Inflation Convergence and Divergence Within the European Monetary Union*

Fabio Busetti,[a] Lorenzo Forni,[a] Andrew Harvey,[b] and Fabrizio Venditti[a]
[a]Research Department, Bank of Italy
[b]Faculty of Economics, Cambridge University

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 subsample by means of unit-root tests on inflation differentials, arguing that for testing absolute convergence, a power gain is achieved if the Dickey-Fuller regressions are run without an intercept term. We find evidence for the convergence hypothesis over the period 1980–97 and a clear indication of the important role played by the Exchange Rate Mechanism (ERM) in strengthening the convergence process. We then investigate whether the second subsample is characterized by stable inflation rates across the European countries. Using stationarity tests on inflation differentials, we find evidence of diverging behavior. In particular, we can statistically detect two separate clusters, or stability clubs: (i) a lower-inflation group that comprises Germany, France, Belgium, Austria, and Finland and (ii) a higher-inflation one with Spain, the Netherlands, Greece, Portugal, and Ireland. Italy appears to form a cluster of its own, standing between the other two.

JEL Codes: C12, C22, C32, E31.

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1. Introduction

Inflation differentials among the countries of the European Monetary Union (EMU) 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 a percentage point in 1999 and nearly doubled in 2003. While the slowdown of prices and the converging behavior of inflation rates were remarkable successes of the process that led to the adoption of the single currency, the subsequent dynamics of national rates of inflation has raised some concern. Persistent differences in (actual and expected) inflation among members of a monetary union may 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[1] and therefore experience a relatively high real interest rate; this in turn could add further to the divergence of inflation.[2] On the other hand, in the absence of exchange rate flexibility, inflation differentials may work as an adjustment mechanism:

here are those of the authors, not the Bank of Italy. Andrew Harvey gratefully acknowledges the hospitality and financial support of the Research Department of the Bank of Italy. Author e-mails: Busetti: fabio.busetti@bancaditalia.it; Forni: lorenzo.forni@bancaditalia.it; Harvey: Andrew.Harvey@econ.cam.ac.uk; Venditti: fabrizio.venditti@bancaditalia.it.

[1]See European Central Bank (2003) for a concise overview on the relationship between price dynamics and the output gap in the euro area. In particular, it is argued that for the larger euro-area economies, a 1-percentage-point increase of the output gap leads to a rise of about 15 to 30 basis points in the annualized inflation rate.

[2]This argument should be qualified. If inflation differentials are due, for example, to administered prices, there is no reason to expect that the differences in real interest rate should lead to different incentives to investment. A similar argument holds if the differentials are due to different import prices or divergent wage growth while profit margins remain unchanged. Furthermore, von Hagen and Hoffman (2004) argue that, for firms selling in all euro-area markets, the relevant measure of real rate of interest is based on average euro-area inflation. This would somehow attenuate the effect of heterogeneity of inflation differentials on investment decisions across countries. Differences in real interest rates, however, would still affect other demand components, like private consumption.

