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Does European Monetary Union make
inflation dynamics more uniform?

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Abstract

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Key words: inflation dynamics, EMU, Markov process, cluster analysis

J.E.L. Classification: C22; C38; E31.

1 Introduction

Several studies show that the adoption of a common currency has led to an increase or to negligible changes in the inflation differentials among the countries belonging to the European Monetary Union (EMU) (see Angeloni and Ehrmann, 2007; Buseti *et al.*, 2007; Beck *et al.*, 2006; Duarte, 2003; Honohan and Lane, 2003 and the papers therein quoted; see also Holmes, 2008 for a partially alternative view). In fact, the convergence of the inflation rates is a prerequisite for the admission to the EMU, but the differentials may persist or increase once the currency union is established, due to a number of structural

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reasons (e.g., different exposure of the countries to the exchange rates movements outside the Eurozone or heterogeneities in the productivity growth of tradables and nontradables sectors, i.e. the Balassa-Samuelson effect).

Despite these results, that seem to conjecture a clusterization of the EMU countries around different steady state inflation rates, the analysis of the short-term inflation dynamics also matters in determining to what extent the EMU is contributing to maintain price stability in Europe, which is one of the major aims of the currency union (Eickmeier and Breitung, 2006; Palomba *et al.*, 2009). As Palomba *et al.* (2009) suggest, in fact, dissimilarities in the short-term inflation may have an impact on the short-term real interest rates across the Euro countries, with consequences on the optimal monetary policy the European Central Bank should pursue (Benigno and López-Salido, 2006; Benigno, 2004), and may condition the successful enlargement of the EMU to the Eastern Europe countries.

Taking as a matter of fact that the inflationary processes may be converging to different steady state values, this paper offers a nonparametric analysis of whether the entrance in the Eurozone has made the price dynamics more homogeneous in terms of short-term trend and volatility. Assuming that inflation dynamics follows a Markovian diffusion process, in Section 2 we propose a nonparametric dissimilarity measure based on an estimate of the Markovian structure of the processes. In Section 3 we examine - via a cluster analysis based on this distance - the similarities of price dynamics inside the Eurozone. In Section 4 a structural change point analysis is conducted in order to show the fragility of the emerging clustering pattern. Figures are collected at the end of the paper.

2 The method

If we assume that inflation dynamics can be modeled as a diffusion process, we can estimate nonparametrically its Markov operator, which is entirely characterized by the drift and volatility of the process itself. A dissimilarity measure between operators is defined, such that a clusterization of the processes can be proposed, that illustrates to what extent belonging to the EMU generates more homogeneous price dynamics. Compared to Palomba *et al.* (2009), this approach is totally nonparametric and hence less model dependent or, in other words, more robust to model misspecification.

2.1 Markov operator

Denote by $X_t = X(t)$ the value of a time series at time t . We assume in this paper that each time series is a diffusion process satisfying the stochastic differential equation of the form $dX_t = b(X_t)dt + \sigma(X_t)dW_t$, $t \geq 0$, where $\{W_t, t \geq 0\}$ is a Wiener process, $b(\cdot)$ (the drift function) and $\sigma(\cdot)$ (the diffusion coefficient or volatility) are sufficiently smooth functions so that the stochastic process is well defined (see Iacus, 2008). Despite being the time series a continuous time process, we assume it is observed at regularly spaced discrete times $t_i = i\Delta$, $\Delta > 0$, $i = 0, 1, \dots, n$: denote by $X_i = X_{t_i} = X(t_i)$ the observation at time t_i . The above process X_t is a Markov process and it is completely characterized by the functions $b(\cdot)$ and $\sigma(\cdot)$ in the sense that the transition density of X_t given $X_s = x$, $s < t$, depends only on $b(\cdot)$, $\sigma(\cdot)$, $t - s$ and x .

