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ABSTRACT

Twin Peaks in Regional Unemployment and Returns to Scale in Job-Matching in the Czech Republic*

The regional distribution of unemployment rates in the Czech Republic during the transition period is shown to be characterised by twin peaks, i.e. a high and a low unemployment equilibrium. The emergence of strong regional disparities at the beginning of the 1990s can, at least partially, be explained by regionally different degrees of competition between the emerging private sector and state-owned enterprises for skilled labour and the role of on-the-job transitions on the parameters of the matching function. This study presents a formalisation of these effects and estimates empirical matching functions for a panel of labour market districts of the Czech Republic between January 1992 and July 1994. When dynamics of unemployment to job exits are taken into account and dynamic panel estimators are applied, the Czech matching function is shown to exhibit increasing returns to scale. This is consistent with the finding of multiple unemployment equilibria.

JEL Classification: E24, J64
Keywords: regional labour markets, matching functions, returns to scale, multiple unemployment equilibria, on-the-job search, job competition, Czech Republic

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NON-TECHNICAL SUMMARY

Despite overall unemployment remaining at very low levels during the early stages of the transition in the Czech Republic, the regional dispersion of unemployment rates dispersed sharply, as in other central and eastern European countries.

This study goes one step further in analysing the pattern of regional unemployment and vacancy rates in the Czech Republic. Instead of looking at the change in the regional distribution of these variables over time, it compares the dynamics and position of unemployment and vacancy rates of each labour market district within the regional distribution over several points in time applying a methodology devised by Quah (1996) in the context of convergence of growth rates among countries. While such intra-distributional dynamics confirm the increasing disparities among Czech regional unemployment rates between 1991 and 1994, the distribution is shown to be characterised by *twin-peaks*, i.e. a low-unemployment and high-unemployment equilibrium which are both stable over time. For vacancies, we find a unimodal regional distribution over time and a weak trend to convergence. Several factors, such as the heterogeneity of Czech labour market districts with respect to industrial structure or limited regional mobility may contribute to this development.

In this study, increasing returns to job-matching are proposed as one explanation for finding a bimodal distribution of regional unemployment over time. As is well known from the literature on job-search and matching theory, increasing returns to scale are a necessary condition for multiple equilibria in such models (Diamond, 1982; Pissarides, 1986). In a theoretical part, which is based on a model of job-competition by Burgess (1993), it is demonstrated how endogenous adjustments in search intensities of employed job seekers with respect to local labour market conditions alter the parameters of a reduced-form matching function and why such effects are particularly relevant in a transition economy such as the Czech Republic. Ultimately such effects may induce increasing returns to scale in job-matching.

The intuition behind these endogenous adjustments of search intensities of employed job seekers lies in the existence of a wage premium, which private sector firms offer workers in the state sector in competition for skilled labour. The existence of such wage premia in transition economies has been reported by Flanagan (1995) among others. Hence, employed and unemployed workers are assumed to sample wage offers from disjointed parts of the same wage distribution, where some high wage jobs are only offered to the employed to motivate job-to-job transitions. In addition, it is assumed that this wage premium varies according to local labour market conditions: a higher

unemployment (vacancy) rate leads to a lower (higher) wage premium offered to employed job seekers. This can be justified by stigma effects when being unemployed at a low overall unemployment rate. After deriving optimal search intensities for both types of workers, these are plugged into a *core* matching relationship, which determines the job-finding probability.

The model demonstrates, that in a transition economy, the described mechanism can lead to a result where the congestion effects of on-the-job search for the parameters of a reduced-form matching function – as in Burgess (1993) – can be reversed. Particularly, if employed job seekers adjust their search intensities very elastically to changes in local unemployment (and to a lesser extent to vacancies), it is likely to find increasing returns to scale in job-matching. Moreover, the model provides one explanation for the inherent instability of the matching function in transition economies (as found by Burda and Profit, 1996 among others).

In the empirical part of the study, matching functions are estimated based on a monthly panel of all 76 labour market districts in the Czech Republic between January 1992 and July 1994. Almost all earlier studies based on this type of data have identified constant or decreasing returns to scale for the matching process in Czech districts. In this paper, it is argued that these empirical studies have largely neglected the dynamics of unemployment to job flows. The process of screening potential workers and workplaces takes time, during which search activities for other trading partners may be suspended. And even when an employment contract is signed, the match may not become productive at the same instant. A more realistic description of labour markets is to assume that contracts fix a starting date for the employment relationship. During the time between signing the contract and starting work, an unemployed person will not be engaged in job search and a vacancy though possibly still posted will not accept further applications. This implies that the elasticities of hires with respect to unemployment and vacancies in a matching function will only gradually adjust to their long-term values. Another explanation for serial correlation in unemployment outflows is the dependence of search intensities on aggregate economic activity which shows strong serial correlation (Baker et al., 1996). Empirical matching functions applied to regional panels, neglecting such dynamics may yield seriously biased estimates of the parameters of interest which has severe implications for predicted unemployment dynamics in regional labour markets.

Therefore this study estimates dynamic matching functions in differences using instrumental variable methods (Anderson and Hsiao, 1982) and general of moment estimators devised by Arellano and Bond (1991). These techniques avoid the well known problems when estimating dynamic models with fixed effects. Also, different sets of instruments are used to check the impact of imperfect exogeneity of instruments on the estimated parameters of

the matching function. As predicted by the theoretical model, one finds a higher coefficient on unemployment – greater than one in some specifications. More importantly, constant returns to scale can be rejected in favour of increasing returns to scale in almost all specifications for the Czech matching function.

Unfortunately, no direct information on employed job-search is available for the Czech Republic, especially not at the district level. In order to relate these findings closer to the results of the theoretical model, a number of interactions are introduced into the empirical specification of the matching function, which allow for non-uniformity of slopes according to a districts relative position in the unemployment and vacancy distributions over time, as described earlier. This classification of labour market districts is obtained by a simple cluster analysis. Especially for districts in the high-unemployment equilibrium, the empirical findings are in line with the predictions of the theoretical model: a coefficient on unemployment clearly above one and a small, even negative coefficient on vacancies. Speaking in terms of the theoretical model this means, that the effect of higher unemployment or lower vacancies on the search intensity of employed workers out-weights the effect of a lower employment rate on the matching probability in these regions. The findings are mixed however: the large coefficient on unemployment in districts with higher relative vacancy rates at the outset of the transition process works against this argument. In addition, heterogeneity of slopes in the matching function is allowed for according to the number of private enterprises relative to total employment and the employment share in the service sector. Again, only in the latter case, the evidence is in line with the predictions of the model.

1. Introduction

Despite a remarkable progress in restructuring the economy and developing the private sector, unemployment in the Czech Republic has remained at surprisingly low levels compared to most other Central and Eastern European economies during the first years of the transition. Moreover the unemployment rate shows none of the persistence known from western European labour markets. Nevertheless, low aggregate unemployment rates hide the fact that the regional dispersion increased sharply during the transition period (see OECD, 1995; 1996). This study proposes increasing returns to job-matching caused by regionally disproportionate endogenous adjustments of search intensities of employed job-seekers as a response to structural changes in the local economy as one possible explanation for increased labour market disparities in a country with a high degree of labour reallocation and low level of overall unemployment.

The paper centres around the analysis of the aggregate matching function, which describes the process of workers and firms contacting each other and eventually forming employment relationships. As such, it captures informational deficiencies concerning the quality of a potential match, time-consuming and costly search, sorting and screening processes of workers and firms, as well as various forms of mismatch in labour markets due to qualificational, sectoral and regional discrepancies. In addition, the institutional environment and legal regulations -- such as the administration and efficiency of labour offices in mediating vacant jobs with job-seekers, or the generosity of unemployment benefits -- may have an influence on search behavior, and impose or alleviate frictions on the outcome of job search activities.

The specification of the matching function commonly adopted in the literature relates labour market stock variables, unemployment, possibly adding those who seek *on-the-job*, and the number of posted vacancies, as *matching factors* to the number of hires during a certain time interval, where the latter is often approximated by unemployment outflows in empirical work (see Pissarides, 1986b; Blanchard and Diamond, 1989). Hall (1977) derived a basic version of the matching function where the instantaneous number of hires is an increasing function of the number of job-seekers and vacancies, and the matching function exhibits constant returns to scale (CRTS): Pissarides (1990) demonstrates that with CRTS the vacancy-to-unemployment ratio is a sufficient statistic to determine the transition rate from unemployment to jobs. Theoretical reasoning for CRTS has found support from theoretical analyses, as

the assumption of constant returns in matching is consistent with constant unemployment rates along a steady-state growth path in theories of equilibrium unemployment. This is in line with empirical evidence of non-trending unemployment rates in the US and UK (Blanchard and Diamond, 1989; Coles and Smith, 1996).

On the other hand, theoretical studies have established the plausibility of increasing returns (IRTS) in matching due to various trading externalities resulting from endogenous adjustments of search activities of labour market participants (Diamond, 1982). Also relaxing the assumption of random search and assuming that workers and firms are able to discriminate between currently arrived job seekers and vacancies, and those who have been in the market already produces similar effects (Coles, 1994; Gregg and Petrongolo, 1997; Coles and Smith, 1998).

Another possible source for IRTS in job-matching is identified as being particularly relevant in transition economies in this study. The ability of employers of private enterprises to discriminate between job offers given to unemployed and employed job seekers creates endogenous adjustments of search intensities of employed job seekers to local labour market condition, which may render a reduced-form matching function with IRTS. Boeri (1995), Burgess (1993a,b) and Pissarides (1994) explore the role of on-the-job search for the matching process. In particular, Burgess (1993b) shows, that endogenous job competition between employed and unemployed job seekers has important consequences for returns to scale in matching and the interpretation of matching function parameters as a whole.

IRTS imply an increased matching efficiency in markets where job-reallocation and turnover is high, limiting their impact on the equilibrium unemployment rate (Courtney, 1992). Moreover, Pissarides (1986) identifies IRTS in matching as a necessary condition for the existence of multiple unemployment equilibria. Hence, modelling and estimation methods of matching functions have important implications for unemployment dynamics in a macroeconomic framework. This is particularly true for regional labour market dynamics where the existence of multiple equilibria may give scope to permanent effects of regional or active labour market policies.

The scope of this paper is to show how endogenous adjustments in search intensities of employed job seekers with respect to local labour market conditions may be responsible for increasing returns to matching, when potential employers are allowed to discriminate between

employed and unemployed job seekers. This effect is particularly relevant for labour markets in Central and Eastern European transition economies, where emerging private enterprises compete with state enterprises for skilled labour. The empirical part of the paper explores regional labour market dynamics in the Czech Republic over the transformation period and presents estimates of matching functions from a monthly panel of unemployment, vacancies and unemployment-to-job transitions for 76 labour market districts between January 1992 and July 1994 taking into account the dynamic properties of unemployment-to-job transitions.

The results show that, in contrast to previous evidence, the emerging pattern of regional unemployment in the Czech Republic is consistent with the finding increasing returns to job-matching.

The following section illustrates regional dynamics of unemployment and vacancy rates in the Czech Republic over the transformation process applying nonparametric smoothing techniques. Section 3 introduces a stylised model of job competition establishing the plausibility of increasing returns to job-matching in a transition economy like the Czech Republic. In section 4, I highlight econometric problems involved in estimating dynamic specifications of the matching function with panel data and apply GMM techniques. The robustness of matching function estimates across various specifications is discussed, particularly with respect to the validity of instruments. Section 5 is a tentative analysis of the effects described in the model of section 3, and section 6 concludes.

