homogeneous convergence rates to be false. First, all of the $\rho_i < 0$. In that case, rejection of the unit root null correctly provides evidence of convergence. Second, some of the $\rho_i < 0$ and some of the $\rho_i = 0$. In that case, there is a mixed panel and rejection of the unit root null does not correctly provide evidence of convergence. We consider the performance of our test with mixed panels below.

O'Connell (1998), Maddala and Wu (1999), and Lopez (2009), among others, show that panel unit root tests estimating the residual covariance matrix should rely on simulated critical values to reduce any size distortions due to the cross-sectional correlation, while Chang (2004) proves the asymptotic validity of a sieve bootstrap procedure for non-pivotal homogeneous panel unit root tests. As a result, the bootstrap critical values are generated using a non-parametric resampling method with replacement. First, the coefficient estimates ($\hat{\phi}_{ij}$) and the fitted residuals $(\hat{u}_t)$ are estimated from $\Delta y_{it}^{diff} = \sum_{j=1}^{k_i} \phi_{ij} \Delta y_{i,t-j}^{diff} + u_{\mathrm{it}}$.[4] Then, the bootstrap samples (u_{it}^{*}) are drawn from the centered fitted-residuals. More specifically, to

[4] The initial lag selection uses BIC.

preserve the contemporaneous correlation, the (u^*_{jt}) are resampled as a vector $u^*_t = (u^*_{1t}, u^*_{2t}, ..., u^*_{Nt})$ from the empirical distribution of $\left(\hat{u}_t - \frac{1}{T}\sum_{t=1}^{T}\hat{u}_t\right)_{t=1}^{T}$. Next, the bootstrap samples $\left(\varepsilon^*_{it}\right)$ are recursively generated using the estimated parameters ($\hat{\phi}_{ij}$) and the bootstrap samples $\left(u^*_t\right)$ as $\varepsilon^*_{it} = \sum_{j=1}^{k_i}\hat{\phi}_{ij}\varepsilon^*_{i,t-i} + u^*_{it}$, starting from $\left(u_{i0}, ..., u_{i,-k_j+1}\right)$. Finally, the pseudo-data y^{diff*}_{it} are obtained by taking the partial sum of $\left(\varepsilon^*_{it}\right)$ as $y^{diff*}_{it} = y^{diff*}_{i0} + \sum_{j=1}^{k_i}\varepsilon^*{}_{ij}$.[5] For each generated set of series, the estimation procedure previously explained is applied.

Davidson and G. MacKinnon (2006) explain that "imposing the restriction […] yields more efficient estimates of the nuisance parameters upon which the distribution of the test statistics may depend. This generally makes bootstrap test more reliable, because the parameters of the bootstrap DGP are estimated more precisely". Since the restriction is true by construction, we expect the restricted test to perform better in small samples than the unrestricted one.

2.4 Impact of the constraint in small samples

In order to analyze the impact of the restriction $\sum_{i=1}^{N} c_i = 0$, a set of simulations investigates the finite sample performance of the ADF-SUR test with and without the restriction. Let consider the following data generating processes:

$$y_{it} = \rho y_{i,t-1} + u_{i,t} \qquad \text{with } i=1, ..., N \text{ and } t=1, ..., T$$

The innovations $\{u_{it}\}$ are drawn from *iid* normal distributions with mean zero and a diagonal covariance matrix $\sum$.[6] The panel dimensions are N = 5, 10, and 20 and T = 25, 50, 100, and 200. For each experiment, the finite sample critical values and the empirical rejection probabilities calculated at a 5% nominal level are based on 2000 iterations.[7,8] Since we are

---

[5] Each pseudo-data $\left(y^{diff*}_{it}\right)$ is generated with T+50 observations, then the first 50 observations are discarded, hence each $\left(y^{diff*}_{i0}\right)$ is random.

[6] Similar simulations have been reproduced using non-diagonal matrix covariance, that is including and accounting for contemporaneous correlation, without any significant change regarding the impact of the restriction on the intercept.

[7] Davidson and McKinnon (1999) advise a minimum of 1500 bootstraps when analyzing the performance of the test at 1%.

using randomly generated data, each experiment is repeated 20 times, hence Tables 1 and 2 report the average rejection probabilities.

Table 1 reports the finite sample properties of both restricted and unrestricted ADF-SUR tests. The data sets are generated under the null hypothesis (ρ = 1.00) for the size and under the alternative (ρ = 0.99, 0.97, 0.95 and 0.90) for the size adjusted power.[9] Both tests report almost no size distortion with a probability of rejecting the unit root null when the data have one, close the nominal size of 5%. However, the tests significantly differ in their ability of rejecting accurately the null hypothesis when analyzing stationary data. For example, for highly persistent data such that (N, T, ρ) = (10, 100, 0.97), the restriction increases the size-adjusted power of the ADF-SUR test from 0.384 to 0.595. Similarly, for moderately persistent data such that (N, T, ρ) = (20, 50, 0.95), the restriction increases the power from 0.337 to 0.539. As expected, these improvements disappear as N and T increase and the data is less persistent, that is in the cases where the ADF-SUR test performs well. In addition, the restriction has only a moderate impact when the panel has a small time dimension, T = 25 and 50, and the data is extremely persistent, ρ = 0.99. In sum, the restriction significantly enhances the test's performance for persistent data ($\rho > 0.9$) and small to medium data spans ($T < 200$).

Table 2 focuses on the test's performance when the data is not generated under the alternative hypothesis of homogeneous and stationary rates of convergence but as a mix of stationary and non-stationary processes. More specifically, some series converge at a same rate ($\rho_i = \rho =$ 0.97, 0.95, 0.90 and 0.8 for $i = 1,\dots,k$) while others follow a non-stationary process (ρ_j = 1.0 for $j = k+1, \dots, N$).[10] The data length T is equal to 100 for N = 5, 10 and 20.

Such an experiment allows us to investigate whether the improved finite sample performance of the restricted test leads to an increase in unwanted rejections of the null hypothesis over the unrestricted test. Indeed, Taylor and Sarno (1998) and Breuer, McNowan, and Wallace (2001) have provided evidence that, in the general case where the sum of the intercepts is not constrained to equal zero, the unit root null can be rejected by panel methods with homogeneous rates of convergence even when the panels contain only a few stationary

---

[8] Davidson and McKinnon (2006) define and discuss this probability for the power and size of bootstrap tests

[9] The case ρ = 0.8 is not reported as it does not provide any new insights on the test's behavior.

