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Unemployment and Worker-Firm Matching: Theory and Evidence From East and West Europe

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UNEMPLOYMENT AND WORKER-FIRM MATCHING: THEORY AND EVIDENCE FROM EAST AND WEST EUROPE+

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UNEMPLOYMENT AND WORKER-FIRM MATCHING: THEORY AND EVIDENCE FROM EAST AND WEST EUROPE

1. Introduction

During the two decades after the fall of the Berlin Wall, unemployment has been a major problem in the post-communist economies of the former Soviet bloc, including the new members of the European Union (EU). High unemployment has also been a serious issue in many western European countries. The question therefore arises as to whether similar or different factors bring about the high unemployment outcome in the two sets of economies and to what extent their labor markets converge to similar patterns.

In policy discussions in Central-East Europe, three hypotheses have emerged as leading explanations for this phenomenon, namely that high unemployment is the result of (1) problems related to the economic structures (mismatch) in these countries, (2) macroeconomic policies or major external shocks, or (3) ongoing (unfinished) transition from plan to market in the presence of globalization.[1] The discussion complements that in Western Europe, where the focus in explaining unemployment has been on the relative importance of (a) structural (mismatch) shocks, (b) aggregate demand shocks, and (c) hysteresis (e.g., Jackman, Pissarides and Savouri, 1990, and Jackman and Layard, 2004). The nature of appropriate policies for tackling unemployment obviously depends on identifying the nature of the problem.

In this paper, we use new data to address this issue, while advancing the theoretical and applied econometric literature on matching functions. Our strategy is to compare the evolution of unemployment dynamics and analyze the efficiency of matching of workers with job vacancies in four different transition economies and one geographically close West European market economy. In particular, we use

[1] A fundamental systemic feature of the Soviet-type economies was the nonexistence of open unemployment. An equally distinguishing feature of the transition during the early-to-mid 1990s was the emergence of double digit unemployment rates in all the rapidly transforming economies except for the Czech Republic.

newly collected 1991-2005 *district-level monthly panel data* on the unemployed *U*, vacancies *V*, inflow *S* into unemployment (reflecting labor turnover in firms), and outflow *O* from unemployment in five former communist economies (the Czech Republic, Hungary, Poland, Slovakia, and eastern part of Germany – hereafter "East Germany") and in the western part of Germany (a benchmark western economy – hereafter "West Germany") to examine the three hypotheses in the context of labor market turnover and the efficiency of matching of the unemployed and vacancies.[2]

The comparison of the transition economies with an otherwise similar and spatially close market economy is useful because it enables us to identify the main differences and similarities in the evolution of the key variables and thus draw conclusions as to whether different or similar factors are at work. From an analytical standpoint we are also comparing an interesting set of transition economies. East Germany, Czech Republic and Slovakia were until the end of communism close adherents to the centrally planned, state-ownership system, with East Germany subsequently abruptly merging with a mature market economy (our benchmark) and its functioning institutions and the Czech and Slovak republics pursuing an independent path developing market institutions steadily from scratch. In contrast, communist Hungary and (to a lesser extent) Poland had already introduced some market oriented reforms and Poland had a non-negligible private sector (especially in agriculture) throughout the communist era.

[2] These countries constitute an appropriate set of economies in which to examine these issues. In the Czech Republic, the unemployment rate remained at mere 3-4 percent throughout the (transformation) recession of first half of the 1990s and only rose to 6-9 percent during the second recession of 1997-99. In the 2000s, the unemployment rate remained in the very high 14-20 percent range in the rapidly growing economy of Slovakia (as well as Poland), and stabilized in the high 7-10 percent range in the moderately growing Czech Republic and Hungary. In Western and Eastern Germany, which we examine as comparison economies, the unemployment rate has since the early 1990s fluctuated around 10 percent and 15 percent, respectively. An important part of the answer to the above questions is that from the time unemployment started appearing in CEE in the early 1990s, the Czech Republic has had extraordinary low inflow rate and higher outflow rate of individuals from the unemployment state to employment than did the other CEE economies (see e.g., Boeri, 1994, Boeri and Scarpetta, 1995 and Ham, Svejnar and Terrell (HST), 1998, 1999). For instance, in 1993 inflow rates were 0.7 in the Czech Republic but 1.5 in Slovakia, 1.9 in East Germany, and 1.13 in Poland (see table 1). Similarly, in 1993 the outflow rate in the Czech Republic was 21.0, 8.1 in Slovakia, 4.9 in Poland, and 4.3 in Hungary (from HTS). Moreover, possible causes of the less rapid rise of unemployment in the Czech Republic in the early 1990s, such as lower inflow rates into unemployment due to higher government subsidies to Czech firms or to greater declines in Czech labor force participation, are not borne out by the data. These basic findings suggest that one needs a better understanding of the determinants of outflow from unemployment and matching of the unemployed and vacancies in the Czech Republic and the other CEE

We are hence able to assess if the outcomes differ systematically with the diverse initial conditions and subsequent paths.

Turning to the aforementioned hypotheses, which follow formally from equation (4) below, hypothesis H1 implies that high unemployment is caused by inefficient matching. This hypothesis contains two sub-hypotheses because matching efficiency may be low in terms of either low or negative trend in disembodied efficiency (H1a) or low returns to scale in matching (H1b). This inefficiency may be brought about for example by inadequate labor market institutions leading to decreasing search effort, skills depreciation, rising reservation wage of the unemployed, or geographical or skill mismatch (see also Jurajda and Terrell, 2006). If this hypothesis is correct, one would observe both U and V being simultaneously high, but not necessarily in the same districts or skill groups. In fact, spatial mismatch measured by a standard mismatch index indicates that in general the extent of mismatch has not increased over time.[3] If one finds support for H1, the policy should focus on labor market institutions and measures to foster residential labor mobility, create appropriate skills, and stimulate job search effort.

Hypothesis H2 states that high unemployment is caused by low demand for labor (e.g., due to restrictive macroeconomic policies, overvalued exchange rate, or globalization shocks). The manifestation of this would be low V relative to S causing high U, and the policy implication would be that macroeconomic policies are key.

Hypothesis H3, namely that high unemployment implies that restructuring is at work, is consistent with the observation that inflow S (presumably from old jobs) is high. The manifestation of this situation would be high U brought about by high S and the policy implication would be that restructuring needs to be completed. In the case of West Germany, the phenomenon would not represent the transition from

countries. While HST (1998,1999) examine the outflow issue using individual unemployment duration data, in the present paper we analyze the process of matching using long monthly panels of district-level data.

[3] In West Germany, for instance, the spatial mismatch index declined during the 1990s by about one-third and remained stable afterwards. After the early 1990s, mismatch remained stable in all the transition countries that we study except

plan to market but rather restructuring brought about by globalization and other forces.

In Table 1 we provide time series statistics on unemployment, inflow, outflow and vacancies in the six economies. In the left panel of the table, we give the inflow rate, outflow rate and *U/V* ratio (a measure of labor market tightness), while in the right panel we express unemployment and vacancies as a share of the labor force in each country. As may be seen from the table, the six economies differ markedly in terms of their unemployment, flows and vacancy levels and rates.[4] West Germany is in the intermediate range, displaying between 1991 and 2005 an unemployment rate that increases from 5 to 10 percent, inflow rate that rises from 0.9 to 1.6 percent, outflow rate that declined only slightly, and a vacancy rate (as a share of the labor force) that fluctuated between 0.7 and 1.4 percent. The changes in these variables occur mostly in two waves, reflecting the business cycles and also a notable shift toward a service economy with higher natural labor turnover that translates into steadily rising inflows into unemployment (a rise by two-thirds over the 1991-2005 period). East Germany, in contrast, registers an open unemployment rate rising from near zero to 18.6 percent, inflow rate almost doubling from an already high level of 1.7 to 3.0 percent, outflow rate rising to rather high 13-15 percent by the mid 1990s and fluctuating around this level ever since, and vacancies as a share of labor force rising from 0.4 percent in 1991 to about 1 percent in the late 1990s and remaining at that level in the 2000s. For most of the 1991-2005 period, the East German part of the German economy hence displays a much higher unemployment rate driven primarily by extraordinarily high inflow rate (labor turnover in firms resulting in registered unemployment). Note however, that East Germany also has higher outflow rates relative to the number of unemployed and a similar vacancy rate as the Western part of Germany. In particular, the East German economy operates with a higher unemployment rate in the presence of very sizable active labor market

Hungary, which experienced a steady growth during last decade, and Slovakia, which experienced a noisy but discernible decline during the same period.

[4] Numbers presented are country aggregates based on our working district level database. Because some districts for some countries are excluded from our analysis due to changes of district borders, data in Table 1 could slightly differ from

policies that lead to relatively high outflows out of unemployment, but unfortunately do not prevent high (subsequent) inflows into unemployment. Slovakia and Poland represent two transition economies that, like East Germany, operate with very high unemployment rates but, unlike East Germany, have not experienced an administratively set high wage level and cross-border subsidies. For most of the 1990s and 2000s, these two economies have experienced an unemployment rate in the 14 to 20 percent range, accompanied by moderately high inflow and low outflow below 10 percent. In most years, they have also had vacancy rates significantly below 1 percent. The Czech Republic is an intermediate case, with unemployment rising from the low rate of 3-4 percent in the early to mid 1990s to 8-10 percent range since then. Its inflow rate has risen from extraordinary low levels of about 0.6-0.8 percent in the early-to-mid 1990s to a still relatively low level of 1.1-1.2 percent since then. Its vacancies as a share of the labor force have declined from very high levels of 1.4-1.9 percent in the early-to-mid 1990s to 0.8-1.1 percent since then. Finally, Hungary has achieved the lowest and rather stable level of unemployment rate. After reaching an unemployment rate of about 11 percent in the mid-to-late 1990s, Hungary succeeded to lower the rate to around 8 percent in the mid 2000s, reduced its inflow rate to 1.4 percent, raised the outflow rate to 14-16 percent and kept the vacancy rate at 1.0-1.1 percent. Hungary's relative success is hence brought about by keeping the outflow rate relatively high and inflow rate relatively low.

The Hungarian and Polish unemployment outcomes must be interpreted soberly in light of a relatively high outflow of people out of the labor force. In particular, between 1992 and 2004 the ratio of employed to the population in the 15-59 year old cohort declined by 9.9 percentage points. The corresponding decline in the Czech and Slovak Republics was 4.3 and 7.1 percentage points, respectively. The situation in Poland is especially serious because the 9.9 percentage point decline occurred from an already relatively low base.[5]

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official aggregate statistics published. German series exclude East and West Berlin.

