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DAVID KATRENCIK
JOANNA TYROWICZ
PIOTR WÓJCIK

UNEMPLOYMENT CONVERGENCE IN TRANSITION

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Unemployment Convergence in Transition

David Katrencik

Technical University of Ostrava

Joanna Tyrowicz

Faculty of Economic Sciences, University of Warsaw

Central Bank of Poland

Piotr Wójcik

Faculty of Economic Sciences, University of Warsaw

Abstract

In this paper an attempt is made to inquire the dynamics of regional unemployment rates in transition economies. We use policy relevant NUTS4 unemployment rates for transition economies characterised by both relatively intense (Poland, Slovakia) and relatively mild labour market hardships (namely Czech Republic). We apply diverse analytical techniques to seek traces of convergence, including beta - and sigma-convergence as well as time-series approach. Results in each of the countries suggest no support in favour of beta-type convergence, i.e. convergence of levels. Even controlling for nation-wide labour market outlooks (conditional convergence) does not provide any support to this hypothesis. Further, regions with both very high and very low unemployment show signs of high persistence and low mobility in the national distribution, while the middle ones tend to demonstrate higher mobility and essentially no regional unemployment differentials persistence. This diagnosis is confirmed by sigma-convergence analysis which indicates no general divergence or convergence patterns. Transitions seem to be more frequent, but at the same time less sustainable for middle range districts, while movements up and down the ladder occur frequently for the same districts. Findings allow to define the patterns of local labour market dynamics, pointing to differentiated divergence paths. Importantly, these tendencies persists despite cohesion policies financing schemes, which allocate relatively more resources to deprived regions in all these countries.

Keywords:

unemployment, beta-type convergence, sigma-type convergence, stochastic convergence, social cohesion

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1 Introduction

The reasons for inquiring regional unemployment convergence are twofold. Firstly, socio-economic cohesion constitutes the main purpose of many policies, frequently supported by considerable resources. This is true on both national and supranational level, with EU targeting cohesion with several specialised funds (namely Cohesion Fund and European Social Fund). Differences in regional unemployment rates are often used to describe regional economic inequality, while relative labour market hardships often serve explicitly as discriminating factors in resources allocation. Understanding the persistency of regional unemployment differences helps to assess how effective regional policies have been.

The second reason draws back to Blanchard and Katz (1992) paper. For labour markets adjusting towards equilibrium in the long run, there exist two main channels for convergence of regional unemployment rates. Either unemployed workers can undertake employment in regions where labour demand exceeds supply or capital can flow to low-wage locations to take advantage of lower labour costs. Naturally, the speed of adjustment may indeed be very slow, leading to relatively persistent unemployment disparities, as forcefully argued by Armstrong and Taylor (2000). In addition, new asymmetric shocks may appear, contributing further to initial regional unemployment rate differentials.

Empirical strategy for verifying the convergence hypothesis developed so far are varied. The most obvious is the test of β convergence (unconditional and conditional). Finding β convergence corresponds to proving that *levels* of unemployment converge to a common rate, while levels themselves may be conditioned on structural parameters characterising particular local labour market. Consequently, unconditional β convergence describes one common level for all regions, whereas conditional one allows for differentiated levels for groups of structurally similar communities. One can also inquire if the *dispersion* of unemployment lowers over time and this may be approached testing for σ convergence. Finally, one can try to investigate how persistent the regional unemployment rate *differentials* are, by applying the concept of stochastic convergence.

Transition economies typically experienced rapid growth of the unemployment rates due to profound restructuring. Naturally, these processes affected local labour markets asymmetrically, since regions were diversified with respect to industry composition and economic outlooks. In this paper we analyse unemployment rates for three transition economies experiencing very different unemployment evolutions over the past years: Czech Republic, Poland and Slovakia. The two latter are consistently scoring highest in the EU in as far as labour market hardships are concerned. Conversely, Czech Republic enjoys a more favourable situation. We resort to policy relevant NUTS4 level monthly data covering the time span of 1999-2007 for Poland and 1995-2007 period for Czech Republic and 1997-2004 for Slovakia.

By applying the variety of convergence analysis tools we intend to inquire about the dynamics of local labour markets evolutions. We demonstrate that these distributions are highly stable over time. Some evidence in favour of "convergence of clubs" is supported by the data, but only for high unemployment regions. Moreover, regional differentials seem to be highly persistent. Whereas this last finding can be attributed to relatively short time horizon, the conclusions concerning the dynamics do not seem to be all driven by temporary shocks.

The paper is structured as follows. The next section focuses on the brief literature review in order to justify the use of multiple empirical strategies. These are outlined in section 3, while results for respective analyses are presented in section 4. Section 5 concludes with some indications of future research directions.

2 Literature review

Generally, in the literature regional unemployment disparities have been more at the core of interests for regional researchers than for economists (see: Pehkonen and Tervo (1998) for the Finland, Dixon, Shepherd and Thomson (2001) for Australia and Gray (2004) for the UK). Most recently, Huber (2007) surveys empirical literature on the regional labour market developments in transition countries. Boeri and Terrell (2002) inquire if these differentials could be explained on the grounds of the optimal speed of

transition theory (see: Ferragina and Pastore (2005) for an extensive review). At the same time, Buettner (2007) compares empirical evidence on regional labor market flexibility in Europe and, in particular, in the EU-accession countries in Central and Eastern Europe. Whereas substantial regional disparities in unemployment are found for pre-accession EU member countries as well as for accession countries, an empirical analysis taking account of spatial effects shows that regional wage flexibility is significantly higher for accession countries. In an impressive volume on the evolution of the Czech labour market Flek, Galuscak, Gottvald, Hurnik, Jurajda and Navrati (2004) argue that over the period of ten years, a transition from over-employment to under-employment may have occurred.

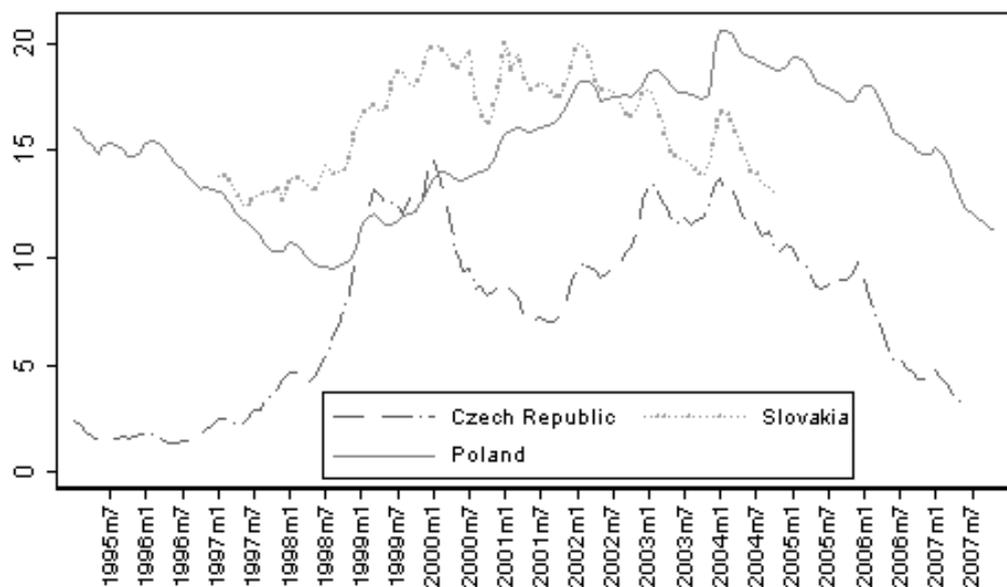


Figure 1: Unemployment rates in Czech Republic, Poland and Slovakia

The process of employment restructuring in most formerly centrally planned economies consisted mainly of the reductions in employment with growing average job tenure as well as average time spent in unemployment or inactivity (cfr. Svejnar (2002)). Dismissals - if compensated at all - found their outcome with hiring of young, better educated workers, but with standard obstacles youth faces when entering the labour markets in Europe. People who lost their employment usually became permanently unemployed or inactive (Grotkowska 2006). In Poland for example, currently less than 13% of the unemployed still retain the right to unemployment assistance, thus suggesting that most of the unemployed are either long-term unemployed (above 12 months) or have a long record of unstable employment (with less than 6 consecutive months of work).