[10] The case ρ = 0.99 is not reported as it does not provide any new insights on the test's behavior.

series. Breuer, McNowan, and Wallace (2002), Sarno and Taylor (2002), and Taylor and Taylor (2004) go further, arguing that the unit root null can be rejected even if only one of the series is stationary. To address this concern, we first look at the bottom row of Table 2, for N = 5, 10, and 20, that reports the (correctly sized 0.05) rejection frequencies when all series have a unit root. Going up one row, the rejection frequencies for both the restricted and the unrestricted tests are depicted when one of the series is stationary, that is $(\rho_i, \rho_j) = (\rho, 1.00)$ for $i = 1$ and $j = 2,\ldots,$N. For $N = 5$, they range from 0.07 ($\rho = 0.97$) to 0.11 ($\rho = 0.8$), for N = 10, they range from 0.06 ($\rho = 0.97$) to 0.08 ($\rho = 0.8$) and, for $N = 20$, they range from 0.06 ($\rho = 0.97$) to 0.07 ($\rho = 0.8$). Hence, it seems very unlikely that the inclusion of one stationary series will produce a rejection of the unit root null with any of these tests.[11]

While the argument that inclusion of one stationary series will produce rejections using panel unit root tests with homogeneous rates of convergence seems overstated, the results confirm that one needs to be careful about interpreting rejections of the null as evidence that all of the series are stationary. For example, with $N = 10$, both tests report a rejection frequency of about 0.50 with 8 stationary series if $\rho = 0.95$. Since the result of rejection or non-rejection would be analogous to the outcome of a coin flip, one would not want to conclude in favor or against the null hypothesis.

Yet, it is worth noting that, for all three panels with a mix of unit root and less persistent ($\rho = 0.9$ and 0.8) stationary series, the rejection frequencies for the restricted test are smaller than those for the unrestricted test. Hence, one would be less likely to falsely reject the unit root null hypothesis for most of the cases when using the restricted ADF-SUR test. For the panels with a mix of unit root and more persistent ($\rho = 0.95$ and 0.97) stationary series, the rejection frequencies for the restricted tests are still smaller or equal to those for the unrestricted tests except in presence of very few (up to three depending the panel) unit roots.

In practice, however, one is much less likely to falsely reject the unit root null with restricted than with unrestricted ADF-SUR tests. This is because, with highly persistent processes and $N = 5$ or $N = 10$, the tests do not have much power to reject the unit root null even when all of the series are stationary. Taking the most extreme example $(N, \rho) = (5, 0.97)$

[11] Some of our rejection frequencies without the constraint are lower than in Breuer, McNowan, and Wallace (2001) for identical panels. The differences appear to be due to their use of Levin, Lin, and Chu (2002) critical values which do not account for serial correlation. Papell (1997) discusses this issue.

for emphasis, the 5% size adjusted power is only 0.41 for the restricted test and 0.23 for the unrestricted test when all of the series are stationary. With one stationary series, the fact that the rejection frequency is larger for the restricted (0.22) than the unrestricted (0.16) test is unlikely to cause an inappropriate conclusion as the restricted test under rejects the null hypothesis around 80% of the time.

A very different picture emerges with less persistent processes where the tests have good power to reject the unit root null when all of the series are stationary. We will focus on a comparison of the rejection frequencies between the two tests for the smallest number of stationary series for which the rejection frequency of the unrestricted test is 0.50 or higher. For N = 5, the rejection frequency is 0.58 for the restricted test and 0.65 for the unrestricted test with 4 stationary series and ρ = 0.9 and is 0.40 for the restricted test and 0.51 for the unrestricted test with 3 stationary series and ρ = 0.8. With N = 10, the rejection frequency is 0.57 for the restricted test and 0.66 for the unrestricted test with 7 stationary series and ρ = 0.9 and is 0.54 for the restricted test and 0.64 for the unrestricted test with 6 stationary series and ρ = 0.8. When N = 20, the rejection frequency is 0.46 for the restricted test and 0.56 for the unrestricted test with 11 stationary series and ρ = 0.9 and is 0.42 for the restricted test and 0.53 for the unrestricted test with 9 stationary series and ρ = 0.8. In the above examples, both tests have high power to reject the unit root null when all of the series are stationary, so they represent cases where it is plausible that the unit root null might be rejected with a mixture of stationary and non-stationary series. While other examples could be chosen, the pattern is clear. For mixed panels that contain less persistent stationary series with ρ = 0.8 or ρ = 0.9, one is less likely to mistakenly reject the unit root null hypothesis with the restricted than with the unrestricted tests.

When the data is, by construction, restricted so that the sum of the intercepts is equal to zero for each period, the gain in efficiency obtained by imposing the restriction in the estimation has two main impacts on the ADF-SUR test.[12] First, the more precise estimation and resulting bootstrap procedure leads to a more powerful size-adjusted test for the most commonly encountered panel dimensions in macroeconomics. Second, the rejection frequencies are generally smaller for mixed panels of stationary and non-stationary processes.

---

[12]The gain in efficiency refers to the more precise of the estimation that leads to smaller variance of the error terms.

Combining the results, the restriction improves the overall behavior of the test, enhancing its ability to correctly reject the unit root null hypothesis when all series are stationary and to correctly fail to reject the unit root null when a subset of the series are non-stationary.

3. Inflation convergence within the Euro Area, 1979-2006

The Maastricht treaty, signed in 1992, states that "the achievement of the high degree of price stability...will be apparent from a rate of inflation which is close to that of, at most, the three best performing member States in terms of price stability."[13] In practice, the inflation rate of a given country is measured by the CPI and must not be greater than 1.5 percentage points above of the three EU countries with the lowest inflation. Hence, the criterion imposes that the inflation rates converge toward a common value.

In light of the achievement of the Maastricht criteria, the fixing of Euro Area exchange rates in mid-1998, and the establishment of the Euro in January 1999, one would expect Euro Area inflation rates to have converged during the period immediately preceding the advent of the Euro. This expectation is confirmed by numerous studies, including Rogers, Hufbauer and Wada (2001), Engel and Rogers (2004), Weber and Beck (2005), Busetti, Forni, Harvey and Venditti (2007) and Rogers (2007), which agree that prices were less dispersed and inflation rates among Euro Area countries converged in the mid-1990s. In contrast, research investigating the post-1998 period, including ECB (2003), Honohan and Lane (2003), Engel and Rogers (2004), Weber and Beck, (2005), Rogers (2007), and Fritsche and Kuzin (2008), concludes that the advent of the single currency resulted in the weakening of inflation convergence among the Euro Area countries and in an increase in their price dispersion. An exception is Honohan and Lane (2004), who report sharp convergence in inflation rates since 2002.

Our study focuses on time series measurement of inflation convergence. We examine the behavior of inflation differentials with respect their cross-sectional means. First, we investigate the evolution of convergence over time, starting with a period prior to the European Monetary System and ending with the post Euro period. Next, we highlight the impact of the Euro by comparing the estimated speeds of convergence before and after the adoption of the single currency.

