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Testing similarities of short-run inflation dynamics among EU countries after the Euro

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Abstract

In this paper we introduce new definitions of pairwise and multivariate similarity between short-run dynamics of inflation rates in terms of equality of forecast functions and show that in the context of invertible ARIMA processes the Autoregressive distance introduced by Piccolo (1990) is a useful measure to evaluate such similarity. Then, we study the similarity of short-run inflation dynamics across EU-15 area countries during the Euro period. Consistent with studies on inflation differentials and inflation persistence, our findings suggest that after seven years from the launch of the Euro the degree of similarity of short-run inflation dynamics across EU countries is still weak.

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Giulio Palomba *Emma Sarno* *Alberto Zazzaro*

1 Introduction

When optimum currency area conditions are not met, countries belonging to a monetary union must achieve an adequate degree of structural similarity in real and nominal economic quantities in order to be politically and economically stable (Bayoumi and Eichengreen 1993; Feldstein 1997). It was this widely accepted principle which lay behind the convergence criteria fixed by the Maastricht treaty in February 1992 with the aim of forcing less disciplined countries towards the best practices of the more virtuous ones. In particular, to place the European Central Bank (ECB) in the position of effectively implementing a non-inflationary monetary policy, the treaty provided that, in order to enter the monetary union, the inflation rates of member countries should converge to a lower and steady common level[1].

As Figure 1 makes clear, after the creation of the European Monetary System in March

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[1] To be precise, the Maastricht Treaty provided that for an accession country to be eligible for full participation in EMU, its inflation rates, during the three years preceding the beginning of EMU, cannot be more than 1.5 percentage points higher than the average inflation rate of the three less inflationary accession countries.

1979, European Union (EU) countries experienced a steady convergence of their inflation rates both in terms of levels and dispersion. This pattern continued throughout the 90s, but has slightly reversed since the beginning of Stage Three of European Monetary Union (EMU) and the introduction of the Euro when inflation differentials among EMU countries area somewhat increased. This trend was particularly marked for Ireland, the Netherlands and Portugal but also affected other countries like Greece and Spain (ECB 2003; Honohan and Lane 2003).

Descriptive evidence apart, a large number of inferential studies have analyzed inflation convergence within the EU. Obviously, the results reached in these studies vary with the notion of convergence, statistical methodologies, inflation measures, time periods and countries (or regions) considered in the analysis[2]. All in all, however, econometric findings have confirmed the presence of inflation convergence among EU countries towards German and area average levels. This convergence was more intense during the "hard" Exchange Rate Mechanism period (1987-1992). It continued, albeit at a lower pace, after the exchange rate crises in 1992 and 1993 and essentially stopped after the birth of the Euro in 1999, when inflation differentials showed a diverging behaviour and persisting heterogeneity at country and regional level (Mentz and Sebastian 2003; Beck, Hubrich and Marcellino 2006; Busetti, Forni, Harvey and Venditti 2006)[3].

[2]Different approaches for measuring convergence have been followed in the literature: cross-section tests for β- and σ-convergence (Weber and Beck 2003; Ball and Sheridan (2003); Hyvonen 2004); unit-root and stationarity tests (Kočenda and Papell 1997; Holmes 2002; Busetti, Forni, Harvey and Venditti 2006); Johansen tests for the number of cointegrating vectors and common stochastic trends (Amian and Zumaquero 2002; Siklos and Wohar 1997; Westbrook 1998; Mentz and Sebastian 2003); time-varying parameter models (Hall and Wickens 1997; Holmes 1998).

[3]In particular, convergence in price levels, differences in productivity growth between tradable and non-tradable good sectors, different degree of openness towards non-EMU countries, asymmetries in business cycles across economies not yet perfectly integrated have been indicated as major factors responsible for the diverging

Typically, current research steers attention towards the differential of inflation levels across EMU countries and to the asymptotic properties of their inflation series. However, the convergence of inflation rates towards lower and common levels is not sufficient to allow the ECB to take on a common monetary policy, avoiding nationalistic tensions and asymmetric regional effects. In fact, so long as national short-run inflation dynamics remain reciprocally dissimilar, the appropriate stance of monetary policy can differ across State members of the European System of Central Banks (Aksoy, De Grauwe and Dewachter 2002; Benigno 2004; Benigno and López-Salido 2006).

On this point the evidence available is not encouraging. First, while a similarity in the sign of the price responses to monetary policy impulses clearly emerges (Mojon and Peersman 2001; Angeloni and Ehrmann 2004; Benigno and López-Salido 2006), the magnitude of responses varies considerably across EMU countries. Second, the degree of inflation persistence in the Euro area (*i.e.*, the speed of convergence of inflation rates towards some reference level after a shock) appears to be significantly different across countries, although it is moderately low on average and declining over time (Gadzinski and Orlandi 2004; Levin and Moessner 2005; Angeloni, Aucremanne and Ciccarelli 2006).

In this paper we focus on the short-run inflation dynamics of EU countries after the introduction of the Euro by analysing the degree of dissimilarities of their data generating processes (DGPs). Specifically, in the context of univariate time series analysis we estimate ARIMA models for inflation rates of EU-15 area countries in the period 1999-2006. Therefore, for each pair of countries, we evaluate the dissimilarity between their inflation rate DGPs by using the autoregressive (AR) metric introduced by Piccolo (1989, 1990), and test the hypothesis that dissimilarity between AR representations is statistically equal to zero. Finally, we cluster inflation series on the basis of the statistical significance of the dissimilarity test and estimated

dynamics of EMU country inflation rates after the introduction of the Euro. See ECB (2003) and Hofman and Remsperger (2005) for overviews of causes and consequences of inflation differentials in the Euro area.

distances.

The rest of this paper proceeds as follows. In Section 2, we discuss why it matters to have similar short-run inflation dynamics in a currency area. In Section 3, first we provide a definition of similarity of short-run inflation dynamics in terms of equality of forecast functions and then present the notion of AR distance between two time series as a useful measure for such similarity. In Section 4, we outline the statistical test we use to assess the similarity of inflation dynamics. In Section 5, we present results on short-run inflation dynamics in EU-15 and Euro area countries. Section 6 concludes.

2 Why study similarity of short-run inflation dynamics?

According to the Treaty on European Union and the Statute of the European System of Central Banks (ESCB), the primary objective of ECB's monetary policy is to maintain price stability in the Euro area. To this aim, the ECB has established that price stability is guaranteed if the yearly area-wide aggregate inflation rate (in terms of the Harmonised Index of Consumer Prices) is below, but close to, 2% over the medium term. Although ECB's monetary strategy explicitly ignores the short-run inflation dynamics of member countries, dissimilarities in the short-run dynamic properties of inflation series across Euro area countries might *(i)* make the aggregate Euro-wide price index worthless for short-run forecasting inflation rates of Euro area countries and *(ii)* give rise to a different dynamics of short-term real interest rate and affect diversely the optimal monetary policy of each ESCB State member.

In order that the aggregate Euro-wide price index summarizes the behaviour of its components and accurately informs monetary policy, members' inflation series have to share similar short- and long-run dynamic properties (Patell and Zeckauser 1990; Martín-Álvarez, Cano-Fernández and Cáceres-Hernández 1999). Otherwise, the informative power of the aggregate index is weak and its use for policy decisions problematic. In particular, forecasts from Euro-wide models of inflation rates might become misleading and may be different and less accurate than forecasts built by pooling forecasts of country inflation rates (Marcellino, Stock and Watson 2003).

Another reason why differentials of short-run inflation dynamics across Euro area countries matter is that they create potential conflicts within ECB in deciding the common monetary policy and affect the way in which it should be optimally conducted. In this spirit, Aksoy, De Grauwe and Dewachter (2002) analyze how decision procedures within the ECB's Governing Council influence the conduct of common monetary policy and members' welfare, where its propagation mechanisms on output and prices are dissimilar across State members. They show that when nationalistic perspectives prevail within the Governing Council at least partly, differences between the interest rate desired by each State member and the interest rate jointly indicated by the Council can arise, generating high welfare losses, mostly for small countries. In the presence of differentials of short-run inflation dynamics, shared preferences about inflation stabilization among national central bankers are therefore not enough to prevent country-specific positions concerning monetary policy and strained terms within ECB.

