Regional Dynamics of Unemployment in Poland - A Convergence Approach

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Abstract

In this paper we approach the regional unemployment dynamics in Poland. Using policy relevant NUTS4 level data from 1999 to 2006, we employ tools typically applied to income convergence analyses to inquire the patterns of unemployment distribution. We apply diverse analytical techniques to seek traces of convergence, including β and σ convergence as well as pass-through analysis.

We demonstrate that the unemployment rate distribution is highly stable over time, while only weak 'convergence of clubs' is supported by the data and only for the high unemployment regions. Results 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 those in the middle tend to demonstrate higher mobility and essentially no persistence of regional unemployment differentials. This diagnosis is confirmed by σ-convergence analysis which indicates no general divergence or convergence patterns. Transitions seem to be slightly more frequent, but at the same time less sustainable for middle range districts, while movements up and down the ladder occur predominantly for the same districts.

This methodology allows to define the patterns of local labour market dynamics, pointing to differentiated divergence paths. Importantly, these tendencies prevail despite cohesion financing schemes, which allocate relatively more resources to deprived regions.

Key words: regional unemployment rate differentials, convergence analysis, Poland

JEL Codes: J43, R23, R58, E64, J18

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1 Introduction

There are at least two main motivations to inquire the dynamics of local labour markets. First, most macro-level models implicitly assume homogeneity and symmetry of shock response at least within countries (for example, national average is informative in a sense that it describes a fairly homogenous process). Typically, it is the international heterogeneity and asymmetry of shock response that receives most of the explicit scholarly attention. Do we know that for a fact, or do we assume that for a lack of better options?

In the case of EU for example, studies find income convergence between nations and divergence on more disaggregated levels, both within and across countries, eg. Egger and Pfaffermayr (2005) or Paas and Schlitte (2007). If the same held for income dynamics, any analysis of real convergence would be under the imperative to use regional instead of national data. Having an explicit verification of whether the same holds for wages or employment outlooks would provide insights into the relevance of the homogeneity assumption. This problem seems to have received more attention in recent years. For example, Armstrong and Taylor (2000) argue that one should focus on the problem of adjustment speed as well as persistence of potential unemployment differentials, instead of aggregate convergence *per se*, since some effects may accumulate over time, effectively altering the direction of adjustments. Further on the theoretical grounds, Boeri and Terrell (2002) inquire whether these differentials could be explained on the grounds of optimal transition speed theory (Ferragina and Pastore (2008) provide an extensive review of this issue).

The second reason is more rooted in the policy choice area. Within Europe, cohesion and catching up of the regions lagging behind are not only one of the main policy objectives but also a constituent expression of Community values. These values are frequently transferred to national levels, where cohesion, equal access and convergence receive attention both explicitly in constitutions and laws and implicitly in financing algorithms. For example, in Poland any labour market policy financing is distributed with preference to areas with an above-average unemployment rate, above-average share of long-term unemployed and above-average number of unemployed. Consequently, regions facing relative hardships receive more resources to alleviate their impact. Do we observe any impact of these policy measures?

These two questions constitute the main motivation for choosing this area of analysis. The choice of Poland is dictated by the fact that this country is rather exceptional within EU. Over the past two decades unemployment rates have already swung twice between 10% and 20% thresholds. Despite recent improvements, the situation of the Polish labour market has been extremely difficult over the past years, with the unemployment rates consistently above 16-18% thresholds, while the chances of continuing to long-term unemployment still exceeded 50%. Reported demographic data exhibits that approximately 700 000 young women (roughly 5% of labour force) have no or negligible work experience, with gender employment gap among young and 50+ women double the EU average. At the same time labour activity in the 50+ group is among the lowest in the whole EU oscillating around 26%.

In this paper we apply β- and σ-convergence analysis to the registered unemployment data covering the 1999-2007 period. For the former - convergence of levels - parametric econometric estimation techniques were applied, with the main innovation of demonstrating the dynamic properties of estimated β coefficients. To the latter problem - that of distribution stability - the nonparametric technique of conditional kernel density estimates was applied. This tool is traditionally applied to the problems of income convergence and we are not aware of any study that has employed it to local level unemployment data[1]. Nonetheless, there seem to be virtually no methodological arguments against its application, while it has many advantages from the interpretational point of view. The basic research question behind this study is to inquire the dynamics of local labour market evolutions and test whether they exhibit any convergence/divergence patterns in levels (β) and in dispersion (σ). Taking into account the challenges implied by very high national unemployment rates as well as the still continuing transition process we inquire, whether any regional difference in development patterns may be observed. More explicitly, we want to test the hypothesis of weather any traces of differentiated response to adverse movements in the labour market may be observed.

NUTS4 level data were chosen, since this is the policy-relevant level of analysis. Another policy-relevant level (a level at which authorities exist and policies are implemented) is NUTS2 in the case of Poland (NUTS3 and NUTS1 are only statistical units). However, as of 1999 and the so-called decentralisation reform - labour market policy is designed, implemented and evaluated at NUTS4 level, while NUTS2 units have neither passive nor active instruments at their disposal. Indeed, NUTS2 units are so big and heterogeneous that only slight differentiations of unemployment rates may be observed (lowest to highest ratio amounts to only 1:1.5 at NUTS2 and as much as 1:25 at NUTS4). Therefore, instead of inquiring the convergence between NUTS2 units, we focused on the processes within each of them. The period we chose captures the so-called 'second wave of unemployment' commencing in 2001 as well as introducing ALMPs on a relatively comprehensive scale (from less than 10% up to 25% of unemployed subjected to active instruments). These two occurrences are convenient in a sense that prior to 1999 data on NUTS4 level do not exist in a consistent manner.

This paper is organised as follows. We briefly discuss the literature, subsequently proceeding to describing the methodology in section 3. Section 4 covers data. Section 5 presents the main findings with reference to distribution dynamics, while Section 6 focuses on the main findings of β-convergence analysis. Section 7 concludes.

2 Literature review

Unemployment convergence at regional level has received a lot of attention from the academia. Buettner (2007) compares empirical evidence on regional labor market flexibility in Europe (but uses different aggregation levels for different countries, which makes the results weaker). Marelli (2004) as well as Huber (2007) provide an overview of similarities and disparities across European Union regions. In particular, it seems that CEE countries exhibit higher regional wage flexibility (Buettner 2007). At the same time, despite phenomenal migrations emerging after 2004, labour mobility is still assessed to be low (Kaczmarczyk and Tyrowicz 2008), while Fihel (2004) demonstrates that effectively in the local scale unemployment is not significant as pushing factor (these issues have been surveyed, among others, by Huber (2007)). In the case of CEECs, the role of transition processes may indeed still be significant, (Svejnar 2002b).

In Poland, the employment restructuring process consisted mainly in the reductions in employment with growing average job tenure as well as average time spent in unemployment or inactivity (Svejnar

[1]Kernel density estimates (KDE) were employed, among others, by Bianchi and Zoega (1999), Lopez-Bazo, del Barrio and Artis (2002) and Lopez-Bazo, del Barrio and Artis (2005). In fact, our approach differs significantly in that we have KDE conditional on the distribution (again, KDE) in the previous period, which makes this technique so suitable for analysing σ-convergence.

2002a). Dismissals - if compensated at all - were geared towards hiring of young, better educated workers, but the youth unemployment rate for a long time continued to be the highest in Europe (as well as across age groups in Poland). People who lost their employment usually became permanently unemployed or inactive (Grotkowska 2006) with currently less than 13% of the unemployed still retaining the right to benefits[2], thus suggesting that most of the unemployed are either long-term unemployed or have a long record of unstable short-term employment.

