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In memory of my father

Abstract. When job search takes place across labour markets, the standard flow approach to labour market analysis fails to uncover the effectiveness at which workers are matched to available jobs. A spatially augmented matching function is backed by a spatial search model with endogenous search intensity. Recent studies deal with the issue of spatial externalities by assuming the process of job matching to be homogenous across space. This study shows that this supposition is not valid for the unified Germany. Particularly differences in labour mobility give reason for the existence of West-East regimes of the matching process. Spatial heterogeneity is additionally found on the level of German macroregions. Though matching efficiency is affected by labour market characteristics, its cyclical pattern is closely related to business cycle fluctuations. Variation of regional mismatch over the business cycle can only explain a relatively small fraction of matching inefficiency.

Keywords: Matching function, regional mismatch, spatial spillovers, spatial heterogeneity

JEL: C21, C23, E24, J23

1. Introduction

Since the early 1990s, the spatial pattern of the unemployment rate in Germany has been relatively stable. High unemployment areas in East Germany in particular coexist with relatively low unemployment regions in West Germany. Classical theory of migration predicts that regional disparities will be reduced by labour moving to regions offering the highest real wages. However, opposite effects will occur because of the difficulty for depressed regions to attract investments. Inflows of migrants into prosperous regions can trigger cumulative expansionary processes thereby reinforcing regional disparities (Armstrong and Taylor, 2003, pp. 162). Thus, considerable spatial movements of workers may come along with persistent regional unemployment disparities. Job search across labour markets is addressed in job search models of migration (Hughes and McCormick, 1994). The decision to stay or leave the home region is, however, completely ignored in the standard matching or flow approach to labour market analysis (see Petrongolo and Pissarides, 2001). Burda and Profit (1996) were the first to have substantiated a spatially augmented matching function by a model of nonsequential search over space. They provide evidence of the relevance of spatial interaction in job search for the Czech economy.

Externalities in job matching across travel-to-work areas in the United Kingdom are proved by Burgess and Profit (2001). Petrongolo and Wasmer (1999) found weak cross-regional spillovers for Britain and France. Recently, López-Tamayo et al. (2006) established evidence for the relevance of the spatial dimension in matching workers to vacant jobs for Spanish NUTII and NUTIII regions. Fahr and Sunde (2006a, 2006b) investigated spatial interaction in the matching process for West German planning regions in the period 1980-1997.

As space adds to search frictions in the labour market (Wasmer and Zenou, 2006), involving a spatial dimension in the econometric analysis of matching workers and employers is a necessary step. In previous studies, internal and external externalities in job matching are uniformly inferred from the aggregated matching function. Burda and Profit (1996) explain instationarities in the Czech matching function by nonuniform spatial dependence. Structural instability of parameters is, however, usually considered as indicative for spatial heterogeneity (cf. Anselin, 1990). The long-ranging process of integration of the two formerly independent German countries may give reason for the existence of West-East regimes in job matching. Spatial heterogeneity may, however, even be present on a regional scale. At present, there is a lack of knowledge on the combined effect of interregional dependency and spatial heterogeneity in the job matching process.

The purpose of the present paper is to close this gap for the unified Germany. We employ a panel of monthly data for 180 travel-to-work areas on flows from unemployment to employment, stocks of unemployed and vacancies for the period 1998 – 2004 to investigate the above issues. Additional labour market factors like the proportions of the older unemployed, high-skilled unemployed, long-term unemployed and tightness are involved by reason of their potential effect on matching efficiency. In order to allow for both spatial and temporal dependencies, the relationships are modelled in a spatial SUR system. By means of switching regressions of spatial regime models spatial heterogeneity in matching workers and firms is tested. All econometric models control for regional and time effects as well as for seasonal variation.

At the outset, we discuss various properties of the aggregate German matching function. In all specifications, the constant-returns-to-scale (CRS) hypothesis is rejected in favour of decreasing returns to scale. Regional job matches come along with highly significant internal and external spillovers. Negative externalities by non-resident unemployed workers and positive externalities from job openings in other regions are well in line with economic reasoning. Switching regressions give strong evidence for the existence of West-East regimes in the matching relationship. Spatial interaction turns out to be considerably more strongly marked across East German labour markets than between West German travel-to-work areas. This backs the hypothesis that migration is an essential part in the job search process of workers living in high unemployment regions.

The impacts of labour market factors are mostly as expected. Conditioned to these factors, changes in matching efficiency turn out to be closely related to business cycle fluctuations. Evidence of parameter instability at the level of macroregions is revealed by switching regressions of multiple-regime SUR models. Based on aggregate and macroregional matching functions, an indicator of regional mismatch is evaluated for West and East Germany as well as for Germany as a whole. As regional mismatch only explains a relatively small portion of cyclical inefficiency, the influence of fluctuations of economic activity on matching effectiveness is clearly evidenced (cf. Wall and Zoega, 2002; Kosfeld et al., 2006).

The paper proceeds as follows. Section 2 introduces the tool of the spatially augmented matching function. In section 3, spatial panel econometric methods employed in this study are outlined. Section 3 deals with the definition of regional labour markets and data issues. In section 5, we discuss the empirical findings on matching unemployed workers and firms. First, we interpret the internal and external spillovers in job matching derived from the aggregate German matching function (section 4.1). We additionally analyse the influence of labour market factors on matching efficiency and its relation to business cycle fluctuations. Next, we examine parameter instability in the process of job matching across West-East regimes. (section 4.2). Finally, we deal with spatial heterogeneity at the level of German macroregions and regional mismatch. (section 4.3). Sector 5 provides a summary and conclusions.

2. The spatially augmented matching function

Due to frictions like imperfect information, distance and mobility barriers, the process of job matching is costly and time consuming (Pissarides, 2000; Petrongolo and Pissarides, 2001). In analysing the effectiveness of job search in view of all kind of frictions, the concept of the matching function has proved itself a useful modelling tool. Although hires of firms result from stocks of unemployed, employed and out of the labour force, we focus our attention on matches of unemployed workers to available jobs. The regional matching function (1) relates the number of new hires of region r in period t, M_{rt}, to the stocks of unemployed persons, $U_{r,t-1}$, and job openings, $V_{r,t-1}$, at the end of the previous period t-1:

$$(1) \qquad M_{rt} = M(U_{r,t-1}, V_{r,t-1}), \quad \partial M_{rt} / \partial U_{r,t-1} > 0, \ \partial M_{rt} / \partial V_{r,t-1} > 0 \ .$$

The standard matching function (1) is assumed to be a concave and increasing function of the arguments U and V. Although the premise of constant returns to scale (CRS) is attractive for formal derivations, this property is often empirically rejected (Petrongolo and Pissarides, 2001). As for the dependent variable, we restrict new hires to the flows from unemployment to employment in order to ensure a close correspondence between the flow and stock variables entering the function. The probability of an unemployed job seeker living in region r finding a job during a period in his home region is given by $M(U_{r,t-1},V_{r,t-1})/U_{r,t-1}$, while a vacancy is filled with probability $M(U_{r,t-1},V_{r,t-1})/V_{r,t-1}$. With constant returns to scale, the probability $M(U_{r,t-1},V_{r,t-1})/U_{r,t-1}$ will decrease with rising labour market tightness, $\theta=V/U$, while the probability $M(U_{r,t-1},V_{r,t-1})/V_{r,t-1}$ varies positively with θ.