Thus, on an individual level one can easily point the ideal type of winners and losers in the transition process. However, in terms of regional analysis the "conventional wisdoms" are no longer comparably relevant. Some of the highest unemployment regions are located relatively close to the "growth poles", while regions typically considered to lag behind exhibit average labour market indicators. Scarpetta and Huber (1995), Góra and Lehman (1995), Lehmann and Walsh (1998) and more recently Newell and Pastore (1999), suggest restructuring and heavy industry location as main differentiation factors. In the similar spirit Buettner (2007) demonstrated that regional unemployment disparities are indeed profound across most of the CEECs countries, including Czech Republic, Poland and Slovakia. However, this research used differentiated levels for desaggregation (pre-reform NUTS2 for Poland¹, current NUTS3 for Slovakia and current NUTS4 for Czech Republic).

Importantly, most of these findings are not consistent with recent labour market developments.

¹This suggests that the data set ends with December 1998, essentially only the middle of transition.

Studies from the beginning of the previous decade used fairly aggregate and not policy-relevant level data². Moreover, if we follow any of these approaches in understanding the unemployment dynamics on the policy relevant level (NUTS4 in each of these countries), none of the argued suggested hypotheses hold. Rural NUTS4 regions tend to exhibit a whole spectrum of unemployment rates, with averages fairly similar to industrialised NUTS4 regions. In addition, regions experiencing restructuring in the beginning of transformation perform currently both very well and very bad.

Finally, none of these studies takes into account that over the past fifteen years local labour markets were subject to many other context-specific shocks, positive (*e.g.* active labour market policies, sometimes specifically targeting one particular group of unemployed in a particular location) as well as negative (*e.g.* currency crisis in Czech Republic, Russian embargo on Polish exports, etc.)³. Figure (1) presents the evolution of the unemployment rates in the three countries considered in our study.

Consequently, in this paper we attempt to fill the gap in the literature on the evolution of local labour markets in transition by inquiring the dynamics at the policy relevant NUTS4 levels in Czech Republic, Poland and Slovakia.

3 Empirical strategies

One can imagine four main dynamic evolution patterns, which is depicted by the Figure (2). The first is suggested by unconditional convergence, implying evolutions becoming alike both in terms of levels and in terms of the deviations from these levels (scenario A). If the local unemployment rates exhibited such property, at national level cohesion would be fostered despite even relatively intense labour market hardships.

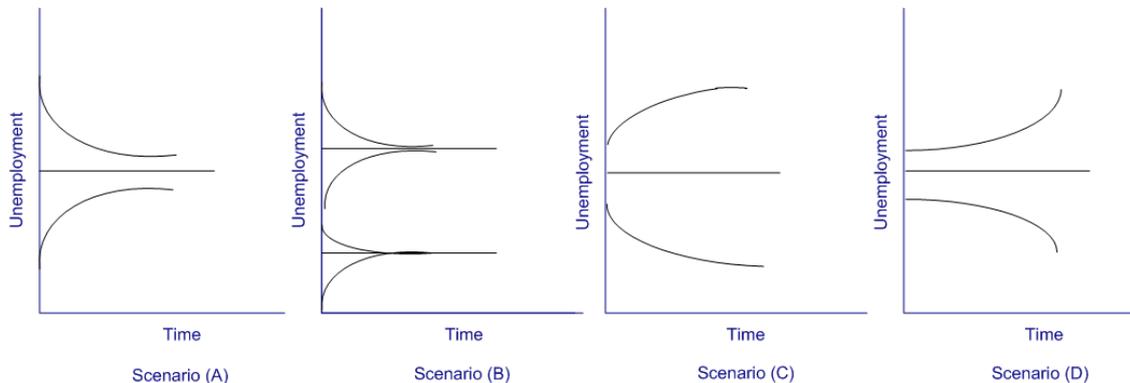


Figure 2: Dynamic evolution patterns

Secondly, convergence may still occur but to differentiated levels. Consequently, total sample deviations from average might persist, but within groups both levels and deviations converge (scenario B). In terms of policy evaluation, such findings may be interpreted as partially successful cohesion efforts (within groups) or a lack of success (if differences are driven by structurally dislike fundamentals).

Finally, data may exhibit divergence, either limited or explosive in terms of deviations (scenario C and D, respectively). Importantly, in the case of scenarios C and D, computed average contains no useful information concerning the behaviour of the unemployment rates. In addition, even in the case of scenario B one should understand that the average computed is nonexistent, it is a statistical artefact of two (or more) group averages.

²With the exception of Newell and Pastore (1999) who work on LFS data, all of the above papers for Poland use 49 voivodships (comparable to the current 44 regions at NUTS3 level), which currently are not even equipped in any authorities, let alone public employment service bureaus.

³See Martin (2006) for a case study of the impact decline in Danube transport is believed to have had on employment in all riparian countries, namely: Austria, Bulgaria, Croatia, Germany, Hungary, Moldova, Romania, Slovakia, Ukraine, Serbia and Montenegro as well as some nonriparian ones.

3.1 Convergence of levels

Convergence of levels is typically investigated through the β analysis. In its unconditional form (scenario A), in principle β convergence implies that the higher the level was in the beginning of the period, relatively the lower it should become throughout time. Consequently, one expects a negative β coefficient in a regression

$$x_{i,t} = \alpha_i + \beta x_{i,0} + \epsilon_{i,t}. \quad (1)$$

However, imposing the constraint that all i 's need to converge to a unique level may be too demanding if i 's differ substantially in the underlying fundamentals. Therefore equation (1) may be tested in a conditional form, *i.e.*

$$x_{i,t} = \alpha_i + \beta x_{i,0} + \gamma Z_{i,t} + \epsilon_{i,t}, \quad (2)$$

where $Z_{i,t}$ denotes a set of variables differentiating the i 's. Finding a negative β coefficient in equation (2) is equivalent to scenario B. Insignificant estimates of β would suggest no distinct pattern, while positive values definitely provide evidence in favour of divergence.

Unfortunately, regression analysis does not allow to discriminate between scenarios C and D. Neither is it possible to effectively approach the situations in which some groups of i 's would exhibit convergence, while some other divergence. This is where σ approach can provide useful insights.

3.2 Convergence of variance

Kernel density estimates in general approximate an unknown density function for a random variable, basing on a finite number of observations drawn from this distribution. This estimator is continuous equivalent of the histogram. The values of the density function at some point are calculated as relative frequency of the observations in the nearest surrounding of this point (bandwidth window), while this relative frequency is estimated using a density function (kernel).

Although the choice of the kernel function has evident but in fact only slight impact on the way the unknown density functions are estimated, it is the bandwidth window that essentially drives the results. The imposed size predetermines the degree of the curve or surface smoothing. Too wide bandwidth window will hide the real data distribution, while too narrow might misleadingly result in function with multiple vertices - not necessarily true in reality and rather troublesome in terms of interpretation. Silverman (1986) provides the procedures for finding optimal bandwidth, subject to differentiated kernel functions, basing on standard deviations and inter-quartile differentials (independently for all vectors in the case of multidimensional distributions). Another way to avoid the problems associated with choosing the bandwidth of the windows can also be solved by adaptive kernel density estimation, which allows for differentiated bandwidths for each observation and this is the method we employ in the paper.