Consequently, on an individual level it is usually easy to identify the ideal type of winners and losers in the transition process. However, in terms of regional analysis, the 'conventional wisdom' of Eastern Poland generally lagging behind finds no support in data, while some of the highest unemployment regions are located relatively close to the 'growth poles', (Gorzelak 1996), which stays in contrast to the categorisation suggested by the previous literature. For example, Scarpetta and Huber (1995) construct a measure that captures both the degree of economic development and the structure of industry in a single index - economic development is proxied by an index of industrial diversification, where regions are classified in six groups: I - developed agricultural; II - other agricultural; III - developed heavily industrialised; IV - other industrialised; V - developed diversified; VI - other diversified. They naturally find, that modern voivodships are hit least, while in the case of industrial/agricultural ones it depends mostly on the type of industry.

In a similar vein, Góra and Lehman (1995) classify voivodships by the degree of economic development of a region, but build on employment shares of services and industry in 1990, relative change in total employment and that of employment in services and the relative per capita income of municipalities in 1992. In addition, Lehmann and Walsh (1998) build an economic classification of voivodships with an intention to produce an index reflecting the degree of employment restructuring with the use of seven indicators: share of services in employment; share of short-tenured men (i.e. with tenure less than ten years) in total male employment; number of telephones per capita; voivodship shares of domestic and direct foreign investment, normalised on population; share of construction in total employment; and share of agriculture in total employment. However, although these indices correlate reasonably well among each other, correlation with voivodship unemployment rates is highly unsatisfactory (Newell and Pastore 1999). Moreover, these findings no longer hold if one disaggregated to NUTS4 level (units with on average over 6 000 unemployed in the labour market distress period and over 40 000 in the labour force).

Taking a different perspective, Newell and Pastore (1999) argue that it is the hazard of job loss differentiating for employees with longer tenure that drives the regional differences, but these findings cover the 1995-1999 time span (a period of gradual improvement in both economic and labour market outlooks) and are no longer consistent with more recent (2001-2005) developments (Grotkowska 2006), namely, explanatory power of these hazard differentials disappears with the

[2]In Poland the entitlement to unemployment benefit is temporary and lasts only 12 months after the registration (18 months in regions with more labour market hardships). After this period, unemployment benefits may be replaced with social assistance benefit (which is lower and based on family income rather than labour market status). The entitlement to unemployment benefit is re-established for an unemployed who obtains legal employment for a period above 6 months. No publicly available statistics report the effective share of long-term unemployed or unemployment duration. For example, if one does not confirm 'willingness to undertake employment', one is de-listed from the unemployment registries. However, 3 months after this occurrence, one may register again (the basic incentive is free access to public health care system for the unemployed and his/her family) and then, the unemployment tenure is calculated from the scratch. However, benefit entitlements are not. Consequently, the share of unemployed still retaining the right to the benefit is a reliable measure of actual rate of long-term unemployment.

general growth of unemployment rate. Thus, the persistence of high unemployment rate regional differentials remains as intriguing as the persistence of high unemployment itself.

In the empirical literature of unemployment rate characteristics, one can find a number of differentiated approaches towards the unemployment rate dynamics and persistence as well as distribution (*cfr.* Decressin and Fatas (1995), Obstfeld and Peri (1998) or more recently Armstrong and Taylor (2000)). Perugini, Polinori and Signorelli (2005) use NUTS2 level data and inquire the regional differentiation of Poland and Italy. Marelli (2004) focuses on specialisation for NUTS2 EU regions with tripartite desaggregation (industrial, agricultural and service sectors) reaching the conclusion that convergence in economic structures occurs, while income does not. However, Marelli (2004) analyses predominantly income and economic convergence and not explicitly the underlying fundamentals (like, for example, labour market performance).

Suggesting a different angle, Bayer and Juessen (n.d.) perform a unit-root test on regional unemployment rate differentials using Mikrozensus data for West Germany between 1960 and 2002. By differentiating between the theoretically motivated imperative of convergence itself (Blanchard and Katz 1992) and the speed of adjustment (as argued by Armstrong and Taylor (2000)) they focus on the concept of stochastic convergence (Carlino and Mills 1993)[3]. In this framework, convergence is present only if shocks to the unemployment differential are temporary, thus erasing disparities between regions, providing a testable hypothesis of regional and national unemployment rate cointegration. Bayer and Juessen (n.d.) find moderate evidence in support of the convergence hypothesis. A similar technique has been applied by Gomes and da Silva (2006) for the regions of Brazil, finding strong evidence of hysteresis and the persistence of regional unemployment differentials.

However, one can put forward a strong argument against these results, namely that stationarity of the regional unemployment rate differentials can happen under both convergence and divergence scenarios, let alone trend stationarity. Notably, with some regularity in the cycles, unemployment rate differentials can positively pass the unit-root test even if real differentials are growing (some regions still suffering harder during the crisis and recovering less with the good economic outlooks). Thus, in this paper a different approach is followed, namely we analyse the conditional density functions with kernel estimates, assessing the changes in each region's position in the nation-wide unemployment rate distribution. Bianchi and Zoega (1999) use non-parametric kernel density methods to test the hypothesis of multimodality in regional unemployment rate distribution across counties in the UK, thus analysing the patterns of variance. They found that regional transition probabilities are similar for both high and low unemployment counties with the persistence of 97%.

3 Methodology

The 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 a continuous equivalent of the histogram. At each point the values of the density function are calculated

[3]Testable hypothesis of local and national unemployment rates cointegration can be formulated as $\forall t$: $\lim_{s\to\infty} E(U_{i,t+s} - U_{j,t+s}|I_t) = constant$, where $U_{i,t}$ denotes respective unemployment rates and I_t is the conditioning information set. To be precise, this is a conditional stochastic convergence formula. Unconditional versions would require the limit to approach 0. However, such a condition would discriminate between dispersion convergence scenarios to differentiated levels (so called convergence of clubs) classifying it as non-convergence. Allowing a non-zero constant, permits to account for regional differentiation. This is empirically approached by testing for a unit root in $u_{i,t} = lnU_{i,t} - ln\bar{U_t}$, where $\bar{U_t}$ is the corresponding national average.

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 an evident yet 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 smoothening. Too wide a bandwidth window will hide the real data distribution, while a too narrow one might misleadingly result in a function with multiple vertices - not necessarily corresponding to the 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 is $x+1$, the distribution of $x+1$ conditional on x may be written down as:

$$f[x+1|x] = \frac{f[x, x+1]}{f_x[x]}, \tag{1}$$

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 the numerator and the denominator of 1 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.:

$$\hat{f_x^A}[x] = \frac{1}{n}\sum_{i=1}^{N} \frac{1}{h_x w_i} K \frac{x - x_i}{h_x w_i}, \tag{2}$$

where n is the number of observations, h_x is the bandwidth window for the initial unemployment rate and $K[.]$ represents the kernel function[4]. In 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 1, is thus estimated by:

$$\hat{f_{x_t,x_{t+1}}^A}[x] = \frac{1}{n}\sum_{i=1}^{N} \frac{1}{h_x h_{x+1} w_i^2} K \frac{x-(x+1)_i}{h_x w_i} K\left[\frac{(x+1)-(x+1)_i}{h_{x+1} w_i}\right], \tag{3}$$

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

Importantly, in 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 differentiations of the bandwidth windows are calculated according to:

$$w_i = \left(\frac{1}{n}\sum_{i=1}^{N} \frac{1}{h_x h_{x+1} w_i^2} K\left[\frac{x - x_i}{h_x w_i}\right] K\left[\frac{(x+1)-(x+1)_i}{h_{x+1} w_i}\right]\right)^{1/2}. \tag{4}$$

[4]With the large number of observations (over 400 units for Poland) 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.