[13] The text of the Maastricht Treaty can be found at www.europa.eu.

3.1 Data and estimation results

This study aims to describe the evolution of Euro Area inflation rates from the first stage of the European Monetary System up to 2006, while isolating the post Euro period. Annual inflation rates with monthly data π_{it} for the i^{th} country at time t are calculated such that: $\pi_{it} = \ln(CPI_t) - \ln(CPI_{t-12})$.[14] The differentials y_{it} are generated so that: $\pi_{it}^{diff} = \pi_{it} - \bar{\pi}_t$ where $\bar{\pi}_t$ is the cross-sectional average inflation rate.[15]

The monthly CPI data are from International Financial Statistics (CD June 2007) from 1979:1 to 2006:12. Euro 11 (E11) countries are Austria, Belgium, Finland, France, Germany, Greece, Italy, Luxembourg, Netherlands, Portugal and Spain, and Euro 10 (E10) countries are E11 without Greece.[16]

The descriptive statistics of the data provides us with some useful insights. The cross-country means and standard deviations of the inflation rates, reported in Figure 1, show an overall decrease in the cross-sectional mean and variance throughout the entire period for both the E10 and E11 panels. More specifically, this decrease occurs in three phases: 1982-1987 has the steepest slope, followed by 1990-1999 with a flatter slope and 2000-2006 with no visible change in the slope.[17] Furthermore, the stable standard deviation of the 2000-2006 period confirms that the inflation rates seem to observe a stable relation between each other.

The enhanced performance of the new estimation procedure enables us to consider relatively short periods while retaining good size and power of the test. We isolate the Euro period by considering an eight-year estimation window with 96 monthly observations, which corresponds to the data for 1999:1-2006:12. The window is then rolled from 1979:1-1986:12 to 1999:1-2006:12, one month at a time. This approach limits the impact of potential changes in the parameters on the estimation results while depicting the evolution of the results through time. In contrast to studies which use a recursive (expanding) estimation window to study convergence, our results are not affected by the fact that the power of panel unit root tests increases with the number of observations as well as the size of the panel.

---

[14] The data is seasonally adjusted

[15] Yearly inflation with monthly data and annualized monthly average inflations yield to similar results.

[16] The monthly data for Ireland is available only starting 1998:1, hence we do not consider E12 as we could not compare the periods before and after the advent of the Euro.

[17] Lopez (2008b) shows that the Euro-zone inflation rates are regime-wise stationary.

Figure 2 reports p-values for both the E10 and E11 panels using restricted and unrestricted ADF-SUR tests. Both groups of countries lead to similar conclusions. A comparison between the restricted and the unrestricted estimations emphasizes the impact of the previously discussed gain in precision with the new estimation procedure. While the results observe a similar pattern, the restricted approach consistently leads to lower p-values. The findings based on restricted estimation show rejections of the unit root hypothesis for pre-Euro windows ending in 1990-1994 and 1997-1998. The evidence of convergence is stronger for the E11 panel than for the E10 panel for the windows ending in the early 1990s. For E11, the unit root null is rejected (at the 10 percent level) for all windows ending in 1990:3 – 1994:4 except for 1991:9-1999:11. For E10, convergence is only found for 1990:3 - 1991:7. The opposite holds for the windows ending in the late 1990s. For E10, the unit root null is rejected for all windows ending in 1997:4 - 1998:6 while, for E11, convergence is only found for windows ending in 1997:12 - 1998:4.

The strongest evidence of convergence comes from windows ending in 2000 – 2006, after the advent of the Euro. Again focusing on the restricted estimation, the unit root null is rejected (at the 10 percent level) with E11 for all windows ending in 2002:3 - 2006:12 and with E10 for all windows ending in 2000:9 - 2006:12 except for 2001:9-2002:1 and for 2005:7-2005:11. The impact of imposing the restrictions is very clear for the Euro period. For the estimates that do not impose the restrictions, evidence of convergence is sporadic after 2004 for E11 and almost disappears for E10. It should perhaps be emphasized that, for the particular case of testing for group-wise convergence, there is no question that imposing the restriction that the sum of the intercepts in Equation (3) is equal to zero is the correct procedure. Unlike the usual case of imposing restrictions, which may or may not be correct, this restriction is correct by construction.

Figure 3 plots the values of ρ for the restricted model for E10 and E11 from 1979:1-1986:12 to 1999:1-2006:12. In accord with our definition of group-wise convergence, variations of ρ can be interpreted as a measurement of the strength of inflation convergence toward Maastricht's common target. As a result, a more persistent differential (higher value of ρ) would correspond to weaker inflation convergence as any shock would have a longer lasting impact, and a less persistent differential (lower value of ρ) would correspond to stronger inflation convergence. In contrast to the p-values, the rate of convergence remains relatively stable up to the window ending in 2002:1. Starting in 2002:2, the rate increases (lower value of ρ) and decreases (higher value of ρ) sharply before returning to more stable behavior near the end of the sample. The rate of convergence at the end of the sample is faster than the rate that characterizes the sample for the windows ending before 2002. The lower values of ρ for the windows starting in 2002:2 are consistent with Honohan and Lane's (2004) evidence of convergence in Euro Area inflation rates since 2002.[18]

The initial phase, ending in 1997:3, reports a rate of convergence close to 0.96 for both panels. The E10 panel then shows a slow decrease in ρ from 0.951 to 0.939, between the windows ending in 1997:4 and 2002:1, while the E11 panel remains highly persistent. Both panels report drastic changes in the rate of convergence for the windows ending from 2002:2 to 2005:1. The windows ending between 2002:2 and 2003:12, first, report a significant reduction in persistence (ρ decreases from 0.939 to 0.839 for E10 and from 0.945 to 0.866 for E11), which is then partially compensated by a strengthening of the persistence (ρ increases from 0.842 to 0.898 for E10 and from 0.866 to 0.906 for E11) for the windows ending in 2004:1-2004:12. Following this period of transition, a period of stability concludes the sample: the windows ending between 2005:1 and 2006:12 report an average value for ρ of 0.908 for E10 and 0.904 for E11.

The behavior of the rates of convergence closely follows the European Monetary Union (EMU) timetable. The mechanism that led to the single currency included three major steps: from 1990:7 to 1993:12 (windows ending in 1997:7-2000:12), capital was allowed to move freely within the European Economic Community, from 1994:4 to 1998:12 (windows ending in 2001:4-2005:12) the Treaty of Maastricht was implemented and in 1999:1 (window ending in 2006:12), the single currency was introduced.