If the initial unemployment rate is defined by x , while the one for the current period by $x + 1$, the distribution of $x + 1$ conditional on x may be written down as:

$$f[x + 1|x] = \frac{f[x, x + 1]}{f_x[x]}, \quad (3)$$

where $f_x[x]$ is the marginal distribution of the initial unemployment rate, while $f[x + 1|x]$ represents the combined distribution of x and $x + 1$. Estimating the conditional density function, both numerator and denominator of (3) are replaced by non-parametric estimators. By stating that adaptive kernel estimation is employed to estimate marginal distribution of the initial unemployment rates we mean specifically that one-dimensional distributions are applied, *i.e.*:

$$\hat{f}_x^A[x] = \frac{1}{n} \sum_{i=1}^N \frac{1}{h_x w_i} K \frac{x - x_i}{h_x w_i}, \quad (4)$$

where n is the number of observations, h_x is the bandwidth window for the initial unemployment rate and $K[\cdot]$ represents the kernel function⁴. At the first stage, weights w_i take the value of 1 for all observations. The combined distribution of initial and final unemployment distribution i.e. the denominator of equation (3), is thus estimated by:

$$f_{x_t, x_{t+1}}^A \hat{[x]} = \frac{1}{n} \sum_{i=1}^N \frac{1}{h_x h_{x+1} w_i^2} K \left[\frac{x - (x+1)_i}{h_x w_i} \right] K \left[\frac{(x+1) - (x+1)_i}{h_{x+1} w_i} \right], \quad (5)$$

where h_{x+1} is the bandwidth window for the final unemployment rate distribution, while subscript A signifies the use of adaptive technique.

Importantly, at the first stage combined density function is estimated with the optimal bandwidth window, while weights are uniform for all observations. Subsequently, basing on these estimates, local differentiation of bandwidth windows are calculated according to:

$$w_i = \left(\frac{1}{n} \sum_{i=1}^N \frac{1}{h_x h_{x+1} w_i^2} K \left[\frac{x - x_i}{h_x w_i} \right] K \left[\frac{(x+1) - (x+1)_i}{h_{x+1} w_i} \right] \right)^{1/2}. \quad (6)$$

In this expression, the denominator of the formula in the parentheses is the combined density function estimator calculated with the use of uniform weights and bandwidth window⁵, while the numerator gives the geometric average of this estimator for matching couples of both variables. The final conditional density function is found basing on the weights from equation (6) to equations (4) and (5) (calculating their quotient), according to equation (3).

This methodology has shorthand interpretative advantages. First of all, convergence / divergence may be easily detected from the graphs of the conditional density functions. Namely, vertical shape of this function suggests divergence, while horizontal vertical alignment is consistent with the convergence hypothesis. If the conditional density function follows the 45° line, overall density function exhibits stability, i.e. an observation drawn randomly at one point in time is highly unlikely to move towards relatively higher or lower values in any preceding or subsequent point in time.

3.3 Stochastic convergence

Carlino and Mills (1993) suggest a time-series approach to the theoretically motivated imperative of convergence (Blanchard and Katz 1992). They argue that a crucial condition is required for a stochastic convergence, namely that shocks to relative local levels should be temporary only. Consequently, a testable hypothesis of local and national unemployment rates cointegration can be formulated:

$$\forall t : \lim_{s \rightarrow \infty} E(U_{i,t+s} - U_{j,t+s} | I_t) = \text{constant}, \quad (7)$$

where U denotes respective unemployment rate and I_t is the conditioning information set⁶. This is empirically approached by testing for a unit root in

$$u_{i,t} = \ln \frac{U_{i,t}}{\bar{U}_t}. \quad (8)$$

⁴With the large number of observations (approximately 400 for Poland, and 80 for Slovakia and Czech Republic) we uniformly used the Gaussian kernel function, thus implicitly assuming normal distribution. However, Gaussian assumption is by far the most frequently used one, while it only concerns the properties of the nearest surrounding of each point (within the bandwidth windows) and not the distribution as a whole.

⁵Fixed window kernel estimate.

⁶To be precise, this is a conditional stochastic convergence formula. Unconditional version would require the limit to approach 0. However, such a condition would discriminate between scenarios A and B, classifying B as non-convergence. Allowing a non-zero constant, permits to account for across regions differentiation. Although we do not subscribe to the idea of regional amenities by Marston (1985) as laid out also in (Blanchard and Katz 1992) that more interesting regions will be burdened with higher unemployment because people have higher utility of living there anyway. However, we acknowledge that there are numerous structural reasons for regional disparities to persist, especially in the 10 years time horizon, as is the case in this paper.

Armstrong and Taylor (2000) suggest that if the speed of adjustment is slow while external shocks strong, divergence may emerge as a statistical artefact in spite of effective convergence exhibited by the processes. Therefore, cointegration tests should encompass considerations for possible structural breaks. This last approach is applied by Bayer and Juessen (2006) for Germany and Gomes and da Silva (2006) for the case of Brazil. Bayer and Juessen (2006) perform a unit-root test on regional unemployment rate differentials using Mikrozensus data for West Germany over the 1960-2002 time span. They find moderate evidence in support of the convergence hypothesis, namely when controlling for structural breaks unit-root is rejected for the majority of regions. Similarly, Gomes and da Silva (2006) for the six metropolitan regions of Brazil find strong evidence of hysteresis and unemployment regional differential persistence, especially strong for the case of Rio de Janeiro.

Unit-root tests are typically troubled by weak power. To circumvent this problem panel data unit-root tests are applied. In particular, standard regression for these tests bases on:

$$u_{i,t} = \rho_i u_{i,t-1} + z'_{i,t} \gamma + \epsilon_{i,t}, i = 1, \dots, N; t = 1, \dots, T, \quad (9)$$

where $z_{i,t}$ is the deterministic component and ϵ should be a stationary error term under null. Three most recently developed approaches to test stationarity in panel data include Breitung and Meyer (1994) (henceforth BM) Levin, Lin and Chu (2002) (henceforth LLC) and Im, Pesaran and Shin (2003) (henceforth IPS). The first two assume that each unit in the panel shares the autoregressive coefficient (*i.e.* $\rho_i = \rho \forall i$), while BM has better asymptotic properties for larger N and smaller T , while LLC has the opposite characteristics. When compared to the single ADF tests, both BM and LLC enjoy higher power by exploiting the cross-equation parameter restriction on the autoregressive parameter ρ . By contrast, IPS assumes heterogenous adjustment paths, by formulating the alternative hypothesis to imply at least one non-stationary variable, but not necessarily all of them.

Unfortunately, each of these tests requires a balanced panel, which is not always feasible due to relatively frequent administrative changes in transition economies. Therefore, whenever forced to do so by the data, we will resort to a Fisher test that combines the p -values from N independent unit root tests, as developed by Maddala and Wu (1999). Based on the p -values of individual unit root tests, Fisher's test assumes that all series are non-stationary under the null hypothesis against the alternative that at least one series in the panel is stationary. Test allows to specify either Phillips-Perron test or Augmented Dickey-Fuller test for each individual case.

4 Data and results

In this paper we employ policy relevant NUTS4 level unemployment data using official registry data for Czech Republic, Poland and Slovakia. In total we use 374 units for Poland⁷, 77 units for Czech Republic and 79 units for Slovakia. Since these are registry data, they suffer from many well-known shortcomings, including underreporting or overreporting (*e.g.* either due to forced passivity or in order gain access to social transfers, respectively). Unfortunately, for none of these countries LFS data can be reliably disaggregated to the NUTS4 level - they are only representative for NUTS3 in the case of Czech Republic and Slovakia and NUTS2 in the case of Poland.

Data cover periods January 1995 till June 2007 for Czech Republic, Jan 1999 till August 2007 for Poland⁸ and January 1997 till October 2004 for Slovakia. The choice of time boundaries was dictated by

⁷Administrative reform of 1999 has introduced the current structure of NUTS4 levels with the exemption of large cities, whose administrative units were separated from the non-agglomerations only as of January 2001. Consequently, prior to 2001 for some district data cover both municipal and rural areas, while after 2001 two districts are formed instead of one, with two separate unemployment rates reported. Since it would be impossible to estimate data for these districts for the period prior to January 2001, for the purposes of analysis we kept both the combined districts and separated ones in relevant periods of time. Thus, the number of units under analysis is inflated to 428.