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 window[5], while the numerator gives the geometric average of this estimator for matching couples of both variables. The final conditional density function is found using the weights from equation 4 in evaluating the equations 2 and 3 (calculating their quotient), according to equation 1[6].

This methodology has shorthand interpretative advantages. First of all, convergence/divergence may be easily detected from the graphs of the conditional density functions. Namely, a vertical shape of this function suggests divergence, while the horizontal alignment is consistent with the convergence hypothesis. If the conditional density function follows the 45° line, an 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. Stability implies directly that neither divergence nor convergence of distribution can be tracked.

4 Data

In the paper monthly data covering the period from January 1999 to August 2007 were used at the lowest available administration level of *poviats*[7]. However, the choice was only marginally affected by data availability, with the main reason being the fact that labour market policy is actually performed on this level exactly. At the same time, this period covers the so-called 'second wave of unemployment', commencing with the economic slowdown from the end of 2001 onwards as well as the recovery period of 2005-2007 which allows us to explore the symmetry of response to macroeconomic changes on a local level. Figure 1 demonstrates the unemployment developments in Poland over this period.

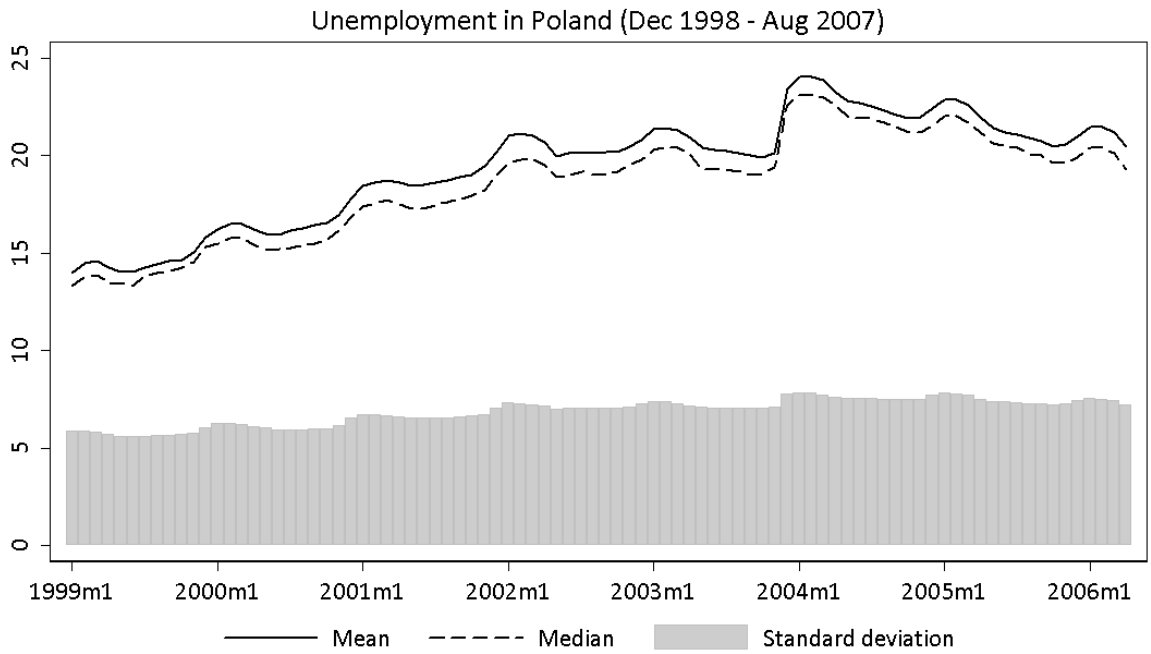

Figure 1: Unemployment rate in Poland (1999-2007), Source: registry data, ML&SA

[5]Fixed window kernel estimate.

[6]An approach similar to ours was taken by Overman and Puga (2002) with the main difference that they consider two distinct points in time - namely 1986 and 1996 - for NUTS2 level EU regions.

[7]Due to the administrative changes in Poland in 1999 no prior data are available at NUTS4 level.

In this paper we employ policy relevant NUTS4 level unemployment data using official registry data for Poland. In total we use 374 units[8]. These are registry data, which implies they suffer from many well-known shortcomings, including underreporting or overreporting (*e.g.* either due to forced passivity or in order to gain access to social transfers, respectively). Unfortunately, LFS data can only be reliably disaggregated to the NUTS2 level.

Observing Figure 1 one sees a significant increase in the unemployment rate in December 2003. As of January 2004 new census data from 2002 were applied to calculate the size of the labour force. Thus, although the above unemployment rates are bases on registered unemployment recorded by local Public Employment Services (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 Central Statistical Office 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.

The choice of time boundaries was dictated by the data availability and seems to bear no serious limitations for the possible results except one. Namely, labour market evolutions have commenced in Poland in early 1990s. Unfortunately, NUTS4 data do not exist prior to 1999, while separate metropolitan municipalities were established only in 2001. Hence, although this paper inquires the dynamics by testing β, σ and stochastic convergence, the data analysed commence roughly in the middle of the dynamic evolution patterns. Nonetheless, the data set covers periods of both increases and decreases in the national unemployment rates as depicted by Figure 1.

The distribution seems quite volatile since the beginning of 1999, with obvious seasonal fluctuations of the maximum unemployment rate. Over the whole period the average has been larger than the median indicating that generally poviats with higher unemployment rate are larger (national average is population weighted), which is depicted by Figure 1. This is an important observation, since generally municipal units are larger than rural ones, but at the same time they typically experience better labour market outlooks (large cities). Consequently, these are larger non-urban local labour markets, which drive this result, suggesting that their employment prospects may indeed be dramatic.

More importantly, as can be inferred from Figure 1, the dispersion of the unemployment rates has been constantly growing over the entire time span - especially in the down cycles, be they seasonal effects or general trends in the labour market evolution (the solid line demonstrates the non-weighted average standard deviation for the whole period). This observation suggests that whenever job prospects worsen in general throughout the country, more deprived regions are hit harder. On the other hand, although rather worrying as a labour market phenomenon, this is rather fortunate from the 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-directional trends.

[8] An administrative 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 onwards. Consequently, prior to 2001 for some districts data cover both municipal and rural areas, while after 2001 in each of these cases two districts were formed instead of one, with two separate unemployment rates reported. Since units comprising cities and rural areas were divided to two separate *poviats*, each with a different labour market structure and potential. Therefore, subsequent to the change, both these units are treated as new in our sample.

