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Revisiting inflation in the euro area allowing for long memory.[*]

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Abstract

We analyse inflation and inflation differentials in the euro area allowing for long memory and a new type of limiting theory denoted fixed-bandwidth. Our results differ from the traditional ones based on standard normal asymptotics and the short memory assumption, and we also find that the inflation differentials between "core" and "peripheral" countries are strongly persistent. "Core" economies appear to have less persistent differentials and may be more integrated, while "peripheral" countries with high inflation may find themselves under competitive pressure for a long time.

Keywords: Inflation persistence, inflation differentials, euro area, fractional integration and cointegration.

JEL classification: C32, E31

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1. INTRODUCTION

In the last few years, inflation and inflation differentials in the euro area have raised the attention of many theoretical and empirical researchers. One of the main reasons behind this recent interest is that the European Central Bank (ECB) is a sovra-national institution to which member states have transferred powers: therefore it is important to monitor that this institution is acting in accordance with its mandate, that is targeting inflation as stipulated. However, monitoring national levels of inflation is equally important: even if the ECB succeeds in stabilizing inflation for the whole euro area, if one member state is significantly distant from the target, its participation to the euro may be destabilizing. In support of this view, Fendel and Frankel (2009) found that the ECB took inflation differentials into account, for example holding back a restrictive monetary policy if this could have caused deflation in member states with low inflation.

The existence of inflation differentials poses a threat to the sustainability of the euro within the European Union because, without the possibility of devaluating the currency of the country with higher inflation, and given the currently narrow scope for fiscal transfers or other forms of adjustments, inflation differentials can eventually push some countries into extreme recession or in bankruptcy, see for example Wickens (2010), or Coudert, Couharde, and Mignon (2013), who showed that, after considering the role of macroeconomic fundamentals, currency misalignments and differential persistence have increased after the introduction of the euro. Quint (2016) argues that the tension generated from having a common monetary policy, instead of one designed to target inflation at national level, is not more than what is observed within the US or what had been observed within Germany when Bundesbank run its monetary policy. However, this is not reassuring, as the euro area, contrary to the US or Germany, does not have area-wide mechanisms, such as relevant fiscal transfers, to counter this stress.

Much literature takes the existence of inflation differentials in the euro area as a stylized fact and focuses on investigating the causes. The reasons for inflation differentials in the euro area are many: they may reflect long term convergence of productivity and prices, especially in the early stages of the monetary union, or structural factors, such as differences in the labour market and different participation to trade outside the euro area, and also divergences originated by different macroeconomic policies, see for example ECB (2003) and ECB (2005); Andersson, Masuch and Schiffbauer (2009) mentioned the effect of administered prices and the lack of synchronization in the economic cycle. These differences may make it more probable for some countries to be hit by idiosyncratic shocks: for example, Honohan and Lane (2003) found that higher openness to trade outside the euro area was a major source of inflation differentials for Ireland.

The role of structural differences in the labour or goods markets was also discussed by Pirovano and Van Poeck (2011); for Angeloni and Ehrmann (2007), demand, supply or exchange rate shocks, in this order, have been the main sources of inflation differentials. We refer to de Haan (2010) for a recent survey and further discussion. Interestingly, although the determinants of inflation differentials are many and possibly specific to each country, structural factors generating inflation persistence are themselves a source of persistent inflation differentials: even if two countries were hit by the same shock, different inflation persistence would eventually generate a visible inflation differential. Indeed, Angeloni and Ehrmann (2007) identify these as the main vehicle propagating inflation differentials.

