Active Labour Market Policies and Unemployment Convergence in Transition

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In this paper we approach the issue of social cohesion across local labour markets in Poland. We analyse regional dynamics of unemployment rates and try to evaluate the impact of Active Labour Market Policies (ALMPs) in observed trends. Using data from 1999 till 2008 we employ tools typically applied to income convergence analyses to test the stability of unemployment distribution - both unconditionally and taking into account explanatory power of unemployment structure and ALMPs in Polish regions.

Results give no support to the hypothesis of unconditional convergence understood both in terms of levels and in terms of dispersion. Among the highest unemployment regions, however, data seem to suggest "convergence of clubs". The analysis included also accounting for potential impact of ALMPs, controlling for differentiated unemployment structure. We find no evidence that cohesion efforts contribute to more of the convergence or less of the divergence phenomena.

Keywords: regional unemployment rate differentials, convergence analysis, Poland

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

1   Introduction

Under the conditions of stark unemployment rate disparities - as is the case of Poland - so-cial cohesion necessitates that more resources are allocated to the higher unemployment rate

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regions[1]. Indeed, the algorithm allocating funding to active labour market policies (ALMPs) across regions favours more troubled local labour market, giving a premium to higher than average unemployment rate, number of unemployed and worse than average structure of unemployed (e.g. share of long-term unemployed). However, despite general decrease in national unemployment, some regions still struggle with 40% unemployment rate thresholds. Should the ALMPs be implemented in an efficient manner, positive impact on convergence would be observed.

When evaluating the effectiveness of differentiated active labour market policies (ALMPs), two basic approaches exist in the literature. Firstly, concentrating on individual data allows for the estimation of the treatment effect for differentiated instruments, taking into account the developments in a control group. This approach requires not only relatively detailed micro-level data, but also observing individuals after the completion of activisation programmes, which most transition countries lack in general.

The latter approach focuses on regional data instead. The obvious shortcoming is that either quite strong assumptions need to be made concerning the distribution of unemployed among regions (essentially imposing homogeneity during the estimation procedures), or one needs relatively large datasets and considerable heterogeneity to sustain underpinnings for policy implications of the findings. On the other hand, an extensive theoretical framework for the effects of ALMPs on employment has been developed by Calmfors (1993), and recently in a stochastic framework by Lechner and Vazquez-Alvarez (2006). As stated by Hagen (2003), raising the efficiency of matching process is usually regarded as the main aim of ALMPs, and can be reached by adjusting the human capital of job seekers to the requirements of the labour market and by increasing the search intensity (as well as search capacity) of (former) programmes participants. These aims are especially pronounced in transition countries with large structural mismatches.

The aim of this paper is to inquire the *time* stability of the *geographical* distribution pattern, assessing additionally to what extent the interplay of active labour market policies (ALMPs) and economic outlook have influenced the local labour market developments. The situation in the labour market in Poland has been extremely difficult for the past years, with the unemployment rates consistently above 16-18% thresholds (Polish Labour Force Survey, 2006), while the odds to become long-term unemployed exceed 50%. At the same time, as noted by Munich, Svejnar and Terrel (1998), Poland had one of the lowest among the Central and Eastern European Countries (CEECs) outflow rates. Vacancies ratios were dramatically low throughout the whole transition period, with averages around a thousand job-seekers per one offer[2].

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[1]Prior to the recent global financial crisis, Poland was among the highest unemployment rate countries in the European Union. Financing of active labour market policies has been intensified gradually as of 2004, reaching over 0.2% of GDP in 2006. Nonetheless, these policies covered barely 20% of the unemployed, with some evidence of "creaming", (Tyrowicz 2006).

[2]The principal studies in this area are carefully surveyed in Münich, Svejnar and Terrell (1997).

In this paper we apply β- and σ-convergence analyses to the registered unemployment data covering the period 1999-2008. We use a unique data on active labour market policies and unemployment structure, which have not been analysed before. The technique of kernel density estimates for convergence analysis is applied to unemployment rates on NUTS4 (*poviat*). The period we chose captures the so-called "second wave of unemployment" commencing in 2001 as well as implementing ALMPs on a more comprehensive scale (from less than 10% up to 25% of unemployed subjected to active instruments). Taking into account the emphasised structural character of Polish unemployment as well as challenges this situation implies, we employed a technique typically applied to income analysis allowing to inquire, whether any regional differences in development patterns may be observed. Notably, we test the hypotheses of convergence of levels as well as the stability of unemployment rates distribution. The first of the questions is approached with the use of econometrics, whereas the latter is addressed with the use of nonparametric kernel density estimates[3].

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 main findings of β-convergence analysis. Section 7 concludes.

2 Literature Review

There are at least two main motivations to inquire the dynamics of local labour markets. First, most macro-level models assume implicitly homogeneity and symmetry of shock response at least within countries. This assumption is not always rooted in data. In the case of EU for example, studies find income convergence between nations and divergence on more disaggregated levels, both within and across countries (for example, Egger and Pfaffermayr (2005) or Paas and Schlitte (2007)). This problem seems to receive more attention in the recent years[4]. 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,

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[3]This tool is traditionally applied to the problems of income convergence, while Overman and Puga (2002) are probably the only attempt to employ it for unemployment data in a spirit similar to ours. In more general attempts, 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.

[4]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).

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 policies financing is distributed with preference to areas with (i) higher than average unemployment rate, (ii) higher than average share of long-term unemployed and (iii) higher than average number of unemployed. Consequently, regions facing relative hardships receive more resources to alleviate their impact.

In Poland, the process of employment restructuring consisted mainly of the reductions in employment with growing average job tenure as well as average time spent in unemployment or inactivity (Svejnar 2002a). Dismissals - if compensated at all - found their outcome with 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 by age groups in Poland. People who lost their employment usually became permanently unemployed or inactive for durations of many years, (Grotkowska 2006).

Using labour force survey data, it is easy to identify the ideal type of individual 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 literature before.