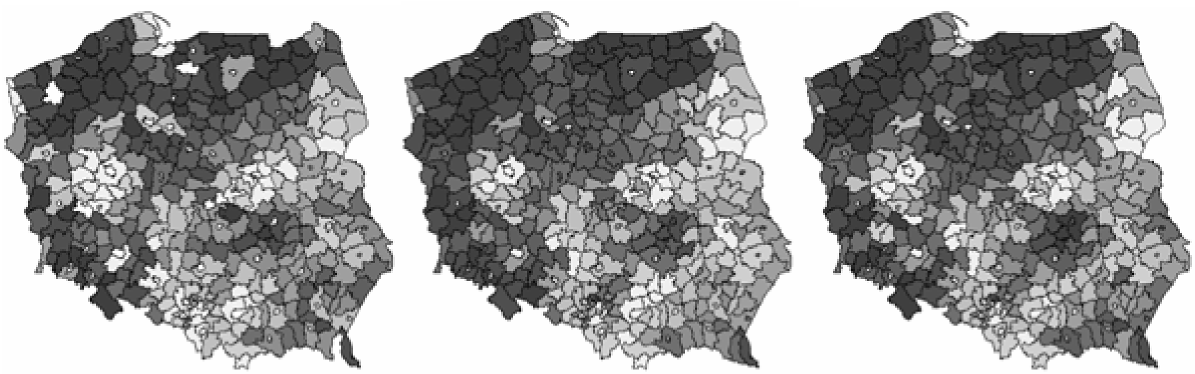

Figure 2: Unemployment in Polish *poviats* (Dec 1998 in left panel, Dec 2002 in the middle and Dec 2006 in the right panel, the darker the shade, the higher relative unemployment rate). Source: registry data, ML&SA

The maps in Figure 2 demonstrate December unemployment rates on a *poviat* level for the 1998, 2002 and 2006, with the shades darkening with the relative labour market hardships. Data demonstrates that the discrepancies at the regional level are even 25-fold (from 0.11 of the 50% percentile to 2.8 of this value in December 1998).

5 Results - distribution dynamics

The analysis of σ-convergence - as covered in Section 3 - allows to inquire the dynamics of the local unemployment rates distribution. In principle, this analysis may be treated as observing the 'ranking' of *poviats* at each point in time and verifying, whether a position in this ranking (measured by the relative distance to the average) changes or not with respect to previous period ranking. In other words, if all *poviats* were moving towards the average, one would expect a horizontal alignment of the resulting contour plot of the conditional density function (in the 'ranking' the relative distance between the lowest and the highest is shrinking). If *poviats* are moving away from the average, one would observe a vertical shape (the relative distance is growing).

Figure (3) presents contour plots of the density functions showing distribution dynamics for relative unemployment rates in *poviats* over the whole period for which data is available (December 1998 to August 2007) - monthly changes on the left panel and yearly rolling [9] in the right panel. These figures are a two-dimensional depiction of the distribution of the current relative unemployment rate (vertical axis) conditioned on the relative unemployment rate in previous period (horizontal axis). Monthly relative unemployment rates seem to be very stable (the shape is positioned along the diagonal, which suggests that only small changes in unemployment occur on a monthly basis). For the highest relative unemployment rate (2.5-3.0 of the average) the shape lies slightly below the diagonal which suggests that highest unemployment rates were slightly decreasing from month to month - although they are still around 2.5 to 2.7 of the average.

The yearly relative unemployment rate (right panel) shows that more changes occur on yearly basis than on a monthly basis (the shape is thicker), but unemployment is still quite stable (shape is mainly positioned along the diagonal). However there are two peaks on the opposite ends of the figure that seem to position more along the horizontal axes. This suggests that separately the

[9] Yearly rolling change means a change in a given month with respect to the same month in previous year for all months over subsequent years.

poviats with the highest unemployment rates (above 2.5 times the average) and those with the lowest unemployment rates (below 0.25 times the average) are becoming similar, so there is an indication of convergence of highest and lowest unemployment *poviats* separately. Therefore - if any - convergence of clubs may be observed for highest unemployment *poviats*.

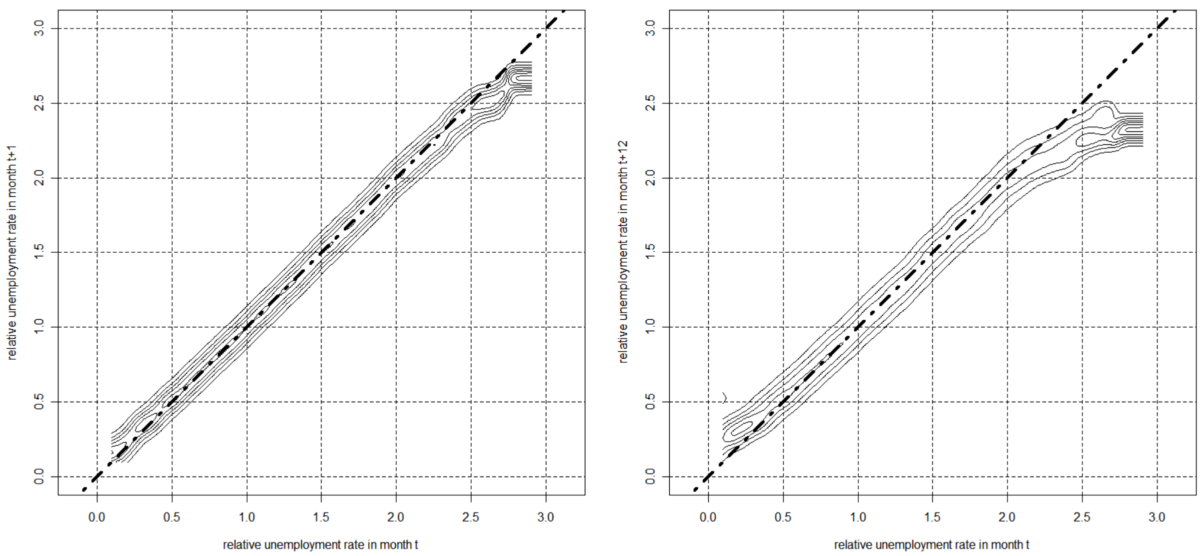

Figure 3: Kernel density estimates - levels, NUTS4 units in relation to the national average, 1998-2007. Source: own calculation based on registry data.

Since the national unemployment rate is rather high for Poland, even reaching the thresholds of 20%, the distribution is condense - we do not observe levels higher than threefold. Nonetheless, the shape is located strongly along the diagonal with no traces of convergence/divergence for monthly transitions (left panel). In the case 12-month rolled ones (right panel), for the highest unemployment regions some convergence may be traced (convergence of 'clubs'). Regions with initially higher unemployment rates tend to exhibit lower relative unemployment rates in the following year (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 rates. Particularly in the case of regions whose unemployment rates already exceeded 40%, one might expect some boundaries as to how much more this rate may still increase[10]. 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 the national unemployment rate rather than the decrease in local differences.

Although ordering of *poviats* seems fairly stable over time, within the last decade only convergence of clubs could be observed, with high unemployment and low unemployment poles of gravitation. Computing the transition matrices intuitively confirms these findings. Transition matrices report probabilities of moving from one decimal group to the other calculated at every point in time. They are a discrete equivalent of the kernel density estimates discussed above. At the beginning of the sample (December 1998) *poviats* were allocated to ten equal sized groups with respect to initial values of the relative unemployment rate. The transition matrix for *poviats* from each decile group reports the probability of staying in the same decile group or moving up or down the relative unemployment rate scale. This procedure, like the kernel density estimates, was applied for the monthly and 12-

[10] Over the analysed time horizon Polish unemployment rate moved between 10% to 20% thresholds.

month rolling changes (left and right panel of Table 1 respectively).

The diagonal values show the probability of staying in the same decile group. Values above the diagonal denote the likelihood of moving to a higher unemployment rate group - conversely, below the diagonal values represent the odds of moving to a lower unemployment group. Ergodic values give information about the percentage of *poviats* that would be found in every decimal group if in the long run the unemployment rate dynamics were characterized by the estimated transition matrix. This should not be interpreted as a long run forecast - rather as a simple summary of tendencies observed in the period for which transition matrix is estimated. In the initial period all groups were equal in size (10% of total sample). Therefore values in the ergodic vector higher than 10% imply that there are tendencies for *poviats* of moving to that group[11]. Values above the diagonal denote the likelihood of moving to a lower unemployment group - conversely, below the diagonal values represent the odds of moving to a higher unemployment group.