2. Regional Unemployment-Vacancy Dynamics in the Czech Republic

Before analysing the properties of the matching function in the Czech Republic, "the outcome" of the job-matching process -- the prevailing regional dispersion of unemployment and unfilled vacancies -- is explored. The phenomenon of low overall registered unemployment combined with a strong increase of regional disparities in the Czech Republic over the transformation period is widely documented and discussed in the literature (see Boeri, 1994; Munich et al. 1995).

The nonparametric density estimates for the dispersion of relative deviations of districts' unemployment rates from the overall average of all 76 labour market districts of the

Czech Republic for each month between December 1990 and June 1994 in Figure 2.1 confirm this evidence[1]. It reveals that while a large fraction of district unemployment rates is concentrated around a single peak until 1991, the cross-sectional distribution becomes much flatter and skewed to the right in subsequent years. The pattern of increasing regional disparities does not carry over to the demand side of the labour market. As displayed in Figure 2.2, there is no obvious trend in the regional cross-section distribution of vacancy rates despite some seasonal variation. Table 2.1 confirms the finding of diverging regional unemployment rates. Additionally, coefficients of variation reveal a converging trend in unfilled vacancies.

Table 2.1 Evolution of Distributions over the Transition Period

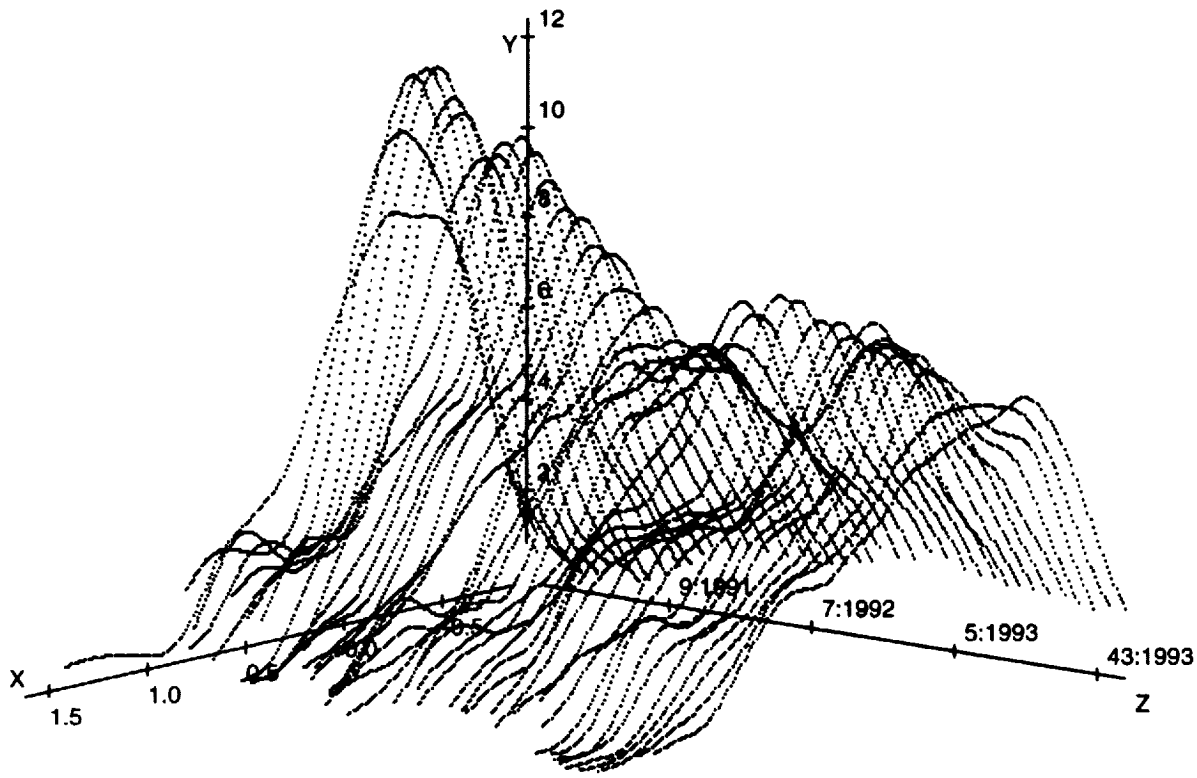
	Unemployment Rate Deviations					
	Min	1. Quartile	Median	3. Quartile	Max	C.V.
6:1991	-0.84	-0.25	-0.04	0.22	1.00	0.38
6:1992	-0.86	-0.31	-0.02	0.46	1.49	0.49
6:1993	-0.90	-0.35	-0.04	0.55	1.39	0.53
6:1994	-0.92	-0.39	-0.03	0.53	1.46	0.54

	Vacancy Rate Deviations					
	Min	1. Quartile	Median	3. Quartile	Max	C.V.
6:1991	-0.86	-0.60	-0.29	0.06	1.40	0.63
6:1992	-0.79	-0.44	-0.18	0.10	1.08	0.52
6:1993	-0.84	-0.45	-0.18	0.21	1.03	0.52
6:1994	-0.70	-0.34	-0.06	0.23	1.23	0.40

The evolution of relative deviations of unemployment rates suggest that some districts were hit harder by the transformation process. Regions like Northern Moravia or parts of Northern Bohemia experienced comparably strong increases in unemployment as a result of reallocation of resources and labour shedding in industries which were given priority in the centrally planned economy. The stability of regional unemployment diffusion after 1991 implies a limited role of labour mobility in overcoming such regional disequilibria, probably due to shortages in rental housing, and increasing cost of public transport. The convergence trend

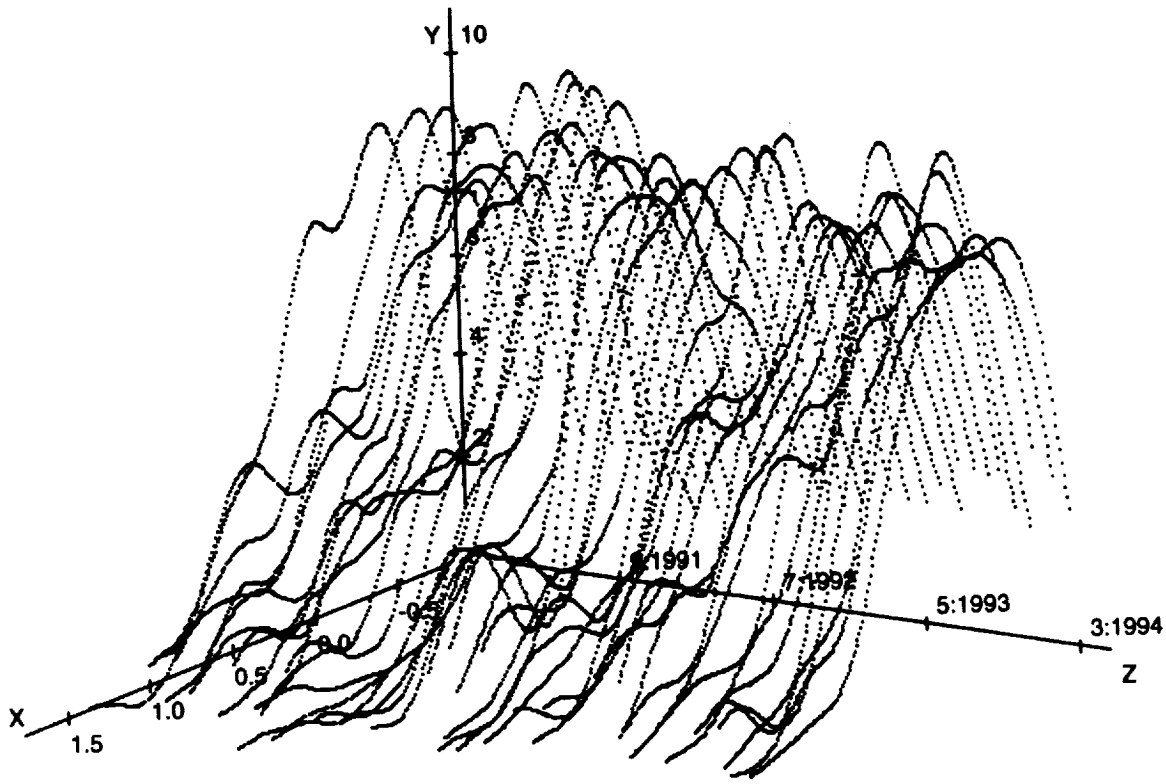
[1] Due to data limitations unemployment and vacancy rates are calculated on the basis of labour force figures from yearend 1992. The data used in this study are registered unemployment, vacancies, and unemployment-to-jobs exits collected from all labour market districts in the Czech Republic, and provided by the Czech Ministry of Labour and Social Affairs. I am grateful to Miroslav Pribyl for providing the data. All nonparametric estimations were done using XploRe. See Härdle (1990). Note that a value of one on the x-axis indicates a local unemployment rate twice as high as the national mean.

Figure 2.1 Dispersion of unemployment rates in the Czech Republic, 12:1990-6:1994



X: Regional deviation from national unemployment rate
Y: Frequency
Z: 12:1990 - 6:1994, Bandwidth h= 0.35

Figure 2.2 Dispersion of vacancy rates in the Czech Republic, 12:1990-6:1994



X: Regional deviation from national vacancy rate
Y: Frequency
Z: 12:1990 - 6:1994, Bandwidth h = 0.35

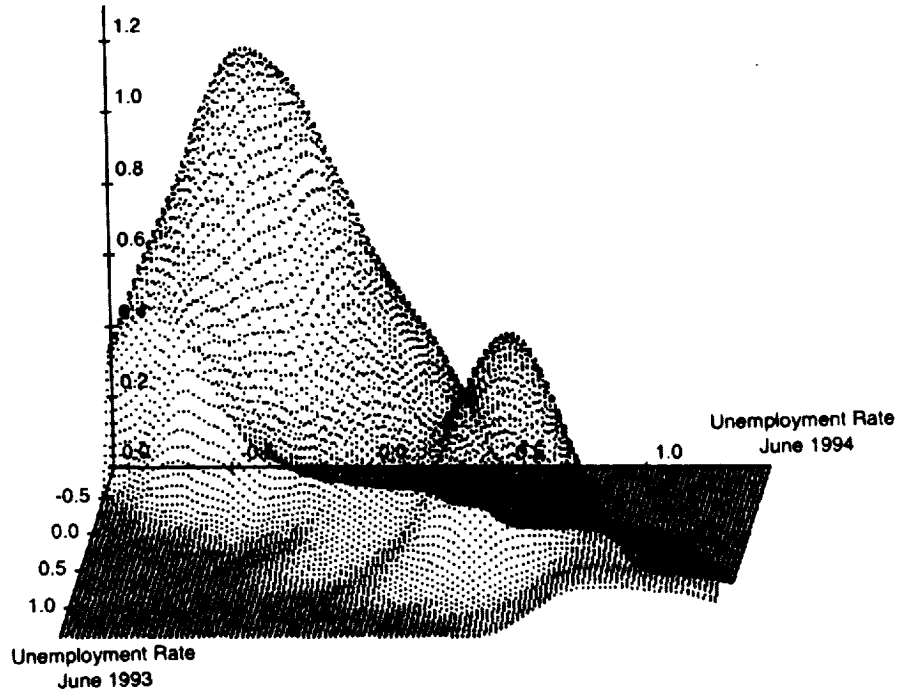
in the regional distribution of vacancies may be due to a proportionate emergence in small business dynamics or capital mobility even to depressed regions (Burda and Profit, 1996).