Search efficiency depends on individual characteristics of the unemployed and firms. As they are expected to vary among different groups of unemployed job seekers, it is justified that demographic, educational and other structural variables enter the matching functional on a regional level. Petrongolo and Pissarides (2001), for instance, point to the proportions of young and long-term unemployed as potential control variables. Other structural variables used in empirical studies are the proportions of the older unemployed, unemployed women, high and low educated unemployed and labour market tightness (Fahr and Sunde, 2006a; Andersson and Burges, 2000; Münich et al., 1999).

Usually, the matching function (1) is specified by a Cobb-Douglas technology:

$$(2) \qquad M_{rt} = A_{rt} U_{r,t-1}^{\beta_1} V_{r,t-1}^{\beta_2} \,.$$

β_1 and β_2 are the matching elasticities with respect to unemployment and vacancies, respectively. The scale parameter A can be interpreted as an efficiency parameter. It may be differentiated by space and by time. In case of increasing returns of scale, $\beta_1+\beta_2>1$, more than one labour market equilibrium could occur due to externalities (Howitt and McAffee, 1987). Although the Cobb-Douglas form of the matching function has no convincing micro-foundation, particularly in empirical research it has become the 'standard' formalisation of the matching model (Petrongolo and Pissarides, 2001).

Regional job matching does not only depend on local stocks of unemployed workers and job openings. Unemployed workers of neighbouring or other labour markets will compete with their local job searchers for vacant posts. As both commuting and migration is a possible outcome of the job search process of workers, are spatial externalities involved in the matching process. In order to allow for regional spillovers, Burda and Profit (1996) consider a search model of nonsequential search over space. The spatially augmented matching function of the form

$$(3) \qquad M_{rt} = M(U_{r,t-1}, V_{r,t-1}, WU_{r,t-1}, WV_{r-t-1}), \quad \partial M_{rt} / \partial U_{r,t-1} > 0, \; \partial M_{rt} / \partial V_{r,t-1} > 0$$

is obtained as an approximation from the spatial search model. WU and WV denote spatially lagged unemployment and vacancies variables. As search costs will rise with increasing distance from the origin travel-to-work area (Armstrong and Taylor, 2003, p. 157), the spatial lags WU and WV are conceived to be some distance-weighted sums of unemployed workers and job openings from all other region. In general, beneficial or unfavourable externalities may be triggered by WU and WV. With hires restricted to the flows of regions' unemployed workers to firms in any labour market, however, WU is expected to negatively affect regional matching, $\partial M_{rt} / \partial WU_{r,t-1} < 0$, due to competition between local and 'foreign' job applicants. Workers from other regions will partially crowd out domestic unemployed. This means that the negative impact of spatially lagged unemployment on regional matches renders congestion effects. On the other hand, spatially lagged vacancies will enrich the employment opportunities for domestic workers, thereby creating beneficial effects on regional matches ($\partial M_{rt} / \partial WV_{r,t-1} > 0$). In this way WV is capable to generate positive spatial spillovers of job matching.

For a Cobb-Douglas matching technology, the spatially augmented matching function (3) reads

$$(4) \quad M_{rt} = A_{rt} U_{r,t-1}^{\beta_1} V_{r,t-1}^{\beta_2} WU_{r,t-1}^{\beta_3} WV_{r,t-1}^{\beta_4} .$$

While β_1 and β_2 reflect the internal matching elasticities with respect to unemployment and job openings, β_3 and β_4 denote the respective external elasticities. Spatial spillovers exist if at least one of the elasticities β_3 and β_4 is different from zero. They manifest themselves in matches of unemployed workers to vacancies of different travel-to-work areas. In the spatially augmented matching function (4), returns to scale have to be concluded from the sum of all four matching elasticities. In order to test for spatial heterogeneity and assess regional mismatch, we allow for spatially varying matching elasticities. The simplest case of spatial heterogeneity is captured by a switching model with two spatial regimes (West-East regimes). A generalisation leads to a multi-regime model consisting of German macroregions.

3. Spatial panel models

The spatially augmented matching function (4) is estimated with monthly data for German travel-to-work areas. Given R regions and T periods (R > T), we employ a spatial seemingly unrelated regressions (spatial SUR) model that allows for temporal and spatial dependencies (Anselin, 1988, pp. 139). In order to avoid regressions with "artificial" data possibly resulting from seasonal adjustments, we capture seasonal variation by seasonal dummy variables S. For monthly data the seasonal effects are measured relative to the first month of a year:

$$(5) \quad S_j = \begin{cases} 1 \text{ for } j\text{th month} \\ 0 \text{ otherwise} \end{cases}, \; j=2,3,\ldots,12.$$

Changes in matching efficiency are accounted for by annual dummy variables T. When measured relative to the first year, 1998, they are defined by

$$(6) \quad T_i = \begin{cases} 1 \text{ for } i\text{th year} \\ 0 \text{ otherwise} \end{cases}, \; i=2 \; (1999), \; 3 \; (2000), \ldots, \; 7 \; (2004).$$

As with the seasonal dummies (5), one period has to be dropped with annual dummy variables to avoid multicollinearity. We additional control for region-specific shocks on matching efficiency by including dummy variables (MR) for each of the twelve German macroregions defined in Table A1 (see Appendix).

As the numbers of unemployed and vacancies, U and V, are registered at the end of the month, matches M in period t result from the stocks of U and V in period t-1. Using the one period lagged values of U and V as well as WU and WV as explanatory variables in the panel model also avoids a possible endogeneity bias. The spatial lag variables WU and WV are defined as weighted sums of unemployed workers and job openings, respectively, of the surrounding regions:

$$\text{(7)} \qquad WU_{rt} = \sum_{s=1}^{R} w_{rs} \cdot U_{st}$$

and

$$\text{(8)} \qquad WV_{rt} = \sum_{s=1}^{R} w_{rs} \cdot V_{st}$$

with $w_{rr} = 0$. The spatial weights w_{rs} quantify the spatial interaction between the travel-to-work areas. If spatial interaction is assumed to be restricted between contiguous regions, binary weights w_{rs} will be an adequate choice (Anselin, 1988, pp. 17; Haining, 2004, pp. 82). As migration is not confined to neighbouring regions (Puhani, 2001), the more general weights w_{rs} have to be chosen. According to the gravity model of migration (Armstrong and Taylor, 2003, p. 157), we define the spatial weights by the distance decay function

$$\text{(9)} \qquad w_{rs} = \begin{cases} 1/d_{rs}^{\alpha} \text{ for } r \neq s \\ 0 \text{ for } r = s \end{cases}.$$

with the decay parameter α=2. d_{rs} denotes the distance between the centres of the regions r and s. The weight function (9) reflects the decay of regional interactions in virtue of rising costs of migration (e.g. transport cost) with increasing distances. We normalise the spatial lags by forcing the sums of U and WU as well as those of V and WV to be equal.

Regional matching efficiency is expected to be affected by the local structure of the labour market. Some labour market variables, for example, the proportions of young, high-skilled and long-term unemployed are available on a highly disaggregated regional level. We include these labour market variables as control variables in the panel model in order to test their impacts on effectiveness of job matching. Let l_k be the k*th* auxiliary labour market variable and θ_k its unkown regression coefficient. With p metric control variables, the panel model takes the form

$$\log(M_{rt}) = \beta_1 \log(U_{r,t-1}) + \beta_2 \log(V_{r,t-1}) + \beta_3 \log(WU_{r,t-1}) + \beta_4 \log(WV_{r,t-1})$$

$$(10) \qquad + \sum_{k=1}^{p} \theta_k l_{krt} + + \sum_{\ell=1}^{12} \alpha_\ell MR_\ell + \sum_{i=2}^{7} \gamma_i T_i + \sum_{j=2}^{12} \delta_j S_j + \varepsilon_{rt}$$

r=1,2,..,R; t=1,2,…,T. The regression coefficients β_j, j=1,2,3,4, give the internal and external matching elasticities with respect to the unemployed workers and vacancies. The α-and γ-coefficients render the region-specific varieties and annual changes of matching efficiency, respectively; the δ-coefficients reflect the seasonal variation.

