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# Unit labor cost growth differentials in the Euro area, Germany, and the US: lessons from PANIC and cluster analysis

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# Unit labor cost growth differentials in the Euro area, Germany, and the US: lessons from PANIC and cluster analysis[§]

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**Abstract**

Inflation differentials in the Euro area are mainly due to a sustained divergence of wage developments across the Euro area, and narrower differences in labour productivity growth (Alvarez et al., 2006). We investigate convergence of inflation using unit labour cost (ULC) growth and applying PANIC (Bai and Ng, 2004) and cluster procedures (Hobijn and Franses, 2000, Busetti et al., 2006) to Euro area countries as well as US States, US Census Regions and German *Länder*. Euro area differs in that dispersion in general (and its fraction due to idiosyncratic factors in specific) is larger and common factors are much less important in explaining the variance of ULC growth. We report evidence for convergence clusters in all countries.

**Keywords**: Unit labor costs, inflation, European Monetary Union, Germany, United States of America, convergence, convergence clubs, panel unit root tests, PANIC
**JEL classification**: E31, O47, C32, C33

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# 1 Introduction

Inflation differentials in the Euro area are a matter of concern for central bankers (European Central Bank, 2005, Trichet, 2006, Gonzalez-Paramo, 2005, and Issing, 2005) as well as researchers (Alvarez et al., 2006, Angeloni et al., 2006, Cecchetti and Debelle, 2006). It is argued, that the observation of a continuing divergence in price dynamics might be due to structural rigidities or differences in labour or product market structures, which in turn reduce the speed of the adjustment process (Angeloni and Ehrmann, 2004, Benigno and Lopez-Salido, 2002, Campolmi and Faia, 2006) or due to an inappropriate wage-setting process (Fritsche et al., 2005). It has furthermore been argued that the cross-section dispersion in inflation and output growth rates are connected (Lane, 2006). This in turn might lead to dangerous imbalances in EMU if amplified (Belke and Gros, 2006).

The research results of the "ECB Inflation Persistence Network" indicate, that the most important source of inflation differentials across EMU can be found in internal factors, namely a sustained differential in wage growth and narrower differences in productivity growth.[1] The whole argument became a matter of even greater concern because some of the large countries in the Euro area – Germany and Spain – showed remarkable deviations from the average EMU inflation rate for a number of consecutive years after the introduction of the Euro.

To illustrate the relevance of the point, we first calculated measures of variability in unit labor costs and unit labor cost growth. Denoting $\pi_j$ the respective unit labor cost (ULC) or ULC growth in country/ region $j$ ($j = 1, \ldots, N$), $\pi^*$ the currency area average of the respective measure and $\varpi_j$ a certain weight of country/ region $j$, then the root of the somehow weighted squared distance of $j$ is given by:

$$s_j = \sqrt{\varpi_j \left(\pi_j - \pi^*\right)^2} \qquad (1)$$

We calculated unweighted ($\varpi_j = 1/N$) as well as real GDP weighted ($\varpi_j = GDP_j/GDP^*$) variability measures for levels and growth rates. The respective cross-section distributions of $s_j$ for each year are plotted using Box-Plots

[1] This, in turn became an argument in statements of ECB officials quite recently: "[I]n most countries [of the Euro area], domestic factors dominate external factors in generating inflation differentials. In particular, we have witnessed a sustained divergence of wage developments across the euro area, and narrower differences in labour productivity growth. As a result, differentials in the growth of unit labour costs have been persistent."(Trichet, 2006

[2] for the Euro area (EU 12), Germany, the United States (States and Census regions). [3]

First, we look at the cross-section dispersion in ULC growth rates:

*Insert figures 1 and 2 about here.*

As can be seen in figures 1 and 2, the dispersion was high after the EMS crisis in the early 1990s and diminished in the process of accession to EMU. If we look at the evidence in other currency areas, we notice a historically higher dispersion in ULC growth in Europe than in Germany or the US – which remains slightly higher also after the introduction of a common currency.

Second, we look at variability measure in ULC levels. The result is somewhat different:

*Insert figures 3 and 4 about here.*

For the EU 12 countries after 1999, level dispersion increased and is nowadays as high as it was in the early 1990s – at least when considering the weighted measure. This is mainly due to the relative large weights of deviations in ULC from average in relatively large countries like Spain or Germany. The respective dispersion data for the panel of Western German *Länder* show a remarkable increase in both – unweighted and weighted – dispersion measures since the re-unification boom, however dispersion is not as high as in the EMU. For the United States – irrespective if we have a look at the states level or at a higher level of aggregation (e.g. Census regions) and irrespective from the weigting scheme – the exercise reveals a very stable dispersion of unit labor cost growth over time.

Last, we compare the calculated measure for ULC growth and inflation to illustrate that there is a relationship between both measures:

*Insert figure 5 about here.*

The figure reveals, that both – inflation and ULC growth – dispersion measures show peaks around the same periods. Remarkable increases can be

[2]The median is plotted by a line in the center of a box together with shaded areas denoting a significance area, a box denoting the borders to the first and third quartile, and a whisker denoting the inner fences (1.5 times the interquartile range). Data points with a circle denote near outliers, stars indicate a far outlier.

[3]See section 3 for details.

associated with the two oil price shocks of the 1970s/1980s and the EMS crisis in the early 1990s (shaded areas).

Analyzing the topics of diverging unit labor costs dynamics in a common currency area needs some theoretical underpinnings. It seems to be appropriate to refer to the concepts of $\beta$- and $\sigma$-convergence, when analyzing the topic. According to Barro and Sala-i-Martin (1991), $\beta$-convergence is present if different cross-sectional time series show a mean reverting behavior to a common level. In contrast, $\sigma$-convergence measures the reduction of the overall dispersion of the time series. As pointed out by Quah (1993), the absence of $\sigma$-convergence – as we see from figures 3 and 4 can not be taken as indicating the absence of $\beta$-convergence. In the following, we will mainly base our arguments on $\beta$-convergence. A further distinction, which has to be considered in the context of convergence analysis refers to the distinction between *absolute* and *relative* convergence (Bernard and Durlauf, 1996). Absolute convergence implies, that the ULC growth rates converge towards the same rate, whereas relative convergence means that the relative distance between the growth rates is stationary. This distinction has important implications when applied to inflation rates: relative convergence within a currency union implies, that the competitive position of each country/ region deteriorates on average with a stable rate, whereas absolute convergence implies a stabilization of the competitive position at a given point. Such a situation might be indicated by comparing the dispersion in levels and growth rates for the EMU.

In general, there are several ways to test for $\beta$-convergence empirically. One line of research aims to estimate the average growth rate as a function of the deviation from equilibrium at a given starting point (Beck et al., 2006). A second line of research analyzes common trends between inflation (or – in our case – ULC growth rates) *in levels* within a cointegration framework as e.g in Mentz and Sebastian (2003). A third line of research is based on the analysis of the stationarity properties of inflation differentials (Beck et al., 2006, Busetti et al., 2006).

As preliminary unit root tests – see the results in section 2.2 – indicate, the ULC level data can best be described as I(2) processes. However to avoid a more complicated I(2) analysis and to deliver comparable results with the exiting literature on inflation differentials, we decided to analyze the ULC data in growth rates.

To get a better understanding of the sources of underlying divergence, we add therefore to the existing literature on inflation divergence in the following way:

- First, we employ the PANIC approach as developed in Bai and Ng (2004) and used to analyze comovement and heterogeneity in Euro area data by Eickmeier (2006). The question of interest refers to the distinction of common and idiosyncratic factors driving the variance in panels of ULC growth dynamics. If the non-stationarity in the panel is mainly due to common factors but not to the idiosyncratic components, we cannot reject the hypothesis of convergence.

- As a second approach, we analyze the case for convergence clusters (clubs), using the procedure of Hobijn and Franses (2000) based on all possible bivariate ULC growth rate differentials.

- Third, we aim to make sense out of the factor analysis by means of an SVAR approach using long-run restrictions (Blanchard and Quah, 1989).

- In general, we compare the the evidence for the EMU countries under all methods with the evidence for the States and census regions of the United States of America as well as the German *Länder*.