Unemployment convergence at regional level received recently considerable attention from the academics. Buettner (2007) compares empirical evidence on regional labor market flexibility in Europe. 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)). For the case of CEECs, the role of transition processes may indeed still be significant (Svejnar 2002b). In fact, 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 1995-1999 time span - a period of gradual improvement in both economic and labour market outlooks. This period was followed by a five year period of stark unemployment increase with a culmination at 20% of unemployment in the middle of 2003 and noticeable decreases only observable as of the second half of 2005. Thus, the persistence of high differentiation in regional unemployment rates 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 (see Decressin and Fatas (1995), Obstfeld and Peri (1998) or more recently Arm-

strong 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 markets performance). Overman and Puga (2002) perform conditional kernel density analyses of European unemployment rates taking into account the distributions of underlying fundamentals (eg. the skills, the regional specialisation as well as the growth rates of population and the labour force).

Suggesting a different angle, 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. 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). 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 cointegration of regional and national unemployment rates. Bayer and Juessen (2006) find moderate evidence in support of the convergence hypothesis. Similar technique has been applied by Gomes and da Silva (2006) for the regions of Brazil finding strong evidence of hysteresis and unemployment regional differential persistence.

However, one can put forward a strong argument against these results, namely that stationarity of the regional unemployment rates differentials can happen both under 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 rates distribution across counties of 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%.

The effects of ALMPs in a transition context have been analysed already in mid 1990s, albeit with scarce data: including Boeri (1994), Lehmann (1995), Gora, Lehmann, Socha and Sztanderska (1996), Kwiatkowski and Tokarski (1997) and Puhani (1999) as well as summary by Dar and Tzannatos (1999), Martin (2000) and Grubb and Martin (2001). Typically, unlike micro-level studies, the findings were rather discouraging in terms of value for money or sometimes even lack of visible ALMPs effects. Frequent defence argument bases on the fact that

some ALMPs effects take longer to appear or may not be discounted in the period of labour market contraction but will eventually boost employment with the change of business outlooks.

Vodopivec, Wörgötter and Raju (2003) review also the effects of the passive component of the labour market policies, finding some expected negative spillovers and interrelations between active and passive labour market policies. While a new wave of research sprung recently, incorporating Balkan and CIS countries with the availability of World Bank labour market surveys, the findings of the post-transition period are only slightly more discouraging. e.g. Vroman (2002), Godfray (2003), Betcherman, Olivas and Dar (2004), Hujer, Thomsen and Zeiss (2006), Fares and Tiongson (2008).

3   Methodology

Empirical strategies 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 decreases over time and this may be approached by testing for σ convergence. Finally, one can try to investigate how persistent the regional unemployment rate *differentials* are, by applying dynamic distribution analysis[5]. Importantly, β-convergence is a necessary, but not a sufficient condition of σ-convergence. In principle, in the case of conditional β-convergence one could expect that within the groups of units approaching the same *level*, also dispersion around it decreases leading to σ type convergence of clubs.

Kernel density estimation as the dynamic distribution analysis tool can be used to approach σ type convergence. In general, it approximates 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).

The choice of the kernel function has an 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 window size predetermines the degree of smoothening of the resulting curve or surface. Too wide bandwidth window will hide the real data distribution, while too narrow might misleadingly result in function with multiple vertices - not necessarily true

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[5]Pritchett (1997), Temple (1999), de la Fuente (2000) as well as Barro and Sala-i Martin (2003) provide excellent literature reviews of the subject and empirical evidence thusfar.

in reality and rather troublesome in terms of interpretation. Silverman (1986) provides the procedures for finding optimal bandwidth for different kernel functions, basing on standard deviations and inter-quartile range (independently for all vectors in the case of multidimensional distributions). An extension of this approach is to use 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 (previous) unemployment rate is denoted 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:

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

where $f_x[x]$ is the marginal distribution of the initial unemployment rate, while $f[x, x + 1]$ represents the combined distribution of x and $x + 1$. In estimation of the conditional density function both numerator and denominator of (1) are replaced by non-parametric estimators. The adaptive kernel estimator[6] of marginal distribution of the initial unemployment rates is given by:

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

where n is the number of observations, h_x is the bandwidth window for the initial unemployment rate and $K[.]$ represents the kernel function[7]. At the first stage of the adaptive estimation 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, at the first stage of the adaptive estimation joined density function is estimated with the equal (optimal) bandwidth window for all observations – weights are uniform. Subsequently, basing on these estimates, local weights are calculated according to:

$$w_i = \left(\frac{\hat{f}_g}{\hat{f}_k(y_i, x_i)} \right)^{\frac{1}{2}} \tag{4}$$

In this expression, the denominator of the formula in the parentheses is the joined density function estimator calculated with the use of uniform weights and bandwidth window[8], while

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[6]Two stage method of density estimation that allows for better approximation of real structure of the data.

[7]We used the Gaussian kernel function.

[8]Fixed window kernel estimate.

the numerator is the geometric mean of this estimator for matching pairs of observations of both variables. The final conditional density function is estimated using the weights from equation (4) in equations (2) and (3) and calculating their ratio, according to equation (1)[9].

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 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 subsequent point in time. Stability implies directly that neither divergence nor convergence of distribution can be tracked.

The distribution used in the analysis is obtained by regressing growth rates of the unemployment rate on a distributed lags of the single conditioning variable and extracting the residuals for subsequent analysis. The residuals from the regression contain the part of the growth rate of unemployment that is not explained by the analyzed factor. Prior to the analysis we conduct Granger causality tests, to take into account only these variables that seem to have impact on unemployment rate dynamics, at the very least in the sense of being leading indicators of subsequent changes.

Finally, these residual growth rates are used to calculate the distribution of unemployment rates with the impact of the analyzed variable. In our causality analysis we assume a horizon of half a year - taking six monthly lags of the growth rate of the independent variable to explain the regional unemployment growth rate. The method derives from that suggested by Sims (1972) and implemented subsequently by Quah (1996). Consequently, we implicitly assume one directional causality of analyzed factor at the unemployment rate. While not in all cases this would be justified, the main concern of this paper is to account for the dynamics of the unemployment rate distributions. Therefore, even if some reverse causality was in play, the analysed direction would still be active, whilst kernel density analysis is non-parametric and hence not susceptible to eventual inconsistency issues.