In the panel model (10), we allow the numbers of unemployed workers and job openings to be endogenously determined, but claim the temporally lagged variables to be weakly exogenous:

$$E[\log(U_{is})\cdot\varepsilon_{it}] \neq 0,\ E[\log(V_{is})\cdot\varepsilon_{it}] \neq 0,\ E[\log(WU_{is})\cdot\varepsilon_{it}] \neq 0,\ E[\log(WU_{is})\cdot\varepsilon_{it}] \neq 0 \text{ for } s\geq t$$

and zero otherwise. Provided that interregional dependencies arise from spatial spillovers as suggested by search theory (Burda and Profit, 1996),[1] the disturbances ε_{rt} will be contemporaneously uncorrelated:

(11) $\quad Cov(\varepsilon_{tr}, \varepsilon_{th}) = E(\varepsilon_{tr}\cdot\varepsilon_{th}) = 0, r\neq h.$

If the equations (10) are unrelated over time, the condition

(12) $\quad Cov(\varepsilon_{tr}, \varepsilon_{sr}) = E(\varepsilon_{tr}\cdot\varepsilon_{sh}) = 0, t\neq s$

will hold as well. Provided that spatial heterogeneity is negligible, pooled OLS could then be a feasible estimation strategy. As generally the equations (10) will not be unrelated over time, the loss of efficiency and invalidation of inference with pooled OLS is avoided by spatial SUR estimation (Anselin, 1988, p. 140).

In the case R>T, the covariances $Cov(\varepsilon_{tr}, \varepsilon_{sh})$ can be estimated from the data by imposing some restrictions. In replacing assumption (12) by the condition

(13) $\quad Cov(\varepsilon_{tr}, \varepsilon_{sr}) = E(\varepsilon_{tr}\cdot\varepsilon_{sr}) = \sigma_{ts},$

the panel model (10) defines a spatial SUR model of the matching function. The spatial SUR system consists of equations for different time periods which are related to each other by condition (13).

Let $\mathbf{m}_t$ be a Rx1 vector of the observations on the dependent variable log(M), $\mathbf{X}_{t-1}$ a Rx4 vector of the observations on the explanatory variables log(U), log(V), log(WU) and

[1] This supposition turns out to be questionable for administrative regions (e.g. NUTS regions), as they are not defined with respect to the range of the underlying spatial process. This reflects the modifiable areal unit problem (MAUP), that is a potential source of spatial error autocorrelation (Arbia, 1986; Unwin, 1996).

log(WV), $\mathbf{Z}_t$ a Rx(p+29) matrix of the control variables (incl. dummy variables) and ε_t a Rx1 vector of the disturbances. In compact form, the t*th* equation of panel model (10) reads

$$\text{(14)}\qquad \mathbf{m}_t = \mathbf{X}_{t-1}\boldsymbol{\beta} + \mathbf{Z}_t\boldsymbol{\gamma} + \boldsymbol{\varepsilon}_t,\ t=1,2,\ldots,T,$$

with the coefficient vectors $\beta' = (\beta_1\ \beta_2\ \beta_3\ \beta_4)$ and $\gamma' = (\theta_1 \ldots \theta_p\ \alpha_1 \ldots \alpha_{12}\ \gamma_2 \ldots \gamma_7\ \delta_2 \ldots \delta_{12})$.[2] Dependencies over time are modelled by defining the nxn covariance matrix Σ:

$$\text{(15)}\qquad \Sigma = E(\varepsilon_t \cdot \varepsilon'_s) = \sigma_{ts} \cdot I_n,$$

where $\mathbf{I}_R$ is the RxR identity matrix. With this the time-dependent autocorrelation structure of the spatial SUR system (14) is given by the RTxRT covariance matrix

$$\text{(16)}\qquad \Sigma^* = \Sigma \otimes \mathbf{I}_R.$$

that can be estimated directly form the data.

Although the panel model (14)-(16) accounts for temporal and spatial dependencies, it assumes the matching process to be homogenous over space. The steep East-West gradient in the unemployment rate may cast doubt upon this presupposition. Boeri and Scarpetta (1995) point to variations in sectoral composition as a potential cause of parameter instability. In the German case, differences in effectiveness of regional labour agencies and different interregional mobility of workers in both parts of the economy may gain additional relevance. Given the West and East regimes of the matching process, the extended matching function (4) is econometrically adequately depicted by a spatial switching model (Anselin, 1988, pp. 132),

$$\text{(17)}\qquad \begin{bmatrix} \mathbf{m}_t^W \\ \mathbf{m}_t^E \end{bmatrix} = \begin{bmatrix} \mathbf{X}_{t-1}^W & \mathbf{0} \\ \mathbf{0} & \mathbf{X}_{t-1}^E \end{bmatrix} \begin{bmatrix} \boldsymbol{\beta}_W \\ \boldsymbol{\beta}_E \end{bmatrix} + \begin{bmatrix} \mathbf{Z}_t^W & \mathbf{0} \\ \mathbf{0} & \mathbf{Z}_t^E \end{bmatrix} \begin{bmatrix} \boldsymbol{\gamma}_W \\ \boldsymbol{\gamma}_E \end{bmatrix} + \begin{bmatrix} \boldsymbol{\varepsilon}_t^W \\ \boldsymbol{\varepsilon}_t^E \end{bmatrix},\ t=1,2,\ldots,T,$$

with the covariance structure (15) and (16). While the data submatrices in (17) have the same numbers of columns as the corresponding matrices in (14), the numbers of rows are R_1=133 in the Western regime (W) and R_2=47 in the Eastern regime (E). The coefficient subvectors for both spatial regimes are of the same dimension as the corresponding coefficient vectors β and γ. The spatial regime model (17) is especially backed, when the coefficient vectors β_W and β_E are different, which can be ascertained by the Wald test.

[2] We additionally considered a model with partial adjustment of hires (cf. Burda and Profit, 1996; Hujer and Zeiss, 2005). While the regression coefficient of lagged outflows takes a value in the range of 0.4 to 0.6 in the pooled OLS model, it drops to values around zero in the spatial SUR model. Thus, various forms of inertia in the matching process are captured by allowing for the covariance structure of the disturbances.