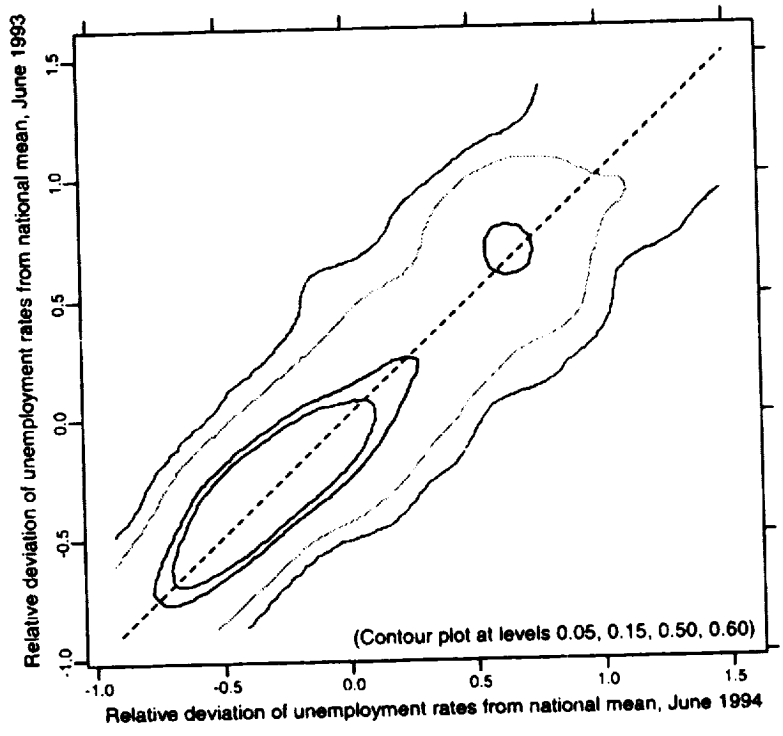
If the dynamics of regional distributions are considered over time only, important movements within the distribution are neglected. Identifying the relative position of a district's unemployment and vacancy rate within the regional distribution over two or more points in time is crucial to the understanding of the forces driving the transition process to a market economy. Such intra-distribution dynamics may evolve as the result of *mobility* or *churning* of districts within the distributions, possibly due to properties of the matching technology.[2] Quah (1996) has applied a methodology to explore transition patterns between cross-sectional densities through different points in time. He interprets such bivariate distributions as continuous versions of a Markov transition probability matrix. Suppose that, in a discrete setting, K classes of unemployment (or vacancy) rate deviations are given, $k = 1,...,K$ and transition probabilities for districts moving between or within these classes during the time interval t and $t+n$ can be calculated. If $K \rightarrow \infty$ and each class becomes infinitesimally small, one obtains a continuous transition function from one labour market state in period t to any labour market state in $t+n$.

The intuition behind this methodology is demonstrated in the top panel of Figure 2.3, which shows intra-distribution transitions for regional unemployment rate deviations between June 1993 and June 1994 in the Czech Republic. The bottom panel in Figure 2.3 shows the corresponding contour plot. Two of the three axes show relative unemployment rates compared to the national mean in two points in time. The bivariate density shows the dynamics of relative unemployment in each labor market districts within the regional distribution over time. Considering a district with a specific relative deviation from the national unemployment rate Δu_t and cutting through the distribution parallel to the $t+n$ axis gives the marginal density $g(\Delta u_{t+n} | \Delta u_t)$, which can be interpreted as a measure of the conditional probability of a transition to another position in the regional unemployment distribution.

If the distributions were perfectly stable over time, the contour plot of the bivariate distribution would degenerate to the main diagonal as illustrated in Figure 2.4. This is the case of *full distributional persistence* of regional unemployment rates over time. With complete convergence among districts, the ridge of the two-dimensional distribution should form par-

Figure 2.3 One-year transitions in the cross-district distribution of relative unemployment rate deviations, June 1993 - June 1994





allel to the t-axis, whereas in the case of divergence, the ridge is a horizontal line. Finally, different modes along the main diagonal indicate the existence of *convergence clubs* or multiple equilibria among regional unemployment or vacancy rates.

Figures 2.5 and 2.6 present the results of a nonparametric kernel estimator for the bivariate densities of regional unemployment and vacancy rate deviations at the beginning of the transformation process (June 1991) and three years later (June 1994)[3].

In Figure 2.5, the main peak of the bivariate distribution lies on the main diagonal slightly below the national unemployment rate. The ridge of the distribution is clearly flatter than the 45°-line indicating a diverging trend in the regional distribution of unemployment consistent with the evidence form inter-distribution unemployment dynamics in Figure 2.1. In addition, I find a *twin-peaked* distribution with a second persistent mode which gathers districts showing an unemployment rate more than 50% above the national rate. Figure 2.1 demonstrated that regional unemployment disparities mainly emerged at the outset of the transformation. Figure 2.3 shows one-year transitions between June 1993 and June 1994 and supports this impression: the bivariate distribution of relative unemployment rates is fairly stable along the main diagonal between 1993 and 1994. Both panels clearly support the bimodality (*twin peaks*) and *distributional persistence* of the two unemployment equilibria during the respective period. The evidence on the dynamic evolution of vacancy rates in Figure 2.6 shows that the distribution is single peaked with a converging pattern (i.e. a vertical ridge).

Such dynamic "sorting" processes towards district steady-states of high and low unemployment across regions can have a variety of explanations, such as the heterogeneity of districts with respect to industrial structure, or limited mobility of the labour force. A model of job competition and endogenous job search intensity in Section 3 will propose that increasing returns to scale in job-matching may also be a candidate to explain labour market disparities in the Czech Republic.

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[2] Bianchi and Zoega (1998) analyse intra-distributional dynamics on regional labour markets for Britain.

[3] The densities were estimated using a quartic kernel. Analytically correct bandwidths were estimated using Silverman's rule of thumb (see Silverman, 1986). In practice, these bandwidths rendered density estimates which were considerably oversmoothed; hence, the results reported below were estimated with bandwidths of 0.35. Contour line levels are given at the bottom of the figures. Bianchi (1997) developed a nonparametric test for multimodality based on critical bandwidths but only for univariate distributions.

Figure 2.4 Convergence, divergence and persistence of the cross-district distributions

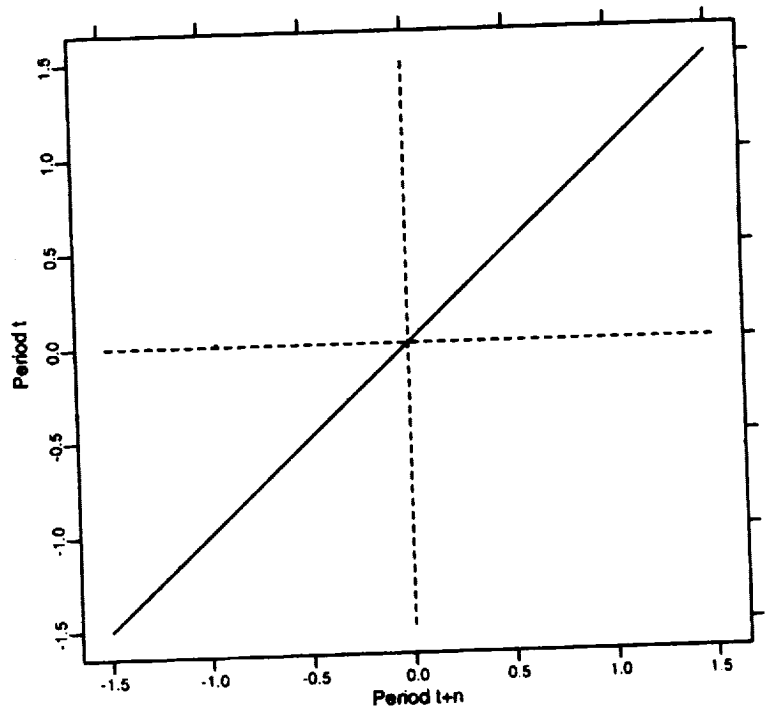


Figure 2.5 Three-year transitions in the cross-district distribution of relative unemployment rate deviations, June 1991 - June 1994

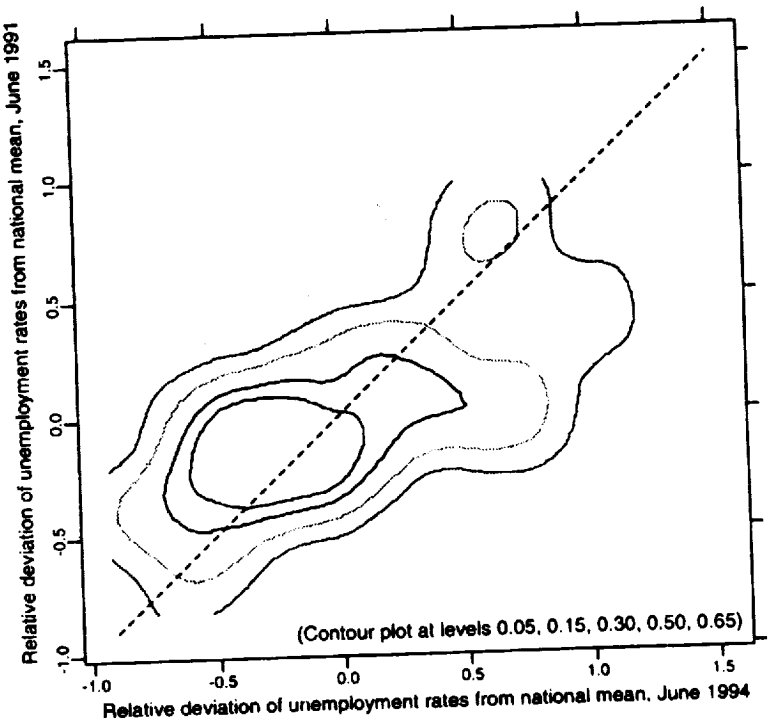
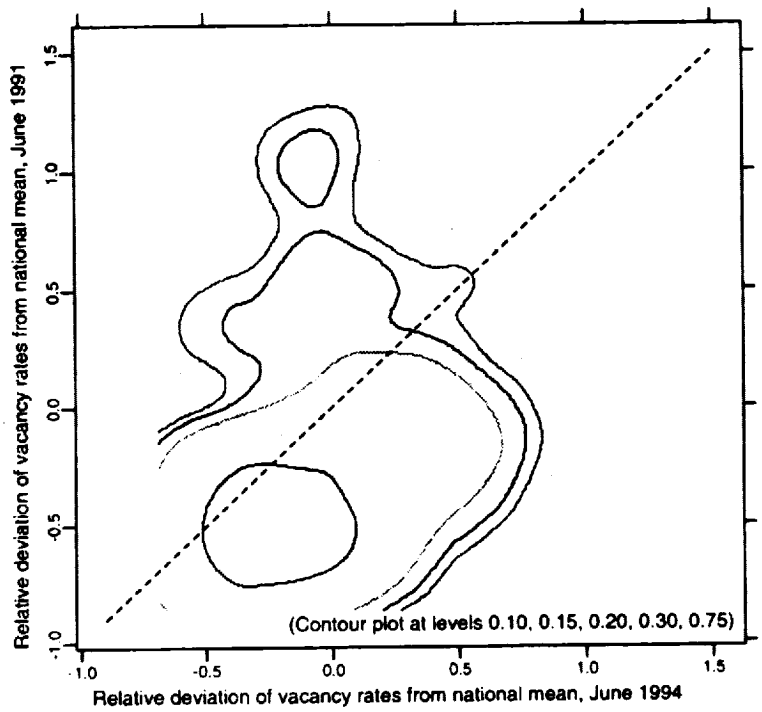


Figure 2.6 Three-year transitions in the cross-district distribution of relative vacancy rate deviations, June 1991 - June 1994



3. A Stylised Model of Job Competition and Endogenous Search Intensity of Employed Job Seekers

Trading externalities in job-matching can either originate in the mechanical component of the matching process, in feedback effects working through search intensities, or in endogenous effects in the matching technology (relating to institutional characteristics of the labour market and the availability of informational services)[4].

