Structural Convergence of Macroeconomic Time Series: Evidence for Inflation Rates in EU Countries[1]

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Abstract

In this paper we introduce a new concept of structural convergence and propose an index of dissimilarity among time series as a measure of global convergence of macroeconomic phenomena. The index is built up from the autoregressive distance estimator. The index has the suitable characteristic of being a continuous measure that allows the evaluation of the overall convergence of several time series by using the information on the mutual convergence of single pairs. In this paper, we apply the index to the series of inflation rates of 13 European Union countries. We find that the convergence of the average level of inflation rates, as required by the Maastricht treaty to enter the monetary union, was only partly accompanied by the convergence in time of inflation dynamics. Moreover, such process of convergence did not concern all countries.

Keywords. Convergence, autoregressive metric, inflation dynamics. J.E.L. C23, E31

1 Introduction

When optimum currency area conditions are not met, for a monetary union to be politically and economically stable countries belonging to it must achieve an adequate degree of economic convergence (Bayoumi and Eichengreen, 1993). It was this widely accepted principle which was behind the convergence criteria fixed by the Maastricht treaty in February 1992. In particular, to place the European Central Bank (ECB) in the position of effectively implementing a non-inflationary monetary policy pursuing the goal of price stability, the treaty provided that in order to enter the monetary union the average inflation rates of member countries converged to a lower and steady common level.

Actually, after Maastricht, European Union (EU) countries experienced an alignment of their inflation rates, in terms of average levels and dispersion (see figure 1), as well as a significant increase in their cross correlations (Angeloni and Dedola, 1999).

[Insert Figure 1]

However, to allow the ECB to run a common monetary policy successfully, avoiding distortionary and asymmetric regional effects, the convergence towards lower levels and common cycles of inflation rates is not enough. In fact, a positive correlation among inflation rates might occur along with appreciable differences in response to policy interventions of monetary authority. This means that the process of convergence of EU-wide inflation rates cannot be confined to their average levels or their common trend, but should also concern their dynamic structure.

As for the latter point, the available evidence is less promising. From simulation exercises, conducted both from VAR (Mojon and Peersman, 2001) and macro-structural models (van Els, Locarno, Morgan and Villetelle, 2001), there clearly emerges a similarity in the sign of the price responses to monetary policy impulses across EU countries, but a substantial difference in their size. These types of exercises, analyzing the response paths of the equilibrium inflation rates, are appropriate for investigating the differences in monetary transmission mechanisms. However, since they do not take into account the data generating processes of the inflation rates, they do not allow us to draw a definite measure of the convergence in the inflation dynamics across EU countries. In other words, the present empirical research does not allow us to give an answer to the central question that we pose in this paper: During the 1990s, did the inflation rate dynamics of EU countries become more similar? Specifically, after Maastricht, did the convergence of the average inflation rates occur along with a greater similarity of their data generating processes?

To address this issue, in section 2 we will introduce a notion of structural convergence and, in section 3, propose, as its measure, an index of dissimilarity among time series which is built up on the autoregressive metric between ARIMA models. Then, in section 4, we will employ this index to evaluate whether or not the EU countries experienced a common convergence process in their inflation rate dynamics. Section 5 concludes.

2 A notion of structural convergence

In macroeconomic analysis, and in empirical growth literature in particular, different notions and measures of convergence have been proposed, ranging from cross section (as in the β-convergence and σ-convergence approach) to time series tests (as in the cointegration and distribution-dynamics approach)[1]. All of them, however, refer to a more general concept of convergence which considers that two, or a group of economic time series converge if the expected level of their random components becomes identical or differs in some constant value, as the time horizon goes to infinity. Therefore, the whole dynamic structure of the series is not taken into account, except for analyzing the existence of common trends, as occurs, for example, in the cointegration approach (Durlauf, 1989; Quah, 1992; Bernard and Durlauf, 1995).

Such a concept of convergence is pertinent when a long run equilibrium is theoretically identifiable and the empirical issue is to ascertain whether or not the economic series really tend to such equilibrium. It is much less appropriate, instead, when a well-defined concept of equilibrium is not available, and the empirical issue is to verify whether or not two series respond similarly if hit by the same shock.

