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Byrne, Joseph P.; Fiess, Norbert

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Euro area inflation: aggregation bias and convergence

Joseph P. Byrne · Norbert Fiess

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Abstract Monetary policy of the European Monetary Union targets aggregate euro area inflation. Concerns are growing that a focus on aggregate inflation may cause national inflation rates to diverge. While different explanations for diverging aggregate euro area inflation have been brought forward, the very impact of aggregation on divergence has, however, not been studied. We find a striking difference in convergence depending on the level of aggregation. While aggregate national inflation rates are diverging, disaggregate inflation rates are converging. We find that aggregation appears to bias evidence towards non-convergence. Our results are consistent with prominent theoretical and empirical evidence on aggregation bias.

Keywords Euro area inflation · Aggregation bias · Convergence

JEL classification C12 · C22 · E31

1 Introduction

> "One significant feature of euro area inflation differentials is their persistence"
>
> Jean-Claude Trichet, President of the European Central Bank, 31 March 2006.

Since the inception of European Monetary Union (EMU), the European Central Bank (ECB) has focused on aggregate euro area inflation. There are growing

J. P. Byrne (✉) · N. Fiess
Department of Economics, University of Glasgow, Adam Smith Building,
Glasgow G12 8RT, UK
e-mail: j.byrne@lbss.gla.ac.uk

N. Fiess
World Bank, Washington, DC, USA

concerns that by focusing on aggregate data, national inflation rates may actually diverge. Such divergence, if persistent, could prove harmful to monetary union in the long term. While different explanations for diverging euro area inflation rates have been brought forward, to date, the very impact of aggregation on this persistence has not been studied. Aggregation is important since it may act as a veil and suggest persistent inflation differentials when disaggregate inflation rates actually converge. This issue is highly relevant to current euro area members, as adherence to an inappropriate model can be costly. It is also of particular interest to potential future EMU entrants as well as others interested in emulating an EMU style currency union.

The primary objective of the ECB is to achieve stable inflation and initial evidence suggests that it has been successful in achieving this aggregate goal (Wyplosz 2006). The ECB targets a single euro area inflation rate, calculated as annual changes of a weighted average of national consumer prices indices (CPIs). A major concern, however, is that by targeting a euro area aggregate, inflation rates of individual member countries may actually diverge. While identifying national inflation divergence is in principal straightforward, evidence of divergence may be driven by aggregation bias over heterogeneous dynamic cross section (see Pesaran and Smith 1995; Altissimo et al. 2006, 2009; Imbs et al. 2005a, b, 2009).

Economic theory is ambiguous about the impact of monetary union on inflation divergence. The optimum currency area literature would suggest that in the absence of nominal exchange rate flexibility, inflation differentials will take on the role of shock absorbers. Inflation divergence could also originate from the Balassa–Samuelson effect, as countries with accelerating productivity growth may experience higher inflation. Furthermore, monetary union produces lower real interest rates for countries with higher inflation, which generates a tendency to perpetuate inflation differentials.[1] A growing empirical literature examines divergence of inflation in EMU. Recent studies have identified some divergence of aggregate inflation since the beginning of monetary union. Lane (2006) identifies high level of persistence in inflation differentials across EMU members. Busetti et al. (2007), using formal methods, find evidence of recent divergence in euro area inflation based on panel tests for stationarity which are robust to cross-sectional correlation. We seek to extend Busetti et al. (2007) with respect to sectoral inflation data.

Our main concern is whether aggregate inflation data overstates evidence of inflation divergence, as the process of aggregation can affect the dynamic properties of the data and as a result can make common effects more pervasive and less stationary. Empirical work by Imbs et al. (2005a),[2] Altissimo et al. (2006, 2009) support evidence of aggregation bias. Imbs et al. (2005a) find evidence of

[1] However, Sinn and Reutter (2001) do not find strong evidence that difference in productivity explain inflation differentials in the euro area. Honohan and Lane (2003) identify differential exposure to euro exchange rate movements, output gaps and price level convergence as factors behind euro area inflation divergence. Angeloni and Ehrmann (2007) explain inflation differentials on the basis of inflation persistence itself and Canzoneri et al. (2006) suggests that demand shocks explain inflation differentials. Benigno and López-Salido (2006) using an optimizing model illustrate that heterogeneity in inflation persistence, and hence persistence in inflation differentials, matters for the design of monetary policy.

[2] See the further discussion on this paper by Chen and Engel (2005) and Imbs et al. (2005b).

purchasing power parity (PPP) when they utilise disaggregate Eurostat price data. PPP evidence is absent at the aggregate level because of aggregation bias. Impediments to price arbitrage are expected to vary with different goods' characteristics. This gives rise to high relative price persistence at the aggregate level but lower relative price persistence at the disaggregate level. In a paper that summarizes work by the euro area Inflation Persistence Network, Altissimo et al. (2006) investigate sectoral and aggregate inflation dynamics in the euro area. Altissimo et al. (2006) suggest that aggregation bias can be found when examining the level of inflation and report a very high degree of inflation persistence at the aggregate level, but find a considerably reduced level of inflation persistence at sectoral level. The authors attribute this to the influence of transitory sector-specific shocks that are "smoothed" out through aggregation. Given the importance of disaggregation for convergence stories of inflation, it seems surprising that the literature has not yet tested the inflation divergence hypothesis using both aggregate and disaggregate data in EMU. This paper intends to fill this gap.