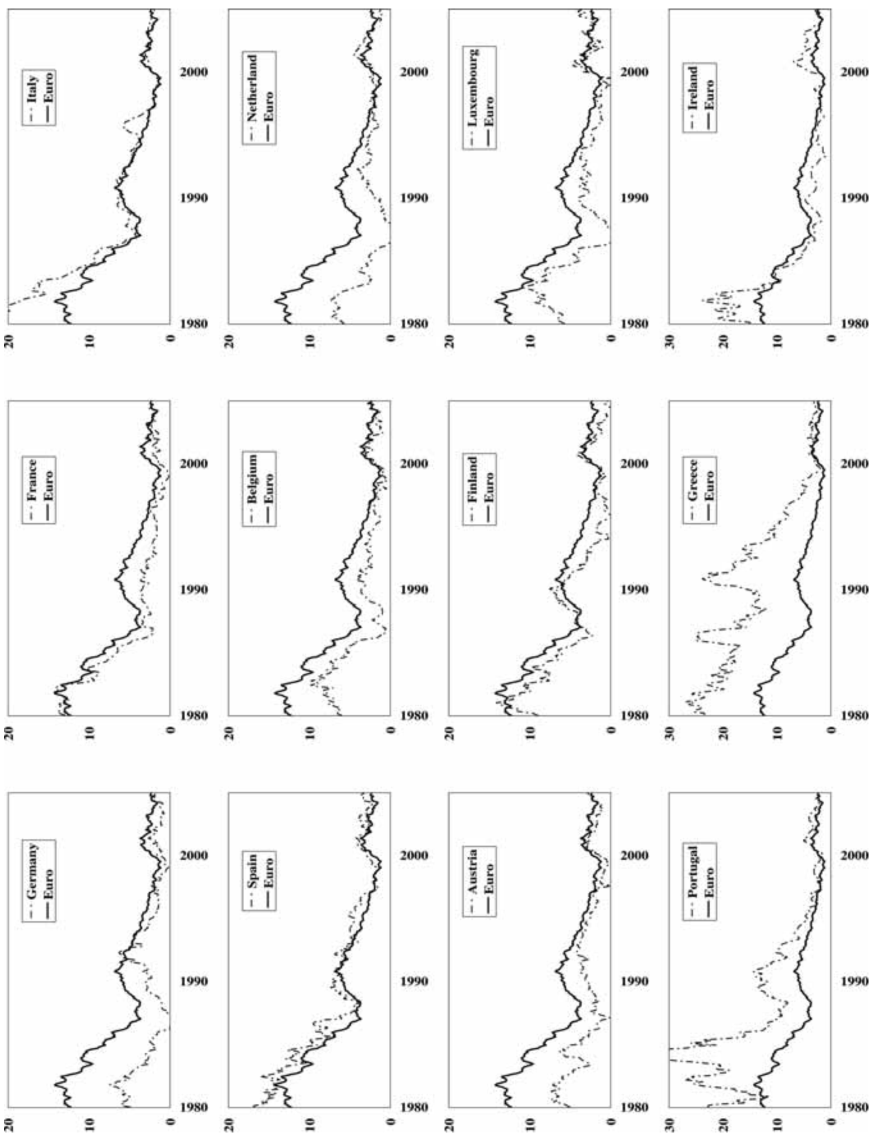
countries with higher productivity or lower wage growth than others would experience a depreciation of the real exchange rate and thus a gain in trade competitiveness. Overall, whether the expansionary effects associated with a real-interest-rate reduction or the contractionary ones induced by real-exchange-rate appreciation due to a positive inflation differential 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. However, part of the differences in inflation could also be due to country heterogeneities in the relative productivity growth of the tradable versus the nontradable sector (the so-called Balassa-Samuelson effect), and therefore they might last as long as these persist.

In this article 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[3] 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 was in action since the early eighties at least until the beginning of the common monetary policy. We study convergence in the pre-euro subsample by 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

[3]One of the Maastricht criteria for joining the EMU required each country's inflation differential (with respect to the average of the three best performers) to be less than 1.5 percentage points in 1997. The three lowest inflation countries turned out to be Austria (1.2 percent), Ireland (1.2 percent), and France (1.3 percent).

Figure 1. Year-on-Year Rates of Inflation for European Countries and EMU Average, 1980–2004



the ERM played an important role in strengthening the convergence process.

Having obtained evidence in favor of convergence before the start of the common monetary policy, we investigate whether the second subsample 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, while figures 3 and 4 show the dispersion of inflation in terms of cross-country standard deviation and coefficient of variation. The tendency for inflation differentials to increase appears quite clear. The minimum for the standard deviation occurred in 1999; the coefficient of variation, on the other hand, reached its lowest levels in the biennium 2000–01 but edged up afterward. Using stationarity tests on the inflation differentials, we find evidence of diverging behavior. From our analysis we can statistically detect two separate clusters: (i) a low-inflation group that comprises Germany, France, Belgium, Austria, and Finland and (ii) a higher-inflation one with Spain, the Netherlands, Greece, Portugal, and Ireland. Italy appears to form a cluster of its own, standing 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[4]) and in which the move to stage 3 of EMU reduced considerably nominal interest rates (Ireland, Portugal, Spain, and, later, Greece), therefore contributing to sustained price dynamics.[5]

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 *stability clubs*) are characterized by inflation dynamics stable within their group but statistically different from other groups, where the difference may be due to either nonstationary

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

[5]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 nineties.