In this paper we analyse the dynamics of inflation for each country, and with respect to the other member states, discussing both the existence and the persistence of inflation differentials. It is commonly taken for granted in the applied literature that some of these differentials are different from zero in the long run, and indeed standard asymptotic theory, according to which the inflation series are treated as weakly autocorrelated processes, would broadly support this assumption, as we also find in our analysis. Note that weak autocorrelation is characterized by a fast decay of the autocorrelations of the processes, therefore displaying the so-called short memory property. A traditional example of this type of behaviour is the stationary and invertible finite order autoregressive and moving average process. However, conclusions from standard analyses based on the assumption of weak autocorrelation need to be taken with caution for at least two reasons. First, assuming short memory, as it is done in most empirical works, might be in some cases appropriate, but it is important to emphasize that this is a very particular case of dependence. More general processes, like the fractional integrated proposed by Granger and Joyeux (1980), have recently featured prominently: the dependence on those processes is mainly driven by a memory parameter δ, so they might display short memory ($\delta = 0$), long memory ($\delta > 0$), for which the autocorrelations decay slowly, or be antipersistent ($\delta < 0$). Thus, taking an agnostic approach about the type of dependence which characterizes the process under study might avoid a problem of misspecification, especially noting that long memory in the series would distort the size of the standard test procedures based on the short memory assumption, leading for example to spurious rejections of the corresponding null hypotheses, see for example Wright (1998). Moreover, allowing for long memory in this empirical investigation seems of particular importance also because of the emphasis given to inflation persistence in the empirical literature: in the long memory framework the order of integration provides a natural, simple and intuitive measure of persistence, that we can be easily compared across countries.

The second piece of warning is that standard inference for weakly autocorrelated processes is typically based on testing procedures where the corresponding studentized (by means of an adequate nonparametric estimator of the long run variance) and centered sample means are assumed to have a $N(0,1)$ limiting behaviour. This is a direct consequence of relying on a consistency argument for the estimator of the long run variance, which is typically achieved by means of an increasing smoothing. As Kiefer and Vogelsang (2005) indicate, this might lead to tests suffering from an important size distortion.

Recently, Hualde and Iacone (2017) have dealt with both concerns. From one side, they proposed a test procedure for the mean of a general covariance stationary process which might exhibit short memory, long memory or antipersistency, where it is not required to make a priori assumptions about the type of dependence characterizing the data. From the other, their test procedure is based on a smoothed periodogram estimator of the long run variance, where, unlike in the traditional approach, the degree of smoothing is kept fixed. This leads to a limiting distribution, denoted fixed-bandwidth, which appears to be closer than the more traditional one (large-bandwidth) to the true sampling distribution of the studentized sample mean. This finding complements similar results achieved by McElroy and Politis (2012, 2103) and, interestingly, opens the door to revisiting previous empirical evidence, allowing now for an agnostic view about the type of dependence characterizing the series under study and using a more accurate limiting result. Thus, in the present paper we analyze along these lines inflation levels and inflation differentials in the euro area, with the aim of confirming or questioning previous well established evidence.

In the next section we carry out the empirical analysis and in Section 3 we present the main conclusions.

2. AN APPLICATION TO INFLATION DIFFERENTIALS IN THE EURO AREA

We collected monthly inflation at annual rates (i.e., 12 times the monthly inflation rate, defined as the 100 times the difference of the logarithm of the price index) of the euro area countries, for the period January 1999 - October 2015. We report in Table 1 the sample means of those inflation rates for each country ($\overline{\pi}$). Specifically, we construct the inflation rate from the series of the Harmonized Index of Consumer Prices (All Items), which were collected from the FRED database, at the Federal Reserve Bank of St. Louis, and are from Eurostat. In the tables we use the acronyms EU, BG, FR, OE, FN, GR, BD, IT, IR, LX, NL, ES, PT for the euro area, Belgium, France, Austria, Finland, Greece, Germany, Italy, Ireland, Luxemburg, the Netherlands, Spain, Portugal, respectively (the series for the euro area is adjusted according to changes in the membership).