4   Data

In the paper monthly data covering the period from January 1999 till June 2008 were used at the lowest available administration level of *poviats*[10]. This paper uses for the first time unique administrative data on unemployment structure (as of January 2000) and ALMPs spendings (as of January 2001) reported consistently by labour offices at NUTS4 level. We treat structural vari-

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[9]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.

[10]NUTS (Nomenclature of Territorial Units for Statistics) is a common clasification of territorial units for statistical and administrative purposes, with higher numbers indicating lower level of aggregation. NUTS4 correspond to local level (as opposed to NUTS5, which correspond to community level)

ables as control factors while financing of active labour market policies plays the role of proxy for policy measures. We construct three measures: spendings per person in any treatment by public employment services, spendings per unemployed (irrespectively of whether he/she was in treatment or not) and share of unemployed in any treatment offered by public employment services, *i.e.* coverage. These measures were meant to capture intensity, depth and extensiveness of active labour market policies use, respectively[11].

In this paper we employ policy relevant NUTS4 level unemployment data using official registry data for Poland. In total we use 374 units[12]. 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.

Using more aggregate level would diminish the value added of the analysis in at least two ways. First of all, NUTS2 level (Polish *voivodships*) do not perform any active labour market policies, which implies that outcomes are not dependent upon efforts made at this level, but rather aggregate of different (approximately 20-30) lower level administrative units. Secondly, NUTS2 units are relatively large and therefore less diversified both when it comes to comparing unemployment levels and unemployment structure. Indeed, NUTS2 units are so big and heterogeneous that only slight differentiation 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).

Due to the administrative changes in Poland in 1999 no data before that moment are available at NUTS4 level. 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 2004-2006 which allows us to explore the regional symmetry of response to nation-wide macroeconomic changes. Figure (1) demonstrates the unemployment developments in Poland over this period.

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 base on the registered unemployment recorded by local labour 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

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[11]Despite numerous efforts by non-governmental organisations, public employment services still hold virtually a monopoly over activisation instruments, since providing such services is subject to licensing while public financing is essentially unavailable to non-public providers.

[12]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. 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 units comprising cities and rural areas were divided into two separate *poviats*, each with a different labour market structure and potential, subsequent to the change, both these units are treated as new in our sample.

Figure 1: Unemployment Rate in Poland (1999-2008)
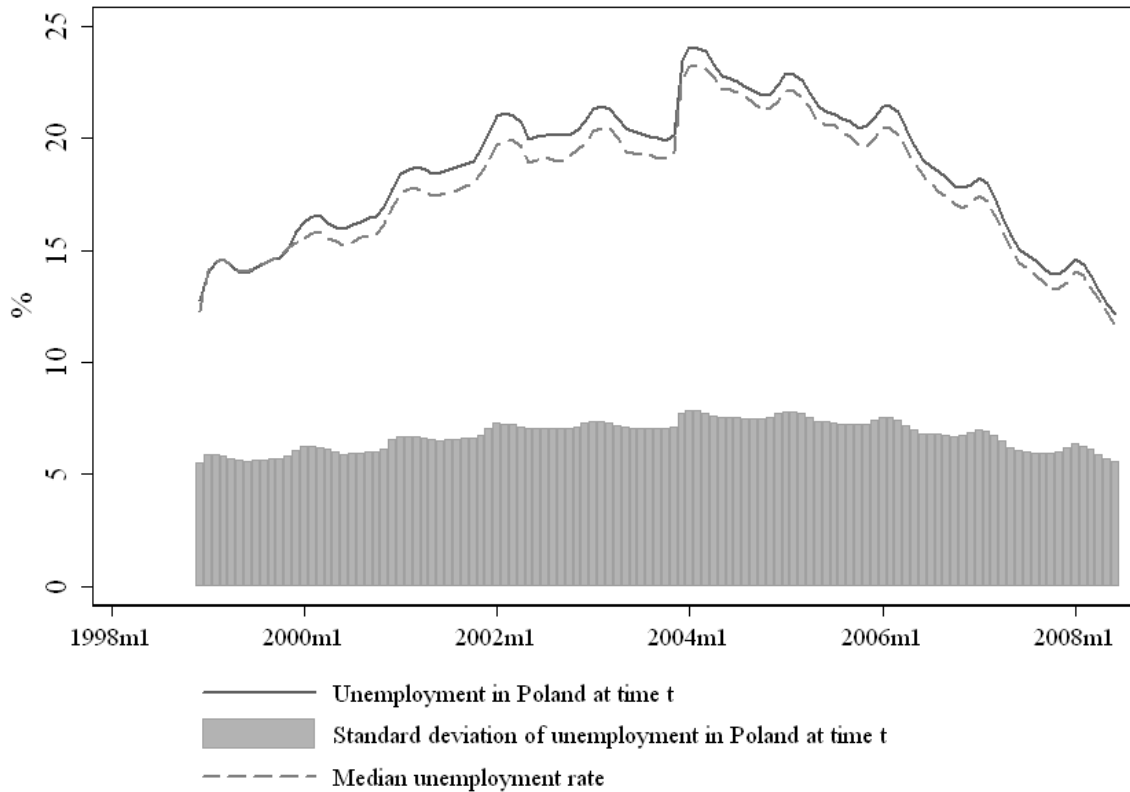
Source: CSO, registered unemployment data, median for all *poviats* at each point in time, standard deviation in percentage points.

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 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 which stems both from the skewness of the distribution and the fact that generally *poviats* with higher unemployment rate tend to have larger population (the average is weighted by population), Figure (1).

More importantly, dispersion of the unemployment rates has been constantly growing over the entire time span - especially in the down cycles, be it 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-direction trends.