Besides nationalistic tensions, asymmetries in the degree of price stickiness across Euro area countries might also have an impact on the optimal monetary policy at the aggregate level. This issue has been recently investigated by Benigno (2004) and Benigno and López-Salido (2006) in a two-region general equilibrium framework with monopolistic competition. In this framework, disparity in the degree of nominal rigidities and inflation dynamics between regions inefficiently affects the terms of trade and the allocation of resources following asymmetric shocks. To minimize distortions and deadweight losses the optimal inflation targeting rule should provide for weighing more the inflation rate of the country where nominal rigidities are stronger. However, such an inflation targeting rule adjusted for country price rigidities might generate destabilizing incentive problems by lowering the urges of sticky countries to introduce reforms aimed at cutting down their nominal rigidities.

3 Measuring similarities of inflation dynamics

Let us begin by introducing two new definitions of pairwise and multivariate similarity of short-run inflation dynamics.

Definition 1 ***(Similarity of short-run inflation dynamics)***. *Countries i and j have similar short-run inflation dynamics if they share the same forecast function, i.e., if inflation forecasts at a fixed time t and at step h, with $h = 1, 2, \ldots, \infty$, are equal for the two countries:*

$$E(y_{i,t+h} - y_{j,t+h} | \mathcal{F}_t) = 0 \qquad \forall\ h \geq 1$$

Definition 2 ***(Similarity of multivariate short-run inflation dynamics)*** *Countries $z = 1, 2, \ldots, n$ have similar short-run inflation dynamics if they share the same forecast function, i.e., if inflation forecasts at a fixed time t and at step h, with $h = 1, 2, \ldots, \infty$, are equal for all z countries:*

$$E(y_{i,t+h} - y_{z,t+h} | \mathcal{F}_t) = 0 \qquad \forall\ i,\ \forall\ z \neq i \text{ and } \forall\ h \geq 1$$

Definitions 1 and 2 extend to the short run the Bernard and Durlauf (1996) definitions of pairwise and multivariate convergence in terms of equality of long-run forecasts at a fixed time. More exactly, when the series have identical initial values, the forecast functions of the two inflation series coincide in the short- as well as in the long-run. Instead, when past values of inflation differ, Definitions 1 and 2 only imply equality of long-run inflation forecast.

Obviously, to make these notions of similarity operational, it is necessary to refer to a forecasting method and to a measure of similarity between statistical time series models. In this paper, we restrict our analysis to the class of invertible ARIMA models. As recent research suggests, traditional univariate linear models show a good short-run forecasting performance for macroeconomic series, which is hardly improvable by more complex multivariate or non-linear models (Meese and Geweke 1984; Canova 2002; Marcellino, Stock and Watson 2003). Moreover, the statistical literature provides a number of criteria for measuring similarities of univariate linear time series models on the basis of their dynamic properties[4]. In particular, a useful measure for evidencing similarity of data generating processes can be constructed from the autoregressive (AR) metric proposed by Piccolo (1989, 1990).

3.1 The AR distance

Consider two mean-zero invertible ARIMA processes $X_{i,t}$ and $X_{j,t}$. In accordance with the classical Box and Jenkins (1976) notation,

$$X_{i,t} \sim \text{ARIMA}(p_i, d_i, q_i)(P_i, D_i, Q_i)$$

if

$$\phi_i(B)\Phi_i(B^s) \bigtriangledown^{d_i} \bigtriangledown_s^{D_i} X_{i,t} = \theta_i(B)\Theta_i(B^s)a_{i,t}, \tag{1}$$

where B indicates the backward operator, $\bigtriangledown = (1 - B)$, s represents the seasonality, $\bigtriangledown_s = (1-B^s)$, $\phi_i(B)$ and $\theta_i(B)$ are polynomials in B of order p_i and q_i respectively, $\Phi_i(B)$ and $\Theta_i(B)$ are seasonal polynomials in B^s of order P_i and Q_i, and finally $a_{i,t}$ is a Gaussian white noise process (of course, an analogous representation holds for $X_{j,t}$).

The AR distance is defined as the Euclidean distance between the sequences of the autoregressive coefficients of the pure AR(∞) representations of $X_{i,t}$ and $X_{j,t}$, which is given

[4]Statistical criteria to evaluate the similarity or dissimilarity between stochastic processes can be distinguished into two broad classes of *(i)* time-domain measures, which the AR metric belongs to (Thomson and De Souza 1985; Peña 1990; Piccolo 1990; Tong and Dabas 1990; Maharaj 2000, 1996), and *(ii)* frequency-domain measures (Shumway and Unger 1974; Alagon (1989); Kakizawa, Shumway and Taniguchi 1998; Caiado, Crato and Peña 2006). For a survey see Corduas (2003) and Caiado, Crato and Peña (2006).

by

$$\pi_i(B)X_{i,t} = a_{i,t} \tag{2}$$

for $X_{i,t}$ (and analogously for $X_{j,t}$), where

$$\begin{aligned}\pi_i(B) &= \phi_i(B)\Phi_i(B^s) \bigtriangledown^{d_i} \bigtriangledown_s^{D_i}\theta_i^{-1}(B)\Theta_i^{-1}(B^s)\\ &= 1 - \pi_{i,1}B - \pi_{i,2}B^2 - \dots\end{aligned}$$

Therefore, in symbols, the AR distance is

$$d(X_{i,t}, X_{j,t}) = \sqrt{\sum_{k=1}^{\infty}(\pi_{i,k} - \pi_{j,k})^2}. \tag{3}$$

Since for invertible processes the coefficients of the AR(∞) representations converge, $d(X_{i,t}, X_{j,t})$ is always a finite number and assumes value zero if and only if $\pi_{i,k} = \pi_{j,k}$ for any k. Moreover, being defined upon the AR(∞) representation, the AR distance is robust to the quasi-cancellation of AR and MA operators and therefore it is not misled in the case of over-parametrization. Finally, contrary to other distances applied in time series analysis, such as the Mahalanobis distance used by Peña (1990), the AR distance does not take into account the white noise variances of the ARIMA processes, which represents a scale factor that does not affect the ARIMA structure of the model.

Except for initial values, the sequence of π-weights fully specifies the dynamic structure of an invertible ARIMA model and thereby the corresponding forecast function. For example, the optimal one step ahead forecast of X_i at time $t-1$ is given by the expectations of $X_{i,t}$ conditional upon its past history,

$$\hat{X}_{i,t} = \sum_{k=1}^{\infty}\pi_{i,k}X_{i,t-k}.$$

Therefore, for identical initial values, the AR distance between two processes is zero if and only if their forecast functions coincide (Piccolo 1990). Where the forecasting method employed in the analysis of inflation rates is based on linear autoregressive models, similarity of short-run inflation dynamics requires that $d(X_{i,t}, X_{j,t}) = 0$, and similarity of multivariate short-run inflation dynamics that $d(X_{i,t}, X_{z,t}) = 0$ for any i and any $z \neq i$.

4 Testing similarity of inflation DGPs

In this section we introduce two useful statistical tests to assess the pairwise similarity of short-run inflation dynamics.

4.1 The similarity test

On real data, an AR distance estimator can be obtained by considering finite versions truncated at lag L of the pure autoregressive representations of two estimated ARIMA processes:

$$\hat{d}(X_{i,t}, X_{j,t}) = \sqrt{\sum_{k=1}^{L} (\hat{\pi}_{i,k} - \hat{\pi}_{j,k})^2}. \tag{4}$$

For inferential purposes, some results are available on the asymptotic properties of the squared AR distance estimator between independent stochastic processes. In particular, for ARMA and ARIMA models based on Maximum Likelihood (ML) estimates, Piccolo (1989) and Corduas (1996) show that the sample distribution of $\hat{d}^2(X_{i,t}, X_{j,t})$ is a linear combination of independent chi-squared random variables, whereas Sarno (2001) shows a similar result in case of Least Squares estimates.

