THE WILLIAM DAVIDSON INSTITUTE
AT THE UNIVERSITY OF MICHIGAN BUSINESS SCHOOL

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Working Paper Number 107
January 1998

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THE WORKER-FIRM MATCHING IN TRANSITION ECONOMIES: (WHY) ARE THE CZECHS MORE SUCCESSFUL THAN OTHERS?[1]

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[1] In preparing the paper, the authors were in part supported by a grant from the National Science Foundation Grant No. SBR-9512001). Münich and Svejnar's research was also supported by an institutional grant from the Grant Agency of the Czech Republic (GACR No. 402/95/0789). Münich and Terrell's research was in part supported by the National Council for Soviet and East European Research (Contract No. 812-32). Finally, Münich's research was in part supported by the Research Support Scheme (RSS/HESP No.:865/1995). The authors would like to thank Miroslav Lizal and participants of the November 1994 OECD/IHS Technical Workshop on "Regional Unemployment in Central and Eastern Europe", and October 1997 Wiliam Davidson Institute conference on Labor Markets in Transition Economies for valuable comments. The paper also greatly benefitted from comments by Elie Applebaum, Robert Dunn, Jan Kmenta, and Dale Mortensen. The usual disclaimer applies.

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Abstract

In this paper we compare the nature and determinants of outflows from unemployment, taking as a case the Czech and Slovak Republics which in the early 1990's experienced a process close to a controlled experiment. Overall, our study suggests that the exceptionally low unemployment rate in the Czech Republic as compared to that in Slovakia and the other Central and East European economies has been brought about principally by (1) a rapid increase in vacancies along with unemployment, resulting in a balanced unemployment-vacancy situation at the aggregate as well as district levels; (2) a major part played by vacancies and the newly unemployed in the outflow from unemployment; (3) a matching process with strongly increasing returns to scale throughout (rather than only in parts of) the transition period; and (4) the ability to keep the long term unemployed at relatively low levels.

Using the framework of matching functions, we find that in many years the usual Cobb-Douglas specification and the hypothesis of constant returns to scale are rejected. A translog matching function with weak separability between the existing and newly unemployed is found to be the functional form best supported by the data. Our theoretical analysis also indicates that by not adjusting data for the varying size of districts or regions, previous studies may have generated estimates of the returns to scale of the matching function that were biased toward unity.

Non-technical Summary:

The Czech-Slovak comparison is useful for understanding the difference between the Czech experience and that of other economies in the region. This is because during the transition the Czech Republic experienced very low unemployment while the Slovak labor market indicators have resembled those of Poland, Hungary and other Central and East European economies with two-digit unemployment rates.

The differences in unemployment rates are in large part brought about by a high outflow rate from unemployment. Our estimates indicate that the process of outflow from unemployment in the Czech Republic was relatively stable between 1992 and 1995, while in Slovakia the process developed from having virtually insignificant parameters in 1992 to being driven by the number of unemployed in 1994. An important part of our explanation of the Czech-Slovak differences hence lies in much better matching of unemployed individuals and vacant jobs offered by employers in the Czech Republic than in Slovakia in the crucial early period of the transition.

JEL Classification: E24, J64

Keywords: unemployment flows, matching functions, returns to scale, multiple-employment equilibria, Czech Republic, Slovak Republic.

1. Introduction

A fundamental systemic feature of the Soviet-type, centrally planned economies was the absolute nonexistence of open unemployment. An equally distinguishing feature of the transition to a market economy has been the emergence of double digit unemployment rates in all the rapidly transforming economies except for the Czech Republic, where the rate remained at mere 3-4 percent. The gravity of the unemployment problem and the discrepancy between the unemployment experience of the Czech Republic and the other Central and East European (CEE) economies, depicted in Figure 1, poses a fundamental academic as well as policy-related puzzle. Why has the Czech Republic been exhibiting so much lower unemployment rate than its traditional counterpart republic, Slovakia, and the other CEE economies? Until 1993, Slovakia and the Czech Republic have after all shared the same currency, legal system and institutional framework. Moreover, except for unemployment, aggregate economic indicators of the two countries were very similar in the early 1990s (Dyba and Svejnar, 1994, 1995). Why then have the Slovak labor market indicators been similar to those of Poland, Hungary and other transition economies, rather than those of the Czech Republic? What are the policy implications?

The simple answer to the above questions is that the Czech Republic has a dramatically higher outflow rate of individuals from the unemployment state than the other CEE economies (see e.g., Boeri, 1994 and Boeri and Scarpetta, 1995).[2] This basic finding has focused the attention of researchers on the determinants of outflow from unemployment and the matching of the unemployed and vacancies in the CEE countries.

