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In 2006 all ECB publications will feature a motif taken from the €5 banknote.

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CONTENTS

Abstract

We study the convergence properties of inflation rates among the countries of the European Monetary Union over the period 1980-2004. Given the Maastricht agreements and the adoption of the single currency, the sample can be naturally split into two parts, before and after the birth of the euro. We study convergence in the first sub-sample by means of univariate and multivariate unit root tests on inflation differentials, arguing that the power of the tests is considerably increased if the Dickey-Fuller regressions are run without an intercept term. Overall, we are able to accept the convergence hypothesis over the period 1980-1997. We then investigate whether the second sub-sample is characterized by stable inflation rates across the European countries. Using stationarity tests on inflation differentials, we find evidence of diverging behaviour. In particular, we can statistically detect two separate clusters, or convergence clubs: a lower inflation group that comprises Germany, France, Belgium, Austria, Finland and a higher inflation one with Spain, Netherlands, Greece, Portugal and Ireland. Italy appears to form a cluster of its own, standing in between the other two.

Keywords: Absolute Convergence, Inflation Differentials, Stability, Unit Root Tests.

JEL Classification: C12, C22, C32, E31.

Non-technical summary

Inflation differentials among the countries of the European Monetary Union have shown a tendency to increase after the introduction of the common currency. The cross-country standard deviation of European inflation rates reached its minimum in the second half of 1999, picked up in 2000 and remained relatively stable thereafter; the mean absolute differential between each country's inflation rate and the European average was around half percentage point in 1999 and nearly doubled in 2003. While the slowdown of prices and the converging behaviour of inflation rates were a remarkable success of the process that brought to the adoption of the single currency, the subsequent dynamics of national rates of inflation has raised some concern. Persistent differences in inflation among members of a monetary union may, in fact, lead to disparities in real interest rates, given the common monetary policy. These diversities may be exacerbated by cyclical considerations: a country where economic activity is relatively subdued is likely to have weak inflationary pressures and therefore experience a relatively high real interest rate; this in turn could add further to the divergence of inflation. A different view holds that, in the absence of exchange rate flexibility, inflation differentials are an adjustment mechanism: countries with higher productivity or lower wage growth than others would experience a depreciation of the real exchange rate and thus a gain in trade competitiveness. Part of the differences in inflation could also be due to country heterogeneities in the relative productivity growth of the tradable vs. the non-tradable sector and therefore they might last as long as these persist.

This paper analyzes the convergence properties of inflation rates of euro area countries using monthly Consumer Price Index data from 1980 up to 2004. Given the Maastricht agreements and the adoption of the single currency, the sample can be naturally split into two parts, before and after the birth of the euro. We address two distinct questions regarding the dynamics of national inflation rates in the euro area. The first is whether convergence actually occurred by 1997 (the reference year for the Maastricht criteria to be satisfied in order to qualify for the euro) and whether the Exchange Rate Mechanism (ERM) helped in accelerating the convergence process. The second is whether inflation rates have significantly drifted apart afterwards. We use unit root tests on inflation differentials to address the first issue and stationarity tests on the differentials to address the second one. Unit root tests are mostly useful to establish whether two (or more) variables *are in the process of converging*, with large part of the gap between them depending on initial conditions. Stationarity tests, on the other hand, are the most appropriate tool to verify whether the series *have converged*, i.e. whether the difference between them tends to remain stable. For inflation differentials the main interest is to test the hypothesis of absolute convergence (that is, whether or not inflation differentials were converging to zero by the start of the common monetary policy and whether they tended to drift away from zero in subsequent periods). We also show that, for detecting absolute convergence, it is appropriate to run unit root and stationarity tests without intercept terms, otherwise their lower power might provide spurious evidence for the no convergence hypothesis in the case of unit root tests and for convergence in the case of stationarity tests.

The results of univariate unit root tests for the pre-euro sub-sample show that the countries that joined the ERM from the beginning and never defected from the narrow band displayed strong convergence with each other, while the countries that joined at a later stage (Spain, Portugal and Greece) experienced inflation rates persistently higher than the average. Based on our estimates, the median half-life of the inflation differentials among the early joiners is 3.7 months, while that involving late joiners is 11.2 months. The application of multivariate unit root tests provides evidence that, overall, euro area inflation rates were converging absolutely during the period 1980-1997.

We then use stationarity tests in order to investigate the stability of inflation rates among the member countries over the second sub-sample 1998-2004. We find evidence of diverging behaviour. In particular, we statistically detect three separate clusters, or convergence clubs: a low inflation group that comprises Germany, France, Belgium, Austria, and Finland and a high inflation one with Spain, Greece, Portugal and Ireland. Italy, Netherlands and Luxemburg form a cluster standing in between the previous two. If Luxemburg is excluded from the sample, Italy appears to form a cluster of its own, while Netherlands would cluster with the high inflation group. It is worth emphasizing that we find divergence in the sense that countries belonging to different clusters are characterized by inflation dynamics stable within their group but statistically different from other groups, where this difference may be due to either non stationary behaviour or to different underlying means (or both). Our results indicate that differences in the underlying means may explain the divergence. We interpret this evidence as suggesting that, while the single monetary policy has so far successfully stabilized member countries' inflation rates, a significant degree of country heterogeneity still pervades the euro area.

1 Introduction

Economic integration within Europe has progressed rapidly over the past two decades. In 1979 the European Monetary System was established, in 1990 the Stage I of the European Monetary Union started, leading to Stage II in 1994 and to Stage III (the introduction of the single currency) in 1999. In this context the issue of inflation convergence within the countries of the EMU has received considerable attention. Persistent differences in inflation among members of a monetary union may in fact lead to disparities in real interest rates, given that the policy interest rate is set equal for all member countries.[1]

The differences in real interest rates due to inflation differentials may be exacerbated by cyclical considerations: a country where economic activity is relatively subdued will likely have weak inflationary pressures[2] and therefore will experience a relatively high real interest rate; this in turn could add further to the divergence of inflation.[3] A different view holds that, in the absence of exchange rate flexibility, inflation differentials are an adjustment mechanism: countries with higher productivity or lower wage growth than others would experience a depreciation of the real exchange rate and thus a gain in trade competitiveness; see for example European Central Bank (2003). Still, part of the differences in inflation could be due to country

[1]The real interest rate can be computed subtracting from the nominal rate either actual or expected inflation. Persistent differences in actual inflation are likely to produce significant differences in expected inflation too.

[2]There is a consistent body of evidence that supports the view that inflationary pressures in euro area countries are correlated with the output gap, defined as the difference between actual and potential output. See, for a brief overview on the topic ECB (2003), where it is stated that "empirical estimates appear to suggest that a 1 percentage point increase in the positive output gap leads typically to an increase in the annualised inflation rate of about 15 to 30 basis points in the larger euro area economies" (p. 35).

[3]This argument should be qualified. If inflation differentials are due for example to administered prices, there is no reason to expect that the related differences in real interest rate should lead to different incentives to investment. A similar argument should hold if the differentials are due to different import prices or divergent wage growth while profit margins remain unchanged. It has also been argued that the reference to the national inflation rate in order to compute the real interest rate might not be entirely accurate: if firms sell only in the domestic market, the domestic inflation rate might be a good proxy of average sales prices; but if they sell to all markets in the euro area, it is the euro area inflation that they should use to compute the real interest rate. Therefore, the extent to which inflation differentials at the national level induce differences in the real rate of interest relevant for investment expenditure will depend also on the degree of market integration, which is increasing but still far from complete. Differences in real interest rates, however, would still be effective on other demand components, in particular private consumption.

heterogeneities in the relative productivity growth of the tradable vs. the non-tradable sector (the so called Balassa-Samuelson effect) and therefore they might last as long these persist.

Overall, whether the expansionary effects associated with a real interest rate reduction or the contractionary ones induced by real exchange rate appreciation due to a positive inflation differential[4] would dominate, and the horizon at which this might happen, is an empirical question. The answer will depend to a large extent on the magnitude of inflation differentials and on their persistence.

In this paper we analyze the convergence properties of inflation rates of euro area countries using monthly consumer price index (CPI) data from 1980 up to 2004. Given the Maastricht agreements and the adoption of the single currency, the sample can be naturally split into two parts, before and after the birth of the euro. We address two separate questions regarding the convergence properties of the inflation rates of the euro area countries. The former is whether convergence actually occurred by 1997[5] and whether the Exchange Rate Mechanism (ERM) actually helped in accelerating the convergence process. The latter is whether inflation rates significantly drifted apart after the introduction of the single monetary policy.

Figure 1 shows the year-on-year rate of inflation for each country together with the euro area average. It seems clear that some convergence process has been in action from the early eighties at least until the beginning of the common monetary policy. We study convergence in the pre-euro subsample by means of univariate and multivariate unit root tests on inflation differentials, arguing that the power of the tests is considerably increased if the Dickey-Fuller regressions are run without an intercept term. Overall, we are able to accept the convergence hypothesis and to show that the ERM has played an important role in strengthening the convergence process.

