Has the Euro Era Facilitated Inflation Convergence?

Mark J. Holmes*

University of Waikato, New Zealand

Abstract This paper investigates the long-run convergence of national inflation rates in the European Union. The application of a new test is proposed that involves unit root testing of the first principal component based on national-EU inflation differentials. Using inflation data based on harmonized consumer price indices for 1999-2007, the first principal component is stationary. This suggests that all EU national inflation rates are driven by a single common stochastic trend. Further analysis suggests that the national degrees of persistence with respect to deviations in inflation from long-run equilibrium are varied.

Keywords: Inflation, convergence, unit roots, cointegration, principal components

JEL Classification: C5, E3, F4

1. Introduction

The behavior and stability of national inflation rates has been of key concern for European Union (EU) policymakers as member countries move towards ever-closer union. The 1992 Maastricht Treaty states that individual members’ inflation rates should not be more than 1.5% higher than the average of the three lowest inflation rates in the European Monetary System. More recently, the European Central Bank (ECB) has argued that price stability is guaranteed if the yearly area-wide aggregate inflation rate (in terms of the harmonized index of consumer prices) is below, but close to, 2% over the medium term. Embodied in these conditions is the desire for inflation convergence among member states. Moreover, an adequate degree of structural similarity in real and nominal economic quantities among countries belonging to an optimum currency area or monetary union is required for political and economic stability [Fratianni and von Hagen (1992), Feldstein (1997), Palomba et al. (2007)].

Early studies by Hall et al. (1992), Koedijk and Kool (1992), Caporale and Pittis (1993), Thom (1995), Holmes (1998) and others provide mixed evidence that the Exchange Rate Mechanism (ERM)- a forerunner of moves towards a single currency- facilitated varying degrees of convergence in EU inflation rates. Furthermore, these studies find that evidence in favor of German leadership determining EU inflation rates is limited. More recent work has suggested that inflation differentials have shown a divergent behavior and heterogeneity in persistence at country and regional level after the birth of the Euro in 1999 [Mentz and Sebastian (2003), Beck et al. (2006), Busetti et al. (2006)].

This paper examines long-run inflation convergence among fifteen EU members over the Euro period from January 1999 onwards. The key contribution offered by this study to the literature is in terms of the econometric methodology that is employed. The tests for inflation convergence are on the basis of whether the first principal component (FPC), based on
national benchmark deviations from the EU average, is stationary or not. If this principal component- which explains the dominant portion of inflation rate divergence from the EU average- is stationary, we can conclude that national inflation rates converge to the EU average following some process of dynamic adjustment to shocks. The adjustment processes towards long-run equilibrium may, however, be country-specific and feature asymmetries in terms of varied speeds of adjustment across countries.

The use of factor structures to test for unit roots and common trends is reflected in a growing literature that includes Snell (1996), Hall et al. (1999), Moon and Perron (2002), Phillips and Sul (2002), Bai (2004) and Bai and Ng (2004), Holmes and Grimes (2005) and others. On the basis of this literature, one can argue that dynamic factor models are useful in several areas of economic analysis. One relates to index modeling where factors are regarded as unobservable economic indices that capture the co-movement of many variables. Related to this use, a major source of cross-section correlation in macroeconomic data is common shocks. This use is directly relevant to a focus on whether national inflation rates react to the same degree and with the same timing to common shocks that impact across countries. Dynamic factor models are capable of modeling cross-section correlations allowing for heterogeneous responses to common shocks through heterogeneous factor loadings. In a further related application, factor models can be used to study cross-section cointegration in nonstationary panel data. If time series share a common trend, this implies strong cross-section correlation. The standard assumption of cross-section independence is violated, rendering many existing panel unit root tests invalid. Rather than treat cross-section correlation as a nuisance in econometric applications, these co-movements can be exploited to produce new univariate test statistics.