Figure 2. Year-on-Year Rates of Inflation for European Countries and EMU Average, 1998–2004

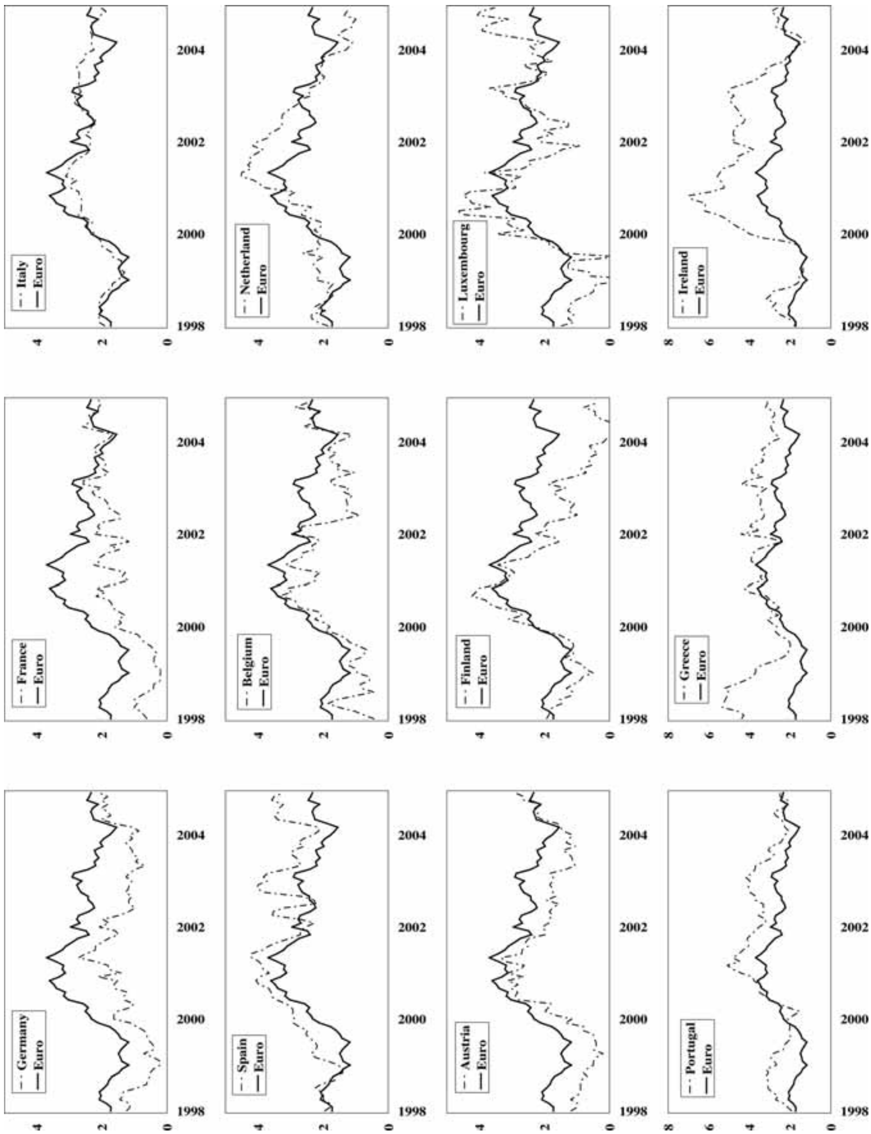
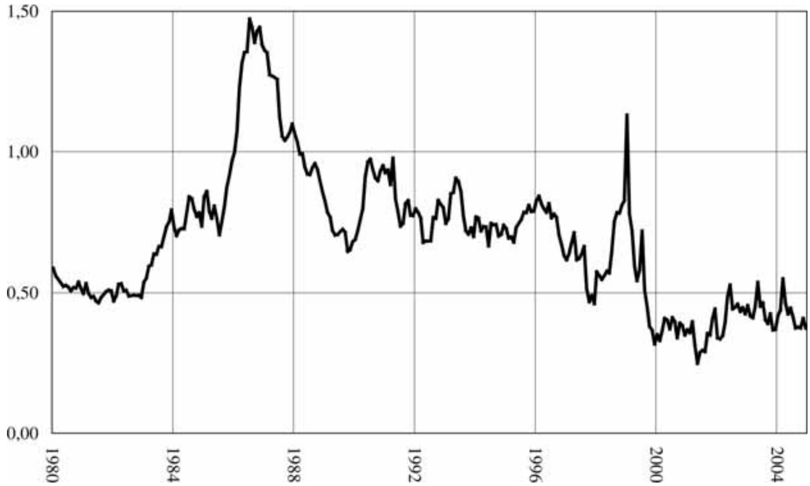


Figure 3. Cross-Sectional Standard Deviation of European Inflation Rates

behavior 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 already 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 and find evidence in favor of inflation convergence, in particular among the countries participating from the start of the 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 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 nineties 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 the United States,

Figure 4. Cross-Sectional Coefficient of Variation of European Inflation Rates



Japan, and Europe over the period 1981–2001, showing that inflation dispersion among European regions is higher than in the United States 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 multicountry 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.