[2] The Czech Republic has had a similar or only somewhat lower rate of inflow of the employed into unemployment than the other transition economies. Burda (1994) for instance reports 1992 monthly inflow into unemployment (as well as outflow from unemployment to job) rates, computed as a percentage of the population of origin, to be .5 (17.4) Czech Republic, .8 (10.4) Slovak Republic, .5 (5.2) Hungary and .6 (4.4) Poland.

The western literature has usually approached these issues in the context of matching functions, assuming that the outflow (number of individuals flowing) from unemployment to employment O is a function of the number of unemployed U and the number of posted vacancies V, $O=f(U,V)$. The process of matching is seen as a search process, with both the unemployed and employers with vacant positions striving to find the best match, given exogenous factors such as skill and spatial mismatch, as well as availability of information. Some authors (e.g., Blanchard and Diamond, 1989, Pissarides, 1990, and Storer, 1994) expect the matching function f 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 (see e.g., Diamond, 1982, Coles and Smith, 1994, Profit, 1996, and Mortensen 1997). Increasing returns are conceptually important because they constitute a necessary condition for multiple equilibria and a rationale for government intervention. In this paper we show that this may be an important phenomenon in the Czech Republic.

In specifying the matching process, the most frequently used functional form is Cobb-Douglas

$$\log O_{i,t} = c + \beta_U \log U_{i,t-1} + \beta_V \log V_{i,t-1} \tag{1}$$

where, $U_{i,t-1}$, and $V_{i,t-1}$ represent the number of unemployed and vacancies in area i at the end of period $t-1$, respectively, $O_{i,t}$ denotes the outflow to jobs in area i during period t (the number of successful matches between the currently unemployed and current vacancies) and constant c captures the efficiency of matching.

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 very 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 production 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. 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.[3]

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[3]In the first study in this area, Burda (1993) uses monthly Czech and Slovak district-level data from October 1990 to May 1992 (with considerable gaps for Slovakia) to estimate simple static matching functions, regressing the logarithm of monthly gross exits from unemployment into employment on previous month's unemployment and vacancies. His OLS estimates of Cobb-Douglas function parameters using pooled sample indicate that the coefficient on unemployment is about twice as high as that on vacancies and that the matching function displays constant or decreasing returns to scale. In Slovakia, the coefficient on vacancies is small and statistically insignificant in the cross-sectional estimates.

Boeri (1994) uses regional panel data for varying periods for the Czech Republic (1991:3-1993:5), Hungary (1991:10-1993:3), Poland (1992:9-1993:3), and Slovakia (1992:5-1993:6) to estimate a Cobb-Douglas matching function in vacancies and unemployment, with unemployment entered as a CES function of short and long term unemployment. His pooled OLS estimates also suggest that vacancies have a relatively small effect on outflow to jobs and that the impact of long term unemployed is significantly lower than the impact of short term unemployed.

Svejnar, Terrell and Münich (1995) estimate augmented matching functions in order to assess whether factors other than unemployment and vacancies systematically affect outflow. The authors use annual 1992 and 1993 data from the Czech and Slovak Republics and regress for each annual cross section the logarithm of the average monthly outflows in each district on the logarithm of district-level unemployed and number of vacancies, demographic characteristics of the district, district demand variables, structural variables and the level of per capita expenditures on active labor market policies. In their seemingly unrelated regressions (SUR) across the two years, they find the coefficient on unemployment to be about 0.8 in the Czech Republic and about 0.4-0.6 in Slovakia, while the coefficient on vacancies is about 0.14-0.17 in Slovakia but insignificant in the Czech lands. In assessing the impact of active labor market policies (ALMPs), such as job subsidies, public jobs creation and training of the unemployed, the authors find that a 1% increase in per capita expenditures on ALMPs increases outflows by 0.17% in the Czech Republic but has no statistically significant effect on outflows in Slovakia in these two years. In addition, the analysis points to demand conditions (captured by variation in industrial production rather than just by the number of vacancies) as important determinants of the larger Czech outflow rate. In particular, decreases in industrial output in the district are found to bring about a larger decrease in the outflow rate in the Slovak Republic than in the Czech Republic. Tests of equality of coefficients across the two years lead to a rejection of the null hypothesis, indicating that the transition indeed changes the regime and that it is thus important to allow for changes in the structure of the underlying model over time.

Svejnar, Terrell and Münich (1994) re-estimate the aforementioned augmented matching functions in a dynamic distributed lag model as well as in static first difference framework. They find general support for their earlier findings, although some of the results are sensitive to model specification.