[5] In 1992 the ratio of employed to the population of 15-59 year olds was only 64.8 percent in Poland, as compared to 74.7

As mentioned above, during the last two decades the Western part of Germany, like other market economies, has been undergoing significant adjustments in response to globalization. At the firm level, one observes a relatively sizable increase in the labor turnover rate, which is reflected in a two-thirds increase in the inflow rate into unemployment between 1991 and 2005. The rise in the inflow rate in market economies such as West Germany appears to be driven in part by a decline of some traditional industries and rise of the service sector, with the former having a lower and the latter a higher rate of turnover. These different rates of turnover in turn seem to be caused by lower competition and greater firm-specific human capital in the declining relative to the rising sectors of the economy. Another part of the explanation for the rising inflow rate is growing international competition and greater frequency of shocks that result in permanently higher rates of job destruction and job creation.[6]

Models of transition from plan to market assume that the turnover (inflow) rate would rise dramatically as the old state sector sheds workers who go through unemployment into new jobs that are being created in the emerging private sector (e.g., Aghion and Blanchard, 1994, Blanchard, 1997, and Castanheira and Roland, 2000). The models predict that the inflow rate would be temporarily very high and gradually decline and approach the level observed in otherwise similar market economies such as West Germany. Interestingly, data from the five transition economies, presented in Table 1, indicate that the inflow rate trajectories have been very different from the theoretical scenario. First, except for East Germany (to be discussed presently), none of the transition countries that we study had a considerably higher inflow rate than West Germany during the entire 1991-2005 period. In fact, some of the countries had a lower inflow rate than West Germany for extended periods of time -- the Czech Republic in the early-to-mid 1990s being a notable example. Second, by the mid 2000s the inflow rate in all economies except East Germany converges to a similar range (1.1-1.6). Third, by the mid-2000s the West German

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percent in the Czech Republic.
[6] For more on the issue of shocks and labor market institutions, see Ljungqvist, L. and T. Sargent, (1998), den Haan et al.

inflow rate actually exceeded the rate observed in the Czech Republic, Hungary and Slovakia, and was similar to that in Poland.[7]

In terms of our hypotheses, the raw data in Table 1 suggest that the West German economy displays elements of all three hypotheses. Both the unemployment and vacancy rates are relatively high (H1), the vacancy rate has declined (H2) and the unemployment rate has risen with increasing inflows (H3). The East German economy has conformed to H3, having generated high unemployment through relatively but not extremely high outflow rates (much going into training programs) and extremely high inflow rates (much being probably re-inflow from training programs). East Germany is also consistent with H2 in that its vacancy rate has been low. Slovakia and Poland reflect primarily H2 (low vacancies) throughout the 1990s and 2000s, although they have also experienced rising inflow rates consistently with H3. The Czech economy had virtually no unemployment problem in the early-to-mid 1990s, as inflow rate was extremely low and vacancies remained high. While this could be interpreted as a sign of slow restructuring, various studies (e.g., Jurajda and Terrell, 2008) suggest that labor mobility through job-to-job moves from old to new jobs was substantial. Indeed, many of these job-to-job moves were shifts from state-sector employment to private-sector employment. In Poland, for example, such job-to-job shifts were in 1992-93 more than twice as large as flows from public sector employment to unemployment: almost 9 percent of state sector employment moved directly to the private sector compared with a modest 4 percent becoming unemployed (Boeri, 2000). In the Czech Republic during the 1991-1996 period, 19 percent of all those who left an old state job went directly to a new sector job, while 3 percent became unemployed or left the labor force (Jurajda and Terrell, 2008). However, with the onset of a recession in 1997 and gradual elimination of fiscal subsidies to firms through the banking sector, the Czech Republic has increasingly conformed to H2 and H3. Finally, Hungary has an element of all three hypotheses. Its

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(2005), Hornstein et al. (2007).

[7] The Slovak inflow rate profile has a concave part and it could be argued that it resembles the model prediction. However,

unemployment and vacancies are relatively high (H1), the vacancy rate is low relative to inflow (H2) and inflow is relatively sizable (H3).

In view of this background, one could provide analytical information on unemployment and its dynamics by focusing on either inflows (job destruction in firms) or outflows (matching of the unemployed and vacancies). In our analysis, we use the newly collected district-level data on individuals and vacancies to identify the extent to which the countries of Central Europe exhibit different levels of efficiency in matching. In a companion paper (Münich and Svejnar, 2007), we focus on inflows and shifts in unemployment and vacancies.

The paper is structured as follows: We start in Section 2 by presenting our conceptual framework of a matching function model and a brief survey of the literature. In Section 3 we discuss our estimating framework and explain how we overcome some of the principal problems of the existing studies. In Section 4 we describe our data and the implementation of our econometric model. In Section 5 we present basic statistics and our econometric estimates. We conclude in Section 6.

2. Conceptual Framework and Existing Literature

It is useful to start by noting that most of the relevant literature focuses on the matching function – a relationship that describes the complex pairing of the unemployed U and vacancies V in creating outflow O from unemployment into jobs:

$$O = M(U,V). \tag{1}$$

Theories of search and matching generally do not imply a particular functional form of the matching function and most studies use the Cobb-Douglas function, which may be written in a deterministic form of discrete observations as[8]

the rise occurs only in the late 1990s rather than at the start of the transition in the early 1990s.

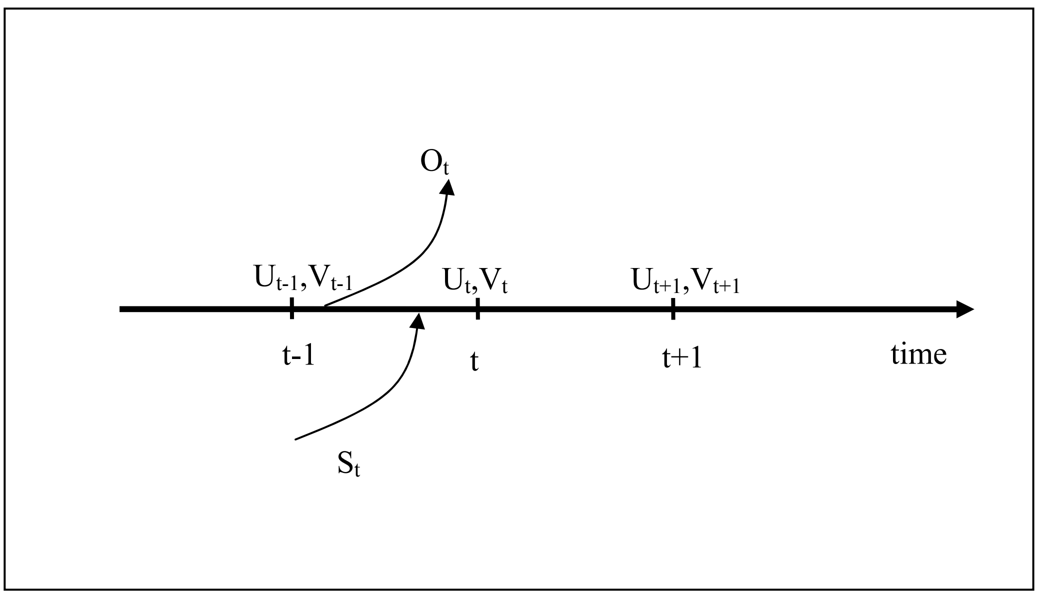
[8] There are of course exceptions. Pissarides (1990) for instance shows that in his theoretical model the Cobb-Douglas

$$\ln O_t = \beta \ln U_{t-1} + \gamma \ln V_{t-1} + \ln A \,, \qquad (2)$$

where, U_{t-1}, and V_{t-1} are the number of unemployed and vacancies at the end of period *t-1,* respectively, O_t denotes the outflow to jobs during period t (the number of successful matches between the currently unemployed and current vacancies) and constant ln*A* captures part of the efficiency of matching.

In general, matching may be viewed as a search process in which the unemployed and employers with vacancies strive to find an acceptable match. The process is conditioned by exogenous factors such as skill and spatial mismatch, as well as costly access to information and foregone income in cases of protracted duration of search. The timing of this stock-flow matching process is depicted on Chart 1.

Chart 1: Stock-flow process of matching



Some authors (e.g., Blanchard and Diamond, 1989, Pissarides, 1990, and Storer, 1994) assume the matching function *M* to display constant returns to scale (CRS), while others have identified reasons such

function could represent a useful approximation. In the empirical work, Boeri (1994) estimates a Cobb-Douglas matching function of unemployment and vacancies, with unemployment entering as a CES function of short and long term unemployed. Warren (1996) also uses more complex specifications in the U.S. context.

as externalities in the search process, heterogeneity in the unemployed and vacancies and lags between matching and hiring, in explaining why increasing returns to scale (IRS) may prevail (e.g., Diamond, 1982, Coles and Smith, 1994, Profit, 1996, and Mortensen 1997). IRS are conceptually important because in some models they constitute a necessary condition for multiple equilibria and provide a rationale for government intervention, striving for instance to shift the economy from a high to a low unemployment equilibrium. The radical economic reforms introduced in Slovakia in the early 2000s follow this line of reasoning.

In this paper we show that IRS appear to be an important phenomenon in all the transition countries, especially in the later (1997-2005) period, they are more pronounced in some of the economies than others and have a negative effect on the unemployment rate. We also show that there are plausible conceptual reasons for IRS in matching based on the relationship between steady-state unemployment U^* (a stock variable) and two flow variables: inflow into unemployment (labor turnover) S and vacancies V. In particular, while the high level of U^* is the focus of analytical and policy interest, the parameters of the matching function do not reflect directly the effect of S and V on U^*. We show, however, that the estimates of the matching function permit one to derive a causal relationship between U^* on the one hand and S and V on the other hand. To see this, note that a U^* is fully determined by S and O. The unemployment rate remains constant (steady state) if inflow equals outflow, so that $S = O$, and this flow identity implies that $U^*_r \equiv U/LF = 1/(1 + O_r/S_r)$, where r denotes a rate, $S_r = S/E$ is the inflow rate and $O_r = O/U$ is the outflow rate. Note that while flows S and O may change discontinuously, as may the stock variable U^*, observed U responds to instantaneous shifts in S and O by changing continuously.[9] Observed changes in U hence constitute movements either around a steady state or, in case of trends in S and O, from one steady state to another. In the conceptual framework it is useful to think of S and O as being net of seasonal and random disturbances.

