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

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1. Introduction

Almost two decades after the fall of the Berlin Wall, unemployment is a major problem in the post communist economies of the former Soviet bloc and Yugoslavia, including the new members of the European Union (EU). In policy discussions, three hypotheses have remerged as leading explanations for this phenomenon, namely that high unemployment is the result of (a) ongoing (unfinished) transition from plan to market in the presence of globalization, (b) macroeconomic policies or major external shocks, (c) problems related to the economic structures of these countries.[1] The discussion complements that in Western Europe, and the nature of appropriate policies for alleviating 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. In particular, we use 1991-2005 *district-level panel data* on the unemployed *U*, vacancies *V*, inflow *S* into unemployment, and outflow *O* from unemployment in five former communist economies that are now members of the EU (the Czech Republic, Hungary, Poland, Slovakia, and eastern part of Germany) and in the western part of Germany (a benchmark western economy) to examine the three hypotheses in the context of inflows into unemployment and the efficiency of matching of the unemployed and vacancies.[2] Hypothesis 1, namely that high unemployment

[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.

[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

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 that *U* is high because *S* is high 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 but rather restructuring brought about by globalization and other forces.

Hypothesis 2 is 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* and *U*, and the policy implication would be that macroeconomic policies are key.

Hypothesis 3 is that high unemployment is caused by inefficient matching, brought about for example by inadequate labor market institutions or geographical or skill mismatch (see also Jurajda and Terrell, 2006). Here we would observe both *U* and *V* being simultaneously high, but not necessarily in the same districts or skill groups. In this case, the policy should focus on labor market institutions and measures to stimulate labor mobility and create appropriate skills.

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 (commonly use 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

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 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.

markedly in terms of their unemployment, flows and vacancy levels and rates.[3] West Germany is in the intermediate range, displaying between 1991 and 2005 an unemployment rate that increases from 5% to 10%, inflow rate that rises from 0.9% to 1.6%, outflow rate that declines from 17-24% to 10-11%, and a vacancy rate (as a share of the labor force) that varies between 0.7% and 1.4%. The changes in these variables occur mostly in two waves, reflecting the business cycles and possibly also a shift toward a service economy with higher natural labor turnover. East Germany, in contrast, registers an unemployment rate rising from 11.5% to 18.6%, inflow as a share of the labor force rate almost doubling from already high level of 1.5% to 2.5 %, outflow as a share of the labor force rising to about 2-2.3% by the mid 1990s and fluctuating around this level ever since, and a vacancy rate rising from 0.4% in 1991 to about 1% 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. Note however, that East Germany has lower outflow rates relative to the number of unemployed and a similar vacancy rate as the Western part of Germany. The East German economy therefore operates with a higher unemployment rate in the presence of very sizable active labor market 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% range, accompanied by relatively high inflow (1.2-1.3% of the labor force) and low outflow (1.0-1.4% of the labor force). In most years, they have also had vacancy rates significantly below 1%. The Czech Republic is an intermediate case, with unemployment rising from the low rate of 3-

[3] 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

4% in the early to mid 1990s to 8-10% range since then. Its inflow and outflow as a share of the labor force have both risen from extraordinary low levels of about 0.5-0.7% in the early-to-mid 1990s to a still relatively low level of 1.0-1.1% since then. Its vacancy rate has declined from very high levels 1.4-1.9% in the early-to-mid 1990s to 0.8-1.1% since then. Finally, Hungary has achieved the lowest level of unemployment. After reaching an unemployment rate of about 11% on the mid-to-late 1990s, Hungary has succeeded to lower the rate to around 8% in the mid 2000s, reduced its inflow rate to 1.4%, raised the outflow rate to 14-16% and kept the vacancy rate at 1.0-1.1%. Hungary's success is hence brought about by keeping the outflow rate relatively high and inflow rate relatively low.