The rationale behind such proofs may be briefly summarised as follows[5]. Consider two independent processes that can be modelled as ARMA, $X_{i,t}$ and $X_{j,t}$. Let $\hat{\theta}_z$ be the ML estimator of the ARMA parameters, for the property of invariance, $\hat{\pi}_z = f(\hat{\theta}_z)$ is the ML estimate of the AR(∞) representation coefficients vector, for $z = i, j$. Hence, as the length of the time series increases, the asymptotic distribution is $\hat{\pi}_z \sim MN(\pi_z, \hat{\Sigma}_z)$, where Σ_z indicates

[5]For details, see the references cited above.

the covariance matrix. Since $f(\hat{\theta}_z)$ is continuous, the covariance matrix Σ_z can be analytically computed as $\Sigma_z = J_z \hat{V}_z J_z'$, where $\hat{V}_z$ is the estimated covariance matrix of $\hat{\theta}_z$ and $J_z = \partial\hat{\pi}_z / \partial\hat{\theta}_z$ is the Jacobian matrix. Then, under the hypothesis $H_0 : \hat{\pi}_i = \hat{\pi}_j$, $\Sigma_i = \Sigma_j = \Sigma$ and $(\hat{\pi}_i - \hat{\pi}_j) \sim MN(0, 2\hat{\Sigma})$, so that, defining $\hat{\eta} = (2\hat{\Sigma})^{1/2}(\hat{\pi}_i - \hat{\pi}_j)$, it follows that $\hat{\eta} \sim MN(0, I)$, where $\hat{\Sigma}$ can be estimated by $0.5(\hat{\Sigma}_i + \hat{\Sigma}_j)$. Therefore, the estimated squared AR(∞) distance is

$$\begin{aligned} \hat{d}^2(X_{i,t}, X_{j,t}) &= 2\hat{\eta}'\hat{\Sigma}\hat{\eta} \\ &= (\hat{\pi}_i - \hat{\pi}_j)'(\hat{\pi}_i - \hat{\pi}_j) \\ &= 2\sum_{h=1}^{r} \lambda_h \chi_1^2, \end{aligned}$$

where λ_h, with $h = 1, 2, \ldots, r$, are the positive eigenvalues of $\hat{\Sigma}$ and $r < p + q + P + Q$.

The value identifying the critical region to reject H_0 at a level α can be obtained by implementing an exact procedure to elicit percentiles of a quadratic form in normal variables (Farebrother 1990). However, the analytical derivation of the matrix may be computationally cumbersome. To simplify the analysis, Corduas (1996) suggests that under H_0 the distribution of $\hat{d}^2(X_{i,t}, X_{j,t})$ can be approximated by a single chi-squared random variable with c degrees of freedom

$$\hat{d}^2(X_{i,t}, X_{j,t}) = a + b\chi_c^2, \tag{5}$$

where parameters a, b and c are evaluated via method of moments estimation. More exactly, recalling that $t_r = 2^r tr(\hat{\Sigma}^r)$, where $\hat{\Sigma}$ represents the covariance matrix of the AR distance, we have $a = t_3/t_2$, $b = t_1 - t_2^2/t_3$ and $c = t_2^3/t_3^2$. As Corduas (1996) shows, approximate and exact critical regions are quite similar and both satisfactory in terms of significance and power far from the non-invertibility region. In this paper we use a test procedure based on the Corduas approximation and implemented in Ox programming[6].

[6]The test procedure is available upon request from the authors. Corduas (2000) presents a similar test procedure in GAUSS.

4.2 The Diebold-Mariano test

The independence between data generating processes involved by the similarity test is a quite demanding assumption in the case of inflation rate series of countries belonging to a currency area. To relax this assumption, we follow the strategy suggested by Otranto and Triacca (2002). First, they show that a necessary condition for $d(X_{i,t}, X_{j,t}) = 0$ is the equal forecastability of processes $Y_{i,t} = (1 - B)^{d_i} X_{i,t}$ and $Y_{j,t} = (1 - B)^{d_j} X_{j,t}$ in the sense of Granger and Newbold (1976), *i.e.*

$$R_i^2 - R_j^2 = \frac{Var(a_{j,t})}{Var(Y_{j,t})} - \frac{Var(a_{i,t})}{Var(Y_{i,t})} = 0.$$

Then, they propose to use a test procedure that compares the predictive accuracy of two different estimated models based on the Diebold and Mariano (1995) test statistic, which does not require the independence of the DGPs, by using as loss functions the quantities R_i^2 and R_j^2.

Formally, given two time series $Y_{i,t}$ and $Y_{j,t}$, the h-step ahead forecast errors conditional upon the information set $\mathcal{F}_t$ are given by $\{a_{i,T+h}|\mathcal{F}_t\}$ and $\{a_{j,T+h}|\mathcal{F}_t\}$ respectively. To determine whether the two models do not have a different forecasting accuracy, the null hypothesis to be tested is

$$H_0 : E(f_t) = E\left[\frac{a_{j,t}^2}{Var(Y_{j,t})} - \frac{a_{i,t}^2}{Var(Y_{i,t})}\right] = 0.$$

Hence, the Diebold-Mariano test statistic is

$$S = \frac{\bar{f}}{\hat{V}(\bar{f})^{1/2}}. \tag{6}$$

A consistent estimator of the asymptotic variance of $\sqrt{T}\bar{f}$ is given by

$$\hat{V}(\bar{f}) = T^{-1}\left[\hat{\gamma}_0 + 2\sum_{k-1}^{g-1} \hat{\gamma}_k\right],$$

where $\bar{f}$ is the average value of f_t, $\hat{\gamma}_k = Cov(f_t, f_{t-k})$ and g is the Barlett lag window. Under the null, Diebold and Mariano (1995) show that $S \sim N(0, 1)$.

5 Similarities of short-run inflation dynamics among EU countries

5.1 Data and sample periods

Our dataset consists of the monthly seasonally unadjusted all-item consumer price index (CPI) from 1999:01 to 2006:12 for each Euro country and for the Euro-wide area. Moreover, we also consider inflation series for Denmark, Sweden and the United Kingdom. The data source is the OECD and they are drawn from Datastream. Inflation rates are computed as the monthly log-differences of CPIs, $p_t = 100[\ln(\mathrm{CPI}_t) - \ln(\mathrm{CPI}_{t-1})]$.

5.2 Modelling short-run inflation dynamics

The first step of our empirical analysis was to remove the seasonal component from the inflation rate series. In this context, we have chosen to treat seasonality as a deterministic component of the series, and therefore we regressed inflation rates against 11 monthly dummy variables. In this way we can model the (de-seasoned) inflation rates without reducing the number of sample observations.

The second step was to assess the degree of integration of our seasonally adjusted series. For this purpose, for robustness, we used four different unit root tests. The standard Augmented Dickey-Fuller test (hereafter ADF) was employed as a benchmark for evaluating the stationarity of inflation rates. The ADF was carried out for two specifications, with and without a constant term[7] (in Table 1, ADF_c and ADF_{nc} respectively). Following Hall (1994), the number of lags was selected according to a general-to-specific approach.

The third test we used is the Elliott, Rothenberg and Stock (1996) efficient DFGLS test for

[7]The specifications with the linear and quadratic trend were omitted because preliminary graphical inspection of series excluded the presence of significant patterns.

an autoregressive unit root. This test is similar to the ADF test, but reveals best performance and higher power in small samples. Once again, Hall's generic-to-specific approach was used for the lag selection.

Finally, we applied the KPSS test (Kwiatkovski, Phillips, Schmidt and Shin 1992). This is a nonparametric test in which the null hypothesis is the stationarity of the time series. The KPSS test statistic has a non-standard asymptotic distribution; it was computed without including a time trend in the estimated average value of inflation rates and by using 16 lags[8].

In Table 1 we report the outcome of these tests and the critical values for the DFGLS and KPSS tests. On the whole, there is robust evidence in favour of the hypothesis of stationarity of inflation rate series. Specifically, the ADFs and the KPSS tests always support the hypothesis of stationarity, with the only exception of the ADFc for Finland. The DFGLS test rejects the null hypothesis for seven countries (AUT, BEL, GER, EIR, NED, DEN and UKD). For the remaining countries the test accepts the hypothesis of a unit root at a confidence level of 10%, even if the test statistic is quite similar at the critical value of -1.6175 for Finland and Portugal.

Once it was checked the stationarity of seasonally adjusted inflation series, the third step of our analysis was estimation of ARMA models for EU-15 area countries and the Euro-wide inflation rate. By using adjusted inflation rates for seasonality, we avoid the risk of estimating overdifferentiated time series and we can also assume that inflation rates follow a stationary ARMA(p,q) model. Hence, equation (1) reduces to the form

$$\phi_i(B)X_{i,t} = \theta_i(B)a_{i,t}. \tag{7}$$

The lag selection was carried out by minimising the Hannan and Quinn (1979) information criterion. Since the AR(∞) representation does not admit for the constant term in the model estimation (Piccolo 1990), we use the deviations of inflation rates from their average values.

[8]The outcome of the KPSS test does not change when the number of lags is augmented to take autocorrelations of higher order into account.