Our results point to the clearly identifiable existence of one non-stationary and one stationary common factor for Germany and the United States. This is in line with the hypothesis of $\beta$-convergence around common factors (which could e.g. be associated by country-wide factors like supply or demand shocks like oil price hikes or monetary policy actions) for those countries. For the Euro area, however, it is quite difficult to identify common factors – idiosyncratic factors dominate in explaining the bulk of variance – and clustering seems to be present. The idiosyncratic components in all currency areas are found to be stationary – however the respective persistence properties are quite different. Wheras in the case of Germany and the US, idiosyncratic components show a white noise behaviour, in the case of EU 12 we find strong serial correlation.

The paper is organized as follows. Having introduced the topic in section 1, in section 2 the applied methods are explained and the results presented. Section 3 discusses and summarizes the findings.

# 2 Empirical analysis

## 2.1 Data

The data refer to nominal unit labor costs, defined as the ratio of a nominal compensation of employees numbers to the respective real gross domestic - or gross state - product numbers.[4] All data are annual data – however the available time span differs a lot. The longest available data set covers the EMU countries. The data (1960 to 2007 as we included the commissions forecast as two extra data points) are directly available from the AMECO data base of the EU commission.[5]

For Germany, the numbers were calculated using the data from the website of the Länder's network for economic statistics ("Arbeitskreis VGR der Länder").[6] Unit labour costs have been computed by dividing the (nominal) compensation for employees by the real gross regional product for each of the 11 Länder. The SNA classification was changed quite recently in Germany and the backward calculated numbers cover the time span from 1970 to 2004 only. As the data for the old federal republic is only available until 1990, and from 1991 only data for all of Germany is provided, the pan-German unit labour cost index is calculated from the old Länder data until 1990 and from pan-German data from 1991 onwards.

For the United States, the necessary data on gross state products and total compensation of employees has been taken from the Bureau of Economic Analysis' database on regional and state GSP.[7] The change from the SIC industrial classification to the NAICS classification in 1997 has created however a slight problem: As data on employees' compensations has not been published for the first years after the statistical change and have only been resumed in 2001, the time series can only be constructed from 1977 to 1997.

Last but not least, to calculate the inflation rate variability for EMU (see figure 5), we used the private consumption deflator as available in AMECO).

[4]We thank Sebastian Dullien for making his data available.
[5]Please follow the link.
[6]Please follow the link.
[7]Please follow the link.

## 2.2 Determining the order of integration

Before starting with our convergence investigation, we conducted an analysis of the stationarity properties of the time series under investigation. We considered the following tests:[8]

- Tests based on a common unit root process: here the methods of Levin et al. (2002) and Breitung (2000) were considered.
- Tests based on individual unit roots: here an augmented Dickey and Fuller (1979) test and an Phillips and Perron (1988) test in panel versions as proposed by Maddala and Wu (1999) and Choi (2001) were considered.
- All the tests mentioned before are based on the null of a unit root. However we furthermore considered the test described by Hadri (2000), which is based on the null of no unit root.

Table 1 summarize the results.

*Insert table 1 about here.*

Considering the contradictory results when comparing the tests with opposing null hypotheses, the overall evidence can be interpreted as in favour of a level of integration higher than 1 for nominal unit labor costs. This is not surprising, since several studies found I(2) properties for nominal variables (Juselius, 1999). This calls for an I(2) analysis of ULC *level* convergence – which we leave for a further paper. For the further conduct of this study, we decided to analyze the convergence issue in terms of ULC growth rates – a variable which is at highest I(1). The reasoning is twofold: on the one hand there is a direct link to the discussion about appropriateness of a unique EMU wide inflation rate as the target of monetary policy of the ECB within the Euro area and on the other hand this makes our results comparable with existing studies dealing with inflation differentials in currency unions (Mentz and Sebastian, 2003, Beck et al., 2006, Busetti et al., 2006).

[8]The panel unit root tests were performed using EViews 5.1 and the respective standard settings with regard to lag length (BIC) and bandwidth selection (Newey-West using Bartlett kernel).

## 2.3 A PANIC attack on ULC growth rates

Bai and Ng (2004) suggest a very useful approach to test for panel unit roots in the presence of stationary or nonstationary common components, known as PANIC – Panel Analysis of Nonstationarity in the Idiosyncratic and Common components. PANIC approach allows both idiosyncratic and common components to be integrated of order one, which makes it very flexible in testing panel unit roots. Since we investigate growth rates of unit labor costs, we assume a model with an intercept but without linear trend and following the notation of Bai and Ng (2004) our model is:

$$X_{it} = c_i + \lambda_i' F_t + e_{it} \tag{2}$$

where $X_{it}$ are $i = 1, \ldots, N$ observed growth rates, $F_t$ is an unobserved vector of common factors and $e_{it}$ are unit specific idiosyncratic components. Both $F_t$ and $e_{it}$ are allowed to be I(1) and for this reason the model has to be estimated in differences, where $x_{it} = \Delta X_{it}$, $f_t = \Delta F_t$ and $z_{it} = \Delta e_{it}$, so we estimate the model:

$$x_{it} = \lambda_i' f_t + z_{it} \tag{3}$$

employing the method of principal components. However, we standardize the first differences before estimating in order to avoid possible distortions by volatile series in calculating principal components, see Bai and Ng (2001). In particular, we divide differenced time series by their cross empirical cross-sectional standard deviations. Estimated common factors and idiosyncratic components are then obtained via cumulating for $t = 2, \ldots, T$ and $i = 1, \ldots, N$

$$\hat{e}_{it} = \sum_{s=2}^{t} \hat{z}_{is} \tag{4}$$

$$\widehat{F}_{it} = \sum_{s=2}^{t} \hat{f}_s \tag{5}$$

where $\hat{z}_{it} = x_{it} - \hat{\lambda}_i' \hat{f}_i$ are estimated residuals. Bai and Ng (2004) show that estimated factors and idiosyncratic components are consistent, in particular $T^{-1/2}\hat{e}_{it} = T^{-1/2}e_{it} + o_p(1)$ and $T^{-1/2}\widehat{F}_t = T^{-1/2}HF_t + o_p(1)$, where $H$ is a full rank matrix. This rate of convergence is fast enough to leave the asymptotic distribution of the ADF-test unchanged, if applied to estimated series $\widehat{F}_t$ and

$\hat{e}_{it}$. So we can apply the univariate ADF-test as well as pooled unit root tests to estimated factors and idiosyncratic components respectively. In case of estimated factors we allow for a constant in a test regression and test without any deterministic terms in the panel case of idiosyncratic components.

An important issue regards our sample sizes in terms of $T$. Due the use of annual data we have relatively few data points, but, on the other side, it is known from simulation studies of Shiller and Perron (1985) and Perron (1989) that the span of the data has more influence on the power of unit root tests than the number of observations. In this respect we posses a typical data span for macroeconometric estimation at least for Germany and Europe - more than 30 and 40 years respectively.

Another important issue is determining the number of factors in PANIC framework. Bai and Ng (2002) suggest some information criteria, in particular $IC_{p1}$, $IC_{p2}$ and $IC_{p3}$, to determine the number of factors. But accordingly to Bai and Ng (2002) our sample size is too small to work with suggested criteria and actually there is no closed minima of the criteria ($k < k_{max}$) or they choose too many factors compared to our sample size. Therefore we decided to calculate fractions of total variation in the differenced data explained by individual common factors and set $k = 2$ on the basis of table 2, because the first factor explains the main bulk of variance in Germany and the US, and on the other side the two factors are almost equally important in Europe.

*Insert table 2 about here.*

Estimation results for all four panels can be seen in figures 6, 7, 8 and 9. In case of Germany we obtain one factor that has clearly non-stationary patterns and on the other hand estimated idiosyncratic components look stationary, see figure 6. Unlike the first picture, there are no clear differences in patterns of European factors and idiosyncratic components, see figure 7. Finally, figures 8 and 9 reveals different patterns between US factors and idiosyncratic components, where the two estimated factors appear to be non-stationary.

*Insert figure 6, 7, 8, and 9 about here.*

The visual impression is confirmed by unit root tests. In table 3 we see the results of ADF tests on estimated common factors. The tests were performed with EViews 5 by employing the Schwarz information criteria (SIC) to determine the lag length. The null of unit root is not rejected for Germany in case of the first factor and is rejected in case of the second one. The test results for the EU 12 reveal also no rejection of the null for the first factor and rejection for the second factor. Finally, in the US case the first factor turns out to be non-stationary and the null is rejected in the case of the second factor at 10% level. However, the small sample size in the US case makes the reliability of non-panel unit root tests highly questionable.