In terms of our hypotheses, the raw data in Table 1 suggest that the West German economy displays elements of all three hypotheses. The unemployment rate has risen with increasing inflows (H1) over business cycles (hysteresis), the vacancy rate has declined (H2) and both the unemployment and vacancy rates are relatively high (H3). The East German economy has conformed to H1, having generated high unemployment through relatively but not extremely high outflow rates (much going into training programs) and extremely high inflow rates (much being 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 relative to inflow) throughout the 1990s and 2000s, although they have also experienced rising inflow rates consistently with H1. The Czech economy had virtually no unemployment problem in the early-to-mid 1990s, as inflows and outflows as a share of the labor force were extremely low and vacancies remained high. While this could be interpreted as a sign of slow restructuring, various studies (e.g., Jurajda and Terrell, 2000) suggest that labor mobility from old to new jobs was substantial. 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 H1 and H2. Finally, in part because it has relatively low unemployment, Hungary does not fit clearly into any of the three

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official aggregate statistics published.

hypothesized scenarios. It has an element of all three hypotheses in that its inflow is relatively sizable (H1), the vacancy rate is low relative to inflow (H2) and unemployment and vacancies are relatively high (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.

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

Conceptually, it is useful to start by noting that the unemployment rate as a stock variable is determined by the inflow rate into and outflow rate from unemployment, with the variation in the two flows determining unemployment dynamics. In a steady state, inflow equals outflow so that $S = O$ and the steady state unemployment rate, u, is given by $U_r = U/LF = 1/(1 + O_r/S_r)$, where $S_r = S/E$ is the steady state inflow rate and $O_r = O/U$ is the steady state outflow rate. For the purposes of this paper, we assume that S is given exogenously by job destruction. This provides the conceptual rationale for the focus of our paper since the only other determinant of the steady state unemployment rate is the matching function describing complex pairing of unemployed U and vacancies V,

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

The fact that unemployment has remained stuck at high levels in many transition economies points to the need to understand the properties and shifts of the matching function and take into account possible differences in the rate of inflow S_r that enters the formula for a steady-state unemployment rate. The literature on matching functions usually assumes that outflow from unemployment to employment *O*.

Matching is seen as a search process, with both the unemployed and employers with vacant positions striving to find the best acceptable match, given exogenous factors such as skill and spatial mismatch, as well as costly access to information. Relationship between determinants of matching and matching function parameters on one hand and equilibrium unemployment on the other can be found noting that in a steady state, inflow into unemployment equals outflow and equilibrium unemployment is determined by implicit function

$$S - M(U,V) = 0. \tag{1b}$$

In section 5, presenting our results, and in Appendix B we employ analytical solution for equilibrium unemployment for our estimable functional form of matching function. Some authors (e.g., Blanchard and Diamond, 1989, Pissarides, 1990, and Storer, 1994) expect the matching function *M* to display constant returns to scale, while others have identified reasons such as externalities in the search process, heterogeneity in the unemployed and vacancies and lags between matching and hiring, why increasing returns may prevail (e.g., Diamond, 1982, Coles and Smith, 1994, Profit, 1996, and Mortensen 1997). Increasing returns are conceptually important because in some models they constitute a necessary condition for multiple equilibria and provide a rationale for government intervention. The radical labor market and other reforms introduced in Slovakia in the early 2000s follow this line of reasoning. In this paper we show that rising returns appear to be an important phenomenon in all transition countries, especially in the later (1997-2005) than the earlier (1993-96) period, and that it is more pronounced in some of the economies than others and has an important negative impact on unemployment rate.

In view of the serious 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 (1a) 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 adjusts the data 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.[4] Unlike 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).

3. The Estimating Framework

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[4] 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 Munich (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 (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 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 Munich et al. (1997).

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 which may be written in a deterministic form as[5]