Spatial heterogeneity of job matching may be present at a higher level of disaggregation. Burda and Profit (1996), for instance, exposed spatial instationarities in non-spatial matching functions of macroregions in the Czech Republic. Because of their claim to capture parameter instability by allowing for spatial dependence, the authors refrained from estimating spatially augmented regional matching functions. We test for spatial heterogeneity at the level of German macroregions by spatially disaggregating the switching model (17). For this, we estimate a SUR system in form of a multiple-regime switching model,

$$
\begin{aligned}
(18)\quad \begin{bmatrix} \mathbf{m}_t^{SH/HH} \\ \mathbf{m}_t^{NI/HB} \\ \vdots \\ \mathbf{m}_t^{TH} \end{bmatrix} &= \begin{bmatrix} \mathbf{X}_{t-1}^{SH/HH} & \mathbf{0} & \cdots & \mathbf{0} \\ \mathbf{0} & \mathbf{X}_{t-1}^{NI/HB} & \cdots & \mathbf{0} \\ \vdots & \vdots & \ddots & \vdots \\ \mathbf{0} & \mathbf{0} & \cdots & \mathbf{X}_{t-1}^{TH} \end{bmatrix} \begin{bmatrix} \boldsymbol{\beta}_{SH/HH} \\ \boldsymbol{\beta}_{NI/HB} \\ \vdots \\ \boldsymbol{\beta}_{TH} \end{bmatrix} \\
&+ \begin{bmatrix} \mathbf{Z}_t^{SH/HH} & \mathbf{0} & \cdots & \mathbf{0} \\ \mathbf{0} & \mathbf{Z}_t^{NI/HB} & \cdots & \mathbf{0} \\ \vdots & \vdots & \ddots & \vdots \\ \mathbf{0} & \mathbf{0} & \cdots & \mathbf{Z}_t^{TH} \end{bmatrix} \begin{bmatrix} \boldsymbol{\gamma}_{SH/HH} \\ \boldsymbol{\gamma}_{NI/HB} \\ \vdots \\ \boldsymbol{\gamma}_{TH} \end{bmatrix} + \begin{bmatrix} \boldsymbol{\varepsilon}_t^{SH/HH} \\ \boldsymbol{\varepsilon}_t^{NI/HB} \\ \vdots \\ \boldsymbol{\varepsilon}_t^{TH} \end{bmatrix}, \quad t=1,2,\ldots,T,
\end{aligned}
$$

of which disturbance variable is subjected to the covariance structure (15) and (16). The macroregions mainly consist of German states (*Länder*) (see Appendix Table A1). City states are merged with a bordering or surrounding state. Because of their small sizes, the neighbouring states of Rhineland Palatinate and Saarland are combined into one macroregion.

4. Regional labour markets and data

Monthly data on outflows from unemployment to various states like employment, training, out of the labour force and others are provided by the German Federal Employment Agency from December 1997 to December 2004 for 439 German districts (*Landkreise* and *kreisfreie Städte*). Because of some assignment problems in the earlier months, we restrict our econometric analysis of regional job matching to the period from April 1998 to December 2004. The outflows from unemployment to employment reflect the numbers of unemployed residing in a specific district that are matched with an open vacancy in the local or another district.

Panel data on the stocks of unemployed and vacancies are reported by the German Federal Employment Agency for overall 870 local employment agencies. The numbers of unemployed cover registered unemployed who are searching for a job. Data on job openings have to be interpreted carefully. As the vacancies notified cover only a fraction of the actual stock, detailed knowledge of the reported shares (*Meldequoten*) would be desirable for adjusting the raw data. However, only annual shares are surveyed by the German Federal Employment Agency for West and East Germany in all. While the shares in West Germany range from 30 to 37% during the period 1998-2004, they vary between 24 and 42% in the eastern part of the country (IAB, 2005). We have applied these reported share series for an East-West adjustment of vacancies.

Both employment agencies (*Dienststellenbezirke*) and districts (*Kreise*) do not meet the requirements of functional regions. On average, 53% of the employees in the social security system in local authority areas are commuters who travel to their workplaces across administrative boundaries. Since the employment agencies are in most instances parts of the districts, the percentage of commuters across these spatial units will be even larger. Thus, the modifiable areal unit problem is expected to affect the spatial study of matching functions by generating artificial spatial patterns (cf. Arbia, 1986; Unwin, 1996).

Spatial autocorrelation due to inadequate delineation of areal units is largely avoided by working with travel-to-work areas. In this study, we therefore refer to labour market regions that are defined by commuter flows. Using data on job commuters across German districts, Eckey (2001) defined 180 regional labour markets of which 133 are located in the western and 47 in the eastern part of Germany. With these functional regions the average share of commuters decreases from 53 to 21%. On average a regional labour market consists of 2.4 districts and 4.8 agencies.

Several demographical, educational and labour market variables that are expected to affect efficiency of matching in a systematic way (cf. Petrongolo and Pissarides, 2001) are included in the regressions as control variables. With one exception all quantitative control variables are defined as shares of the unemployed. We particularly control for job applicants younger than 25 years and older then 50 years, unemployed with no secondary school education, unemployed with a university degree (incl. university of applied sciences, UAS), long-term unemployed and labour market tightness. Long-term unemployment is defined by unemployment spells of one year or longer. The tightness variable, that relates the number of vacancies to the number of unemployed, is added as an indicator of the strength of competition of firms for job applicants. In Table 1, descriptive statistics are depicted for the labour market variables used in this study.

Table 1: Descriptive statistics of labour market variables for 180 German labour markets

Variable	Mean	S.D.	Min	Max
Flow from unemployment to employment	1413	1869	94	24451
Number of unemployed	22840	34995	1680	419312
Adjusted number of vacancies	6901	9573	143	106041
Share of young unemployed	0.121	0.020	0.062	0.237
Share of elderly unemployed	0.290	0.053	0.171	0.509
Share of lowly qualified unemployed	0.361	0.117	0.081	0.614
Share of highly qualified unemployed	0.040	0.022	0.008	0.152
Share of long-term unemployed	0.340	0.071	0.080	0.579
Labour market tightness	0.390	0.340	0.005	4.334

Notes
Monthly data for the period from April 1998 to December 2004.
Source: German Federal Employment Agency

In order to examine the role of business cycle fluctuations on inefficiencies in job matching, we will regress the annual dummy variables of the panel model of the matching function on the output gap. As the difference between potential and actual GDP it measures cyclical fluctuations in the economy. To allow for a nonlinear GDP trend, we employ the Hodrick-Prescott filtered series as a proxy for potential GDP (Hodrick and Prescott, 1981). Annual GDP data is taken from the Statistical Yearbook of the German Federal Statistical Office.

5. Econometric analysis of the aggregate German matching function

We start our econometric analysis of job matching in Germany by focusing on the aggregate matching function. By estimating the panel model (14), we assume the regression coefficients to be stable over space. Table 2 shows new hires to rise with increasing numbers of unemployed and vacancies of the same region in all specifications. According to the SUR results, a rise in the local stocks of the unemployed by 1% comes along with an increase in outflows of unemployment to employment by 0.9%, while a 1% rise in the local stocks of vacancies entails an increase in matches by about 0.06%. The higher value of the unemployment elasticity compared to the vacancy elasticity of matching with outflow from unemployment as the variable to be explained is well-known from previous studies (cf. Petrongolo and Pissarides, 2001). With the used outflow variable, the number of unemployed is the stock "at risk" that matters most for job matches (Kangarsharju et al., 2003). The finding of a very high unemployment elasticity of matching (β_1) combined with an elasticity for vacant posts (β_2) well below 0.10 is confirmed for the sub-period 2003-2004 with different panel methods for Germany by Hujer and Zeiss (2005). However, the authors inferred extremely large values for the unemployment elasticity (β_1>1) from a non-spatial matching model. By contrast, we find decreasing returns to scale with the aggregated German matching function in all cases.