While a Markov chain is characterized by a discrete state space (i.e. the sets of possible values taken by the process) and discrete time evolution, the diffusion process takes values on a continuous state space and evolves in continuous time. Therefore, while a Markov chain is associated to its transition matrix, the diffusion process is associated to its *Markov operator*. Let $f(\cdot)$ be any integrable function, then the Markov operator of X_t can be defined as¹

$$P_\Delta f(x) = E\{f(X_t)|X_s = x\}.$$

Now, if one can estimate the Markov operator from the data, one can also possibly identify the process and in particular the invariant law $\mu_{b,\sigma}$ of X_t , which depends only on the couple (b, σ) . For Markov chains, $f(x) = 1_{\{X_i=j\}}$, then $E\{f(X_i)|X_{i-1} = k\} = P(X_i = j|X_{i-1} = k) = p_{jk}$ is the usual transition probability. For diffusion processes, the state k is replaced by the real number x and the transition is given for all possible time lags $t_i - t_{i-1}$, such that the Markov operator can be intuitively associated to a matrix with an (uncountable) infinite number of rows and columns.

De Gregorio and Iacus (2010) proposed a dissimilarity measure between two Markov processes, based on the distance between the estimators of P_Δ for two time series. In fact, for a given L^2 -orthonormal basis of functions $\{\phi_j(\cdot), j \in J\}$, where J is an index set, it is possible to obtain the matrix $\hat{\mathbf{P}}_\Delta(\mathbf{X}) = [(\hat{P}_\Delta)_{j,k}(\mathbf{X})]_{j,k \in J}$, where

$$(\hat{P}_\Delta)_{j,k}(\mathbf{X}) = \frac{1}{2N} \sum_{i=1}^N \{\phi_j(X_{i-1})\phi_k(X_i) + \phi_k(X_{i-1})\phi_j(X_i)\}, \quad j, k \in J. \quad (1)$$

¹Notice that P_Δ depends on the transition density from X_t to X_s , so we put explicitly the dependence on $\Delta = t - s$ in the notation.

The matrix $\hat{P}_\Delta(\mathbf{X})$, which is a nonparametric estimator of P_Δ (Gobet *et al.*, 2004), can be used as a “proxy” of the probability structure of the model.

Therefore, we can introduce the following dissimilarity measure (De Gregorio and Iacus, 2010). Let \mathbf{X} and \mathbf{Y} be discrete time observations from two diffusion processes: the *Markov operator dissimilarity* is

$$d_{MO}(\mathbf{X}, \mathbf{Y}) = \left\| \hat{P}_\Delta(\mathbf{X}) - \hat{P}_\Delta(\mathbf{Y}) \right\|_1 = \sum_{j,k \in J} \left| (\hat{P}_\Delta)_{j,k}(\mathbf{X}) - (\hat{P}_\Delta)_{j,k}(\mathbf{Y}) \right|, \quad (2)$$

where $(\hat{P}_\Delta)_{j,k}(\cdot)$ is calculated as in (1) separately for \mathbf{X} and \mathbf{Y} .

2.2 Dissimilarity between cluster solutions

In order to compare the clusterings proposed in different studies about the inflation dynamics in the Eurozone (Busetti *et al.*, 2007; Palomba *et al.*, 2009), we use the cluster similarity index proposed in Gavrilov *et al.* (2000) defined as follows. Given two clusterings $C = C_1, \dots, C_K$ and $C' = C'_1, \dots, C'_{K'}$, we compute the following quantities:

$$\text{sim}(C_i, C'_j) = 2 \frac{|C_i \cap C'_j|}{|C_i| + |C'_j|}, \quad i = 1, \dots, K, j = 1, \dots, K',$$

and the final cluster similarity index is given by the formula

$$\text{Sim}(C, C') = \frac{1}{K} \sum_{i=1}^K \max_{j=1, \dots, K'} \text{sim}(C_i, C'_j). \quad (3)$$

Both $\text{sim}(\cdot)$ and $\text{Sim}(\cdot) \in [0, 1]$. Note, besides, that the number of units and groups may differ between the two clusterings.

3 Clustering the short-run inflationary processes

We examine the inflation processes of the EU-27 countries over the period December 1996 to November 2012, using the monthly HICP (Harmonized Index of Consumer Prices) data provided by the ECB Statistical Data Warehouse².