Table 1: Dispersions - distribution dynamics for relative unemployment rate (transition matrix)

	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 percent. 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.

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 groups (groups of higher unemployment) is more likely. Importantly, the majority of transitions on an annual basis happens around the fourth to the sixth decimal groups, mostly among themselves. For high unemployment regions the probability of remaining in the same decimal group reaches 80%-90% thresholds over the analysed period.

The ergodic values confirm the above statements. Namely, although the size of this effect is not very large, lower unemployment groups lose districts, while the higher ones gain. Since each decimal group had approximately 37 *poviats* on average, 1% to 2% differences translate to approximately a half of a district (or approximately 40 changes of group allocation over the whole period). In addition, out-of-diagonal values are small, which suggests that the distribution is very stable. Graphically, this was exhibited by the thickness of the kernel density estimates - they are very thin.

Although we argue that NUTS4 is exactly the policy relevant level, one may consider it is too disaggregated for meaningful analyses. As we already mentioned, unlike NUTS4 and NUTS2, there

[11] Please note that after the initial period the boundaries for decimal groups may change together with the distribution.

are no authorities at NUTS3 in Poland, which makes any analyses there only academically interesting. Therefore, we have focused on NUTS2, but instead of analysing them between each other, we have focused on what happens within each of them. Namely, unemployment rate calculated at NUTS2 is essentially an average of unemployment rates recorded at NUTS4 level. If we traced convergence of NUTS4 units within NUTS2 units, this average would have held its informative power. However, in case of either divergence or distribution stability, the average becomes meaningless. Therefore, instead of inquiring the dynamics of NUTS2 averages vis-*à*-vis each other, we attempted to see how relevant they are.

For the purpose of brevity, we only report graphical representations of kernel density estimates for 16 NUTS2 units in Poland (calculated in the same way as in the previous analyses, with the only exception that for each NUTS4 unit relative unemployment rate was calculated with reference to the relevant NUTS2 level instead of national average)[12]. As may be observed, convergence may be observed only in three cases and only for low-end and high-end clubs. These tend to be higher unemployment regions.

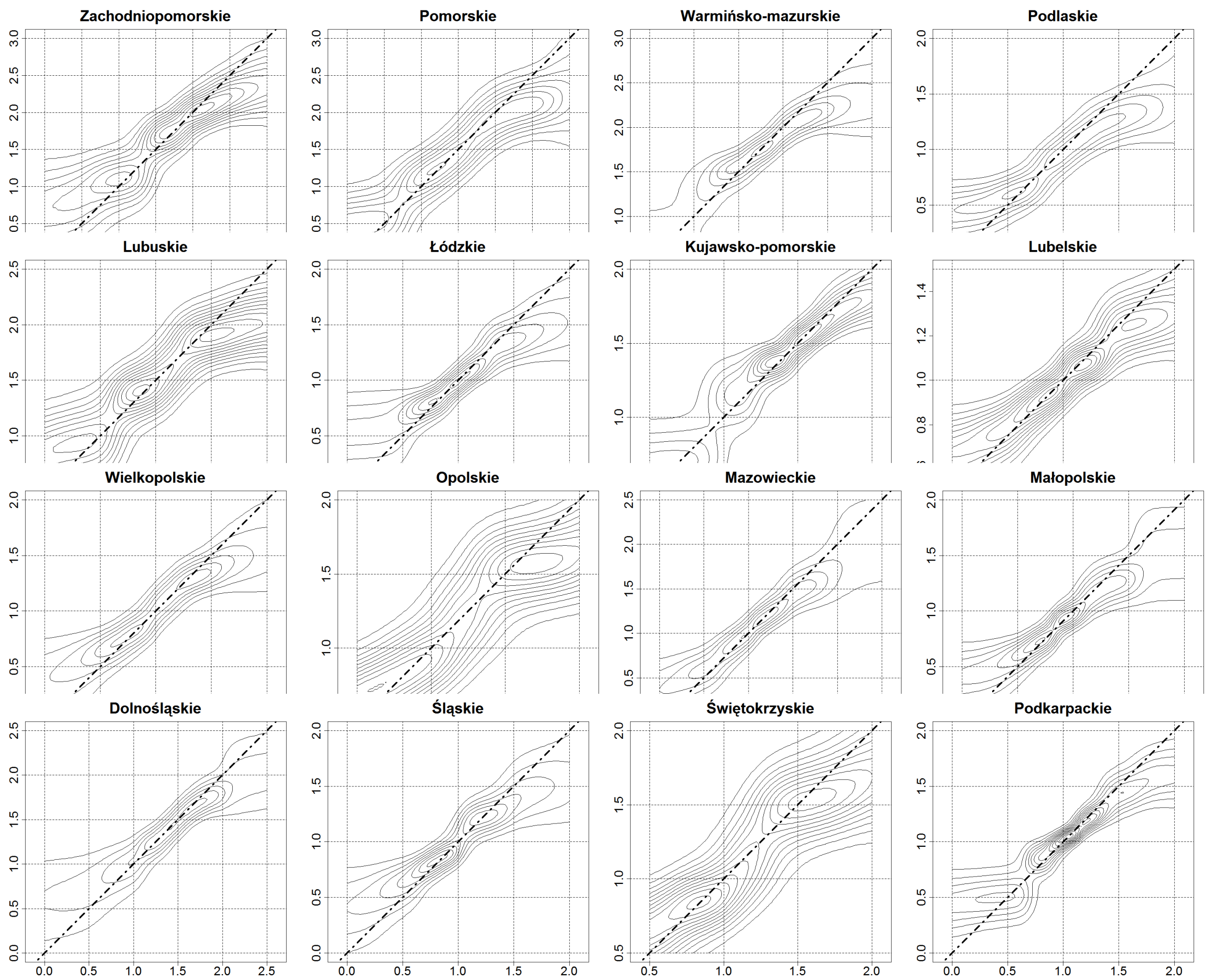

Figure 4: Kernel density estimates - regional convergence, NUTS4 units in relation to the respective NUTS2 average, 1998-2007. Source: own calculation based on registry data.

[12]Graphs for NUTS2 regions were located in order which resembles their boundaries in Poland (a four by four quadratic shape) with the intention facilitating the reader locating NUTS2 aggregates to the maps of NUTS4 presented earlier in the paper.

Figure 3 suggested high-end and low-end clubs. This finding does not hold for all of the regions. Namely, only in the case of four regions (pomorskie, kujawsko-pomorskie, lubelskie and podkarpackie) convergence among lowest unemployment *poviats* may be observed. For highest unemployment regions any convergence is only observed in one region (wielkopolskie), while one region demonstrates to have both high- and low-end clubs (opolskie). In other words, within the majority of regions no support for 'convergence of clubs' was found, let alone general convergence. Consequently, since convergence is only small and considers border cases of very high relative unemployment, regional averages do not seem to be informative about the unemployment rate dynamics within their administrative borders[13].

Naturally, using these NUTS2 level averages, one could also perform an exercise of inquiring the convergence among them. This is depicted in Figure 5. As one would expect, trends are similar to those calculated at NUTS4 level. Namely, in general no convergence may be confirmed, with some evidence of 'clubs' in the high- and low-employment groups. Naturally, since instead of 374 units at *poviat* level, only 16 units at *voivodship* level are now used, the graph is thicker and demonstrates significantly less departures from the national average (the range now is only twofold, whereas both Figure 3 and 4 exhibited three-fold dispersions)[14]. The thickness increases towards the upper and lower 'tails' of the contour plot, which implies that 'movements along the ladder' among highest and lowest unemployment voivodships - naturally, as two separate groups - are less stable than in the middle of the distribution. However, since these are relative values (in reference to the national average), in this context more mobility essentially implies instability of improvements - not that of aggravations.