The stationarity or non-stationarity of the principal components yields important economic insights that are directly related to long-run inflation convergence. Consider the case where national inflation rates are each non-stationary. One can form the inflation differentials of each national inflation rate index less the EU average, and then obtain the principal components of the set of these national inflation differentials. The FPC encapsulates the greatest portion of orthogonalised variances in these data. If there are one or more sources of non-stationary shocks, the FPC will reflect these shocks since the variance of a non-stationary series is unbounded whereas the variance of a stationary series is finite. If the FPC of the inflation differentials is non-stationary, we can conclude that non-stationary shocks are present that cause divergence in at least some of the underlying national inflation rates relative to the EU average. Conversely, if the FPC is stationary, we can conclude that no non-stationary shocks exist that cause national divergence. In this latter case, the vector of inflation differentials must converge to a constant vector, consistent with long-run inflation convergence.

Unit root testing of the FPC based on inflation differentials offers a number of key advantages over existing tests for convergence. In investigating the number of common shared trends, the advantage of examining the stationarity of the FPC is that, unlike the Johansen (1988) maximum likelihood procedure and the Stock and Watson (1988) common trend framework, it does not require the estimation of a complete vector autoregression system (VAR). The size and power of this test is not affected by the VAR being constrained to an unreasonably low order on account of data limitations. This method also avoids the need for an entire sequence of tests for the stationarity of a multivariate system. As indicated by Snell (1996), even if each test in the sequence had a reasonable chance of rejecting the false null, the procedure as a whole is likely to have low power. Snell (1999) presents a small
Monte Carlo study where the size and power of the test based on principal components are computed. Experimentation is based on a variety of data generation processes and the test based on principal components is confirmed as having acceptable size and reasonable power compared to the Johansen likelihood ratio cointegration test. In cases of marginal cointegration (i.e. when the cointegrating combination is only borderline stationary) and in small sample sizes, the Johansen (1988) likelihood ratio test has little ability to discriminate between no cointegration and cointegration whereas estimation based on principal components does have discriminatory power in such circumstances.

The paper is organized as follows. The following section discusses the data and econometric methodology. The third section reports and analyses the results. There is evidence that the FPC based on national inflation deviations from the EU average is indeed stationary. This section also offers some results concerning the variability of national adjustments towards long-run equilibrium. The final section concludes.

2. Methodology

Principal components represent transformations of the original dataset. If the original dataset comprises $n$ series, the first principal component will be a linear combination of these $n$ series combined in such a way that it summarizes the maximum possible variation in the original series. The second principal component then summarizes the maximum variation in the unexplained portion of the original series after extracting the influence of the first principal component, and so on up to the $n$’th principal component. This study concentrates on the first principal component of the deviations of national inflation rates from the EU average.

This study employs a two-stage testing procedure for inflation convergence. Stage One involves computing $n$ principal components using the $n$ inflation differentials. If the FPC is stationary, then all remaining principal components will also be stationary thereby confirming strong convergence among all the $n$ countries. If the FPC based on the $n$ inflation differentials is non-stationary, strong convergence among all inflation rates is rejected. Stage Two involves the case where the FPC is confirmed as stationary. The presence of a single common shared trend driving all inflation rates implies that all bivariate national pairings are themselves stationary. This study tests the hypothesis that shocks to inflation originating in Germany spread across the EU at different speeds. Using a seemingly unrelated regression (SUR) estimator, an error correction model based on inflation differentials between each country and Germany is analyzed. This enables the degree of persistence in inflation differentials to be measured.

More formally, suppose the $n$ countries constitute the sample of a given group. The benchmark deviations are defined as

$$(r_i - r^*)_t = u_{it}$$  \hspace{1cm} (1)$$

where $r_i$ denotes the inflation in country $i$ for $i = 1, 2, ..., n$ countries and $r^*$ denotes the EU or base inflation. Let $X_i$ be an $(n \times 1)$ vector of random variables, namely $u_{it}$ for each of the $n$ inflation rates, which may be integrated up to order one. The principal components technique addresses the question of how much interdependence there is in the $n$ variables contained in $X_i$. We can construct $n$ linearly independent principal components which
collectively explain all of the variation in \( X_t \) where each component is itself a linear combination of the \( u_t \)'s. Since I(1) variables have infinite variances, whereas stationary, I(0), variables have constant variances, it follows that the FPC, which explains the largest share of the variation in \( X_t \), is the most likely to be I(1) and so corresponds to the notion of a common trend [Stock and Watson (1988)]. However, if the FPC is I(0) then all the remaining principal components will also be stationary and there are no common trends which suggests that the \( u_t \)'s contained in \( X_t \) are themselves stationary. This will confirm strong convergence with respect to the base across the sample of \( n \) benchmark deviations. If long-run convergence holds for each country with respect to the EU base then it must the case the convergence holds between all national pairs.