Consider first a cluster analysis of the Markovian operators on the whole period, in

²See <http://sdw.ecb.europa.eu/browse.do?node=9138811>. Data are seasonally unadjusted, since seasonality is not relevant to the techniques adopted in the paper.

order to let the difference between Euro and non-Euro countries emerge (See the dendrogram in Fig. 1). The separation of the non-Euro Eastern Europe countries is quite neat, with the only exception of Estonia, which still joined the Eurozone only in 2011. Inside the main group we find, in a separate small cluster, two countries that do not belong to the Eurozone (United Kingdom and Czech Republic). Three other clusters can be identified:

- the first covers the largest part of the Central and Northern EU. Sweden and Denmark, though not belonging to the EMU, are included in this cluster;
- the second is formed by Southern Europe countries (Italy, Portugal, Cyprus, Malta), with the addition of the Netherlands;
- the third is made by Southern and Eastern countries and also includes Luxembourg and - as a sort of singleton - Ireland. Poland is the only non-Euro country included.

Figure 1 about there

The evolution of the pattern can be followed by decomposing the analysis into sub-periods. We have analyzed separately the pre-Euro period Dec. 1996 to Dec. 1998 (Fig. 2-(a)) and the subsequent Jan. 1999 to Nov. 2012 period (Fig. 2-(b)), which has been furtherly decomposed into Jan. 1999 to Dec. 2004 (Fig. 2-(c)) and Jan. 2005 to Nov. 2012 (Fig. 2-(d)).

Two persistent features characterize the clusterization:

- an Eastern Europe cluster, composed by non-Euro countries, is formed. Slovenia and Slovakia leave the cluster when they enter EU and adopt Euro³. As a partial exception, Estonia - which adopt Euro only in 2011 - is assimilated to the Central-Western countries in the 1999-2004 period, but is still grouped with the Eastern countries in the subsequent period.
- among the Central-Western countries, two main clusters can be identified, grouping the old EU-15 countries: the first include Germany, France, Belgium, Austria, Finland, Sweden, Denmark, Uk; the second is composed by Southern Europe countries together with Ireland, the Netherlands and Luxembourg. It is worth noting the strict similarity with the clusterization proposed by Buseti *et al.* (2007) (notice, in Tab.1, the high values of the cluster similarity index between our Fig. 1 and Fig.

³The two countries entered the EU in 2004. Slovenia adopted the Euro in 2007, Slovakia in 2009.

2-a,b,c, and their clusters): they apply stationarity tests to investigate whether inflation differentials among the EMU countries changed since 1998, i.e. due to the adoption of Euro, identify the same two groups and name them respectively “low inflation rates” and “high inflation rates” countries⁴. The new EU members - which in some cases enter the Eurozone between 2007 and 2009 - are included, with few exceptions, into the Southern Europe cluster.

Figure 2 about there

Comparing the results to Palomba *et al.* (2009), who clusterize the EU-25 countries according to a similarity measure of short-run inflation rates in the 1999-2006 period, we found only a partial correspondence: in fact, while both studies identify a separate Eastern countries cluster, Palomba *et al.* (2009) classify the Central-Western countries into three subgroups with no particular geographical proximity. Consequently, the value of the cluster similarity index is never higher than 0.5 (see Tab.1).

It should be noted, anyway, that the clusterization of the Central and Western countries is not as much sharp in the last period (2005-2012), when several countries seem to belong to the “wrong” group. The cluster similarity index, in fact, comes to its lowest values (0.49) comparing the 1999-2004 and the 2005-2012 periods (see Tab.1). This raises the suspect of a recent increase in the inflation volatility, which we test evaluating the change points of the inflationary processes of the single countries.

4 Structural change point analysis

We want to test whether the volatility of the single-country inflationary processes has significantly changed after the entrance of the country in the Eurozone. This means to calculate the “change points” of the processes. In other words, we have to verify whether at some time τ the structure of the stochastic model changes.