The model presented here introduces an externality of the second type and describes endogenous effects in search intensities of labour market participants, namely those who search on-the-job, and their impact on the parameters of empirical matching functions. It generalises the model in Burgess' (1993a,b) which describes interactions in search intensities of unemployed and employed workers, where the likelihood of finding a job depends on a job offer probability which is given to both groups of labour market participants as well as on the shape of wage offer distributions. Burgess' model has two main implications: on-the-job search raises the overall number of matches (which can be interpreted as an increase in labour demand), but on the other hand more job search on part of the employed crowds out unemployed job seekers, rendering an elasticity of the job finding probability with respect to changes in the number of total matches of smaller than one. Second, given the validity of this job competition model, the parameters estimated from a standard matching function cannot be interpreted in a usual way, but rather as the outcome of a reduced-form relationship. Therefore Burgess (1993a,b) delivers an explanation, why empirical estimates of the matching function relationship which ignore on-the-job search underestimate the returns to scale of the underlying matching function (see also Boeri, 1995).

In this partial equilibrium of the model, however, the first argument crucially depends on the assumption that the process of vacancy creation is exogenous. Crowding-out effects of job-to-job movements on unemployment transitions hinge on the assumption that vacancies

[4] See Blanchard and Diamond (1992) and Courtney (1992) for a decomposition of the matching function. Examples of externalities which work directly through the aggregate matching technology are endogenous increases in the effectiveness of labour market intermediation and information services, the provision of active employment policies (Boeri and Burda, 1996), the degree of specialisation in thick labour markets (Hall, 1989), or the intensity of reallocation (Blanchard and Diamond, 1992).

left by successful employed job seekers are destroyed[5]. In the following, I accommodate Burgess' model to account for characteristics of a labour market in transition. This will reveal an additional effect of search on-the-job, which works against the congestion effect and implies the possibility that an empirical matching function which only includes unemployed searchers actually has increasing returns to scale. The empirical implications of this model are examined in section 4.

The economic background which motivates the model is characterised by intensive reallocation processes combined with limited labour mobility in Central and Eastern European economies that have caused tightness in booming local labour markets and excess labour supply in others[6]. Such regional labour market mismatch has led to significant wage differentials between state-owned and private enterprises (see Flanagan, 1995), which also exhibit large regional variation. In contrast to Burgess (1993b), it is assumed that employed and unemployed job seekers – despite sampling from the same (known) wage offer distribution – obtain offers from partially disjoint ranges of the wage distribution. The reasoning behind this assumption is that potential employers -- in particular those from the emerging private sector -- discriminate between types of job seekers and offer a wage premium to attract skilled workers from the state enterprises. In addition, I assume that the size of wage premium depends on local labour market conditions: when unemployment is low, unemployed workers are stigmatised by employers as being bad quality workers. Consequently employers are willing to offer a high wage premium to attract employed job seekers. However, when local unemployment is high, the stigmatisation effect is less severe and hence the wage premium decreases (see Omuri (1997) for empirical evidence supporting this assumption).

Turning to the empirical evidence on the private-state-owned enterprise wage premium in Central and Eastern transition economies, Flek (1996) emphasises the large share of job-to-job transitions in total labour reallocation and argues that the existence of a wage premium offered by expanding private enterprises is the result of the educational composition of the unemployment pool together with continued labour hoarding of state-owned or privatised state enterprises. Flanagan (1995) presents data from the Czech Survey of Economic Expectations and Attitudes, showing that in November 1994, the state sector comprises 40%, the pri-

[5] For a model which include the supply decision of firms with respect to vacancy creation, see Pissarides (1994) and Boeri (1999).
[6] See Section III.3.3 for an analysis of regional unemployment and vacancy dynamics in the Czech Republic.

vate sector 28% and privatised state enterprises 32% of total employment. Moreover, earnings of full-time employees in the private sector are roughly 25% above those paid in the state sector. After controlling for human capital variables (education, experience and sex) the wage differential even rises to 46%[7]. Vecerník (1995) shows evidence based on the same data suggesting that the earnings gap is mainly due to self-employed workers whose earnings were almost 60% above average earnings. However, earnings of workers in other private enterprises are still more than 15% above those in the state sector.

Although the model is partial equilibrium, I relax the exogeneity assumption for vacancy creation by stating that the range of the wage offer distribution employed and unemployed job seekers sample from differs for both types. It would be possible to extend the model towards a general equilibrium setting along the lines of Burdett and Mortensen (1998). However, since the externality to be derived here is related to the heterogeneity of workers, its effect would not disappear in a general equilibrium. Due to this heterogeneity the upper support of the distribution, from which employed job seekers obtain wage offers, is greater than for the unemployed. The difference between the upper support of the wage offer distribution for employed and unemployed searchers constitutes the wage premium, which depends on local labour market conditions. Hence, the wage offer distribution is assumed to be identically shaped for both types of job seekers except for different truncation values, i.e. different maximum available wage offers[8]. The expected benefit from job search is given by

$$B(w_i, \overline{w}_i) = \mu\sigma_i \int_{w_i}^{\overline{w}_i} \left[V(\omega) - V(w_i)\right] dF(\omega), \qquad i = e, u, \tag{1}$$

where $\mu(= m/s)$ is the *base* job offer probability, which is equal to the ratio of job-matches m to the number of total job seekers s. σ_i is the search intensity of a type i worer, either employed (e) or unemployed (u). The labour force is assumed to be given. In contrast to Boeri

[7] Flanagan (1995) does not control for selectivity bias. OECD (1995) reports smaller or even negative wage differentials between private and state-owned enterprises, which may be due to a composition bias from three sources: (1) whereas private enterprises are mainly created in services, state-owned enterprises consist mostly of industry paying relatively higher wages. (2) since official wage statistics only consider workers of firms above 25 employees they do not cover most emerging private enterprises which are mainly of very small size, and (3) small firms were exempted from wage controls agreed in the Tripartite Commission at the beginning of the transition process.

[8] An alternative strategy would be to assume equal shapes of offer distributions for both types with a positive shift parameter for those searching on-the-job. See Mortensen (1986) for the effect of a shift in mean in the wage offer distribution.

(1995,1999) job finding probabilities are equal among employed and unemployed, and independent of unemployment duration. $V(\omega)$ is the value function of the state characterised by a payoff ω with $V'(\omega)>0$ and $V''(\omega)<0$[9], and $F(\omega)$ is the cumulative wage offer distribution, which is exogenously given. The range of the distribution is bounded from below by w_i which is equal to the unemployment income b or current income w if employed. The upper support, i.e. maximum available wage offer $\overline{w}_i$ also differs by type: (private) firms offer a positive wage premium to attract employed job seekers (from the state sector). These assumptions generate segments of the wage offer distribution which are characterised by different degrees of job competition, depending on the value of w. Figure 3.1 shows that wages in segment II will only be offered to employed job seekers, whereas segment I is the relevant region where Burgess' job competition model applies. In this part of the wage offer distribution job search activity of employed workers crowds out unemployed job seekers.

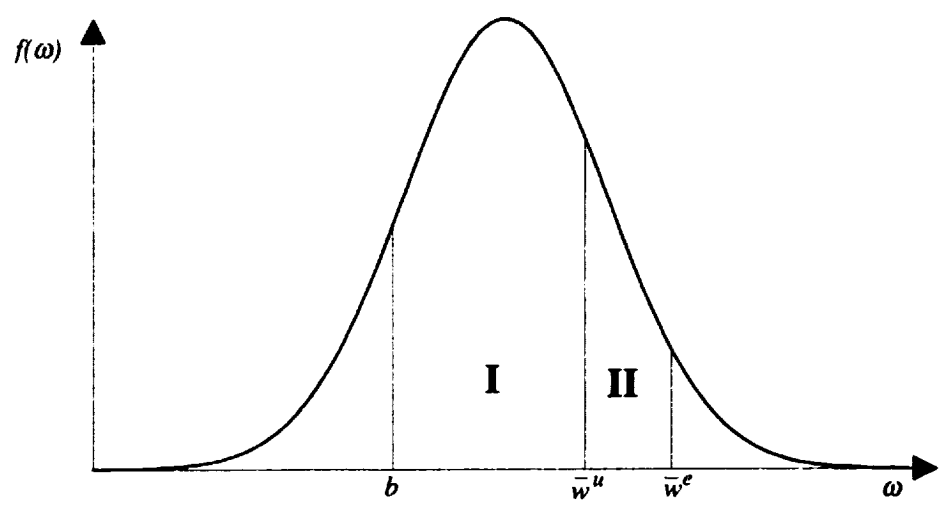


Figure 3.1: Segments of the wage offer distribution

Whereas the acceptance decision of unemployed job seekers is independent of local labour market conditions, suppose that the upper support for wage offers to employed job seekers (the wage premium) is a function of local labour market conditions, the vacancy rate v and the unemployment rate u given a constant labour force. Hence the highest wage offer to employed job seekers is given by $\overline{w}_e = \overline{w}_u + \delta(u,v)$, with $\delta(0,v), \delta(u,0) \geq 0$, $\delta'_u < 0$ and

[9] This formulation implicitly assumes a non-sequential search strategy of job-seekers.

$\delta_v' > 0$. The higher the number of unemployed (vacancies) in a labour market, the lower (higher) the wage premium potential employers are prepared to offer employed job seekers to motivate job-to-job transitions[10]. This can be explained by stigmatisation effects as mentioned before.

Boeri (1995) shows that such a behaviour may indeed play an important role in transition economies. In a cross-country comparison of workers searching on-the-job, which includes mature market and transition economies, he compares the fraction of those being pushed to search, e.g. through the expectation of losing their job, and those who are motivated to search by the prospect of a better paid job. Boeri's analysis shows that (1) the fraction of job searchers being pulled correlates with labour market conditions and (2) this fraction is the highest in the Czech Republic.