For the purposes of this paper, we take S to be given exogenously by job destruction. This provides the conceptual rationale for our focus on matching because the only other determinant of U^* is O, which is determined through the matching function. The matching function may in turn be augmented to include the newly unemployed (given by inflow, S) to account for the fact that the propensity to match is higher at the time of entry into unemployment when the newly unemployed search through all existing vacancies.[10] Later on, those who remain unemployed may not search through existing vacancies that they have already explored and may instead search only through the newly posted vacancies.[11] The Cobb-Douglas matching function then becomes

$$\ln O = \beta \ln U + \gamma \ln V + \delta \ln S + \ln A \;, \tag{3}$$

and as we show in Appendix B, steady-state unemployment is given by

$$\ln U^* = \eta_s \ln S + \eta_v \ln V + \eta_A \ln A \,, \tag{4}$$

where the elasticities of U^* with respect to S and V are given by the matching function parameters as follows:

$$\eta_s = (1-\delta)/\beta > 0, \;\; \eta_v = -\gamma/\beta < 0, \quad \text{and } \eta_A = -1/\beta < 0 \,. \tag{5}$$

Steady-state unemployment U^* is hence determined by the levels of S and V, in addition to the matching function parameters contained in the elasticities of steady-state unemployment with respect to inflow and vacancies, ´$_S$ and ´$_v$. These elasticities measure the causal impact of inflow and vacancies on steady-state unemployment. Function (4) is homogenous of degree $(1 - \gamma - \delta)/\beta$,[12] implying that it is homogeneous of degree 1 only if the matching function exhibits CRS, i.e. $\beta + \gamma + \delta = 1$. IRS of the matching function, $\beta + \gamma + \delta > 1$, imply DRS of the steady-state unemployment formula because $(1 - \gamma - \delta)/\beta < 1$. For example,

9 This is analogous to the water level in a lake adjusting continuously to a sudden change in either inflow or outflow.
10 Evidence on this has been presented by numerous studies such as Coles and Smith (1994).
11 The newly unemployed may also have not yet experienced depreciations of skills and psychological scarring but this is being reflected by matching function parameters.
12 Find detailed exposition of RTS of both functions in Appendix B.

doubling *S* and *V* in a labor market exhibiting IRS in matching will result in less than doubling of steady state unemployment, which is intuitively plausible as more of both newly unemployed and vacancies is likely to result in better matching. As this example based on the steady-state unemployment equation indicates, the assumption of IRS in matching appears much more natural than that of CRS. In other words, there appears to be little rationale for assuming that doubling of both *S* and *V* has to result in doubling of steady state unemployment.

From an empirical standpoint, our theoretical analysis indicates that there is a major advantage in using the matching function because its parameters may be estimated on disequilibrium data series and yet one may infer from its coefficients the causal impact of inflow and vacancies on steady-state unemployment.

In view of the unemployment problem in the transition economies, the literature on the matching of unemployed and vacancies in these economies has grown rapidly. It has also produced contradictory results, in part because the studies use different methodologies and data. Methodologically, the studies differ especially with respect to the specification of the matching function and treatment of returns to scale, the inclusion in equation (1) of other variables that might affect outflows and the extent to which they use static or dynamic models, and with respect to whether and how they account for endogeneity of explanatory variables. In terms of data, the studies differ in whether they use annual, quarterly or monthly panels of district-level or more aggregate (regional) data and whether they cover short or long time periods. None of the studies accounts explicitly for the varying size of the unit of observation (district or region) which, as we show presently, may generate biased estimates of the returns to scale in many studies.[13]

[13] The principal studies in this area are Burda (1993), who uses monthly Czech and Slovak district-level data from 1990 to1992, Boeri (1994), who uses 1991-93 regional panel data for the Czech Republic, Hungary Poland, and Slovakia, Svejnar, Terrell and Münich (1994, 1995), who use annual 1992 and 1993 data from the Czech and Slovak Republics, Lubyova and van Ours (1994), who use 1990-93 monthly data for Slovakia and 1991-93 data for the Czech Republic, Boeri and Scarpetta

Unlike most other studies, we use a more up-to-date econometric methodology and superior data. In particular, unlike other studies we a) control for the endogeneity of explanatory variables, b) account for the presence of a spurious scale effect introduced by the varying size across units of observation (districts), and c) use long panels of comparable monthly data from all districts in the countries that we analyze. Unlike most studies, we also employ both static and dynamic specifications and estimate on contiguous panels to allow for dynamic adjustment and regime changes. Like other studies, we do not address the issue of the matching of vacancies with employed individuals (job-to-job mobility), an issue that should be taken up in future research.

3. The Estimating Framework

Theories of search and matching generally do not imply a particular functional form of the matching function. Like most studies, we use the Cobb-Douglas function given by equations (2) and (3) above. Taking equation (2) for simplicity, using lowercase letters for logarithms of variables and introducing unobserved (time invariant) district specific effects $\alpha_i = \ln A_i$ as well as an idiosyncratic error term $\varepsilon_{i,t}$, we can write (2) as

$$o_{i,t} = \beta u_{i,t-1} + \gamma v_{i,t-1} + \alpha_i + \varepsilon_{i,t} \quad \text{for} \quad t = 1,...,T, \text{ and } i = 1,...,N. \tag{6}$$

In estimating (6), one has to take into account the specific features of the matching model. Estimating by the ordinary least squares (OLS) method is not appropriate if the unobserved district specific effects α_i are correlated with explanatory variables $u_{i,t-1}$ and $v_{i,t-1}$. This correlation exists on account of structural differences between districts caused by unobserved factors that affect both α_i and $u_{i,t-1}$ or $v_{i,t-1}$. One important factor is district size, although this factor and its impact on cross-sectional estimates can be

(1995), who use monthly data for districts/regions in Poland (1992 93), Hungary (1991-94), the Czech Republic (1991-94), and Slovakia (1990 93), Burda and Lubyova (1995) who use monthly and quarterly Czech and Slovak data from 1992 to 1994, Boeri and Burda (1995), who use Czech district-level data over the period 1992 1994, Burda and Profit (1996), who use alleviated by adjusting all time varying variables in equation (6) by appropriate measure of district size.

If panel data are available, as in our case, suitable *within transformation*s of (6) can be used to remove the unobserved α_i. Both deviations from district specific means (fixed effects) and first differences remove the α_i, but the mean deviations transformation is not suitable if the model contains regressors that are only weakly exogenous. This is a relevant issue in matching functions because the explanatory variables (unemployment and vacancies) are predetermined by previous matching processes through the flow identities. In particular, omitting district subscripts for simplicity, the stock-flow identities imply that[14]

$$\begin{aligned} &U_t \equiv U_{t-1} + S_t - O_t \\ &U_{t-1} \equiv U_{t-2} + S_{t-1} - O_{t-1} \\ &U_{t-2} \equiv U_{t-3} + S_{t-2} - O_{t-2} \\ &\ldots \end{aligned} \tag{7}$$

Lagged outflows in (74) are in turn given by a lagged version of matching function (6) as

$$\begin{aligned} &\ln O_{t-1} = \beta \ln U_{t-2} + \gamma \ln V_{t-2} + \alpha_i + \varepsilon_{t-1} \\ &\ln O_{t-2} = \beta \ln U_{t-3} + \gamma \ln V_{t-3} + \alpha_i + \varepsilon_{t-2} \\ &\ldots \end{aligned} \tag{8}$$

Since any district mean is computed from all district observations over time, the means contain all values of the error term $\{\varepsilon_\tau: \tau = T, T-1, \ldots, 1\}$. This creates correlation between transformed explanatory variables and transformed error terms and hence leads to biased estimates.

The first difference transformation contaminates the transformed variables only with recent error terms $\{\varepsilon_\tau: \tau = t-1, t-2\}$, as may be seen by rewriting (6) in a first difference form as

$$\Delta o_t \equiv o_t - o_{t-1} = \beta(u_{t-1} - u_{t-2}) + \gamma(v_{t-1} - v_{t-2}) + \varepsilon_t - \varepsilon_{t-1}, \tag{9}$$

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district and regional 1992-94 data from the Czech Republic, and Profit (1996), who uses Czech district data during 1992-94. For a brief survey of the principal studies se Münich et al. (1997).

[14] These identities assume that all matches are brought about by the reported unemployed and vacancies (there being no out-of-register matching). Other forms of matching may create more complicated identities but will not eliminate the

which may in turn be expressed in a simplified notation as

$$\Delta o_t = \beta \Delta u_{t-1} + \gamma \Delta v_{t-1} + \Delta \varepsilon_t . \tag{10}$$

From equation (7) it follows that u_{t-1} and u_{t-2} contain ε_{t-1} and ε_{t-2}, respectively, through outflows as in equation (8).[15]

The first difference transformation thus leaves further lags of $\{\Delta u_t : \tau = t\text{-}2, t\text{-}3, \dots, 2\}$[16] uncorrelated with $\Delta\varepsilon_t$ (i.e., with ε_t and ε_{t-1}) and these further lags of Δu_t can be used as instrumental variables. Vacancies in (6) and (9) are predetermined in the same statistical sense and may be treated the same way. Available instruments are therefore given by $\{\Delta u_\tau, \Delta v_\tau : \tau = t\text{-}2, t\text{-}3, \dots, 2\}$.

There are several additional features of the matching function model that need to be taken into account in estimation. First, identities in (7) show that lagged changes in inflows $\{\Delta s_\tau : \tau = t\text{-}1, t\text{-}2, \dots, 2\}$ are available as additional instruments because they codetermine the explanatory variables but do not affect current outflows directly.

Second, rather than using changes in lagged values as instruments, we can use lagged levels because differences are simply specific linear combinations of levels.

Third, it is desirable to include month and year specific dummy variables as regressors to control for the sizeable seasonality typically contained in the unemployment flows.

Fourth, although idiosyncratic errors $\{\Delta\varepsilon_\tau : \tau = t\text{-}1, t\text{-}2, \dots, 2\}$ may in principle be uncorrelated, their first differences $\Delta\varepsilon_t$ will by definition be autocorrelated due to $\Delta\varepsilon_t$ which is contained in both $\Delta\varepsilon_t = \varepsilon_t - \varepsilon_{t-1}$ and $\Delta\varepsilon_{t-1} = \varepsilon_{t-1} - \varepsilon_{t-2}$.[17] As a result, to obtain unbiased standard errors for the estimated coefficients, we use a robust variance-covariance matrix.

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problem of weak exogeneity.

[15] It does not change the essence of the argument that (74) is defined in levels and (85) in logs.

[16] Note that the first observation (for $t = 1$) of the first differenced variables is not available because observations for $t = 0$ do not exist.

[17] Note that $COV(\varepsilon_t - \varepsilon_{t-1}, \varepsilon_{t-1} - \varepsilon_{t-2}) = -VAR(\varepsilon_{t-1}) < 0$.

Fifth, further complications arise if ε_t has an autoregressive form. In this case the problem is that current ε_t and lagged ε_{t-1} (and therefore $\Delta\varepsilon_t$) contain previous values $\{\Delta\varepsilon_\tau\text{: } \tau = t\text{-}1, t\text{-}2, \ldots, 2\}$ and this disqualifies past lags as instruments. This problem is fortunately substantially weakened by first differencing. Moreover, the correlation of the instruments and error term declines quickly with the length of lags unless the correlation is very high. In our case further lags of the explanatory variables remain reasonably good instruments.