$$\ln O_{i,t} = \beta \ln U_{i,t-1} + \gamma \ln V_{i,t-1} + \alpha \qquad (2)$$

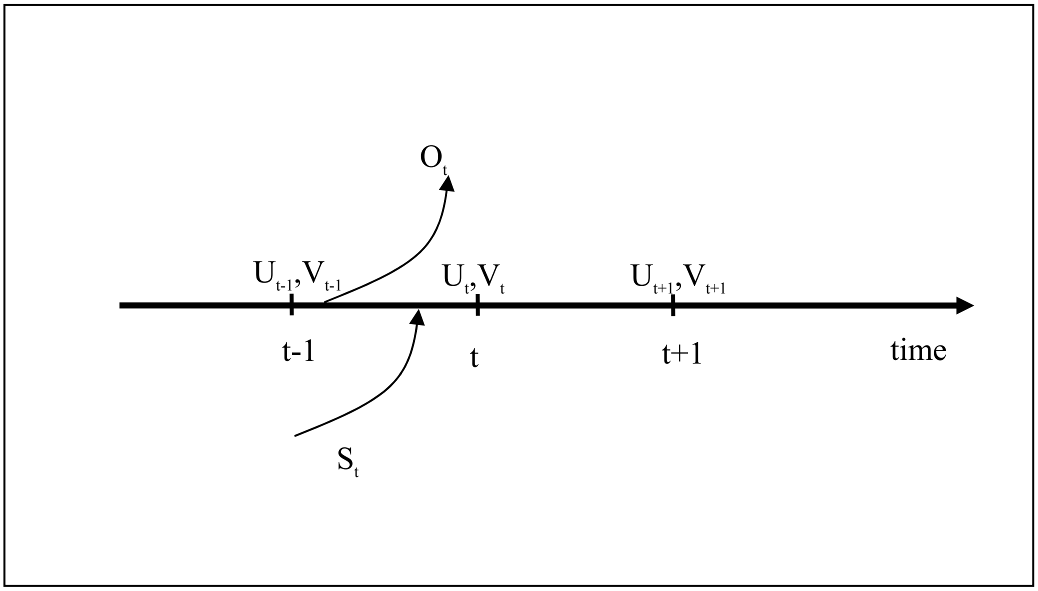
where, $U_{i,t-1}$, and $V_{i,t-1}$ are the number of unemployed and vacancies in district i at the end of period $t-1$, respectively, $O_{i,t}$ denotes the outflow to jobs during period t (the number of successful matches between the currently unemployed and current vacancies) and constant α captures the efficiency of matching. The stock-flow process is depicted on Chart 1. Using lowercase letters for logarithms of variables and introducing unobserved (time invariant) district specific effects α_i and an idiosyncratic error term $\varepsilon_{i,t}$, we can write (2) as

$$o_{i,t} = \beta u_{i,t-1} + \gamma v_{i,t-1} + \alpha_i + \varepsilon_{i,t}\,. \qquad (3)$$

In estimating (3), 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 and v. This correlation is very likely to exist on account of structural differences between districts caused by factors which affect α_i, u and v. One important factor is district size, although this factor is observable and its impact can be alleviated by adjusting all variables in equation (3) by some measure of district size.

Chart 1: Stock-flow process of matching

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



If panel data are available, as in our case, suitable *within transformation*s of (3) 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[6]

$$
\begin{aligned}
&U_t \equiv U_{t-1} + S_t - O_t \\
&U_{t-1} \equiv U_{t-2} + S_{t-1} - O_{t-1} \\
&U_{t-2} \equiv U_{t-3} + S_{t-2} - O_{t-2} \\
&\ldots
\end{aligned}
\tag{4}
$$

Lagged outflows in (4) are in turn given by a lagged version of (3) as

[6] 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 problem of weak exogeneity.

$$\log O_{t-1} = \beta u_{t-2} + \gamma v_{t-2} + \varepsilon_{t-1}$$
$$\log O_{t-2} = \beta u_{t-3} + \gamma v_{t-3} + \varepsilon_{t-2} \quad (5)$$
$$\ldots$$

Since any district mean is computed from all district observations over time, the means contain all values of the error term $\{\varepsilon_t\text{: } t = 1, 2, 3,\ldots,T\}$. 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 $\{\varepsilon_t\text{: } t = T\text{-}1, T\text{-}2\}$, as may be seen by rewriting (5) in a first difference form as

$$\Delta o_t \equiv o_t - o_{t-1} = \beta(u_{t-1} - u_{t-2}) + \gamma(v_{t-1} - v_{t-2}) + \varepsilon_t - \varepsilon_{t-1}, \quad (6)$$

which may in turn be expressed in a simplified notation as

$$\Delta o_t = \beta \Delta u_{t-1} + \gamma \Delta v_{t-1} + \Delta \varepsilon_t \ . \quad (7)$$

From equation (4) it follows that u_{t-1} and u_{t-2} contain ε_{t-1} and ε_{t-2}, respectively, through outflows as in equation (5).[7]

The first difference transformation thus leaves further lags of Δu_t $\{u_t\text{: } t = 2, 3,\ldots, T\text{-}4, T\text{-}3\}$[8] uncorrelated with $\Delta\ \varepsilon_{t-1}$ (i.e., with ε_{t-1} and ε_{t-2}) and these further lags of Δu_t can be used as instrumental variables. Vacancies in (6) are predetermined in the same statistical sense and may be treated the same way. Available instruments are therefore given by $\{\Delta u_t, \Delta v_t\text{: } t = 2, 3, \ldots T\text{-}3, T\text{-}2,\}$.