While the primary aim of this paper is to inform theoretical work on the stylised facts regarding inflation differentials, we nevertheless aim to provide some lessons for the conduct of monetary policy. The current debate on the persistence of inflation differentials in the euro area appears to offer a wide range of policy implications: from no action at all to calls for national accommodating fiscal and structural policies and even recommendations to increase the safety margin around the ECB inflation rate target or to raise the inflation target itself to avoid deflation in some countries, see Sinn and Reutter (2001). Our paper focuses explicitly on the impact of aggregation on the nature of inflation differentials. Evidence of aggregation bias would have wide-ranging policy implications: rather than changing policies, it may be advisable to change the way we construct the aggregate or at least be aware of sectoral performance. In the presence of aggregation bias, the use of a different set of weights (e.g., giving more weight to countries with more persistent dynamics) may improve effectiveness of the ECB in achieving its goal. Issing (2005) discusses if inflation series less responsive to shocks or policy changes should receive a larger weighting in the euro area inflation aggregate, and this might then prove even more to the point. This all highlights the information content of disaggregate data.

Our study tries neither to identify the determinants of euro area inflation differentials nor to uncover to what extent determinants of inflation differentials differ for aggregate and disaggregate data. Our primary concern here is to establish to what extent aggregation explains the observed persistence of euro area inflation differentials. As aggregation bias is associated with the dynamic properties of time series, see, e.g., Pesaran and Smith (1995), conditioning on other explanatory variables will not remove it. Unless we can rule out aggregation bias, the identification of other determinants of inflation differentials (either at aggregate or disaggregate level) seems to us therefore of secondary importance.

We use a novel data set of disaggregate euro area harmonized indices of consumer prices (HICP) to systematically investigate if the level of persistence of inflation differentials is related to the level of aggregation. We test for a unit root in relative inflation in the euro area using the *Panel Analysis of Non-stationarity in*

Idiosyncratic and Common components (PANIC) approach suggested by Bai and Ng (2004) with the number of estimated common factors using Bai and Ng's (2002) information criteria. This approach accounts for cross-sectional correlation and a multiple factor structure, it further provides evidence to what extent idiosyncratic and common factors explain non-stationarity in inflation differentials. Using aggregate inflation data we do not find evidence of pervasive convergence.

However, when we examine disaggregate inflation differences we do find evidence of pervasive convergence. Our results are robust to cross-sectional correlation and structural breaks. A Monte Carlo experiment further confirms that our results are not due to the different cross-sectional dimensions in the panel data sets. The greater evidence of non-stationary aggregate inflation differentials is consistent with theoretical and empirical evidence on aggregation bias of inter alia Pesaran and Smith (1995), Imbs et al. (2005a, b) and Altissimo et al. (2006, 2009). Difficulties do arise in the examination of the euro area given the short time span since its inception. Our panel approach seeks to compensate the paucity of time series data on euro area by fully utilising the cross-sectional dimension. And given that both panels have the same time dimension the different results at the aggregate and disaggregate levels are not dependent upon only 10 years of data on the euro area.

The rest of the paper is laid out as follows: Sect. 2 provides a brief review of relevant literature; Sect. 3 discusses the econometric methods used in this paper; Sect. 4 discusses the aggregate and disaggregate data from Eurostat; Sect. 5 provides results based on our panel data set and a Monte Carlo simulation of the statistical properties of our main estimator; and Sect. 6 concludes and makes policy recommendations.

2 Inflation convergence

There may be reasons to expect divergence in inflation performance across member countries since the beginning of EMU. For example, if inflation outcomes are asymmetric across the euro area, countries with higher (lower) local inflation and a constant ECB nominal interest rate may have lower (higher) local real interest rates. This may affect consumer and investment spending and consequently lead to further divergence in inflation rates. While some inflation differences are to be expected during the early stages of a currency union as the result of convergence process to a common price level, Honohan and Lane (2003) identify circumstances that might give rise to more persistent inflation differentials, which might actually be harmful. If initial inflation differentials fuel a fear of persistent excessive inflation, an exaggerated national policy response could destabilize the rest of the union. Further weaker adjustment mechanisms imply more frequent and more pronounced price misalignments that could induce boom-bust cycles.[3] An analysis of national

[3] Also see Nickell (2006) for a cautionary note. Although productivity differences and hence the Balassa–Samuelson effect could explain these differences, Honohan and Lane (2003) emphasize the response to the exchange rate. There may be reasons to expect divergences to persist to a greater extent in EMU rather than US monetary union. Angeloni and Ehrmann (2007) find an important role for inflation persistence itself.

business cycles under EMU is provided in Honohan and Leddin (2005) and López-Salido et al. (2005).

Papers that have considered national inflation convergence within the European Union (EU) based on panel unit root approaches include Kočenda and Papell (1997) and Holmes (2002).[4] These papers rely on first generation unit root tests and provide evidence of convergence for data up until the 1990s. This is consistent with macroeconomic policy convergence in the run up to EMU. Evidence of pre-EMU inflation convergence is also provided by Busetti et al. (2007) based on a multivariate homogeneous Dickey–Fuller (MHDF) test. Although they identify convergence for a broad sample of countries, it is worth noting that a multivariate version of the Kwiatkowski et al. (1992) unit root test, which has a null hypothesis of stationarity, rejects convergence for the more recent sample period from 1998 to 2004.[5]

From an econometric point of view three issues might bias an empirical analysis of inflation convergence: panel cross-sectional residual correlation, structural breaks and aggregation. Most studies to date deal with these issues insufficiently, which seriously biases results. For one, studies based on first generation unit root tests do not account for cross-sectional residual correlation. This introduces a size distortion (see O'Connell 1998; Breitung and Pesaran 2008) and a tendency to reject the null hypothesis of unit root in favour of convergence.

Structural breaks can bias findings in favour of divergence, as shifts in the mean of inflation in one country can introduce breaks in the overall convergence process. Non-stationarity can result due to breaks of this kind (Perron 1989). Angeloni et al. (2006) examine sectoral inflation rates for six euro area countries both before and after the inception of monetary union. Using individual price records underlying the CPI, they find evidence of a slight increase in inflation persistence. However, in what follows we find evidence that our results are robust to the possibility of structural breaks.