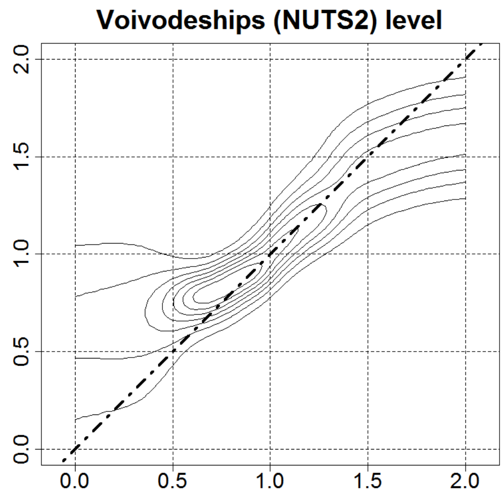

Figure 5: Kernel density estimates - convergence of NUTS2 to national average, 1998-2007. Source: own calculation based on registry data

Comparing Figures 3, 4 and 5 one could ask the question if cohesion policy is more effective at national or at NUTS2 level. Based on the results of Figures 3 and 4[15] we argue, that in fact none of these statements finds confirmation in data. Both national and NUTS2 authorities dispose of

[13]We also observe, that graphs are definitely thicker than in the case of nation-wide analyses, which suggests there is more mobility in the "intra-regional" rankings than in the national ones.

[14]This is due to the fact that NUTS4 labour markets averaged for NUTS2 aggregation produce statistically less diversified outcomes.

[15]Transition matrices available upon request.

financial instruments implemented at NUTS4 level, while the allocation mechanisms are - in principle - targeting more funds to more deprived areas. σ-convergence in the high-end 'club' implies that after reaching certain thresholds, unemployment cannot grow much higher. Much less frequently observed σ-convergence among low-end 'clubs' shows the growth poles - most frequently metropolitan regions attracting most investment and enjoying the best development outlooks. In this sense, cohesion policy may be perceived as inefficient at both national and NUTS2 level.

Please note that this type of analysis is not geographically sensitive. Consequently, theoretically *poviats* within the high and low unemployment poles of gravitation do not necessarily have to be neighbouring or close geographically *poviats*, while the specific processes might differ significantly in the underpinnings. As maps in Figure 2 suggest, this is in fact the case, i.e. there are regions where poor labour market performance spreads across the *poviats* (North and especially Northwest). At the same time improvements in relative local unemployment rates seem to have two main roots. On one hand, they follow from a statistical artefact: the increase of the overall average with the constant local unemployment rate leads to lower relative rates. In fact, the labour market situation in real terms did not improve in these particular *poviat*s. Alternatively, improvements may owe to the idiosyncratic positive shocks due to, for example, localisation of new investments, cfr. Gorzelak (1996).

To corroborate statistically this statement, spatial autocorrelation techniques can be applied. In principle, this requires rather sophisticated data (digital mapping of all processes, eg. investment allocations, labor mobility, activity rates by region, etc.), which are not available at NUTS4 level. Moreover, a theoretical framework allowing for the interpretation of the findings seems crucial, while designing it extends beyond the scope of this paper. Nonetheless, based on a relatively simple inputs including dummy variables denoting units which share borders (neighbourhood regions) global Moran I statistics can be computed, which was performed. A typical symmetric weights matrix is a binary one, where neighbours are coded as 1 and others as 0. Without loosing generality, it can be row standardised (all elements of one row should add up to unity). This index is defined as:

$$I = \frac{n}{\sum_i^n \sum_j^n w_{i,j}} \cdot \frac{\sum_i^n \sum_j^n w_{i,j}(u_i - \bar{u})(u_j - \bar{u})}{\sum_i^n (u_i - \bar{u})^2}, \tag{5}$$

where n is the total number of units, $w_{i,j}$ is the weight matrix (with zeros for adjacent units), while u_i and $\bar{u}$ are unemployment rates at $i-$th location and on average, respectively. If there was no spatial clustering, the expected value of this statistic is 0. Therefore, the value of such index can be tested against a null hypothesis of $I = 0$, with the main obstacle emerging from the fact that under null, Moran's I statistic does not follow a normal distribution (Monte-Carlo experiments needed for testing). Obviously, this is only a global statistic, which implies it only provides a limited set of spatial association measurements. Consequently, clustering as such may be diagnosed, but one cannot ascertain for which values (high/medium/low). The Moran's I statistic for Polish *poviats* is depicted in Figure 6.

Values are statistically significant at a 1% level[16], but fluctuate around a fairly moderate level of 0.25. The interpretation of Moran's I statistic therefore suggests some, albeit weak spatial clustering. Interestingly, the values change over time (increasing from 0.22 to 0.28, these changes are statistically significant at 1% level) despite the fact that as of 2005 improvement in labour market conditions throughout the country may be observed. This observation may be justified by the findings of σ convergence analysis. Namely, we observe that highest and lowest unemployment *poviats* converge in

[16]Detailed results of Monte Carlo experiment available upon request.

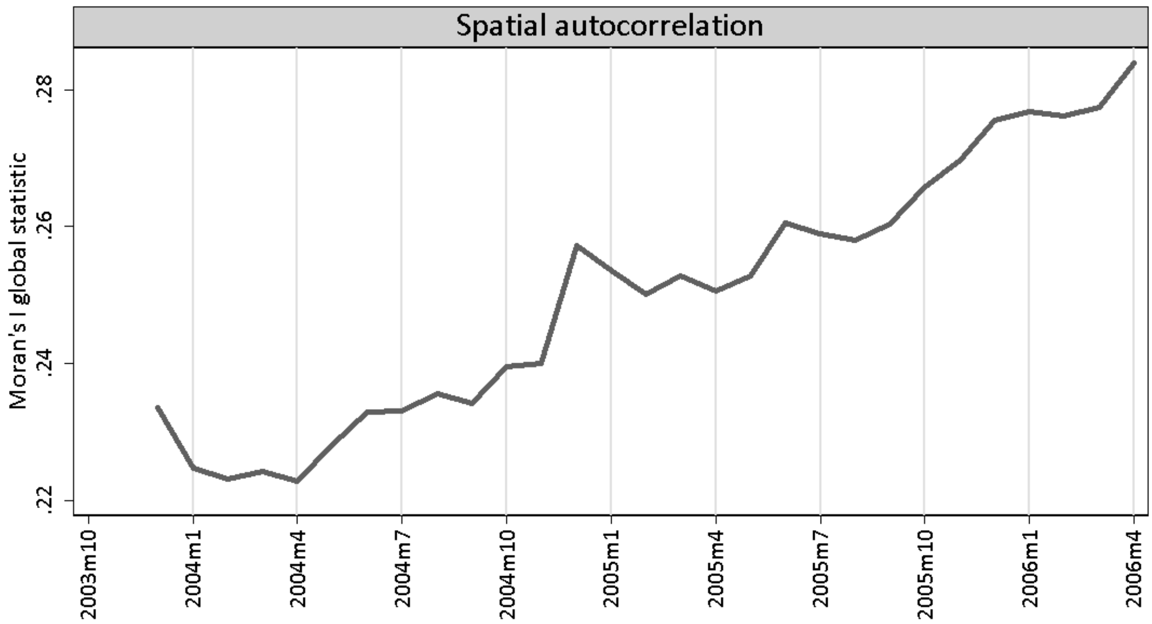

Figure 6: Spatial autocorrelation, Moran's I global statistic. Source: own calculation based on registry data.

distributions. With the improvements in the labour market, employment takes off earliest in larger cities and neighbourhood regions, while the decline in unemployment is slowest in peripheral areas. Therefore, growing intensity of spatial clustering seems in conformity with the previous findings.