It is also worth noting that evidence of stationarity of the differentials, rejection of the unit root null, occurs several years before the processes reach a steady level of persistence. While inflation rates start converging with the 1995-2002 window, they do not attain a stable level of convergence until the 1997-2004 window. This final degree of convergence is

[18] While the value of ρ is biased downward, the focus in the section is on a comparison across time periods which would not be affected by bias correction. In the next section, we conduct median-unbiased estimation for several windows.

significantly higher (ρ is significantly lower) than the one estimated for the first two phases of the EMU.

3.2 Measuring persistence

A closer investigation of the impact of the Euro requires a rigorous assessment of the speed of convergence for the inflation differentials. In order to provide an accurate measure of persistence, we apply median unbiased corrections to the restricted and the unrestricted estimates. We focus on three windows: 1982:7-1990:6, which covers the pre-Maastricht era, 1990:7-1998:6, which covers the pre-Euro period and ends six months before the exchange rates were fixed, and 1999:1-2006:12, which covers the Euro period.

Following Murray and Papell (2005), we use an extension of the Andrew and Chen (1994) method to the panel framework. The originality of our approach, however, consists of generating median unbiased estimates of the homogeneous rate of convergence for the restricted model. The iterative procedure used to generate the approximately median unbiased estimate, ρ_{AMU}, of ρ in (3) starts with the estimation of ϕ_{ij} in (3) via the new procedure. Then, assuming the estimates of ϕ_{ij}'s are true, the first median unbiased estimate $\rho_{1,AMU}$ is obtained by finding the median-unbiased estimator that corresponds to the value of $\rho_{\text{SUR-restricted}}$. We then assume $\rho_{1,AMU}$ to be the true value of ρ and obtain a new set of estimates for the ϕ_{ij}'s. Conditional on these news estimates, we obtain the new median unbiased estimates $\rho_{2,AMU}$. The iterative process continues until convergence occurs and median unbiased estimates of $\rho_{\text{SUR-restricted}}$ and the ϕ_{ij}'s are obtained.

Table 3 reports the rates of convergence for the differentials, the median unbiased estimates (point estimates and 95% confidence intervals of ρ), and the corresponding half-lives. The median unbiased point estimates are (as expected) higher than the GLS estimates. The Euro period is characterized by the fastest rates of convergence, followed by the pre-Maastricht period, with the pre-Euro period displaying the slowest convergence rates. This pattern holds for the E10 and E11 panels and the restricted and unrestricted estimates.[19] For example, using the restricted model, E10 demonstrates a strengthening in group-wise inflation

[19] The only exception is for the unrestricted E10 panel, for which the value of ρ is slightly lower for the pre-Euro than for the pre-Maastricht period.

convergence as ρ_{MU} decreases from 0.970 for the pre-Maastricht period and 0.975 for the pre-Euro to 0.950 for the Euro period.

As expected, there is no difference between the restricted and unrestricted GLS estimates, because the restriction is respected by the data. However, the restriction should lead to a smaller variance of the estimates which is confirmed by the lower restricted median-unbiased estimates compared to the unrestricted median-unbiased estimates across all periods. Similarly, all the confidence intervals when the restriction is imposed are narrower than the unrestricted confidence intervals, confirming the gain in precision from the restrictions discussed above.[20]

The 95 percent confidence intervals for the Euro period confirm the stronger evidence of inflation convergence from the point estimates. The confidence intervals for the E10 panel with the restricted model widen between the pre-Maastricht (0.950 to 0.988) and the pre-Euro (0.946 to 0.996) periods. In contrast, the confidence interval for the post-Euro period (0.914 to 0.975) has a smaller upper bound and a much smaller lower bound than the confidence intervals for the two earlier periods.

The most common measure of persistence of an economic time series is the half-life, the number of periods it takes for a shock on the inflation differential to dissipate by 50 percent. The half-life is approximated by the ratio ($\ln(0.5)/\ln(\rho_{MU})$).[21] The median unbiased estimates and corresponding confidence intervals for the half-lives provide a more explicit illustration of the speed of convergence. A larger half-life would imply slower decay and weaker inflation convergence.

Our results once again illustrate the gain in information due to the use of the restriction, with the restricted HL_{MU} point estimates consistently lower that unrestricted estimates. More importantly, the gain in precision leads to narrower restricted confidence intervals, with a noticeable difference for the upper boundaries. For the half-lives, every

[20] While the E11 panel for the pre-Maastricht period appears to be an exception, with the width of the confidence interval equal to 0.68 for the restricted and 0.53 for the unrestricted estimates, that interpretation is not correct. The upper point of the confidence interval for the unrestricted model is 1.00. Since the confidence intervals are constrained not to exceed unity, no comparison can be made in this case.

[21] While it is generally preferable to compute half-lives from impulse response functions, the panel model used allows for different serial correlation across series, hence there is no common impulse response function on which the half life could be based.

restricted confidence interval is narrower than the corresponding unrestricted confidence interval.

Since the restrictions are valid by construction, we will focus on the median-unbiased estimates of the restricted model. The half-lives of the point estimates of the differentials decrease by more than 40 percent between the pre-Maastricht and Euro periods and by more that 50 percent between the pre-Euro and Euro periods. For the E10 panel, the half-lives rose from 22.76 months in the pre-Maastricht period to 27.38 months in the pre-Euro period, followed by a decline to 13.51 months in the Euro period. For the E11 panel, the half-lives rose from 23.55 months in the pre-Maastricht period to 98.67 months in the pre-Euro period, followed by a decline to 12.98 months in the Euro period. The half-lives for the E10 and E11 panels are very similar for the pre-Maastricht and Euro periods. They are, however, very different for the pre-Euro period. The E11 panel, but not the E10 panel, displays a drastic slowdown of the speed of convergence after the Maastricht treaty, which highlights the impact of Greece and its difficulties in meeting the convergence criteria.

The differences between the Euro and earlier periods are highlighted by the median-unbiased estimates of the confidence intervals of the half-lives for the restricted model, which are both smaller and narrower in the later period. This confirms that not only are inflation differentials less persistent in the Euro period, but that we are more confident about the precision of our estimates of persistence. Going from the pre-Maastricht to the pre-Euro periods, the confidence intervals of the half-lives widen for the E10 panel, which seems coherent with the numerous changes Europe had in the early 1990s (German reunification, differing economic policies) and its evolution toward the more rigorous structure defined by the Maastricht treaty. Similarly, for the same periods, the confidence intervals for the E11 panel increase and widen, again reflecting the influence of the inclusion of Greece. For the Euro period, the confidence intervals of the half-lives are very close for the E10 and E11 panels. The robustness of the results for the Euro period to the panel composition is coherent with the convergence criterion as it sets an identical inflation target for all countries.