Sixth, equation (6) represents a full adjustment model, which assumes that changes in the stocks of unemployed and vacancies translate into outflow immediately. Since in practice there may be frictions and reporting delays, it is useful to consider also a partial adjustment model of matching, where outflow reacts to changes in the explanatory variables only partially during each period:

$$o_{i,t} = \varphi o_{i,t-1} + \beta u_{i,t-1} + \gamma v_{i,t-1} + \alpha_i + \varepsilon_{t,t}\,. \tag{11}$$

Transformed into first differences, Δo_{t-1} on the right hand side is endogenous by definition and contains ε_{t-1} and ε_{t-2}. It is therefore correlated with $\Delta\varepsilon_t$ and leads to the same problem as that caused by the endogeneity of u_{t-1} and v_{t-1}. Lagged difference in outflow hence has to be instrumented and one can consider further lags $\{\Delta o_\tau\text{: } \tau = t\text{-}2, t\text{-}3, \ldots, 2\}$ or corresponding levels as additional instruments.

Seventh, since the within transformation removes fixed-effects α_i, these are not estimated in the first stage. If explanatory variables contain a time-invariant district specific mis-measurement component, this component is removed by the within transformation too. This is the case of vacancies in our special case. The extent to which all vacant jobs are reported to labor office depends on country-specific legislation and could also depend on district-specific administrative effort. While the problem of mis-measurement of vacancies is in general controlled for by the within transformation, the problem would remain if we wanted to estimate the district fixed-effects. This is because the commonly used estimates of fixed-effects[18] are biased upward as a result of the (time-invariant) under-reporting of vacancies. Given that the extent of under-reporting has district-specific and country-specific components, a comparison of estimated fixed-effects across districts and across-countries, although very desirable, would not allow us to distinguish between the actual efficiency of matching and extent of mis-measurement.

As mentioned above, model (6) does not account for the fact that the propensity to match is higher at the time of entry into unemployment when the newly unemployed search through all existing vacancies. To reflect this so called "stock-flow" matching, we include $lnS_t = s_t$ as an additional explanatory variable in (3), while ensuring that current inflow S_t is not part of U_{t-1}. Assuming that job destruction is exogenous with respect to actual matching, o_t, no additional instruments are needed.

Adjusting for Varying Size of Districts

Since the empirical literature on matching has not taken into account the variation in the size of the unit of observation (in our case the district), many of the existing studies have probably generated biased estimates.[19] The reason for the bias, explained in detail in Appendix A, relates to the fact that the size of a district, measured for instance by its labor force L_i, is positively correlated with the levels of O_i, U_i, S_i, and V_i simply because of different sizes of districts. In this situation, when district-level variables are not adjusted for the size of the district labor force, inter-correlations between O_i, U_i, S_i, and V_i are born by economic relationships and also due to variation in district size. The later contributes positively to the overall correlation and tend to dominate the former one (see Appendix Table A2 for an illustration). The

[18] When imputing matching function intercept from equation (6) as $\hat{\alpha}_i = \overline{o}_i - \hat{\beta}\overline{u}_i + \hat{\gamma}\overline{v}_i$ using estimated parameters and actual levels of variables, under-reporting in vacancies biases estimated intercept [illegible] upward and the bias increases with higher parameter [illegible]. In other words, larger under-reporting and larger marginal impact of vacancies implies higher bias in the intercept. Since the scope of under-reporting is not known and differs across districts and countries, one cannot distinguish the economic component from the one brought about by under-reporting.

[19] There are several possible measures of district size. We use the district labour force, but the results would not be materially affected by using other measures.

usual Cobb-Douglas specification estimated on cross-sectional data then provides estimates of coefficients biased unless the returns to scale are constant or the unadjusted U_i and V_i are uncorrelated with the district size. The direction of the bias of $\beta(\gamma)$ is negative if U_i (V_i) is positively correlated with L_i, (measure of district size) which is the most common case, and matching displays increasing returns to scale. Either decreasing returns to scale or negative correlation (but not both) in turn lead to positive bias. Therefore, if the matching process does not exhibit constant returns, the bias is likely to cause an incorrect acceptance of the constant returns hypothesis. The bias is greater, the greater is the portion of the correlation of U_i and V_i with L_i that is due to pure differences in the size of the district labor force. As we show in Appendix A, the bias is specific case of classical omitted variable problem. In what follows, we call this phenomenon the *spurious scale effect*.[20]

It can be shown that the spurious scale effect is avoided if one estimates a Cobb-Douglas function with panel data and accounts properly for the presence of fixed effects, as district size represents one of them. In that case, the within transformation removes the spurious scale effects together with all other unobserved district-specific time-invariant effects captured by [illegible]$_i$'s.

4. The Data and Variable Definitions

In order to produce the best possible parameter estimates, we have assembled an extensive panel of data on 74 Czech, 38(79) Slovak, 20 Hungarian, 49(16) Polish**,** 34 East German and 140 West German districts. The data for all countries except Hungary cover the period from January 1991[21] - July 2005,

[14] An interesting question for future research is whether the size of districts and regions, the usual units of observation in the matching function studies, tends to be determined by an arbitrary administrative fiat or an endogenous optimization process of population settlements, based on historical economic forces that are in principle similar to an optimization process determining the size of firms.

[21] In January 1997, three new Czech districts were formed from two original districts. These three districts are excluded from the analysis. German data exclude districts of Berlin due to data inconsistencies. The structure of Slovak districts was thoroughly changed in 1997 and we use Slovak data as two separate panels. District level data for 49 Polish voivodships are available only till the end of 1998.Afterwards, data are available only by 16 Polish regions and we use Polish data as two separate panels. The Hungarian data at our disposal start in January 1995.

while for Hungary they cover January 1995 – December 2004. The data sets contain monthly observations for the following variables:

$O_{i,t}$ = the number of individuals flowing from unemployment in district *i* during period *[t-1,t]*;

$U_{i,t-1}$ = the number of unemployed in district *i* the end of period *t-1* (i.e. beginning of period *t*);

$S_{i,t}$ = number of individuals flowing into unemployment (the newly unemployed) in district *i* during period *[t-1,t]*;

$V_{i,t-1}$ = the number of vacancies in district *i* at the end of period *t-1* (i.e. beginning of period *t)*;

Although outflow to jobs is a theoretically preferred variable to total outflow, the data on outflow to jobs are available only for the Czech Republic, while data on total outflow are available for all the countries in our study. We have first carried out the estimation for the Czech Republic using both measures and found that the estimates based on total outflow and outflow to jobs are similar.[22] As a result, we assume that the lack of data on outflow to jobs in other countries does not have a dramatic impact on our results (see also Petrongolo and Pissarides, 2001, for similar evidence from other countries).

5. Econometric Estimates

5.1 Basic Statistics

The basic statistics are described in Table 1 and Table 2, with the unemployment and flow data of Table 1 having been discussed above. In Table 2 we provide information on the extent to which shocks and restructuring altered in each country the rankings of districts in terms of various labor market indicators over time. In particular, we present the Pearson's rank correlation coefficients of specific

[22] Total outflows and outflows to jobs are positively correlated, with the latter representing about 75 percent of the former in the Czech Republic.

district-level outcomes (the unemployment, inflow, outflow, and vacancy rates, and U/V ratio) between a given year in early and late transition. A coefficient close to 1 indicates that the shocks and restructuring affect district-specific labor markets in different ways in the sense that for the given indicator the ranking of districts changed little between the two periods, a coefficient of 0 indicates that the ranking of districts is completely unrelated in the two periods and a negative coefficient would indicate reversal patterns in the ranking. In Table 2, we show the rank correlation coefficients for 1992-96 and 1992-99 in the left hand side panel, and correlations between 1999-2002 and 1999-2005 in the right hand side panel.[23] The correlation coefficients suggest that there were shocks and restructuring (in terms of district rankings) taking place in the transition economies during the early-to-mid 1900s and that the speed of this restructuring slowed down during the late 1990s and early-to-mid 2000s. Except for the outflow rate, the extent of restructuring was less pronounced in Poland than in the other transition countries in the 1990s, and in this sense Poland resembles West Germany, which also generates correlation coefficients of less than unity but higher than in the other economies. East Germany represents the other extreme with the lowest values of the correlation coefficients, especially in the 1990s. The correlation coefficients also indicate that in all six economies the shocks and restructuring affected most the vacancy rates – suggesting that there were major inter-regional changes on the demand /job creation side of these economies over time. This pattern holds in the 1990s as well as 2000s and is also to a lesser extent reflected in the other potential measure of demand, namely the outflow rate.[24] Interestingly, the results in Table 2 indicate that shocks and restructuring have not affected in a major way the district ranking for unemployment, inflows and U/V rates, with East Germany again standing out and showing the smallest rank correlation coefficients for these variables over time. Hence, both in the benchmark market economy

[23] The year 1996 is the last year for which we have Slovak data based on the initial classification of districts. Similarly, 1999 is the first year for which reduced structure of Polish districts is available.
[24] The outflow rate measure measures total outflows but in some countries (especially East Germany) it contains also significant outflows into training programs.

(West Germany) and in all the transition economies other than East Germany, we observe considerable persistence in the relative standing of districts with respect to unemployment, job destruction (inflows) and labor market tightness (*U/V ratio*). Transition-related shocks and restructuring alone are hence not an explanation for the persistence in the relative standing of districts along these dimensions of the labor market. Moreover, transition-related factors are also relatively less important for all five labor market indicators in the 2000s since during this period West Germany registers equal or smaller correlation coefficients than the transition economies other than East Germany.

5.2 Econometric Estimates

We start our discussion by presenting in Table 3 various estimates of the matching function (3) for the West German districts during the 1997-2005 period. The West German estimates provide a benchmark for a mature market economy against which we compare the estimates from the five transition economies, including East Germany. In Table 3 we first compare results generated by the techniques used in the literature that may generate biased estimates, followed by estimates that correct for the aforementioned problems. In particular, we first present OLS estimates in Panel A, followed by standard panel data estimates in Panel B, and finally what we consider to be the most appropriate estimation method, namely first-difference IV estimates in Panel C.

The OLS estimates of coefficients on unemployment and vacancies in panel A of Table 3 are low and as may be seen from the p-values for the test of constant returns to scale, they imply decreasing returns to scale ($\beta + \gamma < 1$). Including monthly (or annual) time dummy variables has only a negligible impact on the estimates. The OLS estimates of β and γ based on variables adjusted for district size are smaller than those from unadjusted OLS, which is not surprising given that the spurious scale effect biases the coefficients toward constant returns. As discussed earlier, both sets of OLS estimates are inconsistent due to the presence of unobserved fixed effects and if, as is likely, these unobserved effects are negatively correlated with unemployment and vacancies, the estimated coefficients are downward biased. As we will see presently, this appears to be the case for all cross-sectional estimates in Table 3.