Furthermore, we perform panel unit root tests for estimated panels of idiosyncratic components. Two type of tests are calculated: under assumption of a common unit root process suggested by Levin et al. (2002) as well as by Breitung (2000), and under assumption of individual unit root processes proposed by Maddala and Wu (1999) employing a Fisher-type procedure of combining p-values, see table 4. In all cases we employ SIC to select the lag length and Andrews bandwidth selection with quadratic spectral kernel to estimate the long run variances. In all cases we reject the null of panel unit root. So on the basis of the test results we consider all panels of idiosyncratic components as stationary processes.

*Insert tables 3 and 4 about here.*

Combining our PANIC evidence with a visual inspection of estimated factors and idiosyncratic components we conclude that there are a lot of similarities between panels of German *Länder*, US States and US Census Regions. Firstly, in both cases we get at least one non-stationary common factor and stationary idiosyncratic components. Secondly, if we consider the loadings of this first factor ($\lambda_{1i},\ i = 1, \ldots, N$, see (2)), we observe that they reveal not a lot of variation and always posses the same sign. On the other hand, in the European case many individual idiosyncratic components seem not to be very different from the common factor itself in terms of variance and also their individual course. It can be also seen if we consider fractions of the total variation in the data explained by individual factors, see table 2. The fraction of the first factor in the European case is quite small, compared to the results for Germany and the US, and it is almost equal to the fraction of the second factor. Moreover, the individual loadings of the first common factor are very different in the case of EU 12. They clearly appear to form clusters. In particular, there are some countries with relatively large positive loadings and on the other hand units with negative loadings. Last but not least, we can see differences in the persistence of idiosyncratic components. Figure 10 shows cross sectional means of estimated ACFs of the idiosyncratic components. There is much more persistence in the parts of ULC dynamics unexplained by common factors than in the US or Germany.

*Insert figure 10 about here.*

## 2.4 Convergence clubs in ULC growth?

The New Growth Theory allows for the possibility, that countries may not converge to the same level of per capita GDP, productivity or prices but instead sub-groups may form convergence clubs. Hobijn and Franses (2000) propose an algorithm for the identification of convergence clubs based on multivariate stationarity tests. The procedure has recently been applied to regional EMU inflation rates (Busetti et al., 2006). Applying the algorithm using a version of stationarity test which does not allow for an intercept is equivalent to identifying clusters around the same mean (Busetti et al., 2006, p. 15). The procedures has the nice feature that it is independent of the ordering of the series. It is however, not invariant to the number of series in that sense that including additional series may change the composition of clusters.

The clustering algorithm (Hobijn and Franses, 2000, Busetti et al., 2006) is applied to a panel of all possible bivariate differentials in ULC growth rates and can be described as follows:[9]

1. Denote $k_i$ as a set of indices of variables in cluster $i$, $i \leq n^*$, where $n \leq n^*$ denotes the number of clusters. Define $p^*$ as a significance level for the inclusion of a series in the cluster. Proceed with the following steps.

2. Initialize $k_i = \{i\}$, $i = 1, \ldots, n = n^*$ so that each country/ variable is a cluster.

3. For all $i, j \leq n^*$, such that $i < j$ perform a test whether $k_i \cup k_j$ form a cluster according to the criterion of a multivariate stationarity test on the contrast (here: by means of a multivariate version of the Kwiatkowski et al. (1992) test) and let $p^{i,j}$ the resulting p-value of the test. Decide: If $p^{i,j} > p^*$ for all $i, j$ then go to the end of the procedure.

4. Replace cluster $k_i$ by $k_i \cup k_j$ and drop $k_j$, where $i, j$ correspond to the most likely cluster (maximum p-value of the previous step); replace the number of clusters by $n^* - 1$ and go one step back.

[9]The programs for this exercise are available on Bart Hobijn's homepage. We thank Bart Hobijn for helpful comments.

5. The resulting $n^*$ clusters are labeled "convergence clubs" (convergence to a common mean))

6. The procedure proceeds in testing for relative convergence (convergence to a stationary distance) by applying the same procedure with different p-values.

Due to comptutational errors – probably due to the fact, that the cross-sectional dimension is much larger in this case than the time dimension – we were not able to conduct the test for the US states, however, we applied the procedure succesfully for the US census regions. For all tests, we applied a p-value of 0.01 and a bandwidth of 4 for the Bartlett window used to perform the Kwiatkowski et al. (1992) test. The results are however robust with regard to the choice of p-value.[10]

The tables 5 and 6 summarize the result.

*Insert table 5 and table 6 about here.*

There is evidence for 2, respective 3 clusters in the United States, 3 clusters in Germany and 2 cluster in EU 12. There is furthermore no difference between the results for absolute and relative clustering with the exception of the US census regions.[11] We can learn two things from the exercise: First, even in established currency regions we can find evidence for convergence clustering and the existence of stable clusters does not *per se* hinder the functioning of a currency area.

Second, the clusters in EMU confirms the finding that historically there was a "hard currency" block – lead by Germany – where countries like Austria or Netherlands were anchored too and a "soft currency" block (mainly all other countries). Quite astonishing is the finding that Spain is counted as a member of the "hard currency" club. However, looking at the data reveals that Spain for long periods "overshot" the criterion for being a member in the "good boys club" in terms of more inflationary policy, a strategy which was however from time to time interrupted by sharp (nominal) devaluations. Seen over a long period, this is in line with the hypothesis of absolute convergence. It might create a problem however if the mechanisms of (nominal) devaluation are not available anymore and the real exchange rate has to adjust by differences in inflation rates only.

[10]We did not experiment with the bandwidth, since the value was proposed in the paper of Hobijn and Franses (2000).

[11]This can be interpreted as evidence for absolute convergence clusters, because each of them is a relative convergence cluster with a stationary distance of zero as well.

## 2.5 A structural interpretation of factors driving ULC growth

It is a well-known drawback of the principal component method, that the factors stemming from this analysis are not identified in a structural way. However, we can try to identify the structural shocks driving the dynamics in the panel of ULC growth rates indirectly. To do so, we rely on long-run restriction in a structural VAR framework as proposed by Blanchard and Quah (1989).

For each panel with $i$ entities, we estimate $i$ unrestricted VARs, consisting of the two estimated common factors, and the ULC growth rate of entity $i$ in the respective panel. Due to the very few data points, in all cases we allowed 2 lags to enter – without a rigorous testing.

To identify the shocks, we assume, that the first common factor – which is tested to be non-stationary in all cases – is the driving force of non-stationarity (we can e.g. think of a permanent shock which hits the whole currency area) and therefore not affected by any other variable in the long-run. Furthermore we assume, that the second factor (we can e.g. think of a non-permanent but possibly long-lasting shock hitting the whole currency area) is not affected by country-specific (idiosyncratic) factors in the long-run.[12]

Assuming a lag order of one for reasons of simplicity, the structural model can be formulated as:

$$BX_t = \Gamma_0 + \Gamma_1 X_{t-1} + \varepsilon_t \quad \varepsilon \sim N(0, \Sigma_\varepsilon). \tag{6}$$

The estimated reduced form VAR takes the form:

$$X_t = A_0 + A_1 X_{t-1} + e_t \quad e \sim N(0, \Sigma_e) \tag{7}$$

with $A_0 = B^{-1}\Gamma_0$, $A_1 = B^{-1}\Gamma_1$ and $e_t = B^{-1}\varepsilon_t$. To identify the VAR, we re-write the process in the following moving average representation:

$$X_t = \mu + C(L)\varepsilon_t \tag{8}$$

[12]Formally, in the case of EMU, the second factor was found to be non-stationary. In all other cases, the second factor was found to be stationary. However, due to the difficult properties of unit root test in small sample cases, we decided to deal with all panels and VARs equally.

with $\mu = \left[I - \left(B^{-1}\Gamma_1\right)L\right]^{-1} B^{-1}\Gamma_0$ and $C(L) = \left[I - \left(B^{-1}\Gamma_1\right)L\right]^{-1} B^{-1} = \sum_{i=0}^{\infty} A_1^i B^{-1} L^i = \sum_{k=0}^{\infty} C(k) L^k$.