4. Conclusion

This paper proposes a new estimation procedure that can be used when investigating convergence of a group of series toward a common target. Group-wise time-series convergence is commonly measured using panel unit root tests on differentials generated as the difference between each series and the cross-sectional average. Hence, each resulting differential has a non-zero mean, but the cross-sectional mean of the group of differentials is equal to zero for each period. Our method uses that information in order to increase the size adjusted power of the test. Monte Carlo simulations report noticeable improvements of the test's power, especially when the data is persistent data ($\rho > 0.9$) or when the data has a limited length ($T < 200$). Both of these characteristics are commonly featured in macroeconomic time series. Furthermore, the restricted ADF-SUR test also shows a greater ability of rejecting the unit root null solely when all the series are stationary, which is welcome improvement on one of the most acknowledge drawback of the panel unit root approach.

Using the new approach, we analyze inflation convergence within the Euro Area countries. More specifically, we investigate when the inflation differentials become stationary and if the Euro has had an impact on inflation differential persistence. The increase in size adjusted power from the imposition of the restriction that the cross-sectional mean of the differentials is equal to zero, which is true by construction, allows us to estimate the model for all eight-year rolling windows and separately for the pre-Maastricht, per-Euro, and Euro periods.

We conduct panel unit root tests for rolling windows from 1979:1-1986:12 to 1999:1-2006:12, and report evidence of group-wise convergence if the unit root null can be rejected for the inflation differentials. While sporadic evidence of inflation convergence begins with the period starting shortly after the implementation of the Maastricht treaty, consistent evidence of convergence only occurs with windows ending during the Euro period.

In order to sharpen our focus on the speed of convergence, we calculate median-unbiased point estimates, half-lives, and confidence intervals for the pre-Maastricht, per-Euro, and Euro periods. The rate of convergence is much faster and the confidence intervals are considerably narrower for the Euro period than for the two earlier periods. The half-lives of the point estimates of the differentials, the number of periods that it takes for a shock to the inflation differentials to decrease by one-half, falls by more than 40 percent between the pre-Maastricht and Euro periods and by more that 50 percent between the pre-Euro and Euro periods.

It is commonly accepted that inflation convergence in the Euro Area weakened after the advent of the Euro. We have presented compelling evidence that this view is not correct, based on estimates that, to our knowledge, are the first to solely isolate the Euro period. The statistical evidence of group-wise convergence is much stronger and the rate of convergence much faster for the Euro period than for the earlier periods. Finally, we show clear evidence of group-wise inflation convergence for the post Euro period.

References

Afonso, A. and C. Rault, 2007, What do we Really Know about Fiscal Sustainability in the EU? A Panel Data Diagnostic, European Central Bank, WP n. 820.

Andrews, D. W. K., and H-Y. Chen, 1994, Approximately Median-Unbiased Estimation of Autoregressive Models, *Journal of Business & Economic Statistics*, 12 (2): 187-204.

Bai, J. and S. Ng, 2004, A PANIC Attack on Unit Roots and Cointegration, *Econometrica*, 72: 1127-1177.

Ben-David, D., 1993, Equalizing Exchange: Trade Liberalization and Income Convergence, *Quarterly Journal of Economics*, 108(3): 653-79.

Ben-David, D., 1996, Trade and convergence among countries, *Journal of International Economics*, 40(3-4): 279-298.

Bernard, A. B. and S.N. Durlauf, 1995, Convergence in international output, *Journal of Applied Econometric*, 10: 97-108.

Bernard, A. B. and S.N. Durlauf, 1996, Interpreting tests of the convergence hypothesis, *Journal of Econometrics*, 71: 161-174.

Breitung, J. and M. H., Peseran, 2005, Unit Roots and Cointegration in Panels, L. Matyas and P. Sevestre (eds), *The Econometrics of Panel Data: Fundamentals and Recent Developments in Theory and Practice*.

Breuer, J., R. Robert and M. Wallace, 2001, Misleading Inferences from Panel Unit-Root Tests with an Illustration from Purchasing Power Parity, *Review of International Economics*, 9(3): 482-93.

Breuer, J., R. McNown and M. Wallace, 2002, Series-Specific Unit Root Tests with Panel Data, *Oxford Bulletin of Economics and Statistics*, 64(5):527-46.

Busetti F., L. Forni, A. Harvey and F. Venditti, 2007, Inflation Convergence and Divergence within the European Monetary Union, *International Journal of Central Banking, International Journal of Central Banking*, 3(2): 95-121.

Chang, Y. 2004, Bootstrap Unit Root Tests in Panels with Cross-sectional Dependency, *Journal of Econometrics*, 120: 263-293.

Culver, S.E. and Papell, D.H., 1997, Is There a Unit Root in the Inflation Rate? Evidence from Sequential Break and Panel Data Models, *Journal of Applied Econometrics*, 12(4): 435-44.

Davidson, R., and J.G MacKinnon, 1999, The Size Distortion Of Bootstrap Tests. *Econometric Theory*, 15:361-376.

Davidson, R. and J.G. MacKinnon, 2006, Bootstrap methods in econometrics, *Palgrave Handbooks of Econometrics: Volume 1 Econometric Theory*, 812-838.

Durlauf, S. and Quah, 1999, The new empirics of economic growth, Handbook of Macroeconomics, in: J. B. Taylor & M. Woodford (ed.), Handbook of Macroeconomics, ed.1, 1 (4): 235-308.

European Central Bank, 2003, Inflation differentials in the euro area: potential causes and policy implications

Engel, C. and J. Rogers, 2004, European Market Integration After the Euro, *Economic Policy*, 19: 347-384.

Evans, P. and G. Karras, 1996, Do Economies Converge? Evidence from a Panel of U.S. States, *The Review of Economics and Statistics*, 78(3): 384-88.

Evans, P. 1998, Using Panel Data to Evaluate Growth Theories, *International Economic Review*, 39(2): 295-306.

Fleissig, A.R. and J. Strauss, 1999, Is OECD real per capita GDP trend or difference stationary? Evidence from panel unit root test, *Journal of Macroeconomics,* **21**(4), 673-690.

Fleissig, A.R. and J. Strauss, 2001. Panel unit-root tests of OECD stochastic convergence, *Review of International Economics*, 9: 153–162.

Fritsche, U. and V. Kuzin, 2008, Analyzing Convergence in Europe Using a Non-linear Single Factor Model, Macroeconomics and Finance Series, Hamburg University.

Greene, W.H, 2008, Econometric Analysis, Pearson Hall, 6th edition.

Hegwood, N.D., and D.H. Papell, 2007, Are Real GDP Levels Trend, Difference, or Regime-Wise Trend Stationary? Evidence from Panel Data Tests Incorporating Structural Change, *Southern Economic Journal*, 104-113.

Honohan, P. and P.R Lane, 2003, Divergent inflation rates in EMU, Economic Policy, 12 (37): 357-394.

Honohan, P. and P.R Lane, 2004, Exchange Rates and Inflation Under EMU: An Update, CEPR Discussion Papers 4583.