For each country, Table 2 reports the estimated ARMA models, the log-likelihood, the number of the estimated parameters and the values of the Bayesian Information Criterion (BIC) and the Hannan-Quinn Information Criterion (HQC). It can be worthwile to note that, in spite of the seasonal adjustment, the models selected for Belgium, Denmark, the Netherlands and Sweden require high number of lags in the MA component.

5.3 AR distances and pairwise similarity of short-run inflation dynamics

For each estimated ARMA model, we derive the AR(L) representation, with $L = 200$, and compute the AR distance for each pair of series considered. Table 3 reports the values of estimated AR distances, while Table 4 displays the results of the similarity test based on $\hat{d}^2(X_{i,t}, X_{j,t})$ and the Diebold-Mariano test.

On the whole, our results show the presence of a high degree of dissimilarity of inflation dynamics among EU countries. Assuming a critical region at level 0.05 (0.1), for only 20 (11) pairs of comparisons out of 120 the similarity test indicates that the structural dynamics of short-run inflation rates are not significantly dissimilar.

When we limit our attention to the twelve Euro area countries, the number of cases in which the null of zero AR distance is accepted drastically reduces to 12 (5). The AR distance calculated between Greece and Italy is the lowest (0.14), while it assumes the largest values for the pairs France-Luxembourg (1.03), Austria-Spain (0.96) and Luxembourg-Spain (0.95). The DGP of the Euro-wide area inflation rate mimics the inflation dynamics of Finland with good statistical significance (the similarity test accepts the null at 10% confidence level). It is similar to Belgium, Germany, Greece and Portugal DGPs at the level 0.01, while it is largely dissimilar to the inflation dynamics prevailing in Austria, France, Ireland, Italy, Luxembourg and Spain.

Rather surprisingly, EU countries that have preserved their national currency (Denmark, Sweden and the United Kingdom) display a high degree of similarity of their own short-run inflation dynamics with Euro area countries. In particular, the AR distance between inflation series of Sweden and the Netherlands is almost zero, while it is very low between the pairs Belgium-Denmark, Belgium-United Kingdom, Germany-United Kingdom and Portugal-United Kingdom. Moreover, inflation dynamics of Denmark, Sweden and United Kingdom are not significantly different from that of the Euro-wide area at least at a significance level 0.01.

These results are fully corroborated by the Diebold-Mariano test which, as we stated above, does not call for the independence of the inflation time series. Since the equal forecastability of inflation rates is a necessary condition for similarities of inflation dynamics, if the Diebold-Mariano test is rejected we can reject the hypothesis $d(X_{i,t}, X_{j,t}) = 0$. However, if the Diebold-Mariano test is accepted we cannot say that the AR distance is zero. As shown in Table 4, in all cases in which the similarity test is accepted, the Diebold-Mariano test is consistently not rejected at least at the 5% level. Moreover, also according to the Diebold-Mariano test, the DGP of the aggregate inflation of the Euro-wide fails to reproduce the short-run dynamics of the inflation of all State members. In particular, the hypothesis of equal forecastability of the Euro-wide inflation rate is clearly rejected against France, Luxembourg and Spain.

5.4 Multivariate similarity of short-run inflation dynamics

Typically, AR distance has been used as a criterion for clustering time series by standard algorithms (Piccolo 1990). However, the strategy of clustering time series from the estimated AR distances has two major weaknesses. First, in this way we waste the valuable information arising from the inferential analysis on the squared AR distance. Secondly, the sample distribution of $\hat{d}^2(X_{i,t}, X_{j,t})$ depends on the parameter space, showing an increasing mean and variance as one gets closer to non-invertibility regions. Therefore, making comparisons between estimated distances, corresponding to different points of the parameter space, can be misleading (Sarno and Zazzaro 2002).

In this paper, we use two alternative grouping procedures. First, we group countries on the basis of the density properties of a graph built on the statistical significance of the pairwise similarity tests. Second, after making distances commensurable by discounting their squares with the corresponding standard errors $\mathrm{SE}(\hat{d}^2)$, as an inverse measure of the estimates' precision, we execute a traditional cluster analysis upon these quantities.

A. Graph analysis

Consistent with Definition 2, we can group short-run inflation dynamics of EU countries on the basis of the similarity test on $\hat{d}^2(X_{i,t}, X_{j,t})$. In particular, following graph theory, from a pairwise distance matrix Δ we can build up an adjacency matrix $A = \{a_{ij}\}$ where

$$a_{ij} = \begin{cases} 1 & \text{if } p\text{-value} \geq \text{significance level } \alpha \\ 0 & \text{if } p\text{-value} \leq \text{significance level } \alpha. \end{cases} \tag{8}$$

Therefore, from A we can derive the graph $\mathcal{G}$, where each node represents a country and connecting lines (edges) between nodes designate short-run inflation dynamics that are not significantly dissimilar[9].

When the number of comparisons is large, a graph is a powerful tool to investigate and visualize a number of interesting properties that either unite or separate inflation rate DGPs. In particular, we can elicit: *(i)* the number of countries with which a country i shares the same

[9] This grouping procedure was introduced by Sarno (2005). In a similar vein, Maharaj (1996, 2000) suggests to use a standard hierarchical clustering procedure based on the p-values associated with a test on the equality of the autoregressive parameters. The advantage of following the Maharaj's procedure is that one can get a perfect partition of the EU countries' inflation dynamics. Once again, however, in this way we would lose valuable information, failing to highlight all the inflation dynamics that are reciprocally statistically non-dissimilar, albeit with a slightly lower p-value.

DGP, that is, in graph theory terminology, the degree of a vertex[10], *(ii)* the number, size and composition of country subsets for which every pair of countries shares the same inflation DGP, that is the size and composition of cliques[11], *(iii)* size and composition of the country subset for which each element has a different inflation DGP, that is, the size and composition of the independent set[12].

Obviously, both cliques and the independent set depend on the choice of the significance level for the dissimilarity test. In Figure 2, we report four graphs. Panels A and B include Euro area countries: in the former we set the significance level at 0.05, in the latter at 0.01. Panels C and D include all EU-15 area countries, with a significance level at 0.05 and 0.01, respectively[13].

As displayed in Panel A, during the first seven years of the Euro life, two groups of countries with the same inflation dynamics emerge. On one side, there are France, Greece, Finland and Italy that form the largest clique. On the other side, there are Austria, Finland and Germany whose inflation dynamics are reciprocally similar according to both the similarity and Diebold-Mariano tests. Finland is at the centre of these two groups, showing the highest degree of similarity with other Euro partners. Astonishingly, for eight of the twelve Euro area countries (namely, Austria, Belgium, France, Ireland, Luxembourg the Netherlands, Portugal and Spain) the short-run inflation dynamics is statistically dissimilar. Among these countries, Luxembourg

[10]According to graph theory, the degree of a vertex is the number of edges which it is incident with.

[11]According to graph theory, a clique of a graph $\mathcal{G}$ is a sub-graph of $\mathcal{G}$ whose vertices are pairwise adjacent, that is, incident to a common edge.

[12]According to graph theory, an independent set of graph $\mathcal{G}$ is a subset of vertices such that no two vertices in the subset are adjacent.

[13]Graphs and their metrics are elicited by Graph Interface (GRIN), a software by V. Pechenkin freely available on the Web at *www.geocities.com/pechv_ru/*.

and Spain display an idiosyncratic inflation dynamics without any linkages with other Euro area countries, while inflation dynamics in the Netherlands is statistically similar only to that of Finland at level 1%.

When we reduce the critical region for the dissimilarity test at level 0.01, one more four-country (Belgium-Finland-Greece-Portugal) and four more three-country cliques (Belgium-Finland-Germany; Belgium-Portugal-Germany; Greece-Finland-Ireland; Greece-France-Portugal) emerge (see Panel B). Even so, the number of countries in the independent set remains very large, with only Portugal less with respect to Panel A. Taking in mind the strict relationship between the AR distance and the forecast functions, the strong dissimilarity of inflation dynamics after the introduction of the Euro can give reason for Marcellino, Stock and Watson (2003) findings that inflation forecasts constructed by aggregating the country-specific models are more accurate than forecasts built on aggregate Euro-wide data.