Our purpose is twofold. First we analyze whether there are significant deviations from the long run inflation target. Then we explore whether inflation differentials between pairs of countries are significant. Specifically we test for the null hypotheses that the corresponding population mean deviations from the target or that population mean inflation differentials are zero. For this purpose we will use three different testing procedures where, in all the cases, we consider two sided alternatives. In the first test procedure, we use a studentized sample mean, where the long run variance is estimated by a weighted periodogram as in Hualde and Iacone (2017). Letting x_t, $t = 1, ..., T$, be the difference between the inflation and the target or the differential between the inflation rates of two countries, denote by $w_x(\lambda) := (2\pi T)^{-1/2} \sum_{t=1}^{T} x_t e^{i\lambda t}$ the Fourier transform of x_t and by $I(\lambda) := |w_x(\lambda)|^2$ the periodogram. We use a simple estimator of the long-run variance σ^2, given by $\widehat{\sigma}^2 := 2\pi m^{-1} \sum_{j=1}^{m} I(\lambda_j)$, where $\lambda_j := 2\pi j/T$ are Fourier frequencies, for a user-chosen m which will be specified below. Letting $\overline{x} = T^{-1} \sum_{t=1}^{T} x_t$ be the sample mean, the first testing strategy, denoted as ν, confronts the statistic $\sqrt{T}\overline{x}/\widehat{\sigma}$ with the $N(0, 1)$ limiting distribution. This strategy is based on the assumption that x_t is short memory and also that $m \to \infty$ (but $m/T \to 0$) as $T \to \infty$: this leads to consistency of $\widehat{\sigma}$, so the replacement of σ by $\widehat{\sigma}$ in the test statistic does not affect its standard normal asymptotic distribution.

However the test ν is only appropriate if $\delta = 0$, where δ denotes the memory of x_t. Otherwise, when $\delta > 0$ and $m \to \infty$, then $\sqrt{T}\overline{x}/\widehat{\sigma}$ diverges as $T \to \infty$ even if the null hypothesis is true, giving rise to spurious rejections of the null hypothesis (cases with $\delta < 0$ are also a concern, as the size goes to 0, with an adverse effect on power). For this situation, Robinson (2005) proposed the feasible Memory Autocorrelation Consistent (MAC) standardized sample mean instead, which again is asymptotically standard normal. In what follows we refer to this test procedure as θ.

Despite the appealing asymptotics, Hualde and Iacone (2017) showed that the tests ν and θ might suffer from size distortion in finite samples. To address this problem, we also use Hualde and Iacone's (2017) approach, where the null limiting distribution of $\sqrt{T}\overline{x}/\widehat{\sigma}$ is derived keeping m fixed, allowing also for the possibility that x_t might display short memory, long memory or be antipersistent. This third test strategy will be referred as τ, and can be viewed as an improved version of ν, τ enjoying a better finite sample behaviour than ν. When $\delta = 0$ the fixed-bandwidth limit distribution is a Student-t with $2m$ degrees of freedom; when $\delta \neq 0$ the limit distribution is not standard, but critical values can be computed by simulation. The Monte Carlo exercise in Hualde and Iacone (2017) finds that these alternative critical values help improving the size, and that the empirical size best approximates the theoretical one when $m = 1$. However, as the power increases with m, there is a trade off between correct size and the best power. Tests θ and τ are appropriate for $\delta \in (-0.5, 0.5)$: for $\delta \geq 0.5$ the sample average $\overline{x}$ is not a consistent estimate of the population mean.

2.1. Inflation targeting

The ECB is committed to inflation below, but close to, 2% in the medium term. In the first analysis we check whether this target is met not only for the euro area as a whole, but for all the euro area countries separately. While the definition of the inflation target is only formulated for the euro area, we also check whether the target holds for all the member states individually because the deviation from the 2% is a rough measure of an inflation differential with respect to the rest of the union. Thus we test whether the mean inflation of euro area countries meets the 2% target, against the two sided alternative that it deviates from it, performing the test with three different procedures, ν, θ and τ.

As discussed before, all tests require the use of bandwidths. In particular ν was implemented for the choice $m = 15$ (for $m \in [10, 14]$, ν leads to identical conclusions). Our bandwidth choice is driven by the aim of avoiding the interference due to seasonality: monthly data typically lead to a seasonal peak in the periodogram at a frequency between Fourier frequencies λ_{16} and λ_{17}, so choosing a bandwidth smaller or equal than 15 limits the impact of this peak on the estimation of the long run variance.