There are several additional features of the matching function model that need to be taken into account in estimation. First, identities in (4) show that lagged changes in inflows $\{\Delta s_t\text{: } t = 2, 3, \ldots,T\text{-}2, T\text{-}1)$ are available as additional instruments because they codetermine the explanatory variables but do not directly affect outflow.

Second, rather than using changes in lagged values as instruments, we can use lagged levels because

[7] It does not change the essence of the argument that (4) is defined in levels and (5) in logs.
[8] Note that the first observation (for $t = 1$) of the first differenced variables is not available because observations for $t = 0$

differences are simply 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 $\{\varepsilon_t$: $t = 1, 2,\ldots, T)$ may in principle be uncorrelated, their first differences $\Delta\varepsilon_t$ will by definition be autocorrelated. To see this, note that $\Delta\varepsilon_t = \varepsilon_t - \varepsilon_{t-1}$, $\Delta\varepsilon_{t-1} = \varepsilon_{t-1} - \varepsilon_{t-2}$, and both of them contain ε_{t-1}.[9] As a result, to obtain unbiased standard errors for the estimated coefficients, we use a robust variance-covariance matrix.

Fifth, further complications arise if ε_t has an autoregressive form. In this case the problem is that current ε_t and ε_{t-1} (and therefore $\Delta\varepsilon_t$) contain all previous values $\{\varepsilon_t$: $t = 2, 3, \ldots,T-2, T-1)$ 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 (3) represents a full adjustment model, which assumes that outflow reacts immediately and fully to changes in the stocks of unemployed and vacancies. Since in practice there may be frictions, 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:

$$o_{i,t} = \varphi o_{i,t-1} + \beta u_{i,t-1} + \gamma v_{i,t-1} + \alpha_i + \varepsilon_{t,t}\,. \tag{8}$$

In this model, the lagged value of outflow 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 and v. Lagged outflow hence has to be instrumented and one can consider further lags $\{\Delta o_t$: $t = T-2,T-3, \ldots)$ or their levels $\{o_t$: $t = T-3, T-4,\ldots)$ as instruments.

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do not exist.

Finally, Model (3) assumes that propensity to match is the same independent of unemployment duration. Yet, studies such as Coles and Smith (1994) suggest that the propensity to match is higher at the time of entry into unemployment when the newly unemployed search through all existing vacancies. Later on, those who remain unemployed may not search through existing vacancies (those already explored) and may tend search through only the newly posted vacancies.[10] To reflect this so called "stock-flow" matching, we also include inflow into unemployment, s_t, as an additional explanatory variable in (3). 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 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.[11] 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 values 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, the inter-correlations among O_i, U_i, S_i, and V_i tend to be biased upward on account of the variation in the size of districts (see Appendix Table A2 for an illustration). The usual Cobb-Douglas specification based on cross-sectional data then provides biased estimates of coefficients unless there are constant returns to scale or the unadjusted U_i and V_i are uncorrelated with the labor force L_i. The direction of the bias of $\beta(\gamma)$ is negative if U_i (V_i) is positively correlated with L_i and matching displays increasing returns to scale. Either decreasing returns to scale or

[9] Note that $\text{CORR}(\Delta\varepsilon_t, \Delta\varepsilon_{t-1}) = \text{COV}(\varepsilon_t - \varepsilon_{t-1}, \varepsilon_{t-1} - \varepsilon_{t-2})/\text{VAR}(\varepsilon_t - \varepsilon_{t-1})\ \text{VAR}(\varepsilon_{t-1} - \varepsilon_{t-2}) = \ldots = -0.5$, so that $\Delta\varepsilon_t = -.5\Delta\varepsilon_{t-1} + \text{error}_t$.

[10] The newly unemployed may also have not yet experienced depreciations of skills and psychological scarring but this is being reflected by matching function parameters.