In order to deal with this issue, in this paper, we introduce a new definition of convergence between time series, as the tendency of their forecast functions to coincide.

Definition 1 (Structural convergence) *Two stochastic processes $X_{i,t}$ and $X_{k,t}$ structurally converge if their forecast functions become more similar over time. This implies that, given identical initial values, the processes $X_{i,t}$ and $X_{k,t}$ structurally converge if, considering two spells of time T and T', where T precedes T', the following inequality holds: $\left|\mathcal{F}_{T,j}^{X_i}\left(I_T^{X_i}\right) - \mathcal{F}_{T,j}^{X_k}\left(I_T^{X_k}\right)\right| > \left|\mathcal{F}_{T',j}^{X_i}\left(I_{T'}^{X_i}\right) - \mathcal{F}_{T',j}^{X_k}\left(I_{T'}^{X_k}\right)\right|$, $\forall j = 1, 2, ...,$ where $\mathcal{F}_{T,k}^{X_z}\left(I_T^{X_z}\right)$ denotes the forecast for the process $X_{z,t}$ at time $T+j$ derived from the information set $I_T^{X_z} = \{X_{z,T-n},\, n = 0,\, 1,\, ...T\}$, with $z = i,\, k$.*

[1] For useful surveys, see Bernard and Durlauf (1996); Durlauf and Quah (1999); Hall, Robertson and Wickens (1992, 1997).

In some ways, Definition 1 may resemble Bernard and Durlauf's (1996) definition of convergence as equality of long-term forecasts at a fixed time. If the forecast functions of two series coincide, as well as their initial values, so certainly do their long-term forecasts. The critical difference between the two convergence notions is that the former takes into account the dynamic properties of the two series considered, while the latter only looks at their long-term expected values, examining the co-movements of the series over time.

Usually, however, the empirical issue that one is called to tackle is to evaluate whether, and to what extent, a group of time series shows a tendency towards converge. Also from this point of view the existing convergence measures are not entirely satisfactory.

First of all, since economic convergence is treated as a limit concept, the convergence measures employed in the economic literature only allow an all-or-nothing evaluation. More specifically, convergence tests are able to say whether or not series converge, and, at most, the speed of this process, but they do not offer any measure of how much series diverge, at a certain moment.

Secondly, convergence tests either say too little or are too stringent. Cross section studies, for example, give an intuitive measure of the process of global convergence. However, they do not permit us to identify exactly which series actually converge, nor to attribute different weights to series when their importance differs. By contrast, time series tests, as in the cointegration approach, are very demanding, since they require that all the series considered two by two converge. An adequate notion of global convergence, instead, ought to allow us to find out which series are converging and ascertain to what extent they do it on average. As for this issue, we suggest extending structural convergence to the case of several time series in the following way.

Definition 2 (Global structural convergence) *Given a finite number of stochastic processes* $\{X_{i,t}\}_{i=1,\dots,N}$*, they globally converge over time if, on average, each forecast function corresponding to a single process becomes more similar to those of the others. This implies that, given identical initial values, N stochastic processes structurally converge if* $\left|\sum_{i=1}^{N}\sum_{k=i+1}^{N}\left[\mathcal{F}_{T,j}^{X_i}\left(I_T^{X_i}\right)-\mathcal{F}_{T,j}^{X_k}\left(I_T^{X_k}\right)\right]w_{i,k}\right| > \left|\sum_{i=1}^{N}\sum_{k=i+1}^{N}\left[\mathcal{F}_{T',j}^{X_i}\left(I_{T'}^{X_i}\right)-\mathcal{F}_{T',j}^{X_k}\left(I_{T'}^{X_k}\right)\right]w_{i,k}\right|$, $\forall\, j = 1,\, 2,\, ...,$ *where* $w_{i,k}$, $\sum_{i=1}^{N}\sum_{k=i+1}^{N} w_{i,k} = 1$*, are weights reflecting the importance assigned to each pair of processes considered.*

3 An index of dissimilarity among time series

To operationalize our notion of convergence, it is necessary to refer to a forecasting method and to a measure of similarity between statistical time series models. In the following we restrict the analysis to the class of invertible ARIMA models. First of all, as the recent research has confirmed, traditional univariate linear models show a good short-run forecasting performance for macroeconomic series, hardly improvable by more complex multivariate or non-linear models (Meese and Geweke, 1984; Canova, 2002; Marcellino, 2002; Marcellino, Stock and Watson, 2003). Moreover, statistical literature provides several parametric measures of similarity between univariate linear models[2].