Having obtained evidence in favour of convergence before the start of the common monetary policy, we investigate whether the second sub-sample is characterized by stable inflation rates across the member countries. The year-on-year inflation rates over the period 1998-2004 are graphed in Figure 2: the mean absolute differential between each country's inflation rate and the European average was around half percentage point in 1999 and nearly doubled in 2003. Figure 3 and 4 show the dispersion of inflation in terms of

[4] A symmetric but reversed argument holds for a negative inflation differential.

[5] Among the Maastricht criteria for joining the EMU, each country's inflation in 1997 had to be less than 1.5 percentage points above the average of the three best performers. These turned out to be Austria (with inflation rate as low as 1.2%), Ireland (1.2%) and France (1.3%).

cross country standard deviation and coefficient of variation. It is seen that the standard deviation reached its minimum level in the second half of 1999, picked up in 2000 and remained relatively stable thereafter; a similar pattern is followed by the coefficient of variation. Using stationarity tests on the inflation differentials, we find evidence of diverging behaviour. In particular, we can statistically detect two separate clusters, or convergence clubs: a low inflation group that comprises Germany, France, Belgium, Austria, Finland and a higher inflation one with Spain, Netherlands, Greece, Portugal and Ireland. Italy appears to form a cluster of its own, standing in between the other two. To the high inflation cluster belong countries whose convergence process started rather late in the nineties (i.e. Portugal, Spain and Greece[6]) and in which the move to Stage Three of EMU reduced considerably nominal interest rates (Ireland, Portugal, Spain and, later, Greece), therefore contributing to sustained price dynamics.[7]

It is worth emphasizing that, since stationarity tests (that do not allow for an intercept term) are applied to inflation differentials, each cluster contains inflation rates that are found to be stationary around the same mean. Thus the evidence for divergence over the period 1998-2004 is in the sense that countries belonging to different clusters (or convergence clubs) are characterized by inflation dynamics stable within their group but statistically different from other groups, where the difference may be due to either non stationary behaviour or to different underlying means (or both). In fact, our results suggest that differences in the underlying means may explain the divergence result.

The issue of inflation convergence within European countries has indeed been analyzed in several papers, mainly within the framework of unit root and cointegration tests for panel data. Kocenda and Papell (1997) use panel unit root tests to find evidence in favour of inflation convergence, in particular among the countries participating from the start into the Exchange Rate Mechanism (ERM), and they argue that the convergence process was not substantially affected by the 1992/1993 ERM crises. Siklos and Wohar (1997) run cointegration tests for several European countries to obtain evidence for the presence of a single stochastic trend, a result that

[6]For example, in 1995 (two years before qualifying for the Euro) Portugal and Spain had respectively 4.2% and 4.7% inflation rate, while Greece, that entered two years later, in 1997 recorded an inflation of 5.4%; these numbers must be compared with the 1997 threshold.

[7]Note that in Portugal, Spain and Greece inflation rates have been above the euro area average since 1990. On the other hand, Ireland had negative inflation differentials (relative to the euro average) during most of the '90s.

is consistent with the hypothesis of convergence. Holmes (2002) finds that inflation convergence was strongest during the years 1983-90 whereas the turbulence experienced within the ERM in the early '90s conferred some degree of macroeconomic independence to certain member countries. More recently, Beck and Weber (2003) have performed a beta and sigma convergence analysis of regional inflation data for US, Japan and Europe over the period 1981-2001, showing that inflation dispersion among European regions is higher than in the US or in Japan. Honohan and Lane (2003) argue that the increase in inflation differentials immediately after the start of the single monetary policy is partly due to the differential impact of the depreciation of the euro. Angeloni and Ehrmann (2004) estimate a stylized multi country structural model for the euro area to analyze the response of inflation differentials to a number of different shocks. Their model suggests that the persistence of inflation differentials is mainly determined by the level of inflation persistence at the country level.

Our paper differs from the existing literature in a number of ways. Firstly, we show that, for testing absolute convergence in the rates of inflation, unit root and stationarity tests should be run without intercept terms, otherwise their low power will tend to favour the no convergence hypothesis in the case of unit root tests and the hypothesis of stability in the case of stationarity tests. Secondly, we study convergence of groups of countries by using multivariate versions of unit root and stationarity tests. In particular, the multivariate Dickey-Fuller test, introduced by Abuaf and Jorion (1990) and developed further by O'Connell (1998), Flôres et al (1999) and Harvey and Bates (2003), takes account of the cross-correlation among countries. On the other hand, most of the empirical literature on convergence has used panel unit root tests (such as those proposed in Levin et al., 2002, and Im et al., 2003) that assume cross-country independence, a condition that is unlikely to be satisfied for most macroeconomic variables. Bornhorst (2003) and O'Connell (1998) have investigated the resulting size distortion and power loss of these tests under cross-section dependence and shown that it can be considerable. Thirdly, as we now have a sufficient number of observations, we are able to run formal tests of stability of inflation rates after the inception of the common monetary policy; furthermore, using an algorithm proposed by Hobjin and Franses (2000) in the context of stationarity tests, we are able to identify separate clusters of countries, or convergence clubs, characterized by relatively high and low inflation dynamics.

The paper is organized as follows. Section 2 provides the theoretical background: it describes our definition of convergence and stability and the testing methodology by means of univariate and multivariate unit root and stationarity tests. It also shows the power gains that can be attained by not including an intercept term when testing for absolute convergence. Section 3 describes the results for the "convergence" sub-period (1980-1997), while section 4 explores the issue of whether inflation rates have started drifting apart after the adoption of the single currency. Concluding remarks and a brief summary are contained in Section 5.

2 Convergence and stability of inflation rates: definition and tests

If $\pi_{t,i}$ denotes the series of inflation rate in country i, $i = 1, ..., n$, the convergence properties between countries i and j can be studied from the time series properties of the inflation differential between them,

$$y_t^{i,j} = \pi_{t,i} - \pi_{t,j}, \qquad i, j = 1, ..., n$$

which we call the *contrast* between i and j. In order to simplify the notation we drop the superscript i, j in the remainder of this section.

In the time series literature on convergence there is often some confusion on the role played by unit root and stationarity tests for detecting convergence. The two types of tests are in fact meant for different purposes and they cannot be arbitrarily interchanged. Unit root tests are mostly useful to establish whether two (or more) variables are in the *process of converging*, with large part of the gap between them depending on the initial conditions. Stationarity tests, on the other hand, are the more appropriate tool to verify whether the series *have converged*, i.e. whether the difference between them tends to remain stable. In other words, there is the need to distinguish between *convergence* and *stability*, the former analyzed by testing the null hypohesis of unit root, the latter by testing the null of stationarity.[8] In the subsections below we describe in detail the proper methods to adopt to study the convergence properties of several series.

2.1 Convergence and stability: univariate tests

A suitable model for convergence will be asymptotically stationary, satisfying the condition that

$$\lim_{\tau \to \infty} E(y_{t+\tau} | Y_t) = \alpha, \qquad (1)$$

[8]In this paper we also refer, perhaps with some abuse of terminology, to *divergence* as being associated to rejection of stationarity tests: this is reasonable within our empirical investigation of inflation rates in the post-euro period, when inflation differentials are typically close to zero at the start of the sample and tend to widen thereafter.

where Y_t denotes current and past observations. Convergence is said to be *absolute* if $\alpha = 0$, otherwise it is *relative* (or conditional); see, for example, Durlauf and Quah (1999). The simplest such convergence model is the AR(1) process

$$y_t - \alpha = \phi\left(y_{t-1} - \alpha\right) + \eta_t, \quad t = 1, ..., T, \tag{2}$$

where η_t's are martingale difference innovations and y_0 is a fixed initial condition. By rewriting (2) in error correction form as

$$\Delta y_t = \gamma + (\phi - 1)y_{t-1} + \eta_t, \tag{3}$$

where $\gamma = \alpha(1 - \phi)$, it can be seen that the expected growth rate in the current period is a negative fraction of the gap in the two regions after allowing for a permanent difference, α. We can therefore test for convergence by a unit root test, that is, a test of $H_0 : \phi = 1$ against $H_1 : \phi < 1$. The power of a unit root test will depend on the initial conditions, that is how far y_0 is from α; see for instance Muller and Elliott (2003).

For inflation differentials, the interest in most cases is in testing the hypothesis of absolute convergence. If α is known to be zero, the test based on the Dickey-Fuller $t-$statistic, denoted τ_0 when there is no constant, is known to perform well, with a high value of $|y_0|$ actually enhancing power. The test based on τ_0 is also more powerful, for detecting absolute convergence, than the popular GLS-based alternative of Elliott et al. (1996). Monte Carlo experiments in Harvey and Bates (2003) quantify the power properties of many unit roots tests for different initial conditions; their findings are in line with the arguments of Muller and Elliott (2003).