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 unit-root and stationarity tests. Section 3 describes the results for the "convergence subperiod" (1980–97), 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 $i, i = 1, \ldots, n$, the convergence properties between countries i and j can be studied from the time-series properties of the inflation differential between them,

$$y_t^{i,j} = \pi_{t,i} - \pi_{t,j}, \quad i, j = 1, \ldots, n,$$

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 sometimes 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 cannot be arbitrarily interchanged. Unit-root tests are mostly useful for establishing whether two (or more) variables are in the process of converging, with a large part of the gap between them depending on the initial conditions. Stationarity tests, on the other hand, are the more appropriate tool for investigating whether the series have converged—that is, whether the difference between them is stable. In the presence of strong serial correlation—which typically characterizes converging paths—it is known that the actual size of stationarity tests is well above the significance level; see table 3 of Kwiatkoski et al. (1992), henceforth KPSS, and Müller (2005). It is therefore important to distinguish between convergence and stability, the former analyzed by testing the null hypothesis of unit root, the latter by testing the null of stationarity.[6] The subsections below formally describe how to test for convergence and stability and to detect stability clubs; see also Busetti, Fabiani, and Harvey (2006).

[6]In the present paper we also refer, with some abuse of terminology, to *divergence* as being associated with rejection of stationarity tests. However, this seems reasonable within our empirical investigation of inflation rates in the post-euro period, as inflation differentials were typically close to 0 at the beginning of the sample and tended to widen thereafter.

2.1 Convergence and Stability

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

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

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

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

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.

For inflation differentials, the interest in most cases is in testing the hypothesis of absolute convergence. If α is known to be 0, 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, Rothenberg, and Stock (1996). Monte Carlo experiments in Harvey and Bates (2003) and Busetti, Fabiani, and Harvey (2006) quantify the power properties of many unit-root tests for different initial conditions; the findings are in line with the arguments of Müller 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.

Regarding stability, we say that 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 Hobijn and Franses (2000), KPSS, and Busetti and Harvey (2007) are then the appropriate instrument for testing whether convergence has already taken place.

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 detrending) the series as in KPSS. Thus the stationarity test will reject for large values 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 nonparametric estimator of the long-run variance of y_t—that is,

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

with $w(\tau, m)$ 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. The 5 percent and 1 percent critical values are 1.656 and 2.787, respectively; see Nyblom (1989).

If the null hypothesis of interest is stationarity around a nonzero mean, the stationarity test must be computed using the demeaned observations. We will denote the resulting statistic as ξ_1; the limiting distribution is now in terms of a Brownian bridge instead of Brownian motion, with critical values provided in KPSS. Note that the statistic ξ_0 will asymptotically diverge if the data are generated by a stationary process with a nonzero mean; see Busetti and Harvey (2007).

2.2 Multivariate Tests and Detection of Stability Clubs

If interest lies in studying stability and 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., $\mathbf{x}_t = (y_t^{1,n}, y_t^{2,n}, \ldots, y_t^{n-1,n})'$ if the benchmark is the n-th country. Multivariate tests of convergence can be obtained by a generalization of the Dickey-Fuller methodology as proposed in Abuaf and Jorion (1990) and Harvey and Bates (2003); see the working paper version Busetti et al. (2006) for details.

For investigating stability, a generalization of the KPSS test can be applied to $\mathbf{x}_t$ to test whether the n countries have converged. The statistic is

$$\xi_0(N) = Trace(\widehat{\boldsymbol{\Omega}}^{-1}\mathbf{C}), \tag{7}$$

where $\mathbf{C} = \sum_{t=1}^{T}(\sum_{j=1}^{t}\mathbf{x}_j)(\sum_{j=1}^{t}\mathbf{x}_j)'$ and $\widehat{\boldsymbol{\Omega}}$ is a nonparametric 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 is invariant to the benchmark country. Nonrejection of the null hypothesis would imply overall evidence of stability, in the sense that the n countries should have converged absolutely.[7]

Hobijn and Franses (2000) have proposed a clustering algorithm that utilizes a sequence of multivariate stationarity tests to identify *stability clubs*, where each club will be formed by series that are found to be stationary around the same mean. The algorithm, described in the appendix, is independent of the ordering of the series because of the invariance properties of the tests. However, it is not independent of the number of countries included in the sample, and including additional countries may alter the composition of clusters.