We follow the way proposed by Blanchard and Quah (1989) and impose $n(n+1)/2$ restrictions by the assumption of the orthonormality of the structural innovations: that means $\Sigma_\varepsilon = I$ and $n(n-1)/2$ long-run restrictions. The matrix of long-run multipliers $C(1)$ for each of the VARs takes the following form:

$$\begin{bmatrix} \Delta F_1 \\ \Delta F_2 \\ \Delta ULC_i \end{bmatrix} = \begin{bmatrix} C_{11}(1) & 0 & 0 \\ C_{21}(1) & C_{22}(1) & 0 \\ C_{31}(1) & C_{32}(1) & C_{33}(1) \end{bmatrix} \begin{bmatrix} \varepsilon_1 \\ \varepsilon_2 \\ \varepsilon_3 \end{bmatrix} \tag{9}$$

To analyze the issue of convergence, we used forecast error variance decompositions (FEVD) of the structurally identified VARs. At different forecast horizons, the FEVD give answer to the question, which portion (in per cent) of the variance of the time series' stochastic part can be explained by each of the structural shocks. Specifically we report evidence, how much of the variance in individual ULC growth rates can be explained either by the first structural shock ($\varepsilon_1$)– which has the highest degree of exogeneity since it is not affected by any other shocks in the long run –, or by the second structural shock ($\varepsilon_2$)– which is not affected by country-specific effects in the long run –, or by the country-specific shock ($\varepsilon_3$)itself.[13]

In case of convergence – in the traditional sense, not in the sense of convergence clubs – , we would expect the variance in the panel of time series to be very much driven by exogenous permanent shocks which are influencing each of the panel members in the same way. In contrast, in the case of convergence clubs or divergence, we could have the finding, that different panel members are driven by different forces – might this either be different common factors or idiosyncratic factors.

The results are shown in tables 7, 8, and 9.[14]

*Insert tables 7, 8, and 9 about here.*

In the case of Germany and the US, we confirm the findings from the PANIC analysis. The variance of the individual ULC growth rates is mainly driven

[13]Given the limited number of observations, the point estimates might be misleading. However, we did not bootstrap confidence bounds since the interest is more in the general picture instead of assessing the uncertainty around the point estimates.

[14]Again we skip the results for the US federal states. These results are available from the authors on request.

by the first common factor – which is in line with the hypothesis of convergence. There are some exceptions, where either the second factor and/ or the idiosyncratic shock seems to play a role but these cases are exceptions rather than the rule.

In the case of EMU, we however find very differing effects of the shocks on ULC growth rate variance for different panel members. In some countries, the variance to a considerable extent driven by the first shock (e.g. Belgium, Germany, Luxemburg, Netherlands, Austria) – which seems to grasp lasting shocks to the hard currency club. Other countries are much more connected to the second shock (e.g. Greece, Spain, Portugal, Ireland, Finland) – which seems to grasp common movements in catching-up countries. All in all, the variance explained by idiosyncratic shocks is much higher than in the US or Germany. These results are in line with the high dispersion in the factor loadings in the PANIC analysis and the relatively high persistence of idiosyncratic components compared to the US and Germany. More specifically, we can confirm to some extent the result from the convergence club analysis.

# 3 Discussion

Our analysis of ULC growth dynamics in selected countries/ regions points to the existence of one non-stationary common factor for the United States (States and Census regions) and Germany. The idiosyncratic components in all currency areas are found to be stationary. This is in line with the hypothesis of $\beta$-convergence around common factors (which could e.g. be associated by country-wide factors like supply or demand shocks like oil price hikes or monetary policy actions). For the Euro area, however, it is quite difficult to identify common factors – idiosyncratic factors dominate in explaining the bulk of variance – and clustering seems to be present. It cannot be rejected that idiosyncratic factors in the Euro area are stationary, however the persistence is much stronger than in other currency areas – which points to long-lasting adjustment processes. There is little sign of change in that respect in the second half of the sample. The case for convergence clusters is confirmed when using the procedure as in Hobijn and Franses (2000) for all currency areas. The differences between individual ULC growth rates are, however, much smaller for the United States and Germany than within the EU 12 (table 10).

These results were confirmed by an SVAR analysis to give the unidentified factors a more structural interpretation. Forecast error variance decomposition reveals that for the United States and Germany one non-stationary common factor can be identified as the main source of forecast error variance of almost all individual ULC growth rates. The picture differs dramatically with regard to the EMU countries. Here, the finding is more in line with the existence of convergence clubs and a very strong influence of idiosyncratic components.

Why should a central bank like the ECB be concerned about this finding? Prices in one country would be a little higher than in the rest of the union for a number of years and below the average for another number of years. The phenomenon, one could argue, is a purely nominal one. However, there are a number of theoretical arguments to cast doubt on this view. Instead, if divergences persist for a prolonged periods, they might cause misallocations and even long-term detrimental effects to growth. There are different reasons for that: Domestic inflation makes the financing of consumption and capital input cheaper while investment in the tradable sector becomes less attractive with the loss of competitiveness, it might lead to excessive investment in the housing sector – especially if financial markets are incomplete and suffer from asymmetric information problems. Not only might an excessive amount of capital be allocated to this sector which contributes relatively little to long-term productivity growth but thus might shift the Beveridge curve outwards and increase structural (long-term) unemployment. Furthermore, persistent deviations in the price trend might lead to a strong overvaluation of one country within the monetary union. Whereas undervaluation leads to increasing exports and income, increasing import prices lead to a deterioration in the trade balance and adjustment occurs in the long run – as the exchange rate channel dominates the interest rate channel. Adjustment processes might however be asymmetric with regard to speed and intensity, due to hysteresis phenonenom: Once trapped in a situation of overvaluation, profits might suffer and investment contract, leading to a longer period of subtrend economic growth until the real appreciation is corrected again. These boom-and-bust-periods might not only bring about negative welfare effects due to increased income volatility but might also lower the potential output of a single country if there are good arguments for hysteresis in the labour market, meaning that unemployment is at least to a certain extent path-dependent (Fritsche and Logeay, 2002, Gottschalk and Fritsche, 2005). Last but not least, political economy arguments hint that prolonged boom-and-bust cycles as a result from divergences might actually endanger the political stability of the euro-area. Leaving the union would allow the country to depreciate sharply and forego the adjustment costs of relative wage deflation. If the country's politicians have a sufficiently high personal discount rate, the short-term benefits of leaving EMU might actually be perceived larger than the long-run costs of the forgone membership in the monetary union such as lower long-term interest rates. This might in the end lead to single countries pulling out of EMU.

One might argue, that the perceived results might be due to the time span under investigation. Due to a lack of evidence after the introduction of the Euro and given the existence of long-lasting adjustment processes, we can only informally test for structural change. However, preliminary stability investigations as well as a visual inspection of the factor decomposition analysis results gives rise to serious concern for the Euro area. The behaviour of idiosyncratic components does not seem to have changed and shows strong persistence. The same is true when looking at the cluster procedure results – there is still evidence for inflation clubs in the EMU (12). This finding is in line with the findings of Busetti et al. (2006) – where in contrast to our study monthly inflation numbers were used. The discussion has clear implications for the conduct of economic policy within the Euro area. The lasting evidence for persistent inflation differentials calls for re-organization of macroeconomic policy at a European level – e.g. newly designed fiscal transfer mechanisms or wage policy coordination (Fritsche et al., 2005)– or for increased labor mobility and productivity adjustment (Belke and Gros, 2006, Blanchard, 2006). Both solution have their own advantages and drawbacks – which goes beyond the scope of this paper.

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# Appendix

## Figures and Tables

| Currency area | Test | on... log($y$) p-val. | $\Delta \log(y)$ p-val. | $\Delta(\Delta \log(y))$ p-val. |
|---|---|---|---|---|
| Euro area | | | | |
| | Levin et al. (2002) | 0.32 | 0.00 | 0.00 |
| | Breitung (2000) | 0.90 | 0.00 | 0.00 |
| | Im et al. (2003) | 1.00 | 0.00 | 0.00 |
| | Dickey and Fuller (1979) (Fisher $\chi^2$) | 0.99 | 0.00 | 0.00 |
| | Phillips and Perron (1988) (Fisher $\chi^2$) | 1.00 | 0.00 | 0.00 |
| | Hadri (2000) | 0.00 | 0.00 | 0.73 |
| Germany | | | | |
| | Levin et al. (2002) | 0.00 | 0.00 | 0.00 |
| | Breitung (2000) | 1.00 | 0.00 | 0.00 |
| | Im et al. (2003) | 0.00 | 0.00 | 0.00 |
| | Dickey and Fuller (1979) (Fisher $\chi^2$) | 0.00 | 0.00 | 0.00 |
| | Phillips and Perron (1988) (Fisher $\chi^2$) | 0.00 | 0.00 | 0.00 |
| | Hadri (2000) | 0.00 | 0.00 | 0.01 |
| USA (States) | | | | |
| | Levin et al. (2002) | 0.00 | 0.00 | 0.00 |
| | Breitung (2000) | 1.00 | 0.00 | 0.00 |
| | Im et al. (2003) | 0.00 | 0.00 | 0.00 |
| | Dickey and Fuller (1979) (Fisher $\chi^2$) | 0.00 | 0.00 | 0.00 |
| | Phillips and Perron (1988) (Fisher $\chi^2$) | 0.00 | 0.00 | 0.00 |
| | Hadri (2000) | 0.00 | 0.00 | 0.72 |
| USA (Regions) | | | | |
| | Levin et al. (2002) | 0.00 | 0.07 | 0.00 |
| | Breitung (2000) | 0.96 | 0.01 | 0.00 |
| | Im et al. (2003) | 0.00 | 0.11 | 0.00 |
| | Dickey and Fuller (1979) (Fisher $\chi^2$) | 0.00 | 0.19 | 0.00 |
| | Phillips and Perron (1988) (Fisher $\chi^2$) | 0.00 | 0.53 | 0.00 |
| | Hadri (2000) | 0.00 | 0.00 | 0.95 |