Finally, and most importantly, it is increasingly realized that the process of aggregation affects time series dynamics. Consequently, aggregation by itself can drive the rejection of inflation convergence. Blanchard (1987) emphasizes different speeds of price adjustment at the aggregate and disaggregate level. Aucremanne and Dhyne (2005) and Altissimo et al. (2006, 2009) suggest price setting is heterogeneous across product categories. This heterogeneity is likely to lead to bias in aggregation and in particular to greater evidence of persistence of inflation differentials (see Altissimo et al. 2006, 2009; with respect to inflation and Imbs et al. 2005a, b; Pesaran et al. 2009 with respect to PPP). We adopt a time series approach

[4] Espana et al. (2002) consider the euro area inflation rate disaggregated by sector, although not by country as in this paper, and then test for convergence using non-stationary methods. They do not report evidence of convergence between sectors but disaggregation helps improve inflation forecasts. Our paper has a much broader and deeper data set and additional we utilize recent developments in panel econometrics.

[5] Busetti et al. (2007) also experiment with excluding a constant from their augmented Dickey–Fuller (ADF) regression to improve power. For the second sample period they find evidence that they can reject the null of convergence almost five times as frequently as with a constant. Our panel approach will have improved power over existing univariate tests. Also convergence can be conditional in our set up, when inflation differential can be positive.

to convergence which was first suggested by Bernard and Durlauf (1995) in the context of output convergence. Two countries i and j converge if

$$\lim_{k\to\infty} E\left(\pi_{i,t+k} - \pi_{j,t+k}|I_t\right) = 0. \tag{1}$$

If differences converge upon zero, convergence is absolute. In addition, if inflation differences converge to a constant, convergence is conditional. We allow for the possibility of non-zero differences in inflation on average by including a heterogeneous constant in out panel estimation methods which are discussed in the next section.

3 Econometric issues

In Sect. 2 we identified the importance of aggregation bias and cross-sectional correlation when testing for inflation convergence. Here we review empirical methods for testing non-stationarity in inflation differentials at the aggregate and disaggregate level. First, we introduce panel methods that account for cross-sectional heterogeneity but assume cross-sectional residual independence. We then introduce the Bai and Ng (2004) PANIC methodology which by imposing a factor structure allows us to test for, and take account of, cross-sectional residual dependence.

3.1 Panel ADF tests

To examine whether national inflation differentials ($\tilde{\pi}_{it} = \pi_{it} - \pi_{jt}$) are stationary, and hence converge, we first apply the Im et al. (2003) (IPS) panel version of the augmented Dickey–Fuller (ADF) test. This test is based on the following panel autoregressive model:[6]

$$\tilde{\pi}_{it} = \alpha_i + (\varphi_i + 1)\tilde{\pi}_{it-1} + \varepsilon_{it} \quad i = 1,\ldots,N; \quad t = 1,\ldots,T \tag{2}$$

where α_i is a heterogeneous constant, $(\varphi_i + 1)$ is a heterogeneous autoregressive parameter and ε_{it} is a cross-sectionally independent residual. Equation (2) can be transformed to a simple panel regression as follows:

$$\Delta\tilde{\pi}_{it} = \alpha_i + \varphi_i\tilde{\pi}_{it-1} + \varepsilon_{it} \tag{3}$$

The panel ADF unit root test of Im et al. (2003) has a null hypothesis that all time series processes in Eq. 3 are random walks with drift.

$$H_0 : \varphi_1 = \varphi_2 = \ldots = \varphi_N = \varphi = 0 \tag{4}$$

And a heterogeneous alternative hypothesis of:

$$H_1 : \varphi_1 < 0,\ldots,\varphi_{N_1} < 0, \quad N_1 \le N \tag{5}$$

[6] For expositional purposes we abstract from lagged first-difference terms in Eq. 2, although as is standard we include them in our estimations to deal with residual serial correlation.

The *heterogeneous* alternative hypothesis contrasts with the *homogeneous* alternative in Levin et al. (2002), which assumes that the autoregressive parameter (φ_i) in Eq. 3 is equivalently stationary for all cross-sections and hence is more restrictive. In the context of our paper, accepting the null hypothesis is evidence of pervasive non-convergence of inflation in the euro area.

The panel unit root test statistic for Im et al. (2003) is constructed as follows

$$Z_{IPS} = \frac{\sqrt{N}[\bar{\tau} - E(\tau_i)]}{\sqrt{Var(\tau_i)}} \rightarrow \mathrm{N}(0, 1) \tag{6}$$

Therefore the test statistic has by standard limit theory a standard normal distribution where $\bar{\tau} = 1/N \sum_{i=1}^{N} t_i$and t_i is the test statistic of the individual ADF tests. Im et al. (2003) provide first ($E(\tau_i)$) and second moment ($Var(\tau_i)$) corrections.

3.2 Panel Lagrange multiplier tests

To test for inflation convergence we further use a panel Lagrange multiplier (LM) test based on Schmidt and Phillips (1992) and developed in Im et al. (2005). This test allows for structural breaks in the mean of the processes. It is based on the Im et al. (2003) ADF Eq. 3 and shares the same null and alternative hypotheses. Assuming independent errors with a zero mean and constant variance σ^2, we obtain the pooled log-likelihood function (where SSE_i are the squared standard errors):

$$\ln L = \sum_{i=1}^{N} \left(\frac{-T}{2} \ln 2pi\sigma_i^2 - \frac{1}{2\sigma_i^2} SSE_i \right). \tag{7}$$

The standardised test statistic Γ_{LM} itself is calculated by first averaging individual regression test statistics and then adjusting with simulated first and second moments. Im et al. (2005) provide evidence that the test has good size properties and is more powerful than the panel unit root approach of Im et al. (2003). Lee and Strazicich (2001) suggest that omitting a shift in the mean α_i in Eq. 2 may lead to invalid results following the approach of Perron (1989). Consequently, we incorporate a test statistic which allows for a shift in the mean of the autoregressive process. Here the break date is determined endogenously for each cross-section based on a grid search procedure utilising t-statistics on the autoregressive coefficients. The asymptotic distribution of the LM test statistic is unchanged when break dummies are included, Amsler and Lee (1995). LM panel unit root test statistics (with or without breaks) are distributed as asymptotic standard normal.