Figure 2: Unemployment in Polish *poviats*
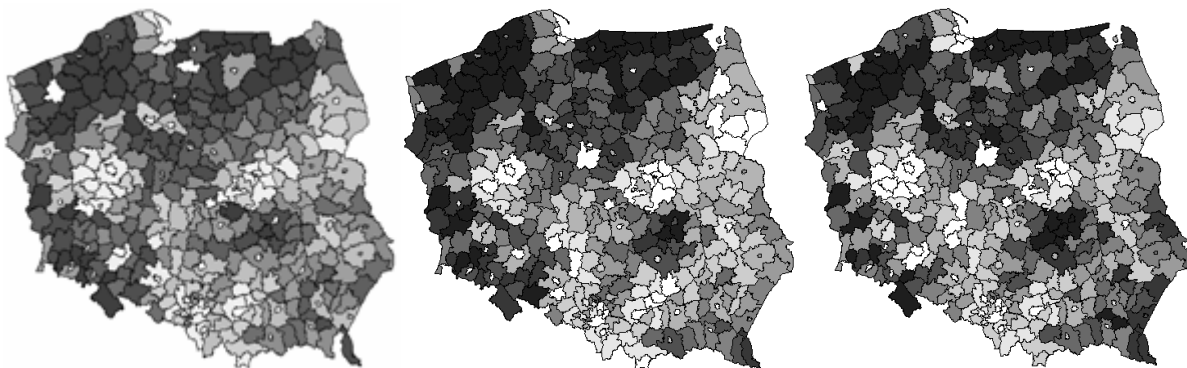
Source: registered unemployment data, CSO. Dec 1998 in left panel, Dec 2003 in the middle and Jun 2008 in the right panel, the darker the shade, the higher relative unemployment rate.

The maps on Figure (2) demonstrate December unemployment rates on *poviat* level for the 1998, 2003 and the most recent available June 2008, with the shades darkening with the relative unemployment rate [13]. In fact, analysing detailed data, one finds that the discrepancies on the local level are even 25-fold (e.g. from 0.11 of the average to 2.8 of this value).

5 Results - Distribution Dynamics

The analysis of σ-convergence - as covered in Section 3 - allows to inquire the dynamics of 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

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[13]White spots follow from the changes in the structure of *poviats* in Poland (benchmarked to the national average). As of January 2001 municipal units were created, while past data referring to these units cannot be inferred from CSO datasets.

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 - June 2008) - monthly changes on the left panel and 12-month rolling changes in the right panel[14]. These figures depict in two dimensions 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 (figure is positioned along the diagonal, which suggests that only small changes in unemployment occur on a monthly basis).

Yearly relative unemployment rate (right panel) shows that more changes occur on yearly basis than on monthly basis, but unemployment is still quite stable (figure 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 - especially the one for the high unemployment rate values. This suggests that separately the *poviats* with highest unemployment rates (above 2.3 of the average) and those with lowest unemployment rates (below 0.25 of 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*.

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 groups 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. Transition matrix for poviats from each decile group reports probabilitiy of staying in the same decile group or moving up or down the relative unemployment rate scale. This procedure similarly to kernel density estimates was applied for monthly and yearly rolling changes (left and right panel of Table 1 respectively)[15].

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[14]12-month rolling change means annual changes for all available months in the sample.

[15]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, values below the diagonal represent the odds of moving to a lower unemployment group. Ergodic values inform about the percentage of *poviats* that would be found in every decimal group if in the long run the unemployment rate dynamics was 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 sized (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. Please note that after the initial period the boundaries for decimal groups may change together

Figure 3: Kernel Density Estimates - Levels, NUTS4 units in relation to the national average, 1998-2008
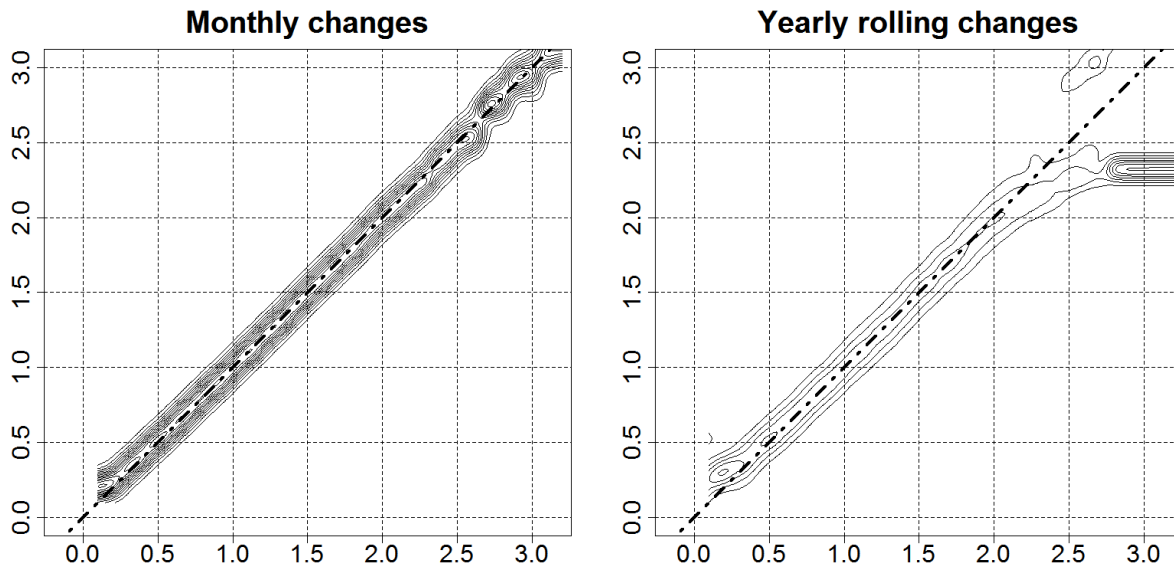
Source: own calculation based registered unemployment data.

On average 91% of *poviats* remain in the same group on the monthly basis, while 62% are likely not to change the decimal group for rolled 12-monthly changes. Probabilities above the diagonal are in most cases 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. For high unemployment regions the probability of remaining in the same decimal group reaches 80%-90% thresholds over the analysed period. Generally, out-of-diagonal values are rather 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.

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. It is interesting to observe that in case of 12-month transitions (right matrix) the ergodic distribution indicates polarization with relatively stable group of lowest unemployment rate, increasing number of poviats at high unemployment groups and diminishing middle unemployment groups.

Table (2) demonstrates the values of ergodic vectors when one takes into account the structure of unemployment in Polish *poviats* as well as financing and coverage of ALMPs. According to the procedure described in Section (3) we used monthly data for the period January 2000 -

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with the distribution.