Tests θ and τ require the estimation of δ. These estimates (reported in Table 1 as $\widehat{\delta}_l$) were computed by the local Whittle method, see Robinson (1995), with bandwidths $l \in [10, 15]$. Given that $\widehat{\delta}_l$ are consistent estimators of δ whenever $\delta \in (-0.5, 1)$, see Velasco (1999), the local Whittle method is implemented for an optimizing set chosen as $[-0.499, 0.999]$. In most cases the estimated values $\widehat{\delta}_l$ are positive and in all cases the estimated values are below the upper bound of the optimizing interval, suggesting mean reversion but at a speed that is much slower than it is commonly considered. For example, for $l = 15$, in two cases (GR and IT) the estimated values fall just behind the stationarity threshold $\delta = 1/2$, and in one case (IR) it falls further ahead. However, for $l \in [10, 14]$ all the estimated values for IT and four out five for GR were below the stationarity threshold, while all the outcomes were beyond 1/2 in the case IR. Larger values of δ may signal more difficulty to revert to the long term mean: this may occur because these countries are less willing or less able to counter inflation differentials, or because they entered the euro area with inflation larger than the average. This "structural" interpretation for δ is also motivated by an empirical fact: grouping the European countries as "core" (BD, FR, BG, NL, LX, OE) and "periphery" (FN, GR, IT, ES, PT and IR), the estimates of δ are always larger in the "periphery" (with the exception of NL), and, interestingly, among those, the five largest estimates correspond to the five countries that have been most under pressure during the financial crisis.

In addition, apart from an estimate of δ, test θ requires the estimation of the contribution of the short memory component to the frequency zero of the spectral density of x_t. Again, this requires the choice of a bandwidth which we also fixed to $l = 15$ (choices $l \in [10, 14]$ lead to identical conclusions). Under regularity conditions (see Robinson, 2005), θ has a $N(0, 1)$ null limiting distribution.

For test τ we used $m = 1$ and fixed-m null limiting distribution (we also experimented with $m = 2$, $m = 3$ and $m = 4$ but the outcomes did not change). Note that the null limiting distribution of τ depends on δ. In order to compute the critical values we use the plug-in approach of McElroy and Politis (2012, 2013) using $\widehat{\delta}_{15}$ as the estimate of δ. Finally, we will use a 5% significance level throughout.

Testing whether the inflation means meet the 2% target, we only found significance for BD, and just with the standard ν test procedure. Overall, the results corresponding to tests τ and θ suggest that we cannot reject the hypothesis that the ECB succeeded in stabilizing inflation near 2%, although this may have been at a very slow speed, especially for FN, GR, IT, IR, NL, ES and PT. In fact, it can be shown that the power of θ and τ, diminishes for larger δ, although this appears to be less worrisome than the spurious rejections that might originate $\delta > 0$ when using ν. The results of procedure ν would be received with more concern, as it would lead to the conclusion that BD, whose mean inflation over the period considered is clearly the smallest one (1.52%), has gained a competitive advantage over the years, and may be putting pressure on the European partners. Taking persistence into account, as it is done in tests τ and θ, is therefore very important, because it might indicate that the deviation of BD from the 2% benchmark hinted by the test ν could be just due to a slow convergence to the target.

Additionally, these results also provide an alternative interpretation of findings in Busetti et. al. (2007), who identified two different "convergence clubs" approximately similar to the "core" and "periphery" groups listed here. Busetti et. al. (2007) suggested the convergence to two different means in the euro area: we do not find a strong evidence supporting this claim in our extended sample (although, as will be seen, the analysis of inflation differentials will offer a somewhat different picture), but it is true that the speed of convergence to the mean differs substantially in "core" or "periphery" countries, being much slower for the latter ones (with the possible exception of NL).