[11] There are several possible measures of district size. We use the district labour force, but the results would not be materially

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 differences in the size of the district labor force. As we show in Appendix A, the bias is very similar to that stemming from an omitted variable problem. In what follows, we call this phenomenon the *spurious scale effect*.[12]

It can be shown that the spurious scale effect is avoided if one uses panel data and estimates a Cobb-Douglas function accounting in some way 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.

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[13] - July 2005, 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;

affected by using other measures.

[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.

[13] 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. At this moment, Hungarian data at our disposal start in January 1995.

$U_{i,t}$ = the number of unemployed in district i the end of period t;

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

$V_{i,t}$ = the number of vacancies in district i at the end 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.[14] 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, 2000, 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 already been discussed above. The Pearson's rank correlation coefficients reported in Table 2 provide information on the persistence of regional (inter-district) differentials in the unemployment, inflows, outflows, vacancies, and U/V rates in the early and late transition. In particular, the left hand side panel shows persistence between 1992 and 1996, while the right hand side panel gives persistence between 1999 and 2005.[15] The rank correlation coefficients indicate that in all six economies there is a major decline over time in the persistence of regional differentials in vacancies. Hence, there

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

[15] 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.

have been major inter-regional changes on the demand or job creation side of these economies over time, as captured by the Pearson's correlation coefficient of the vacancy rate. The decline is most pronounced in East Germany, but it is substantial in all the other economies, including West Germany. Moreover, the pattern holds in the 1990s as well as 2000s. Interestingly, there is much more persistence in the ranking of districts with respect to the other potential measure of demand, namely the outflow rate, but this measure is less accurate since it measures total outflows and in some countries it contains significant outflows into training programs. Nevertheless, the rank correlation coefficients in Table 3 indicate that there has been considerable persistence in regional differentials for all the other variables -- unemployment, inflows, outflows, and U/V rates, with East Germany again standing out and showing the most substantial declines in the correlation coefficients for all these variables over time. Hence, both in the benchmark market economy (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), movement out of unemployment, and labor market tightness (*U/V ratio*). Transition alone hence does not seem to be an explanation for the persistence in the relative standing of districts along these dimensions of the labor market.

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 post-unification 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 present results generated by the various techniques used in the literature that may generate biased estimates, and then we present estimates that correct for the aforementioned deficiencies. 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 does not affect the estimates. The OLS estimates of β and γ based on variables adjusted for district size are lower, which is not surprising given that the spurious scale effect is biasing 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 levels, 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 greater than the corresponding OLS estimates, and γ coefficients that are somewhat smaller than the OLS estimates. 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 causes 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.

Our final estimates, in Panel C of Table 3, come from an IV estimation based on first differences of variables. Since these are our preferred estimates, we report them 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.[16] When the newly unemployed are included in the regression (second to last row in Table 3), we find that their coefficient δ is significant at 0.12, indicating that the newly unemployed indeed display a high propensity to match. Indeed, when one converts the estimated elasticities β and δ into the probability that additional unemployed and a newly unemployed flow out at the mean of U and S, one finds that these probabilities are not statistical different (see the appendix for the formula). 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 0.2 and it is highly statistically significant. The matching process hence appears to be better captured as a dynamic phenomenon, but it is notable that the estimated monthly extent of adjustment is estimated to be relatively high at 0.8 = 1.0 - 0.2. 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 early-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

[16] The instruments explain 20% to 75% of variation in 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 decreases significantly the variance of the explanatory variables, while doubling the variance of the error term and potential measurement error, and leads to higher standard errors of estimated parameters. However, since the measurement error causes negative bias, our estimates represent the lower bound of actual coefficients. Measurement error is more likely to be present in vacancies and cause negative bias in the coefficient on vacancies.

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 the inflow into unemployment (newly unemployed) as regressors.