Table 1: ULC panel unit root tests

| | Factor 1 | Factor 2 |
|---|---|---|
| Germany | 0.654 | 0.098 |
| EU 12 | 0.227 | 0.211 |
| US States | 0.410 | 0.108 |
| US Census Reg. | 0.739 | 0.123 |

Table 2: Fractions of the total variation in the differenced data explained by individual common factors.

| | Factor 1 | Factor 2 |
|---|---|---|
| Germany | -1.88 (0.34) | -8.39 (0.00) |
| EU 12 | -2.35 (0.16) | -3.34 (0.02) |
| US States | -1.80 (0.37) | -2.91 (0.06) |
| US Census Reg. | -1.86 (0.34) | -3.08 (0.05) |

Table 3: ADF test result for the estimated common factors, performed in EViews 5.0 with MacKinnon (1996) critical values. SIC is used to determine lag length of test regressions, where $p = 0$ results for all tests regressions and $p_{max}$ is set automatically by EViews. P-values are in brackets.

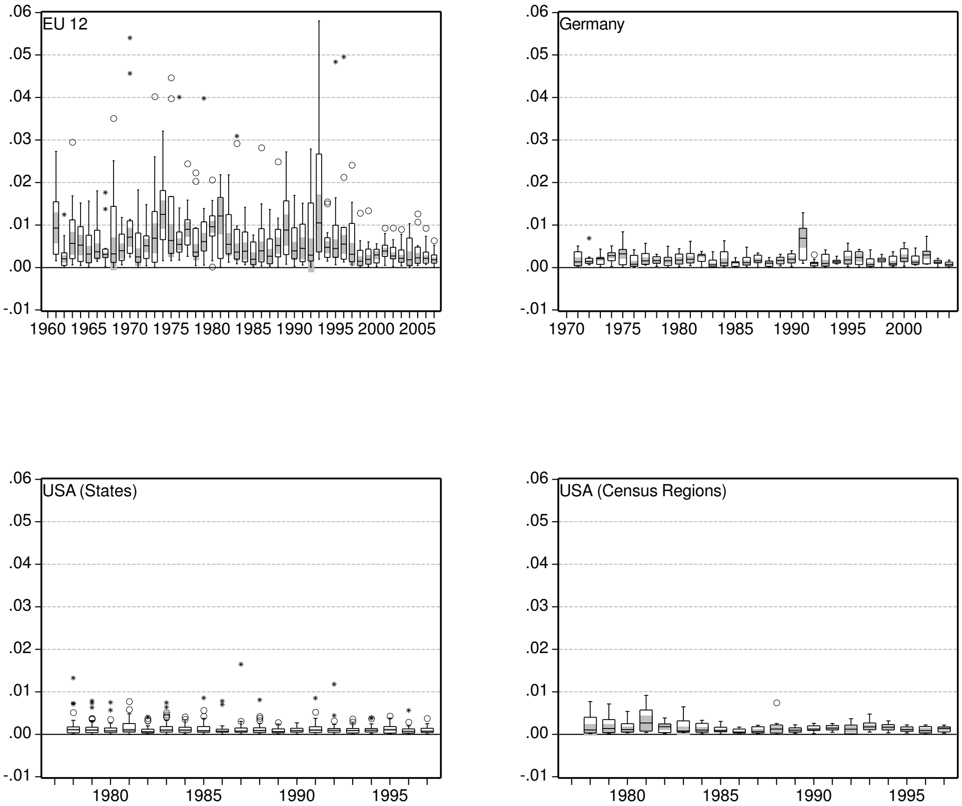


Figure 1: Unweighted ULC growth variability: Box-Plots

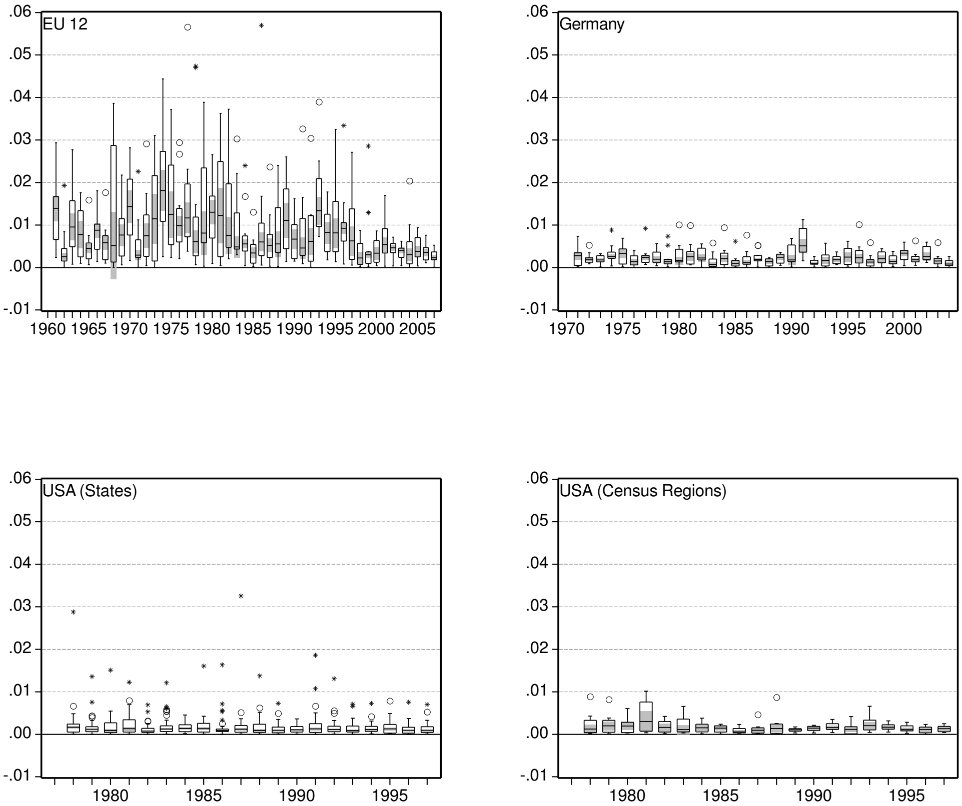


Figure 2: Weighted ULC growth variability: Box-Plots

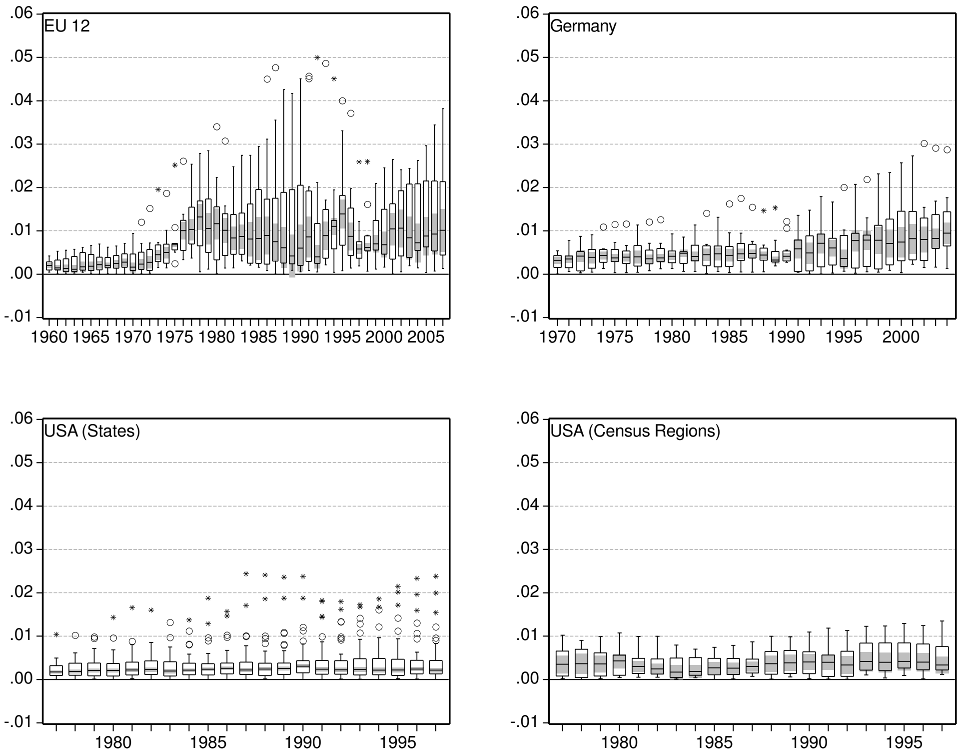


Figure 3: Unweighted ULC level variability: Box-Plots

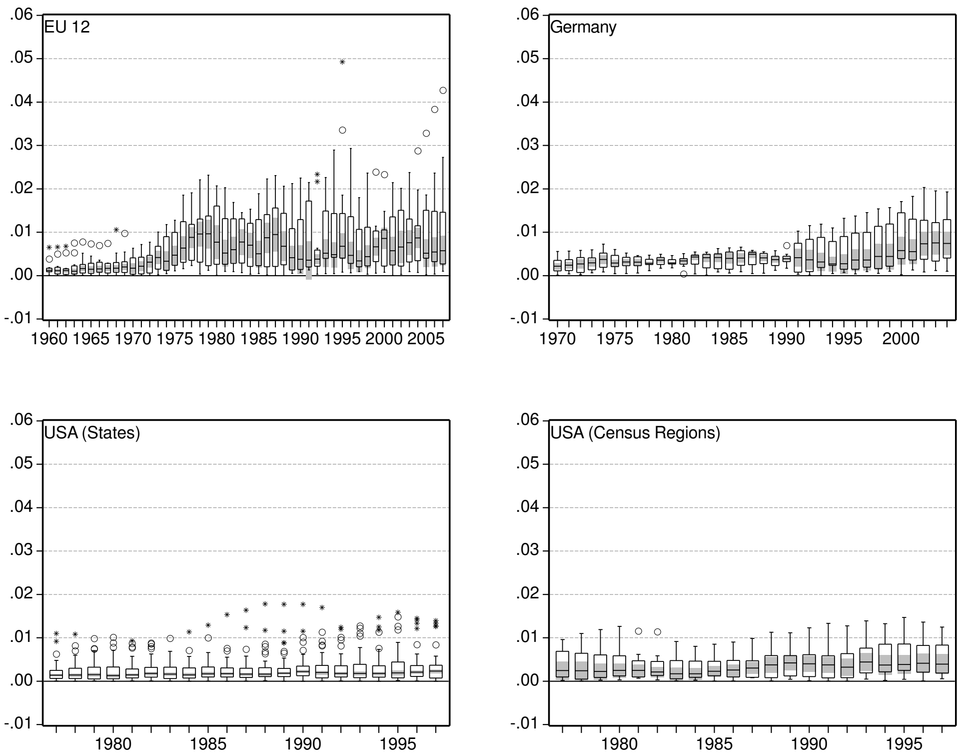


Figure 4: Weighted ULC level variability: Box-Plots

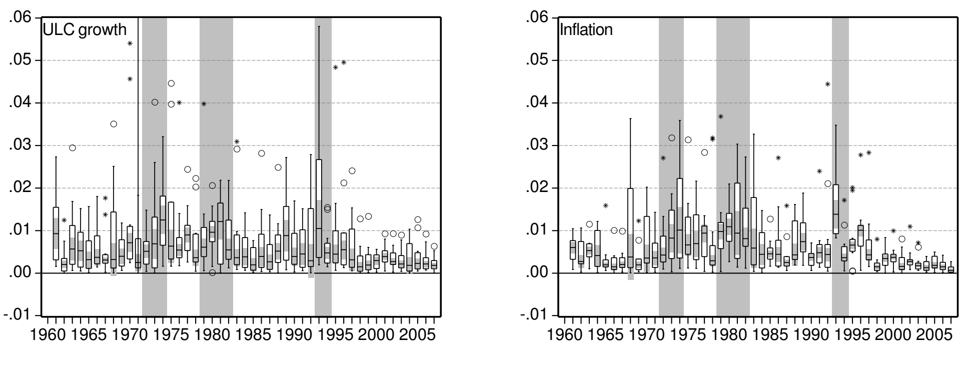


Figure 5: ULC growth variability and inflation variability in the Euro area: Box plots

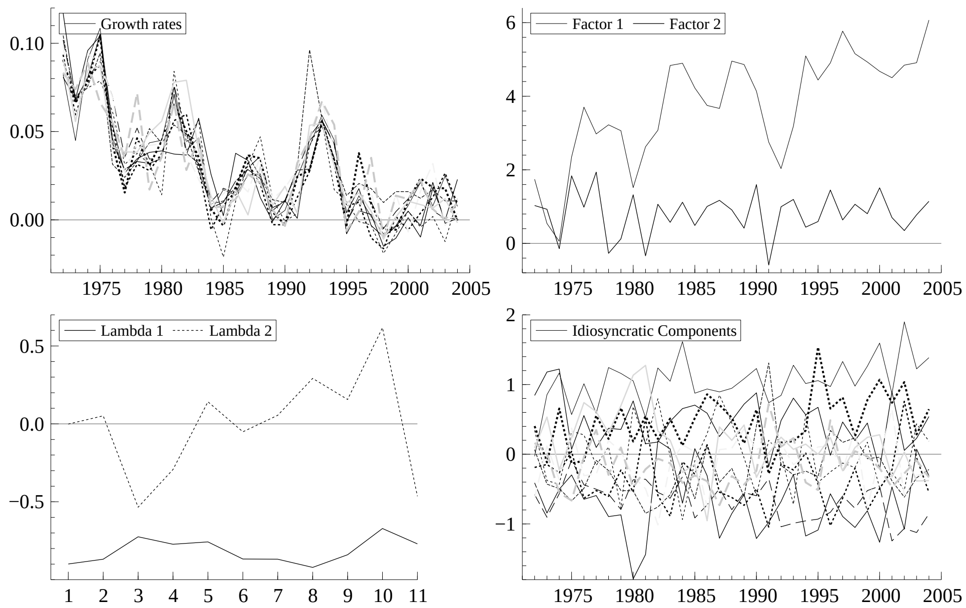


Figure 6: Observed ULC growth rates, estimated factors, loadings and idiosyncratic components for German *Länder*.

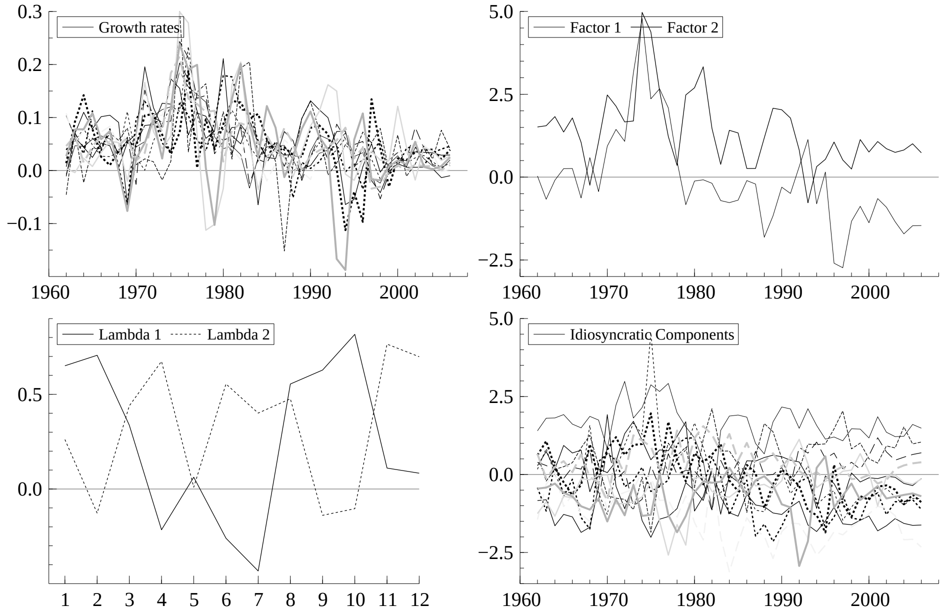


Figure 7: Observed ULC growth rates, estimated factors, loadings and idiosyncratic components for the European Union.