We may now consider stage one of the FPC methodology in relation to the identification of common trends. Following Stock and Watson (1988) we can argue that each element of \( X_t \) may be written as a linear combination of \( k \leq n \) independent common trends which are I(1), and \((n-k)\) stationary components which correspond to the set of \((n-k)\) cointegrating vectors among the \( u_t \)'s. Let \( \tau_t \) denote a \((k \times 1)\) vector of common trends and \( \xi_t \) denote an \((n-k) \times 1 \) vector of stationary components,

\[
\tau_t = \alpha X_t \\
\xi_t = \alpha'_\perp X_t
\]

where \( \alpha \) is an \((n-k)\) matrix of full column rank, \( \alpha_\perp \) is an \( nx(n-k) \) matrix that forms the \((n-k)\) cointegrating vectors, \( \alpha'\alpha = I \) and \( \alpha'_\perp \alpha_\perp = 0 \). If there are \( k \) common trends, it can be shown that the \( k \) largest principal components of \( X_t \) may be written as a \((k \times 1)\) vector \( \tau_t^* \) as follows

\[
\tau_t^* = X_t^* \alpha^*
\]

where \( X_t^* \) is an \((n \times 1)\) vector of observations on the \( u_t \) in mean deviation form, \( \alpha^* \) represents the \( k \) eigenvectors corresponding to the largest eigenvalues of \( X_t \) and is defined as \( \alpha R \) where \( R \) is an arbitrary, orthogonal \((k \times k)\) matrix of full rank. This relationship guarantees that under the null hypothesis of \( k \) common trends, each of the \( k \) principal components will be I(1). Similarly, for the \((n-k)\) remaining principal components, it can be shown that these smaller principal components may be written as a \((n-k) \times 1 \) vector \( \xi_t^* \) as follows

\[
\xi_t^* = X_t^* \alpha^*_\perp
\]

where \( \alpha^*_\perp \) corresponds to the \((n-k)\) eigenvectors that provide the \((n-k)\) smallest principal components and is defined as \( \alpha_\perp S \) where \( S \) is an arbitrary orthogonal \((n-k) \times (n-k)\) matrix.
The FPC will be I(1) provided there is at least one common trend among the $u_t$’s contained in $X_t$. We can therefore test the null hypothesis that the FPC is non-stationary against the alternative hypothesis that the FPC is I(0). Rejection of the null means that all principal components are stationary and so there are no common trends among the $u_t$’s contained in $X_t$. This confirms strong convergence with respect to the EU base across all countries. To test the stationarity of the FPC, this study employs the familiar ADF unit root test as well as the GLS-based univariate unit root test advocated by Elliott et al. (1996) that offers higher power and less size distortion. In addition to this, panel data unit root tests advocated by Levin et al. (2002) and Im et al. (2003) that test the null hypothesis of joint non-stationarity across all components are also utilized. One might argue for simply applying these panel data unit root tests to panels comprising inflation differentials rather than principal components. However, the advantage of the procedure followed here is that the components are orthogonal to each other and therefore the validity of the panel tests that employed are not undermined by the presence of cross-sectional dependency.