Formally, the existence of a change point at time $\tau \in [0, T]$ means that the process X_t satisfies $dX_t = b(X_t)dt + \sigma(X_t, \theta_1)dW_t$ for $t \in [0, \tau)$ and $dX_t = b(X_t)dt + \sigma(X_t, \theta_2)dW_t$ for $t \in [\tau, T]$, where for simplicity we assume that the diffusion coefficient has a multiplicative form, i.e. $\sigma(X_t, \theta) = \sqrt{\theta}\sigma(X_t)$, so that the change in volatility is captured by the

⁴Italy is a sort of singleton in Busetti *et al.* (2007), while is included in the Southern countries cluster in our analysis.

scaling factor θ . The coefficients $b(\cdot)$ and $\sigma(\cdot)$ are supposed to be unknown and estimated nonparametrically as in De Gregorio and Iacus (2010).

Figure 3 about there

The results show a recent increase in price volatility, particularly for the EU-15 countries (see Fig. 3). For 21 out of the EU-27 countries, in fact, the change point corresponds to an increase in volatility (i.e. $\theta_1 < \theta_2$); for 18 countries the change points are distributed between 2004 (the year of EU enlargement to 10 new countries, mostly from Eastern Europe) and 2008 (the beginning of the global economic crisis).⁵ In particular, 11 out of the EU-15 countries exhibit an increase in inflation volatility between 2004 and 2008.

We can conclude that, starting from 2004, a higher instability has characterized the European price dynamics and has perturbed the short-run inflation pattern previously registered.

Table 1: Cluster similarity index

	FIG1	FIG2a	FIG2b	FIG2c	FIG2d	Busetti <i>et. al.</i>	Palomba <i>et. al.</i>
FIG1	1.00	0.60	0.78	0.58	0.66	0.75	0.43
FIG2a		1.00	0.60	0.76	0.50	0.85	0.50
FIG2b			1.00	0.53	0.68	0.87	0.43
FIG2c				1.00	0.49	0.75	0.50
FIG2d					1.00	0.42	0.45
Busetti <i>et. al.</i>						1.00	0.50
Palomba <i>et. al.</i>							1.00

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⁵See Eickmeier and Breitung (2006) for a comparison.

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Figure 1: Cluster analysis of period December 1996 - November 2012