An AR(p) process provides a natural generalization of (2) that allows for richer dynamics, i.e.

$$\Delta y_t = \gamma + (\phi - 1)\, y_{t-1} + \gamma_1 \Delta y_{t-1} + \gamma_{p-1} \Delta y_{t-p+1} + \eta_t, \tag{4}$$

parameterized in error correction form, with $0 < \phi < 1$. The Augmented Dickey-Fuller (ADF) test is based on such a regression. Again a constant term should not be included if the hypothesis of interest is that of absolute convergence.

As regards stability, following Harvey and Carvalho (2002) we say the *countries i and j have converged* if the inflation differential y_t is a stationary process (with strictly positive and bounded long run variance). Stationarity tests as proposed in Kwiatkowski et al.(1992), denoted as KPSS, Hobjin and Franses (2000) and Busetti and Harvey (2005) are then the appropriate instrument for testing whether convergence has taken place, i.e. for testing stability.

For the case of zero mean stationarity (the most relevant for the analysis of inflation differentials) the test statistic should be computed without demeaning (or de-trending) the series as in KPSS. Thus the stationarity test will reject for large value of

$$\xi_0 = \frac{\sum_{t=1}^{T}\left(\sum_{j=1}^{t} y_j\right)^2}{T^2\widehat{\sigma}_{LR}^2}, \tag{5}$$

where $\widehat{\sigma}_{LR}^2$ is a non-parametric estimator of the long run variance of y_t, that is

$$\hat{\sigma}_{LR}^2 = \hat{\gamma}(0) + 2\sum_{\tau=1}^{m} w\left(\tau, m\right)\hat{\gamma}(\tau), \tag{6}$$

with $w\left(\tau, m\right)$ being a weight function, such as the Bartlett window, $w(\tau, m) = 1-|\tau|/(m+1)$, and $\widehat{\gamma}(\tau)$ the sample autocovariance of y_t at lag τ; the bandwidth parameter m must be such that, as $T \to \infty$, $m \to \infty$ and $m^2/T \to 0$; see Stock (1994). Under the null hypothesis of zero-mean stationarity of y_t, $\xi \xrightarrow{d} \int_0^1 W^2(r)dr$, where $W(r)$ is a standard Brownian motion process; critical values are provided in McNeill (1978), Nyblom (1989), Hobjin and Franses (2000). As showed in Busetti and Harvey (2005), the test based on (5) is powerful also against a non-zero mean stationary process, that is against the hypothesis that the series have converged relatively.

2.2 Convergence and stability: multivariate tests

If interest lies in studying convergence across a group of countries, a multivariate test is appropriate. Let $\mathbf{x}_t$ be the $N = n - 1$ vector of contrasts between each of n countries and a benchmark, e.g. $\mathbf{x}_t = \left(y_t^{1,n},\, y_t^{2,n},\, \ldots,\, y_t^{n-1,n}\right)'$ if the benchmark is the n-th country. The simplest multivariate convergence model is the zero mean VAR(1) process

$$\mathbf{x}_t = \Phi\mathbf{x}_{t-1} + \boldsymbol{\eta}_t, \tag{7}$$

where $\mathbf{\Phi}$ is a $N \times N$ matrix and $\boldsymbol{\eta}_t$ is a N dimensional vector of martingale differences innovations with constant variance $\mathbf{\Sigma}_\eta$. The model is said to be *homogeneous* if $\mathbf{\Phi} = \phi\mathbf{I}_N$. Following Abuaf and Jorion (1990), Harvey and Bates (2003) propose the use of the multivariate unit root test from the homogeneous model. This is given by the Wald-type statistic on $\rho = \phi - 1$, that is

$$\tau_0(N) = \frac{\sum_{t=2}^{T} \mathbf{x}'_{t-1} \widetilde{\mathbf{\Sigma}}_{\eta}^{-1} \Delta \mathbf{x}_t}{\left(\sum_{t=2}^{T} \mathbf{x}'_{t-1} \widetilde{\mathbf{\Sigma}}_{\eta}^{-1} \mathbf{x}_{t-1}\right)^{\frac{1}{2}}}, \tag{8}$$

and referred to as the *multivariate homogeneous Dickey-Fuller* (MHDF) statistic; $\widetilde{\mathbf{\Sigma}}_{\eta}$ is initially estimated by the sample covariance matrix of first differenced data and then re-estimated by iterating the estimation of ϕ to convergence. Under H_0 of $\rho = 0$,

$$\tau_0(N) \xrightarrow{d} \frac{1}{2} \frac{\sum_{i=1}^{n-1} \left(W_i(1)^2 - 1\right)}{\left(\sum_{i=1}^{n-1} \int_0^1 W_i(r)^2 dr\right)^{\frac{1}{2}}}$$

where $W_i(r)$ are independent standard Brownian motion processes, $i = 1, ..., N$; if N is large, $\tau_0(N)$ is approximately Gaussian. The test rejects for $\tau_0(N)$ less than a given critical value, tabulated in Harvey and Bates (2003); see also O'Connell (1998).

One of the attractions of the MHDF test is that it is invariant to pre-multiplication of $\mathbf{x}_t$ by a non-singular $N \times N$ matrix and thus, in our context, it is invariant to which country is chosen as benchmark; such invariance is lost in a heterogeneous model where $\mathbf{\Phi}$ is assumed to be diagonal; see Taylor and Sarno (1998) and Phillips and Sul (2003).

Serial correlation in the innovations can be accounted for by the VAR(p) process

$$\Delta \mathbf{x}_t = (\mathbf{\Phi} - \mathbf{I})\, \mathbf{x}_{t-1} + \mathbf{\Gamma}_1 \Delta \mathbf{x}_{t-1} + ... + \mathbf{\Gamma}_{p-1} \Delta \mathbf{x}_{t-p+1} + \boldsymbol{\eta}_t, \tag{9}$$

written in error correction form. The analogous of the homogeneous model has $\mathbf{\Phi} = \phi \mathbf{I}_N$. In this case the test will be computed by the statistic (8) where $\Delta \mathbf{x}_t$ and $\mathbf{x}_{t-1}$ are replaced by the OLS residuals of regressing each of them on $\Delta \mathbf{x}_{t-1}, ..., \Delta \mathbf{x}_{t-p+1}$. The same limiting distribution and critical values apply.

In the context of stability analysis a generalization of the KPSS test can be applied to $\mathbf{x}_t$ to test whether the n countries have converged. The statistic is now given by

$$\xi_0(N) = Trace\left(\widehat{\mathbf{\Omega}}^{-1} \mathbf{C}\right), \tag{10}$$

where $\mathbf{C} = \sum_{t=1}^{T} \left(\sum_{j=1}^{t} \mathbf{x}_j\right) \left(\sum_{j=1}^{t} \mathbf{x}_j\right)'$ and $\widehat{\mathbf{\Omega}}$ is a non-parametric estimator of the long run variance of $\mathbf{x}_t$ (obtained by a straightforward multivariate extension of (6)). Under the null hypothesis of zero mean stationarity

$\xi_0(N) \xrightarrow{d} \sum_{i=1}^{n-1} \int_0^1 W_i(r)^2 dr$; critical values are provided in Nyblom (1989) and Hobijn and Franses (2000).

The multivariate stationarity test has also invariance properties and thus it does not depend on the benchmark country. Non-rejection of the null hypothesis would imply that the n countries have converged absolutely.

2.3 Convergence Clubs

When the analysis involves more than two countries there is also the possibility that convergence has occurred only for some subsets of them (so-called convergence clubs). A practical approach for identifying the clubs consists of looking at the evidence on all pairwise stability tests, possibly supplied by some knowledge in advance of which countries are expected to behave similarly.

An algorithm for the identification of convergence clubs has been proposed by Hobjin and Franses (2000) in the context of multivariate stationarity tests.[9] This is described in the appendix. If the algorithm is applied using stationarity tests that do not allow for an intercept term, then each club will be formed by series that are found to be stationary around the same mean. The algorithm is independent of the ordering of the series because of the invariance properties of the multivariate stationarity tests on which it is based. However it is not invariant to the number of series in the sample, in the sense that including additional variables may, in general, alter the composition of clusters.

2.4 Power gains from testing without an intercept

When the relevant hypothesis is that of absolute convergence, to enhance power unit root and stationarity tests should be run without allowing for an intercept term. We use the local limiting power functions of the tests, that is the rejection probabilities from small deviations from the null hypothesis of unit root and stationarity respectively, to quantify the power loss from introducing a redundant intercept. These power functions are graphed in Figure 5 for the (augmented) Dickey-Fuller tests and in Figure 6 for the stationarity tests.