The second stage of the testing procedure involves computing the half lives of deviations from long-run convergence where stage one confirmed strong convergence. Using SUR, the following equation is estimated for each country

$$
\Delta r_t = \zeta_t + \rho_t (r_t^* - r^*)_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta r_{t-j} + v_t
$$

where $r^*$ denotes German (as opposed to average EU) inflation and $\rho_t$ indicates the speed of adjustment toward long-run equilibrium. The justification for choosing Germany as $r^*$ is that interest in the German dominance hypothesis that has been debated elsewhere in the literature. Since cointegration is a transitive concept, stationarity of all inflation differentials with respect to the EU average implies stationarity of all differentials with respect to Germany. The half life of a deviation from long-run equilibrium may be approximated as $\ln(0.5)/\ln(1+\hat{\rho}_t)$. The presence of differing speeds of adjustment for national inflation rates will be reflected in national variations in $\hat{\rho}_t$.

Before proceeding to the results discussion, it is important to highlight some caveats associated with this methodology. The advantages over existing methods of testing for long-run inflation convergence have been discussed above. A standard criticism of principal component estimation, and indeed of common stochastic trends, is that they are linear combinations of economic variables and so the economic interpretation of a given component can be problematic. Since this study is explicitly not concerned with isolating the determinants of inflation rates themselves, and instead using the methodology only to test for stationarity, this criticism is not directly relevant. Also, testing the null of non-stationarity of the first FPC leaves one vulnerable to the standard criticisms concerning the low power attached to unit root tests making it difficult to reject the null of non-stationarity. The low power of the tests is likely to bias against finding long-run convergence of inflation rates. As shown subsequently in our results, this criticism actually strengthens our findings since non-stationarity of the FPC is rejected.
3. Data and Results

The data examined are monthly observations on inflation for fifteen EU countries. The inflation data are computed from seasonally-adjusted harmonized consumer price indices. The countries included are Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden and the UK. The sample comprises existing EU members from 1995 or earlier and includes twelve countries now participating in the single currency. The study period covers January 1999 to October 2007 where the start date coincides with Euro members irrevocably fixing their nominal exchange rates for a period of three years before the introduction of the Euro was completed in January 2002. Figure 1 presents the fifteen annual inflation rates expressed as differences with respect to the EU average.

Table 1 reports unit root tests on inflation differentials with respect to the EU average. With the exception of Greece, the ADF unit root tests cannot reject non-stationary at the 10% significance level. The application of Elliott et al. (1996) DF-GLS unit root tests reject non-stationarity of inflation differentials involving Belgium, Denmark, France, Greece, Ireland, Italy and Spain, but rejection at the 5% significance level is only present in the cases of Greece and Spain. Despite the application of unit root tests that are relatively more powerful than ADF unit root testing, the initial analysis points to a lack of convergence with the possibility of multiple stochastic trends driving EU inflation rates.

Alternatively, inflation convergence may be examined in a multivariate setting using the familiar Johansen (1988) cointegration testing procedure which is more powerful than the regression-based tests. Caporale and Pittis (1993) and Thom (1995) use this approach and find evidence of multiple stochastic trends driving EU inflation rates over study periods covering the early ERM years. An earlier application of the Johansen procedure to the datasets employed in our study also indicated the presence of multiple stochastic trends driving the inflation differentials.

The central theme of this paper is that the univariate unit root and multivariate cointegration tests suffer from low test power making rejection of the null of non-stationary or non-cointegration difficult. To address this issue, the previously described FPC methodology is applied to the inflation differentials. Table 2 reports that application of the DF-GLS unit root test enables us to reject the null hypothesis of non-stationarity of the FPC. Since the FPC explains the largest variation in the behavior of inflation differentials, it is the principal component that is most likely to be non-stationary. However, since the FPC is in fact stationary at the 10% significance level, it follows that all other principal components will also be stationary. This implies that all inflation differentials are stationary and all bivariate national pairs share the same common stochastic trend with a long-run coefficient of unity.

An issue with univariate unit root testing may be low test power making it difficult to reject a non-stationary null hypothesis. Further support to our findings is offered through the application of the Levin et al. (2002) and Im et al. (2003) panel data unit root tests reported in Table 3. These tests are applied to panels comprising all principal components. An advantage of working with panels comprising the set of principal components, as opposed to direct inflation differentials, is that the avoidance of cross-sectional dependencies that could otherwise undermine asymptotic normality and lead to size distortion and over-rejection of the null hypothesis of joint non-stationarity. Both tests reject joint non-stationarity at the 1% significance level. In addition, the null of joint stationarity is tested using the procedure
advocated by Hadri (2000). With a test statistic of 0.806 and associated $p$-values of 0.210, the null is not rejected. Thus all our tests based on principal components are consistent with strong convergence of inflation rates across the EU.