[2] Among these, the measures most applied are the Mahalanobis distance (Peña, 1990), the Kullback-Lieber divergence (Shumway and Unge, 1974; Alagon, 1989), the Bhattacharya distance (Chaudury *et al.*, 1991) the *cepstral* coefficients distance (Thomson and De Souza, 1985) and the autoregressive distance (Piccolo, 1990; Corduas, 1990; Maharaj, 1996; Sarno, 2002)

For our purpose, a useful measure for evidencing convergence over time of a group of data generating processes can be constructed from the autoregressive (AR) metric proposed by Piccolo (1989; 1990).

The AR metric is a measure of structural discrepancy between two invertible ARIMA processes, $X_{i,t}$ and $X_{k,t}$. It represents the Euclidean distance between the sequences of the autoregressive coefficients of their pure AR(∞) representations which, according to classical notation, are given by: $\pi_{X_i}(B) X_{i,t} = a_{X_i,t}$ and $\pi_{X_k}(B) X_{k,t} = a_{X_k,t}$, where the polynomial $\pi_{X_z}(B) = \phi_{X_z}(B) \Phi_{X_z}(B^s) \nabla^d \nabla_s^D \theta_{X_z}^{-1}(B) \Theta_{X_z}^{-1}(B^s) = 1 - \pi_{X_z,1} B - \pi_{X_z,2} B^2 -$ ...., and $a_{X_z,t}$ is a Gaussian white noise process, for $z = i, k$. In symbols:

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

As is well known, the sequence of $\pi-$weights fully specifies the dynamic structure of an invertible model and thereby its corresponding forecast function. Therefore, the AR distance allows us to compare the dynamic structure of two ARIMA processes and evaluate the similarity of their forecast functions. As a matter of fact, for given initial values, the AR distance between two processes decreases when their forecast functions become more similar, and is null if and only if their forecast functions coincide (Piccolo, 1990).

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

$$\widehat{d}_T(X_{i,t}, X_{k,t}) = \sqrt{\sum_{j=1}^{L} (\widehat{\pi}_{X_i,j} - \widehat{\pi}_{X_k,j})^2}, \tag{1}$$

where T denotes the observational sample period.

Thus, we can unambiguously say that the forecast functions of two stochastic processes $X_{i,t}$ and $X_{k,t}$ structurally converge if their estimated AR distances calculated over successive spells of time decrease. In other words, $X_{i,t}$ and $X_{k,t}$ converge over time if $\widehat{d}_T\left(X_{i,t}, X_{k,t}\right) > \widehat{d}_{T'}\left(X_{i,t}, X_{k,t}\right)$, with T preceding T'.

The asymptotic properties of the squared AR distance estimator $\widehat{d}^2$ are known under ML estimates (Piccolo, 1989; Corduas, 1996) and LS estimates (Sarno, 2001). In particular, it was shown that the sample distribution of the $\widehat{d}^2$ is a linear combination of independent Chi-squared random variables.

Since the AR distance satisfies the properties of a metric, our notion of convergence between two stochastic processes can be immediately extended to the case of a group of time series by means of a summary statistics based on distances calculated on single pairs. Hence, we provide a measure of global convergence referred to a group of stochastic processes, where traditional convergence tests employed in economics only allow an all-or-nothing evaluation.

Unfortunately, the sample distribution of $\widehat{d}^2$ depends on the parameter space, showing an increasing mean and variance as one gets closer to non-invertibility regions. A consequence of such dependence is that making comparisons between estimated distances, corresponding to different points of the parameter space, can be misleading. In order to make distances commensurable, we suggest normalizing the estimated squared distances by their standard errors (SE), this representing an inverse measure of the estimates' precision. Such as they are, these quantities belong to the class of the "Pearson distances" (Mardia, Kent and Bibby, 1979, p. 377). Of course, the normalized squared AR distances do not satisfy all the properties of a metric anymore. For instance, the triangular property does not hold. However, since we are interested in evaluating convergence and divergence of macroeconomic phenomena, $\frac{\widehat{d}^2}{SE}$ remains a suitable measure of dissimilarity for our purposes.