3.3 Stationarity as the null hypothesis

We also examine whether the data rejects the null hypothesis of stationarity of inflation differentials based on Hadri (2000). The Hadri test statistic is based upon a panel regression of the random walk plus noise model:

$$\tilde{\pi}_{it} = \mu_i + u_{it} \tag{8}$$

where inflation differentials ($\tilde{\pi}_{it}$) are a function of a constant (μ_i) and a residual error term (u_{it}). The Hadri (2000) test statistic (Z_μ) uses averages of the univariate Kwiatkowski et al. (1992) KPSS test ($\overline{LM}_\mu$) and is distributed as standard normal:

$$Z_\mu = \frac{\sqrt{N}\left(\overline{LM}_\mu\right)}{\varsigma_\mu} \to N(0,1) \tag{9}$$

where N is the number of cross-sections and asymptotic moments are $\xi_\mu = 1/6$ and $\zeta_\mu = \sqrt{1/45}$. An average of individual KPSS tests statistics is constructed as follows:

$$\overline{LM}_\mu = \frac{1}{N}\sum_{i=1}^{N}\left(\frac{\frac{1}{T^2}\sum_{t=1}^{N} S_{it}^2}{\hat{\sigma}_{u,i}^2}\right). \tag{10}$$

where $\hat{\sigma}_{u,i}^2$ is an estimator of the long-run variance of the panel time series under investigation and S_{it} is the partial sum of the estimated residuals u_{it} in Eq. 8. Busetti et al. (2007) suggest that when inflation rates are already somewhat aligned, a test for convergence should rely on a null hypothesis of stationarity. In this context, the null hypothesis of stationarity also tests for non-divergence. Following convention we, however, place more weight on the unit root evidence. It further provides evidence of whether shocks are permanent.

3.4 Panel analysis of non-stationarity in idiosyncratic and common components

The PANIC approach introduced by Bai and Ng (2004) uses a factor structure to understand the nature of non-stationarity in large dimensional panels. Bai and Ng factor model is set out for the case where only an intercept is included:

$$\tilde{\pi}_{it} = c_i + \lambda_i' F_t + e_{it} \tag{11}$$

$$F_t = \alpha F_{t-1} + u_t \tag{12}$$

$$e_{it} = \rho_i e_{it-1} + \varepsilon_{it} \tag{13}$$

The series $\tilde{\pi}_{it}$ is the sum of a cross-section-specific constant (c_i), a common component $\lambda_i{}' F_t$ and an error, e_{it}, which is the idiosyncratic component. The series $\tilde{\pi}_{it}$ is non-stationary if the common factors ($\alpha = 1$) and/or the idiosyncratic component ($\rho_i = 1$) are non-stationary. PANIC allows us to identify whether non-stationarity is pervasive (due to the common factor) or series-specific (due to the individual series). The correct number of factors is determined by the information criteria procedure developed in Bai and Ng (2002). It is recommended that a panel Bayesian information criteria is used to identify the number of factors since it is more robust to cross-sectional correlation in the idiosyncratic errors. Gengenbach et al. (2004) suggest PANIC allows non-stationarity to arise in either the common or idiosyncratic component, whilst Moon and Perron (2004) and Pesaran (2007a) assume common and idiosyncratic non-stationarity under the null hypothesis. It is particularly useful in our context that PANIC determines explicitly whether the non-stationarity in a series is pervasive or variable-specific. PANIC also provides consistent estimates of the space spanned by F_t (denoted $\hat{F}_t$) and the idiosyncratic component (denoted $\hat{e}_{it}$).

We make use of two test statistics from Bai and Ng (2004). Firstly, an ADF test on the common factor ($ADF^c_{\hat{F}}$) and secondly a Fisher-type pooled ADF test on the idiosyncratic individual errors ($ADF^c_{\hat{e}}(i)$). Bai and Ng (2004) Theorem 4 test statistic on the idiosyncratic element is distributed as standard normal as follows:

$$P^c_{\hat{e}} = \frac{-2\sum_{i=1}^{N} \log p(i) - 2N}{\sqrt{4N}} \to N(0,1). \tag{14}$$

$p(i)$ is the p-value associated with ($ADF^c_{\hat{e}}(i)$) of the ADF test for the i cross section, where ρ_i is the autoregressive parameter of the independent error processes in Eq. 13. The test statistic examines whether H_0: $\rho_i = 1$ $\forall i$ against H_0: $\rho_i < 1$ for some i. In addition we report tests with a null of stationarity as suggested by Bai and Ng (2005). Here the null hypothesis is that all cross-sections are stationary and the alternative is that some may be non-stationary.

The PANIC approach has a number of useful facets which encourage us to use it in our empirical study. O'Connell (1998) suggests cross-sectional correlation causes the standard pooled panel tests to over reject the null hypothesis of unit root. However, O'Connell's generalized least squares (GLS) data transformation requires that the common component is stationary. This may not always be the case. The PANIC approach is advantageous since the common factors and idiosyncratic components are consistent irrespective of whether they are stationary or not: the unobserved components are estimated by first-differencing the data and then accumulating the estimates. Additionally, Jang and Shin (2005) provide Monte Carlo evidence that Bai and Ng's (2004) second generation panel unit root test has preferable statistical properties to tests based on principle components such as Moon and Perron (2004) and Phillips and Sul (2003). Due to the nature of subtracting the factor in Bai and Ng (2004) there are more stable sizes under cross-sectional dependency and also OLS estimation (Jang and Shin 2005). Gengenbach et al. (2004) also report evidence of more favourable statistical properties for Bai and Ng (2004) than alternatives including Moon and Perron (2004) and Pesaran (2007a). Additionally Bai and Ng (2004) allows us to impose the correct number of factors based on the Bai and Ng (2002) information criteria rather than arbitrarily imposing a factor structure which might not be supported by the data. This is important in what follows.