More precisely, for the case of unit root tests we consider model (2) with $\alpha = 0$ and $\phi = 1 - \frac{c}{T}$, where c is a fixed non negative constant. The limiting power function is given in terms of the local-to-unit root parameter

[9] Given that these clubs are obtained by grouping the series according to the outcome of stationarity tests, they should perhaps more appropriately be labelled *stability clubs*.

c; $c = 0$ yields the size of the test. Let τ_0 and τ_1 denote the (augmented) Dickey-Fuller t-test statistics computed respectively without and with an intercept. The limiting power is obtained by the asymptotic representation of the statistics under the local-to-unit root assumption that $\phi = 1 - \frac{c}{T}$, that is

$$\tau_i \xrightarrow{d} \frac{\int_0^1 U_i(r;c)dU_i(r;c)}{\left[\int_0^1 U_i(r;c)^2 dr\right]^{\frac{1}{2}}}, \qquad i = 0,1 \tag{11}$$

where $U_0(r;c) = \int_0^r e^{c(s-r)} dW(s)$ is the Ornstein-Uhlenbeck process, with $W(r)$ being a standard Brownian motion, and

$$U_1(r;c) = U_0(r;c) - \int_0^1 U_0(s;c)ds \tag{12}$$

see for example Stock (1994, p.2770-2776). Thus the limiting power function for the tests with α significance level is given by $\Pr\{\tau_i \leq q_i(\alpha)\}$, $i = 1,2$, where $q_i(\alpha)$ is the α-quantile of the distribution of the random variable at the right hand side of (11) for $c = 0$ (that is the Dickey-Fuller distribution); the power is monotonically increasing with c, the magnitude of the deviation from the null hypothesis. Figure 5 shows that the power of the ADF t-test without an intercept, τ_0, is higher by a large amount than that of τ_1. For example, for $c = 10$ the use of τ_0 would allow to achieve a power as high as 0.75 against a value of 0.30 from τ_1.

For the case of stationarity tests, the limiting power functions can be obtained by the data generating process $y_t = \mu_t + \varepsilon_t$, with $\mu_t = \sum_{j=1}^t \eta_j$, where ε_t is a zero mean stationary process with strictly positive and bounded long run variance σ_{LR}^2 and η_t is a white noise process with zero mean and variance equal to $d^2\sigma_{LR}^2/T^2$. Therefore, if $d = 0$, y_t is a stationary process, while if $d > 0$ it contains a local random walk component, where the local nature is due to the variance depending on the sample size T. Here the limiting power is in terms of the parameter d. Let ξ_0 be defined by (5) and ξ_1 be obtained by the same formula replacing y_j by the demeaned observations $y_j - T^{-1}\sum_{t=1}^T y_t$ (that is the OLS residuals from fitting an intercept to the data). The asymptotic representation of the test statistics is given by

$$\xi_i \xrightarrow{d} \int_0^1 V_i(r;d)^2 dr, \qquad i = 0,1 \tag{13}$$

where $V_0(r;d) = W(r) + d\int_0^r W^*(s)ds$,

$$V_1(r;d) = W(r) - rW(1) + d\int_0^r \left(W^*(s) - \int_0^1 W^*(u)du\right) ds, \tag{14}$$

and $W(r)$, $W^*(r)$ are independent standard Brownian motions; see proposition 1 of Busetti and Harvey (2005). The computation of the limiting power for the stationarity tests then proceeds essentially as described for the ADF tests. Figure 6 shows that the advantage of using ξ_0 can be significant, allowing a power increase that can be as large as 0.13. Furthermore, Busetti and Harvey (2005) show that ξ_0 (but not ξ_1) is also powerful and consistent against the case of stationary process with a non-zero mean; in this case the limiting power of ξ_0 is not much lower than that of the Wald t-test on the mean of the observations (and *vice versa* the Wald t-test can be used to detect the presence of a random walk component).

3 Convergence of European inflation rates: 1980.1-1997.12

In this section we analyze the convergence properties of European inflation rates in the pre-euro subsample (1980.1-1997.12). The data are the (monthly) log-differences of the national Consumer Price Indexes (CPI); the source is the Bank for International Settlements.[10] Seasonality has been removed using the STAMP software by Koopman et al. (2000). In general we would expect to see a stronger rejection of the unit root hypothesis in the contrasts between countries that were part of the Exchange Rate Mechanism (ERM).[11] The same data, but over the post-euro subsample, will be the object of the empirical investigation of next section.

The results of the ADF tests on the pairwise contrasts are displayed in the left hand panel of Table 1, in the eight columns jointly labelled *1st subsample:1980.1-1997.12*. The first two columns report the ADF t-test statistic (obtained by the ADF regression without an intercept) and the estimated autoregressive parameter ϕ, respectively. The third column contains the outcome of the ADF test (whether the null hypothesis is rejected at the 1, 5 or 10% significance level or it is not rejected), while the fourth column shows the number of lags in the ADF regression selected according to the modified AIC criterion of Ng and Perron (2001). The following 4 columns

[10] Notice that while figures 1 and 2 show the year-on-year price changes, the tests are computed on the monthly rates of inflation.

[11] The Exchange Rate Mechanism was established by the European Community in March 1979 as part of the European Monetary System (EMS) to reduce exchange-rate variability among member countries. The system was reformed in 1993 to allow for wider fluctuation bands. Spain and Portugal joined in 1989 and 1992 respectively. Austria and Finland joined in 1995 and 1996 respectively, while Greece, although participating to the European Community since 1981, only entered the ERM in 1998.

(5^{th} to 8^{th}) refers to the ADF test with an intercept and it is organized in the same way. The contrasts are ordered by countries according with their GDP weights in the Euro area: Germany (GE), France (FR), Italy (IT), Spain (SP), Netherlands (NE), Belgium (BE), Austria (AU), Greece (GR), Finland (FI), Ireland (IR), Portugal (PT) and Luxemburg (LU).

As a first remark we notice that the null hypothesis of non convergence is rejected much more frequently when the ADF regression is run without allowing for an intercept, which may be a reflection of the power loss from testing with an unnecessary intercept term as seen in the previous section (cf. figure 5). In particular we obtain that, for these series of European inflation differentials, the ADF test without an intercept rejects (at least at 10% significance level) the null hypothesis of non-convergence 58% of the times against only 23% with an intercept. In the following we only comment on the tests without an intercept.

The results of Table 1 have a clearer interpretation when we separate the European countries that joined the ERM since the beginning (Germany, France, Italy, Netherlands, Belgium, Ireland, Luxemburg) from the ones that joined at a later stage (Spain, Portugal, Greece). To allow an even clearer interpretation, we include Austria and Finland in the first group of countries even though they entered the ERM in 1995 and 1996 respectively. This can be justified by the fact that the fluctuations of the Austrian Schilling have consistently been closely related to those of the German mark, while the movement in the Finnish currency significantly departed from those of the German mark only in the last 4-5 years of the sample. Notice that the nine countries in the first group are, in general, characterized by lower inflation than the others.

The evidence in favour of convergence in the group of the lower inflation countries is very strong: the ADF test rejects the null hypothesis at least at 10% significance level for 33 out of 36 inflation differentials. On the other hand, in the remaining contrasts that involve countries of the late joiners group (Spain, Portugal and Greece), the null is rejected at least at the 10% level of significance only in 4 out 30 cases.

To complement the evidence documented in Table 1 we report, in Table 2, the minimum, maximum, median and average estimates of the persistence parameter ϕ. The results that apply to the pre-euro period are contained in the three columns jointly labelled *1*st*subsample:1980.1-1997.12.* For the differentials among ERM members the estimated persistence parameters ranges from 0.14 to 0.94, its median value being equal to 0.83: for monthly data this value corresponds to very fast convergence, as it implies a half life of 3.7 months. For the contrasts involving Spain, Portugal and Greece the estimated persistence parameter ϕ ranges from 0.34 to 0.96 with a median value of 0.94 (and median half life of 11.2 months). In line with previous findings (e.g. Kocenda and Papell, 1997) these results appear to grant an active role to the ERM in speeding up inflation convergence among European countries. In particular, it appears that countries that were not part of the ERM suffered from inflation rates persistently higher than the average, while countries that never defected from the narrow ERM bands displayed stronger convergence with each other.

Finally we consider multivariate evidence for convergence by running the MHDF test on the 11-dimensional vector of all inflation differentials. No matter the number of lags used in the autoregressions, we obtain strong evidence against the null hypothesis for both cases of including and not including an intercept term. Here the evidence for convergence extend to the countries late joiners to the ERM, as a reflection of the higher power properties of multivariate tests.[12]Thus we can conclude that, overall, European inflation rates appeared to be absolutely converging in the pre-euro subsample 1980.1-1997.12.