Table 2: Aggregate German matching function

	Pooled OLS				Spatial SUR			
	Ia	IIa	IIIa	IVa	Ib	IIb	IIIb	IVb
$\log(U_{t-1})$	0.773	0.863	0.775	0.866	0.907	0.909	0.903	0.904
	(211.530)	(170.589)	(208.159)	(168.667)	(157.161)	(162.361)	(155.009)	(160.813)
$\log(V_{t-1})$	0.128	0.089	0.128	0.091	0.039	0.055	0.040	0.057
	(40.257)	(19.336)	(39.492)	(19.841)	(20.033)	(18.062)	(19.485)	(19.440)
$\log(WU_{t-1})$	-	-	-0.134	-0.140	-	-	-0.093	-0.121
			(-18.715)	(-21.891)			(-22.490)	(-29.007)
$\log(WV_{t-1})$	-	-	-0.051	0.043	-	-	0.042	0.054
			(-6.951)	(6.329)			(7.180)	(8.751)
Young unemployed	-	-0.154	-	-0.190	-	-0.404	-	-0.412
		(-1.320)		(-1.664)		(-5.432)		(-5.549)
Older unemployed	-	-0.194	-	-0.268	-	-0.915	-	-0.958
		(-3.597)		(-4.805)		(-17.722)		(-17.811)
Low qualification	-	-0.593	-	-0.649	-	-0.485	-	-0.522
		(-16.121)		(-17.563)		(-9.373)		(-10.146)
High qualification	-	-0.765	-	-0.955	-	-0.132	-	-0.214
		(-8.538)		(-10.844)		(-0.860)		(1.409)
long-termed unemployed	-	-1.958	-	-1.841	-	-0.680	-	-0.673
		(-59.782)		(-56.795)		(-20.815)		(-20.645)
Tightness	-	0.002	-	-0.016	-	-0.024	-	-0.026
		(0.261)		(-1.948)		(-4.306)		(-4.822)
R^2	0.945	0.970	0.949	0.962	0.990	0.995	0.990	0.995
P-DW	1.265	1.166	1.261	1.108	1.889	1.862	1.860	1.862
RS	0.901^{**}	0.952^{**}	0.718^{**}	0.860^{**}	0.946^{**}	0.964^{**}	0.892^{**}	0.894^{**}

Notes
t-values are given in parenthesis below the regression coefficients.
R^2: Coefficient of determination; P-DW: Panel Durbin-Watson statistic
RS: Returns to scale (**: H_0: CRS rejected at the 1% level in favour of decreasing returns to scale)

While the elasticity of a region's own stock of unemployment is underestimated with pooled OLS, the inverse holds for the local vacancy elasticity. The bias is partly due to ignored structural and cyclical labour market variables. Different to this, the SUR estimates turn out to be relatively robust. The high unemployment elasticity β_1 of about 0.9 means that the negative externality for job searchers caused by competitors living in the same region is rather small (β_1-1≈-0.10). Thus, negative externalities arising from congestion are relatively weakly marked (1-β_1≈0.10). Positive externalities of local firms on job applicants are very limited as well (β_2≈0.06). From this, no evidence for a thick-market effect can be concluded. By contrast, high positive ($\beta_1 \approx 0.9$) and negative ($\beta_2 - 1 \approx 0.94$) externalities come along with job creation of the firms.

Although the internal elasticities for the unemployed and vacancies are almost unchanged by spatially augmenting the matching function, their external counterparts are highly significant. Unemployment in other regions has a negative effect on local job matching. If the number of distance-weighted unemployed in 'foreign' regions increases by 1%, the number of local matches will decrease by about 0.1% on average. Non-resident unemployed compete with a region's own applicants for open vacancies. They will partially crowd-out local job seekers and thereby decrease intraregional matches. On the other hand, local unemployed workers contact job centres or apply directly for job openings in 'foreign' regions. Thus, an increasing number of vacancies offered by non-resident firms will raise the chances for job matches with a region's own applicants. On average, a 1% increase of distance-weighted 'foreign' vacancies comes along with a rise of local matches of about 0.05%. This positive spatial externality of job creation is concealed, however, when ignoring structural factors and temporal dependencies in the job matching process (model IIIa). While the values of the panel Durbin-Watson statistic signify strong positive residual autocorrelation for pooled OLS, they tend to approach the range of uncorrelatedness with spatial SUR.[3]

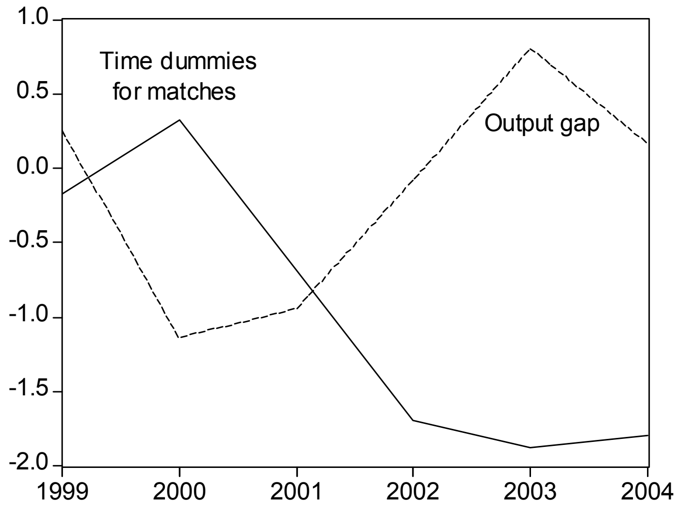
Although no other studies on a spatially augmented matching function for Germany as a whole are available, we can compare our results with the findings for other countries. Our outcomes are qualitatively well in line with those brought about by panel analyses of Burgess and Profit (2001) for Britain. With both contiguity- and distance-based weights Burgess and Profit established internal effects that are somewhat lower compared to our SUR estimates, while their external effects are larger. Based on pooled OLS with time and macro-regional dummies, however, Burda and Profit (1996) estimated spillover elasticities with the "wrong" sign for the Czech economy. By replacing the dummies for macroregions by district fixed effects, both external elasticities become positive. For different distance ranges, significant spatial effects with varying signs emerged in the panel model with macroregional dummies. Petrongolo and Wasmer (1999) found weak neighbourhood effects of unemployment on matching for Britain and for France. Lopez-Tamyo et al. (2006) established negative and positive spatial externalities from unemployment and vacancies, respectively, with inverse

[3] Bhargava et al. (1982) provide critical values for the panel DW test for some combinations of R and T. The critical values refer, however, to small time dimensions ($T \leq 10$).

distance weights for Spanish NUTS III regions. The opposite holds for NUTS II regions. Their findings on external effects between neighbouring areas are difficult to interpret.

In the above studies structural characteristics and cyclical factors are not explicitly taken into account. Table 1 reveals that the signs of the matching variables are not influenced by labour market controls in the spatial SUR models, although the external effects gain in strength. With two exceptions, both estimation methods coincide in assessing the influences of the structural and cyclical variables on job matching. The negative regression coefficients for the shares of young and older unemployed as well as that of the fraction of long-term unemployed is also found by Hujer and Zeiss (2005). While the share of low-skilled workers is negatively related to matching efficiency, no significant effect from high-skilled unemployed on the chance for a job match is concluded by spatial SUR estimation. More insight on the role of this factor may be obtained from switching regressions of West-East regimes. The negative influence of labour market tightness on the efficiency of the matching process may result from frictions due to increasing coordination problems. With higher tightness it becomes more and more difficult for firms filling open posts with suitable job applicants. This outbalances the improved chances of the unemployed in finding a job.

Figure 1: Matching efficiency and business cycle fluctuations



Annual dummy variables in the panel model capture changes of matching efficiency over time after partialling out the effects of the labour market variables. In Figure 1 the series of time dummies of the saturated spatial SUR model IVb is plotted along with the output gap. From the negative relationship between both series, improvements of matching efficiency in upturns and deteriorations in downturns are suggested. Labour market tightness does at most partially account for changes in matching efficiency in the course of the business cycle. The partial correlation coefficient between matching efficiency and the output gap amounts to –0.7. This outcome is well in line with the findings on the role of cyclical factors for shifts of the German Beveridge curve (see Kosfeld et al., 2006).