[7]If interest lies in relative convergence and stability, the demeaned observations, $\mathbf{x}_t - \overline{\mathbf{x}}$, must be used and the limiting distribution is different; see Hobijn and Franses (2000).

2.3 Power Gain from Testing without an Intercept

When the relevant hypotheses are absolute convergence and stability, to enhance power, unit-root and stationarity tests should be run without allowing for an intercept term. In the working paper version Busetti et al. (2006), the limiting local power function of the tests was used to compute the power gain of the tests without intercept. Furthermore, Busetti and Harvey (2007) show that ξ_0 (but not ξ_1) is also powerful and consistent against a stationary process with a nonzero 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, similarly, the Wald t-test can be used to detect the presence of a random walk component).

3. Convergence of European Inflation Rates: January 1980–December 1997

In this section we analyze the convergence properties of European inflation rates in the pre-euro subsample (January 1980–December 1997). The data are the (monthly) log-differences of the national CPIs; the source is the Bank for International Settlements.[8] Seasonality was removed using the STAMP software of Koopman et al. (2000). In general, we would expect to see a clearer rejection of the unit-root hypothesis in the contrasts between countries that were part of the ERM.[9] The same data, but over the post-euro subsample, will be the object of the empirical investigation of the 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 labeled "Subsample 1: January 1980–December 1997." The first two columns report the ADF t-test statistic (obtained by the ADF

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

[9]The ERM 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 in the European Community since 1981, only entered the ERM in 1998.

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 percent significance level or 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 four columns refer to the ADF test with an intercept and are organized in the same way. The contrasts are ordered by countries according to their GDP—that is, Germany (DE), France (FR), Italy (IT), Spain (ES), the Netherlands (NL), Belgium (BE), Austria (AT), Greece (GR), Finland (FI), Ireland (IE), Portugal (PT), and Luxembourg (LU).

The first point to note is that the null hypothesis of no convergence is rejected much more frequently when the ADF regression is run without an intercept. This is a reflection of the power loss from testing with an unnecessary intercept term as noted in section 2.3. In particular, we find that, at the 10 percent significance level, τ_0 rejects the null hypothesis of no-convergence 58 percent of the time as opposed to only 23 percent for τ_1. In what follows, we only comment on the results for τ_0.

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, the Netherlands, Belgium, Ireland, and Luxembourg) from the ones that joined at a later stage (Spain, Portugal, and 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 four to five 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 favor of convergence in the low-inflation group is very strong: the ADF test rejects the null hypothesis at the 10 percent significance level for thirty-three out of thirty-six inflation differentials. On the other hand, in the contrasts that involve countries of the late-joining group (Spain, Portugal, and Greece), the null is

Table 1. Unit-Root Tests (Subsample 1) and Stationarity Tests (Subsample 2) on Pairwise Inflation Contrasts

	Subsample 1: January 1980–December 1997								Subsample 2: January 1998–December 2004			
	ADF—No Intercept				ADF—With Intercept				KPSS—No Intercept		KPSS—With Intercept	
	Statistic	ø	Reject	Lags	Statistic	ø	Reject	Lags	Statistic	Reject	Statistic	Reject
DE-FR	−1.76	0.91	10%	13	−1.50	0.91	–	13	0.47	–	0.25	–
DE-IT	−2.39	0.91	5%	14	−1.85	0.89	–	14	11.43	1%	0.16	–
DE-ES	−1.57	0.92	–	9	−1.49	0.87	–	13	17.14	1%	0.07	–
DE-NL	−2.26	0.54	5%	15	−2.36	0.49	–	15	6.44	1%	0.46	5%
DE-BE	−1.38	0.86	–	13	−1.35	0.85	–	14	2.91	1%	0.13	–
DE-AT	−2.69	0.48	1%	15	−2.82	0.40	10%	15	1.29	10%	0.13	–
DE-GR	−1.15	0.96	–	11	−1.24	0.87	–	14	18.61	1%	0.23	–
DE-FI	−1.83	0.84	10%	11	−1.75	0.82	–	13	0.77	–	0.45	10%
DE-IE	−2.14	0.89	5%	9	−2.03	0.88	–	9	8.65	1%	0.22	–
DE-PT	−1.28	0.95	–	12	−1.22	0.92	–	15	14.58	1%	0.19	–
DE-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-ES	−1.36	0.88	–	15	−5.45	0.31	1%	5	13.44	1%	0.16	–
FR-NL	−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-AT	−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%