⁸Observing Figure (1) one sees a significant increase in the unemployment rate in December 2003 in the case of Poland and Slovakia. As of January 2004 new census data from 2002 were applied to calculate the size of the labour force. Thus,

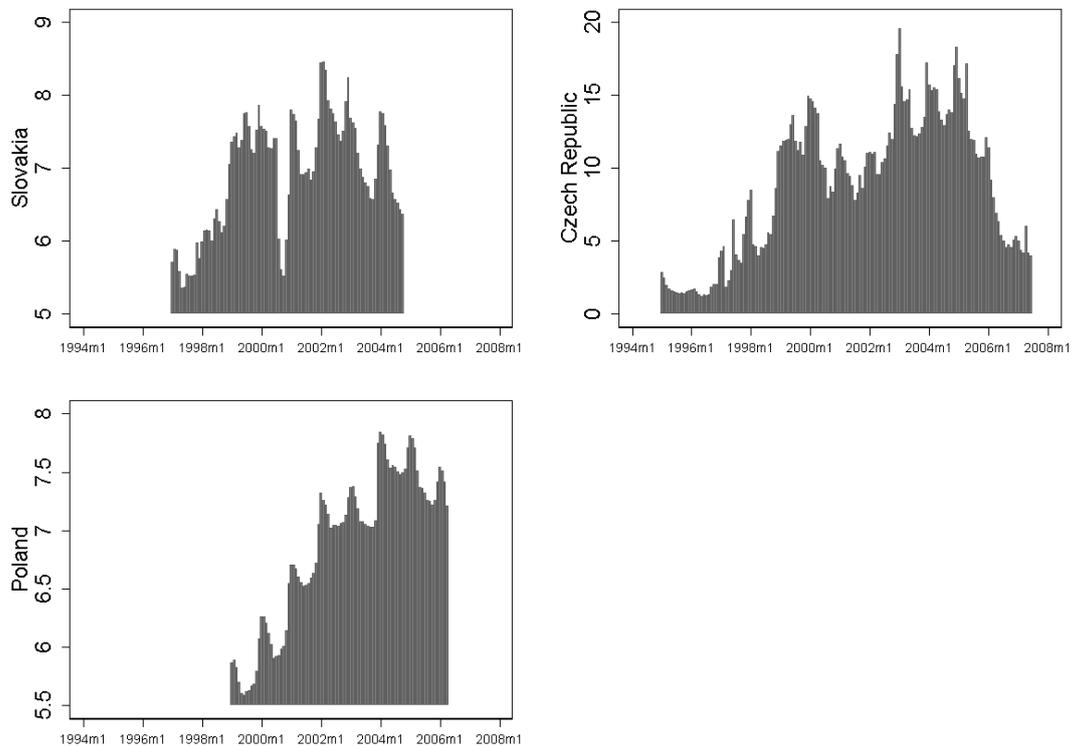


Figure 3: Standard deviations of local unemployment rates distributions

the data availability and seems to bear no serious limitations for the possible results except for one obstacle. Namely, labour market evolutions have commenced in these countries in early 1990s. Unfortunately, NUTS4 data prior to 1999 are not accessible for Poland as the administrative reform establishing this level of local authorities was only implemented as of this year, while only in 2001 metropolitan municipalities were founded. For Czech Republic the consistency of data is destroyed by the change in unemployment definition prior to 1995. For Slovakia, the quality of data prior to 1997 after 2004 is the lowest since the definition of unemployment changed frequently.

Hence, although this paper inquires the dynamics by testing beta, sigma and stochastic convergence, the data analysed commence roughly in the middle of the dynamic evolution patterns. Nonetheless, datasets cover periods of both increases and decreases in the national unemployment rates which is depicted already by Figure (1). Figure (3) demonstrates the standard deviations for the three countries - a measure of the dispersion in the local unemployment rates in every point in time.

Observing these graphs, distributions seems quite volatile over time, with obvious cyclical fluctuations. Over the whole period the averages have been larger than the medians indicating that generally districts with higher unemployment rates are larger. More importantly, as can be inferred from Figure (3), dispersion of the unemployment rates has been growing in the down cycles, be it seasonal effects or general trends in the labour market evolution. This observation suggests that whenever job prospects worsen in general throughout the country, more deprived regions are hit harder in each of the three countries analysed. On the other hand, although rather worrying as a labour market phenomenon, this is rather fortunate from the

although the above unemployment rates base on the registered unemployment recorded by local PES offices, the denominator used for rate calculations at Central Statistical Office has been lowered following the 2002 census. The data have not been re-calculated by CSO for the whole sample, but - for the purposes of comparison from 2004 onwards - December 2003 data were changed, resulting in almost 3,2 percentage point increase in the unemployment rate over only one month. Nonetheless, this change had solely statistical character and does not reflect any labour market process. This effect is controlled for in further research.

empirical point of view, since overall dispersion both increased and decreased in the analysed time horizon. Therefore, obtained results do not risk to be driven by short term uni-direction trends.

4.1 Levels - β convergence

In this section we report the results of a panel regression of unemployment in period t on the unemployment in the initial period (the β -convergence). This is done in both unconditional (simplified) and conditional (extended) framework. For each of the countries estimations were performed separately not to impose logically redundant constraint on the size of the estimated coefficients.

The basic version of unconditional convergence specifies that only the unemployment rate in the initial period needs to be included. However, in the estimation a dummy correcting for the statistical effect of December 2003 in the case of Poland and Slovakia is always present.

In the extended version, to control for low and high unemployment regions, a synthetic proxy was generated, indicating to which of the ten decimal groups a district belonged in the initial period. Since this measure is constructed on the basis of empirical distribution moments, it can take simply the values of 1 to 10, without hazarding the correctness of estimates due to non-linear or non-monotonic effects. To control for cyclicity as well as changing labour market conditions, overall nation wide unemployment rates were incorporated, although from an econometric point of view introducing this variable plays the role of imposing fixed effect on period in the cross-sectional time-series analysis. Finally, some interaction terms were allowed for, to see the extent to which initial distribution and initial unemployment rate effects are symmetric for high and low unemployment regions.

Table 1. Convergence of levels

Dependent variables	Czech Republic			Poland			Slovakia		
Initial unemployment (IU)	1.76*	1.76*	1.76*	1.03*	0.46*	0.44*	1.11*	1.11*	1.11*
National unemployment		1.39*	1.40*		1.11*	1.05*		0.66*	0.49*
Decimal group		0.18*	0.17**		0.53*	0.44*		0.11*	0.09*
Decimal group · IU		0.21*	0.19*		0.02*	0.03*		0.18*	0.11*
Constant	4.51*	-5.97*	-6.02*	4.83*	-10.93*	-9.28*	1.09*	-11.1*	-11.1*
No. of observations	11 538	11 538	11 538	32 578	32 578	32 578	7 365	7 365	7 365
No. of groups	77	77	77	428	428	428	79	79	79
R^2 within	n.a.	0.42	n.a.	n.a.	0.61	n.a.	n.a.	0.44	n.a.
R^2 between	n.a.	0.44	n.a.	n.a.	0.88	n.a.	n.a.	0.90	n.a.
χ^2 statistic	*	*	*	*	*	*	*	*	*
Hausman test	RE	RE	RE	RE	RE	RE	RE	RE	RE
Method	FGLS	OLS	FGLS	FGLS	OLS	FGLS	OLS	OLS	FGLS

Notes: FGLS estimation allows effectiveness even in the presence of AR(1) autocorrelation within panels and cross-sectional correlation and heteroskedasticity across panels. In either case, panel OLS estimates use robust standard errors. Whenever suggested by the data pattern in the conditional analyses time trend and its relevant powers included, but not reported (available upon request).

* and ** denote statistical significance at 1% and 5% levels, respectively. All χ^2 Wald statistics are highly statistically significant, *p-values* available upon request.

Summarising, results are not susceptible to the method of estimation used. The sign and the size of the estimated coefficients remain essentially unaffected when possible heterogeneity in the data is controlled for (cross-sectional time-series FGLS with heteroscedastic panels instead of OLS with robust standard errors).