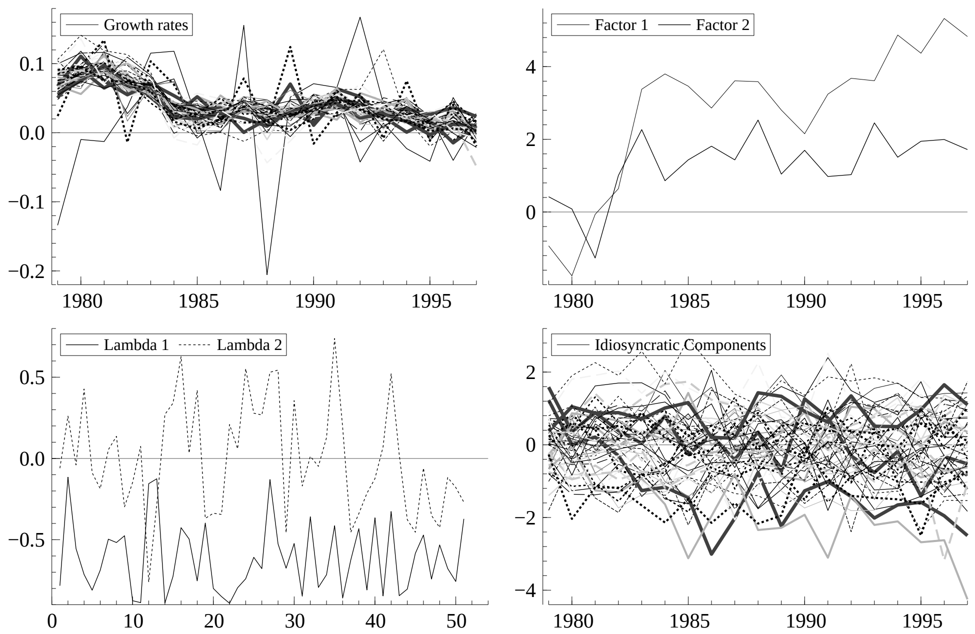


Figure 8: Observed ULC growth rates, estimated factors, loadings and idiosyncratic components for US States.

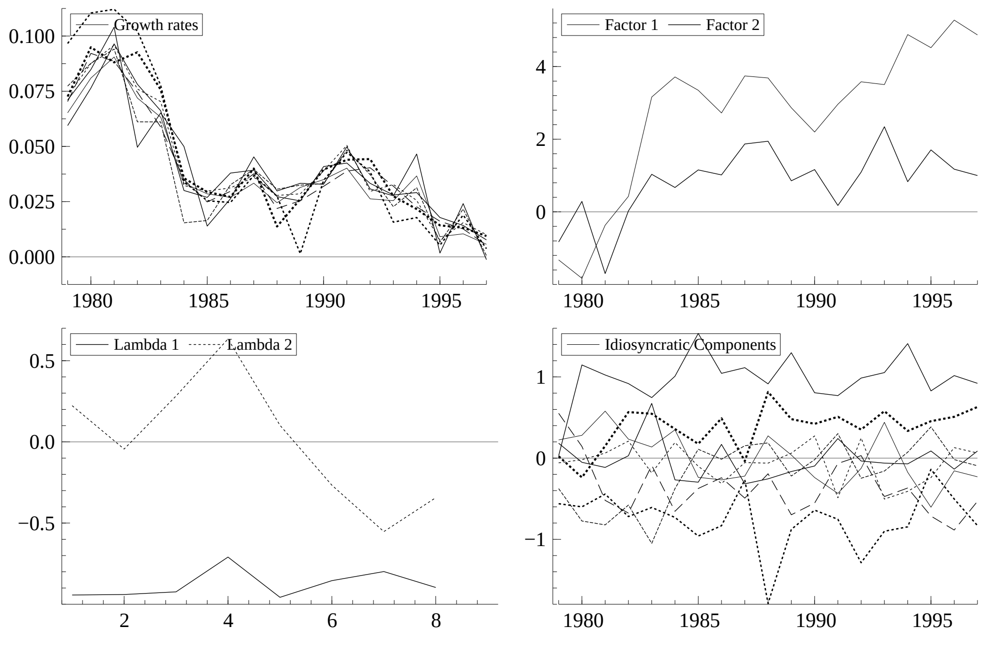


Figure 9: Observed ULC growth rates, estimated factors, loadings and idiosyncratic components for US census regions.

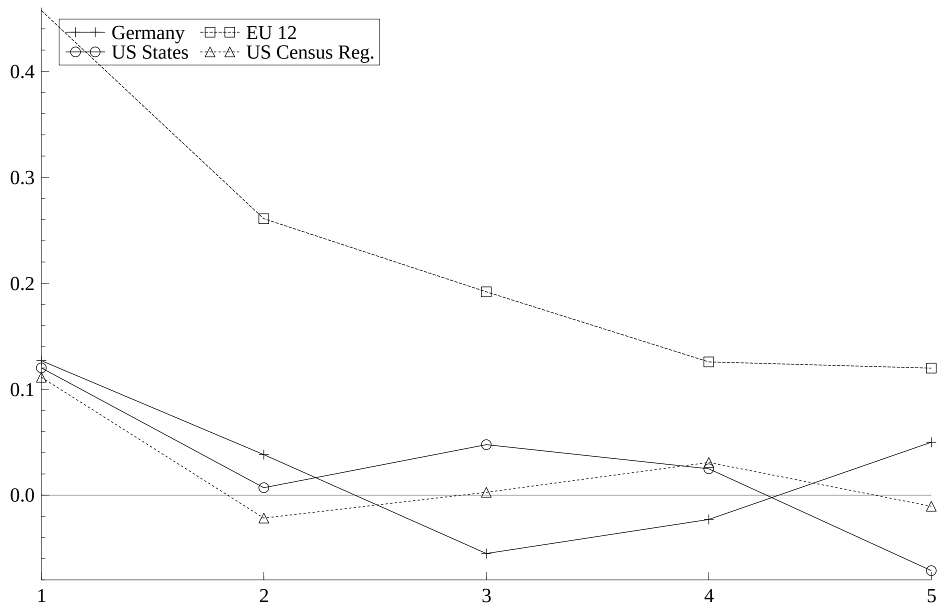


Figure 10: Cross sectional means of estimated autocorrelation functions of idiosyncratic components up to the fifth lag.

| | LLC | Breitung | ADF (Fisher) | PP (Fisher) |
|---|---|---|---|---|
| Germany | -4.45 (0.00) | -4.43 (0.00) | 124.93 (0.00) | 142.82 (0.00) |
| EU 12 | -3.40 (0.00) | -4.70 (0.00) | 93.85 (0.00) | 135.01 (0.00) |
| US States | -7.74 (0.00) | -6.83 (0.00) | 267.74 (0.00) | 353.16 (0.00) |
| US Census Reg. | -3.41 (0.00) | -1.84 (0.03) | 60.69 (0.00) | 62.86 (0.00) |

Table 4: Results of panel unit root tests assuming a common unit root process (LLC denotes Levin, Lin and Chu test, Breitung denotes Breitung t-statistic) and assuming individual unit root process (ADF and PP denote Fisher tests using individual ADF and PP tests), no deterministic is included, maximal lag length is set automatically by EViews 5.0 using SIC, Andrews bandwidth selection using Quadratic Spectral kernel is employed in the PP case. P-values are in brackets and are calculated under assumption of asymptotic normality or $\chi^2$ distribution.

| Region | No. of clusters | Cluster 1 | Cluster 2 | Cluster 3 |
|---|---|---|---|---|
| Euro area | 2 | Belgium | Germany | |
| | | Greece | Spain | |
| | | France | Luxembourg | |
| | | Ireland | Netherlands | |
| | | Italy | Austria | |
| | | Portugal | | |
| | | Finland | | |
| Germany | 3 | Baden-Wuerttembg. | Berlin | Hamburg |
| | | Bavaria | North Rh.-Westphalia | Hesse |
| | | Bremen | Rh.-Palatinate | |
| | | Lower Saxony | | |
| | | Saarland | | |
| | | Schleswig-Holstein | | |
| US (regions) | 3 | Plains | South East | Great Lakes |
| | | Middle East | South West | New England |
| | | Far West | Rocky Mountains | |