Employed and unemployed workers choose their optimal search intensities σ_i to maximise their net present value from job search, trading-off higher search costs against a higher probability of receiving an acceptable wage offer. Similar to Burgess (1993b), I model the offer arrival rate of a type i worker as $\mu\sigma_i$ and search cost as $c_i = c(\sigma_i)$, with $c_i' > 0$, $c_i'' > 0$ (see Pissarides, 1990, and Mortensen, 1986). Equating marginal benefits to marginal costs one obtains

$$c'(\sigma_e^*(\mu, w, \delta)) = \mu \int_{w}^{\bar{w}_u + \delta(u,v)} [V(\omega) - V(w)] dF(\omega) \tag{2}$$

for employed and

$$c'(\sigma_u^*(\mu, b)) = \mu \int_{b}^{\bar{w}_u} [V(\omega) - V(b)] dF(\omega) \tag{3}$$

for unemployed workers given that they search at all. In equilibrium, optimal search intensities σ_e^* and σ_u^* depend on the base offer probability and the reservation wage for the respective job seeker type. In addition, note the search intensity of the employed depends on the degree of labour market tightness.

[10] Van Ours (1995) analyses the degree job competition of employed and unemployed job seekers with respect to the choice and intensity of use of different recruitment channels. A similar argument applies here: it is assumed that within a certain range of the wage offer distributions only employed job seekers are offered jobs, which may coincide with specific recruitment channels not accessible to unemployed.

The total number of job seekers s in a labour market relative to the labour force is given by the pools of employed and unemployed job seekers weighted by their search intensities,

$$(4) \qquad s = u\sigma_u^*(\mu, b) + (1-u)\bar{\sigma}_e^*, \qquad \bar{\sigma}_e^* = \int_b^{w_0} \sigma_e^*(\mu, \omega, \delta) dF(\omega),$$

where w_0 is defined by $\sigma_e^*(w_0) = 0$. Note that for employed job seekers the average optimal search intensity has to be adopted since individual search intensities depend the current individual wage. From (4), the fact that $\mu = m/s$, and assuming a standard Cobb-Douglas specification with constant returns to scale for the *core* matching technology $m = \pi s^{1-\alpha} v^{\alpha}$, the job finding probability becomes

$$(5) \qquad \mu = m(s,v)/s(m,u,v) = \pi\left(\frac{v}{s}\right)^{\alpha}.$$

By applying the implicit function theorem, the following expressions can be derived for the matching parameters of interest. First consider elasticity of the base job offer probability with respect to the unemployment rate

$$(6) \qquad \eta_{\mu u} = \frac{\alpha\left\{\frac{u}{s}\left(\bar{\sigma}_e^* - \sigma_u^*\right) - \frac{1-u}{s}\eta_{\bar{\sigma}_e^*, u}\bar{\sigma}_e^*\right\}}{1 + \alpha\left[\frac{u}{s}\sigma_u^*\eta_{\sigma_u^*, \mu} + \frac{1-u}{s}\eta_{\bar{\sigma}_e^*, \mu}\bar{\sigma}_e^*\right]} \gtreqless 0,$$

where η_{xy} is the elasticity of x with respect to changes in y. Without the possibility of potential employers to discriminate between employed and unemployed job seekers, $\eta_{\bar{\sigma}_e^*, u} = 0$, expression (6) coincides with the finding of Burgess (1993b), and since $\sigma_u^* > \bar{\sigma}_e^*$ (Mortensen, 1986), $\eta_{\mu u} < 0$. An increase in the unemployment rate depresses the base offer rate, since the number of contacts between job seekers and potential employers increases by more than the number of matches, given the number of vacancies. In empirical matching functions, regressing the log number of hires on log levels of unemployment and vacancies (see Blanchard and Diamond, 1989), this effect produces a coefficient on unemployment of less than unity.

By allowing for the possibility of discrimination of wage offers between employed and unemployed job seekers, and endogenizing search intensities of employed job seekers with

respect to labour market conditions, the sign of $\eta_{\mu u}$ may becomes ambiguous depending on the sign of $\eta_{\bar{\sigma}_e^* u}$. For a high proportion of job search among the employed $\bar{\sigma}_e^*$ and a large negative elasticity of employed job search intensities with respect to local labour market conditions $\eta_{\bar{\sigma}_e^* u}$, $\eta_{\mu u}$ may even become positive, implying the possibility of increasing returns to scale in the reduced-form matching function. This formalises the effect described in Baker et al. (1996).

To reveal the forces driving $\eta_{\bar{\sigma}_e^* u}$, suppose for simplicity that the search costs have the form $c_i = 0.5\sigma_i^2$, $i = e,u$, and workers consider the base offer probability μ as given. Hence

$$\left.\frac{\partial \sigma_e^*(\mu,\omega,\delta)}{\partial u}\right|_{\mu=\bar{\mu}} = \mu[V(\bar{w}_u + \delta(u,v)) - V(\omega)]\delta_u \tag{7}$$

$$= \mu \Delta V(\omega)\frac{\delta}{u}\eta_{\delta u} \leq 0,$$

with $\Delta V(\omega) \equiv [V(\bar{w}_u + \delta(u,v)) - V(\omega)]$ and

$$\eta_{\bar{\sigma}_e^* u} = \frac{u}{\bar{\sigma}_e^*}\int_b^{w_0} \frac{\partial \sigma_e^*(\mu,\omega,\delta)}{\partial u} dF(\omega) \tag{8}$$

$$= \frac{\mu\delta}{\bar{\sigma}_e^*}\eta_{\delta u}\int_b^{w_0} \Delta V(\omega) dF(\omega) \leq 0$$

From (8) it is obvious that the elasticity of the average search intensity of employed workers with respect to the unemployment rate $\eta_{\bar{\sigma}_e^* u}$ is unambiguously negative, and depends on the absolute level of the wage premium and its elasticity with respect to unemployment. Plugging this result into (6) reveals, that a higher wage premium offered to those searching on-the-job – due to lower local unemployment – increases the elasticity of the base offer probability with respect the unemployment rate. This is one possible source of increasing returns to matching.

Similarly, a change in vacancy rates at a given unemployment rate changes the base offer probability according to

$$(9) \qquad \eta_{\mu v} = \frac{\alpha\left[1 - \frac{1-u}{s}\eta_{\bar{\sigma}_e^* v}\bar{\sigma}_e^*\right]}{1+\alpha\left[\frac{u}{s}\sigma_u^*\eta_{\sigma_u^*\mu} + \frac{1-u}{s}\eta_{\bar{\sigma}_e^*\mu}\bar{\sigma}_e^*\right]} \gtreqless 0,$$

where

$$(10) \qquad \left.\frac{\partial\sigma_e^*(\mu,\omega,\delta)}{\partial v}\right|_{\mu=\bar{\mu}} = \mu\Delta V(\omega)\frac{\delta}{v}\eta_{\delta v} \geq 0,$$

and hence

$$(11) \qquad \eta_{\bar{\sigma}_e^* v} = \frac{\mu\delta}{\bar{\sigma}_e^*}\eta_{\delta v}\int_b^{w_0}\Delta V(\omega)dF(\omega) \geq 0.$$

Equation (11) shows that a higher maximum wage premium has a dampening effect on the elasticity of base offer probability with respect to vacancies. In contrast to Burgess' (1993b) findings, an increase in vacancy rates may even decrease the probability of obtaining a job offer, if wage premia are sufficiently high to induce a strong positive effect on average search intensities of employed job seekers.

This stylised models illustrates the conditions which indicate, how endogenous effects in the search intensity of employed job seekers affects the estimates of a reduced-form matching function in a transition economy like the Czech Republic. The results and implication for the empirical sections can be summarised as follows:

(1) In contrast to a mature labour market, for newly created labour markets in Central and Eastern transition economies on-the-job search does not necessarily create a negative congestion externality on the outflow probability of unemployed workers. The reason is that the emerging private sector absorbs (skilled) labour from the state sector by offering a two-tier wage policy which discriminates against unemployed job seekers. The wage setting behaviour implies that some job offers go exclusively to employed workers. This creates a positive externality on the transition probability of unemployed workers.

(2) Since the wage premium offered to the employed depends on local labour market conditions, the model indicates a potential source of instability in the empirical matching function in the Czech Republic, due to diverging economic and ownership structures across regions and the adjustments in search behaviour. Note that this is not a criticism in the matching the-

ory itself but builds on the notion of endogenous adjustments of job seekers' search intensities which cannot be observed in empirical applications.

(3) The model predicts a large elasticity of job matches with respect to unemployment and a small elasticity with respect to vacancies in a transition economy. Note that the net effect on returns to scale depends on the sensibility of the search intensity of employed workers with respect to changes in unemployment relative to changes in vacancies. This in turn depends on the way firms adjust the wage premium in reaction to changes in labour market conditions. If a uniform increase in unemployment and vacancies induces firms to lower their wage offers to employed job seekers, the model predicts increasing returns to scale in a reduced-form matching function.

(4) In comparing the matching process across different regions, returns to scale will differ, not only as a result of institutional and geographic characteristics but also by regional labour market conditions. Expression (6) implies that the elasticity of the unemployment exit probability declines with the level of unemployment. However, if the elasticity of the search intensity of employed is sufficiently large, the effect may be reversed.

Section 4 estimates matching functions from a panel of Czech labour market districts over the transition period taking into account the dynamic properties of unemployment-to-job exits, and testing for returns to scale in job-matching. Unfortunately, direct information on private to state-owned enterprise wage premia is unavailable. Therefore, I use the information attained from the analysis of Czech labour market dynamics in section 2 to approximate the impact of the intensity of structural change and on-the-job search in local labour markets in section 5.

4. Consistent Estimation of Regional Czech Matching Functions with Panel Data

Modelling endogenous adjustments in search intensities in the previous section has demonstrated that the assumption of CRTS in job-matching is not necessarily justified when the behaviour of employed job seekers is taken into account. However, the majority of empirical studies have not rejected the hypothesis of constant returns to scale in job-matching. Table 4.1 provides a selection of recent returns to scale estimates from matching functions for vari-

ous countries, time periods and data sets. Some studies have however challenged this view and argue that standard estimation procedures and specifications may render biased estimates of underlying elasticities of matches with respect to unemployment and vacancy changes. A first argument relates to the notion of heterogeneity of pools of job-seekers and job offers. Coles (1994) and Coles and Smith (1998) drop the assumption of pure random search. They argue that, if no successful match is formed, agents only sample through currently arrived job offers or job candidates in subsequent periods. Hence, a correctly specified matching function implies a reduced form where hirings are a function of not only stocks of job-seekers and firms but also of inflows of new job-seekers and vacancies. Other studies question the relevance of the Cobb-Douglas technology of empirical matching functions and analyse the effects of functional misspecification on returns to scale estimates[11]. Aggregation over space, sectors, or time possibly also biases matching function parameters downwards. Anderson and Burgess (1995) use a regional panel of US labour market data at MSE level and find slightly increasing returns to scale. Burda and Profit (1997) demonstrate that a matching function in local labour markets, which considers the importance of spatial spillovers through job-seekers and recruitment activities of firms from other regions, does not necessarily exhibit CRTS. Burdett, Coles, and van Ours (1994) argue that standard estimates of matching parameters may underestimate the underlying coefficients as a result of temporal aggregation.