6. Econometric analysis of West-East regimes of the matching functions

Since the unification in 1990, West and East Germany have been growing together. However, the East-West gap in labour productivity and income per capita is still large. In 2005 East Germany's labour productivity achieves 78.7% of that in West Germany, while the ratio for income per capita amounts to just 69.4% (BMWT, 2006). After a notable catching-up until the mid 1990s, the convergence process came to a halt in the late 1990s. The speed of convergence has been different within both parts of the economy (cf. Kosfeld et al., 2006). Losses in production are particularly linked to high unemployment in East Germany. Large divergence of West-East German unemployment rates places special emphasis on the functioning of labour markets in both parts of the economy.

Table 3 shows that the internal elasticities for unemployment and job openings for West and East German lie close together. They roughly coincide the respective elasticities of the aggregate German matching function. While the unemployment elasticity for West Germany is well in line with the different panel estimates of Burda (1994), most of his estimates of the vacancy elasticity are considerably larger in the early 1990s. For East Germany four of five panel estimates of the unemployment elasticity are distinctly larger than one, whereas the estimated elasticities for job openings are similar to our results. In contrast to Burda (1994), Fahr and Sunde (2006a) estimate West German matching functions with spatial spillovers over a longer period with different hiring flows. With hires as the dependent variable, the weight of vacancies increases at the cost of unemployment. However, in none of the cases, the relevance of both explanatory variables in explaining hires instead of unemployment outflows is reversed. With hires of unemployed in the home region as the dependent variable, the influence of vacancies vanishes (Fahr and Sunde, 2006b).

As for spatial spillovers in job matching, a distinct feature in the Western and Eastern part of the economy becomes evident. Although the estimates vary across different panel models, the effects of spatially lagged numbers of unemployed and vacancies are always much more strongly in East Germany. This may be explained by higher labour mobility in case of extremely poor employment prospects (Mertens and Haas, 2005). In all cases, the congestion effect arising from job applicants of other regions is negative. According to spatial SUR estimates, an increase of unemployment in the surrounding regions by 1% leads to a

Table 3: West-East regimes of the matching functions

	Pooled OLS				Spatial SUR			
	I West	I East	II West	II East	III West	III East	IV West	IV East
$\log(U_{t-1})$	0.723	0.866	0.868	0.844	0.899	0.886	0.883	0.867
	(159.134)	(139.512)	(119.900)	(116.584)	(132.360)	(69.952)	(140.882)	(96.472)
$\log(V_{t-1})$	0.170	0.075	0.092	0.092	0.045	0.030	0.061	0.081
	(42.023)	(15.390)	(13.278)	(13.680)	(12.739)	(12.718)	(11.848)	(18.689)
$\log(WU_{t-1})$	-0.050	-0.292	-0.039	-0.276	-0.060	-0.223	-0.135	-0.228
	(-6.224)	(-20.770)	(-6.219)	(-28.045)	(-13.750)	(-30.969)	(-31.147)	(-32.542)
$\log(WV_{t-1})$	-0.144	0.207	-0.047	0.191	0.023	0.051	0.022	0.104
	(-17.440)	(15.892)	(-6.784)	(18.579)	(2.687)	(6.537)	(3.464)	(13.624)
Young unemployed	-	-	-1.383	-0.477	-	-	-0.599	-0.022
			(-13.519)	(-2.540)			(-7.392)	(-0.188)
Older unemployed	-	-	-0.112	-0.453	-	-	-0.648	-1.404
			(-2.461)	(-3.654)			(-11.024)	(-12.662)
Low qualification	-	-	-0.713	-0.742	-	-	-0.514	-0.577
			(-19.479)	(-10.070)			(-9.203)	(-6.006)
High qualification	-	-	-1.381	1.148	-	-	-0.529	0.483
			(-15.892)	(5.120)			(-3.372)	(1.175)
Long-term unemployed	-	-	-2.731	-1.142	-	-	-1.161	0.390
			(-66.019)	(18.725)			(-25.010)	(6.827)
Tightness	-	-	-0.045	-0.160	-	-	-0.017	-0.177
			(-4.613)	(-6.492)			(-2.638)	(-11.419)
R^2	0.954		0.966		0.991		0.994	
P-DW	1.369		1.193		1.860		1.868	
RS	0.699^{--}	0.866^{--}	0.874	0.851	0.907^{--}	0.744	0.781	0.824
W_U W_V	345.844^{**}	222.486^{**}	5.446^{*}	0.000	0.958	11.406^{**}	2.168	9.187^{**}
W_{WU} W_{WV}	234.510^{**}	557.475^{**}	451.633^{**}	390.303^{**}	389.938^{**}	6.446^{*}	127.628^{**}	72.603^{**}

Notes
t-values are given in parenthesis below the regression coefficients.
R^2: Coefficient of determination; P-DW: Panel Durbin-Watson statistic
RS: Returns to scale (**: H_0: CRS rejected at the 1% level in favour of decreasing returns to scale)
W_U, W_V, W_{WU}, W_{WV}: Wald tests on equality of regression coefficients of $\log(U_{t-1})$, $\log(V_{t-1})$, $\log(WU_{t-1})$, $\log(WV_{t-1})$, respectively, in both regimes

crowding-out effect of about 0.2% in East Germany compared to good 0.1% in West Germany. The differences with the positive externality of ‘foreign’ vacancies are even larger. While for West Germany a rise in a region’s matches of only 0.02% is inferred as a response on a 1% increase of distance-weighted job openings in other travel-to-work areas, matches are expected to rise in East Germany by 0.1%. The latter response is concluded from the saturated spatial SUR model IV East, while the estimate of β_4 is downward biased when labour market variables are omitted.

Since Burda (1994) only estimated non-spatial West-East German matching functions, comparisons can only drawn from Fahr and Sunde (2006a, 2006b) for West Germany. In Fahr and Sunde (2006a) the authors established a strong negative congestion effect of unemployed workers from neighbouring regions, whereas a positive influence on matching occurs by unemployed from non-neighbouring regions. As only hires in the home region are considered, spatial externalities by ‘foreign’ vacancies are not examined. For hirings from unemployment, Fahr and Sunde (2006b) estimate a positive external matching elasticity with respect to unemployed job seekers from neighbouring regions which may be due to the special definition of the dependent variable. The negative effect of unemployed from non-neighbouring areas is found not to be significant.

As with the aggregate German matching function, the coefficients of the proportions of young, older and low-skilled unemployment have a negative sign in both switching regressions. This applies as well to labour market tightness. Differences between West and East German labour markets occur with the shares of high-skilled and long-term unemployed. Job matching for skilled unemployed may become increasingly difficult due to stronger competition. For the long period from 1980 to 1997 such a negative association is not found (see Fahr and Sunde, 2006a). Although marginal productivity of human capital is found to be low in most East German travel-to-work areas (Eckey et al., 2005), the matching chances of highly educated and skilled workers are above average. On the other hand, a higher share of long-term unemployed is expected to reduce the chances of matches. For East Germany the estimation results are in variance with both panel methods. The somewhat “odd” SUR estimate may result from unfavourable behaviour of matches in response to declining long-term unemployment during the upturn in the macroregion Brandenburg/Berlin. As for all other Eastern macroregions this inverse reaction does not occur at all, the positive sign is for the most part resolved by regional disaggregation.

7. Econometric analysis of regional matching functions

By estimating West-East regimes of the matching function, spatial heterogeneity in job matching is allowed for at a highly aggregated level. More insight on spatial nonstationarity in the matching process is obtained by analysing regional matching functions. With the aid of multi-regime switching regressions we are as well able to assess the importance of regional mismatch.