In Panels C and D we also take into account the inflation dynamics of Denmark, Sweden and the United Kingdom. At significance level 0.05, three cliques turned on the UK, and including Belgium, Denmark, Finland and Portugal add to the number of cliques elicited in Panel A. The similarity of short-run inflation dynamics among these countries is still greater if we consider a significance level of 0.01. In this case (Panel D), the largest clique is formed by Belgium, Denmark, Finland, Greece, Portugal and UK. Finally, neither Denmark nor Sweden and UK belongs to the independent sets of graphs C and D.

B. Cluster analysis

As the similarity test does not satisfy transitivity, the approach based on graph theory does not allow us to get a partition of the short run-inflation dynamics. For this purpose, we also run a traditional cluster analysis by the hierarchical complete linkage method. However, in order to make AR distances commensurable, we implement the cluster method on the estimated squared distances discounted by their standard error.

From the dendrogram for the Euro area countries (Figure 3, Panel A) two clusters of countries clearly emerge. The first cluster is formed by Austria, Belgium, Germany, Portugal, and Finland, the second by France, Greece and Italy. The inflation rates of Luxembourg and the Netherlands gravitate towards the first cluster and that of Ireland gravitates towards the second one, while Spain shows a very idiosyncratic inflation dynamics.

Interestingly, our findings are only partly consistent with results on convergence clubs of long-run inflation rates in Euro area countries recently obtained by Busetti, Forni, Harvey and Venditti (2006) in the context of stationarity tests on pairwise inflation differentials. Like Busetti *et al.*, we find strong similarity of short-run dynamics of inflation rates among Austria, Belgium, Finland and Germany. However, while long-run inflation in France and Germany seem to have converged, their short-run behaviour are still quite different. Moreover, although its inflation rate fluctuates around higher levels, Greece displays a short-run inflation dynamics quite similar to that of France and Italy.

The dendrogram reported in Panel B confirms the findings of graph analysis and show that inflation dynamics of Denmark and the UK belong to the cluster of low inflation countries, while inflation dynamics of Sweden is very similar to that prevailing in the Netherlands.

6 Conclusions

The purpose of this paper is to assess the degree of similarity of short-run dynamic properties of inflation rates among EU-15 area countries after the introduction of the Euro in 1999. The question is of primary importance to the design of a common monetary policy in a currency area. For example, similarity of short-run inflation dynamics is required to forecast the impact of European Central Bank (ECB) monetary policy at country level starting from the aggregate Euro-wide price index. Moreover, it reduces potential nationalist tensions within ECB concerning optimal monetary policy.

In this paper, we introduced two definitions of pairwise and multivariate similarity of inflation rates in terms of forecast functions. We then used AR distance (Piccolo 1990) to measure the pairwise similarity of short-run inflation dynamics. Finally, on the basis of inferential analysis conducted on the estimated AR distances, we appraised the multivariate similarity of inflation DGPs among EU countries by means of graph and cluster methods.

On the whole, consistent with studies on inflation differentials and inflation persistence, our findings suggest that after seven years from the launch of the Euro the degree of similarity of short-run inflation dynamics among EU countries is still weak. Within this picture, however, we were able to elicit two groups of countries (namely, France-Greece-Italy and Austria-Belgium-Finland-Germany-Portugal) whose own inflation DGPs showed a statistically similar behaviour.

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Appendix: Tables and Figures

Table 1: Unit roots tests

series	lags	$\mathbf{ADF}_{nc}$	p-value	lags	$\mathbf{ADF}_{c}$	p-value	lags	DFGLS	KPSS
AUT	2	-5.2019	2.80e-7	2	-5.1800	8.62e-6	6	-1.9762	0.0900
BEL	3	-6.4740	0.0000	3	-6.4374	9.90e-9	4	-4.0464	0.0706
FIN	7	-2.0416	0.0395	7	-2.0256	0.2759	7	-1.5774	0.2720
FRA	6	-6.3602	0.0000	6	-6.4466	9.38e-9	7	-0.9347	0.2043
GER	3	-6.0413	3.71e-9	3	-6.0067	1.16e-7	6	-2.4841	0.1079
GRE	4	-7.9211	0.0000	4	-7.9104	0.0000	7	-0.7797	0.1117
EIR	0	-7.0847	0.0000	0	-7.0475	4.95e-8	4	-1.7936	0.1782
ITA	6	-4.2756	2.03e-5	6	-4.3586	0.0003	7	-1.2450	0.2535
LUX	7	-5.6544	2.86e-8	7	-5.7229	5.40e-7	8	-1.1977	0.0820
NED	7	-3.5559	0.0004	7	-3.5076	0.0078	8	-1.6616	0.2639
POR	5	-2.8661	0.0040	5	-2.8615	0.0500	5	-1.5684	0.1712
ESP	6	-4.2971	1.85e-5	6	-4.2666	0.0005	7	-0.4018	0.0984
DEN	0	-10.045	0.0000	0	-9.9909	1.28e-7	1	-6.1056	0.3783
SWE	7	-4.0668	4.91e-5	7	-4.0434	0.0012	8	-1.0108	0.1131
UKD	0	-9.5187	0.0000	0	-9.4675	2.47e-8	1	-6.2968	0.1702
EUR	4	-6.1647	1.91e-9	4	-6.1574	4.96e-8	1	-6.5110	0.1098

Notes: p-values for the ADF tests were computed by using the MacKinnon (1996) algorithm. KPSS test is carried out with a window size of 16. The critical values for the DFGLS test, provided by Elliott, Rothenberg and Stock (1996), are:

1%	5%	10%
-2.5873	-1.9434	-1.6175

The critical values for the KPSS test without trend, provided by Kwiatkovski, Phillips, Schmidt and Shin (1992), are:

10%	5%	2.5%	1%
0.347	0.463	0.574	0.739

Table 2: Estimated ARMA(p,q) models

		AUT					**LUX**		
	coeff	s.e.	t-stat	p-value		coeff	s.e.	t-stat	p-value
AR-1	-0.2278	0.0997	-2.2852	0.0246	AR-7	-0.2701	0.1112	-2.4292	0.0170
MA-3	0.2559	0.1114	2.2971	0.0238	AR-8	-0.4034	0.1124	-3.5891	0.0005
					MA-2	0.3863	0.1115	3.4646	0.0008
	loglik.	par.	BIC	HQC		loglik.	par.	BIC	HQC
	29.8295	2	-0.5264	-0.5582		-34.5932	3	0.8633	0.8156
		BEL					**NED**		
	coeff	s.e.	t-stat	p-value		coeff	s.e.	t-stat	p-value
AR-4	-0.7432	0.1687	-4.4053	0	MA-8	-0.3424	0.1068	-3.2056	0.0018
MA-4	0.6530	0.1783	3.6625	0.0004					
MA-11	0.1621	0.0743	2.1809	0.0317					
	loglik.	par.	BIC	HQC		loglik.	par.	BIC	HQC
	11.3498	3	-0.0938	-0.1416		-25.3418	1	0.5755	0.5596
		FIN					**POR**		
	coeff	s.e.	t-stat	p-value		coeff	s.e.	t-stat	p-value
AR-2	0.8199	0.1520	5.3944	0	MA-6	0.2363	0.1120	2.1100	0.0375
MA-2	-0.8560	0.1631	-5.2480	0					
MA-8	0.2395	0.0677	3.5393	0.0006					
	loglik.	par.	BIC	HQC		loglik.	par.	BIC	HQC
	16.5904	3	-0.2030	-0.2507		12.8384	1	-0.2199	-0.2358
		FRA					**ESP**		
	coeff	s.e.	t-stat	p-value		coeff	s.e.	t-stat	p-value
AR-5	-0.2690	0.1037	-2.5940	0.0110	AR-3	-0.8828	0.0603	-14.6503	0
MA-2	-0.3873	0.1026	-3.7744	0.0003	MA-1	0.2756	0.0881	3.1280	0.0023
MA-6	0.2071	0.1028	2.0145	0.0468	MA-3	0.3470	0.1061	3.2701	0.0015
	loglik.	par.	BIC	HQC		loglik.	par.	BIC	HQC
	35.2634	3	-0.5920	-0.6398		22.4600	3	-0.3253	-0.3730
		GER					**DEN**		
	coeff	s.e.	t-stat	p-value		coeff	s.e.	t-stat	p-value
AR-7	0.1886	0.1007	1.8730	0.0642	MA-10	0.2134	0.0985	2.1676	0.0327
MA-1	-0.2191	0.0984	-2.2253	0.0285					
	loglik.	par.	BIC	HQC		loglik.	par.	BIC	HQC
	31.5649	2	-0.5625	-0.5943		39.6781	1	-0.7791	-0.7950
		GRE					**SWE**		
	coeff	s.e.	t-stat	p-value		coeff	s.e.	t-stat	p-value
AR-7	-0.2274	0.1133	-2.0073	0.0476	MA-8	-0.3162	0.1202	-2.6308	0.0100
MA-2	-0.1998	0.0995	-2.0075	0.0476					
	loglik.	par.	BIC	HQC		loglik.	par.	BIC	HQC
	-22.8061	2	0.5702	0.5384		-18.4635	1	0.4322	0.4163
		EIR					**UKD**		
	coeff	s.e.	t-stat	p-value		coeff	s.e.	t-stat	p-value
MA-1	0.3364	0.0970	3.4664	0.0008	MA-6	0.1142	0.1056	1.0814	0.2823
	loglik.	par.	BIC	HQC		loglik.	par.	BIC	HQC
	5.8773	1	-0.0749	-0.0908		-74.1213	1	1.5917	1.5758
		ITA					**EUR**		
	coeff	s.e.	t-stat	p-value		coeff	s.e.	t-stat	p-value
AR-7	-0.2137	0.1261	-1.6944	0.0935	MA-3	-0.2548	0.1204	-2.1162	0.0370
MA-2	-0.3169	0.1030	-3.0768	0.0027					
	loglik.	par.	BIC	HQC		loglik.	par.	BIC	HQC
	48.3417	2	-0.9120	-0.9439		-0.2548	1	0.0529	0.0369