In addition to matching function parameters, we also present elasticities of equilibrium unemployment with respect to inflow and vacancies, ´$_s$ and ´$_v$. As we show in the Appendix 2, for the C-D functional form of the matching function, analytical solution for steady-state unemployment can be found such that

$$\ln U^* = \eta_s \ln S + \gamma_v \ln V \quad - \ln A$$

where

$$\eta_s = \frac{(1-\delta)}{\beta} \quad \text{and} \quad \eta_v = -\frac{\gamma}{\beta}.$$

Note that steady-state unemployment U^* is determined by actual levels of S and V. It is the great advantage of the matching function tool that its parameters can be estimated on disequilibrium data series. During the early period 1994-1996, elasticity of equilibrium unemployment with respect to inflow (´$_s$) in all transition countries except Poland was higher compared to the period of late transition while it stayed almost unchanged in the West Germany. It implies that steady-state unemployment was more sensitive to growing turnover in earlier stages of transition. It should be also noted that inflows (turnover) were relatively higher in the later period in all countries. This implies that these two simultaneous effects – decreasing elasticity of inflow and growing inflow contributed to unemployment in opposite directions, offsetting each other. The elasticity of equilibrium unemployment with respect to vacancies increased (became more negative) in all transition countries except Poland where it stayed statistically insignificant in both periods. 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 is relatively low, compared to transition countries where unemployment became more reactive to vacancies. It should be also noted that the number of vacancies, contrary to the size of inflow, stayed at similar levels in both periods.

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 v. 0.75) and more precisely estimated coefficient on vacancies in West Germany than in the Czech Republic. In both East Germany and Slovakia, 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 units of observation are larger than the districts 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 suggesting that the newly unemployed match better with vacancies than the existing unemployed. In fact, when we transform δ to be comparable to the coefficient β on the number of unemployed (see Appendix C for details), we see in Table 4 that the transformed coefficient δ^* for the newly unemployed is considerably larger at 2.0 to 4.4 than the coefficient (0.75 to 2.60) on existing unemployed. The only exception is Poland 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 countries. With one exception, returns to scale in matching 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 results point to constant returns, but the estimate is very imprecise.

The results in Table 4 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 highly positive externality from workers to firms (increasing search and matching intensity with rising unemployment). However, unemployment was statistically unimportant 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 δ^* of 2.0 to 4.4 in all these economies, thus suggesting that job creation involved in a major way the newly unemployed.

During the late 1990s and early-to-mid 2000s, the efficiency of matching, as captured by returns to scale, rises in all the transition economies except Poland, but declines somewhat in West Germany. 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. 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 out of the labor force. 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 few vacancies and higher inflows in Slovakia than in the Czech Republic.

The value of matching function intercept A represents disembodied efficiency of matching and its comparison across countries would have high policy relevance. Unfortunately, estimates of this parameter are precluded due to under-reported vacancies.[17] Assuming, that the scope of under-reporting is mainly country specific, we can identify country specific trends in matching function estimates. In 1997-2005, the estimated annual trend (capturing the disembodied change in matching efficiency) 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 1990s, while Poland appears to be improving matching efficiency over both periods. As we show in Figure 1, the negative trend in the Czech Republic is becoming less negative over time, while in East Germany the trend actually oscillates around zero with a diminishing amplitude over time. In the West Germany, the trend turned from negative one in late 90s to slightly positive during 1999-2002, most likely linked to temporary economic upswing in this period. More detailed insight on trends is provided on the top panels of Figure 1. Trends presented there are based on repeated estimations on two years of data (24 months), from which we consecutively remove the oldest month and add the newest one. Estimates are presented for three countries which did not change district structure during whole period 1992-2005, Czech Republic, East and West Germany. Thus observing the difference between a market and a transition economy that share the same historical and cultural background, but with the transition

[17] Systematic under-reporting of vacancies is very common. Measurement error biases are to great extent diluted by the within transformation we use which removes systematic component of under-reporting. However, when imputing matching function intercept from equation (2) as ® = *o* - ¯*u* - °*v* - ±*S*, using estimated parameters and actual levels of variables, under-reporting in vacancies biases estimated intercept upward. In other words, larger under-reporting implies higher intercept. Since the scope of

economy being affected by administratively set subsidies and increased wages as a result of unification. Comparison of the Czech Republic and West Germany eliminates the effect of unification but also common history and culture. The dots in the figures are the individual point estimates for each position of the window and the lines provide one standard error confidence interval smoothed bands around these point estimates.