4 Stability, divergence and clustering: 1998.1-2004.12

The other empirical issue that we want to explore is whether inflation differentials have remained stable since 1998. For this purpose the appropriate instruments are the univariate and multivariate stationarity tests (5) and (10) respectively. We run the tests both without and with an intercept term (the two statistics being ξ_0 and ξ_1, respectively), bearing in mind that not only they have different power properties (as showed in section 2.4) but they also convey different information. The test with an intercept in fact would reject the null hypothesis of stability when inflation differentials display unit root behaviour around a possibly nonzero mean, while without the intercept the test would reject the null if either the differentials contain a unit root or they are stationary around a nonzero mean. Here a rejection of the ξ_0 stationarity test will be taken as evidence for *divergence*, since it implies that inflation differentials, typically very close to zero at the start

[12]Applying the MHDF test to the 2-dimensional vector of inflation differentials among the late joiners to the ERM (Spain, Portugal, Greece) we obtain that the evidence against the null hypothesis is somewhat sensitive to the number of lags included in the autoregressions. In most cases, however, the null is rejected at the 10% level of significance (but often not at the 1%), and, interestingly, the evidence is very similar whether or not we include an intercept term. The power gains from using a multivariate test are not expected to be very significant in this case.

of the post-euro sample, tended to widen thereafter either in a unit root fashion or by stabilizing around a non-zero mean.

The results of the stationarity tests on the pairwise contrasts are displayed in the right hand panel of Table 1, in the four columns jointly labelled *2ndsubsample:1998.1-2004.12*. The first two columns report the values of the statistic ξ_0 and the outcome of the stationarity test without an intercept (whether the null hypothesis is rejected at the 1, 5 or 10% significance level or it is not rejected); the third and fourth columns contain analogous information but for the test with the intercept ξ_1. In all cases the nonparametric spectral estimator of the long run variance is computed for a bandwidth parameter $m = 8$ (that is using autocovariances up to order 8), but the results are very similar for all values of bandwidth between 4 and 12.

A first look at the right hand panel of Table 1 immediately tells that the stationarity test without an intercept rejects the null hypothesis much more frequently (70% of the cases) than the test with an intercept (27%). As already explained, this is coherent not only with the lower power properties of the tests that include a redundant constant term but also with the case of inflation differentials that are stable around a nonzero mean. In the following we will comment on the results for the tests ξ_0, since the main interest is to establish whether inflation differentials have converged absolutely (among all European countries or among subsets of them).

The table shows that there is no evidence for overall stability (around a zero mean) of inflation differentials. However, inflation rates appear to move homogeneously among group of countries, *convergence clubs* (or stability clubs). In particular, the univariate tests of Table 1 show that there is a high degree of stability among the inflation rates of Germany, France, Austria and Finland, countries characterized by relatively low average inflation over the period 1998-2004 (ranging from 1.3% of Germany, in annual terms, to 1.8% in Austria). There also appears a second group of countries, namely Spain, Portugal, Greece and Ireland, where inflation rates are stable but fluctuate around higher levels (from 3.1% of Spain to 3.7% of Ireland).

A formal tool to identify convergence clubs using multivariate stationarity tests is given by the clustering algorithm of Hobjin and Franses (2000), described in the Appendix. Table 3 contains the results of the clustering algorithm applied to the series of m largest countries of the European Monetary Union, with m ranging from 5 to 12. We start by considering Germany, France, Italy, Spain and the Netherlands ($m = 5$), corresponding to around 85% of the euro area GDP. We then progressively add Belgium, Austria, Greece, Finland, Portugal, Ireland, Luxembourg.[13] The tests are computed without fitting an intercept.

Considering the 12 series together ($m = 12$), three convergence clubs are found: a lower inflation group with Germany, France, Belgium, Austria and Finland, a medium group with Italy, Netherlands and Luxembourg, and a higher inflation club with Spain, Greece, Portugal, Ireland. This outcome broadly confirms the finding of the analysis performed using univariate tests. However if Luxembourg is excluded from the sample ($m = 11$) Netherlands turn out to belong to the higher inflation club, and Italy forms a cluster of its own. In particular, Italy stands out by itself for $m = 7, 8, 9, 10, 11,$ while in all cases except $m = 12$ Netherlands belong to the higher inflation club. Figures 7 and 8 graph the average rates of inflation within clusters for $m = 12$ and $m = 11$: the patterns are very similar but it is interesting to see that in the case $m = 12$ average inflation rates never cross each other along the sample.

Thus, overall, there is an indication of divergence of inflation rates in the post-euro subsample. That is, despite the evidence for absolute convergence by the start of the common currency that we obtained in the previous section, it appears that inflation rates began to diverge thereafter (widen apart displaying unit root behaviour or cluster around different average levels). However, it should also be noticed that the persistence of inflation differentials has considerably shrunk since the adoption of the euro. The minimum, maximum, median and average estimates of the persistence parameter, estimated in the post-euro subsample, are reported in the right hand side panel of Table 2, in the three columns jointly labelled *2[nd] subsample:1998.1-2004.12*. The first column refers to the pairwise contrasts involving the countries belonging to the low inflation club (obtained with $m = 11$); the second column considers the inflation differentials involving at least one country of the high inflation club; the third column contains results for all pairwise contrasts. As expected, the estimates of the persistence parameter are lower for countries within the low inflation club and, interestingly, they are also significantly lower than in the pre-euro subsample.

Finally if we apply the multivariate stability test on the 11-dimensional vector of all inflation differentials we obtain that the null hypothesis is clearly rejected when testing without an intercept term while it cannot be rejected

[13] The results are for the period 1998:1-2004:12 and are obtained setting $p^* = 0.05$ and with a bandwidth parameter for spectral estimation set equal to 8. However, very similar output is obtained with the bandwidth ranging between 4 to 12 and with $p^* = 0.01$.

(no matter the value of the bandwidth parameter) if an intercept term is included. Thus, while inflation rates within the EMU can be considered jointly stationary over the period 1998-2004, they appear to fluctuate around different means, forming two or possibly three convergence clubs.

5 Concluding remarks

The paper has provided evidence over the convergence properties of European inflation rates. We have used univariate and multivariate unit root and stationarity tests to show that convergence occurred by the birth of the single currency in 1999. We have provided evidence that the Exchange Rate Mechanism has played a prominent role in favouring convergence among member countries. However inflation rates seem to have begun diverging after 1998. In particular, we have been able to statistically detect two separate clusters, or convergence clubs, over the period 1998-2004, characterized by relatively lower and relatively higher rates of inflation. To the lower inflation club belong Germany, France, Belgium, Austria, Finland, while the higher inflation group comprises The Netherlands, Spain, Greece, Portugal and Ireland. Italy appears to be standing in between the two groups.

Additional empirical results, pertinent to the post-euro sub-sample, were included in an earlier version of the paper, available upon request. By decomposing the changes in the deflators of GDP and final demand, we were able to assess the relative contributions of external factors (such as import prices) and internal factors (mainly wages and productivity) to the inflation differentials observed after 1998. We found that the clusters obtained using the final demand and the GDP deflators closely resemble those obtained in section 4 (based on consumer price indexes) and that these clusters are mainly driven by country differences in the development of per-capita compensations.

Overall, the evidence presented in this paper suggests that while the single monetary policy has, so far, successfully stabilized member countries' inflation rates, a certain degree of cross-country heterogeneity still pervades the euro area.

Appendix: the clustering algorithm for the identification of convergence clubs

Let k_i be set of indices of the series in cluster i, $i \leq n^*$, where $n^* \leq n$ is the number of clusters, and let p^* be a significance level for testing whether some series form a cluster. The algorithm has the following steps:

(1) Initialization: $k_i = \{i\}$, $i = 1, ..., n = n^*$. Each country is a cluster.

(2) For all $i, j \leq n^*$, such that $i < j$, test whether $k_i \cup k_j$ form a cluster (by a multivariate stationarity test on the contrasts) and let $p^{i,j}$ be the resulting p-value of the test. If $p^{i,j} < p^*$ for all i, j then go to step (4).

(3) Replace cluster k_i by $k_i \cup k_j$ and drop cluster k_j, where i, j correspond to the maximum p-value of the previous test (i.e. the most likely cluster); replace the number of clusters n^* by $n^* - 1$. Go to step (2).

(4) The n^* clusters are the convergence clubs.