4 Data

To construct our inflation measure, and hence to test inflation differentials, we use consumer prices from the harmonized index of consumer prices (HICP) from the Statistical Agency of the European Commission, Eurostat. The coverage of the HICP is defined in terms of household final monetary consumption expenditure and is a Laspeyres chain index. The aim of the HICP was to provide a best measure of

Table 1 Descriptive statistics for the euro area inflation performance

	EZ	GE	FR	BG	SP	IT	NL	GR	IR	PO	AT	FN
1997–2006												
Mean	1.90	1.37	1.61	1.80	2.81	2.23	2.39	3.54	3.07	2.75	1.52	1.50
SD	0.50	0.59	0.66	0.77	0.77	0.40	1.18	0.96	1.31	0.91	0.64	0.89
2000–2006												
Mean	2.17	1.58	1.97	2.10	3.23	2.43	2.64	3.34	3.61	3.07	1.85	1.60
SD	0.28	0.51	0.35	0.71	0.51	0.29	1.34	0.51	1.17	0.90	0.45	1.03

household inflation for international comparisons within the euro area and the EU. Consumer prices are also used by Busetti et al. (2007) and provide a broader measure of the prices consumers actually face, rather that producer price or the gross domestic product deflator. The ECB also targets consumer prices. We have monthly inflation data from 1997M1 to 2006M4 ($T = 112$) and for eleven countries in our sample: Germany, France, Belgium, Spain, Italy, the Netherlands, Greece, Ireland, Portugal, Austria and Finland.[7] The disaggregate HICP data set has eleven sectors. We seasonally adjust data using US Bureau of Census X11. We use aggregate year-on-year HICP inflation rates ($\pi_t = 100 \times \ln[p_t/p_{t-12}]$).[8]

Pesaran (2007b) and Pesaran et al. (2009) show that tests of convergence can be highly sensitive to the choice of the base country. To remove any size distortion due to base country effects, Pesaran (2007b) advocates a pair-wise approach, which we follow. Consequently, to test for aggregate (national) inflation convergence we examine $(\pi_{it} - \pi_{jt})$ $\forall i \neq j$, where π_{it} and π_{jt} are aggregate HICP inflation, $i, j = 1, \ldots, N_1$ and N_1 is the number of countries in our sample. For disaggregate (sectoral) inflation convergence we examine the differential $(\pi_{ikt} - \pi_{jkt})$ $\forall i \neq j$, $\forall k$, $i, j = 1, \ldots, N_1$, $k = 1, \ldots, N_2$, and N_2 is the total number of sectors. It should be noted that although this pair-wise approach accounts for much of the cross-sectional correlation it may not necessarily remove all of it, as other common shocks may prevail.

Table 1 presents descriptive statistics and Fig. 1 graphs the data. Mean inflation between 1997 and 2006 range from 3.54% p.a. in Greece to 1.37% p.a. in Germany. Standard deviations range from 0.40 in Italy to 1.31 in Ireland. Figure 1 shows that

[7] Given our sample period encompasses the fixing of nominal exchange rates and the introduction of notes and coins, this reinforces the testing of break dates. 1997 is a natural place to start our analysis since this was the Maastricht criteria reference year where inflation rates were not to deviate by more than 1.5% of the three best performing countries. Additionally we exclude Luxembourg as is standard in country studies.

[8] Applying annual differences is consistent with the policy rate which the ECB targets (i.e., year-on-year inflation). Monthly changes are not as relevant for monetary policy in the euro area and our results are robust to any induced moving average error since we apply our first-differencing approach equally to both aggregate and sectoral inflation.

Sectors included are: Food and non-alcoholic beverages; Alcoholic beverages, tobacco and narcotics; Clothing and footwear; Housing, water, electricity, gas and other fuels; Furnishings, household equipment and routine maintenance of the house; Health; Transport; Communications; Recreation and culture; Restaurants and hotels; and Miscellaneous goods and services.

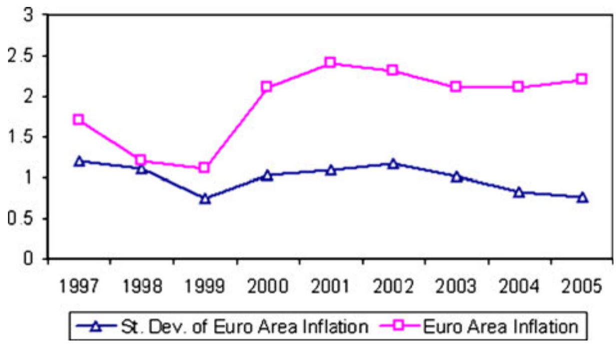

Fig. 1 Euro area inflation

euro area inflation initially increased in 2000 but has since remained close to its 2% target. Inflation dispersion, measured by the standard deviation across countries' inflation rates peaked in 2002 at 1.2 after which it declined. A decline in standard deviation of inflation is associated with increasing sigma convergence.[9]

5 Results

5.1 First generation panel unit root tests

In this section we present evidence on inflation convergence within the EMU based on panel unit root and stationarity tests. These results will serve as an important first benchmark, although their validity is conditional upon cross-sectional residual independence. Evidence is based on the Im et al. (2003) panel unit root test and the Hadri (2000) panel stationarity test. As structural breaks can alter the validity of unit root tests, we also apply the Im et al. (2005) panel LM test, which has a null of a unit root and implicitly accounts for structural shifts. Finally, as an additional sensitivity approach commonly associated with first generation unit root tests, we cross-sectionally demean the data to account for possible common shocks. Table 2 summarizes the results and does not suggest that there are persistent inflation differentials or that 'aggregation bias' is qualitatively important when using a first generation approach.