These findings clearly demonstrate strong divergence in all three countries, while the effect seems to be strongest for a country with lowest unemployment levels, *i.e.* Czech Republic. In the case of Poland the size of divergence seems to decrease significantly for conditional analysis, while for both Czech and Slovak labour markets the estimate for the rate of divergence remains stable independently of including the national labour market variables and decimal group indicators. In the case of Poland the divergence tends to be much more business cycle driven than for the other two countries. The β coefficient drops from 1.03 to approximately 0.5 when conditionality is allowed for, whereas for Czech and Slovak republics inclusion of nation wide trends seems to have no effect on the β estimates. Being located in the 10th decimal group boosts the divergence size by as much as approximately 20 percentage points in the case of Czech Republic, 18 percentage points in the case of Slovakia and additional 2-3 percentage points for Poland.

In principle, β convergence is a necessary but not sufficient condition for σ convergence. Therefore, the above results already rule out finding growing similarities in dispersion. Nonetheless, it is still possible to find either divergence of distributions (also in the form of "clubs") or distributions stability. Thus, we move on to testing the dispersion dynamics.

4.2 Dispersion - σ convergence

To analyse the dynamics of unemployment rates dispersion kernel density estimates were calculated for immediate (month-to-month) and indirect (yearly, *i.e.* 12-month) rolled transitions. Transitions are less likely to demonstrate stability if viewed from a twelve month horizon than directly for two adjacent periods, if the cohesion policy was to work.

In this approach, axes correspond to the unemployment rate vis-a-vis a national average. The horizontal envisages the "current", while the vertical serves the "next" period distributions. The contours demonstrate the relative intensity of the distributions at every point in this space⁹. In principle these shapes correspond to the relative density functions when transformed orthogonally.

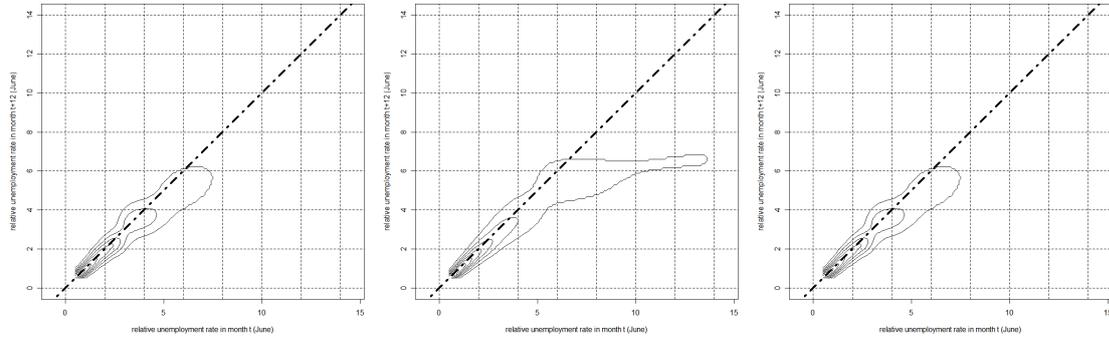


Figure 4: Kernel density estimates - Czech Republic

The graphs in Figure (4) demonstrate estimates of kernel density functions for immediate (left panel) and 12-month rolled data until 2001 and afterwards (central and right panel, respectively) for Czech Republic¹⁰. Visibly, in the case of immediate changes for the majority of the distribution no convergence pattern may be traced. The shape is located along the diagonal, demonstrating that distributions are highly stable over time. At the level of approximately 5-fold the national unemployment rate the shape moves below the diagonal which suggests that highest unemployment districts were converging to slightly lower relative unemployment rates. However, this observation may be just a statistical artefact, since even small cyclical upswings in the national unemployment rates with stable high unemployment in these districts would produce exactly the same pattern. The results are thus consistent with the analysis of β convergence findings.

In order to get a better understanding of the dynamics of these processes, transition matrices were computed. They represent essentially a non-continuous version of kernel density estimates, *i.e.* given the boundaries for the groups of observations probabilities of changing the group to a higher or lower unemployment class are estimated. This may be performed either on month-to-month dynamics or on rolled changes over longer periods of time. Correspondingly to the continuous graphical representation, Table (2) reports the estimated probabilities (left panel for the first subperiod and the right panel for post 2001 data).

⁹The actual value would have to be depicted in a three-dimensional space, but graphs obtained this way are less clear.

¹⁰Computational power of both R-CRAN and SAS are too low to enable calculations over the whole sample period. The split was chosen as to allow both subperiods to cover both increases and decreases in national levels - see Figure (1).

Table 2. Convergence of dispersions - Czech Republic

	1	2	3	4	5	6	7	8	9	10	1	2	3	4	5	6	7	8	9	10
1	72	21	4	2	0	0	0	0	0	0	74	11	6	4	4	0	0	2	0	0
2	22	35	31	11	0	0	0	0	0	0	6	38	32	21	2	0	2	0	0	0
3	4	35	28	13	15	2	0	0	0	0	7	18	29	18	22	5	0	0	0	0
4	0	9	28	24	15	13	2	0	0	0	0	4	6	19	23	13	13	0	0	0
5	0	9	11	22	26	22	7	4	0	0	2	6	20	20	17	20	6	7	2	0
6	0	0	2	19	28	9	26	11	6	0	0	0	0	11	24	28	22	6	7	0
7	0	2	0	2	9	29	33	11	13	0	0	2	0	2	16	24	31	16	9	0
8	0	0	0	4	0	24	17	24	22	9	0	0	2	0	6	6	24	28	35	0
9	0	0	0	0	0	4	11	26	30	28	0	0	0	0	9	4	34	38	15	0
10	0	0	0	0	0	0	2	9	19	70	0	0	0	0	0	0	4	13	83	0
E	12	13	12	10	9	9	9	8	8	10	7	6	10	9	12	12	11	11	12	11

Notes: Table reports the probabilities in percents. Boundaries for the decimal groups were given by 36.4%, 49.6%, 60.7%, 73.9%, 91.5%, 113.2%, 139.2%, 180.9% and 243.2% of the national unemployment rate in the case of prior to January 2001 data. For rolled 12-month transitions in the subsequent subperiod these boundaries were 40.5%, 52.9%, 66%, 79.5%, 98.9%, 123.1%, 154.2%, 197.3% and 297.7%. In either case, they were computed based on the empirical distributions in the initial period.

Numbers do not add up to unity, because they were rounded to whole percentage points.
Line *E* denotes values for ergodic vector.

Based on the figures in the left panel, one can state that so to say "downgrading" is more likely than "upgrading", since larger probabilities are found below than above diagonal. Very high and very low unemployment regions exhibit high persistence, while the middle ones tend to demonstrate higher mobility. This is especially visible when looking at the 6th and 8th decimal groups, which have virtually almost switched places. For both highest and lowest unemployment districts persistence seems to have strengthened over time, as in the right panel values on the diagonal are higher. At the same time, middle groups demonstrate even higher mobility in recent years than in the 1990s. Low values of the ergodic vector suggest high persistence, but it is visible that in the 1990s high unemployment groups were shrinking while low unemployment ones expanding, while in the 21st century the direction of these shifts is reversed. Very low unemployment regions lose approximately 3-4% of the districts, which would suggest that even the districts who performed relatively well can experience worsening of the labour market outlooks (potential convergence but to the higher unemployment rate thresholds). Nonetheless, this rate of divergence is strikingly low if one considers that decimal groups contain on average slightly less than 8 districts.