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	88	11	1	0	0	0	0	0	0	0
2	3	92	5	0	0	0	0	0	0	0	15	64	18	3	0	0	0	0	0	0
3	0	5	89	6	0	0	0	0	0	0	1	20	56	20	3	0	0	0	0	0
4	0	0	6	87	6	0	0	0	0	0	0	2	21	48	24	4	0	0	0	0
5	0	0	0	6	87	7	0	0	0	0	0	0	1	23	46	24	5	0	0	0
6	0	0	0	0	7	87	6	0	0	0	0	0	0	3	24	46	23	4	0	0
7	0	0	0	0	0	6	89	5	0	0	0	0	0	0	2	24	52	19	2	0
8	0	0	0	0	0	0	5	90	5	0	0	0	0	0	0	1	18	60	20	1
9	0	0	0	0	0	0	0	5	91	4	0	0	0	0	0	0	0	17	67	16
10	0	0	0	0	0	0	0	0	3	97	0	0	0	0	0	0	0	0	11	89
E	8	8	8	9	9	9	10	11	12	14	9	7	6	7	7	8	9	11	15	22

Notes: Table reports the probabilities in percents. Boundaries for the decimal groups were given by 66.5%, 81.1%, 91.7%, 102.6%, 113.3%, 125.5%, 139.7%, 156.8%, and 180% of the national unemployment rate in the case of monthly transitions. For rolled 12-month transitions these boundaries were 67.3%, 81.3%, 91.7%, 102.3%, 112.9%, 124.7%, 138.4%, 155.6% and 177.9%. In either case, they were computed based on the empirical distributions in the initial period.
Line *E* denotes values for ergodic vector.

December 2006 for structure and January 2001 - December 2006 for ALMPs. Performing analyses similar to those described in Table (1) above, we computed transition matrices and ergodic vectors controlling for the explanatory power of each of these variables separately.

Comparing the results for varied control factors one should note four important points. First of all, the longer the time span, the more clear emergence of high-end club of *poviats*. This is surprising since, as depicted by Figure (1), 2007-2008 period saw a stark improvement in the general labour market conditions. Dispersion measured by standard deviation decreased as well. Therefore, clear emergence of these "clubs" as well as decreasing size of middle-range decimal groups suggest that improvement was not equally spread, while for regions within 20% of the average (both below and above) relative deprivation seems to be suggested by data.

This assertion is further corroborated when one compares the boundaries for conditional and unconditional analyses. Namely, as of third decimal group boundaries are lower in the case of conditional. This implies that heteroscedasticity of residuals in each separate regression is inversely linked to the unemployment level. This implies that less regions are essentially less "responsive" to changes in structure, but also policy instruments employed. While justifiable by the nature of unemployment dynamics (local labour markets experiencing more hardship are generally less dynamic and therefore have lower "disturbancies"), this finding will impose the necessity to perform β convergence analyses taking into account estimators consistent in this environment.

Table 2: Dispersions - Conditional Ergodic Vectors

|	Group 1	Group 2	Group 3	Group 4	Group 5	Group 6	Group 7	Group 8	Group 9	Group 10
Unconditional										
Monthly [a]	(≤66.5%)	(66.5-81.1%]	(81.1-91.7%]	(91.7-102.6%]	(102.6-113.3%]	(113.3-125.5%]	(125.5-139.7%]	(139.7-156.8%]	(156.8-180%]	(≥180%)
Yearly rolled [a]	9%	7%	6%	7%	7%	8%	9%	11%	15%	22%
Monthly [b]	8%	8%	8%	10%	11%	11%	10%	11%	12%	11%
Conditional										
Inflows rate [b]	(≤64.5%)	(64.5-77%]	(77-86.4%]	(86.4-95.8%]	(95.8-105.8%]	(105.8-116.8%]	(116.8-129.2%]	(129.2-145.3%]	(145.3-165.8%]	(≥165.8%)
	8%	8%	8%	8%	9%	10%	10%	11%	13%	14%
Outflows rate [b]	(≤64.5%)	(64.5-77.1%]	(77.1-86.5%]	(86.5-96.1%]	(96.1-106%]	(106-117.1%]	(117.1-129.3%]	(129.3-145.4%]	(145.4-166%]	(≥166%)
	8%	7%	8%	8%	9%	10%	10%	11%	13%	14%
% of females [b]	(≤64.5%)	(64.5-77%]	(77-86.4%]	(86.4-95.9%]	(95.9-105.9%]	(105.9-116.9%]	(116.9-129.2%]	(129.2-145.3%]	(145.3-166%]	(≥166%)
	8%	8%	8%	8%	9%	10%	10%	11%	13%	14%
% of LTU [b]	(≤64.5%)	(64.5-77%]	(77-86.5%]	(86.5-95.9%]	(95.9-105.9%]	(105.9-117%]	(117-129.3%]	(129.3-145.4%]	(145.4-165.9%]	(≥165.9%)
	8%	7%	7%	8%	9%	10%	10%	11%	13%	15%
% of rural inhabitants [b]	(≤64%)	(64-76.3%]	(76.3-85.1%]	(85.1-94%]	(94-104%]	(104-114.5%]	(114.5-127.2%]	(127.2-142.4%]	(142.4-161.9%]	(≥161.9%)
	18%	14%	12%	12%	10%	9%	8%	7%	6%	5%
% of female LTU [b]	(≤64.5%)	(64.5-77.1%]	(77.1-86.5%]	(86.5-96%]	(96-106%]	(106-117%]	(117-129.3%]	(129.3-145.3%]	(145.3-165.9%]	(≥165.9%)
	8%	7%	8%	8%	9%	10%	11%	11%	13%	15%
% of youth [b]	(≤64.5%)	(64.5-77%]	(77-86.4%]	(86.4-95.9%]	(95.9-105.9%]	(105.9-117.1%]	(117.1-129.3%]	(129.3-145.3%]	(145.3-165.9%]	(≥165.9%)
	8%	7%	7%	8%	9%	10%	10%	11%	13%	15%
% of benefits recipients [b]	(≤64.5%)	(64.5-76.8%]	(76.8-86.4%]	(86.4-95.9%]	(95.9-105.9%]	(105.9-117.1%]	(117.1-129.3%]	(129.3-145.3%]	(145.3-166.1%]	(≥166.1%)
	8%	8%	8%	8%	9%	10%	10%	11%	13%	15%
PLN per treated [c]	(≤65.9%)	(65.9-76.8%]	(76.8-86.6%]	(86.6-96.4%]	(96.4-105.4%]	(105.4-115.7%]	(115.7-127.5%]	(127.5-143.6%]	(143.6-162.7%]	(≥162.7%)
	4%	4%	4%	6%	8%	9%	11%	13%	18%	23%
PLN per unemployed [c]	(≤65.9%)	(65.9-76.8%]	(76.8-86.6%]	(86.6-96.4%]	(96.4-105.4%]	(105.4-115.7%]	(115.7-127.6%]	(127.6-143.6%]	(143.6-162.8%]	(≥162.8%)
	4%	4%	4%	6%	8%	9%	11%	13%	17%	23%
% of treated [c]	(≤65.9%)	(65.9-76.9%]	(76.9-86.7%]	(86.7-96.4%]	(96.4-105.4%]	(105.4-115.8%]	(115.8-127.6%]	(127.6-143.6%]	(143.6-162.7%]	(≥162.7%)
	5%	4%	5%	6%	8%	9%	11%	13%	17%	23%