Another possible source of misspecification in the analysis of matching functions, especially when estimated with regional panel data, arises from neglecting the dynamic properties of unemployment outflows. Estimation results for Czech labour markets will demonstrate that unemployment-to-job flows are serially correlated, even after controlling for unemployment and vacancy stocks at the beginning of the period. Matches between job-seekers and firms do not occur instantaneously. The process of screening potential workers and workplaces takes time, during which search activities for other *trading* partners may be suspended. And even when an employment contract is signed, the match may not become productive at the same instant. A more realistic description of labour markets is to assume that contracts fix

[11] Warren (1996) generalizes the functional form to a more flexible trans-log specification and finds support of locally IRTS in US manufacturing during the 1970s. Using the same data set, Fox (1996) additionally emphasises the necessity of modelling "technical progress" in matching. He finds that returns to scale estimates crucially depend on the functional form assumptions. Storer (1994) applies nonparametric spline techniques, and also stresses the importance of analyzing the functional form of the aggregate matching function, but does not explicitly analyse returns-to-scale.

a starting date for the employment relationship. During the time between signing the contract and starting work, an unemployed person will not be engaged in job search and a vacancy though possibly still posted will not accept further applications. This implies that the elasticities of hires with respect to unemployment and vacancies in a matching function will only gradually adjust to their long-term values. Another explanation for serial correlation in unemployment outflows is the dependence of search intensities on aggregate economic activity which shows strong serial correlation (Baker, et al., 1996). Empirical matching functions applied to regional panels, neglecting such dynamics may yield seriously biased estimates of the parameters of interest and have severe implications for predicted unemployment dynamics in regional labour markets[12].

Table 4.1 Comparison of the Returns-to-scale Estimates in Matching Functions

	Country, Period	Data	Estimation Method	RTS
Anderson and Burgess (1995)	US, 1978-1984	regional panel	LSDV	IRTS
Blanchard and Diamond (1989)	US, 1968-1981	time-series	OLS, NLS, IV	CRTS, (IRTS)
Boeri (1994)	CEECs, ~ 1991-1993	regional panel	LSDV & random effects	DRTS, CRTS
Burda (1994)	East Ger., 1990-1992 CR, 1990-1992	regional panel	LSDV	CRTS DRTS
Burda and Profit (1996)	CR, 1992-1994	regional panel	LSDV	CRTS, DRTS
Coles and Smith (1996)	UK, March 1987	cross-section	OLS	CRTS
Fox (1996)	US, 1969-1974	time-series	OLS	CRTS, IRTS
Gorter and van Ours (1994)	NL, 1980-1993	regional panel	NLS	CRTS
Storer (1994)	CAN, 1972-1978	regional panel	spline regr.	--
Warren (1996)	US, 1969-1974	time-series	OLS, IV	IRTS

Keys: LSDV: least squares dummy variable model, NLS: nonlinear least squares, IV: instrumental variable models.

[12] A recent study by Münch et al. (1998) applies similar techniques as those applied in this study and also find increasing returns to scale in the job-matching in the Czech Republic.

Increased availability of regional and international panel data sets has allowed the identification of cross-section effects to control for unobserved heterogeneity in the data. In contrast to typical microeconomic panel data, macroeconomic panels often have much larger time-series dimensions. Analyses of such data sets have been widely applied to the field of economic growth and convergence between countries and regions, but also to labour markets, especially to the estimation of matching functions[13]. However, the latter category of studies largely ignores the dynamic properties of unemployment-to-jobs exits.

Consider the Cobb-Douglas specification of the matching function in levels in log-linear form, where lower-case letters are logarithms.

$$f_{it} = \alpha_0 + \gamma f_{it-1} + \alpha_1 u_{it-1} + \alpha_2 v_{it-1} + \eta_i + \mu_t + \varepsilon_{it} \tag{12}$$

f_{it} is the log number of outflows from unemployment to jobs in district i over period t, which is regressed on its lagged value f_{it-1}, on the stock of log registered unemployment and on log notified vacancies in district i at the beginning of period t. η_i is a time-invariant group-specific fixed effect and μ_t is a period fixed effect capturing seasonal effects and an aggregate time trend. Let N be the number of cross-sections and T the number of time-series observation in the panel. I assume for the error term ε_{it} to have the usual properties

$$E[\varepsilon_{it} | f_{it-1}, u_{it-1}, v_{it-1}] = 0,$$

$$V[\varepsilon_{it} | f_{it-1}, u_{it-1}, v_{it-1}] = \sigma_u^2 \text{ for all } i \text{ and } t,$$

$$Cov[\varepsilon_{it}, \varepsilon_{js} | f_{it-1}, u_{it-1}, v_{it-1}] = 0 \text{ for all } i \neq j \text{ or } t \neq s.$$

The model in section 3 has established the importance of endogeneity of participation of the employed in the search process and its relevance for the size of α_1 and α_2. Endogenous adjustments of search intensities of those searching on-the-job has been shown to increase the elasticity of job-matches with respect to unemployment on the one hand, and to dampen elasticity of job-matches with respect to vacancies on the other hand. The overall effect on returns to scale of the matching function depends on the strength of both effects. Table 4.2 presents regression results for all 76 local labour markets in the Czech Republic between January 1992

[13] Mankiw, Romer, and Weil (1992), and more recently, Islam (1995) use panel data to estimate rates of convergence in growth between countries.

and July 1994. I estimate a *bare bones* matching function, which does not parametrise exogenous effects on the matching technology such as the impact of active labour market policies, the role of local spillover effects in job-matching, or the heterogeneity of labour market districts due to structural composition[14].

Table 4.2 Regressions in Levels of the Czech Matching Function, # of observations: 2356, N=76, T=31 (1:1992 - 7:1994), Dependent Variable: $\ln f_{it}$

		$\ln f_{it-1}$	$\ln u_{it-1}$	$\ln v_{it-1}$	*RTS*	*SSE*	*Wald*	*B-G(1)*
1	Pooled OLS	--	0.829 (87.9)	0.153 (15.7)	0.982	284.5	1.784	541.6*
2	LSDV, time and district fixed effects	--	0.774 (24.3)	0.134 (7.31)	0.908	111.7	5.642*	193.6*
3	Pooled OLS, dynamic	0.418 (25.4)	0.500 (32.4)	0.072 (7.77)	0.990	223.4	0.745	0.261
4	LSDV, time and district fixed effects, dynamic	0.276 (15.2)	0.623 (19.6)	0.099 (5.63)	0.998	101.2	0.002	2.754

Keys: Absolute t-values are given in parentheses. Intercept is not reported. Asterisks indicate rejection of the Null hypothesis at 5% significance. Under the Null hypothesis of (long-run) constant returns to scale the Wald statistic is distributed $\chi^2(1)$. The Breusch-Godfrey statistic under the Null of no first-order serial correlation is also $\chi^2(1)$.

Regression 1 in Table 4.2 reports the benchmark results of model (12) from pooled OLS and confirms the theoretical prediction of significant positive elasticities of unemployment exits with respect to unemployment and vacancies. With a Wald test statistic of 1.784 the null hypothesis of constant returns to scale ($\alpha_1 + \alpha_2 = 1$) cannot be rejected at 5% significance. A Breusch-Godfrey test reveals clear evidence of first-order serial correlation in regression residuals. I include lagged unemployment exits to account for partial adjustment in job-matching in regression 3 of Table 4.2. Again the hypothesis of (long-run) constant returns to scale ($\alpha_1 + \alpha_2 + \gamma = 1$) cannot be rejected. A Breusch-Godfrey test statistic shows no further evidence of first-order serial correlation[15]. These estimates neglect the possibility of heterogeneity of districts and seasonality in unemployment exits. In regressions 2 and 4 of Table 4.2, a fixed effects model (LSDV) for time and districts accounts for these effects. Compared to the

[14] See Burda and Profit (1997), Burda and Lubyova (1995), Boeri and Burda (1996), and Boeri and Scarpetta (1995).

[15] Burda and Lubyova (1995) and Burda and Profit (1997) show that this partial adjustment process may be of higher order. However, they also show that about 60% of the adjustment occurs within the first month. Hence, I restrict the analysis to a first-order process.

dynamic OLS regression, the partial adjustment parameter drops sharply whereas short-term coefficients on unemployment and vacancies increase slightly. While the Wald statistic even indicates decreasing returns to scale for the static fixed effects model, constant returns to scale cannot be rejected in the dynamic model, regression 4. The inclusion of a lagged dependent variable effectively removes first-order serial correlation in residuals.

Doel and Kiviet (1994) demonstrate that estimates obtained from OLS or from a least-squares dummy variables approach (LSDV), as those in Table 4.2, are severely biased and inconsistent when partial adjustment dynamics are neglected. Nickell (1981) shows that even when lagged dependent variables are included, fixed effects models yield inconsistent and biased estimates, and derives an expression for the bias. This expression is shown to disappear as $T \rightarrow \infty$. Whereas studies of economic growth – the main field in economics where regional panel data are used – are mostly concerned with the coefficient of the lagged dependent variable, the *convergence parameter*, long-run coefficients of explanatory variables, unemployment and vacancies, are of interest in the context of job-matching. Nickell (1981) demonstrates how in dynamic fixed effects models estimated with OLS the inconsistency and the bias carries over to coefficients on exogenous variables. This inconsistency may have severe consequences for returns to scale estimates and the implied dynamics of equilibrium unemployment.

Judson and Owen (1996) present Monte Carlo evidence that even in the presence of relatively long time-series the bias in autoregressive fixed effects models estimated with OLS (or LSDV) may still be important. They find that even with T in the range of 30 observations the bias still accounts for 30% of the true values of γ, whereas the bias in the estimates of α_i is found to be relatively small. Even though Judson and Owen (1996) find the LSDV estimator to perform well with large T, they advise alternative techniques which produce consistent estimates for partial adjustment model using panel data sets.

Anderson and Hsiao (1982) have proposed an estimator which removes individual fixed effects by differencing (4.1),

$$f_{it} - f_{it-1} = \gamma(f_{it-1} - f_{it-2}) + (\mathbf{x}_{it-1} - \mathbf{x}_{it-2})'\alpha + (\varepsilon_{it} - \varepsilon_{it-1})$$

(13) or

$$\Delta f_{it} = \gamma \Delta f_{it-1} + \Delta \mathbf{x}_{it-1}'\alpha + \Delta\varepsilon_{it},$$

where $\mathbf{x}_{it-1} = (u_{it-1}, v_{it-1}, t)'$ and $\alpha_t = (\alpha_1, \alpha_2, \mu_t)'$. Since the disturbance $\Delta\varepsilon_{it}$ in (13) is correlated with Δf_{it-1}, Anderson and Hsiao (1982) recommend to instrument the latter with Δf_{it-2} and estimate with 2SLS. Arellano (1989) proposes f_{it-2} as an instrument, since it can be shown to render more efficient estimation results for some combinations of parameters. Arellano and Bond (1991) suggest a more efficient estimator which exploits a larger set of moment conditions. This estimator is "most semi-asymptotically efficient" among available instrumental variable estimators, which use lagged values of the dependent variable as instruments (Sevestre and Trognon, 1992 ; Harris and Mátyás, 1996) [16]. The formal expressions for the Anderson-Hsiao (AHIV) and Arellano-Bond (GMM(1)) estimator are given in Appendix. In the presence of heteroskedasticity, Arellano and Bond (1991) show that applying a two-step procedure yields more efficient results: first, regression residuals are obtained from a consistent one-step GMM estimator. The regression residuals are then exploited to construct the optimal weighing matrix for the GMM(2).