(continued)

Table 1 (continued). Unit-Root Tests (Subsample 1) and Stationarity Tests (Subsample 2) on Pairwise Inflation Contrasts

	Subsample 1: January 1980–December 1997								Subsample 2: January 1998–December 2004			
	ADF—No Intercept				ADF—With Intercept				KPSS—No Intercept		KPSS—With Intercept	
	Statistic	ø	Reject	Lags	Statistic	ø	Reject	Lags	Statistic	Reject	Statistic	Reject
FR-FI	−1.72	0.72	10%	14	−1.74	0.71	–	14	0.46	–	0.59	5%
FR-IE	−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-ES	−4.01	0.34	1%	10	−3.97	0.34	1%	10	2.69	5%	0.14	–
IT-NL	−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-AT	−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-IE	−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	–
ES-NL	−1.41	0.94	–	13	−1.18	0.90	–	10	0.21	–	0.40	10%
ES-BE	−1.97	0.85	5%	10	−2.78	0.50	10%	14	4.91	1%	0.10	–
ES-AT	−1.54	0.92	–	13	−1.54	0.84	–	13	11.93	1%	0.09	–

(continued)

Table 1 (continued). Unit-Root Tests (Subsample 1) and Stationarity Tests (Subsample 2) on Pairwise Inflation Contrasts

	Subsample 1: January 1980–December 1997								Subsample 2: January 1998–December 2004			
	ADF—No Intercept				ADF—With Intercept				KPSS—No Intercept		KPSS—With Intercept	
	Statistic	ø	Reject	Lags	Statistic	ø	Reject	Lags	Statistic	Reject	Statistic	Reject
ES-GR	−1.10	0.94	–	11	−2.04	0.69	–	14	0.73	–	0.19	–
ES-FI	−1.50	0.85	–	14	−2.33	0.56	–	14	3.42	1%	0.52	5%
ES-IE	−1.73	0.84	10%	15	−2.57	0.70	10%	15	1.38	10%	0.23	–
ES-PT	−1.54	0.90	–	13	−2.06	0.80	–	13	0.45	–	0.21	–
ES-LU	−1.63	0.88	10%	14	−2.36	0.62	–	14	2.18	5%	0.16	–
NL-BE	−1.48	0.83	–	12	−1.43	0.80	–	14	2.03	5%	0.25	–
NL-AT	−1.76	0.72	10%	15	−2.00	0.58	–	15	2.77	1%	0.48	5%
NL-GR	−1.12	0.96	–	14	−0.90	0.91	–	14	1.32	10%	0.20	–
NL-FI	−1.62	0.87	–	13	−1.42	0.85	–	14	2.64	5%	0.10	–
NL-IE	−2.26	0.90	5%	14	−1.95	0.90	–	14	2.03	5%	0.07	–
NL-PT	−1.09	0.96	–	15	−0.99	0.94	–	15	1.32	10%	0.23	–
NL-LU	−1.69	0.84	10%	14	−1.74	0.80	–	14	0.49	–	0.35	10%
BE-AT	−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%

(continued)

Table 1 (continued). Unit-Root Tests (Subsample 1) and Stationarity Tests (Subsample 2) on Pairwise Inflation Contrasts

	Subsample 1: January 1980–December 1997								Subsample 2: January 1998–December 2004			
	ADF—No Intercept				ADF—With Intercept				KPSS—No Intercept		KPSS—With Intercept	
	Statistic	ø	Reject	Lags	Statistic	ø	Reject	Lags	Statistic	Reject	Statistic	Reject
BE-IE	−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	–
AT-GR	−1.15	0.96	–	14	−1.20	0.87	–	14	6.83	1%	0.25	–
AT-FI	−1.88	0.78	10%	14	−1.88	0.75	–	14	0.28	–	0.71	5%
AT-IE	−2.19	0.89	5%	14	−1.99	0.89	–	13	8.16	1%	0.32	–
AT-PT	−1.05	0.96	–	15	−1.00	0.93	–	15	7.55	1%	0.26	–
AT-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	–
GRIE	−0.91	0.96	–	15	−1.63	0.83	–	14	0.33	–	0.18	–
GRPT	−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-IE	−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%
IE-PT	−1.01	0.95	–	13	−1.59	0.86	–	13	0.75	–	0.12	–
IE-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