The Im et al. (2003) Z_{IPS} panel unit root test reject the null hypothesis of a unit root for both aggregate and disaggregate data; evidence of convergence is present for both aggregate and disaggregate data, however, it is considerably stronger for the disaggregate data. The Hadri (2000) Z_{μ} panel stationarity test fails to provide evidence of divergence for both data sets. Results based on the panel LM (Im et al. 2005) also reject the null hypothesis of panel unit roots. As the LM tests accounts for structural shifts and the results are not qualitatively different from the other tests, we take this as evidence that our results are robust to structural breaks. Our results

[9] Lane (2006) suggests dispersion in annual inflation not out of line with US regions. Lane (2006) also suggest that there may be twice as much persistence of inflation differentials in EMU versus the US.

Table 2 First generation panel evidence on inflation differentials

	Z_{IPS}	Z_{μ}	Panel LM with level shift
Aggregate differentials ($\pi_{it} - \pi_{jt}$)	−8.361* (−8.798*)	0.196 (0.192)	−4.119* (−5.321*)
Disaggregate differentials ($\pi_{ikt} - \pi_{jkt}$)	−29.241* (−29.148*)	0.063 (0.062)	−19.957* (−19.640*)

The panel dimension of the aggregate data are ($N = 55$, $T = 112$) and disaggregate data ($N = 605$, $T = 112$). For these tests statistics an asterisk (*) denotes rejection of the null hypothesis. Z_{IPS} is the Im et al. (2003) panel unit root test, see Eq. 6, contains a constant and has a critical value at the 5% significance level of −1.65. The Hadri (2000) Z_{μ} test statistic corresponding to Eq. 9 has a null hypothesis of stationarity is distributed as standard normal. The panel LM from Im et al. (2005) corresponds to Eq. 7 in the test and allows for one break which is endogenously determined by max |td|. Results in parentheses (.) have cross-sectional means removed

appear further robust to removing cross-sectional means. Demeaned results are presented in parentheses in Table 2. Importantly this approach imposes the restrictive assumption that common factors impact homogeneously across cross-sections and this assumption is put to the test in what follows.

5.2 PANIC panel unit root tests

First generation panel unit root tests, as presented in Sect. 5.1, do not correct for cross-section correlation beyond incorporating time dummies. As these force common shocks to have a common impact across cross-sections, these earlier tests potentially suffer a size distortion (O'Connell 1998) and have a tendency to reject the null hypothesis of unit root. This questions the robustness of our previous results. To explicitly take account for potential cross-sectional correlation we test for the number of common factors using the information criteria developed in Bai and Ng (2002). We then test for panel unit root and stationarity using Bai and Ng's (2004, 2005) PANIC approach. As already mentioned, in this model non-stationarity can arise from the idiosyncratic component, the common factor(s), or both. PANIC is therefore an important improvement over our earlier results as well as many other second generation panel unit root tests, e.g., Phillips and Sul (2003), Moon and Perron (2004) and Pesaran (2007a), as both the number of factors as well as the time series properties of the underlying components are data determined and not arbitrarily imposed. To recall, PANIC decomposes times series into, and performs unit root tests on, idiosyncratic and common component(s).

Test results of the time series properties of aggregate and disaggregate inflation differentials are presented in Table 3. For *aggregate inflation data*, PANIC identifies five common factors and further establishes different time series properties for the idiosyncratic and the common components. The idiosyncratic component, which refers to the part of inflation differentials which are country-specific are found to be stationary; the idiosyncratic unit root tests $P^{c}_{\hat{e}}$ suggest that we can reject the null hypothesis of a unit root in national inflation differentials. With respect to what is common to all euro area inflation differentials, however, the Bai and Ng (2002) information criteria identifies five factors, four of which are

Table 3 PANIC evidence on inflation differentials

	$P_{\hat{e}}^{c}$	Number of factors	$ADF_{\hat{F}}^{c}$
Aggregate differentials ($\pi_{it} - \pi_{jt}$)			
Unit root	3.064*	5	−2.550, −2.377, −3.293*, −2.484, −1.165
Stationarity	2.161*	5	0.178, 0.373, 0.391, 0.299, 0.613*
Disaggregate differentials ($\pi_{ikt} - \pi_{jkt}$)			
Unit root	40.120*	0	−3.312*
Stationarity	0.473*	0	0.057

Asterisk (*) denotes rejection of the null hypothesis. In our factor model $ADF_{\hat{F}}^{c}$, the factor unit root test, has a 5% asymptotic critical value of −2.86 (see Bai and Ng 2004, p. 1135). The idiosyncratic unit root test, $P_{\hat{e}}^{c}$ is distributed as standard normal, hence the critical value at the 5% level is 1.64. We utilise Bai and Ng's (2002) third information criteria to determine the number of factors. For Bai and Ng's (2005) stationarity test the critical value on the factor is 0.463, while the critical value on the idiosyncratic is 0.324. Lag lengths are determined by the formula $4[T/100]^{1/4}$ following Bai and Ng (2004). The dimension of the aggregate are ($N = 55$, $T = 112$) and for the disaggregate ($N = 605$, $T = 112$)

non-stationary ($ADF_{\hat{F}}^{c}$ test statistic is less than the 5% critical value). As the overall time series properties of the data are the result of the joint time series properties of its components, we interpret the PANIC results for aggregate inflation data as pervasive non-convergence of aggregate inflation in the euro area.