Table 4 shows that for the most part regional disaggregated matching functions behave as expected. In all cases, a rise in local stocks of unemployed and job openings goes with an increase of a region’s own job matches. Moreover, the competition hypothesis with regard to ‘foreign’ unemployed is well corroborated. Spatial SUR brings out significant negativeestimates of the external matching elasticity with respect to unemployed workers for all twelve macroregions. The hypothesis of positive spatial externalities by the provision of

Table 4: Macroregional matching functions

	SH/HH	NI/HB	NW	HE	RP/SL	BW	BY	MV	BB/BE	SN	LSA	TH
					Pooled OLS							
Macroreg.	-0.241	0.037	0.830	0.968	-0.661	-0.637	-0.542	0.530	-0.766	-1.527	-0.082	0.530
Dummy	(-0.648)	(0.179)	(4.204)	(2.944)	(-3.207)	(-4.972)	(-4.479)	(1.172)	(-2.117)	(-5.447)	(-0.236)	(1.172)
$\log(U_{t-1})$	0.736	0.895	0.900	0.788	0.912	0.818	0.915	0.805	0.774	0.967	0.719	0.879
	(9.532)	(46.892)	(24.783)	(21.081)	(31.205)	(34.595)	(64.646)	(29.637)	(44.037)	(50.548)	(24.108)	(31.256)
$\log(V_{t-1})$	0.328	0.055	0.049	0.125	0.052	0.114	0.087	0.156	0.056	0.027	0.252	0.054
	(4.659)	(3.281)	(1.385)	(3.234)	(2.024)	(5.373)	(5.850)	(8.458)	(4.637)	(1.809)	(8.609)	(2.112)
$\log(WU_{t-1})$	-0.072	-0.175	-0.070	-0.102	-0.001	0.077	-0.062	-0.323	-0.143	-0.154	-0.264	-0.250
	(-1.746)	(-9.224)	(-3.762)	(-3.024)	(-0.046)	(4.538)	(-4.343)	(-8.588)	(-4.335)	(-6.745)	(-8.248)	(-8.599)
$\log(WV_{t-1})$	0.076	0.097	-0.169	-0.092	-0.076	-0.154	0.054	0.176	0.174	0.147	0.166	0.064
	(1.722)	(4.317)	(-7.850)	(-3.839)	(-3.040)	(-8.043)	(3.896)	(7.432)	(6.998)	(7.608)	(4.936)	(1.440)
RS	1.068	0.872**	0.710**	0.719**	0.887**	0.855**	0.994	0.814**	0.861**	0.987	0.873**	0.747**
	R^2	0.970	P-DW	1.308	W_U	115.1**	W_V	89.3**	W_{WU}	245.2**	W_{WV}	369.2**
					Spatial SUR							
Macroreg.	-1.277	-0.885	-0.705	-1.268	-1.277	-0.899	-1.033	1.530	-0.622	-0.803	2.074	0.902
Dummy	(-3.414)	(-3.900)	(-3.175)	(-4.422)	(-4.914)	(-4.686)	(-6.670)	(4.007)	(-1.950)	(-3.000)	(6.633)	(2.389)
$\log(U_{t-1})$	0.884	0.911	0.885	0.942	0.913	0.925	0.868	0.888	0.839	1.044	0.704	0.881
	(23.754)	(56.155)	(35.754)	(34.617)	(43.374)	(53.426)	(76.003)	(29.606)	(36.635)	(52.145)	(27.019)	(33.448)
$\log(V_{t-1})$	0.144	0.052	0.058	0.029	0.048	0.039	0.080	0.047	0.055	0.032	0.218	0.074
	(4.196)	(4.893)	(2.587)	(1.202)	(2.953)	(2.939)	(8.355)	(4.693)	(8.676)	(4.183)	(13.923)	(5.284)
$\log(WU_{t-1})$	-0.042	-0.088	-0.128	-0.071	-0.061	-0.046	-0.025	-0.293	-0.204	-0.206	-0.252	-0.230
	(-2.057)	(-8.039)	(-12.328)	(-4.769)	(-4.783)	(-4.874)	(-2.870)	(-16.426)	(-12.177)	(-14.744)	(-15.497)	(-13.212)
$\log(WV_{t-1})$	0.008	0.021	0.084	0.024	0.043	-0.044	0.016	0.085	0.111	0.034	-0.034	0.029
	(0.226)	(1.123)	(3.699)	(0.977)	(1.986)	(-2.616)	(1.170)	(4.518)	(7.428)	(3.016)	(-1.556)	(1.185)
RS	0.994	0.896**	0.899**	0.924*	0.943*	0.872**	0.939*	0.727**	0.802**	0.904**	0.636**	0.754**
	R^2	0.997	P-DW	1.884	W_U	129.9**	W_V	134.5**	W_{WU}	492.9**	W_{WV}	75.8**

Notes

Both regressions were run with seasonal dummies and macroregional-specific time dummies. t-values are reported in brackets below the regression coefficients.

The German macroregions are defined in Table A1 (Appendix).

R^2: coefficient of determination; P-DW: Panel Durbin-Watson statistic; RS: Returns of scale, * (**): significant deviations from the CRS hypothesis at the 5% (1%) level; W_U, W_V, W_{WU}, W_{WV}: Wald tests on the equality of regression coefficients of $\log(U_{t-1})$, $\log(V_{t-1})$, $\log(WU_{t-1})$, $\log(WV_{t-1})$, respectively, in multi-regimes models additional places of work in 'foreign' regions is not generally confirmed. For the macroregions Schleswig Holstein/Hamburg, Lower Saxony, Hesse, Bavaria and Thuringia, the elasticity for external vacancies has the expected positive sign without being significant. The workplaces in these macroregions attract workers from other areas, but the job search of their own unemployed persons mainly takes locally place. The sign of the external vacancy elasticity is reversed for Baden-Württemberg and – without being significant – Saxony-Anhalt. The larger external elasticities for Eastern macroregions point to higher cross-regional labour mobility of the East German workers.

Apart from Schleswig-Holstein/Hamburg, the CRS hypothesis is rejected for all macroregions in favour of decreasing returns to scale in job matching. Thus, the main features of the aggregate and West/East German matching functions are as well founded on the level of regional disaggregation adopted here. However, Wald tests clearly reject the hypothesis of equal matching elasticities across macroregions with respect to local and non-local unemployment and vacancies. The test results do not change within West and East German macroregions. This unambiguously points to spatial nonstationarities in the process of job matching that cannot be entirely captured by the West-East regimes.