The Anderson-Hsiao estimator in regression 5 in Table 4.3 includes lagged difference in log exits from unemployment whereas regression 6 uses lagged log levels of unemployment exits as instruments. Regressions 7 to 8 apply variants of GMM estimators. To reduce the dimension of the instrument matrix for the GMM in the presence of a large time-series dimension, I restrict the number of instruments for the exogenous variables to lagged first differences as proposed by Sevestre and Trognon (1992), and the triangular expansion matrix to a maximum of two lags in unemployment exits.

All *difference* estimators of the Czech matching function in Table 4.3 yield very similar results, in particular significantly higher elasticities of unemployment exits with respect to unemployment stocks compared to Table 4.2. Most importantly, the Wald test soundly rejects long-run constant returns to scale in all cases. The Sargan test for overidentifying restrictions reported in the right hand column of Table 4.3 cannot reject the hypothesis of instrument validity. However, in contrast to Nickell's (1981) findings the coefficient on lagged unemployment-to-job outflows is smaller compared to regression 4 in Table 4.2 when estimated with AHIV or GMM, indicating additional problems with this specification.

[16] Ahn and Schmidt (1995) work out a full set of restrictions on second moments which in contrast to the Arellano and Bond estimator also leads to non-linear orthogonality conditions in the lagged dependent variable. See also Wansbeek and Bekker (1996) for a discussion of these estimators.

Table 4.3 Regressions in first Differences (IV and GMM), Dependent Variable $\Delta \ln f_{it}$

Instruments: $\ln f_{it-2}$ (resp. $\Delta \ln f_{it-1}$), $\Delta \ln u_{it-1}$, $\Delta \ln v_{it-1}$

		$\Delta \ln f_{it-1}$	$\Delta \ln u_{it-1}$	$\Delta \ln v_{it-1}$	*RTS*	*SSE*	*Wald*	*Sargan*
5	AHIV, time fixed effects, diff. instr.	0.097 (1.91)	1.904 (17.3)	0.048 (1.48)	2.049	163.4	52.3*	--
6	AHIV, first diffs, time fixed effects, lev. instr.	0.169 (1.68)	1.980 (13.7)	0.047 (1.40)	2.196	174.5	26.9*	--
7	GMM(1), time fixed effects, A-B instr. restr. to 2 lags[a)]	0.164 (2.81)	1.926 (14.8)	0.081 (2.16)	2.171	173.7	60.5*	69.5 (60)
8	GMM(2), time fixed effects, A-B instr. restr. to 2 lags[a)]	0.160 (15.8)	1.917 (67.3)	0.085 (9.02)	2.162	173.2	131.4*	69.7 (60)

Keys: See Table 4.2. The Sargan test for over-identifying restrictions is asymptotically $\chi^2(h)$ distributed with h degrees of freedom equal to the number of overidentifying instruments given in parentheses below the test statistic.
[a)] t-values calculated with White's heteroskedasticity robust standard errors. See Arellano and Bond (1991). A-B: uses the Arellano-Bond triangular expansion matrix.

Table 4.4 Regressions in first Differences (GMM), Dependent Variable: $\Delta \ln f_{it}$

Instruments: $\ln f_{it-2}$, $\Delta \ln u_{it-2}$, $\Delta \ln v_{it-1}$

		$\Delta \ln f_{it-1}$	$\Delta \ln u_{it-1}$	$\Delta \ln v_{it-1}$	*RTS*	*SSE*	*Wald*	*Sargan*
9	GMM(1), time fixed effects, A-B instr. restr. to 2 lags[a)]	0.169 (4.06)	0.946 (3.49)	0.071 (1.91)	1.186	182.8	0.48	81.4* (60)
10	GMM(2), time fixed effects, A-B instr. restr. to 2 lags[a)]	0.161 (18.3)	0.997 (13.3)	0.089 (8.15)	1.247	180.7	9.6*	68.0 (60)

Keys: See Table 4.2 and 4.3.

The ability of estimators based on the specification of the matching function in differences in reducing the Nickell-bias crucially hinges the availability of exogenous instruments for lagged unemployment-to-job flows and the assumption of an uncorrelated error term in the equation 4.1 (Sevestre and Trognon, 1992). However, the definition of flow variables implies that the change in unemployment over a certain time interval equals the number of inflows into unemployment and out of unemployment, $u_{it} = u_{it-1} + i_{it} - f_{it} - g_{it}$, where g_{it} is the flow from unemployment out of the labour force. From $\Delta f_{it} = \Delta u_{it-1} - \Delta u_{it} + \Delta i_{it} - \Delta g_{it}$ and since Δf_{it} contains $\Delta \varepsilon_{it}$, it is clear that $corr(\Delta \varepsilon_{it} \Delta \ln u_{it-1}) \geq 0$, hence the residual in (13) is correlated with the instrument, which produces an upward bias in the coefficient on unemployment

(see Burda, 1994). As an escape route, a twice lagged difference in unemployment is used as instrument in Table 4.4. As expected, regressions 9 and 10 show that the elasticity of unemployment outflows with respect to unemployment drops from 1.9 to about 1. At least for GMM(2), constant returns to matching are still rejected.

Table 4.5 Regressions in first Differences (GMM), Dependent Variable: $\Delta \ln f_{it}$

Instruments: $\left(\ln u_{it-3}\right), \ln v_{it-2}$, 1:1992 - 7:1994

		$\Delta \ln f_{it-1}$	$\Delta \ln u_{it-1}$	$\Delta \ln v_{it-1}$	*RTS*	*SSE*	*Wald*	*Sargan*
11	GMM(1), time fixed effects, A-B instr., complete[a)]	0.281 (2.75)	1.229 (5.47)	0.099 (1.45)	1.609	201.3	7.28*	136.7* (119)
12	GMM(2), time fixed effects, A-B instr., complete[a)]	0.252 (13.6)	1.283 (19.3)	0.103 (4.94)	1.638	194.4	110.9*	73.2 (119)
13	GMM(1), time fixed effects, A-B instr., complete[a)]	0.323 (2.79)	0.647 (1.79)	0.156 (3.04)	1.126	222.3	0.18	72.8 (58)
14	GMM(2), time fixed effects, A-B instr., complete[a)]	0.300 (11.8)	0.757 (9.26)	0.159 (12.5)	1.216	214.4	9.33*	46.8 (58)

Keys: See Table 4.2 and 4.3.

Allowing for serially correlated error terms ε_{it} in equation (12) also invalidates $\ln f_{it-2}$ as a feasible instrument. Hence, I only use triangular expansion matrices for the levels of twice lagged unemployment and lagged vacancy stocks in regressions 11 and 12 of Table 4.5. Again the coefficient on unemployment is lower compared to Table 4.4, but still greater than one. CRTS are rejected at 5% significance. Taking together the implications of residual correlation and the stock-flow identity, even further lagged unemployment is invalid as an instrument for lagged outflows to employment. Hence, regressions 13 and 14 show GMM estimates only taking the triangular expansion matrix on lagged vacancies. The results show a short-term elasticity with respect to unemployment of less than one, but at least for the two-step GMM, long-run returns to scale are still rejected in favor of IRTS in job-matching. In addition, the coefficient on the lagged dependent variable is increased to a value of 0.3 which is higher compared to the OLS estimates in regression 4 in Table 4.2, as predicted by Nickell (1981).

Hence, the finding of long-run IRTS in the Czech matching function is robust against different specifications with respect to the choice of instruments, when consistent estimators are applied. This is in sharp contrast to most earlier studies[17].

It has been argued that IRTS may be due to the endogenous effects in the search intensity of employed as a response to local labour market conditions. Moreover IRTS in job-matching – as a necessary condition for multiple labour market equilibria – are consistent with the bimodality of the regional distribution of unemployment rates in the Czech Republic in section 2.

5. Decomposition of Returns to Job-Matching

The model section 3 provides theoretical underpinning to the importance of job-to-job movements for the matching process: it predicts a larger coefficient on unemployment and a smaller coefficient on vacancies for a higher fraction of employed to total job seekers. The effect of on-the-job-search is difficult to infer directly since data on employed job search is not readily available in the Czech Republic, especially not on a regional level, which is the perspective taken in this study. It is, however, possible to find variables which possibly provide information on the intensity of job-to-job transitions and its impact on job-matching.

A first possible approach to analyse the impact of employed job search on the matching process is related to the analysis of intra-distributional dynamics of regional unemployment and vacancy rates in section 2. A simple cluster analysis which minimises the average distance between two clusters classifies districts into three groups for relative unemployment rates and two clusters for relative vacancy rates, as shown in Figure 5.1. A dummy variable for each of the five clusters is interacted with log unemployment and vacancies, and interaction terms are included as explanatory variables to estimate the reduced-form matching function.

Table 5.1 shows matching function estimates with separate coefficients for each interaction. The method is GMM(2) using lagged vacancies as instruments as in regression 14 in Table 4.5. The results show clear heterogeneity of matching coefficients depending on the relative position of a district within the regional distribution. The matching coefficients in the

[17] Münch et al. (1998) also find IRTS using similar estimation methods.

high unemployment cluster show the expected parameter constellations in the presence of strong employed job search: a coefficient on log unemployment larger than one and, in contrast to the standard matching theory, a negative coefficient on log vacancies. Speaking in terms of the model in section 3 (equation (6)) this means, that the effect of higher unemployment or lower vacancies on the search intensity of employed workers out-weights the effect of a lower employment rate on the matching probability in these regions. The findings however are mixed: the large coefficient on unemployment in districts with higher relative vacancy rates at the outset of the transition process works against this argument.

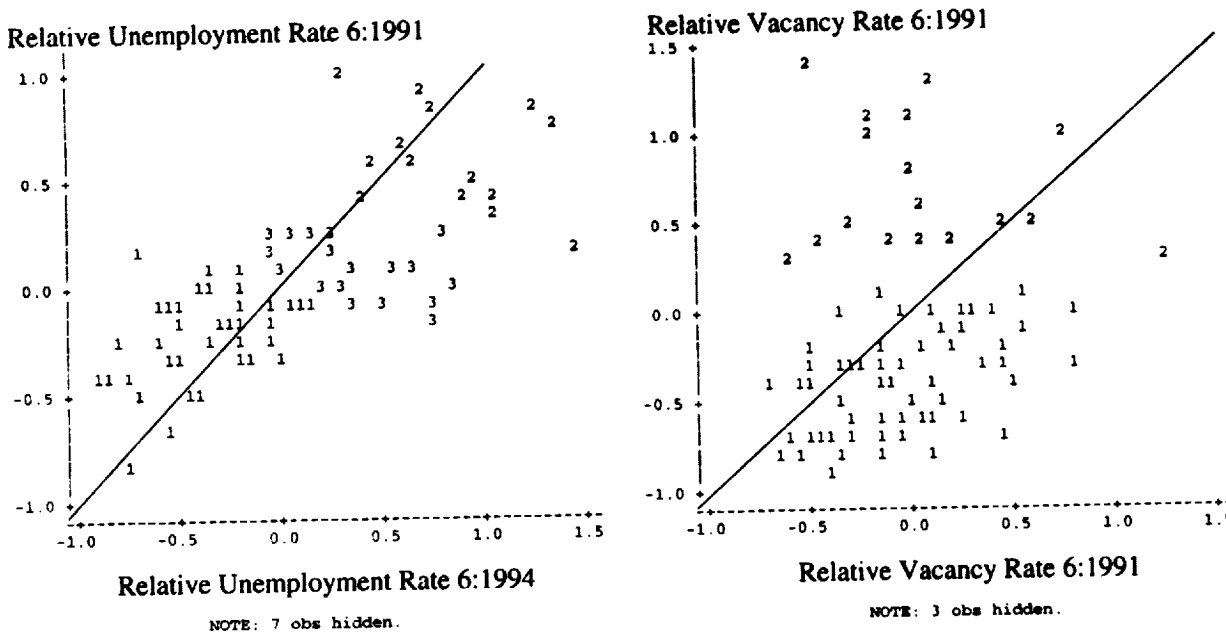


Figure 5.1 Clusters of Districts

A second approach to measure the effects of endogenous on-the-job search on job-matching is to interact the ratio of private enterprises and the ratio of employment in the service sector to total employment at yearend 1994 with log unemployment and vacancies, and to augment the matching function with these terms[18]. The results are shown in regression 16 and 17 in Table 5.1. The value in square brackets gives the short-run elasticity of unemployment-to-job exits with respect to unemployment and vacancy changes evaluated

[18] The ratios of private enterprises and sectoral employment to total employment are provided by the Czech Statistical Office. Service sector employment includes retail, tourism, hotel and restaurants, transport and communication, banking and insurance, and services provided by enterprises.

at mean ratio of private enterprises (21.4%) and service sector employment (27%) to total employment.