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Table 1. Unit root tests (first subsample) and stationarity tests (second subsample) on pairwise inflation contrasts

	1^{st} *subsample: 1980.1-1997.12*								2^{nd} *subsample: 1998.1-2004.12*			
	ADF - NO INTERCEPT				*ADF - WITH INTERCEPT*				*KPSS - NO INTERCEPT*		*KPSS - WITH INTERCEPT*	
	Statistic	ϕ	Reject	lags	Statistic	ϕ	Reject	lags	Statistic	Reject	Statistic	Reject
GE - FR	-1.76	0.91	10%	13	-1.50	0.91	-	13	0.47	-	0.25	-
GE - IT	-2.39	0.91	5%	14	-1.85	0.89	-	14	11.43	1%	0.16	-
GE - SP	-1.57	0.92	-	9	-1.49	0.87	-	13	17.14	1%	0.07	-
GE - NE	-2.26	0.54	5%	15	-2.36	0.49	-	15	6.44	1%	0.46	5%
GE - BE	-1.38	0.86	-	13	-1.35	0.85	-	14	2.91	1%	0.13	-
GE - AU	-2.69	0.48	1%	15	-2.82	0.40	10%	15	1.29	10%	0.13	-
GE - GR	-1.15	0.96	-	11	-1.24	0.87	-	14	18.61	1%	0.23	-
GE - FI	-1.83	0.84	10%	11	-1.75	0.82	-	13	0.77	-	0.45	10%
GE - IR	-2.14	0.89	5%	9	-2.03	0.88	-	9	8.65	1%	0.22	-
GE - PT	-1.28	0.95	-	12	-1.22	0.92	-	15	14.58	1%	0.19	-
GE - LU	-1.47	0.86	-	15	-1.49	0.85	-	15	1.76	5%	0.16	-
FR - IT	-2.01	0.88	5%	14	-8.34	0.48	1%	1	9.55	1%	0.60	5%
FR - SP	-1.36	0.88	-	15	-5.45	0.31	1%	5	13.44	1%	0.16	-
FR - NE	-1.67	0.91	10%	13	-1.29	0.91	-	13	3.58	1%	0.56	5%
FR - BE	-3.04	0.62	1%	13	-3.00	0.57	5%	13	1.27	10%	0.41	10%
FR - AU	-1.89	0.87	10%	15	-1.70	0.86	-	15	0.35	-	0.12	-
FR - GR	-0.92	0.96	-	14	-1.75	0.79	-	14	13.17	1%	0.43	10%
FR - FI	-1.72	0.72	10%	14	-1.74	0.71	-	14	0.46	-	0.59	5%
FR - IR	-2.98	0.71	1%	9	-3.08	0.68	5%	9	7.53	1%	0.33	-
FR - PT	-0.97	0.96	-	15	-1.41	0.87	-	15	9.80	1%	0.34	10%
FR - LU	-2.56	0.72	5%	14	-2.57	0.68	-	14	1.51	10%	0.10	-
IT - SP	-4.01	0.34	1%	10	-3.97	0.34	1%	10	2.69	5%	0.14	-
IT - NE	-2.02	0.94	5%	6	-1.55	0.91	-	14	0.71	-	0.36	10%
IT - BE	-2.98	0.83	1%	9	-3.69	0.59	1%	9	1.45	10%	0.07	-
IT - AU	-2.26	0.91	5%	13	-2.07	0.87	-	13	2.43	5%	0.21	-
IT - GR	-0.87	0.95	-	14	-1.97	0.76	-	14	5.80	1%	0.16	-
IT - FI	-2.03	0.83	5%	14	-2.28	0.69	-	14	0.99	-	0.38	10%
IT - IR	-2.33	0.83	5%	13	-3.68	0.63	1%	8	3.62	1%	0.18	-
IT - PT	-1.68	0.89	10%	15	-2.58	0.77	10%	7	4.38	1%	0.11	-
IT - LU	-2.37	0.89	5%	8	-3.00	0.72	5%	14	0.23	-	0.24	-
SP - NE	-1.41	0.94	-	13	-1.18	0.90	-	10	0.21	-	0.40	10%
SP - BE	-1.97	0.85	5%	10	-2.78	0.50	10%	14	4.91	1%	0.10	-
SP - AU	-1.54	0.92	-	13	-1.54	0.84	-	13	11.93	1%	0.09	-
SP - GR	-1.10	0.94	-	11	-2.04	0.69	-	14	0.73	-	0.19	-
SP - FI	-1.50	0.85	-	14	-2.33	0.56	-	14	3.42	1%	0.52	5%
SP - IR	-1.73	0.84	10%	15	-2.57	0.70	10%	15	1.38	10%	0.23	-
SP - PT	-1.54	0.90	-	13	-2.06	0.80	-	13	0.45	-	0.21	-
SP - LU	-1.63	0.88	10%	14	-2.36	0.62	-	14	2.18	5%	0.16	-
NE - BE	-1.48	0.83	-	12	-1.43	0.80	-	14	2.03	5%	0.25	-
NE - AU	-1.76	0.72	10%	15	-2.00	0.58	-	15	2.77	1%	0.48	5%
NE - GR	-1.12	0.96	-	14	-0.90	0.91	-	14	1.32	10%	0.20	-
NE - FI	-1.62	0.87	-	13	-1.42	0.85	-	14	2.64	5%	0.10	-
NE - IR	-2.26	0.90	5%	14	-1.95	0.90	-	14	2.03	5%	0.07	-
NE - PT	-1.09	0.96	-	15	-0.99	0.94	-	15	1.32	10%	0.23	-
NE - LU	-1.69	0.84	10%	14	-1.74	0.80	-	14	0.49	-	0.35	10%
BE - AU	-1.88	0.69	10%	14	-1.93	0.67	-	14	0.39	-	0.25	-
BE - GR	-1.26	0.95	-	11	-1.60	0.80	-	14	7.07	1%	0.12	-
BE - FI	-2.47	0.65	5%	15	-2.47	0.59	-	14	0.17	-	0.35	10%
BE - IR	-3.20	0.76	1%	9	-3.17	0.73	5%	9	6.25	1%	0.18	-
BE - PT	-1.34	0.94	-	12	-1.45	0.88	-	10	4.98	1%	0.09	-
BE - LU	-3.07	0.14	1%	14	-3.06	0.14	5%	14	0.34	-	0.31	-
AU - GR	-1.15	0.96	-	14	-1.20	0.87	-	14	6.83	1%	0.25	-
AU - FI	-1.88	0.78	10%	14	-1.88	0.75	-	14	0.28	-	0.71	5%
AU - IR	-2.19	0.89	5%	14	-1.99	0.89	-	13	8.16	1%	0.32	-
AU - PT	-1.05	0.96	-	15	-1.00	0.93	-	15	7.55	1%	0.26	-
AU - LU	-2.08	0.74	5%	13	-2.02	0.74	-	14	0.88	-	0.12	-
GR - FI	-1.03	0.95	-	13	-2.00	0.72	-	11	4.63	1%	0.21	-
GR - IR	-0.91	0.96	-	15	-1.63	0.83	-	14	0.33	-	0.18	-
GR - PT	-2.17	0.78	5%	11	-2.49	0.70	-	11	0.36	-	0.14	-
GR - LU	-1.14	0.96	-	11	-1.60	0.81	-	14	2.06	5%	0.25	-
FI - IR	-2.49	0.75	5%	10	-2.45	0.74	-	10	12.33	1%	0.18	-
FI - PT	-1.23	0.93	-	12	-1.36	0.86	-	11	5.09	1%	0.24	-
FI - LU	-2.14	0.72	5%	14	-2.20	0.67	-	14	0.41	-	0.64	5%
IR - PT	-1.01	0.95	-	13	-1.59	0.86	-	13	0.75	-	0.12	-
IR - LU	-2.84	0.80	1%	8	-2.91	0.77	5%	8	3.89	1%	0.38	10%
PT - LU	-1.48	0.93	-	7	-1.36	0.88	-	15	1.69	5%	0.24	-

Table 2. Persistence parameters in pairwise inflation differentials

	1st subsample: 1980.1-1997.12			*2nd subsample: 1998.1-2004.12*		
	Early ERM	Late ERM	All	Low Club	High Club	All
Minimum	0.14	0.34	0.14	0.04	0.26	0.02
Maximun	0.94	0.96	0.96	0.88	0.94	0.94
Median	0.83	0.94	0.88	0.39	0.78	0.67
Average	0.78	0.90	0.84	0.42	0.72	0.61

Table 3. Convergence clubs (1998.1-2004.12)

m	*Clusters*		
5	k_1	=	{ Germany, France }
	k_2	=	{ Italy, Spain, Netherlands }
6	k_1	=	{ Germany, France }
	k_2	=	{ Belgium }
	k_3	=	{ Italy, Spain, Netherlands }
7	k_1	=	{ Germany, France, Belgium, Austria }
	k_2	=	{ Italy }
	k_3	=	{ Spain, Netherlands }
8	k_1	=	{ Germany, France, Belgium, Austria }
	k_2	=	{ Italy }
	k_3	=	{ Spain, Netherlands, Greece }
9	k_1	=	{ Germany, France,, Belgium, Austria, Finland }
	k_2	=	{ Italy }
	k_3	=	{ Spain, Netherlands, Greece }
10	k_1	=	{ Germany, France, Belgium, Austria, Finland }
	k_2	=	{ Italy }
	k_3	=	{ Spain, Netherlands, Greece, Portugal }
11	k_1	=	{ Germany, France, Belgium, Austria, Finland }
	k_2	=	{ Italy }
	k_3	=	{ Spain, Greece, Portugal, Ireland, Netherlands }
12	k_1	=	{ Germany, France, Belgium, Austria, Finland }
	k_2	=	{ Italy, Netherlands, Luxembourg }
	k_3	=	{ Spain, Greece, Portugal, Ireland }