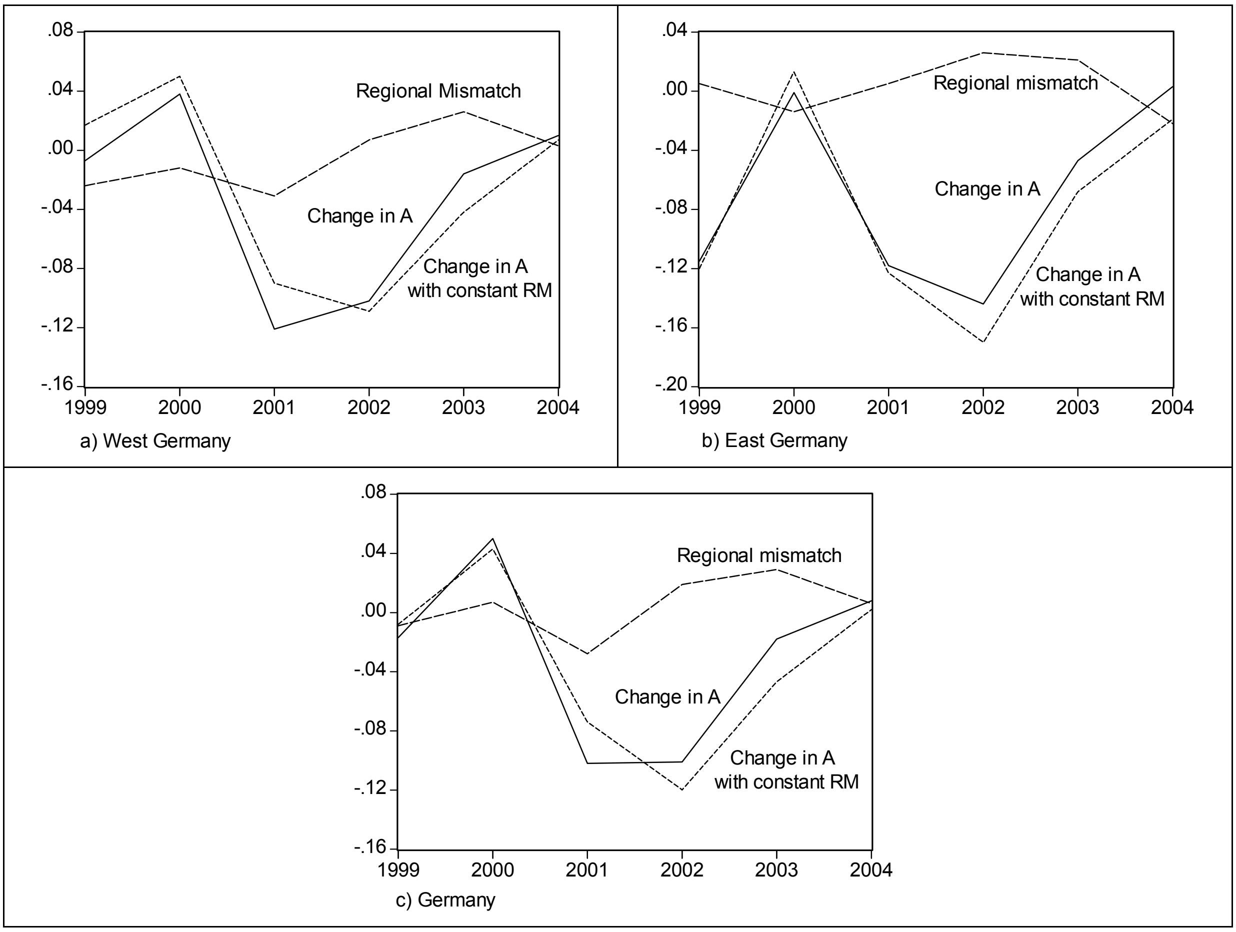
Despite the strong relationship between matching efficiency and business cycle fluctuations, changes in matching efficiency may still be ascribed to regional mismatch. Varying spatial lag coefficients point to differences in labour mobility over space. Imperfect mobility of workers can give rise for regional mismatch being low in peaks and high in troughs. Thus, changes in regional mismatch in the course of the business cycle may account for the cyclical pattern of matching efficiency.

Indicators of regional mismatch are usually based on unemployment rates, unemployment and vacancy rates and employment growth rates (see Jackman et al., 1991; Entorf, 1998). They are, however, not designed to establish the fraction of matching inefficiency that may be attributed to regional mismatch. Such an indicator can be defined by comparing the time dummies of the aggregate and regional matching functions. Let ΔA_t be the difference of time dummy in period t compare to period t-1 in the aggregate matching function. According to the modelling approach, this difference reflects the actual change of matching efficiency. Hypothetical changes holding constant regional mismatch (RM), $\Delta(A_t|RM)$, can be determined from the regional matching functions. Under constant returns to scale, $A_t|RM$ is given by the weighted regional time dummies with the weights equal to the regions' shares of the labour force. Thus, the differences $\Delta A_t - \Delta(A_t|RM)$ will reflect the regional mismatch at period t. A meaningful regional mismatch indicator RMI can then defined by

$$(19)\qquad RMI = \frac{\sum_{t=1}^{T} \left| \Delta A_t - \Delta(A_t|RM) \right|}{\sum_{t=1}^{T} \left| \Delta A_t \right|}.$$

RMI measures the relative extent of changes in matching efficiency that is accounted by regional mismatch in the period 1 to T.

Figure 2: Extent of regional mismatch



Notes
A Matching efficiency, RM Regional Mismatch

The mismatch indicator RMI_{Ger} of 0.331 states that 33.1% of the changes in matching efficiency can be attributed to regional mismatch in Germany after controlling for structural characteristics of the unemployed and labour market tightness. Thus, about two third of the conditioned cyclical behaviour of matching efficiency depicted in Figure 2c cannot be attributed to regional mismatch. This outcome confirms the findings of Wall and Zoega (2002) on the causes for shifts of the British Beveridge curve. That regional mismatch is more serious in West Germany (RMI_{West} = 0.350) than in East Germany (RMI_{East} = 0.217) is well in line with the strength of spatial effects. Both unemployed workers and vacancies of other regions are, in East Germany, much more strongly involved in regional job matching than in West Germany. The higher mobility of East German's workers finds its counterpart in a lower degree of regional mismatch. Annual differences can be traced with the aid of the panels a) and b) in Figure 2. In the period 1999 to 2001, East Germany's curves of the actual and hypothetical changes in matching efficiency lie close together in the first half of the sample period. In the second half of the period of investigation, the divergence of the curves is very similar in both parts of the economy.

6. Conclusions

The standard flow approach of labour market analysis fails to uncover how effective workers are matched to available jobs, when migration plays a significant role in job matching. In a nonsequential search model, Burda and Profit (1996) show how workers and vacancies of other regions are involved in regional job matching. The relevance of spatial interaction in the process of job matching has been empirically corroborated for the Czech Republic, Great Britain, France, Spain and West Germany. However, spatial heterogeneity arising from parameter instability over space, as yet has not been examined.

The study addresses the issues of regional spillovers and spatial heterogeneity in the matching of workers and employers in the unified Germany. The significance of spatial externalities in job matching is clearly confirmed. While 'foreign' job openings will raise the effectiveness of matching, local job applicants are partially crowded out by unemployed living in other travel-to-work areas. Switching regressions show, however, that the parameters of the augmented matching function are not stable across space. West and East regimes of the matching process differ mainly with respect to the strength of spatial interaction. Absolutely larger response coefficients of the spatial lag variables reflect the higher regional mobility of East German workers.

Allowing for spatial heterogeneity on a higher disaggregated regional level leads to a multiple-regimes panel model of German macroregions. Switching regressions give support for macroregional matching functions with spatially varying matching elasticities. Conditioned to regional labour markets structures and tightness, inefficiencies in job matching prove to be closely related to business cycle fluctuations. Although regional mismatch varies as well over the business cycle, it can only explain a relatively small fraction of cyclical behaviour of matching inefficiency. Some labour market reforms may be suitable to advance effectiveness of job matching and thus contribute to reducing unemployment. But the evaluation of such measures is a further subject of research.

Appendix

Table A1: German macroregions

West German States		East German States	
Macroregions	States (*Länder*)	Macroregions	States (*Länder*)
SH/HH	Schleswig-Holstein, Hamburg	MV	Mecklenburg-Western Pomerania
NI/HB	Lower Saxony, Bremen	BB/BE	Brandenburg, Berlin
NW	North Rhine-Westfalia	SA	Saxony
HE	Hesse	LSA	Saxony-Anhalt
RP/SL	Rhineland-Palatinate, Saarland	TH	Thuringia
BW	Baden-Württemberg		
BY	Bavaria		

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