Table 3: AR distances

	AUT	BEL	FIN	FRA	GER	GRE	EIR	ITA	LUX	NED	POR	ESP	DEN	SWE	UKD
BEL	0.4387	-	-	-	-	-	-	-	-	-	-	-	-	-	-
FIN	**0.5404**	**0.4480**	-	-	-	-	-	-	-	-	-	-	-	-	-
FRA	0.6865	0.5636	**0.6083**	-	-	-	-	-	-	-	-	-	-	-	-
GER	**0.3528**	**0.3726**	**0.4814**	0.6237	-	-	-	-	-	-	-	-	-	-	-
GRE	0.4684	**0.3845**	**0.4953**	**0.4387**	0.5006	-	-	-	-	-	-	-	-	-	-
EIR	0.6231	0.4264	**0.5338**	0.5917	0.5937	**0.4207**	-	-	-	-	-	-	-	-	-
ITA	0.5398	0.4566	**0.5515**	0.4037	0.5495	**0.1374**	0.4646	-	-	-	-	-	-	-	-
LUX	0.7816	0.7490	0.8934	1.0330	0.8558	0.7727	0.8277	0.8626	-	-	-	-	-	-	-
NED	0.5084	0.4863	**0.6790**	0.7294	0.4988	0.4766	0.5101	0.5369	0.5591	-	-	-	-	-	-
POR	0.4639	**0.3339**	**0.5063**	**0.5480**	**0.3834**	**0.3980**	0.4329	0.4863	0.7059	0.4378	-	-	-	-	-
ESP	0.9613	0.6783	0.7278	0.7672	0.7788	0.6542	0.5980	0.6999	0.9469	0.7322	0.6092	-	-	-	-
DEN	0.4144	**0.3214**	**0.3602**	0.5682	0.3669	**0.3786**	0.4187	0.4599	0.6716	0.4248	**0.3269**	0.6502	-	-	-
SWE	0.4866	0.4590	**0.6530**	0.7099	0.4743	0.4534	0.4885	0.5168	0.5600	**0.0327**	**0.4124**	0.7165	**0.3985**	-	-
UKD	**0.3920**	**0.2622**	**0.4359**	**0.5388**	**0.3178**	**0.3324**	0.3757	0.4276	0.6870	0.3821	**0.1300**	0.6047	**0.2469**	**0.3526**	-
EUR	0.5634	**0.3541**	**0.4835**	0.6313	**0.3900**	**0.4029**	0.4650	0.4748	0.7439	0.4497	**0.3985**	0.4686	**0.3423**	**0.4249**	**0.3120**

Notes: Values highlighted in bold correspond to pairs for which the similarity test returns p-values higher than 0.01 (see Table 4).

Table 4: Test statistics for pairwise similarity

	Similarity test			Diebold-Mariano	
countries	d^2	d.o.f.	p-value	z-stat	p-value
AUT-BEL	12.9371	3.3037	0.0064	-0.4108	0.6821
AUT-FIN	4.6706	1.6362	0.0684	0.2475	0.8051
AUT-FRA	18.0600	3.0208	0.0004	1.3254	0.1882
AUT-GER	6.4268	2.4841	0.0625	-0.3274	0.7441
AUT-GRE	13.3530	3.0943	0.0043	-0.1986	0.8430
AUT-EIR	18.3564	1.9321	0.0001	0.0139	0.9889
AUT-ITA	16.4059	3.3228	0.0013	0.1489	0.8820
AUT-LUX	24.5385	3.0096	0.0000	1.5940	0.1142
AUT-NED	14.9775	2.5971	0.0012	0.0203	0.9838
AUT-POR	10.9589	2.0066	0.0042	-0.4089	0.6835
AUT-ESP	47.7410	2.3751	0.0000	1.6291	0.1066
AUT-DEN	11.3299	2.4843	0.0060	-0.4579	0.6481
AUT-SWE	12.2740	2.4575	0.0037	-0.0439	0.9650
AUT-UKD	8.2293	1.9146	0.0148	-0.0162	0.9871
AUT-EUR	12.2167	1.5397	0.0012	-0.6393	0.5241
BEL-FIN	2.9516	1.5355	0.1568	1.0402	0.3008
BEL-FRA	12.1001	2.8868	0.0063	2.5889	0.0111
BEL-GER	12.2049	4.1434	0.0177	0.2472	0.8053
BEL-GRE	9.8165	3.1487	0.0229	0.2297	0.8188
BEL-EIR	14.0541	2.9744	0.0028	0.6241	0.5340
BEL-ITA	11.9043	3.1008	0.0085	0.2767	0.7826
BEL-LUX	21.6577	2.6955	0.0001	1.5437	0.1260
BEL-NED	12.4900	2.0375	0.0020	0.1971	0.8441
BEL-POR	8.1301	2.8714	0.0391	0.0108	0.9914
BEL-ESP	37.3796	3.5293	0.0000	4.4781	0.0000
BEL-DEN	9.3078	3.2826	0.0321	-0.0807	0.9359
BEL-SWE	9.4651	1.8347	0.0072	0.1760	0.8607
BEL-UKD	6.1114	3.2804	0.1276	-0.0337	0.9732
BEL-EUR	7.8125	2.5758	0.0352	0.1029	0.9183
FIN-FRA	5.7078	2.0666	0.0613	1.8867	0.0622
FIN-GER	3.5517	1.5355	0.1127	-0.5419	0.5891
FIN-GRE	3.8695	1.6527	0.1065	-0.3958	0.6932
FIN-EIR	4.5584	1.5064	0.0632	-0.3035	0.7622
FIN-ITA	4.9569	1.8049	0.0699	-0.0102	0.9919
FIN-LUX	14.2959	2.1830	0.0010	1.0903	0.2783
FIN-NED	7.7145	1.5858	0.0130	-0.0800	0.9364
FIN-POR	4.0671	1.5240	0.0841	-0.4772	0.6343
FIN-ESP	8.6870	1.5189	0.0072	3.3099	0.0013
FIN-DEN	1.8478	1.4343	0.2692	-0.7012	0.4849
FIN-SWE	7.1879	1.6662	0.0190	-0.1640	0.8701
FIN-UKD	2.8904	1.4460	0.1490	-0.0060	0.9952
FIN-EUR	3.7074	1.5779	0.1080	-0.7016	0.4846
FRA-GER	14.5022	2.7430	0.0017	-2.9352	0.0042

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Table 4 — continued from previous page