2.2. Inflation differentials

We extend the analysis by looking at pairwise mean inflation differentials: again, we estimate the orders of integration first and report them (just for bandwidth $l = 15$) in Table 2. On a few occasions the estimated orders of the differentials are negative: interestingly, this happens for all the combinations of the "core" countries BD, BG, OE, LX, as well as with the FR-LX and IT-LX pairs. In other cases, we at least find that the estimated orders of integration of the differentials are lower than the orders of the two countries taken separately: the largest estimated drop is for BD-BG, and indeed we observe 9 such instances (out of 15 potential ones) when "core" countries are involved; the largest estimated drop outside "core" countries is for ES-PT, and differentials ES-IT, ES-GR and IT-FR also seem to be characterized by a numerically relevant change in persistence. We also experimented with bandwidths between $l = 10$ up to $l = 14$ obtaining qualitatively similar results.

These findings have an interesting economic interpretation: for countries that are economically well integrated, inflation shocks quickly transmit from one to the other, and the persistence of these differentials may well be less than the persistence of inflation itself. In this case, the two inflation rates are cointegrated. Note that we consider here a general version of cointegration, which occurs whenever a linear combination of processes displays a smaller memory than the processes themselves. The estimates in Table 2 seem to suggest that economic integration is stronger at the "core", especially within BD-OE-BG-LX, but the pair ES-PT also enjoys a strong linkage.

Next, we test for zero mean inflation differentials between pairs of countries' inflation rates. We report the corresponding results for the significance of the three tests in Table 3 for the same choice of bandwidths as in Subsection 2.1 (superscripts τ, θ and ν indicate significance at 5% with the corresponding test procedure whereas 0 means that no significance was found). Overall we found 5 significances with the τ test, 8 with θ and 17 with ν. Results for this latter test are robust to bandwidth choice: for example we found 15 and 16 significances with $m = 10$ and $m = 11$, respectively. With $m = 2$ and $m = 3$, the significances with the τ test were 7 for both cases, so even with larger bandwidths or with the test θ, we find weaker statistical evidence of heterogeneity than by the more traditional method ν.

In order to interpret our statistical findings, we begin by comparing the pairwise differentials with BD, as this seems the natural benchmark, given its prominent role in the European trade: we find 2 significances with the τ test (3 significances for both $m = 2$ and $m = 3$), 3 with θ and 8 with ν (7 significances for both $m = 10$ and $m = 11$). Interestingly, the strongest evidence of significant inflation differential with BD is for

BG and LX, which are considered as "core" members of the European Monetary Union, and historically have always been part of the Deutsche Mark area of influence; a third possible significant differential is with OE, again historically a strong trade partner of BD.

Most of the other significant differentials involve LX, again mostly with "core" countries (besides BD, FR and possibly BG and OE, but also with IT), and finally for IT and FR, two founding members of the European Union and well integrated trade partners. On the other hand, we find that the differential between BD and GR is not significant, despite being the second largest in the euro area (and the largest differential altogether, when LX is ignored). Summarizing these results, we find that using θ, τ, all but one significant inflation differentials are for "core" or well integrated countries. This may seem surprising, and it contrasts with the outcome of the conventional ν test with normal asymptotics, for which many rejections of the null hypothesis are indeed for "peripheral" countries versus "core" ones (FR and BD versus GR, ES and PT).

The estimates of the inflation differentials' memories help to explain and interpret these apparently incongruous findings: rejections of the null hypothesis for the τ and θ tests mostly occur when the estimate $\widehat{\delta}$ is close to zero or possibly even negative. From a statistical viewpoint, this simply reflects the fact that tests τ or θ have more power, that is, it is easier to reject an incorrect null hypothesis, the smaller δ is. Note however that with the exception of the IT-LX case, whenever the null is rejected by θ and/or τ, ν reaches the same conclusion. Thus these results appear to be quite robust.