The random effects and mean deviations (fixed effects) estimators presented in panel B of Table 3 yield β coefficients that are somewhat larger than the corresponding OLS estimates, and γ coefficients that are somewhat smaller than their OLS counterparts. As a result, the returns to scale ($\beta + \gamma$) are similar at around 0.8. As discussed above, the estimates based on mean deviations are still biased due to endogeneity. The OLS estimates based on first differences have a notably higher coefficient on unemployment ($\beta = 1.64$), implying increasing returns to scale ($\beta + \gamma > 1$). However, these estimates are also biased because Δu_{t-1} contains $-\varepsilon_{t-1}$ in u_{t-1} through the stock-flow identity, and $-\varepsilon_{t-1}$ is contained also in $\Delta\varepsilon_t$. This induces positive correlation between the transformed error term $\Delta\varepsilon_t$ and both explanatory variables Δu_{t-1} and Δv_{t-1}, and brings about a positive bias and therefore higher coefficients observed in our estimates. We also present estimates based on forward means deviations. Transformation of variables in (6) into deviations from district specific forward means leaves lagged observations as valid instruments. These estimated parameters are close to our preferred estimates that we present in Panel C of Table 3. The preferred estimates come from an IV method based on first differences of variables. We report these preferred estimates in two versions: with and without the newly unemployed (inflow) being included as a regressor. The model without the newly unemployed yields coefficients $\beta = 1.32$ and $\gamma = 0.14$. These estimates are consistent. The instruments used are lagged levels of explanatory variables plus lagged inflows, with close lags for *t-2* and *t-3* being excluded to secure strict exogeneity. In all of our empirical work, we find that the explanatory power of the proposed instruments is adequate.[25] When the newly unemployed are included in the regression (second to last row in Table 3), we find that their coefficient δ

[25] The instruments explain 20% to 70% of variation in the explanatory variables. The lowest explanatory power of instruments was for vacancies (20-30%), and the highest power was for outflows (60-70%). It should also be noted that first differencing removes time invariant component of measurement errors, decreases the variance of the explanatory variables while doubling the variance of the error term and potential idiosyncratic measurement error, and leads to higher standard errors of estimated parameters. Measurement error may also be present in vacancies. While its time invariant component is removed by first differencing, its idiosyncratic part causes negative bias and our estimates hence represent the lower bound of actual coefficients.

is 0.12 and statistically significant. As we show in Table 4, these estimates imply that the newly unemployed display a higher propensity to match than the existing unemployed. Indeed, when one converts the estimated elasticities β (δ) into the probability that an additional unemployed (a newly unemployed) person flows out at the mean of U (S), one finds that these probabilities are statistically different from one another (see Appendix C).

The last row in Table 3 indicates that while the estimates of the basic coefficients remain virtually unaffected, a partial adjustment model is an appropriate specification since the coefficient on lagged outflow is $\varphi = 0.2$ and it is statistically significant. The matching process hence appears to be better captured as a dynamic phenomenon, but the estimated monthly extent of adjustment is estimated to be relatively fast at 0.8.[26] Finally, the disembodied improvement in the efficiency of matching, as captured by the estimated *trend*, is about 1 percent per year.

In Table 4, we present IV first-difference estimates of the parameters of the matching function for the Czech Republic, Hungary, Poland, Slovak Republic, East Germany, and West Germany. In order to capture the potential differences in the functioning of the labor markets in the early-to-mid 1990s and the late 1990s to mid 2000s, respectively, we provide separate estimates for the 1994-96 period in panel A and 1997-2005 period in panel B of Table 4. The earlier period corresponds to the early transition in the post-communist countries and a period of relatively slow economic growth in West Germany. The latter period captures the late transition in the post-communist economies and a period of relative boom and later slowdown in West Germany. For the earlier period, we do not have data for Hungary, but for the latter period we have data on all six economies. In all cases, we present coefficients from the first-difference IV model with unemployment, vacancies and inflow into unemployment (newly unemployed) as regressors. In view of our theoretical discussion, we also present returns to scale (RTS) for both the

[26] Note that including lagged outflow as explanatory variable represents partial adjustment model where $1-\varphi$ is the proportion of gap between actual and equilibrium level of outflow being closed during period *[t-1, t]*.

matching function, $RTS_M = \beta + \gamma + \delta$, and for the steady-state unemployment function, $RTS_U = 1 + (1-RTS_M)/\beta$.

As may be seen from Table 4, the estimated coefficients on unemployment, vacancies and newly unemployed vary considerably across the six economies and, except for West Germany, also across the two time periods. During the 1994-96 period, we observe relatively precisely estimated coefficients in the Czech Republic and West Germany, pointing to moderately increasing (1.24) and highly increasing (1.69) returns to matching, respectively. However, while the returns to scale for West Germany are precisely estimated, for the Czech Republic one cannot reject the hypothesis that there are constant returns to scale. The main difference between the two countries lies in a much higher coefficient on unemployment (1.27 vs. 0.75) and more precisely estimated coefficient on vacancies in West Germany than in the Czech Republic. In both East Germany and Slovakia, in 1994-96 the coefficients on unemployment and vacancies (and hence also returns to scale) are very imprecisely estimated, suggesting that there was a considerable diversity of matching patterns across the districts in these two economies. In Poland, where the districts are substantially larger than those in the other countries, we get a very high coefficient of 2.60 on unemployment and an imprecisely estimated coefficient on vacancies. The high coefficient on unemployment also drives high (2.95) returns to scale. Finally, in all five economies we find a similar (0.17 to 0.27) and precisely estimated coefficient δ on inflows. When we compute the transformed coefficient δ' that is directly comparable to the coefficient β on the number of unemployed (see Appendix C), we see in Table 4 that coefficient δ' for the newly unemployed is 2.0 to 4.4 times larger than the coefficient β on the existing unemployed. The only exception is Poland in the first period and Slovakia in the second period, where the two coefficients are not statistically different from each other.

During the more recent period of 1997-2005, we obtain precisely estimated coefficients in all the countries. With one exception, the returns to scale are increasing in all the countries, with the highest returns being observed in Hungary (2.40) and East Germany (2.14), with the Czech Republic, Slovakia and West Germany coming in next with increasing returns of 1.86, 1.82 and 1.56, respectively. All these estimates are significantly different from 1.0 at the conventional significance test levels. The Polish point estimate suggests that there are constant returns, but the estimate is very imprecise.

The results in Table 4 also indicate that in the early (1994-96) transition period posted vacancies played a negligible part in outflow from unemployment. In contrast, unemployment was an important determinant of outflow in the Czech Republic, which maintained a low unemployment rate, East Germany, which had a high unemployment rate but also a high inflow rate and very sizable active labor market programs, and interestingly also in Poland, where the high estimated coefficient of 2.6 on unemployment suggests that Poland had a highly positive externality from workers to firms (increasing search and matching intensity with rising unemployment). However, unemployment was a statistically unimportant determinant of outflows in Slovakia, which experienced a rapid rise in unemployment during this period. Interestingly, inflow into unemployment generated a similar (0.2-0.3) and statistically significant coefficient δ that translates into an adjusted coefficient of inflows relative to existing unemployed δ'/β of 2.0 to 4.3 in all these economies. This suggests that the incidence of job matching involved in a major way the newly unemployed. In the second (1997-2005) period, the difference in the efficiency of matching of the new and existing unemployed remained, but it diminished somewhat in the Czech Republic, Slovak Republic, East Germany and West Germany, and increased in Poland (lack of data prevents a comparison in Hungary).

In terms of hypothesis H1b, we observe that between the 1994-96 and 1997-2005 period, returns to scale in matching, $RTS_M = \beta + \gamma + \delta$, rise in all the transition economies except Poland, but decline somewhat in West Germany. (Note that growing returns to scale lead to lower steady-state unemployment for given turnover and vacancies.)[27] Leaving Poland aside, we observe that Hungary and East Germany have the highest returns to scale, driven by relatively high coefficients on all three explanatory variables – unemployment, vacancies and inflow into unemployment. In East Germany, this may be in part generated by the very sizable active labor market policies that were targeting high unemployment districts. These are present in Hungary as well, but not on the same scale. The Hungarian results hence suggest that the underlying feature is a relatively efficiently functioning matching system, but one must remember that in Hungary the declining labor force participation means that some exits from unemployment were into the "out of the labor force" state. These findings are consistent with the high unemployment in East Germany and low unemployment in Hungary. The Czech returns to scale are lower than those in Hungary and East Germany, but they are similar to those of Slovakia. However, the Czech-Slovak similarity in returns to scale disguises important differences in terms of a higher estimated coefficient on unemployment and lower coefficients on vacancies and inflow, as well as relatively fewer vacancies and higher inflows, in Slovakia than in the Czech Republic.

Turning to the parameters of the equilibrium (steady state) unemployment equation, we observe in Table 4 that during the 1994-1996 period in all transition countries except Poland the elasticity of equilibrium unemployment with respect to inflow (´$_s$) was higher compared to the period of late transition, while it stayed almost unchanged in West Germany. This implies that steady-state unemployment was more sensitive to growing inflow in the earlier stages of transition. Yet, as may be seen from Table 1, inflow was relatively higher in the later period in all countries – a finding that is somewhat surprising, given earlier expectations (see Münich and Svejnar, 2007). These two simultaneous effects – decreasing elasticity of inflow and growing inflow – had an opposite (offsetting) impact on unemployment. The elasticity of equilibrium unemployment with respect to vacancies (´$_v$) became more

[27] See Appendix B for details.

negative in the Czech Republic, Slovakia and East Germany, but remained statistically insignificant in Poland. Note that in Poland the number of vacancies relative to the size of the market is several times smaller than we commonly find in other countries. In West Germany, the impact of vacancies did not change much and remained relatively low in comparison to the transition countries where unemployment became more sensitive to vacancies. It should be also noted that in West Germany the number of vacancies, unlike the size of inflow, stayed at similar levels in both periods. The resulting returns to scale for the steady-state unemployment function, $RTS_U = 1 + (1- RTS_M)/\beta$, rise in all the countries except for West Germany and Poland.