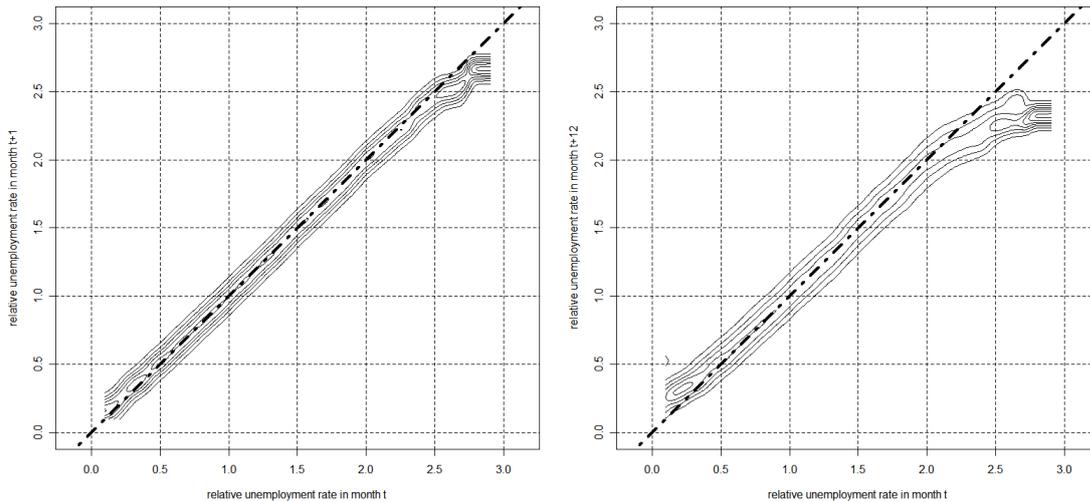


Figure 5: Kernel density estimates - Poland

The situation is very different but at the same time very similar in the case of Poland. Firstly, since the national unemployment rate is consistently higher for Poland reaching even 20% thresholds, the distribution is more condense. Instead of 15-times the average we do not observe levels higher than threefold. Nonetheless, the shape is located strongly along the diagonal with no traces of convergence/divergence for

direct transitions (left panel). In the case 12-month rolled ones (right panel), for highest unemployment regions some convergence may be traced (convergence of "clubs"). Similarly to the case of Czech Republic, higher unemployment regions tend to exhibit slower relative unemployment rates (this part of the shape is located slightly below the diagonal). However, as suggested earlier, this may result from positive trend in the national unemployment rate. Especially in the case of regions, whose unemployment rates already exceed 40% one might expect some boundaries as to how much more this rate may still increase¹¹. Therefore, although the ratio of highest to lowest relative unemployment has decreased from 25 in December 1998 to 7.5 six years later, this effect should be attributed to a general growth in national unemployment rate rather than effectively diminishing local differences.

Computing the transition matrices intuitively confirms these findings. At the beginning of calculations, there were ten groups with poviats evenly distributed. On average 93% of poviats remain in the same group on the monthly basis, while 68% are likely not to change the decimal group for rolled, 12-monthly changes. Probabilities above the diagonal are slightly higher than the ones below, suggesting that moving to higher decimal group (group of higher unemployment) is more likely. Importantly, the majority of transitions on an annual basis happens around 4th to 6th decimal groups, mostly among themselves over nine years. For high unemployment regions the probability of remaining in the same decimal group reaches almost 80% on a eight-year period.

Table 3. Convergence of dispersions - Poland

	1	2	3	4	5	6	7	8	9	10	1	2	3	4	5	6	7	8	9	10
1	97	3	0	0	0	0	0	0	0	0	86	13	1	0	0	0	0	0	0	0
2	3	92	5	0	0	0	0	0	0	0	14	64	18	4	0	0	0	0	0	0
3	0	5	88	7	0	0	0	0	0	0	0	21	54	21	3	0	0	0	0	0
4	0	0	7	87	6	0	0	0	0	0	0	1	22	50	23	4	0	0	0	0
5	0	0	0	6	87	0	6	0	0	0	0	0	1	23	50	22	0	4	0	0
6	0	0	0	7	2	88	5	0	0	0	0	0	0	3	25	48	3	21	0	0
7	0	0	0	0	4	6	90	0	0	0	0	0	0	2	27	15	56	1	0	0
8	0	0	0	0	0	0	0	91	5	5	0	0	0	0	0	0	7	62	19	18
9	0	0	0	0	5	0	0	0	92	3	0	0	0	0	0	0	1	6	72	12
10	0	0	0	0	0	0	0	0	3	97	0	0	0	0	0	1	0	0	12	88
E	9	9	9	10	11	10	11	11	11	10	8	8	8	10	11	11	10	11	12	11

Notes: Table reports the probabilities in percents. Boundaries for the decimal groups were given by 67.3%, 80.9%, 91.2%, 101.4%, 112.6%, 123.6%, 137.1%, 154.5%, and 176.7% of the national unemployment rate in the case of monthly transitions. For rolled 12-month transitions these boundaries were 68.3%, 81.3%, 91.2%, 101.2%, 112%, 123.6%, 136.9%, 154% and 176%. In either case, they were computed based on the empirical distributions in the initial period.

Line *E* denotes values for ergodic vector.

The ergodic values confirm the above statements. Namely, although the size of this effect is not very large, lower unemployment groups loose districts, while the higher ones gain. Since each decimal group had approximately 37 poviats on average are more, 1-2% differences translate to 6 to 8 districts. In addition, out-of-diagonal numbers are considerably smaller in the case of Poland, when compared to Czech Republic. This suggests that the distribution is far more stable. Graphically, this is exhibited by the thickness of the kernel density estimates - they are much thinner in for Poland.

For Slovakia the picture seems to be fairly similar to Poland over the whole period, while the regional differentials seem to be of the slightly higher range. The shape leans to the diagonal with small convergence among the highest and the lowest unemployment districts. For the former, however convergence seems to occur to relatively higher levels (shape lies above the diagonal), while for the latter the opposite seems to hold. Again, this effect should probably be attributed to the general trends, *i.e.* the markets with largest hardships relatively improve with general worsening of the labour market outlooks. At the same time, those least struggling observe some increases in relative unemployment rates in the moments of employment contraction. Finally, when compared to Poland, the shape is considerably thicker suggesting less homogeneity, thus less conformity in responding to nation-wide shocks.

This last conclusion especially is corroborated by the analysis of the transition matrices. Out-of-diagonal percentages are much higher than in the case of Poland. Moreover, especially if 12-month rolled estimates

¹¹Over the analysed time horizon Polish unemployment rate moved from 10% to 20% thresholds.

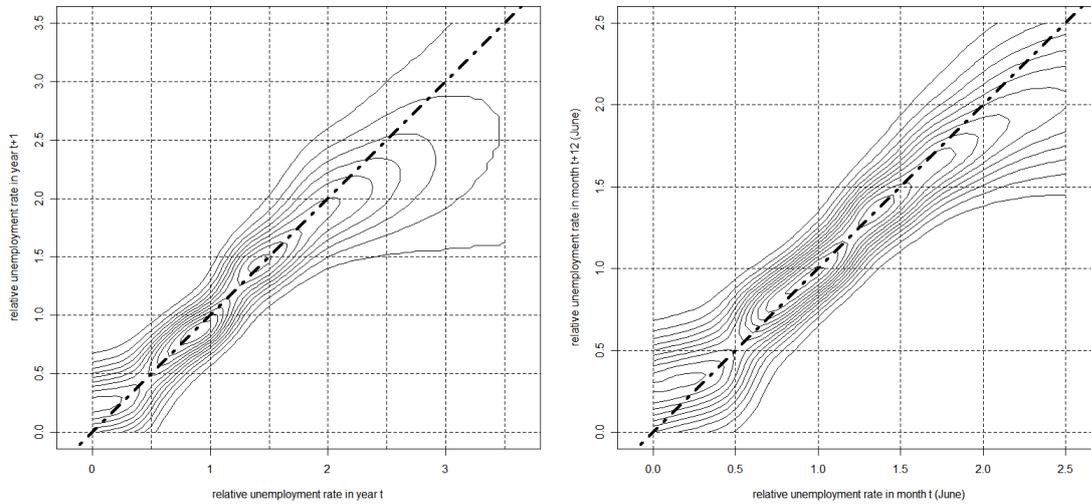


Figure 6: Kernel density estimates - Slovakia

are considered (right panel), lower unemployment decimal groups districts consistently loose, while higher unemployment ones consistently increase. Although, again as in the case of Czech Republic, decimal groups contained on average only few districts, 11% to 15% transition for the 10th group suggests that over the 1997-2004 time span, this group grew from approximately 8 to 12 districts, which is by all means considerable. In addition, there seems to be a lot of rotation in the middle-range groups, with diagonal values of approximately 50%. This effect can stem from two phenomena - either middle range groups experience a lot of volatility or their unemployment rates remain fairly stable over the nation-wide range. For groups 8 and 9 the latter seems to be the case, while for groups 4 to 7 it is rather the former.