Im, K.S., M.H. Pesaran, and Y., Shin, 2003, Testing for Unit Roots in Heterogeneous Panels, *Journal of Econometrics*, 115(1): 53-74.

Islam, N., 1995, Growth Empirics: A Panel Data Approach**,** *Quarterly Journal of Economics*, 110(4): 1127-70.

Judge, G.G., Hill, R.C., Griffiths, W., Lütkepohl, H. and T. Lee, 1988, Introduction to the Theory and Practice of Econometrics, Wiley, 2nd edition.

Kappler, M., 2006, Panel Tests for Unit Roots in Hours Worked, ZEW Discussion WP 06-022

Kočenda, E. and D.H. Papell, 1997. Inflation Convergence Within the European Union: A Panel Data Analysis, *International Journal of Finance and Economics*, 2(3): 189-198.

Lee, H. Y. and J. L. Wu, 2001, Mean Reversion of Inflation Rates: Evidence from 13 OECD Countries, *Journal of Macroeconomics*, 23: 477-487.

Levin, A., C.F. Lin, and C.J. Chu, 2002, Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties, *Journal of Econometrics*, 108: 1-24.

Lopez, C, 2009, Panel Unit Root with Good Power in Small Samples, *Econometric Reviews*, forthcoming.

Lopez, C, 2008, Evidence of Purchasing Power Parity for the Floating Regime Period, *Journal of International Money and Finance*, 27(1), 156-164.

Lopez, C, 2009, Euro-zone Inflation Rates: Stationary or Regime-wise Stationary Processes, *Economics Bulletin*, 29(1): 238-243

Lopez, C., Murray C.J., and D.H. Papell, 2005, State of the Art Unit Root Tests and Purchasing Power Parity, *Journal of Money Credit and Banking*, 37: 361-369 .

Maddala, G.S. and S. WU, 1999, A comparative Study of Unit Root Tests with Panel Data and a New Simple Test, Oxford *Bulletin of Economics and Statistics*, 61: 631-652.

Murray, C.J., and D.H. Papell, 2005, Do Panels Help Solve the Purchasing Power Parity Puzzle?, 2005, *Journal of Business and Economic Statistics,* 23: 410-415.

O'Connell P., 1998, The Overvaluation of Purchasing Power Parity, *Journal of International Economics*, 44: 1-19.

Papell, D.H., 1997, Searching for Stationarity: Purchasing Power Parity under the Current Float, *Journal of International Economics*, 43: 313-332.

Papell, D.H., 2006, The Panel Purchasing Power Parity Puzzle, *Journal of Money, Credit, and Banking,*38 (2): 447-467.

Papell, D.H., and H. Theodoridis, 2001, The Choice of Numeraire Currency in Panel Tests of Purchasing Power Parity, *Journal of Money, Credit and Banking*, 33: 790-803.

Pesaran, M.H., 2007, A Simple Panel Unit Root Test In The Presence Of Cross Section Dependence, *Journal of Applied Econometrics*, 22 (2): 265-312.

Phillips, P.C.B. and D. Sul, 2003, Dynamic Panel Estimation and Homogeneity Testing Under Cross Section Dependence, *Econometrics Journal*, 6: 217-259.

Rapach, D.E., 2002, Are Real GDP Levels Nonstationary? Evidence from Panel Data Tests," *Southern Economic Journal*, 68 (3): 473-495.

Rogers, J.H., 2007, Monetary Union, Price Level Convergence, and Inflation: How Close is Europe to the USA?, *Journal of Monetary Economics*, 54 (3): 785-796.

Rogers, J.H., Hufbauer, G., Wada, E., 2001. Price level convergence and inflation in Europe, Institute for International Economics, WP 01-1.

Sarno, L. and M.P. Taylor, 2002, Purchasing Power Parity and the Real Exchange Rate, *International Monetary Fund Staff Papers*, 49(1), 65-105.

Taylor, M.P, and L. Sarno, 1998, The behavior of real exchange rates during the post-Bretton Woods period, *Journal of International Economics*, 46 (2): 281-312.

Taylor, A.M. and M.P. Taylor, 2004, The Purchasing Power Parity Debate, *Journal of Economic Perspectives*, 18(4):135-158.

Weber A.A and G.W. Beck, 2005, Price Stability, Inflation Convergence and Diversity in EMU: Does One Size Fit All?, CFS WP 2005/30.

Wu, Y. and H. Zhang, 1996, Mean Reversion in Interest Rates: New Evidence from a Panel of OECD Countries, *Journal of Money, Credit and Banking*, 28(4): 604-21.

Table 1: Finite Sample Performance of the Restricted and Unrestricted ADF-SUR test

DGP: $y_{it} = \rho y_{i,t-1} + u_{i,t} \quad with \quad i = 1,...,N \quad t = 1,...,T \quad and \quad u_{\text{it}} \sim WN$

Estimated model: $\Delta y_{it}^{diff} = c_i + \rho y_{i,t-1}^{diff} + \sum_{j=1}^{k_i} \phi_{ij} \Delta y_{i,t-j}^{diff} + \varepsilon_{\text{it}}$

N	T	ρ=1.00		0.99		0.97		0.95		0.90	
		(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
5	25	0.052	0.052	0.078	0.057	0.127	0.071	0.187	0.087	0.302	0.164
	50	0.051	0.053	0.089	0.067	0. 177	0.102	0.325	0134	0.800	0.455
	100	0.052	0.051	0.125	0.081	0.410	0.225	0.769	0.503	0.999	0.960
	200	0.051	0.052	0.187	0.144	0.865	0.679	1.000	0.988	1.000	1.000
10	25	0.051	0.058	0.081	0.061	0.136	0.073	0.202	0.094	0.435	0.184
	50	0.050	0.052	0.102	0.074	0.248	0.138	0.394	0.267	0.918	0.764
	100	0.050	0.052	0.155	0.112	0.595	0.384	0.942	0.827	1.000	1.000
	200	0.050	0.049	0.357	0.223	0.993	0.947	1.000	1.000	1.000	1.000
20	25	0.049	0.048	0.053	0.049	0.057	0.051	0.062	0.050	0.074	0.054
	50	0.050	0.048	0.106	0.078	0.290	0.167	0.539	0.337	1.000	0.918
	100	0.053	0.053	0.228	0.166	0.814	0.664	1.000	0.990	1.000	1.000
	200	0.051	0.051	0.591	0.373	1.000	0.99	1.000	1.000	1.000	1.000

(1) corresponds to the restricted model that uses $\sum_{i=1}^{N} c_i = 0$, while (2) stands for the unrestricted case. Reading illustration: if (N, T,ρ) = (5, 100, 0.97), the size adjusted power of the restricted ADF-SUR test is of 0.410 compared to 0.225 for the unrestricted case.