Finally, while returns to scale in matching represent one element of efficiency (H1b), the value of the matching function intercept A represents the other (disembodied) element of efficiency of matching (H1a).[28] Assuming, that the extent of under-reporting is time-invariant, we can identify country specific trends in matching function estimates. As may be seen in Table 4, for 1997-2005, the estimated annual trend (capturing the change in A) is positive in West Germany and Poland, insignificant in Hungary and Slovakia, and negative in the Czech Republic and East Germany. The positive West German trend constitutes a reversal of a negative trend in the 1990s, while Poland appears to be improving matching efficiency during both periods. In the top panel of Figure 1 we show trends that are obtained from repeated estimations on two years of data (24 months), from which we consecutively remove the oldest month and add the newest one. The estimates are presented for the three countries that did not change their district structure during 1992-2005 -- Czech Republic, East Germany and West Germany. The dots in the figures are the individual point estimates for each position of the window and the line represents

[28] As mentioned earlier, in practice the estimated values of this parameter also contain the effect of systematic under-reporting of vacancies, which may be quite common. Measurement error biases are to great extent diluted by the within transformation which removes the systematic component of under-reporting. However, when imputing matching function intercept from equation (6) as [illegible] = *o* - ¯*u* - [illegible]*v* - ±*s*, using estimated parameters and actual levels of variables, under-reporting in vacancies biases estimated intercept [illegible] upward. In other words, larger under-reporting implies higher intercept. Since the scope of under-reporting is not known and differs across countries, one cannot distinguish the economic component from the one due to under-reporting.

smoothed point estimates. The estimates in Figure 1 suggest that the negative trend in the Czech Republic is becoming less negative, while in East Germany the trend oscillates around zero with diminishing amplitude over time. In West Germany, the trend turned from negative one in late 90s to slightly positive during 1999-2002, a development that was most likely linked to the economic upswing in the latter period. However, since the late 1990s the West German trend is similar to that in East Germany, namely oscillating around zero.

In the bottom panel of Figure 1 we present estimates of returns to scale estimated by the same 24-month wide moving window technique. The figure shows that the returns are most volatile and relatively high in East Germany, with the difference being the most pronounced in the earlier (1994-98) and the most recent (2001-05) periods. The Czech and West German returns to scale are similar, with the Czech returns being below the West German ones in 1994-98 and exceeding the West German ones in the 2002-04 period. Overall, the bottom panel of Figure 1 supports our earlier findings that in terms of matching, the Czech and West German labor markets seem to be more similar than either one of them is to the East German market.

6. Concluding Observations

Our paper is motivated by the three alternative hypotheses about the causes of unemployment in the Central European transition economies and in the benchmark market economy (Western part of Germany). The first hypothesis (H1) is that high unemployment is caused by inefficient matching, in terms of either low or negative trend in disembodied efficiency (H1a) or low returns to scale in matching (H1b) . In this case, the policy should focus on labor market institutions and measures to stimulate labor mobility and create appropriate skills. Hypothesis 2 (H2) is that high unemployment is caused by low demand for labor. The manifestation of this would be low vacancies relative to inflows and unemployment and the policy implication would be that macroeconomic policies are key for resolving the unemployment problem. Hypothesis 3 (H3), namely that high unemployment implies that restructuring is at work, is consistent with unemployment being high because of relatively high labor turnover in firms due to ongoing restructuring, with the policy implication being that restructuring needs to be completed.

Our data and econometric estimates suggest that the situation differs across the sampled economies and over time different hypotheses receive support in different countries. Our benchmark market economy, namely the West German part of Germany, is an economy with slowly rising inflow and unemployment, declining vacancies and relatively efficient matching (high returns to scale in both periods and rising disembodied efficiency of matching in the second period). Its outcome is hence most consistent with H2 and H3. Czech Republic appears to be in a similar situation, and with rising unemployment, as well as inflow and outflow, and a declining vacancy rate and high returns to scale in matching, it increasingly gives support to H2 and H3. East Germany's results are also in line with H2 and H3, in that the region has relatively high unemployment and inflows, a low vacancy rate and very efficient matching in terms of returns to scale (including outflow into the training programs) but a negative trend in (disembodied) efficiency. The Slovak economy displays high inflows and unemployment, low vacancies and outflows, and increasing returns to scale as well as a positive trend in efficiency. It is therefore also consistent with H2 and H3. Hungary has a relatively low unemployment rate, highest increasing returns to matching of all the six economies, and moderate inflow, outflow and vacancy rates. As such it does not fit into any, or alternatively provides limited support to all, of the three hypotheses. It should be noted that since all these economies have pursued a policy of low interest rates and fiscal deficits, the support for H2 implies the presence of negative exogenous demand shocks rather than restrictive macro policies. Finally, Poland is consistent with H1 as it has high unemployment in the presence of low vacancies and outflow, as well as constant returns to matching. It is also the one economy where the Central Bank followed a relatively high interest rate policy.

Overall, our findings suggest that the Central European transition economies contain one broad group of countries and one or two special cases. The group comprises the Czech Republic, Hungary, Slovak Republic and (possibly) East Germany. These countries resemble West Germany in that they display increasing returns to scale in matching and unemployment appears to be driven by restructuring and low demand for labor. The East German case is complex because of its major active labor market policies and a negative trend in efficiency in matching. In some sense, East Germany resembles more Poland, which in addition to restructuring and low demand for labor appears to suffer from a structural mismatch reflected in relatively low returns to scale in matching. The overarching portrayal of the labor market in all of these economies is that it is affected by ongoing long-term restructuring in the presence of limited demand for labor, while regional disparities in unemployment, inflows and outflows are quite persistent over time. Interestingly, relative positions of individual districts within countries according to stock and flow labor market rates are still changing, including West Germany, although the rate patterns in all transitional economies were notably less persistent in the early transition period.

Finally, our data provide evidence that goes counter to one of the main predictions of the theories of transition, namely that the turnover (inflow) rate in the transition countries would rise dramatically at the start of the transition, be temporarily very high and gradually decline and approach the level observed in otherwise similar market economies such as West Germany. First, except for East Germany, none of the transition countries had a considerably higher inflow rate than West Germany during the entire 1991-2005 period. In fact, some of the countries had a lower inflow rate than West Germany for extended periods of time. Second, by the mid 2000s the inflow rate in all economies except East Germany converges to a similar range. Third, by the mid-2000s the West German inflow rate actually exceeded the rate observed in the Czech Republic, Hungary and Slovakia, and was similar to that in Poland.

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Appendix A: The Spurious Scale Effect

For the purposes of exposition, we present a simple case that demonstrates the impact of the spurious scale effect on estimation. Assume that the country is a homogeneous territory divided administratively into districts of different sizes with identical labor market conditions and characterized by a simple Cobb-Douglas matching function with increasing returns to scale *($\beta + \gamma > 1$)*. As a result of homogeneity, the outflow, unemployment and vacancies in each district, O_i U_i *and* V_i, are precisely proportional to national aggregates *O, U, V*, and

$$O_i = l_i O, \quad U_i = k_i U, \quad V_i = k_i V, \tag{A1}$$

where l_i is the share of district *i* in the national labor force, defined as L_i / L. Note that for expositional purpose the variance in district level variables is brought about completely by the administrative variation in district sizes rather than by economic factors. Not taking district size into account and estimating log - transformed matching function on unadjusted cross-sectional data amounts to estimating

$$\ln O_i = \alpha + \beta \ \ln U_i + \gamma \ln V_i + \varepsilon_i \ . \tag{A2}$$

Substituting (A1) into (A2) we get

$$\ln l_i = (\beta \ + \gamma) \ln l_i + (\alpha - \ln O + \beta \ \ln U + \gamma \ln V) + \varepsilon_i \, . \tag{A3}$$

Estimating (A2) is identical to estimating (3). However, (3) represents a regression of l_i on itself plus a constant term. It will therefore tend to estimate constant returns to scale *($\beta + \gamma = 1$)* and a zero constant term *($\alpha = lnO - \beta lnU - \gamma lnV$)*. Note that we have assumed increasing returns to scale. In reality, regions are not perfectly homogenous and model (A2) yields estimates biased toward constant returns to scale. A remedy for this problem is to adjust the variables by the district size in order to obtain the following model:

$$\ln(\frac{O_i}{L_i}) = \alpha + \beta \ \ln(\frac{U_i}{L_i}) + \gamma \ln(\frac{V_i}{L_i}) + \varepsilon_i \ , \tag{A4}$$

which may be rearranged as

$$\ln O_i = \alpha + \beta \ \ln U_i + \gamma \ln V_i + (\ \beta_u + \beta_v - 1) \ln L_i + \varepsilon_i \, .$$

A comparison of the adjusted model (A4) to the unadjusted model (A3) indicates that they are equivalent if and only if at least one of the two following conditions is satisfied:

(i) $\beta + \gamma = 1$ (the underlying matching displays constant returns to scale)

or

(ii) $Cor(lnL_i, lnU_i) = Cor(lnL_i, lnV_i) = 0$.

In our example, neither condition is satisfied because (i) we are assuming increasing returns to scale *($\beta + \gamma > 1$)* and (ii) $U_i = UL_i/L$ *and* $V_i = VL_i/L$, resulting in $Cor(lnL_i, lnU_i) > 0$ *and* $Cor(lnL_i, lnV_i) > 0$. In general, one has no *a priori* information about the returns to scale since they represent a statistic that is to be estimated from the data. The inter-correlations among the unadjusted variables can of course be checked in advance. Judging from the data at our disposal, these inter-correlations are positive and significant.

Appendix B: Matching Function and Equilibrium Unemployment

Comparative Statics

The steady-state inflow into unemployment is an outcome of labor market turnover. In a steady state, unemployment is stable and outflow from unemployment equals inflow. Suppressing time subscripts since steady-state values do not change over time, for a given level of exogenous inflow and vacancies, matching function

$$O = A\, U^{\beta} V^{\gamma} S^{\delta} = S \ , \tag{B1}$$

implies steady-state unemployment

$$U^{*} = \sqrt[\beta]{\frac{S^{1-\delta}}{A\, V^{\gamma}}} \ . \tag{B2}$$

The function determining steady-state unemployment in (B2) is homogenous of degree $RTS_U = (1 - \gamma - \delta)/\beta$. Noting that the matching function is homogenous of degree $RTS_M = \beta + \gamma + \delta$, there is a one-to-one relationship $RTS_U = 1 + (1 - RTS_M)/\beta$. It implies that function (B2) is homogeneous of degree 1 only if the matching function in (B1) exhibits constant returns to scale, i.e. $RTS_M = \beta + \gamma + \delta = 1$. Increasing returns to scale of the matching function, $RTS_M > 1$, imply decreasing returns to scale of the steady-state unemployment formula because $RTS_U < 1$. For example, doubling S and V in a labor market exhibiting increasing returns to scale in matching will result in less than doubling of steady state unemployment. Increasing returns to matching therefore reflect an existence of equilibrium reinforcing market forces.

Note that the matching function does not impose any particular constraint on the relationship between inflow and the number of vacancies. A proportional increase in both S and V is a specific type of shift along a continuum of other simultaneous shifts that are possible. We do not need to specify the inflow-vacancy relationship as long as we consider shifts in S and V to be exogenous and limit our attention to their impact on steady state unemployment and account for their weak endogeneity when estimating matching function parameters.