In order to shed some additional light on this apparent empirical puzzle, we provide an economic interpretation of our findings by means of a brief cointegration analysis. In particular, we first focus on the relationships for which the null of zero mean inflation differential is rejected with θ and/or τ, that is BG-BD, LX-BG, FR-IT, LX-FR, OE-BD, LX-OE, LX-BD and LX-IT. For this purpose we apply the semiparametric procedure of Hualde and Velasco (2008), which tests the null of no cointegration. Two of the main ingredients of this method are the estimates of the (possible) cointegrating parameters and of the memory of the (possible) cointegrating error. We report these estimates in Table 4. Specifically, $\widehat{\beta}_l$ is the narrow band least squares (NBLS) of the (possible) cointegrating parameter for bandwidths $l \in [10, 15]$. As explained by Robinson (1994), a more standard estimation approach like the ordinary least squares might be inconsistent in cases of cointegration between stationary processes (the so-called stationary cointegration), while the NBLS retains consistency. Also, in Table 4 we present the estimates of the memory of the (possible) cointegrating error ($\widehat{\gamma}_l$), estimated by local Whittle applied to the NBLS residuals for choices $l \in [10, 15]$. Note that for each of the above pairs, the inflation corresponding to the first country in the relationship takes the role of dependent variable in the regressions from which the cointegrating parameters are estimated, which also allow for the possibility of a nonzero constant. We omit the results for the pair LX-OE, for which the values of $\widehat{\gamma}_l$ are in general larger than the estimated memories for the inflation of LX and OE (in fact, the test for no cointegration never rejects here). In all other cases, $\widehat{\gamma}_l$ are smaller than at least one of the estimated memories of the corresponding inflations, which supports the likelihood of cointegration.

For the BG-BD case, this is confirmed by the cointegration test which for the six bandwidths considered, rejects (at 1%) in 4 cases. Note also that the estimates of the cointegrating parameter are relatively far from one, which might be coherent with the existence of a nonzero mean inflation differential even if the inflation rates are linked in the long run. Actually, this seems to be also the case for the relations LX-FR (5 significances) and LX-BD (6 significances).

For LX-BG the evidence of cointegration is weaker (2 significances) and the estimates of the cointegrating parameter are much closer to one, which also occurs for OE-BD (again 2 significances). Heuristically, for these two cases, looking at the $\widehat{\beta}_l$'s, the evidence in favour of a nonzero mean inflation differential is weaker and, in fact, τ does not reject. Finally, for FR-IT the test does not show any evidence of cointegration, whereas for LX-IT the evidence is weak (2 significances).

The cointegration analysis complements the conclusions extracted from the results of the test for no mean inflation differential. Inflation rates in "core" economies are characterized by very little persistence, they appear to share a long run equilibrium and adjust towards this equilibrium fairly quickly. Following for example Angeloni and Ehrmann (2007), we can therefore conclude that in these cases, even when statistically significant, an inflation differential may be relatively little and of limited concern; conversely, regardless of whether the long run average of the differential is zero or not, in situations of strong persistence of the inflation differential, the countries with high inflation may find themselves under competitive pressure for a very long time. This might be indeed the cases of GR, IT, for which, apart from the outcome of the ν test (perhaps spurious), their inflation rates do not show any evidence of being cointegrated with that of BD. Similar threat could also affect ES, or even FN, NL and PT, for which the evidence of cointegration with BD is very weak (1 significance in the case of ES, 2 for the rest).

3. CONCLUSION

We analysed inflation and inflation differentials in the euro area allowing for long memory and cointegration: we find that the persistence of these differentials is heterogenous, and the results are consistent with a situation in which the "core" economies are more integrated, and "peripheral" economies are characterized by more persistent inflation differentials. Although we find less evidence of statistically significant differentials than we would if we did not allow for long memory, we also find that countries with high inflation persistence may find themselves under competitive pressure for more time than would have been anticipated without considering long memory techniques.