Figure 1. Year on year rates of inflation for European countries and EMU average, 1980-2004

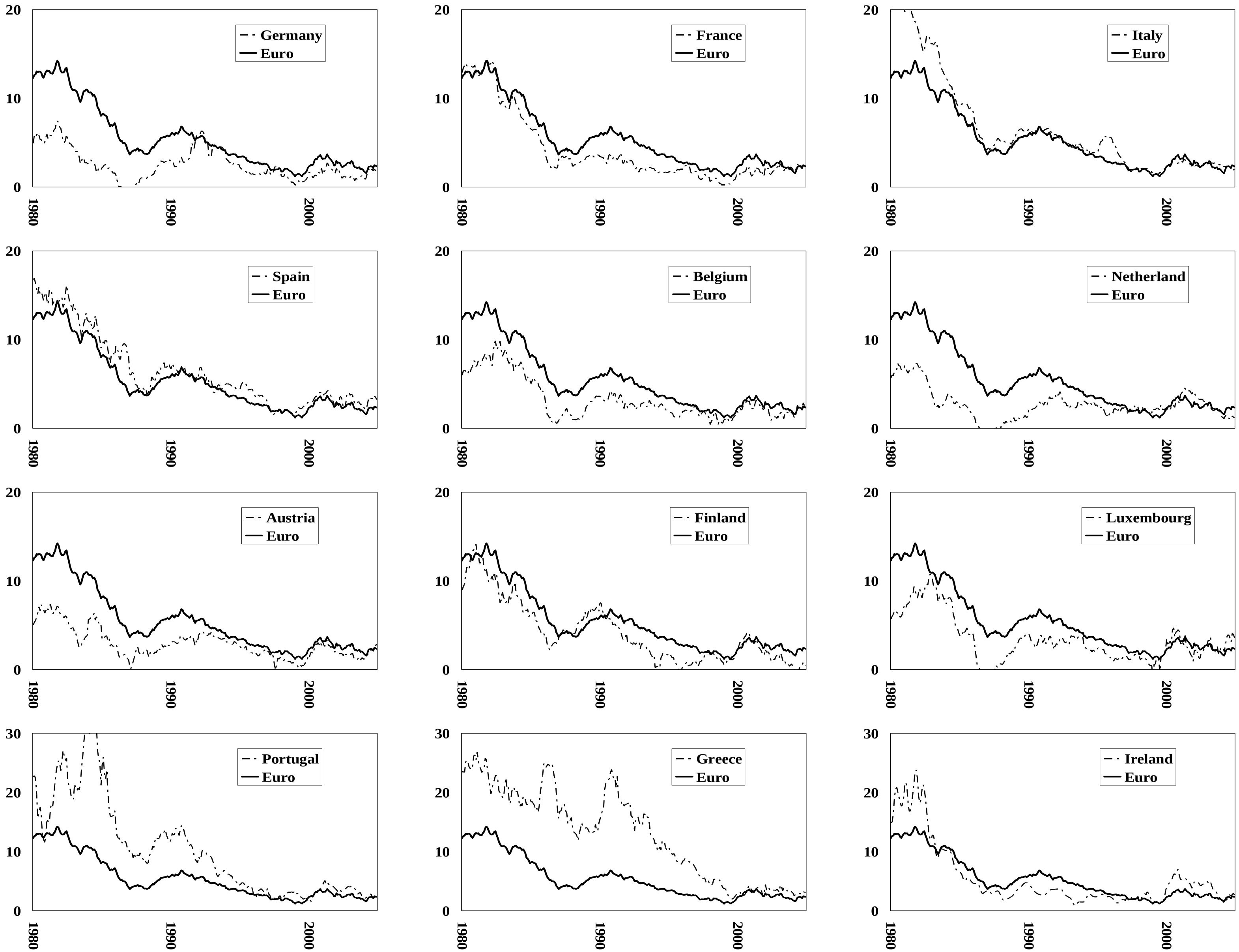


Figure 2. Year on year rates of inflation for European countries and EMU average, 1998-2004