	Subsample 1: Jan. 1980–Dec. 1997			Subsample 2: Jan. 1998–Dec. 2004		
	Early ERM	**Late ERM**	**All**	**Low Club**	**High Club**	**All**
Minimum	0.14	0.34	0.14	0.04	0.26	0.02
Maximum	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

rejected at the 10 percent level of significance in only four out of thirty 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 labeled "Subsample 1: January 1980–December 1997." For the differentials among ERM members, the estimated persistence parameters range from 0.14 to 0.94, with the median 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 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—for example, 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.[10]

[10]Following a reviewer's suggestion, we have checked whether inflation convergence over this subsample has been common to the major industrialized countries. In particular, we have run the tests extending the sample to include seven OECD countries (United States, Canada, United Kingdom, Japan, Denmark, Norway, and Sweden). The results suggest that convergence over the period of analysis has been an international phenomenon.

4. Stability, Divergence, and Clustering: January 1998–December 2004

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 (7). 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 do they have different power properties but they also convey different information. The test with an intercept, in fact, will tend to reject the null hypothesis of stability when inflation differentials display unit-root behavior around a possibly nonzero mean, while without the intercept the test will tend to 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 0 at the start of the post-euro sample, tended to widen thereafter either in a unit-root fashion or by stabilizing around a nonzero 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 labeled "Subsample 2: January 1998–December 2004." 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 percent 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.[11]

A first look at the right-hand panel of table 1 immediately tells us that the stationarity test without an intercept rejects the null hypothesis much more frequently (70 percent of the time) than the test with an intercept (27 percent). As already explained, this is

[11]For these series, it therefore appears that a value of $m = 4$ is sufficient to take care of most of the serial correlation and that the typical power loss induced by adding extra lags is probably negligible.

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 subsequent discussion, we will focus on the results for 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 within groups of countries. 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 percent in Germany, in annual terms, to 1.8 percent in Austria). There is also a second group of countries—namely Spain, Portugal, Greece, and Ireland—where inflation rates are stable but fluctuate around higher levels (from 3.1 percent in Spain to 3.7 percent in Ireland).

We used the clustering algorithm of Hobijn and Franses (2000) to identify, in a formal way, the existence of stability clubs. Table 3 contains the results of the algorithm applied to the series of n largest countries of the European Monetary Union, with n ranging from 5 to 12. We start by considering Germany, France, Italy, Spain, and the Netherlands ($n = 5$), corresponding to around 85 percent of the euro-area GDP. We then progressively add Belgium, Austria, Greece, Finland, Portugal, Ireland, and Luxembourg.[12] The tests are computed without fitting an intercept.

Considering the twelve series together ($n = 12$), three stability clubs are found: (i) a lower-inflation group with Germany, France, Belgium, Austria, and Finland; (ii) a medium group with Italy, the Netherlands, and Luxembourg; and (iii) a higher-inflation club with Spain, Greece, Portugal, and Ireland. This outcome broadly confirms the finding of the analysis performed using univariate tests. However, if Luxembourg is excluded from the sample, so $n = 11$,

[12]The results are for the period January 1998–December 2004 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 $p^* = 0.01$.

Table 3. Stability Clubs (January 1998–December 2004)

n	Identified Clusters
5	k_1 = Germany, France k_2 = Italy, Spain, the Netherlands
6	k_1 = Germany, France k_2 = Belgium k_3 = Italy, Spain, the Netherlands
7	k_1 = Germany, France, Belgium, Austria k_2 = Italy k_3 = Spain, the Netherlands
8	k_1 = Germany, France, Belgium, Austria k_2 = Italy k_3 = Spain, the Netherlands, Greece
9	k_1 = Germany, France, Belgium, Austria, Finland k_2 = Italy k_3 = Spain, the Netherlands, Greece
10	k_1 = Germany, France, Belgium, Austria, Finland k_2 = Italy k_3 = Spain, the Netherlands, Greece, Portugal
11	k_1 = Germany, France, Belgium, Austria, Finland k_2 = Italy k_3 = Spain, Greece, Portugal, Ireland, the Netherlands
12	k_1 = Germany, France, Belgium, Austria, Finland k_2 = Italy, the Netherlands, Luxembourg k_3 = Spain, Greece, Portugal, Ireland

the Netherlands is allocated to the higher-inflation club, and Italy forms a cluster of its own. In particular, Italy stands out by itself for $n = 7, 8, 9, 10, 11$, while in all cases except $n = 12$ the Netherlands belongs to the higher-inflation club. Figures 5 and 6 graph the average rates of inflation within clusters for $n = 12$ and $n = 11$: the patterns are very similar, but it is interesting to see that for $n = 12$ average inflation rates never cross.