All analyses in this sections consider distribution dynamics, *i.e.* evolution of relative unemployment rates. Their levels - if used at all - served the purpose of 'grouping' *poviats*. However, the severity of the unemployment problems follows not only from the distribution, but also from the magnitude of this phenomenon. To address this problem we analyse the convergence in levels of unemployment.

6 Results - β convergence

In this section we report the results of a panel regression of unemployment growth in periods t on the unemployment in the initial period (the β-convergence). To control for low and high unemployment regions, a synthetic proxy was generated, indicating to which of the ten decimal groups a *poviat* belonged in the initial period. Since this measure is constructed on the basis of empirical distribution moments, it can simply take the values of 1 to 10, without hazarding the correctness of estimates due to non-linear or non-monotonic effects. In the estimation a dummy correcting for the statistical effect of December 2003 was additionally included. To control for seasonality as well as changing labour market conditions, the overall unemployment rate in Poland was incorporated, although from an econometric point of view introducing this variable plays the role of imposing a fixed effect on a period in the cross-sectional time-series analysis. Finally, some interaction terms were allowed for to see the extent to which the initial distribution and the initial unemployment rate effects are symmetric for high and low unemployment regions. Consequently, the following equation was under scrutiny:

$$\Delta unemployment_{i,t} = \alpha + \beta \cdot unemployment_{i,T0} + \gamma \cdot control \quad variables_{i,t} + \epsilon_{i,t}. \qquad (6)$$

This equation was estimated in a number of versions:

1. Unconditional β convergence of contemporaneous to initial ($T0$ =December 1998) unemployment rate allowing for annual structural breaks (year-specific dummies) for the whole sample - reported in Table 2, panel A.

2. Conditional β convergence of contemporaneous to initial ($T0$=December 1998) unemployment rate controlling for the evolutions in national average and structural breaks for the whole sample and by decimal groups - reported in Table 2, panel A.

3. Conditional β convergence of contemporaneous to previous 24-month unemployment rate for each *poviat* rolled through the entire time span controlling for the evolutions in national average (more specifically the relation is given by $\Delta u_{i,t} = f(u_{i,t-24}, u_{Poland,t})$ for the whole sample and by decimal groups. Although the choice of 24 month lag may seem arbitrary, these estimates only serve the purpose of tracing if there is any evidence supporting convergence/divergence patterns irrespectively of the starting point. Estimates reported in Table 2, panel B, specific rolling window values averaged for decimal groups depicted in Figure 7.

Control factors include the dummy accounting for the effect of 'December 2003'. We report results for entire sample as well as those where estimates were obtained by decimal group (to account for potential differences in the processes among *poviats* with relatively better/worse labour market outlooks)

To asses that local unemployment rates exhibit β-convergence, the coefficient of β in equation 6) would need to prove statistically significant and negative. A positive value of this coefficient would suggest divergence in levels. However, one must keep in mind that the period we analyse was characterised by a stark increase and a subsequent decrease in the unemployment rates, while nonetheless the final (August 2007) level was higher than the initial (December 1998) for most of the observations (national average was 5 percentage points higher). Therefore, a positive sign of the estimator in the unconditional version would only be a confirmation, that *poviats* with higher unemployment rate in the initial period observe higher unemployment growth rates in subsequent periods - not necessarily that the response is asymmetric among *poviats*. This is why we also include national average unemployment rates in the estimations and verify if they differ across decimal groups[17]. Table 2 reports the findings.

The results of convergence in levels provide similar conclusions as the ones for convergence of dispersions. Namely, the coefficient of main interest - β - turns out largely insignificant most of the time (only for the third group, when its value takes -0.027, *p-value* actually amounts to 0.11, which may be considered marginally significant). We find that this result is robust to the sample selection (total sample and decimal groups essentially provide the same outcomes). Moreover, evidence suggests that indeed the effect of nation-wide unemployment shocks differs across the decimal groups, since all γ estimators are statistically significant. The higher the group (the higher the initial relative unemployment level), the stronger the link between national outlooks and local unemployment rate

[17]Monthly data (relatively high frequency) may exhibit seasonality and autocorrelation. In addition, since units of analysis differ substantially in unemployment levels and changes observed over time, one risks heterogeneity as well. Therefore, our preferred econometric specification is feasible generalised least squares (FGLS) with heteroscedasticity and autocorrelation consistent standard errors and panel-specific autocorrelation structure. More explicitly, we calculate panel-corrected standard error (PCSE) estimates for linear cross-sectional time-series models where the parameters are estimated by Prais-Winsten regression. When computing the standard errors and the variance-covariance estimates, this method assumes that the disturbances are heteroscedastic and contemporaneously correlated across panels.

Table 2: Levels - β convergence analysis

Panel A: Initial unemployment rate

	Full sample	Full sample	Decimal group 1	2	3	4	5	6	7	8	9	10
β	-0.0001	0.0004	0.002	-0.007	-0.027	0.005	0.021	0.002	0.002	-0.002	-0.001	-0.002
	(0.0008)	(0.0005)	(0.005)	(0.011)	(0.017)	(0.02)	(0.024)	(0.018)	(0.017)	(0.18)	(0.015)	(0.004)
γ		0.41*	0.21*	0.28*	0.34*	0.37*	0.37*	0.43*	0.51*	0.51*	0.50*	0.54*
		(0.006)	(0.011)	(0.013)	(0.014)	(0.016)	(0.017)	(0.017)	(0.019)	(0.02)	(0.02)	(0.02)
Constant	0.075	-4.80*	-2.39*	-3.23*	-3.62*	-4.33*	-4.55*	-5.06*	-6.09*	-5.90*	-5.76*	-6.20*
	(.013)	(0.06)	(0.26)	(0.18)	(0.23)	(0.31)	(0.36)	(0.32)	(0.35)	(0.40)	(0.39)	(0.29)
Year (0-1)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N obs.	32 149	32 149	2745	3089	3173	2999	3272	3448	3340	3300	3382	3401
N groups	428	428	43	43	43	42	43	43	42	43	43	43
χ^2	0.02	7131.7	675.9	696.0	828.9	692.7	671.4	849.3	970.0	835.9	756.1	741.2

Panel B: Lagged unemployment rate $U_{i,t-24}$

	Full sample	Full sample	Decimal group 1	2	3	4	5	6	7	8	9	10
β	0.002	0.0004	-0.00009	0.0007	-0.0029	0.001	-0.003	-0.001	-0.002	-0.0004	-0.002	0.003
	(0.0008)	(0.005)	(0.002)	(0.003)	(0.017)	(0.003)	(0.002)	(0.004)	(0.004)	(0.005)	(0.004)	(0.004)
γ		0.40*	0.21*	0.22*	0.29*	0.33*	0.37*	0.37*	0.51*	0.50*	0.48*	0.51*
		(0.006)	(0.011)	(0.012)	(0.014)	(0.015)	(0.016)	(0.019)	(0.021)	(0.023)	(0.022)	(0.025)
Constant	0.075	-4.80*	-4.26*	-5.69*	-6.56*	-7.22*	-7.22*	-8.43*	-9.99*	-9.77*	-9.49*	-10.08*
	(0.013)	(0.06)	(0.21)	(0.27)	(0.28)	(0.32)	(0.37)	(0.32)	(0.40)	(0.44)	(0.43)	(0.48)
Year (0-1)	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N obs.	27 486	27 486	2288	2662	2706	2544	2803	2977	2880	2831	2912	2928
N groups	382	382	32	37	37	35	39	41	39	39	40	43
χ^2	1464.8	6196.6	685.0	605.6	728.1	628.6	678.2	703.5	814.7	679.0	709.0	590.6