	Similarity test			**Diebold-Mariano**	
countries	d^2	d.o.f.	p-value	z-stat	p-value
FRA-GRE	5.1810	2.1069	0.0825	-1.7629	0.0811
FRA-EIR	12.4431	2.4313	0.0033	-1.4467	0.1513
FRA-ITA	4.2700	2.3221	0.1521	-0.2819	0.7787
FRA-LUX	34.0025	3.2636	0.0000	0.4744	0.6363
FRA-NED	19.8674	2.7137	0.0001	-0.4481	0.6551
FRA-POR	8.3662	1.8538	0.0129	-1.1945	0.2352
FRA-ESP	22.7207	2.7974	0.0000	0.9152	0.3624
FRA-DEN	12.2095	2.5013	0.0040	-1.5970	0.1136
FRA-SWE	18.1550	2.7397	0.0003	-0.5149	0.6078
FRA-UKD	8.6953	1.9163	0.0118	-0.0859	0.9317
FRA-EUR	14.0538	2.5293	0.0017	-2.6030	0.0107
GER-GRE	10.8407	1.8809	0.0038	0.0548	0.9564
GER-EIR	19.2471	1.8118	0.0000	0.4082	0.6840
GER-ITA	12.3719	2.1205	0.0024	0.2552	0.7991
GER-LUX	27.8511	2.7550	0.0000	1.3965	0.1658
GER-NED	15.0825	2.3421	0.0008	0.1269	0.8993
GER-POR	10.6965	2.7375	0.0105	-0.1483	0.8824
GER-ESP	40.2340	2.6575	0.0000	2.7812	0.0065
GER-DEN	11.5162	2.9221	0.0086	-0.2722	0.7861
GER-SWE	11.5911	2.0956	0.0034	0.0624	0.9504
GER-UKD	8.6166	2.9250	0.0328	-0.0022	0.9982
GER-EUR	9.2956	2.4144	0.0149	-0.3193	0.7502
GRE-EIR	8.7906	1.8475	0.0103	0.1837	0.8546
GRE-ITA	0.6046	1.9719	0.7329	0.1963	0.8448
GRE-LUX	20.4307	2.5625	0.0001	1.3251	0.1883
GRE-NED	13.0378	2.4447	0.0025	0.1098	0.9128
GRE-POR	9.7001	2.4484	0.0127	-0.2323	0.8168
GRE-ESP	25.2591	2.4930	0.0000	1.7828	0.0778
GRE-DEN	9.4247	2.3348	0.0129	-0.2153	0.8300
GRE-SWE	10.5177	2.3169	0.0075	0.0617	0.9510
GRE-UKD	7.1423	2.3107	0.0381	-0.0092	0.9926
GRE-EUR	9.0494	2.3899	0.0164	-0.2763	0.7829
EIR-ITA	9.9157	2.0786	0.0077	0.1083	0.9140
EIR-LUX	25.8685	2.3872	0.0000	1.5727	0.1191
EIR-NED	15.1601	1.9373	0.0005	0.0149	0.9882
EIR-POR	12.1659	1.9949	0.0023	-0.4178	0.6770
EIR-ESP	15.9067	1.4883	0.0002	2.5505	0.0123
EIR-DEN	12.7108	1.9102	0.0015	-0.6565	0.5131
EIR-SWE	12.1133	1.8221	0.0019	-0.0563	0.9553
EIR-UKD	10.2033	1.9130	0.0055	-0.0717	0.9430
EIR-EUR	11.3043	1.6973	0.0024	-0.5494	0.5840
ITA-LUX	24.5943	2.6791	0.0000	1.1884	0.2376
ITA-NED	15.5978	2.7836	0.0011	-0.1073	0.9147
ITA-POR	12.8467	2.6012	0.0033	-0.6968	0.4876

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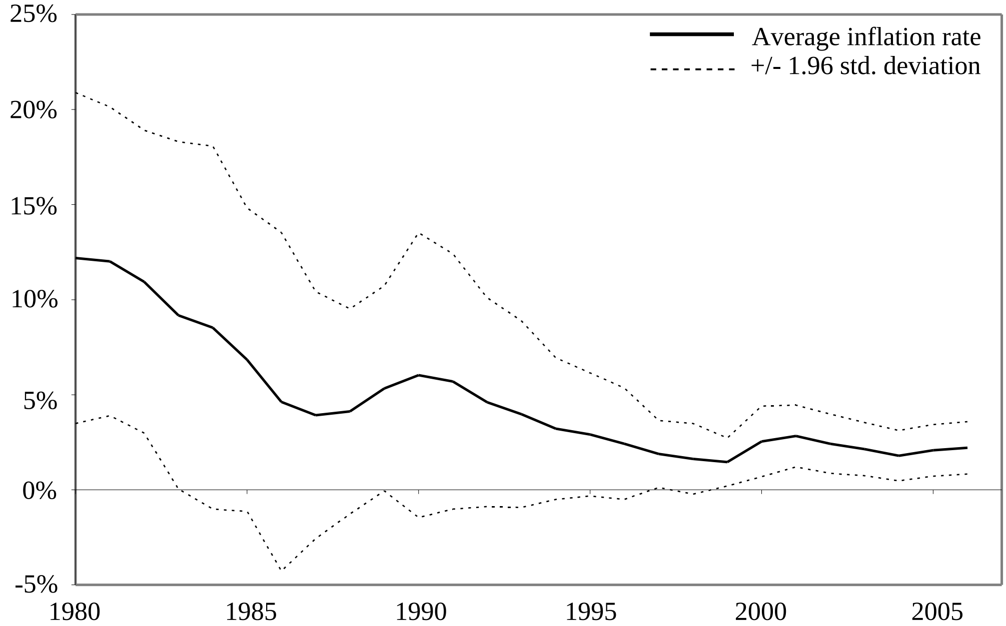
Table 4 — continued from previous page

	Similarity test			Diebold-Mariano	
countries	d^2	d.o.f.	p-value	z-stat	p-value
ITA-ESP	27.8029	2.9450	0.0000	0.5858	0.5594
ITA-DEN	12.3718	2.5586	0.0040	-0.3827	0.7028
ITA-SWE	13.3282	2.7107	0.0029	-0.1435	0.8862
ITA-UKD	10.2082	2.4258	0.0097	-0.2958	0.7680
ITA-EUR	11.9330	2.7335	0.0059	-0.3633	0.7172
LUX-NED	10.7801	2.4411	0.0075	-2.2078	0.0296
LUX-POR	17.8827	2.2793	0.0002	-4.4948	0.0000
LUX-ESP	32.4809	2.3723	0.0000	-0.2312	0.8177
LUX-DEN	16.8241	2.2528	0.0003	-1.3004	0.1966
LUX-SWE	10.1629	2.3918	0.0096	-4.8614	0.0000
LUX-UKD	17.5710	2.2332	0.0002	-0.8706	0.3861
LUX-EUR	21.8050	2.7305	0.0001	-2.0090	0.0473
NED-POR	11.0247	1.9598	0.0038	-0.3241	0.7465
NED-ESP	33.1535	2.3155	0.0000	0.5089	0.6120
NED-DEN	10.8051	1.7432	0.0033	-0.1747	0.8616
NED-SWE	0.0257	1.0000	0.8725	-0.1308	0.8962
NED-UKD	8.7235	1.7498	0.0095	-0.2421	0.8092
NED-EUR	10.8074	1.9995	0.0045	-0.2448	0.8071
POR-ESP	25.1053	2.4021	0.0000	1.3386	0.1838
POR-DEN	7.4428	1.8785	0.0212	-0.0524	0.9583
POR-SWE	8.5810	1.8578	0.0116	0.5140	0.6085
POR-UKD	0.6096	1.0007	0.4352	-0.0180	0.9857
POR-EUR	7.1600	1.4517	0.0147	0.0704	0.9440
ESP-DEN	34.8262	2.6620	0.0000	-2.6435	0.0096
ESP-SWE	27.1581	2.0837	0.0000	-0.8589	0.3925
ESP-UKD	28.4447	2.4994	0.0000	-0.2273	0.8207
ESP-EUR	9.2816	1.5208	0.0053	-2.7490	0.0071
DEN-SWE	8.0203	1.6007	0.0113	0.1467	0.8837
DEN-UKD	5.2071	1.9999	0.0740	-0.0017	0.9986
DEN-EUR	6.8576	1.7270	0.0241	0.1477	0.8829
SWE-UKD	6.2565	1.6072	0.0287	-0.1984	0.8432
SWE-EUR	8.7801	1.9699	0.0120	-0.1533	0.8785
UKD-EUR	4.5075	1.3136	0.0519	0.0133	0.9894