Table 2: Finite Sample Power of the Restricted and Unrestricted ADF-SUR test

Mixed processes, T=100

DGP: $y_{it} = \rho_i y_{i,t-1} + u_{i,t}$ with $u_{\text{it}} \sim WN$

Where ρ_i <1 for $i = 1,...,k$ and ρ_j=1 for $j = k+1,...,N$.

Estimated model: $\Delta y_{it}^{diff} = c_i + \rho y_{i,t-1}^{diff} + \sum_{j=1}^{k_i} \phi_{ij} \Delta y_{i,t-j}^{diff} + \varepsilon_{\text{it}}$

N	5							
ρ_i	0.97		0.95		0.90		0.80	
ρ_j=1	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
N-k								
0	0.41	0.23	0.77	0.50	1.00	0.96	1.00	1.00
1	0.22	0.16	0.37	0.31	0.58	0.65	0.70	0.81
2	0.15	0.12	0.21	0.19	0.31	0.36	0.40	0.51
3	0.09	0.09	0.12	0.12	0.16	0.19	0.22	0.26
4	0.07	0.07	0.08	0.08	0.09	0.10	0.11	0.11
5	0.05	0.05	0.05	0.05	0.05	0.05	0.05	0.05

(1) corresponds to the restricted model that uses $\sum_{i=1}^{N} c_i = 0$, while (2) stands for the unrestricted case.

Reading illustration: if *N-k*=2 then the panel is a mix of 2 unit roots and 3 stationary processes. If then ρ_i =0.97, the size adjusted power of the restricted ADF-SUR test is of 0.15 compared to 0.12 for the unrestricted case.

Table 2 (continue): Finite Sample Power of the Restricted and Unrestricted ADF-SUR test

Mixed processes, T=100

DGP: $y_{it} = \rho_i y_{i,t-1} + u_{i,t}$ with $u_{\text{it}} \sim WN$

Where ρ_i <1 for $i = 1,...,k$ and ρ_j =1 for $j = k+1,...,N$.

Estimated model: $\Delta y_{it}^{diff} = c_i + \rho y_{i,t-1}^{diff} + \sum_{j=1}^{k_i} \phi_{ij} \Delta y_{i,t-j}^{diff} + \varepsilon_{\text{it}}$

N	10							
ρ_i	0.97		0.95		0.90		0.80	
ρ_j=1	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
N-k								
0	0.60	0.38	0.94	0.83	1.00	1.00	1.00	1.00
1	0.43	0.33	0.71	0.65	0.92	0.93	0.97	0.98
2	0.32	0.26	0.51	0.50	0.74	0.82	0.86	0.92
3	0.25	0.21	0.38	0.38	0.57	0.66	0.70	0.80
4	0.19	0.17	0.28	0.29	0.42	0.49	0.54	0.64
5	0.15	0.14	0.21	0.22	0.30	0.36	0.39	0.48
6	0.12	0.11	0.15	0.16	0.21	0.25	0.27	0.33
7	0.09	0.09	0.11	0.12	0.15	0.17	0.18	0.22
8	0.08	0.07	0.09	0.09	0.10	0.11	0.12	0.13
9	0.06	0.06	0.07	0.07	0.07	0.08	0.08	0.08
10	0.05	0.05	0.05	0.05	0.05	0.05	0.05	0.05

(1) corresponds to the restricted model that uses $\sum_{i=1}^{N} c_i = 0$, while (2) stands for the unrestricted case.

Reading illustration: if *N-k*=4 then the panel is a mix of 4 unit roots and 6 stationary processes. If then ρ_i =0.90, the size adjusted power of the restricted ADF-SUR test is of 0.42 compared to 0.49 for the unrestricted case.

Table 2 (continue): Finite Sample Power of the Restricted and Unrestricted ADF-SUR test, Mixed processes, T=100

DGP: $y_{it} = \rho_i y_{i,t-1} + u_{i,t}$ with $u_{\text{it}} \sim WN$

Where ρ_i <1 for $i = 1,...,k$ and ρ_j =1 for $j = k+1,...,N$.

Estimated model: $\Delta y_{it}^{diff} = c_i + \rho y_{i,t-1}^{diff} + \sum_{j=1}^{k_i} \phi_{ij} \Delta y_{i,t-j}^{diff} + \varepsilon_{\text{it}}$

N	20							
ρ_i	0.97		0.95		0.9		0.8	
ρ_j=1 *N-k*	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
0	0.81	0.66	1.00	0.99	1.00	1.00	1.00	1.00
1	0.69	0.58	0.95	0.93	0.99	1.00	1.00	1.00
2	0.60	0.52	0.87	0.86	0.98	0.99	0.99	1.00
3	0.52	0.46	0.79	0.79	0.94	0.97	0.98	1.00
4	0.39	0.40	0.69	0.71	0.88	0.94	0.95	0.99
5	0.33	0.36	0.60	0.63	0.80	0.89	0.90	0.97
6	0.29	0.31	0.52	0.54	0.72	0.82	0.84	0.93
7	0.25	0.27	0.44	0.47	0.62	0.74	0.76	0.88
8	0.23	0.24	0.37	0.40	0.53	0.65	0.67	0.81
9	0.21	0.21	0.27	0.34	0.46	0.56	0.58	0.71
10	0.19	0.18	0.25	0.29	0.39	0.48	0.50	0.63
11	0.16	0.16	0.23	0.25	0.32	0.40	0.42	0.53
12	0.14	0.14	0.19	0.20	0.26	0.33	0.35	0.44
13	0.11	0.12	0.16	0.17	0.22	0.26	0.28	0.35
14	0.10	0.11	0.14	0.15	0.18	0.21	0.23	0.27
15	0.09	0.10	0.11	0.12	0.14	0.16	0.18	0.21
16	0.08	0.08	0.09	0.10	0.11	0.13	0.14	0.16
17	0.07	0.07	0.08	0.08	0.09	0.10	0.11	0.12
18	0.06	0.06	0.07	0.07	0.08	0.08	0.08	0.09
19	0.05	0.06	0.06	0.06	0.06	0.06	0.06	0.07
20	0.05	0.05	0.05	0.05	0.05	0.05	0.05	0.05

(1) corresponds to the restricted model that uses $\sum_{i=1}^{N} c_i = 0$, while (2) stands for the unrestricted case.

Reading illustration: if N-*k*=8 then the panel is a mix of 8 unit roots and 12 stationary processes. If then ρ_i =0.90, the size adjusted power of the restricted ADF-SUR test is of 0.53 compared to 0.65 for the unrestricted case.