Notes: PCSE estimation allows effectiveness even in the presence of AR(1) autocorrelation within panels and cross-sectional correlation and heteroscedasticity across panels. Robust standard errors reported. Year dummies significant (not reported, available upon request).
Standard errors in parentheses. * and ** denote statistical significance at 1% and 5% levels, respectively. Except for pooled unconditional estimation (first column), χ^2 Wald statistics highly statistically significant, *p-values* available upon request.

changes (differences are statistically significant). Naturally, the size of the constant also increases with the decimal group. Importantly, γ coefficients are very significant despite annual dummies inclusion.

Rolling window estimates depicted in Figure 7 reveal even more intriguing results[18]. Namely, indeed, convergence, if any appeared only at the beginning of the sample, except for highest unemployment *poviats* who experienced divergence at the time. Neither the stabilisation of the labour market at the average thresholds 18%-20% brought evidence of levels convergence/divergence, but the subsequent improvements again exhibited divergence among highest unemployment decimal groups (and surprisingly the fourth group) - estimators for groups 7 to 10 are statistically significant and positive. In the lowest unemployment group, initial convergence transforms to stability with time, the former of which can be explained easily (with the deterioration of the situation, any increase of the unemployment rate brings them closer to the mean). The latter however suggests that the improvement happens much faster in low unemployment regions than throughout the country, pointing to - possibly temporary - divergence.

7 Conclusions

Analysing income convergence of EU regions, Boldrin and Canova (2001) reached the conclusion that "[p]roponents of EU support may claim that, had the intervention not been there, the distributions would have spread out further and inequalities become more marked and that policies that aim at

[18] To avoid problems with statistical quality of the estimates, rolling window analysis was performed on data post December 2003 - this significant "shock" to unemployment levels was disastrous to the the quality of estimates.

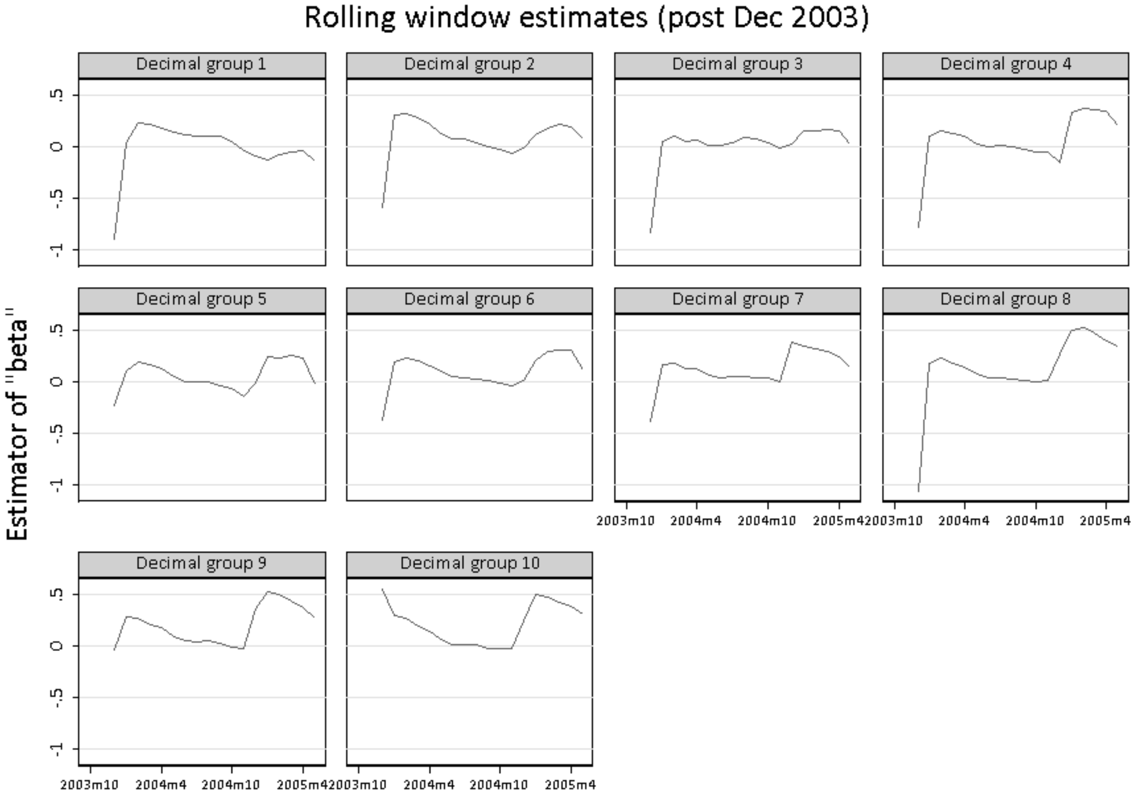

Figure 7: Rolling estimates for β coefficients. Source: own calculation based on registry data.

preventing emigration from poorer to richer areas are the only ways of avoiding the further polarization in income that such would cause. This may be true, as counterfactuals of this kind are almost impossible to test" (p. 242). This statement is generically true for most convergence studies, especially if one is unable to control for the differentiation in financing constraints.

The main purpose of this paper was to inquire the convergence patterns of local labour markets in a transition economy, Poland. We used policy relevant NUTS4 level data, since actual labour market policies - with special emphasis on the active ones - are performed at this specific level. The 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, the sample commences already some years after the beginning of 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 improve throughout the country, already disadvantaged regions benefit less.

In order to inquire the nature of local unemployment rates evolution we employed parametric econometric techniques (convergence of levels, β convergence) as well as nonparametric kernel density estimates (convergence of dispersion, σ convergence). The distribution of unemployment rates in Poland was found to be highly stable over the sample period with only minor evidence in support of the convergence of clubs - high and low unemployment *poviats* separately. In addition, data does not support any conditional β-convergence either, with some evidence of asymmetry between high and low unemployment *poviats*.

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 are of poor quality. Therefore, the time-span is relatively short, especially in the context of stochastic convergence studies in the literature (Bayer and Juessen (n.d.) examine 40 years for Western Germany, Gomes and da Silva (2006) have 22 years at disposal, 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 disaggregated to NUTS4 level. The findings of this paper effectively suggest that the very notions of 'national' or 'regional' unemployment rates are highly uninformative for these countries. Namely, the average is actually only a statistical operation on strongly differentiated processes with sometimes even diverging dynamics.

We also inquired whether local (NUTS4) units demonstrate convergence within regional (NUTS2) units could provide evidence with reference to geographical clustering of relatively more troubled and relatively more favourised areas. Most NUTS2 regions demonstrated no within σ-convergence, while only 'convergence of clubs' in the high- and low-ends of distributions was demonstrated between unemployment evolutions.

This paper has also some important policy implications. Namely, NUTS2 in Poland do not seem to use the fact that they distribute the active labour market policy financing effectively. Each of the Polish NUTS2 regions contain districts from the 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 the most deprived regions.

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