Table 4. Convergence of dispersions - Slovakia

	1	2	3	4	5	6	7	8	9	10	1	2	3	4	5	6	7	8	9	10
1	97	3	0	0	0	0	0	0	0	0	93	7	0	0	0	0	0	0	0	0
2	3	89	7	0	0	0	0	0	0	0	13	67	16	4	0	0	0	0	0	0
3	0	3	83	8	0	0	0	0	0	0	0	20	61	16	4	0	0	0	0	0
4	0	0	9	78	12	0	0	0	0	0	0	2	13	52	27	7	0	0	0	0
5	0	0	0	11	77	10	1	0	0	0	0	2	22	38	25	5	4	0	0	0
6	0	0	0	1	11	77	10	1	0	0	0	0	0	2	31	44	20	4	0	0
7	0	0	0	0	0	10	80	8	1	0	0	0	0	0	5	21	46	23	4	0
8	0	0	0	1	0	1	9	79	10	1	0	0	0	0	4	22	48	22	4	0
9	0	0	0	0	0	0	1	10	80	8	0	0	0	0	0	4	25	42	29	0
10	0	0	0	0	0	0	0	0	7	93	0	0	0	0	0	0	0	0	25	75
E	11	10	10	10	10	10	10	10	10	11	13	8	6	7	9	9	10	11	11	15

Notes: Table reports the probabilities in percents. Boundaries for the decimal groups were given by 43.5%, 70.9%, 83.2%, 93.0%, 103.7%, 117.8%, 132.7%, 146.7%, and 166.5% of the national unemployment rate in the case of monthly transitions. For rolled 12-month transitions these boundaries were 42.8%, 70.4%, 82.8%, 92.5%, 103.9%, 115.9%, 132.3%, 145.3% and 166.1%. In either case, they were computed based on the empirical distributions in the initial period.

Line E denotes values for ergodic vector.

Summarising, one can compare this analysis to the following exercise: considering the ranking of the districts along their relative unemployment rates we tried to inquire whether they switch places in the ranking, like the steps in the ladder. We already know from the β analysis that this ladder - if anything - gets wider in terms of unemployment levels. σ analysis inquired the mobility of local labour markets in nation-wide distributions. In the case of Czech Republic transitions seem to be more frequent, but at the same time less sustainable - movements up and down the ladder occur for the same districts. For Poland, there appear to be virtually no movements - if anything, *poviats* move to higher unemployment levels. For Slovakia analysis suggests reduction of low unemployment clubs and considerable volatility in the middle

range. In all three cases, however, kernel density estimates failed to provide evidence in favour of general convergence.

Unfortunately, σ convergence analysis does not permit to assess the absolute scale of the differentials persistence. This shortcoming of the kernel density estimates comes directly from their nature - one needs distributions (relative unemployment levels) to estimate them. To see how these differentials behave across time, stochastic convergence analysis is applied.

4.3 Stochastic convergence

As discussed earlier, stochastic convergence essentially implies that one should confirm random walk hypothesis in the univariate time series analysis in the panel context. In order not to exclude scenario B from Figure (2), one can impose weaker constraint of trend stationarity with a constant to account for potentially differentiated steady state levels. In order to assure validity of the results one needs to control for sufficient number of lags. In as far as number of lags is concerned, we followed the findings of Bayer and Juessen (2006), who typically found two to maximum four lags on annual data. Hence, we universally imposed 36 monthly lags. In most cases up to 8 lags was supported by data. Table 5 below reports the results of this analysis.

Table 5. Stochastic convergence

	Czech Republic	Poland	Slovakia
Total number of districts	78	374 (428)	77
No of observations	11 550	32 579	7 426
Multivariate ADF (MADF)	0.00	0.00	0.00
Fisher (Phillips-Perron)	0.00	0.00	0.00
Fisher (ADF)	0.78	1.00	0.37
No. of null rejections at 5%	45	71	45
LLC for panel time-balanced	0.00	0.26	1.00
No of observations	10 472 (76)	24 500 (370)	4606 (73)
Trend	No	Yes (2)	Yes (2)
Structural breaks	No	No	No
LLC for panel unit-balanced	0.00	1.00	1.00
No of observations	11 324 (77)	26 305 (379)	5451 (79)
Trend	No	No	No
Structural breaks	No	No	No
IPS for panel time-balanced	0.049	1.00	0.334
No of observations	10 716 (76)	24 420 (370)	6205 (73)
Trend	No	Yes (2)	Yes (2)
Structural breaks	Yes (4)	Yes (2)	Yes (2)
IPS for panel unit-balanced	0.014	0.469	0.872
No of observations	9 856 (77)	16 297 (379)	6715 (79)
Trend	No	No	No
Structural breaks	Yes (4)	Yes (2)	Yes (2)
Non-stationarity	Not rejected	Not rejected	Not rejected

Notes: Optimal number of lags obtained by sequential $t - tests$ as suggested Ng and Perron (1995). Structural breaks forced on data based on the analysis of national unemployment rate behaviour. Indication "yes" with reference to trend and structural breaks implies that specification with trend no longer failed to reject the null of non-stationarity.

To obtain time-balanced panels periods for which not all units are yet available were eliminated (hence, maximum possible time series length). To obtain unit-balanced panels, units for which data is not available for all periods were eliminated (hence, maximum possible abundance of districts). MacKinnon $p - values$ reported.

We first report multivariate augmented Dickey-Fuller (MADF) panel unit root test (as specified by Sarno and Taylor (1998)) on a variable that contains both cross-section and time-series components. The MADF test is a generalization of the test in which a single autoregressive parameter is estimated over the panel¹².

¹²In contrast, it allows for higher order serial correlation in the series and allow the sum of autoregressive coefficients to vary across panel units under the alternative hypothesis. This test involves verifying for each equation if the sum of the coefficients of the autoregressive polynomial is unity. The null hypothesis consists of the joint test that this condition is satisfied over the N equations. Under the null hypothesis, all of the series under consideration are realizations of nonstationary stochastic processes. The test's null hypothesis should be carefully considered. It will be violated if even one of the series in the panel is stationary. A rejection should thus not be taken to indicate that each of the series is stationary.

Findings suggest that the null hypothesis is strongly rejected. Similarly, Fisher’s test as developed by Maddala and Wu (1999) does not require a balanced panel. Combining the p – values from N independent unit root tests, Fisher’s test assumes that all series are non-stationary under the null hypothesis against the alternative that at least one series in the panel is stationary. The results of ADF version of this test clearly demonstrate that the null hypothesis of non-stationarity cannot be rejected for all these three countries. More importantly, in the Phillips-Perron version of this test, when the null hypothesis is that all variable in the panel contain a unit root, and the alternative is that at least one of the variables in the panel was generated by a stationary process, we find strong rejection of the null in the case of all three countries, which suggests that regions are strongly diversified in the underlying dynamics. This is further confirmed if one analyses the number or cases in which null was rejected. Approximately two thirds of Czech and Slovak districts exhibit stationarity, while in the case of Poland this share drops to as low as approximately 20%.

As frequently raised, Fisher-type test may have too little power to effectively reject the non-stationarity in all relevant cases. This is why reportedly more powerful IPS and LLC tests were applied as well. Unfortunately, this had to come at the expense of data reductions, since these tests require balanced panels¹³. Results seem to be consistent with the unbalanced tests outcomes. However, in the case of Czech Republic the null was consistently rejected, while in the case of Slovakia and Poland data suggest strong persistence of regional unemployment rate differentials despite inclusion of trend and allowing for structural breaks.