Table 5: Result of Hobijn and Franses (2000) cluster procedure: test for absolute convergence

| Region | No. of clusters | Cluster 1 | Cluster 2 | Cluster 3 |
|---|---|---|---|---|
| Euro area | 2 | Belgium | Germany | |
| | | Greece | Spain | |
| | | France | Luxembourg | |
| | | Ireland | Netherlands | |
| | | Italy | Austria | |
| | | Portugal | | |
| | | Finland | | |
| Germany | 3 | Baden-Wuerttembg. | Berlin | Hamburg |
| | | Bavaria | North Rh.-Westphalia | Hesse |
| | | Bremen | Rh.-Palatinate | |
| | | Lower Saxony | | |
| | | Saarland | | |
| | | Schleswig-Holstein | | |
| US (regions) | 2 | Plains | South East | |
| | | Middle East | South West | |
| | | Far West | Rocky Mountains | |
| | | Great Lakes | | |
| | | New England | | |

Table 6: Result of Hobijn and Franses (2000) cluster procedure: test for relative convergence

| Country | Forecast Error Variance due to | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | shock 1 at horizon of | | | shock 2 at horizon of | | | shock 3 at horizon of | | |
| | 1 year | 5 year | 10 year | 1 year | 5 year | 10 year | 1 year | 5 year | 10 year |
| Baden-Wuerttembg. | 86 | 62 | 61 | 0 | 13 | 15 | 13 | 24 | 24 |
| Bavaria | 80 | 63 | 62 | 0 | 12 | 13 | 20 | 25 | 25 |
| Berlin | 68 | 66 | 65 | 11 | 15 | 16 | 21 | 19 | 20 |
| Bremen | 44 | 31 | 31 | 9 | 8 | 9 | 47 | 61 | 60 |
| Hamburg | 56 | 44 | 44 | 3 | 12 | 12 | 41 | 44 | 44 |
| Hesse | 70 | 67 | 68 | 0 | 10 | 10 | 29 | 23 | 22 |
| Lower Saxony | 75 | 65 | 65 | 1 | 11 | 11 | 24 | 24 | 24 |
| North Rh.-Westphalia | 77 | 56 | 56 | 11 | 36 | 36 | 12 | 8 | 8 |
| Rh.-Palatinate | 66 | 52 | 51 | 1 | 14 | 15 | 33 | 34 | 34 |
| Saarland | 45 | 33 | 32 | 30 | 53 | 54 | 25 | 14 | 14 |
| Schleswig-Holstein | 70 | 66 | 66 | 8 | 11 | 11 | 22 | 23 | 22 |

Table 7: SVAR Forecast Error Variance Decomposition: Germany

| Region | Forecast Error Variance due to shock 1 at horizon of | | | shock 2 at horizon of | | | shock 3 at horizon of | | |
|---|---|---|---|---|---|---|---|---|---|
| | 1 year | 5 year | 10 year | 1 year | 5 year | 10 year | 1 year | 5 year | 10 year |
| New England | 91 | 84 | 84 | 6 | 12 | 13 | 3 | 3 | 3 |
| M. East | 92 | 79 | 78 | 0 | 14 | 15 | 8 | 8 | 8 |
| Great Lakes | 78 | 53 | 53 | 9 | 32 | 32 | 12 | 15 | 15 |
| Plains | 49 | 34 | 32 | 40 | 50 | 50 | 11 | 17 | 18 |
| S. East | 87 | 68 | 63 | 0 | 13 | 14 | 13 | 18 | 23 |
| S. West | 65 | 65 | 63 | 5 | 8 | 9 | 30 | 26 | 28 |
| Rocky Montains | 77 | 62 | 62 | 16 | 17 | 18 | 6 | 20 | 20 |
| Far West | 81 | 83 | 82 | 7 | 7 | 7 | 13 | 11 | 11 |

Table 8: SVAR Forecast Error Variance Decomposition: United States (Census Regions)

| Country | Forecast Error Variance due to shock 1 at horizon of | | | shock 2 at horizon of | | | shock 3 at horizon of | | |
|---|---|---|---|---|---|---|---|---|---|
| | 1 year | 5 year | 10 year | 1 year | 5 year | 10 year | 1 year | 5 year | 10 year |
| Belgium | 52 | 52 | 52 | 5 | 8 | 8 | 43 | 40 | 40 |
| Germany | 49 | 41 | 40 | 1 | 3 | 3 | 50 | 56 | 56 |
| Greece | 4 | 4 | 4 | 31 | 31 | 31 | 65 | 65 | 65 |
| Spain | 1 | 5 | 5 | 48 | 44 | 43 | 51 | 52 | 52 |
| France | 9 | 10 | 11 | 2 | 33 | 33 | 88 | 57 | 56 |
| Ireland | 5 | 7 | 7 | 30 | 33 | 33 | 66 | 59 | 59 |
| Italy | 31 | 40 | 39 | 15 | 17 | 17 | 55 | 44 | 44 |
| Luxembourg | 45 | 27 | 27 | 12 | 13 | 13 | 43 | 61 | 61 |
| Netherlands | 59 | 52 | 51 | 1 | 6 | 6 | 40 | 42 | 42 |
| Austria | 75 | 69 | 69 | 1 | 3 | 3 | 24 | 28 | 28 |
| Portugal | 5 | 12 | 13 | 51 | 54 | 55 | 44 | 33 | 32 |
| Finland | 0 | 1 | 2 | 50 | 42 | 42 | 50 | 56 | 56 |

Table 9: SVAR Forecast Error Variance Decomposition: EU 12

| | |
|---|---|
| EU 12 before 1998 | 0.070 |
| EU 12 after 1998 | 0.023 |
| Germany | 0.014 |
| United States | 0.010 |

Table 10: Average S.E. in panels of ULC growth differentials

| Identifier | EMU | Germany | US Census Regions | US Federal States |
|---|---|---|---|---|
| 1 | Belgium | Baden-Wuerttembg. | New England | Alabama |
| 2 | Germany | Bavaria | Middle East | Alaska |
| 3 | Greece | Berlin | Great Lakes | Arizona |
| 4 | Spain | Bremen | Plains | Arkansas |
| 5 | France | Hamburg | South East | California |
| 6 | Ireland | Hesse | South West | Colorado |
| 7 | Italy | Lower Saxony | Rocky Montains | Connecticut |
| 8 | Luxembourg | North Rh.-Westphalia | Far West | Delaware |
| 9 | Netherlands | Rh.-Palatinate | | District of Columbia |
| 10 | Austria | Saarland | | Florida |
| 11 | Portugal | Schleswig-Holstein | | Georgia |
| 12 | Finland | | | Hawaii |
| 13 | | | | Idaho |
| 14 | | | | Illinois |
| 15 | | | | Indiana |
| 16 | | | | Iowa |
| 17 | | | | Kansas |
| 18 | | | | Kentucky |
| 19 | | | | Louisiana |
| 20 | | | | Maine |
| 21 | | | | Maryland |
| 22 | | | | Massachusetts |
| 23 | | | | Michigan |
| 24 | | | | Minnesota |
| 25 | | | | Mississippi |
| 26 | | | | Missouri |
| 27 | | | | Montana |
| 28 | | | | Nebraska |
| 29 | | | | Nevada |
| 30 | | | | New Hampshire |
| 31 | | | | New Jersey |
| 32 | | | | New Mexico |
| 33 | | | | New York |
| 34 | | | | North Carolina |
| 35 | | | | North Dakota |
| 36 | | | | Ohio |
| 37 | | | | Oklahoma |
| 38 | | | | Oregon |
| 39 | | | | Pennsylvania |
| 40 | | | | Rhode Island |
| 41 | | | | South Carolina |
| 42 | | | | South Dakota |
| 43 | | | | Tennessee |
| 44 | | | | Texas |
| 45 | | | | Utah |
| 46 | | | | Vermont |
| 47 | | | | Virginia |
| 48 | | | | Washington |
| 49 | | | | West Virginia |
| 50 | | | | Wisconsin |
| 51 | | | | Wyoming |

Table 11: Country/ Region identifiers used in figures