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Table 1. Averages of inflations and estimated memory by country for $l \in [10, 15]$

	EU	BG	FR	OE	FN	GR	BD	IT	IR	LX	NL	ES	PT
$\overline{\pi}$	1.83	1.96	1.63	1.88	1.85	2.36	1.52	2.06	1.99	2.47	2.01	2.32	2.13
$\widehat{\delta}_{10}$	.257	-.175	.032	.041	.568	.516	.257	.309	.863	-.008	.683	.310	.457
$\widehat{\delta}_{11}$	.290	-.125	.085	.058	.449	.487	.319	.350	.796	.004	.555	.303	.551
$\widehat{\delta}_{12}$	.288	-.103	.099	-.072	.517	.492	.266	.440	.686	-.060	.524	.294	.577
$\widehat{\delta}_{13}$	.335	-.030	.159	-.047	.328	.470	.261	.417	.617	-.047	.532	.313	.451
$\widehat{\delta}_{14}$	.367	.052	.172	-.025	.275	.491	.317	.461	.623	-.026	.398	.371	.399
$\widehat{\delta}_{15}$	.368	.147	.183	.020	.305	.510	.189	.520	.597	.000	.246	.427	.417

Table 2. Estimated orders of integration of inflation differentials, $l = 15$

	FR	OE	FN	GR	BD	IT	IR	LX	NL	ES	PT
BG	.104	−.156	.168	.423	−.338	.094	.615	−.161	.149	.282	.302
FR		.087	.301	.448	.028	.007	.691	−.250	.203	.255	.354
OE			.056	.430	−.252	.076	.526	−.037	.225	.446	.293
FN				.399	.202	.485	.624	.243	.146	.436	.618
GR					.471	.397	.528	.156	.546	.190	.271
BD						.171	.665	−.173	.284	.351	.424
IT							.569	−.269	.127	.177	.365
IR								.376	.360	.456	.477
LX									.019	.009	.127
NL										.169	.141
ES											.091

Table 3. Significant inflation differentials with the ν, θ, τ test strategies

	FR	OE	FN	GR	BD	IT	IR	LX	NL	ES	PT
BG	ν	0	0	0	τ,θ,ν	0	0	θ,ν	0	0	0
FR		0	0	ν	0	τ,θ,ν	0	τ,θ,ν	0	ν	ν
OE			0	0	θ,ν	0	0	θ,ν	0	0	0
FN				0	0	0	0	ν	0	0	0
GR					ν	0	0	0	0	0	0
BD						ν	0	τ,θ,ν	ν	ν	ν
IT							0	τ,θ	0	0	0
IR								0	0	0	0
LX									0	0	0
NL										0	0
ES											0

Table 4. Estimates of the (possible) cointegrating errors' memories and parameters

	$\widehat{\gamma}_{10}$	$\widehat{\gamma}_{11}$	$\widehat{\gamma}_{12}$	$\widehat{\gamma}_{13}$	$\widehat{\gamma}_{14}$	$\widehat{\gamma}_{15}$	$\widehat{\beta}_{10}$	$\widehat{\beta}_{11}$	$\widehat{\beta}_{12}$	$\widehat{\beta}_{13}$	$\widehat{\beta}_{14}$	$\widehat{\beta}_{15}$
BG-BD	-.495	-.499	-.429	-.390	-.337	-.362	1.29	1.27	1.29	1.25	1.24	1.10
LX-BG	-.174	-.140	-.164	-.157	-.152	-.158	.848	.868	.913	.935	.949	.952
FR-IT	-.206	-.134	-.149	-.120	-.077	-.134	.808	.806	.817	.799	.811	.808
LX-FR	-.083	-.072	-.092	-.128	-.055	-.242	1.30	1.34	1.41	1.43	1.45	1.37
OE-BD	.005	-.105	-.117	-.286	-.276	-.270	.983	.972	1.04	.946	.957	.904
LX-BD	.061	-.020	.001	.026	-.009	-.185	1.40	1.40	1.50	1.52	1.52	1.30
LX-IT	-.257	-.165	-.298	-.215	-.156	-.227	1.09	1.13	1.16	1.20	1.23	1.19