Equation (B2) may be used to compute the number of additional vacancies needed to keep steady-state unemployment unchanged when turnover (inflow) increases. It can be shown that if turnover is increased k-times, vacancies have to increase $k^{(1-\delta)/\gamma}$ times to maintain unemployment unchanged. Clearly, the number of vacancies needed to compensate for growing turnover (to secure unchanged steady state unemployment) is lower when matching of the newly unemployed is better (higher δ) and matching of vacancies is better (higher γ).

Equation (B2) may be rewritten as

$$U^{*} = [A^{-1} S^{(1-\delta)} V^{-\gamma}]^{1/\beta}$$

and expressed in logs as

$$\ln U^{*} = \Big[(1-\delta)\ln S \ - \gamma \ln V \ - \ln A \ \Big] / \beta. \tag{B3}$$

In the form of (B3) the equation reveals determination of steady-state unemployment by inflow, vacancies, and the matching function parameters. More conveniently, the impact of inflow and vacancies may be expressed in terms of elasticities as

$$\ln U^{*} = \eta_S \ln S + \eta_V \ln V + \eta_A \ln A \tag{B4}$$

where

$$\eta_s = (1-\delta)/\beta,\ \eta_v = -\gamma/\beta,\ \eta_A = -1/\beta. \tag{B5}$$

Theoretical assumptions $\beta \geq 0$, $\gamma \geq 0$, and $0 \leq \delta \leq 1$ imply that $\eta_s \geq 0$, $\eta_v \leq 0$, and $\eta_A \leq 0$. In economic terms, steady-state unemployment increases with the turnover and the impact of turnover is lower when the matching of newly unemployed is higher (captured by δ). Similarly, vacancies have negative impact on steady-state unemployment and the impact increases with γ.

Dynamics – Transitions between Steady States

Dynamic transition from one steady-state to another as a reaction to an exogenous change in inflow $S^{*} \rightarrow S^{**}$ (or $V^{*} \rightarrow V^{**}$) may be better understood by taking into account the fact that a change in unemployment is given by the difference between inflow and outflow. Therefore,

$$\frac{dU}{dt} \equiv \dot{U} = S^{**} - O = S^{**} - AU^{*\beta}V^{\gamma}S^{**\delta} = S^{**}\left(1 - \frac{AV^{\gamma}}{S^{**(1-\delta)}}U^{*\beta}\right) \tag{B6}$$

Noting that the ratio in the bracket is the formula determining new steady-state unemployment as in (B2), we may rewrite (B6) as

$$\dot{U} = S^{**}\left[1 - \left(\frac{U^{*}}{U^{**}}\right)^{\beta}\right]. \tag{B7}$$

Equation (B7) determines the initial speed of adjustment in unemployment when moving to a new steady state U^{**}. Naturally, the direction of adjustment is positive or negative depending on whether $U^{**} > U^{*}$ or vice versa. The speed of adjustment converges to zero as the difference between U^{*} and U^{**} converges to zero. The speed of adjustment is also proportional to the turnover S^{**}. An important part is played by β. The bigger β, the faster the adjustment. But β also determines steady state unemployment. The bigger β, the smaller the difference between the two steady-states of unemployment and the closer the ratio U^{*}/U^{**} to one.

Appendix C: A Comparison of Coefficients on *U* and *S*

The estimated parameters of our matching function represent elasticities giving the percentage changes in outflows as an outcome of a percentage changes in explanatory variables. Therefore, coefficients on log unemployed, *u*, and log of inflow into unemployment, s, cannot be directly compared to investigate possible difference in probabilities of matching (outflow from unemployment). To see this, note that

$$\beta \equiv \frac{\frac{\Delta O^U}{O}}{\frac{\Delta U}{U}} \Leftrightarrow \Delta O^U = \beta \frac{\Delta U}{U} O;$$

$$\delta \equiv \frac{\frac{\Delta O^U}{O}}{\frac{\Delta S}{S}} \Leftrightarrow \Delta O^S = \delta \frac{\Delta S}{S} O;$$

where superscripts denote the marginal effects due to a change in *U* and *S* respectively. Considering unitary change of *U* and *S*, $\Delta U = \Delta S = 1$, the implied marginal effects of ΔU and ΔS on outflow in geometric means of U and S can be compared as

$$\frac{\Delta O^S}{\Delta O^U} = \frac{\hat{\delta}}{\hat{\beta}} \frac{\overline{U}}{\overline{S}}.$$

However, one additional adjustment is needed. Note that individuals who flow into unemployment in the same calendar month enter the registry on different days within the month. This means that they those who are registered later in given month are subject to lower probabilities of finding vacancies during this month. Assuming that the inflow is spread uniformly over the month, the estimated coefficient on inflow based on monthly frequency has to be multiplied by two to adjust for this.

$$\frac{\Delta O^S}{\Delta O^U} = 2\frac{\hat{\delta}}{\hat{\beta}} \frac{\overline{U}}{\overline{S}}.$$

In other words, if we want to compare marginal quantitative effect of newly unemployed to marginal effect of already unemployed, *β*, we should consider *δ'* rather than *δ*,

$$\hat{\delta}' = 2\hat{\delta}\frac{\overline{U}}{\overline{S}}.$$

Table 1: Unemployment Stocks and Flows (annual means)*

	Rates			Shares of the Labor Force [%]			
	Inflow rate [%]	**Outflow rate [%]**	**U/V ratio**	**Unempl. Rate**	**Inflow**	**Outflow**	**Vacancies**
Czech Republic							
1991	0.80	17.2	3.3	2.70	0.70	0.40	0.80
1992	0.60	24.1	2.4	3.10	0.60	0.70	1.40
1993	0.70	21.0	2.1	2.90	0.70	0.60	1.40
1994	0.60	19.6	2.4	3.20	0.60	0.60	1.40
1995	0.60	19.4	1.7	3.00	0.60	0.60	1.70
1996	0.60	18.1	1.6	3.00	0.60	0.50	1.90
1997	0.80	16.0	2.7	4.10	0.80	0.70	1.50
1998	1.10	14.2	5.5	5.80	1.00	0.80	1.10
1999	1.20	11.8	12.2	8.40	1.10	1.00	0.70
2000	1.20	12.4	10.6	9.00	1.10	1.10	0.90
2001	1.10	12.2	7.5	8.40	1.00	1.00	1.10
2002	1.20	11.0	9.5	9.00	1.10	1.00	1.00
2003	1.20	10.3	12.0	9.90	1.10	1.00	0.80
2004	1.20	10.5	11.7	10.20	1.10	1.10	0.90
2005	1.10	11.1	9.1	9.90	1.00	1.10	1.10
Slovak Republic							
1991	1.30	5.4	23.0	6.60	1.20	0.30	0.30
1992	1.10	10.1	23.8	11.30	1.00	1.10	0.50
1993	1.50	8.1	30.0	12.50	1.40	1.00	0.40
1994	1.30	7.4	35.8	14.40	1.10	1.10	0.40
1995	1.40	9.4	23.4	13.80	1.20	1.30	0.60
1996	1.20	10.4	19.1	10.70	1.00	1.10	0.60
1997	1.40	9.7	14.1	12.70	1.30	1.20	0.90
1998	1.60	7.8	22.5	14.20	1.40	1.10	0.70
1999	1.80	6.2	53.2	18.10	1.50	1.10	0.40
2000	1.70	7.5	73.0	19.60	1.40	1.50	0.30
2001	1.90	7.2	57.8	19.50	1.50	1.40	0.40
2002	1.70	7.5	36.5	19.40	1.40	1.50	0.60
2003	1.50	8.4	22.5	16.90	1.30	1.40	0.80
2004	1.40	8.7	29.1	15.60	1.10	1.40	0.50
2005	1.10	9.4	25.6	14.10	0.90	1.30	0.60
East Germany							
1991	1.70	9.9	30.5	11.50	1.50	1.10	0.40
1992	2.00	10.7	37.1	14.80	1.70	1.60	0.40
1993	1.90	10.1	33.4	14.80	1.60	1.50	0.40
1994	1.90	12.3	25.6	14.90	1.60	1.80	0.60
1995	2.30	14.0	19.3	13.50	2.00	1.90	0.70
1996	2.70	14.8	20.8	15.00	2.30	2.20	0.70
1997	3.00	12.9	24.0	17.20	2.50	2.20	0.70
1998	3.00	15.0	17.4	17.60	2.50	2.60	1.10
1999	3.10	14.5	18.5	17.10	2.50	2.50	0.90
2000	2.90	13.7	21.1	17.30	2.40	2.40	0.80
2001	2.80	13.2	20.1	17.50	2.30	2.30	0.90
2002	2.80	12.9	19.6	17.70	2.30	2.30	0.90
2003	2.90	12.8	23.2	18.30	2.30	2.40	0.80
2004	3.10	13.7	27.6	18.10	2.50	2.50	0.70
2005	2.70	13.3	21.4	18.60	2.20	2.40	1.00

Table 1: continued

	Rates			Shares on the LF			
	Inflow rate	Outflow rate	U/V	Unempl. rate	Inflow	Outflow	Vacancies
West Germany							
1991	0.90	18.5	4.9	5.00	0.90	0.90	1.00
1992	1.00	17.3	5.4	5.30	1.00	0.90	1.00
1993	1.20	15.1	9.1	6.70	1.10	1.00	0.70
1994	1.20	14.7	10.8	7.60	1.10	1.10	0.70
1995	1.30	14.8	9.4	7.60	1.20	1.10	0.80
1996	1.40	14.1	10.1	8.20	1.20	1.20	0.80
1997	1.40	13.3	10.5	9.00	1.20	1.20	0.90
1998	1.40	14.7	8.4	8.60	1.20	1.30	1.10
1999	1.30	15.2	7.0	8.20	1.20	1.20	1.20
2000	1.20	15.9	5.4	7.50	1.10	1.20	1.40
2001	1.30	15.5	5.4	7.30	1.20	1.10	1.40
2002	1.40	15.3	6.7	7.80	1.30	1.20	1.20
2003	1.50	15.2	9.2	8.60	1.30	1.30	0.90
2004	1.60	16.3	10.8	8.70	1.50	1.40	0.80
2005	1.60	13.6	10.2	10.00	1.40	1.40	1.00
Hungary							
1991	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.
1992	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.
1993	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.
1994	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.
1995	0.70	9.9	17.1	6.10	0.60	0.60	0.40
1996	1.40	10.8	14.0	11.40	1.20	1.20	0.90
1997	1.40	12.2	11.3	10.70	1.30	1.30	1.00
1998	1.40	14.1	9.2	9.80	1.30	1.40	1.10
1999	1.40	14.0	8.4	9.30	1.30	1.30	1.10
2000	1.40	14.5	8.0	9.00	1.20	1.30	1.10
2001	1.50	16.3	8.2	8.40	1.30	1.40	1.00
2002	1.40	16.2	7.8	7.90	1.30	1.30	1.00
2003	1.40	15.0	8.0	8.10	1.20	1.20	1.10
2004	1.40	14.6	7.7	8.50	1.30	1.20	1.10
2005	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.
Poland							
1991	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.
1992	0.88	4.5	69.7	13.12	0.76	0.59	0.19
1993	1.13	4.9	80.3	15.24	0.93	0.72	0.19
1994	1.27	6.8	89.5	16.26	1.03	1.06	0.20
1995	1.30	7.9	73.8	15.17	1.11	1.20	0.22
1996	1.21	8.3	84.5	14.15	1.04	1.16	0.18
1997	1.08	10.5	86.5	11.57	0.96	1.21	0.15
1998	1.10	10.1	110.5	9.82	0.99	0.99	0.09
1999	1.34	8.2	169.9	12.32	1.20	0.98	0.07
2000	1.39	7.1	220.0	14.37	1.19	1.02	0.07
2001	1.43	6.0	311.1	16.59	1.19	0.99	0.06
2002	1.50	6.5	296.3	18.18	1.22	1.18	0.07
2003	1.55	7.0	207.4	18.34	1.27	1.29	0.10
2004	1.57	7.7	154.8	17.90	1.29	1.37	0.13
2005	1.58	8.8	112.2	16.86	1.32	1.48	0.17

*Annual rates and shares are computed from averages of monthly values of country-wide aggregates.