Thus statistical evidence points toward divergence of inflation rates since the adoption of the euro. However, the persistence of

Figure 5. Inflation Clusters, EMU Countries

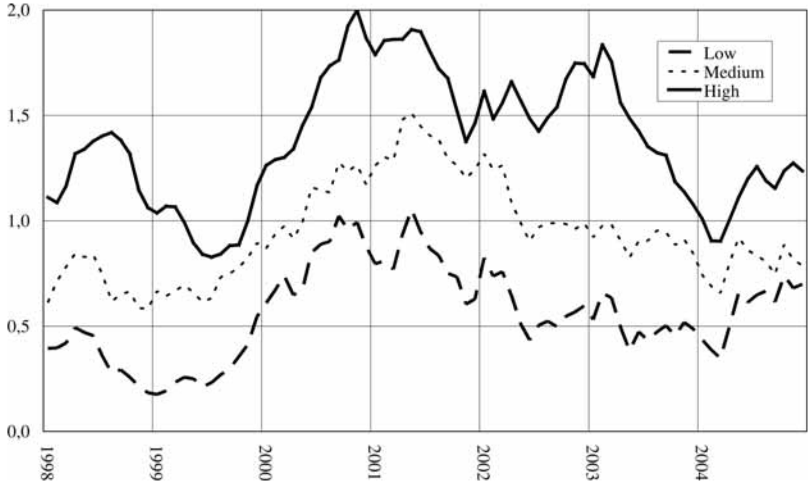
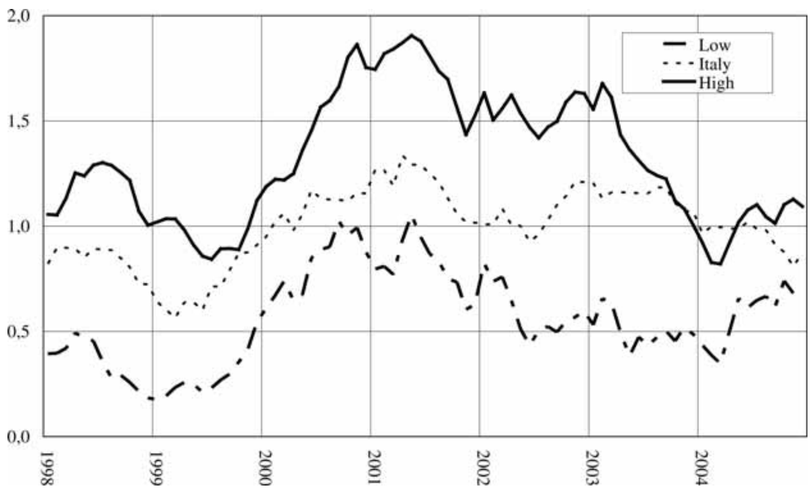


Figure 6. Inflation Clusters, EMU Countries Excluding Luxembourg



inflation differentials has fallen considerably with respect to previous years. The minimum, maximum, median, and average estimates of the persistence parameter, estimated in the post-euro subsample, are reported in the right-hand panel of table 2, in the three columns jointly labeled “Subsample 2: January 1998–December 2004.” The first column refers to the pairwise contrasts involving the countries belonging to the low-inflation club (obtained with $n = 11$); 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 to the vector of all inflation differentials, we find that the null hypothesis is clearly rejected when testing without an intercept term, while it cannot be rejected (for any value of the bandwidth parameter) if an intercept term is included. This is consistent with the idea that 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 stability clubs.

5. Concluding Remarks

We have used unit-root and stationarity tests to show that convergence of European inflation rates had occurred by the birth of the single currency. The Exchange Rate Mechanism seems to have helped convergence. However, inflation rates seem to have begun to diverge after 1998. In particular, we have been able to statistically detect two separate clusters, or stability clubs, over the period 1998–2004, characterized by relatively lower and relatively higher rates of inflation. Germany, France, Belgium, Austria, and Finland belong to the low-inflation club, while the higher-inflation group contains the Netherlands, Spain, Greece, Portugal, and Ireland. Italy appears to stand between the two groups.

Additional empirical results, pertinent to the post-euro subsample, 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 Stability Clubs

Let k_i be a set of indexes 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:

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

(ii) 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 (iv).

(iii) 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 (ii).

(iv) The n^* clusters are the stability clubs.

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