The long-run convergence of EU inflation is not sufficient grounds for the ECB to conduct a common monetary policy. So long as national short-run inflation dynamics remain dissimilar, the appropriate stance of monetary policy can differ across Euro members [Palomba et al. (2007)]. Table 4 reports the results from SUR estimation of the error correction models expressed in equation (6). With the exception of Denmark and the UK who have not participated in the Euro, $\rho_i$ is significantly different from zero at the 10% level or better. For the twelve cases where $\rho_i < 0$, there is evidence that the speeds of adjustment towards long-run inflation equilibrium are varied with estimated half-lives ranging from 2.4 quarters in the case of Austria to 23 quarters in the case of the Netherlands. In controlling a given increase in inflation, relatively slower speeds of adjustment might suggest that the ECB would need to conduct a sharper increase in interest rates than would be the case for countries such as Austria and Belgium.

4. Summary and Conclusion

This study has approached the debate concerning inflation convergence in the EU from a new perspective based on the application of principal components analysis and unit root testing. In contrast to much of the existing literature that employs more familiar unit root and cointegration testing procedures, there is evidence in favor of long-run equilibrium relationships across long-standing member countries based on coefficients of unity within a multivariate setting. Our results suggest that there is constancy in the long-run inflation ratios between all countries. This is an important conclusion in the context of assessing patterns of long run national adjustment to inflation shocks across the EU. Further estimation indicates evidence of considerable variation in the national speeds of adjustment towards long-run equilibrium across the EU. This latter finding suggests that the appropriate stance of monetary policy can differ across Euro members.

Endnotes

* Professor of Economics, Department of Economics, University of Waikato, New Zealand. Email: holmesmj@waikato.ac.nz. Tel. 64-7-838-4454. Fax 64-7-838-4331.

1. See, for example, Child (1970).

2. Currently, Denmark, Sweden and the UK are not members of the Eurozone.

References


Figure 1A. Inflation Levels
Figure 1B. Inflation Differentials (with respect to the EU average)

Notes for Figures 1A and 1B. The notation is as follows: Austria (OE), Belgium (BG), Denmark (DK), Finland (FN), France (FR), Germany (BD), Greece (GR), Ireland (IR), Italy (IT), Luxembourg (LX), Netherlands (NL), Portugal (PT), Spain (ES), Sweden (SD) and United Kingdom (UK). XXEUAVE denotes the inflation differential between country XX and the EU average (EUAVE).
Table 1. Univariate Unit Root of Inflation Differentials

<table>
<thead>
<tr>
<th>Country</th>
<th>ADF (no trend)</th>
<th>DF-GLS (no trend)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Austria</td>
<td>-1.625</td>
<td>-1.136</td>
</tr>
<tr>
<td>Belgium</td>
<td>-2.035</td>
<td>-1.880*</td>
</tr>
<tr>
<td>Denmark</td>
<td>-1.980</td>
<td>-1.730*</td>
</tr>
<tr>
<td>Finland</td>
<td>-1.520</td>
<td>-1.534</td>
</tr>
<tr>
<td>France</td>
<td>-2.039</td>
<td>-1.727*</td>
</tr>
<tr>
<td>Germany</td>
<td>-1.137</td>
<td>-1.065</td>
</tr>
<tr>
<td>Greece</td>
<td>-2.966**</td>
<td>-1.948**</td>
</tr>
<tr>
<td>Ireland</td>
<td>-1.651</td>
<td>-1.642*</td>
</tr>
<tr>
<td>Italy</td>
<td>-2.250</td>
<td>-1.886*</td>
</tr>
<tr>
<td>Luxembourg</td>
<td>-1.909</td>
<td>-1.167</td>
</tr>
<tr>
<td>Netherlands</td>
<td>-1.412</td>
<td>-1.076</td>
</tr>
<tr>
<td>Portugal</td>
<td>-1.824</td>
<td>-1.303</td>
</tr>
<tr>
<td>Spain</td>
<td>-1.649</td>
<td>-1.953**</td>
</tr>
<tr>
<td>Sweden</td>
<td>-1.750</td>
<td>-1.271</td>
</tr>
<tr>
<td>UK</td>
<td>-1.743</td>
<td>-0.920</td>
</tr>
</tbody>
</table>