Source: Registered unemployment data from CSO and ML&SA. Yearly rolled denotes 12-month changes (annual) rolled over the entire span, where step size is one month. [a] denotes time period 1998-2008, [b] denotes time sample 2000-2006, [c] denotes time sample 2001-2006

Third finding is that in general there is little difference between the ergodic vectors in unconditional and conditional analyses. This suggests that there is no clear answer to the question about the drivers of unemployment rate dynamics. None of the structural factors analyzed seems to have significant explanatory power in as far as differing dynamics are concerned.

The fourth conclusion concerns the fact that essentially none of the financial measures seems to contribute to higher cohesion. The size of two high-end groups virtually doubles, while the lowest unemployment regions become half as numerous whichever of the measures is concerned. Data do not seem to suggest that financial measures reach their objectives. Naturally, we lack the clear case-specific counterfactual (what would have happened in each of the units if the financing were not there), but from the *ex post* perspective they seem to suggest that polarization tendencies are strengthened rather than cohesion.

This last conclusion is further corroborated by contour graphs depicting the behaviour of conditional kernel density estimates for policy measures, Figure (4).

Figure 4: Kernel Density Estimates - Impact of Policy Measures, NUTS4 1998-2008
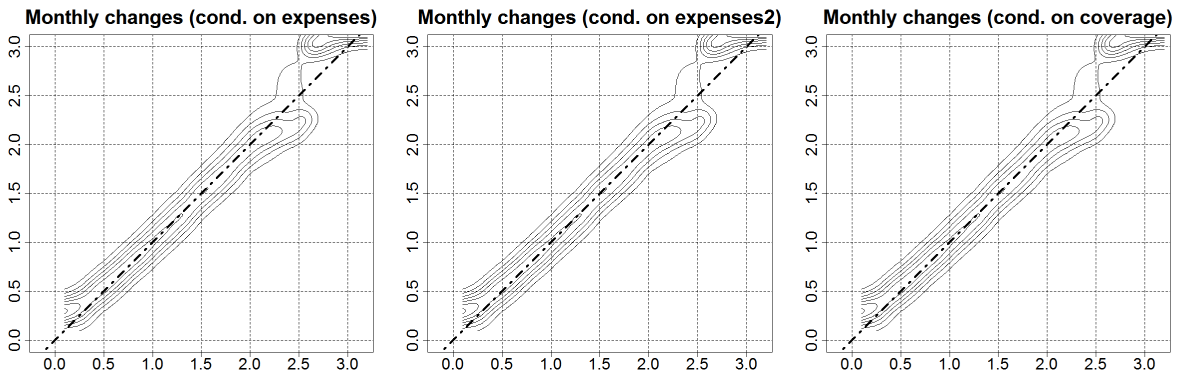
Source: own calculation based on registered unemployment data.

Essentially, unlike in Figure (3), the "club" of high relative unemployment groups lies above the diagonal. This is equivalent to stating that they converge to *higher* unemployment levels when controlling for the effect of policy instruments. In other words, had the financing been not accounted for, the behaviour of the data suggested convergence towards *lower* levels. Although this may suggest that policy instruments are in fact counterproductive, this finding may also result from the nature of the financing algorithm. As was described earlier, local labour markets suffering more absolute and relative hardships receive more financing. We are unable to analytically discriminate between these two explanations, while probably both effects are at play.

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 demonstrated in Figure (2) suggest, that in fact this is the case, i.e. there are regions where poor labour market performance spreads across the *poviats* (North and especially northern West). 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 rate - in fact, 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 (see Gorzelak (1996)).

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 from the magnitude of this phenomenon too. 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). 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, overall unemployment rate in Poland was 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.

Original specification of the convergence hypothesis necessitates the testable equation of the form:

$$\Delta x_{i,t-T0} = \alpha + \beta x_{i,T0} + \gamma \cdot control \quad variables_{i,t} + \epsilon_{i,t}. \tag{5}$$

Naturally, convergence implies negative β coefficient (the lower the level initially, the higher the subsequent growth rates)[16]. However, unlike for example GDP, unemployment is not expected to grow indefinitely. Consequently, the expression in terms of growth rates may be misleading, since changes are indeed frequently negative. Consequently, we inquired if the *level* of unem-

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[16]One could alternatively specify the model in terms of relative unemployment rate, that is unemployment rate in *poviat i* at time *t* divided by unemployment rate in Poland at time *t*. However, previous section demonstrates that no convergence in *relative* unemployment rates can be expected, which implies that using national average as phenomenon would introduce an extra source of variation, while at the same time the results would no longer maintain their interpretative value. Therefore, we use actual unemployment levels.

ployment in the future depends - positively or negatively - on the initial unemployment level. Hence, testing actual convergence requires the β coefficient in the following equation (6) to fall short of unity.