The standard error for the LS squared distance estimator was derived by Sarno (2001) and it is equal to:

$$SE\left(\widehat{d}^2\right) = \sqrt{\widehat{VAR}\left(\widehat{d}^2\right) + \widehat{bias}\left(\widehat{d}^2\right)} = \sqrt{8tr\left(\widehat{V}^2\right) + 2tr\left(\widehat{V}\right)} \tag{2}$$

The matrix V refers to the covariance matrix of the estimator $\widehat{\pi}$ in the autoregression fitting. An estimate is $\widehat{V} = 0.5\left(\widehat{V}_{X_i} + \widehat{V}_{X_k}\right)$, where $\widehat{V}_{X_z} = \left\{n^{-1}v_{X_z,h,j}\right\}$, $v_{X_z,h,j} = \sum_{p=0}^{h-1}\widehat{\pi}_{X,p}\widehat{\pi}_{X,p-h+j}$, with $1 \leq h \leq j \leq L$, for $z = i, k$. The bias in expression (2) refers to the hypothesis $H_0 : d\left(X_{i,t},\, X_{k,t}\right) = 0$.

Therefore, we propose the following dissimilarity index:

$$\widehat{\delta}_T = \sum_{i=1}^{N}\sum_{k=i+1}^{N}\frac{\widehat{d}_T^2\left(X_{i,t}, X_{k,t}\right)}{SE\left(\widehat{d}_T^2\left(X_{i,t}, X_{k,t}\right)\right)}w_{i,k} \tag{3}$$

where N is the number of series considered, and the weights $w_{i,k}$, such that $\sum_{i=1}^{N}\sum_{k=i+1}^{N} w_{i,k} = 1$, reflect the importance assigned to each pair of countries in the set investigated. Clearly, if $w_{i,k} = \frac{2}{N(N-1)}$, $\widehat{\delta}_T$ returns an arithmetic mean of the normalized squared distances.

Given the definition of global structural convergence in section 2 and the dissimilarity index expressed in (3), we can unambiguously state that a group of macroeconomic processes globally converges (diverges) if and only if $\delta_T > (<)\ \delta_{T'}$, as long as the time interval T precedes T'. Then, in order to verify the statistical significance of a reduction (increase) in $\widehat{\delta}$, since the t test cannot be applied because samples drawn at time T and T' are not independent, we suggest employing the Wilcoxon signed-ranks test (Gibbons and Chakraborti, 1992) comparing matched pairs of normalized squared AR distances “before” and ”afterwards”. Hence, we can verify the null hypothesis $H_0 : \delta_T = \delta_{T'}$ against the alternative hypothesis $H_1 : \delta_T > (<)\ \delta_{T'}$ of global structural convergence (divergence).

4 The structural convergence of inflation rates in EU countries

In this section we present evidence on the process of convergence of the inflation dynamics in EU countries after the Maastricht treaty. The dataset consists of the monthly seasonally unadjusted all-item consumer price index (CPI) from 1984:01 to 2001:12 for twelve EU countries (Austria, Belgium, Denmark, Finland, France, Germany, Greece, Italy, Netherlands, Portugal, Spain, Sweden and United Kingdom). The data originate from the OECD. Inflation rates are computed from the $CPIs$ by taking $y_t = 100\log(CPI_t - CPI_{t-1})$. Samples were split into two periods: ante-Maastricht (up to 1993:04) and post-Maastricht (from 1993:05) as suggested in Morana (2000), who detected a different inflation rate regime endogenously through a Markov switching mechanism. Finally, averages of the purchasing power parity GDP in US dollars were employed to elicit the weights $w_{i,k}$ in $\widehat{\delta}_T$ [Source: OECD].

The identification and estimation of ARIMA models for the inflation rates were carried out following the standard Box-Jenkins procedure. In Table 1, we report the estimation results. In a few cases, data showed a strong skewness which forced us to work on subsamples in order to avoid rejection of Normality[3]. All series showed a clear seasonal pattern and, therefore, needed to be differentiated (except for Italian post-Maastricht data).