5 Conclusions and suggestions for further research

The main purpose of this paper was to inquire the convergence patterns of local labour markets in three transition economies: Czech Republic, Poland and Slovakia. The first of the three can be characterised as relatively low unemployment environment, while Slovakia and Poland experienced over the time span very high unemployment levels. We used policy relevant NUTS4 level data, since in all three countries actual labour market policies - with special emphasis on the active ones - are performed at exactly this level. Time span in this study allows to cover both up and down cycles in labour market conditions, which guarantees that the results are not trend driven. Unfortunately, in each of these countries sample commences already some years after the transition, which makes it impossible to establish a direct link between transition and local unemployment rate dynamics. On the other hand, our findings suggest that whenever job prospects get better all the way through the country, already disadvantaged regions benefit less in each of the examined countries.

Results in each of the countries suggest no support in favour of β -type convergence, *i.e.* convergence of levels. Even controlling for nation-wide labour market outlooks (conditional convergence) does not provide any support to this hypothesis. Further, regions with both very high and very low unemployment show signs of high persistence and low mobility in the national distribution, while the middle ones tend to demonstrate higher mobility and essentially no regional unemployment differentials persistence. This is also confirmed by σ -convergence analysis. For Czech Republic transitions seem to be more frequent, but at the same time less sustainable, while movements up and down the ladder occur frequently for the same districts. For Poland and to some extent Slovakia, there appear to be virtually no movements - if anything, districts move to higher unemployment levels.

Returning to the scenarios discussed in the opening of this paper, it seems that this paper provides

¹³LLC imposes a single autoregressive parameter over all units in the panel but utilizes a variant of fixed effect panel estimation. This test may be viewed as an Augmented Dickey-Fuller (ADF) test when lags are included, with the null hypothesis that of nonstationarity. IPS in turn estimates the t – test for unit roots in heterogeneous panels. It allows for individual effects, time trends, and common time effects. Based on the mean of the individual Dickey-Fuller t – statistics of each unit in the panel, the IPS test assumes that all series are non-stationary under the null hypothesis. Lags of the dependent variable may be introduced to allow for serial correlation in the errors. Unlike the LLC test, which assumes that all series are stationary under the alternative, IPS is consistent under the alternative that only a fraction of the series are stationary.

evidence in support of stylised patterns presented in Figure (7). In the case of Czech Republic, rejection of non-stationarity is the weakest. This may follow from both relatively lowest unemployment levels and relatively high proportion of districts not experiencing labour market hardships. There are quite a few districts in case of which non-stationarity is associated with significantly lower than average unemployment, which suggests that differentials persistence found is generally a positive sign. At the same time, some districts with generally more profound labour market problems seem to remain at higher thresholds, with relatively high stability in the composition of the 10th decimal group. In the second subsample (time span 2001-2007), diagonal values are extremely low (ranging between 17% and 38% for 2nd to 8th decimal groups), which suggests there is relatively high mobility in the middle range allowing up and down grading even within four closest groups. While this mobility should be rather attributed to the national unemployment rate movements within 2% to 15% boundaries, Czech districts seem to maintain stability in absolute levels.

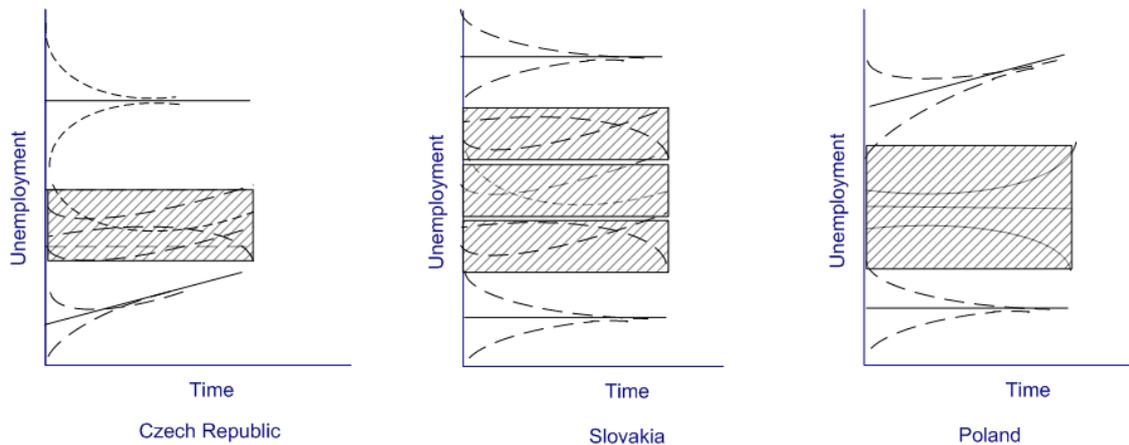


Figure 7: Stylised facts based on findings

The picture seems different for the two countries with with relatively more difficult labour market situation. In the case of both Slovakia and Poland non-stationarity was strongly rejected. For the latter this is true for mainly high unemployment districts, while for the former low unemployment ones demonstrate high differential persistence as well. At the same time, Slovakia seems to demonstrate higher mobility in the mid-range, while for Polish districts diagonal values are higher and off-diagonal values considerably lower, especially for 12-month rolled analysis. Over the analysed period in both these countries highest unemployment regions demonstrate convergence of "clubs". Although the fact that shape is located slightly below the diagonal seems to suggest relatively lower relative unemployment rates, this effect should be attributed to the fact that these levels have decreased significantly in time (national averages moving from 10% to 20% thresholds in the case of Poland and from 12% to 20% for Slovakia).

There are some evident shortcomings of our study, though. Firstly, due to data limitations it was not possible to cover the whole transition period. The relevant district data for earlier years do not exist or have too low quality. Therefore, the time-span is relatively short, especially in the context of stochastic convergence studies in the literature (Bayer and Juessen (2006) use 40 years for Western Germany, Gomes and da Silva (2006) have at disposal 22 years, while Camarero, Carrion-i Silvestre and Tamarit (2006) study the validity of the hysteresis hypothesis with yearly unemployment rates data from 19 OECD countries for the period between 1956 and 2001). Consequently, our results should be interpreted with caution.

At the same time, in search of integrity with actual policy developments, data used are desagregated to NUTS4 level. The findings of this paper effectively suggest that the very notion of "national" unemployment rate is highly uninformative for these countries. Namely, the average is actually only a statistical operation on strongly differentiated processes with sometimes even diverging dynamics. Consequently, however, computations of adjustment speed could not be undertaken in a meaningful way. With potentially 80 different regional evolutions for Czech and Slovak republics and 350 for Poland, obtaining informative and

statistically robust results seems virtually impossible.

What this study did not inquire includes as well other than NUTS4 administrative borders. Namely, NUTS4 regions have supreme authorities on NUTS3 level in Czech and Slovak republics and NUTS2 in Poland. Inquiring whether local (NUTS4) units demonstrate convergence within regional (respectively NUTS3 or NUTS2) units could provide some evidence with reference to geographical clustering of relatively more troubled and relatively more favoured areas. It would be also potentially interesting to perform conditional σ -convergence analysis taking into account either national business cycle trends or unemployment structure. It is evident that local labour markets differ in terms of sectoral composition (agriculture, industry and services) as well as the quality and the range of skill mismatch.

This paper has also some important policy implications. Namely, NUTS3 authorities in Slovak and Czech republics and NUTS2 in Poland do not seem to use the fact that they distribute the active labour market policies financing in an effective way. Each of Polish NUTS2 and most of Czech and Slovak NUTS3 regions contain districts from highest unemployment groups. Financing should be geared towards alleviating the situation in most deprived regions by fostering higher effectiveness. Also, national authorities do not seem to exert sufficient monitoring activities promoting improvements in most deprived regions.

The analysis of standard deviations in all three countries suggests that whenever job prospects get worse all the way through the country, the conditions hit disadvantaged districts more. Furthermore, since β -divergence was strongly confirmed for all three countries even controlling for initial distributions, these results can actually be quite useful in evaluating the efforts to increase the degree of social cohesion throughout the transition process. Although rather indirectly, this research demonstrates that either these activities lack necessary effectiveness, or are largely inappropriate. To inquire this issue in-depth a theoretical framework would need to be developed.

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FACULTY OF ECONOMIC SCIENCES
UNIVERSITY OF WARSAW
44/50 DŁUGA ST.
00-241 WARSAW
WWW.WNE.UW.EDU.PL