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

This equation was estimated in a number of versions. Naturally, we first approach the unconditional form of β convergence, but this may be approached in at least two ways. One of them comprises of comparing initial states with the final ones throughout the entire time span (*i.e. T0 = Dec1998 and t = Jun2008*). Another way is to make use of rolling estimates using, for example, 12-month changes (*i.e. T0 = t − 12 and t = Dec1999, Jan2000, ..., Jun2008*). Moving to conditional form of convergence, one can include control factors (structure and dynamics of local labour markets[17]) as well as policy variables (intensity, depth and extensiveness of ALMPs). Naturally, this can only be done in a rolling version, *i.e.* moving the 12-month window throughout the available sample.

As discussed earlier, to assess that local unemployment rates exhibit β-convergence, the coefficient of β in equation (6) would need to turn out below unity in a statistically significant way. Value of this coefficient exceeding unity would suggest divergence in levels. However, one must keep in mind that the period we analyse was characterised by stark increase of the unemployment rates, while the final level (June 2008) was only approaching the initial one (December 1998) for most of the observations. Therefore, exceeding unity 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 additionally we include national average unemployment rates in the estimations. Please note, that it is equivalent to having time-specific fixed effects in the analysis.

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. On the other hand, however, there are strong arguments in favour of the potential endogeneity. Namely, although financing of the ALMPs follows a backward looking algorithm, if underlying fundamentals exhibit persistence - which they frequently do - statistical issues emerge. The best way to circumvent the potential bias of the estimators is to use the GMM as developed by Arellano and Bond (1991). However, these standard estimators do not permit for autocorrelation. Therefore, we resort to an estimator

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[17]Control factors include the dummy accounting for the effect of "December 2003".

consistent under autocorrelation, as developed by Arellano and Bover (1995) and Blundell and Bond (1998).

Table 3 reports the findings.

Column (1) of Table (3) reports the analysis for basic form of β-convergence analysis. Namely, we regress the unemployment rates at the end of the sample over the initial (December 1998) rate of unemployment. Clearly, no evidence of unconditional convergence is provided by the data, the estimator is statistically significant and exceeds unity (*p-value* of test $\beta = 1$ amounts to 0.0000). The results would imply that 1 percentage point of unemployment in December 1998 "results" effectively in 1.04 percentage points of the unemployment rate at the end of our estimation sample. However, taking into account the evolution in national unemployment average one could suspect that these 9.5-years changes in unemployment level may be indeed a misleading indicator.

Therefore, in column (2) we move to using rolling version of β-convergence analysis. Namely, we use 12-month lags as "initial" unemployment level, while the estimates are calculated over the entire sample. Findings are slightly more optimistic in a sense that statistically significant coefficient below unity is found. The size of this estimator increases slightly when we introduce time trend (both linear and squared) though, as reported in column (3). The inclusion of trend improves the statistical properties of the model, whilst it is justified by the evolution of the national average unemployment rate, as depicted by Figure (1).

Columns (4)-(8) report conditional β-convergence analyses for the entire available sample. We first control for the local unemployment structure (female share of unemployed, the share of long-term unemployed, youth as well as workers above 55 years of age and workers inhabiting rural areas). We include here also measures of local labour market dynamics (inflows and outflows rates). Although these may be affected by policy instruments (e.g. providing financial incentives to registration by relaxing the requirements or offering more activisation opportunities), they are to much higher extent a factor discriminating between "sleepy" and "active" labour markets. All these variables turn out to be significant and with expected signs[18]. Estimators of structural variables are also fairly stable across the specifications.

Columns (7) and (8) report estimators for measures of availability of financing, average treatment cost as well as treatment coverage (*per se* and controlling for unemployment structure and dynamics, respectively). In column (7) neither intensity nor extensiveness of active labour market policies prove to be significant. When unemployment structure is controlled for, coverage remains significant and of a comparative sign. Average treatment cost as well as availability of financing remain insignificant across specifications. The size of coverage coefficient

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[18]Negative sign on the coefficient of the share of LTU is clearly an indication that in the labour markets where activity rates are lower, unemployment rates are also lower. This is predominantly characteristic for the rural areas in Eastern Poland, where large so-called "hidden" unemployment and over-employment in the agricultural sector are characteristic.

Table 3: Levels - β Convergence Analysis

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
Initial unemployment $u_{i,t0}$	1.037***								0.24***	0.13***	0.28***
	(0.033)								(0.007)	(0.02)	(0.009)
Initial unemployment $u_{i,t-12}$		0.81***	0.89***	0.89***	0.89***	0.88***	0.88***	0.87***	0.75***	0.46***	-0.66***
		(0.001)	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)	(0.004)	(0.005)	(0.011)	(0.006)
% of females				-6.99***		-4.39***		-5.58***	-4.24***	-11.34***	-7.55***
				(0.36)		(0.38)		(0.45)	(0.44)	(0.61)	(0.61)
% of LTU				-0.92***		-1.18***		-2.25***	-1.45***	-1.24***	-0.86***
				(0.12)		(0.11)		(0.14)	(0.14)	(0.22)	(0.18)
% of youth				2.93***		3.26***		3.26***	2.91***	3.57***	0.96**
				(0.27)		(0.27)		(0.30)	(0.30)	(0.42)	(0.4)
% of rural				-1.47		-1.33		-1.42	-1.48	-1.32	-1.27
				(1.29)		(1.22)		(1.09)	(1.12)	(1.15)	(1.97)
% of 50+				-2.75***		-3.08***		-3.33***	-3.12***	-2.88***	-2.98***
				(0.20)		(0.20)		(0.23)	(0.23)	(0.32)	(0.29)
Inflows rate					-8.07***	-7.31***		-9.12***	-9.61***	-6.54***	-10.96***
					(0.49)	(0.48)		(0.56)	(0.57)	(0.69)	(0.72)
Outflows rate					-15.19***	-14.16***		-14.28***	-17.61***	-14.99***	-20.13***
					(0.72)	(0.68)		(0.79)	(0.81)	(1.02)	(1.00)
% of treated							-0.10*	-0.11*	-0.10*	-0.16**	-0.22***
							(0.071)	(0.074)	(0.074)	(0.09)	(0.09)
Exp. per unemployed							-0.09	-0.09	-0.09	-0.09	-0.136
							(0.07)	(0.06)	(0.065)	(0.07)	(0.08)
Exp. per treated							-0.09	-0.09	-0.085	0.017	-0.16**
							(0.06)	(0.06)	(0.06)	(0.07)	(0.07)
Time trend	No	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No of observations	374	36 694	36 694	29 185	29 186	29 185	22 719	22 719	21 211	5 983	15 228
Estimation technique											
χ^2	594.8***	2450.6***	67109.3***	39216.03***	30127.5***	40685.29***	20642.8***	31135.2***	30621.3***	11614.2***	25961.2***