These results are reinforced by alternative panel tests (Bai and Ng 2005), which suggest that we can reject the null hypothesis of stationarity in both the idiosyncratic and the pervasive component. Our approach of testing both the null of stationarity and unit root is advantageous in this context. As the panel data set rejects both null hypotheses (unit root and stationarity) for the idiosyncratic test, this presents evidence that some of the cross-sections may be stationary while some may be unit root. Nevertheless, the tests applied to the factor provide evidence of the pervasiveness of non-stationarity.[10] It suggests that while there may be some convergence in the idiosyncratic component it is not broad. In general terms our results are consistent with the aggregate evidence in Busetti et al. (2007) suggesting non-convergence of national inflation. Also, evidence of aggregate convergence presented previously based on Im et al. (2003) Z_{IPS} may be caused by cross-sectional correlation and a size distortion in this first generation test.

Results for *disaggregate inflation data* are very different and suggest euro area inflation differentials are transitory. The idiosyncratic component does not appear to have a unit root and the information criteria fails to identify any common factors; all the evidence seems to point to stationarity, or convergence in disaggregate inflation differentials. This result is not entirely unexpected: Altissimo et al. (2006) suggest that aggregation smoothes out idiosyncratic shocks and makes the series more likely to be dominated by a common factor. Our disaggregate results appear to support this

[10] Additionally we examine whether factor non-stationarity is due to structural breaks. Evidence from Saikkonen and Lütkepohl (2002) unit root tests with endogenously determined level shift dates indicates factor non-stationarity was not due to this kind of model misspecification. Given that the idiosyncratic data is stationary, we do not test this component for structural breaks. The literature clearly indicates that non-stationarity may be due to structural breaks, see Perron (1989), but we are unaware of any evidence which suggests time series stationarity can be due to structural breaks.

argument. Given that other approaches proscribe *one* common factor,[11] we estimate this model with one factor imposed for comparison. We can reject the null hypothesis that the idiosyncratic component is unit root, suggesting that there is no pervasive evidence of non-stationarity in our disaggregate panel. Indeed, there is pervasive stationarity. We find that the factor is stationary unlike in the aggregate case. It is also interesting to note that this factor is stationary under both a unit root null and a stationary null, hence there is pervasive stationarity or convergence. However, it should be noted that this model specification incorrectly imposes a factor when, according to Bai and Ng (2002), none exists. This issue is taken up in more detail below.

5.3 Monte Carlo evidence

Our aggregate inflation convergence results are based on approximately 50 cross-sections and the disaggregate study has approximately 600 cross-sections. To exclude the possibility that our results are driven by size distortion due to the different dimensions of our panel data sets, we conduct a small Monte Carlo experiment and examine rejection frequencies of the Bai and Ng (2004) PANIC test statistics, which are based on a null hypothesis of a unit root. For our Monte Carlo study, we generate random data sets and vary the autoregressive parameter of the idiosyncratic error and common factor. The idiosyncratic error is generated as follows $e_{it} = \rho_i e_{it-1} + \varepsilon_{it}$ and the factor is generated by $F_t = \alpha F_{t-1} + u_t$. Random error terms are denoted by ε_{it} and u_{it}, and ρ_i and α are autoregressive parameters associated with the idiosyncratic errors and the factor, respectively. Consequently we have $y_{it} = \lambda F_t + e_{it}$ and we run 5,000 trials for two panel samples on y_{it} (i.e., $T = 100$, $N = 50$, and $T = 100$, $N = 600$). Our sample sizes in the Monte Carlo study approximate the sample sizes in our empirical study of inflation convergence. The rejection rate of the ADF test applied to the one common factor is reported in column $\hat{F}$. Column $P^c_{\hat{e}}$ provides rejection rate for the idiosyncratic component.

Our results in Table 4 confirm the Monte Carlo results in Bai and Ng (2004) but for the two sample periods in our study. The size of both factor and idiosyncratic tests is close to the nominal size and the idiosyncratic test has greater power. We find that the rejection rates for both sample sizes ($N = 50$, 600) are remarkably similar with only an increase in the rejection frequency on the occasion of a particularly small autoregressive parameter for the common factors ($\alpha = 0$). We therefore conclude that it is highly unlikely that size distortions explain different convergence behaviour of inflation at aggregate and disaggregate level.

5.4 Summary of results

Table 5 summarizes our results, contrasting findings for aggregate and disaggregate data for the first generation panel unit root tests and the PANIC second generation test of Bai and Ng (2004). According to first generation panel unit root tests both aggregate and disaggregate inflation differentials are stationary. Unfortunately,

[11] See Phillips and Sul (2003) and Pesaran (2007a).

Table 4 Rejection rates for the null hypothesis of a unit root

ρ_i	α	$T = 100, N = 50$		$T = 100, N = 600$	
		$P_{\hat{e}}^c$	$\hat{F}$	$P_{\hat{e}}^c$	$\hat{F}$
1.00	0.00	0.06	0.77	0.06	0.99
1.00	0.50	0.06	0.81	0.06	0.94
1.00	0.80	0.06	0.52	0.05	0.58
1.00	0.90	0.06	0.24	0.05	0.26
1.00	0.95	0.06	0.12	0.05	0.12
0.00	1.00	1.00	0.07	1.00	0.06
0.50	1.00	1.00	0.06	1.00	0.06
0.80	1.00	1.00	0.07	1.00	0.07
0.90	1.00	1.00	0.06	1.00	0.06
0.95	1.00	1.00	0.06	1.00	0.06
1.00	1.00	0.06	0.06	0.06	0.06

Our model has one factor. The data are generated as $e_{it} = \rho_i e_{it-1} + \varepsilon_{it}$ and $F_t = \alpha F_{t-1} + u_t$. $P_{\hat{e}}^c$ and $\hat{F}$ are the rejection frequencies for the tests applied to the idiosyncratic and pervasive panel unit root tests. We use 5% critical values