[Insert Table 1 here]

For each estimated model we derived its AR(L) representation, for $L =$

[3] To be precise, six models were estimated over a slightly shorter sample. With regard to the ante-Maastricht period, these are Austria (from 84:02), Denmark (from 86:05), Germany (up to 91:04), Portugal (from 86:02) and Spain (up to 89:12). With regard to the post-Maastricht period, the Netherlands (up to 2000:12) only.

200. As table 1 clearly shows, models belong to different points of the parameter space, thereby attesting the practical relevance of the normalization that we have proposed in the previous section. Therefore, we calculated the normalized squared AR distances for each pair of countries and subsequently calculated the dissimilarity index $\widehat{\delta}$.

In all, we find evidence of a tendency to converge in the dynamics of the inflation rates across EU countries. The index of dissimilarity, when weighted with the GDP share of each pair of countries considered, decreased from 0.1029 in the ante-Maastricht period to 0.0430 afterwards (see Table 2). Also the dissimilarity index calculated as arithmetic mean dropped, but only from 0.0814 to 0.0485. This indicates that, after Maastricht, the convergence process mostly concerned larger countries. However, the reduction in $\widehat{\delta}$ does not appear statistically significant, as suggested by the one-sided Wilcoxon signed-rank test reported in table 2.

[Insert Table 2 here]

Table 3, which reports the normalized squared AR distances for each pair of countries on both periods, allows us to identify countries that increased or decreased their similarity with regard to the inflation dynamics of the other European partners. As one can see, the countries that experienced the strongest convergence were Finland, France, Germany and, above all, the UK that drastically reduced the distance of its inflation dynamics with those of all the other EU partners. An appreciable alignment also occurred for Denmark and the Netherlands, becoming practically identical, and for Greece and Spain. By contrast, Austrian, Belgian, Italian and Portuguese inflation rates showed a clear tendency to diverge both reciprocally and from those of their European partners, France and Germany especially. If we exclude Austria, this result is not entirely surprising: Belgium, Italy and Portugal are among those countries which had the highest public debt and serious structural problems. In some ways, our evidence would confirm the widely held opinion that while the convergence effort of the latter countries was considerable, Maastricht criteria were met without reforming the more structural elements of their economies.

[Insert Table 3 here]

Therefore, the overall convergence process in price dynamics did not affect all EU countries, and was mainly determined by the alignment of the UK towards the rest of Europe. A confirmation of this result may be obtained by computing the dissimilarity index $\widehat{\delta}$ within the Euro zone, i.e. excluding Denmark, Sweden and the UK (see table 4). In this case, the dynamics of the inflation rates show a tendency to structurally diverge from the ante-Maastricht to the post-Maastricht period. Specifically, the weighted dissimilarity index increased from 0.0409 (ante-Maastricht) to 0.0510 (post-Maastricht), while the dissimilarity index as arithmetic mean rose from 0.0399 to 0.0554. Here, the increase in $\widehat{\delta}$ is confirmed by the Wilcoxon test at a significance level of 0.05.

[Insert Table 4 here]

These results on the convergence of the dynamics of inflation rates are in line with recent evidence on price dispersion in the EU countries, which show that the narrowing effect of the EMU was small and restricted to some countries (Sosvilla-Rivero and Gil-Pareja, 2002; Lutz, 2003). To the extent that dissimilarities in the temporal dynamics of inflation rates reflect differences in structural and institutional features of the economies considered, our findings suggest that, with the exclusion of some countries, the process of convergence of EU countries has essentially concerned nominal variables. Hitherto, the integrating impact of the common currency project on the structural characteristics of the EMU countries has been modest and this could represents a great obstacle for the ECB in running a non-distortionary common monetary policy. Of course, one cannot exclude that it is precisely the common conduct of the monetary policy which will help integrating structural effects of the monetary union to emerge. Whatever the case, the structural convergence of EU economies still remains an objective.

5 Conclusions

In this paper we introduced a new concept of structural convergence and we proposed an index of dissimilarity among time series as a measure of global convergence of macroeconomic phenomena. The index was built up from the autoregressive distance measure first introduced by Piccolo (1989; 1990), which compares stochastic dynamic structure of two data generating processes through the sequence of the coefficients of their AR(∞) representations. The index is a continuous measure of dissimilarity and has the convenient characteristic of allowing evaluation of the overall convergence of several time series by using the information on the mutual convergence of single pairs.