Notes for Table 1. These are ADF and DF-GLS [Elliott et al. (1996)] unit roots tests on the national inflation rate minus the EU average. ** and * denote rejection of the non-stationary null hypothesis at the 5 and 10% significance levels respectively with critical values of -2.893 and -2.584 (ADF); 1.944 and -1.615 (DF-GLS).
Table 2. Analysis of the FPC

<table>
<thead>
<tr>
<th></th>
<th>ADF (no trend)</th>
<th>DF-GLS (no trend)</th>
</tr>
</thead>
<tbody>
<tr>
<td>FPC</td>
<td>-1.841</td>
<td>-1.879*</td>
</tr>
</tbody>
</table>

See notes for Table 1.
Table 3. Panel Data Unit Root Tests

<table>
<thead>
<tr>
<th></th>
<th>Statistic</th>
</tr>
</thead>
<tbody>
<tr>
<td>Levin et al. (2002)</td>
<td>-5.604***</td>
</tr>
<tr>
<td>Im et al. (2003)</td>
<td>-12.251***</td>
</tr>
<tr>
<td>Hadri (2000)</td>
<td>0.806</td>
</tr>
</tbody>
</table>

Notes for Table 3. Levin et al. (2002) and Im et al. (2003) test the null hypothesis of joint non-stationarity for all principal components. Levin et al. assume a common unit root process whereas Im et al. assume individual unit root processes. These test statistics are asymptotically normal with a 1% critical value of -2.33. Hadri (2000) assumes a common unit process and tests the null of joint stationarity. This test is also asymptotically normal. *** denotes rejection of the null hypothesis at the 1% significance level.
Table 4. Short-run Adjustment of Inflation rates from Germany

<table>
<thead>
<tr>
<th>Country</th>
<th>$\hat{\rho}_j$</th>
<th>Half-life (quarters)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Austria</td>
<td>-0.271***</td>
<td>2.442</td>
</tr>
<tr>
<td>Belgium</td>
<td>-0.137**</td>
<td>4.696</td>
</tr>
<tr>
<td>Denmark</td>
<td>-0.025</td>
<td>27.263</td>
</tr>
<tr>
<td>Finland</td>
<td>-0.052***</td>
<td>12.943</td>
</tr>
<tr>
<td>France</td>
<td>-0.077***</td>
<td>8.605</td>
</tr>
<tr>
<td>Greece</td>
<td>-0.153***</td>
<td>4.173</td>
</tr>
<tr>
<td>Ireland</td>
<td>-0.078***</td>
<td>8.537</td>
</tr>
<tr>
<td>Italy</td>
<td>-0.058*</td>
<td>11.692</td>
</tr>
<tr>
<td>Luxembourg</td>
<td>-0.059*</td>
<td>11.397</td>
</tr>
<tr>
<td>Netherlands</td>
<td>-0.030**</td>
<td>23.092</td>
</tr>
<tr>
<td>Portugal</td>
<td>-0.072***</td>
<td>9.238</td>
</tr>
<tr>
<td>Spain</td>
<td>-0.124***</td>
<td>5.245</td>
</tr>
<tr>
<td>Sweden</td>
<td>-0.085***</td>
<td>7.820</td>
</tr>
<tr>
<td>UK</td>
<td>-0.018</td>
<td>38.279</td>
</tr>
</tbody>
</table>

Notes for Table 4. All estimates obtained through the non-linear SUR estimation of equation (6). ***, ** and * denotes significance at the 1, 5 and 10% levels based on robust heteroscedasticity-autocorrelation consistent (HAC) standard errors. LR tests indicated that a lag length of $j = 1$ is appropriate.