Table 5 Results summary

	Aggregate inflation differentials	Disaggregate inflation differentials
First generation panel unit root tests	Stationary	*Stationary*
Second generation panel unit root test		
Idiosyncratic component	*Stationary*	Stationary
Common factor	*Non-stationary*	Stationary

Appropriate model according to Bai and Ng (2002) information criteria is in italics

these tests are biased towards stationarity in the absence of cross-section residual independence (see O'Connell 1998; Breitung and Pesaran 2008), which clearly seems to be present in the aggregate data according to the Bai and Ng (2002) Bayesian information criteria. Inference about convergence at the aggregated level should therefore be based on PANIC, which appropriately accounts for cross-sectional dependence in the aggregate data. PANIC identifies non-stationarity in the common factors, and as such evidence in favour of inflation divergence.[12] At the disaggregate level, the Bai and Ng (2002) information criteria does not identify any common factors, which makes the first generation panel unit root test the appropriate model.[13] Therefore, disaggregate data points clearly to inflation convergence in the euro area as a first generation tests are appropriate in this

[12] It should be noted that evidence of permanent differentials in the aggregate may be an artifact of long memory in the differentials or sample-specific. The former is less likely to be the case with our panel approach since we have substantial power improvements over univariate methods.

[13] The disaggregate results from PANIC confirm the absence of a common factor and even if a common factor is (inappropriately) enforced, this factor is returned as stationary.

case and they suggest convergence. There appears to a striking difference in the dynamics of inflation according to the level of aggregation. While we confirm the Busetti et al. (2007) evidence of divergence in euro area inflation for national data, the failure to establish a similar finding in the disaggregate data, compels us to reconcile the difference as evidence of aggregation bias. As a further robustness check, we standardize the data to account for any cross-sectional heterogeneity in the variance of the inflation differentials. Standardization does not have any major impact on results, if anything, it make aggregate inflation differentials appear even more non-stationary.[14]

6 Conclusion

The monetary policy of the European Central Bank targets aggregate euro area inflation. Concerns are growing that such an aggregate focus may actually lead to a divergence of national inflation rates, which could threaten the long-term viability of monetary union. While different explanations for diverging euro area inflation rates have been brought forward in the literature, to date, the impact of aggregation on inflation convergence has not been studied. This has been the primary objective of this paper. We make use of a novel data set of euro area harmonized consumer price indices and base statistical inference on panel econometric tests that are robust to potential cross-sectional correlation. The appropriate factor structure to capture cross-sectional correlation is identified through the Bai and Ng (2002) information criteria. We use a Monte Carlo study to test for the robustness of our results.

We find a striking difference in inflation dynamics depending on the level of aggregation. While we confirm Busetti et al. (2007)'s findings of divergence in euro area inflation for aggregate data, we fail to establish a similar finding in the disaggregate data. The difference between the two data sets appears to be explained by the underlying factor structure. For aggregate data, inflation divergence is linked to the existence of non-stationary common factors, which are not present in the disaggregate data. Our results indicate that aggregation introduces common components which are not present in disaggregate data; such a contention is consistent with theoretical and empirical evidence of inter alia Pesaran and Smith (1995), Imbs et al. (2005a, b), and Altissimo et al. (2006, 2009). We believe this is a new and important finding. It provides further evidence in favour of the positive effects of EMU (Wyplosz 2006), and it goes some way to discrediting claims that euro area inflation differentials are intrinsically persistent (Trichet 2006). The reason we fail to find evidence of divergence of inflation consistent with a "one size fits all" policy interest rate may be due to the internal real exchange rates of the euro area and the constancy of inflation expectations. Real exchange rate movements will act as a buffer for high inflation countries and consequently offset falling real interest rates. Furthermore, Nickell (2006) shows that Euro inflation expectations are highly stable and indeed smoother than actual inflation; as a consequence, real

[14] Standardization will by definition remove some of the heterogeneity in inflation that we wish to test for. Results based on standardized data are available upon request.

interest rates are quite smooth and unlikely to give rise to persistent inflation differentials.

The current debate on the persistence of inflation differentials in the euro area appears to offer a wide range of policy implications: from no action at all to calls for national accommodating fiscal and structural policies and even recommendations to increase the safety margin around the ECB inflation rate target or to raise the inflation target itself to avoid deflation in some countries, see Sinn and Reutter (2001). Policy prescriptions tend to vary if high levels of persistence in inflation differentials are seen as transitory or permanent. No or little policy intervention is advocated if the persistence is seen as short-lived. Such temporary inflation differentials could arise due to one-off transitory factors during the early stages of EMU related to the convergence process itself and would eventually die down naturally, causing inflation series to eventually converge. However, persistent inflation differentials could also be caused by permanent or protracted differences between national economic structures and policies. In that case, inflation series are likely to diverge as economic shocks would trigger different country responses (Gonzáles-Páramo 2005). Structural differences could be linked to rigidities in domestic factor markets; national consumption preferences; differences in productivity and competitiveness; fiscal policy and external effects (e.g., degrees of oil dependency, extra euro area trade integration, exchange rate path-through patterns). While empirical studies do not identify one single source behind the persistence of inflation differentials (ECB 2003; Honohan and Lane 2003), it is interesting to observe that the impact of aggregation on persistence is generally not investigated. A fully structural model, potentially of the form adopted by Imbs et al. (2009) but applied to inflation differentials, may be useful to buttress any policy prescriptions.

One emerging policy lesson from our research is that: rather than changing policies, it may be advisable to change the way we construct the aggregate. In the presence of aggregation bias, the use of a different inflation measure, possibly based on a different set of weights (e.g., giving more weight to countries with more persistent dynamics) may improve effectiveness of the ECB in achieving its goal. Issing (2005) discusses if inflation series less responsive to shocks or policy changes should receive a larger weighting in the euro area inflation aggregate, and this might then prove even more to the point. Any such decision, however, cannot misread how inflation behaves at different levels of aggregation.

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