We applied the index to the series of inflation rates of 13 EU countries. We found that the convergence of the average level of inflation rates, as required by the Maastricht treaty to qualify for entry to the monetary union, was not accompanied by appreciable convergence in time of inflation dynamics for all countries. Some of them, like Belgium, Italy and Portugal, although experiencing a reduction in their inflation levels, had a clear misalignment in the inflation dynamics from that of the remaining European partners.

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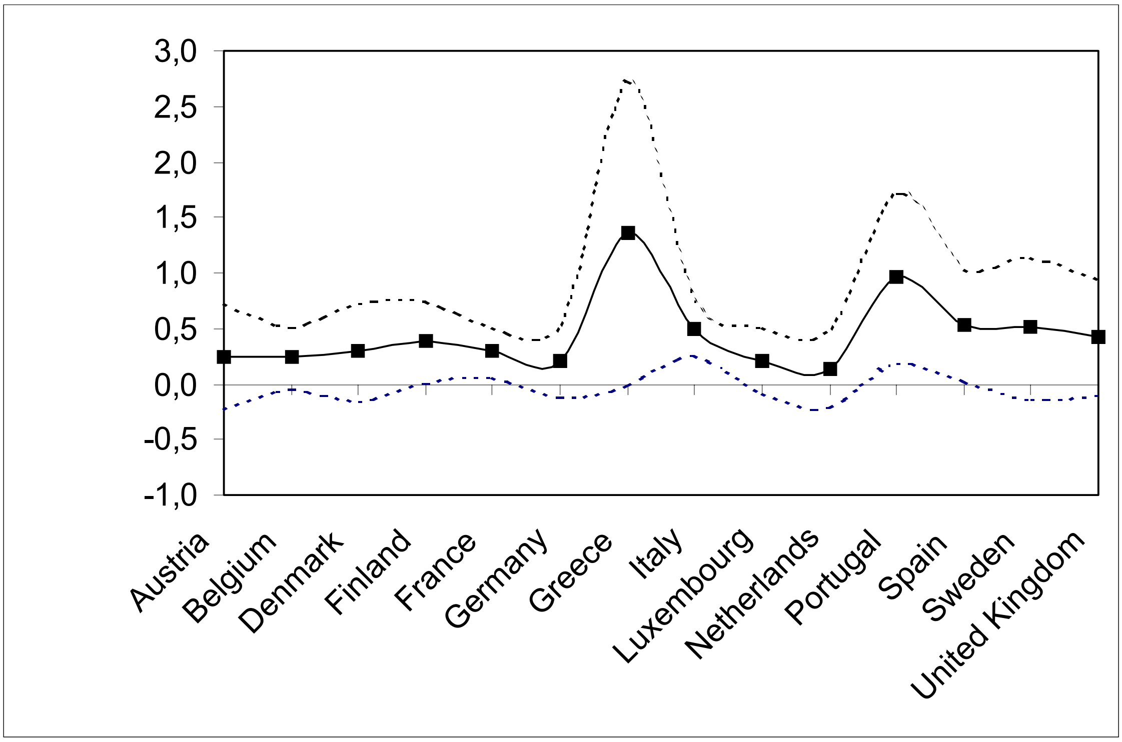
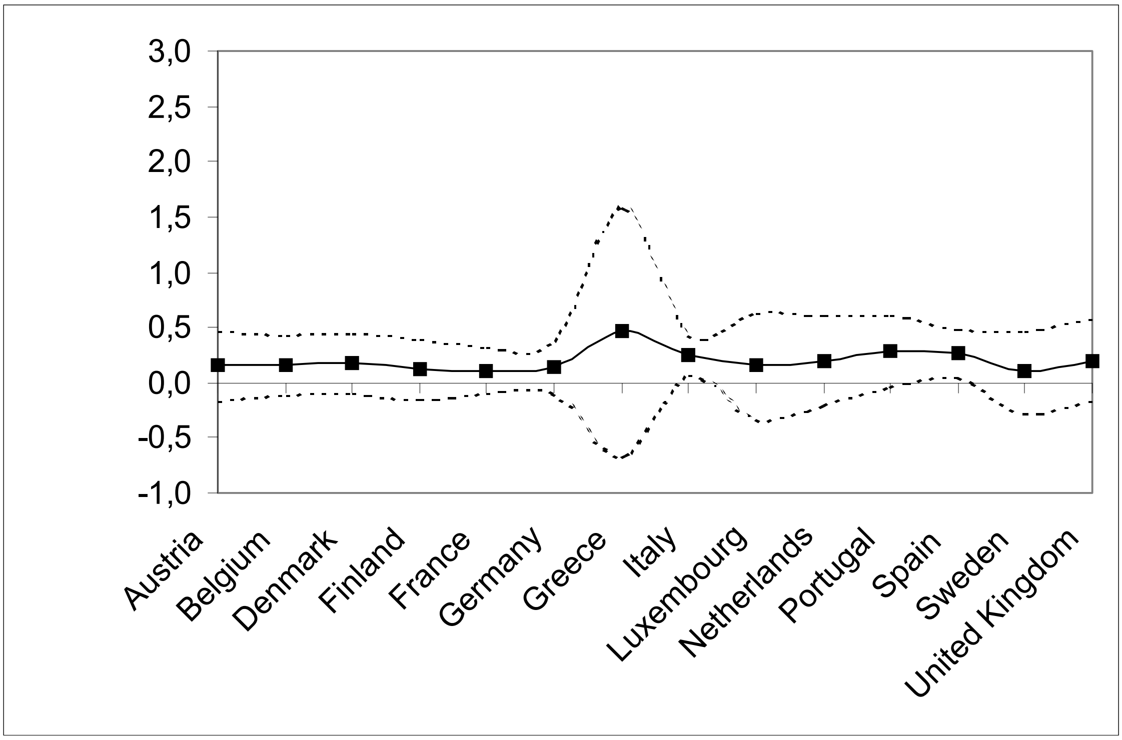