Notes: GMM estimators with robust standard errors. Robust standard errors reported. Time trend, linear and squared, significant (not reported, available upon request).

Standard errors in parentheses. *, ** and *** denote statistical significance at 1%, 5% and 10% levels, respectively. Except for pooled unconditional estimation (first column), χ^2 Wald statistics highly statistically significant, *p-values* available upon request.

suggests that 1 percentage point increase in coverage results in lowering the unemployment rate by roughly 0.1-0.2 percentage points. Indeed, these are not small effects, if one considers the actual current coverage rates in Poland, *i.e.* approximately 20% of the unemployed. However, the estimator of the convergence rate remains unaffected by the inclusion of the ALMPs variables, which implies that these positive labour market effects have in fact no effect on regional unemployment rates dispersion.

Column (9) reports the same regression results as column (8) with the main difference being the inclusion of initial (December 1998) unemployment rate. Technically this plays the role of a panel specific fixed effect estimator for a constant. The estimator is significant and has a positive sign (higher unemployment *level* initially is associated with higher unemployment rates subsequently, throughout the analysis). The below unity value of $u_{i,t-12}$ coefficient is only marginally smaller than in the previous specification, which implies that *de facto* convergence is much weaker than suggested by the relatively large size of estimator from the previous estimations.

Finally, σ-convergence analysis suggested that as far as the behaviour of univariate regression residuals are concerned, lowest three decimal groups display different characteristics than the majority of the sample. Therefore, we performed the analysis as in column (9) for two separate sub-samples: three lowest decimal groups in column (10) and the reminder in column (11). This step was taken to observe if these groups would differ in the sign of influence of policy instruments on regions experiencing less hardship *versus* those who are in a more unfavourable position. As may be inferred from the estimates, convergence patterns are stronger among lowest unemployment regions, while treatment coverage is insignificant in highest unemployment regions. Interestingly, intensity of financing is insignificant in column (10) - lowest unemployment regions - while it retains negative significant coefficient in column (11) - the reminder of the sample - which could imply there is in fact less effectiveness of instruments used in more favoured regions. This is striking because one would typically associate lower unemployment with large, fast growing cities and higher level of human capital, including the level of local public employment services performance.

The β coefficient demonstrates consistently values below unity across conditional convergence specifications, which would suggest convergence. However, in fact, these results are fairly week, which is depicted in Figure (5).

Previously, Tyrowicz and Wójcik (2009) have demonstrated that no convergence (of types β, σ or stochastic) seems to be found in NUTS4 level data on unemployment in Poland. This paper extends these initial findings by incorporating labour market structures and labour market policies into the analysis of convergence patterns. Conclusions from this part of analysis suggest that even though we could find some statistical support to the hypothesis of convergence in levels, it is (i) weak, (ii) weaker among higher initial unemployment regions and (iii) independent or sometimes even negatively affected by active labour market policy measures.

Figure 5: Convergence of Unemployment Levels, 1998-2008

Source: own calculation based on registered unemployment data.

7   Conclusions

For most analyses of convergence one experiences a difficulty of no clear, counterfactual results (Boldrin and Canova 2001). The literature traditionally assumes that diversification of the use and coverage of cohesion policies provides sufficient variation to derive conclusions regarding the effectiveness of cohesion efforts. The main purpose of this paper was to analyse the effects of policy instruments on the convergence patterns of local labour markets in a transition economy. We used policy relevant NUTS4 level data, since 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.

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.

Tyrowicz and Wójcik (2009) demonstrate that diverging unemployment rates' patterns seem to be found in the data for transition countries. This paper demonstrates that much of the observed effect currently may be attributed to the lack of ALMPs effectiveness, i.e. not the consequence of transition hardship, but the "policy mistakes" made every day.

In order to inquire for the nature of the local unemployment rates evolution we employed parametric econometric techniques (convergence of levels, β-convergence) as well as nonparametric kernel density estimates (distribution dynamics). 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 for high unemployment *poviats*. In addition, data support only weak conditional β-convergence, with some evidence of asymmetry between high and low unemployment *poviats*. We cannot confirm that policies contribute to cohesion. In fact, we find that if one statistically controls for the effects of policies, divergence is in fact more apparent. Policy measures are proxied in our study by the extensiveness and intensity of active labour market policies financing (expenses per one unemployed and expenses per one person in treatment) as well as their depth (coverage of treatment).

Unfortunately, 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. Therefore, the time-span is relatively short. Consequently, our results should be interpreted with caution.

On the other hand, our findings are consistent with earlier study by Gora et al. (1996). However, our interpretation of these findings differs in a sense that we do not attempt to judge the adequateness of separate instruments used, rather the overall quality of PES performance. Naturally, weak evidence of convergence may be interpreted in wider terms of general conclusion that such measures do not work. We share the view that in order to provide such strong conclusions one should be first confident that value for money ratio is fairly rational. Currently local labour offices receive financing irrespectively of their performance. Consequently, distribution of financing does not constitute an incentive for offices to improve performance while spending the available budget is the main area of evaluation by NUTS2 authorities as well as Ministry of Labour and Social Affairs. Each of the Polish NUTS2 regions contains districts from highest unemployment groups. Should financing be geared towards alleviating the situation in most deprived regions by fostering higher effectiveness, altering the impact on convergence might be expected.

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