In bottom panels of Figure 1 we present estimates of returns to scale estimated by the same 24-month wide mowing 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, 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), namely that high unemployment implies that restructuring is at work, is consistent with unemployment being high due to relatively high turnover on the labor market due to ongoing restructuring, with the policy implication being that restructuring needs to be completed. 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) is that

under-reporting is not known and differs across countries, one cannot distinguish economic component from the one due to under-reporting.

high unemployment is caused by inefficient matching. In this case, the policy should focus on labor market institutions and measures to stimulate labor mobility and create appropriate skills.

Our results 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 slow trends of rising unemployment and inflow, declining vacancies and relatively efficient matching (high returns to scale). Its outcome is hence most consistent with H1 and H2. 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 matching, it increasingly gives support to H1 and H2. East Germany's results are also in line with H1 and H2, 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 unemployment and inflows, low vacancies and outflows and increasing returns to scale as well as a positive trend in efficiency. It is therefore also consistent with H1 and H2. 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 H3 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 consistently 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 early transitional period in all transitional were notably less persistent.

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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

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

where k_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 (1) into (2), we get

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

Estimation of (A2) is identical to estimation of (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 the problem is to adjust the variables by the district size in order to obtain the following model:

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

which can 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

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. Steady state unemployment, for given level of exogenous inflow and vacancies, is implicitly defined by matching function as

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

implying steady state unemployment

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

Time subscripts are suppressed since steady-state values do not change over time. The function determining steady-state unemployment is homogenous of degree $k^{(\beta+\gamma+\delta-1)}$. It implies that $k=1$ if matching in (B1) exhibits constant returns to scale, i.e. $\beta + \gamma + \delta = 1$.

Note that the matching function does not impose any particular constraint on the relationship between the inflow and number of vacancies. A proportional increase in both *S* and *V* is specific type of shift among continuum of other simultaneous shifts possible. We do not need to specify inflow-vacancy relationship as long as we consider shifts in *S* and *V* 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) can 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. Clearly, the number of vacancies needed to compensate for growing turnover (to secure unchanged steady unemployment) is lower when matching of newly unemployed is better (higher δ) and matching of vacancies is better (higher γ). For maximum δ (= 1), no additional vacancies are needed to keep steady-state unemployment unchanged.

Equation (B2) can be rewritten as

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

and expressed in logs

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

it reveals determination of steady state unemployment by inflow, vacancies, and matching function parameters. More conveniently, the impact of inflow and vacancies can be expressed in terms of elasticities as

$$\eta_s = \frac{(1-\delta)}{\beta} \quad \text{and} \quad \eta_v = -\frac{\gamma}{\beta}. \tag{B4}$$

Theory assumes that $\beta \geq 0$, $\gamma \geq 0$, and $0 \leq \delta \leq 1$ implying that $\eta_s \geq 0$ and $\eta_v \leq 0$. In economic terms, steady-state unemployment increases with the turnover but the impact of turnover is lower when matching of newly unemployed is higher (captured by δ). For maximum value $\delta=1$, turnover has no impact on steady steady-state unemployment. Similarly, vacancies have negative impact on steady-state unemployment and the impact increases with parameter γ.

Dynamics – transitions between steady-state

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

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

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

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

This formula determines the initial speed of adjustment in unemployment when moving to new steady state U^{**}. Naturally, the direction of adjustment is positive or negative depending on whether $U^{**} > U^*$ or otherwise. 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^{**}. Important role is played by β. The greater the β, the faster the adjustment is. But note that β is also determining steady state unemployment. The greater β, the smaller the difference between two steady-states of unemployment and the closer is the ratio U^*/U^{**} to one.

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

Estimated parameters of our matching function represent elasticities. As such they represent percentage changes in outflows as an outcome of a percentage changes in explanatory variables. Therefore, coefficients on unemployed, *U*, and inflow of unemployed, *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 superscript denotes marginal effect due to a change in *U* and *S* respectively. Considering unitary change of *U* and *S*, *¢U = ¢S = 1*, marginal effects of *U* and *S* in their means 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 to obtain comparable effects. Individuals who flow into unemployment in the same calendar month enter on different days within the month. This means that they face different probabilities of finding vacancies during the month of entry. Assuming, that the inflow is spread approximately uniformly over the month, the estimated coefficient on inflow 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 economic (quantitative) marginal effect of inflow with the impact of unemployment, we should consider δ' rather than δ,

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