Table 2: Persistence of Regional Differentials - Pearson's Rank Correlations

Base Year 1992							Base Year 1999						
Year	Czech Republic	Slovak Republic	East Germany	West Germany	Hungary	Poland*	Year	Czech Republic	Slovak Republic	East Germany	West Germany	Hungary	Poland*
						Unempl. Rate							
1996	0.81	0.82	0.54	0.94	n.a.	0.90	**2002**	0.95	0.95	0.87	0.96	0.95	0.86
1999	0.73	n.a.	0.32	0.94	n.a.	0.83	**2005**	0.91	0.92	0.71	0.89	0.92	0.83
						Inflow Rate							
1996	0.79	0.75	0.60	0.89	n.a.	0.90	**2002**	0.94	0.93	0.87	0.97	0.96	0.94
1999	0.70	n.a.	0.47	0.87	n.a.	0.87	**2005**	0.90	0.82	0.51	0.80	0.95	0.92
						Outflow Rate							
1996	0.50	0.57	0.43	0.84	n.a.	0.48	**2002**	0.86	0.88	0.68	0.87	0.86	0.45
1999	0.38	n.a.	0.46	0.84	n.a.	0.47	**2005**	0.83	0.86	0.54	0.78	0.66	0.72
						Vacancy Rate							
1996	0.33	0.25	-0.03	0.51	n.a.	0.55	**2002**	0.69	0.35	0.66	0.81	0.85	0.8
1999	0.29	n.a.	0.01	0.51	n.a.	0.46	**2005**	0.40	0.34	0.00	0.38	0.50	0.6
						U/V ratio							
1996	0.81	0.82	0.54	0.94	n.a.	0.90	**2002**	0.95	0.95	0.87	0.96	0.95	0.86
1999	0.73	n.a.	0.32	0.94	n.a.	0.83	**2005**	0.91	0.92	0.71	0.89	0.92	0.83

Note: Statistics for some countries/years are missing because:
SR changed district structures in 1997
HU data before 1995 are not available
PL changed district structures in 1999
* Hungarian statistics for 2005 correspond to 2004 values.
* Polish statistics for 1999 correspond to 1998 values.

Table 3: Matching Function Parameter Estimates for West Germany During 1997-2005

	Trend	Std.Err.	β	Std.Err.	γ	Std.Err.	δ	Std.Err.	RTS	p-value**	adjR^2	Nobs
Panel A: Cross-sectional estimators												
OLS	0.0122	0.001	0.680	0.003	0.148	0.003	-	-	0.828	0.000	0.853	14734
OLS (Month Dummies)	0.0110	0.001	0.687	0.003	0.129	0.002	-	-	0.817	0.000	0.896	14734
OLS (Size Adjusted)	0.0101	0.001	0.547	0.028	0.029	0.021	-	-	0.576	0.000	0.616	14734
Panel B: Panel data estimators												
Random Coeficients	0.0096	0.000	0.736	0.009	0.075	0.004	-	-	0.811	0.000	0.645	14734
Fixed Effects	0.0097	0.000	0.742	0.011	0.072	0.004	-	-	0.815	0.000	0.664	14734
OLS on 1st Differences	0.0125	0.003	1.643	0.064	0.067	0.013	-	-	1.710	0.000	0.638	14734
Forward Mean Deviations + IV	0.0125	0.003	1.247	0.064	0.124	0.027	-	-	1.466	0.000	0.585	14595
Panel C: Panel data estimators (preferred estimation methods)												
1st Differences + IV (a)	0.0143	0.002	1.315	0.042	0.140	0.031	-	-	1.455	0.000	0.632	14734
1st Differences + IV (b)	0.0123	0.002	1.270	0.038	0.165	0.032	0.121	0.014	1.556	0.000	0.636	14734
1st Differences + IV (c)*	0.0091	0.002	1.280	0.044	0.125	0.027	0.145	0.014	1.550	0.000	0.631	14734

(a) Specification excluding log inflow (s_t) as explanatory variable.

(b) Specification including log inflow (s_t) as explanatory variable.

(c) Specification including lagged log outflow (o_{t-1}) as explanatory variable.

* Estimated coefficient on lagged outflow: Á=.200 (std.error 0.033) from partial adjustment specification o_{it} = Áo_{it-1} + ¯u_{it-1} +°v_{it-1} +±s_{it} + ε_{it}

** p-value for a test of Ho: ¯+°+± = 1 [constant returns to scale]

Table 4: Matching Function Parameter Estimates[1]

Panel A: 1994-1996	Trend	β	γ	δ	δ'/¯	U/S	η_s	η_v	adjR^2	RTS_M [2,3)]	RTS_U [4,5)]	Nobs
Czech Republic	-0.112	0.749	0.231	0.264	3.7	5.2	0.98	-0.31	0.65	1.24	0.67	2661
	(0.027)	*(0.165)*	*(0.110)*	*(0.032)*	*[0.000]*		*[0.000]*	*[0.030]*		*[0.305]*	*[0.010]*	
Slovak Republic	0.045	0.948	0.171	0.174	4.3	11.8	0.87	-0.18	0.31	1.29	0.69	1292
	(0.058)	*(0.583)*	*(0.159)*	*(0.048)*	*[0.000]*		*[0.137]*	*[0.465]*		*[0.598]*	*[0.093]*	
East Germany	0.045	0.915	-0.082	0.264	4.2	7.3	0.80	0.09	0.48	1.10	0.89	1211
	(0.026)	*(0.452)*	*(0.099)*	*(0.065)*	*[0.003]*		*[0.068]*	*[0.495]*		*[0.853]*	*[0.090]*	
West Germany	-0.103	1.272	0.220	0.203	2.1	6.6	0.63	-0.17	0.67	1.69	0.45	5004
	(0.005)	*(0.072)*	*(0.041)*	*(0.022)*	*[0.000]*		*[0.000]*	*[0.000]*		*[0.000]*	*[0.00]*	
Hungary	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.
	n.a.	n.a.	n.a.	n.a.	n.a.		n.a.	n.a.		n.a.	n.a.	
Poland	0.285	2.596	0.163	0.190	2.0	13.7	0.31	-0.06	0.71	2.948	0.25	637
	(0.097)	*(0.770)*	*(0.120)*	*(0.052)*	*[0.095]*		*[0.002]*	*[0.240]*		*[0.014]*	*[0.004]*	
Panel B: 1997-2005	**Trend**	**β**	**γ**	**δ**	**δ'/¯**	**U/S**	η_s	η_v	**adjR^2**	**RTS_M [2,3)]**	**RTS_U [4,5)]**	**Nobs**
Czech Republic	-0.039	1.158	0.510	0.192	2.7	8.0	0.70	-0.44	0.74	1.86	0.26	7770
	(0.008)	*(0.067)*	*(0.062)*	*(0.017)*	*[0.000]*		*[0.000]*	*[0.000]*		*[0.000]*	*[0.000]*	
Slovak Republic	0.004	1.513	0.238	0.066	1.2	13.2	0.62	-0.16	0.49	1.82	0.46	6682
	(0.010)	*(0.144)*	*(0.053)*	*(0.014)*	*[0.490]*		*[0.000]*	*[0.000]*		*[0.000]*	*[0.000]*	
East Germany	-0.021	1.491	0.340	0.305	3.0	7.4	0.47	-0.23	0.68	2.14	0.24	3602
	(0.005)	*(0.115)*	*(0.044)*	*(0.024)*	*[0.000]*		*[0.000]*	*[0.000]*		*[0.000]*	*[0.000]*	
West Germany	0.012	1.270	0.165	0.121	1.3	6.6	0.69	-0.13	0.64	1.56	0.56	14734
	(0.002)	*(0.038)*	*(0.032)*	*(0.014)*	*[0.063]*		*[0.000]*	*[0.000]*		*[0.000]*	*[0.000]*	
Hungary	0.016	1.554	0.512	0.336	3.0	7.0	0.43	-0.33	0.28	2.40	0.10	1920
	(0.016)	*(0.255)*	*(0.115)*	*(0.057)*	*[0.001]*		*[0.000]*	*[0.004]*		*[0.000]*	*[0.285]*	
Poland	0.022	0.756	0.084	0.188	6.9	13.8	1.07	-0.11	0.67	1.03	0.96	1072
	(0.011)	*(0.107)*	*(0.056)*	*(0.033)*	*[0.000]*		*[0.000]*	*[0.164]*		*[0.830]*	*[0.000]*	

Standard Errors in parenthesis ( ), p-values in parenthesis [ ]

1) Estimates are based on instrumental variables estimation of o_{it} = ¯u_{it-1} +°v_{it-1} +±s_{it} + $\varepsilon_{it,}$ transformed into first-differences.

2) Returns to scale of the matching function: RTS_M = ¯+°+±

3) p-value from the test of Ho: $RTS_M = 1$

4) Returns to scale of the steady-state unemployment: RTS_U = (1-°-±)/¯ = 1+(1-RTS_M)/¯

5) p-value from the test of Ho: $RTS_U = 0$

Estimated parameters:

Trend : estimated intercept coefficient represents annual change in matching function intercept.

β, °, δ: matching function slope coeficients on unemployment, vacancies, and inflow.

δ'/β: δ' is the coefficient on inflow, δ, adjusted to be comparable with the coefficient on unemployment, β. Corresponding p-value comes from the test of Ho: δ'/β = 1.

U/S: average unemployment-inflow ratio.

η_s, η_v : elasticities of steady-state unemployment with respect to inflow and vacancies.

Figure 1: Estimated Trends in Intercept and Returns to Scale (24-month wide data moving window estimates)