(a) Ante Maastricht (1984:01 – 1993:04)



(b) Post Maastricht (1993:05 – 2001:12)

Figure 1. Average inflation rates (CPI, all-items seasonally unadjusted)

Source: OECD. Dotted lines are calculated as average inflation rate ± standard deviation.

Table 1. The estimated ARIMA models

	Ante-Maastricht (*1984:01 - 1993:04*)	Post-Maastricht (*1993:05 - 2001:12*)
Austria	∇^{12} AR(4) = 0.191 SMA(12) = -0.886	∇^{12} SMA(12) = -0.340
Belgium	∇^{12} AR(1) = 0.235 SMA(12) = -0.886 AR(4) = 0.307 AR(9) = 0.161	∇^{12} MA(1) = 0.206 MA(4) = -0.409 SMA(12) = -0.646
Denmark	∇^{12} SMA(12) = -0.866	∇^{12} SMA(12) = -0.881
Finland	∇^{12} AR(4) = 0.185 SMA(12) = -0.886 AR(6) = 0.298	∇^{12} SMA(12) = -0.886
France	∇^{12} AR(1) = 0.347 SMA(12) = -0.886 AR(4) = 0.251	∇^{12} SMA(12) = -0.864
Germany	∇^{12} AR(1) = 0.303 SMA(12) = -0.848	∇^{12} AR(1) = 0.134 SMA(12) = -0.886
Greece	∇^{12} AR(2) = 0.207 SMA(12) = -0.867	∇^{12} AR(1) = 0.382 SMA(12) = -0.886
Italy	∇^{12} AR(1) = 0.402 SMA(12) = -0.886 AR(2) = 0.277	AR(1) = 0.293 SMA(12) = 0.886 AR(3) = 0.539
Netherlands	∇^{12} AR(6) = 0.381 SMA(12) = -0.886	∇^{12} SMA(12) = -0.875
Portugal	∇^{12} SMA(12) = -0.866	∇^{12} AR(1) = 0.248 SMA(12) = -0.886 AR(3) = -0.215 SAR(12) = 0.356
Spain	∇^{12} AR(9) = 0.186 SMA(12) = -0.886	∇^{12} AR(1) = 0.469 SMA(12) = -0.885
Sweden	∇^{12} SMA(12) = -0.866	∇^{12} AR(3) = 0.312 SMA(12) = -0.886
United Kingdom	∇^{12} AR(1) = 0.473 SAR(12) = -0.511	∇^{12} AR(1) = 0.224 SMA(12) = -0.868