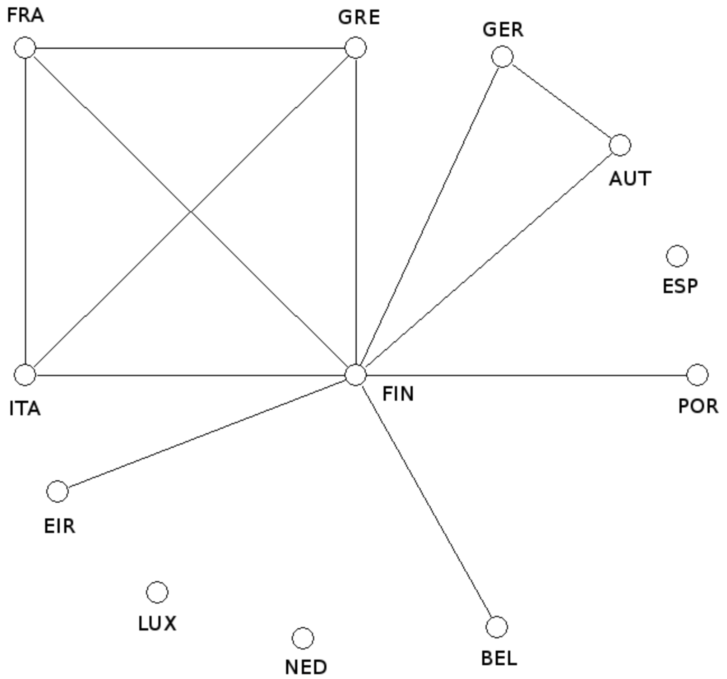
Figure 1: EU-15 area countries yearly inflation rates



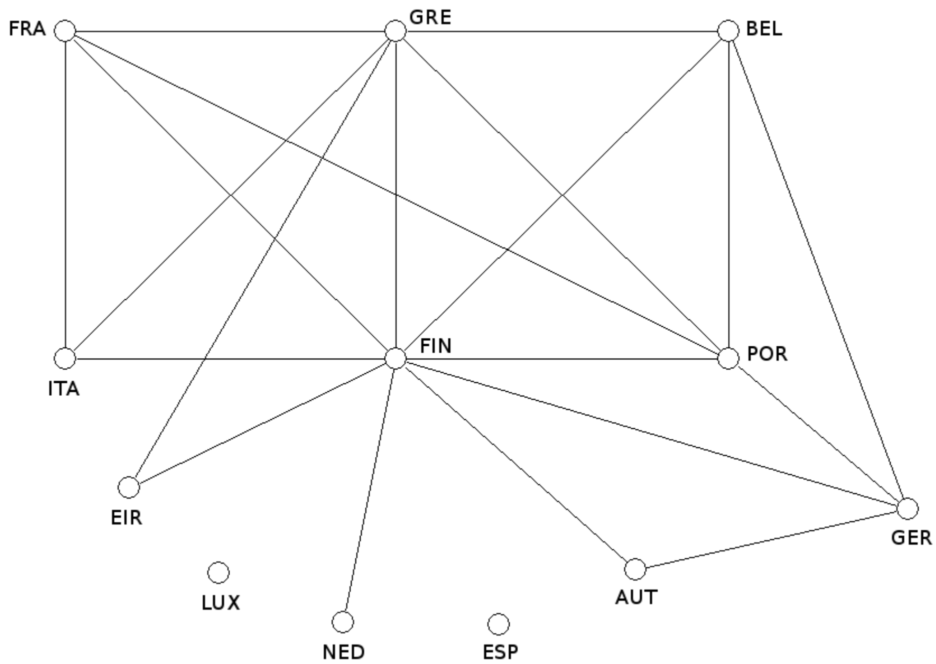
Notes: Inflation rates are calculated by using the yearly seasonally unadjusted all-item consumer price index (Source: OECD).

Figure 2: Multivariate similarity of short-run inflation dynamics: Graph analysis

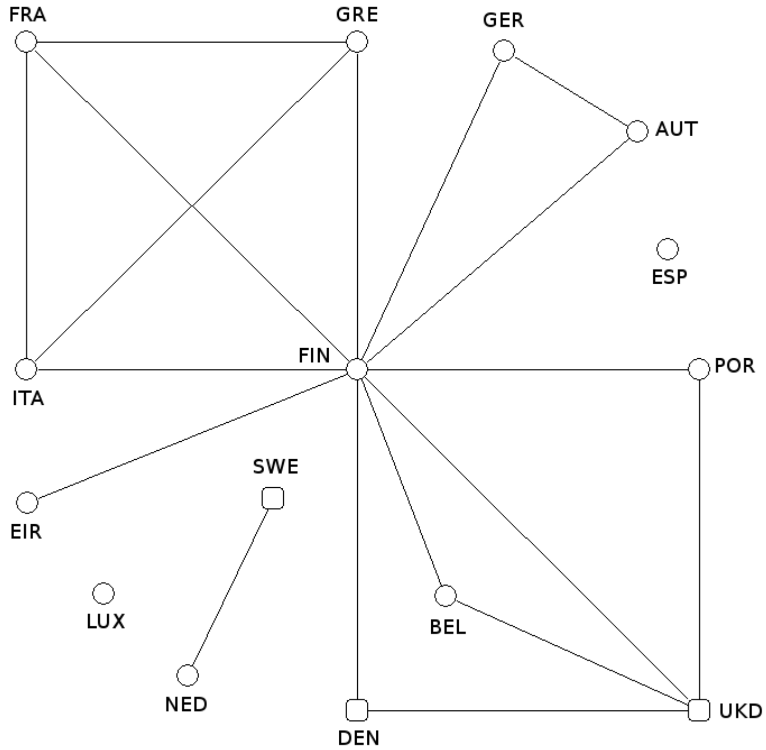
Panel A: Euro area countries (significance level 0.05)



Panel B: Euro area countries (significance level 0.01)



Panel C: EU-15 area countries (significance level 0.05)



Panel D: EU-15 area countries (significance level 0.01)

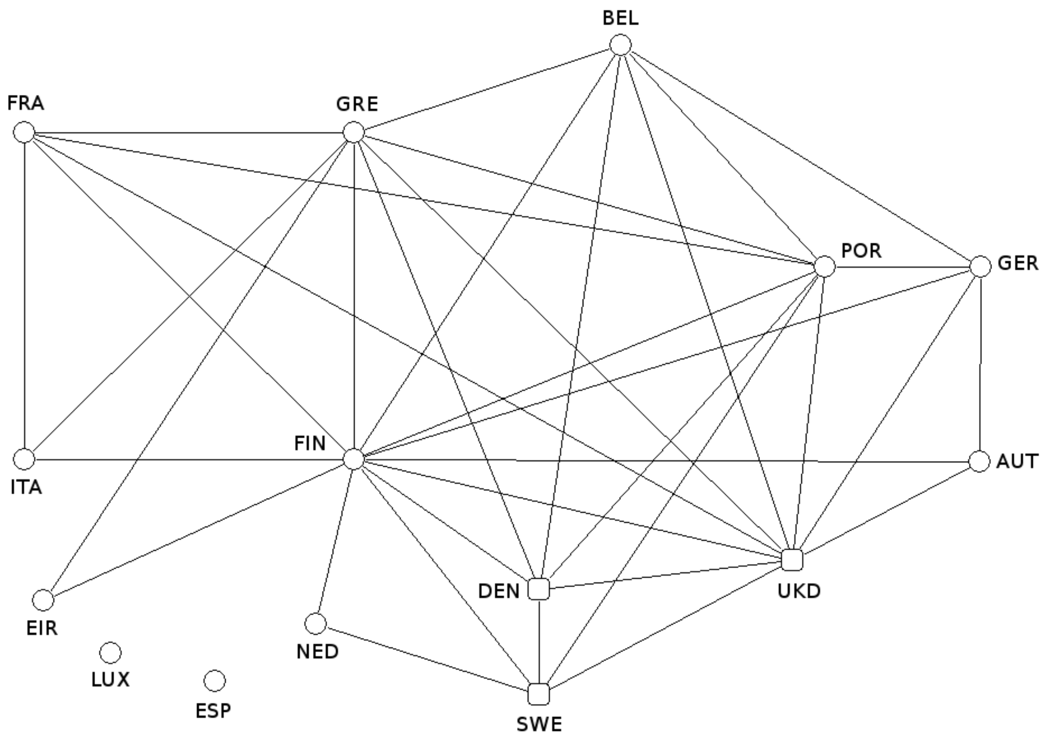


Figure 3: Multivariate similarity of short-run inflation dynamics: Cluster analysis

Panel A: Euro area countries

Panel B: EU-15 area countries