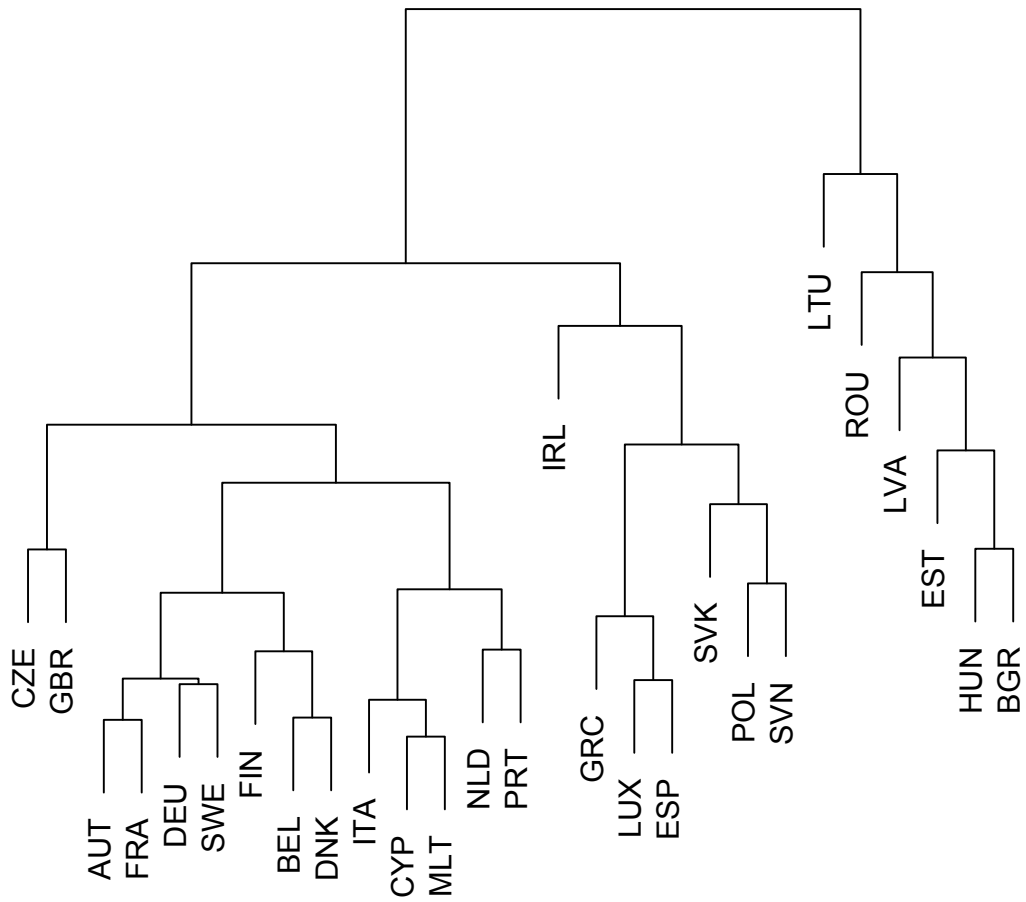


Figure 2: Cluster analysis by subperiods

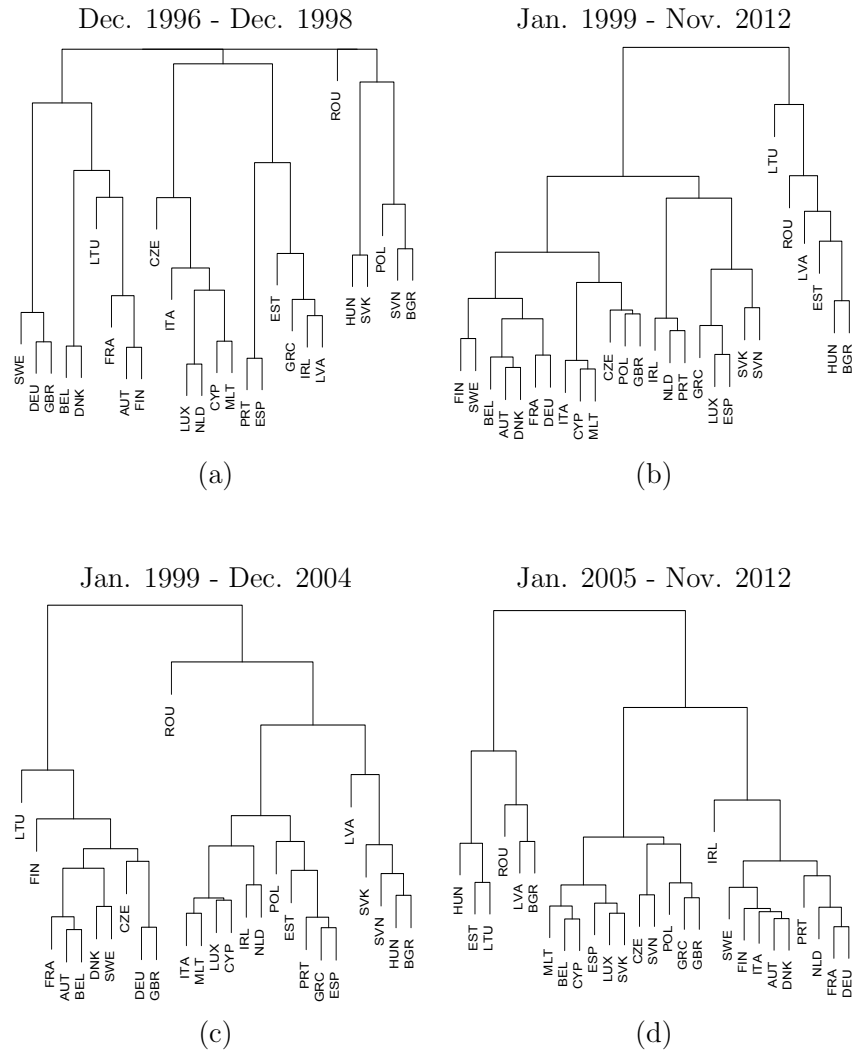


Figure 3: Change points of EU-27 countries in 1997-2012. Countries above the timeline have $\theta_1 < \theta_2$, i.e. volatility increases after the change point. Countries below the timeline have $\theta_1 > \theta_2$, i.e. volatility decreases after the change point. Numbers after the country name indicate the month; e.g., FIN 04: the Finland change point falls on April 2001.