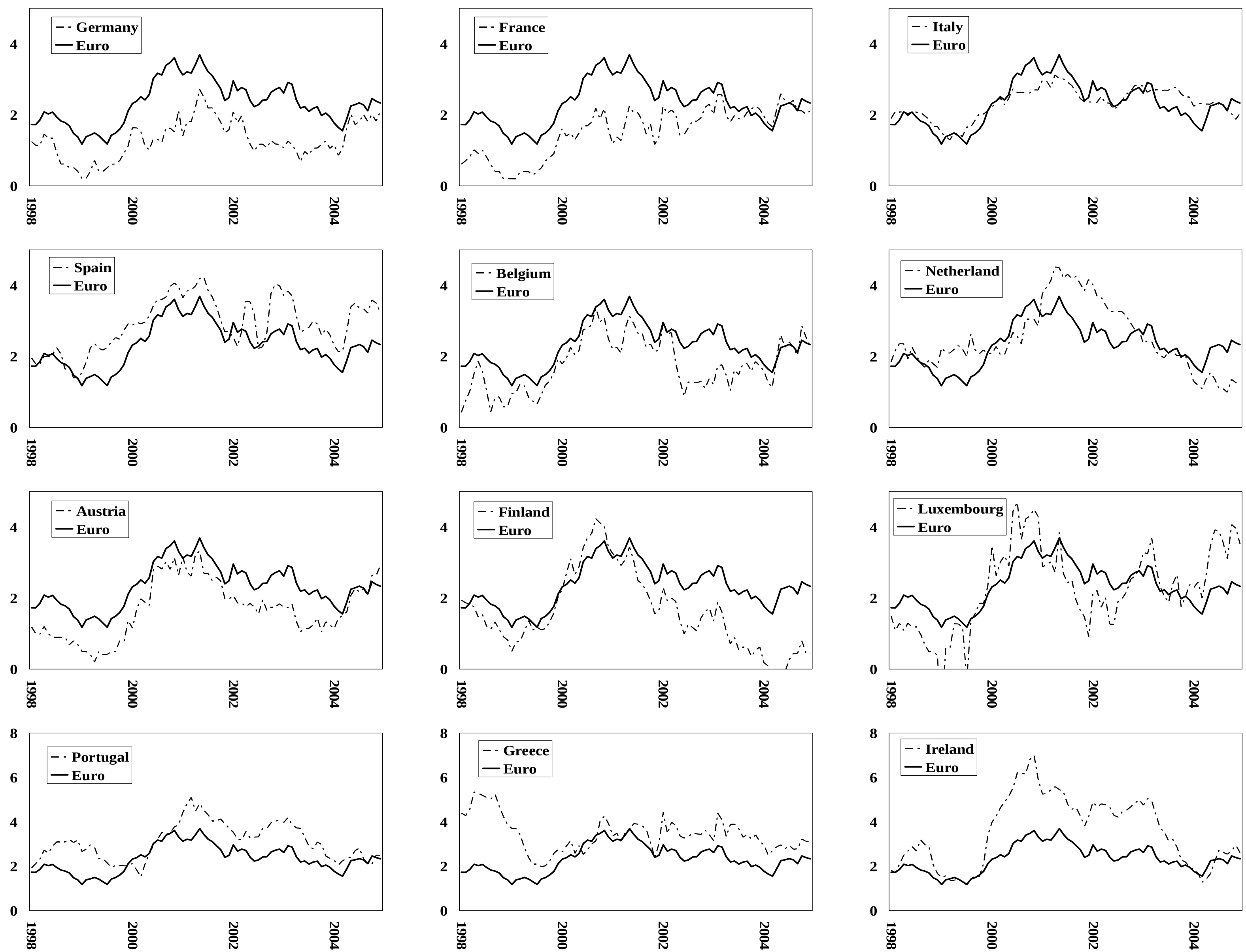


Figure 3. Cross sectional standard deviation of European inflation rates

Figure 4. Cross sectional coefficient of variation of European inflation rates

Figure 5. Limiting power functions of ADF unit root tests

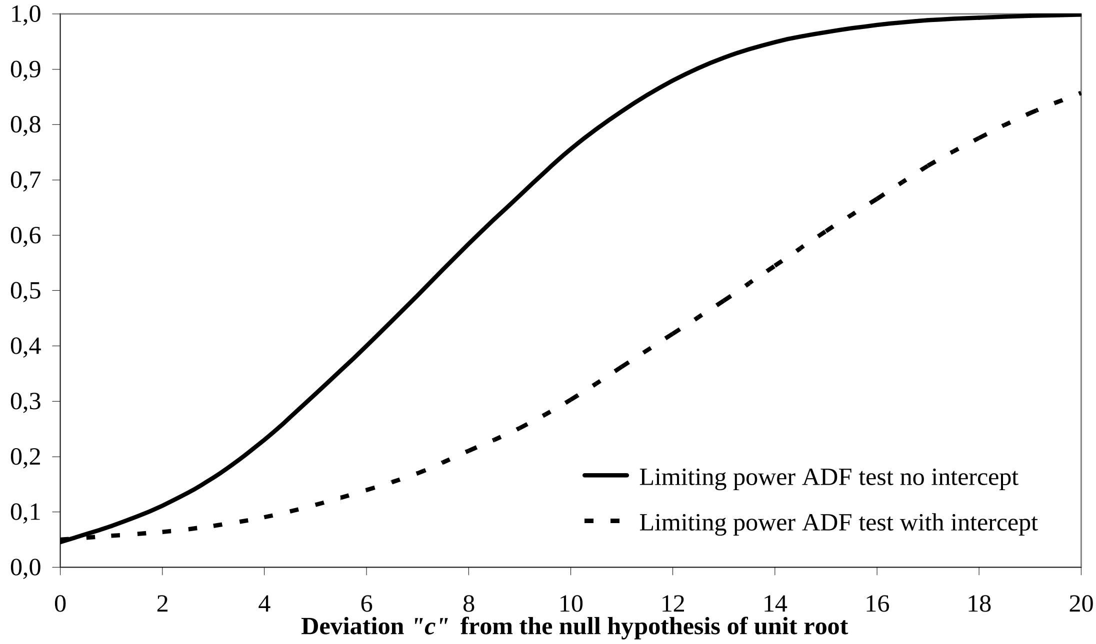


Figure 6. Limiting power functions of stationarity tests

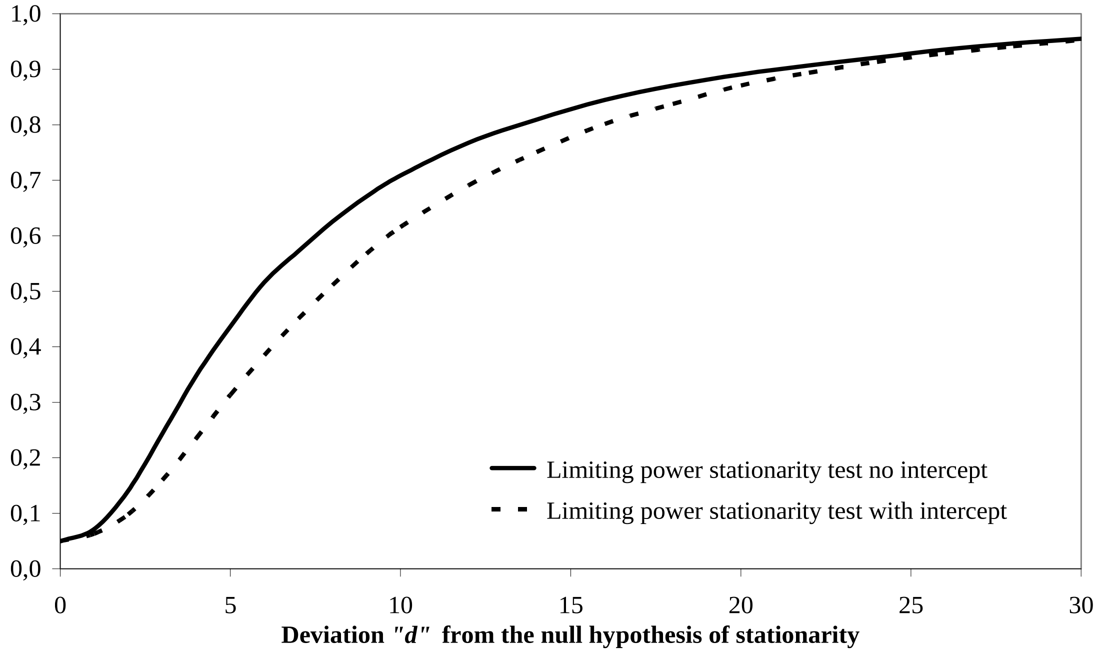


Figure 7. Inflation clusters, EMU countries

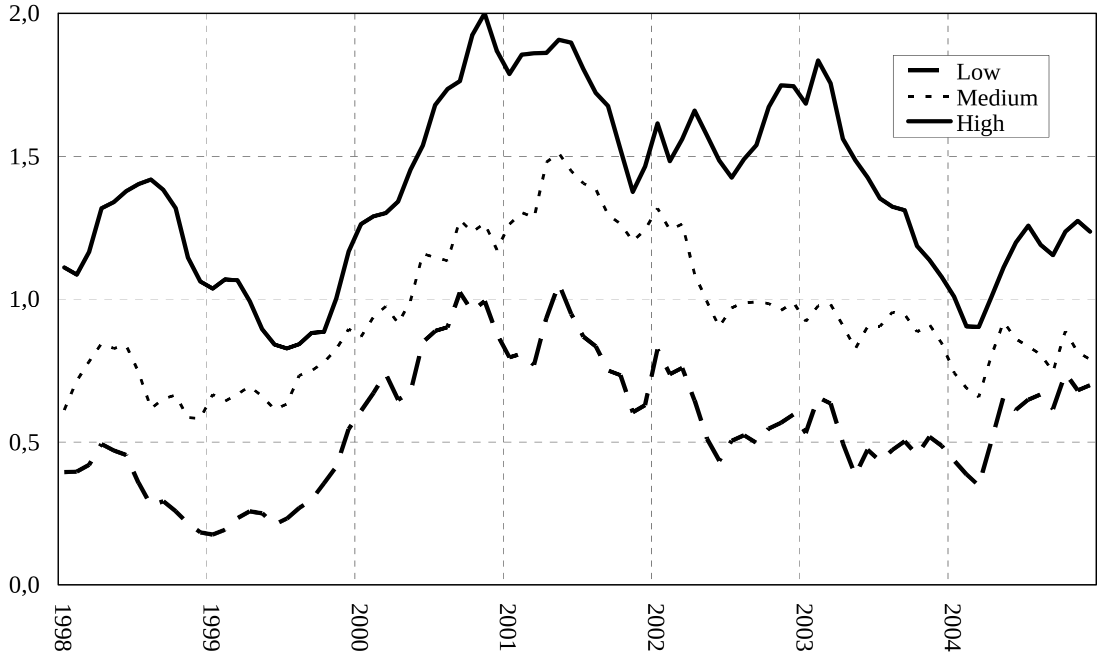
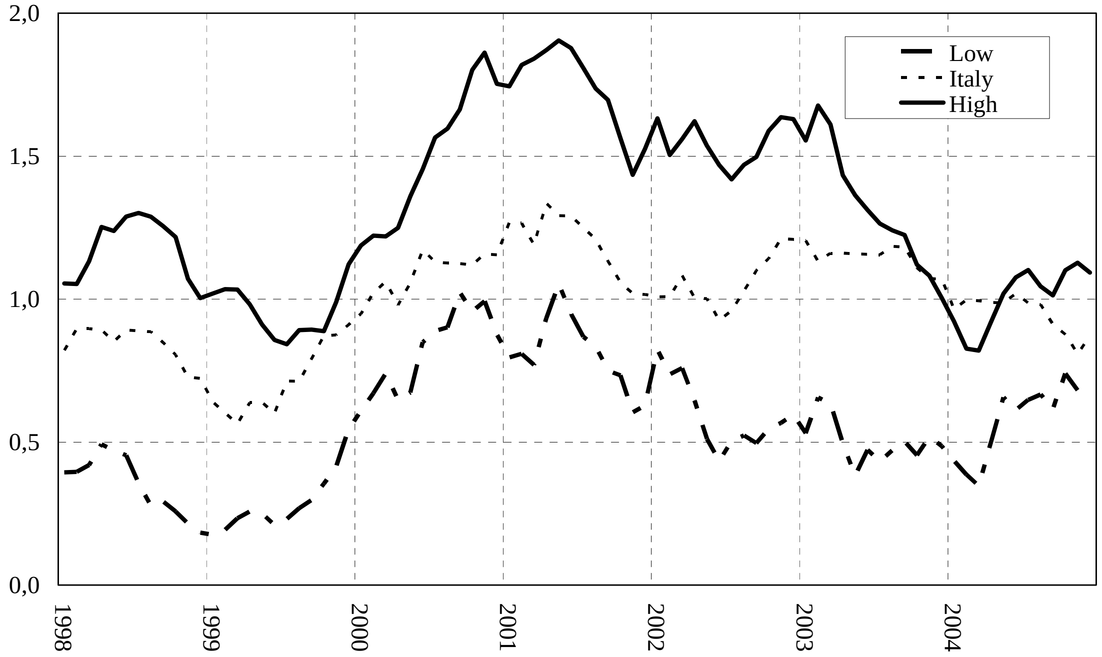


Figure 8. Inflation clusters, EMU countries excluding Luxembourg



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