Table 3: Persistence Measurement: Median Unbiased Estimator (ρ_{MU}) and Half-Life (HL$_{MU}$=ln (0.5)/ln(ρ_{MU}))

$$\Delta y_{it}^{diff} = c_i + \rho y_{i,t-1}^{diff} + \sum_{j=1}^{k_i} \phi_{ij} \Delta y_{i,t-j}^{diff} + \varepsilon_{\text{it}} \text{ , with } i = 1,\dots,N \quad t = 1,\dots,T \text{ and } \varepsilon_{\text{it}} \sim WN$$

	E10					E11				
	ρ	ρ_{MU}	95%CI	HL$_{MU}$	95%CI	ρ	ρ_{MU}	95%CI	HL$_{MU}$	95%CI
Restricted ADF-SUR estimation using $\sum_{i=1}^{N} c_i = 0$										
1982:7-1990:6	0.958	0.970	(0.950; 0.988)	22.76	(13.51; 57.41)	0.952	0.971	(0.921, 0.989)	23.55	(8.42; 62.67)
1990:7-1998:6	0.943	0.975	(0.946; 0.996)	27.38	(12.48; 172.94)	0.957	0.993	(0.965; 0.999)	98.67	(19.46; 692.80)
1999:1-2006:12	0.897	0.950	(0.914; 0.975)	13.51	(7.71; 27.38)	0.895	0.948	(0.905; 0.977)	12.98	(6.94; 29.79)
Unrestricted ADF-SUR estimation										
1982:7-1990:6	0.957	0.979	(0.955; 0.998)	32.66	(15.05; 346.23)	0.952	0.973	(0.947; 1.000)	25.32	(12.72; ∞)
1990:7-1998:6	0.942	0.977	(0.944; 1.000)	29.79	(12.03; ∞)	0.955	0.994	(0.961; 1.000)	172.94	(17.42; ∞)
1999:1-2006:12	0.897	0.957	(0.917; 0. 993)	15.77	(8.00; 98.67)	0.894	0.952	(0.911; 0.987)	14.09	(7.44; 52.97)

Figure 1: Cross-sectional Mean and Standard Deviation

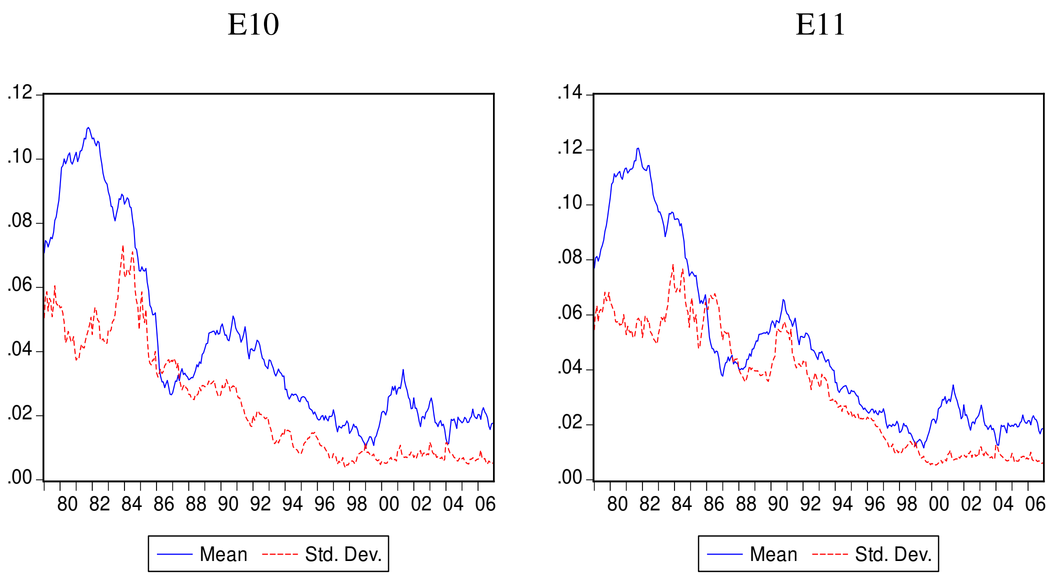


Figure 2: P-values, rolling window from 1979:1-1986:12 to 1999:1-2006:12

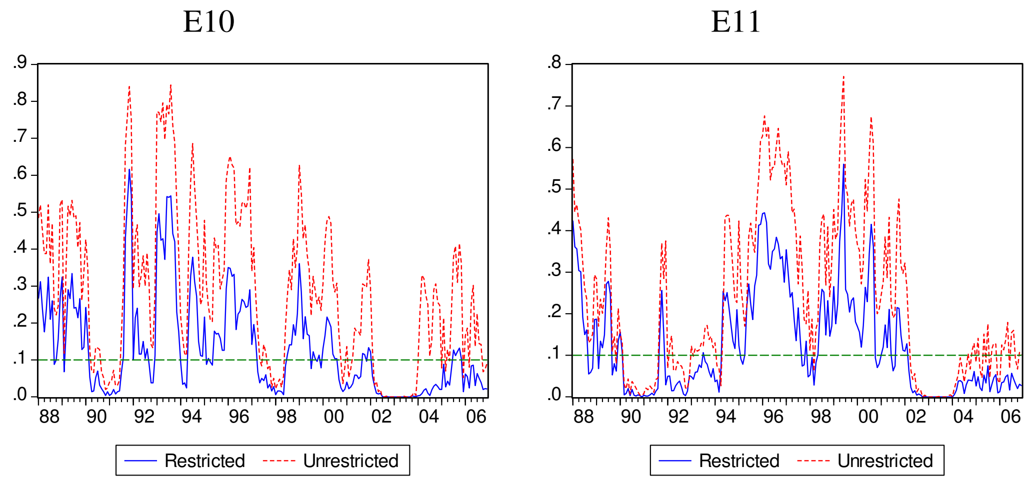
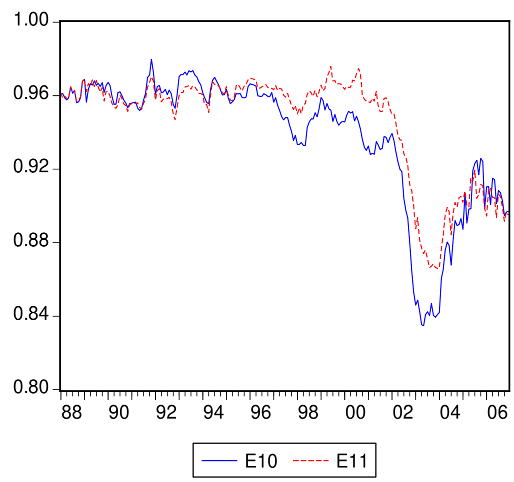


Figure 3: Homogeneous Rate of Convergence, rolling window from 1979:1-1986:12 to 1999:1-2006:12



The x-axes report the end of the period estimated.