Table 2. Global dissimilarity of the inflation dynamics across EU countries

Weights	*Dissimilarity index*		*One-sided Wilcoxon signed-rank test*
	$\hat{\delta}_{AM}$	$\hat{\delta}_{PM}$	$H_0 = \delta_{AM} = \delta_{PM}$
$w_{i,k} = \dfrac{GDP_i + GDP_k}{\sum_{i=1}^{N} GDP_i}$	0.1029	0.0430	$T^+ = 0.5902$ p-value = 0.2775
$w_{i,k} = \dfrac{2}{N(N-1)}$	0.0814	0.0485	$T^+ = -0.4508$ p-value = 0.3261

Table 4. Global dissimilarity of the inflation dynamics in the Euro zone

Weights	*Dissimilarity index*		*One-sided Wilcoxon signed-rank test*
	$\hat{\delta}_{AM}$	$\hat{\delta}_{PM}$	$H_0 = \delta_{AM} = \delta_{PM}$
$w_{i,k} = \dfrac{GDP_i + GDP_k}{\sum_{i=1}^{N} GDP_i}$	0.0409	0.0510	T^+ = -1.7778 p-value = 0.0377
$w_{i,k} = \dfrac{2}{N(N-1)}$	0.0399	0.0554	T^+ = -2.6808 p-value = 0.0037

Table 3. The normalized squared AR distances

	Belgium	Denmark	Finland	France	Germany	Greece	Italy	Netherl.	Portugal	Spain	Sweden	United Kingdom
Austria	**0.0241** **0.0884**	**0.0086** **0.0681**	**0.0231** **0.0695**	**0.0314** **0.0639**	**0.0305** **0.0729**	**0.0217** **0.0955**	**0.0679** **0.1203**	**0.0469** **0.0631**	**0.0088** **0.0376**	**0.0193** **0.1076**	**0.0100** **0.0873**	0.3669 0.0744
Belgium		**0.0396** **0.0811**	**0.0455** **0.0819**	**0.0101** **0.0787**	**0.0281** **0.0730**	**0.0557** **0.0742**	**0.0528** **0.1052**	0.0783 0.0761	**0.0403** **0.0827**	**0.0384** **0.0795**	**0.0456** **0.0979**	0.3030 0.0683
Denmark			0.0283 0.0000	0.0412 0.001	0.0199 0.0047	**0.0104** **0.0356**	**0.0523** **0.0705**	0.0331 0.0000	**0.0000** **0.0573**	**0.0083** **0.0519**	**0.0000** **0.0244**	0.3462 0.0129
Finland				0.0523 0.0003	0.0489 0.0047	0.0434 0.0356	0.0866 0.0706	0.0103 0.0001	**0.0289** **0.0580**	**0.0414** **0.0520**	0.0328 0.0244	0.3706 0.0130
France					0.0148 0.0049	0.0575 0.0357	**0.0334** **0.0703**	0.0799 0.0001	**0.0420** **0.0552**	0.0556 0.0519	0.0475 0.0245	0.2704 0.0128
Germany						0.0315 0.0149	**0.0190** **0.0580**	0.0534 0.0044	**0.0202** **0.0462**	0.0303 0.0263	**0.0226** **0.0286**	0.2579 0.0022
Greece							**0.0412** **0.0541**	0.0487 0.0336	**0.0106** **0.0443**	0.0213 0.0017	**0.0121** **0.0569**	0.3855 0.0060
Italy								0.0911 0.0666	**0.0533** **0.1254**	0.0679 0.0583	0.0601 0.0299	0.2830 0.0533
Netherlands									**0.0338** **0.0535**	**0.0467** **0.0490**	0.0383 0.0230	0.3714 0.0121
Portugal										**0.0084** **0.0492**	**0.0000** **0.1072**	0.3488 0.0403
Spain											**0.0096** **0.0719**	0.3675 0.0139
Sweden												0.3662 0.0362

Notes: In each cell, the first (second) row reports the distance referred to the ante- Maastricht (post-Maastricht) period. Increases in the normalized squared autoregressive distances are highlighted in bold.