Table 5.1 Decomposition of Empirical Matching Functions, 1:1992 - 7:1994, Regressions in first Differences (GMM(2)), Dependent Variable: Log Unemployment-to-Jobs Exits, $\Delta \ln f_{it}$, **Instruments:** $\ln v_{it-2}$

Explanatory variable	*(15)*	*(16)*	*(17)*
$\Delta \ln f_{it-1}$	0.327 (7.3)	0.285 (10.0)	0.158 (4.5)
$\Delta \ln u_{it-1}$	--	0.486 (1.7)	-0.125 (0.3)
- cluster 1: low unempl. rates	0.742 (3.4)	--	--
- cluster 2: high unempl. rates	1.849 (3.9)	--	--
- cluster 3: intermediate unempl. rates	-0.449 (0.9)	--	--
- cluster 1: low vacancy rates	0.773 (3.0)	--	--
- cluster 2: high vacancy rates	1.368 (3.0)	--	--
- priv. enterprises/total employment (1994)	--	0.675 (0.5) [0.630]	--
- empl. in services/total employment (1994)			0.602 (0.6) [0.037]
$\Delta \ln v_{it-1}$	--	-0.563 (2.1)	1.390 (6.4)
- cluster 1: low unempl. rates	0.427 (4.6)	--	--
- cluster 2: high unempl. rates	-0.635 (6.3)	--	--
- cluster 3: intermediate unempl. rates	0.290 (2.8)	--	--
- cluster 1: low vacancy rates	-0.127 (1.1)	--	--
- cluster 2: high vacancy rates	0.208 (1.4)	--	--
- priv. enterprises/total employment (1994)		3.480 (2.7) [0.180]	--
- empl. in services/total employment (1994)			-4.443 (5.3) [-0.192]
SEE	222.2	215.0	204.5
Sargan	46.3 (50)	45.0 (57)	51.6 (57)

Keys: See Table 4.2 and 4.3. Equations (16) and (17) contain the vector of log vacancies multiplied with the share of private enterprises to total employment, and the share of employment in service industries in 1994, respectively. Square brackets contain the total coefficient on unemployment or vacancies evaluated at the mean of the interaction variable, the mean value of the ratio of private enterprises to total employment is 0.2136 across districts, the mean share of employment in service industries is 0.2697.

For the coefficient on unemployment, both interactions are insignificant. For the elasticity of unemployment outflows with respect to vacancies, different levels of private enterprises or service sector employment to total employment have a strong and opposed

effect: a larger relative number of private enterprises increases the coefficient on vacancies, whereas a larger share of employment in services reduces the coefficient. While the latter is consistent with a strong negative effect of endogenous on-the-job search on unemployment outflows, the former effect is probably due to a dominating role of labour demand, since job creation is stronger in private enterprises.

6. Conclusion

The study has presented evidence, that the emergence of strong regional disparities in regional unemployment in the Czech Republic since the outset of the transformation at the beginning of the 1990s can, at least partially, be explained by endogenous processes from local labour markets in this country. In particular, it is shown that competition between the emerging private sector and state-owned enterprises for skilled labour, which gives rise to large job-to-job movements may be an important characteristics of labour markets in transition. Together with low level of unemployment in the Czech Republic, such endogenous adjustments in search intensities of employed job seekers produce external effects on the matching technology implying increasing returns in the reduced-form matching function.

This observation is consistent with the analysis of intra-distributional dynamics of regional unemployment rates between 1991 and 1994, which shows the pattern of a twin-peaked distribution, with a low- and a high unemployment rate equilibrium, and some labour market districts churning between these equilibria. The intra-distributional dynamics for vacancy rates show a clear trend of convergence across Czech districts over the same period of time.

A properly specified and consistently estimated matching function which accounts for autocorrelation in unemployment-to-job exits, the presence of heteroskedasticity, and the validity of instruments renders elasticities of outflows to jobs with respect to unemployment and vacancies which imply increasing returns to matching. Earlier studies have failed to find this effect.

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Appendix: Dynamic Panel Estimators

This appendix contains a formal presentation of the instrumental variable estimator by Anderson and Hsiao (1982) and the generalized method of moment estimator by Arellano and Bond (1991). After stacking observations, I transform (13) to

$$\text{(A.3)} \qquad \Delta f = [\Delta X \vdots \iota_N \otimes D]\beta + \Delta\varepsilon$$

with

$$\Delta f = \begin{pmatrix} \Delta f_{12} \\ \vdots \\ \Delta f_{1T} \\ \vdots \\ \Delta f_{NT} \end{pmatrix}, \quad \Delta X = \begin{pmatrix} \Delta f_{11} & \Delta u_{11} & \Delta v_{11} \\ \vdots & \vdots & \vdots \\ \Delta f_{1T-1} & \Delta u_{1T-1} & \Delta v_{1T-1} \\ \vdots & \vdots & \vdots \\ \Delta f_{NT-1} & \Delta u_{NT-1} & \Delta v_{NT-1} \end{pmatrix}, \quad \Delta\varepsilon = \begin{pmatrix} \Delta\varepsilon_{12} \\ \vdots \\ \Delta\varepsilon_{1T} \\ \vdots \\ \Delta\varepsilon_{NT} \end{pmatrix},$$

and

$$D = \begin{pmatrix} 1 & 0 & & & 0 \\ -1 & 1 & & & \\ 0 & -1 & \ddots & & \\ & & \ddots & 1 & 0 \\ & & & -1 & 1 \\ 0 & & & 0 & -1 \end{pmatrix}, \quad \iota_N = \begin{pmatrix} 1 \\ \vdots \\ 1 \end{pmatrix}, \quad \beta = \begin{pmatrix} \gamma \\ \alpha_1 \\ \alpha_2 \\ \mu_1 \\ \vdots \\ \mu_{T-1} \end{pmatrix}.$$

The $N(T\text{-}1)\times T\text{-}1$ matrix $\iota_N \otimes D$ captures time fixed effects. The instrument matrix Z equals $[\Delta X \vdots \iota_N \otimes D]$ except for the first column which is replaced by Δf_{-2}, or f_{-2} respectively. The Anderson-Hsiao estimator is obtained from

$$\text{(A.4)} \qquad \hat{\beta}_{AH} = \left(Z'[\Delta X \vdots \iota_N \otimes D]\right)^{-1} Z'\Delta f$$

and the covariance matrix is estimated as

$$\text{(A.5)} \qquad \hat{V}\left(\hat{\beta}_{AH}\right) = \hat{\sigma}^2 \left\{[\Delta X \vdots \iota_N \otimes D]' Z(Z'Z)^{-1} Z'[\Delta X \vdots \iota_N \otimes D]\right\}^{-1},$$

where $\hat{\sigma}^2 = 1/(NT-K)(\Delta\varepsilon'\Delta\varepsilon)$ and K is the number of columns of $[\Delta X \vdots \iota_N \otimes D]$. Arellano and Bond (1991) suggested a more efficient estimator which exploits a larger set of moment condi-

tions. This estimator is "most semi-asymptotically efficient" among available IV estimators, which use lagged values of the dependent variable as instruments (Sevestre and Trognon, 1992 ; Harris and Mátyás, 1996). The estimator is given by

$$\text{(A.6)} \qquad \hat{\beta}_{AB} = \left(\left[\Delta X \vdots \iota_N \otimes D \right]' \tilde{Z} \psi \tilde{Z}' \left[\Delta X \vdots \iota_N \otimes D \right] \right)^{-1} \left[\Delta X \vdots \iota_N \otimes D \right]' \tilde{Z} \psi \tilde{Z}' \Delta f ,$$

and the covariance matrix of this estimator is obtained from

$$\text{(A.7)} \qquad \hat{V}\left(\hat{\beta}_{AB}\right) = \hat{\sigma}^2 \left(\left[\Delta X \vdots \iota_N \otimes D \right]' \tilde{Z} \psi \tilde{Z}' \left[\Delta X \vdots \iota_N \otimes D \right] \right)^{-1} .$$

The original proposal of Arellano and Bond (1991) is to construct the instrument matrix $\tilde{Z}_i$ as a triangular expansion matrix for lagged dependent and exogenous variables with the sth block equal to $\left(f_{i0}, \ldots, f_{is}, \Delta \mathrm{x}_{i1}, \ldots, \Delta \mathrm{x}_{is+1}\right)$ with $s = 0, \ldots, T-2$, the row vector $\Delta \mathrm{x}_{it-1} = \left(\Delta u_{it-1} \; \Delta v_{it-1} \; \Delta \mu_t\right)$ and $\tilde{Z} = \left(\tilde{Z}_1, \ldots, \tilde{Z}_N\right)$. For the generalized instrumental variable (one-step) estimator the weight matrix ψ takes the form

$$\text{(A.8)} \qquad \psi = \left(1/N \sum_{1=1}^{N} \tilde{Z}_i' \Sigma \tilde{Z}_i \right)^{-1} \qquad \text{with } \Sigma = \begin{pmatrix} 2 & -1 & & & 0 \\ -1 & \ddots & \ddots & & \\ & \ddots & \ddots & \ddots & \\ & & \ddots & \ddots & -1 \\ 0 & & & -1 & 2 \end{pmatrix} .$$

In the presence of heteroskedasticity a two-step general method of moments estimator is more efficient: first, regression residuals are obtained from a consistent one-step estimator, like (4.6). The weight matrix of GMM(2) is then defined as

$$\text{(A.9)} \qquad \psi = \left(1/N \sum_{1=1}^{N} \tilde{Z}_i' \Delta\hat{\varepsilon}_i \, \Delta\hat{\varepsilon}_i' \, \tilde{Z}_i \right)^{-1} \quad \text{where } \Delta\hat{\varepsilon}_i = \left(\Delta\hat{\varepsilon}_{i2}, \ldots, \Delta\hat{\varepsilon}_{iT}\right)$$