Lubyova and van Ours (1994) estimate matching functions on monthly data for the period 1990:10-1993:12

The contradictory findings of the existing studies and the Czech Republic's high outflow from unemployment relative to Slovakia and the other transition economies have led us to carry out an in-depth comparative analysis of matching in the Czech and Slovak republics. Our focus on the Czech and Slovak Republics is motivated by the fact that they represent a unique laboratory for analysis, in which we can hold constant many initial conditions. Unlike other studies, we also use a more flexible and up-to-date empirical methodology and superior

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in Slovakia and 1991:1-1993:12 in the Czech Republic. In order to allow for matching across districts, they estimate the matching functions on regional as opposed to the more disaggregated district-level data. The authors estimate the matching function separately for each region and search for structural breaks in order to allow for uneven transition process across regions. The estimated coefficients point to strongly increasing returns to scale in all eight regions of the Czech Republic and two of the four Slovak regions. The structural shift coefficients suggest that there was a negative efficiency shift in two regions of the Czech Republic, a positive shift in one Czech and two Slovak regions, and no shift in five Czech and two Slovak regions.

Boeri and Scarpetta (1995), use relatively long panels of monthly data for districts/regions in Poland (1992:1-1993:12), Hungary (1991:1-1994:4), the Czech Republic(1991:1-1994:4), and Slovakia (1990:12-1993:12) to estimate Cobb-Douglas matching functions augmented by variables proxying for the prevalence of agricultural employment and diversification of economic activity. They find that the coefficient on the log of unemployment ranges from 0.52 in Hungary to 0.78 in the Czech Republic, while the coefficient on the log of vacancies ranges from 0.05 in Hungary to 0.28 in Slovakia. Except for Slovakia, one can always reject the hypothesis of constant returns to scale in favor of decreasing returns.

In recent studies, Burda and Lubyova (1995) and Boeri and Burda (1995) use district-level data to estimate augmented matching functions with the aim of quantifying the effect of active labor market policies. Burda and Lubyova (1995) use a combination of monthly and quarterly Czech and Slovak data from 1992:1 to 1993:12 in augmented regressions and 1994:7 in the non-augmented runs. Boeri and Burda (1995) focus on the issue of endogeneity of ALMP measures and use Czech quarterly data for over the period 1992:I-1993:IV in the instrumental variable estimation and 1992:I-1994:II in the OLS runs. Both studies control for district-specific fixed effects by estimating in differences from district means and including time dummies and allowing for dynamics through a partial adjustment model. Both studies find expenditures on active labor market policies to have a significant positive effect on outflows to jobs under all specification (including in Slovakia in the Burda and Lubyova (1995) study). The estimates again display high coefficients on unemployment and low and in Slovakia often insignificant ones on vacancies.

Burda and Profit (1996) note that there is wide dispersion in unemployment rates across regions of the CEE countries and that standard matching functions, estimated with district level panel data, generate different coefficients across districts and regions. They present a model of nonsequential search with endogenous search intensity and show that it can provide a link between the spacial instability of the matching functions and spatial interdependence in matching. The model also induces more complex functional forms and non-constant returns to scale in matching. In their empirical investigation, the authors use district and regional data from the Czech Republic during the period January 1992 - July 1994. They estimate augmented OLS Cobb-Douglas functions with lags in the dependent variable and various proxies for regional interactions in matching. Burda and Profit find a statistically significant, nonuniform impact of surrounding districts on local matching within a district. However, they cannot reject the hypothesis of constant returns in matching.

In a parallel study to the present one, Profit (1996) notes that if matching does not lead to instantaneous hiring, static matching functions with autoregressive fixed effects will yield biased estimates. He estimates a Cobb-Douglas matching functions using panel data from 76 Czech districts during the period January 1992 - June 1994. In order to correct for misspecification, Profit focuses on issues of autocorrelation, heteroscedasticity and validity of instruments and shows that constant returns to scale that are produced by simple models turn into increasing returns in more sophisticated Anderson-Hsiao Instrumental Variable (AHIV) and GMM models.

data. In particular, we a) use a translog rather than the more restrictive Cobb-Douglas specification of the matching function, b) separate the effects of new and longer term unemployed, c) estimate the effects of several potentially factors on outflow from unemployment, d) employ a dynamic specification and estimate on contiguous panels to allow for dynamic adjustment and regime changes during the period of economic transition, e) test and carefully control for the endogeneity of explanatory variables, f) control for heterogeneity of the unemployed searchers, and g) account for the presence of a spurious scale effect introduced by the varying size of units of observation (districts). Our results provide an answer to the fundamental unemployment puzzle in the CEE economies and important new evidence on the theoretically controversial issue of returns to scale in the matching function.

We begin in Section 2 by presenting our estimating framework and explaining how it overcomes some of the principal problems of the existing studies. In Section 3 we describe our data and the implementation of our econometric model. In Section 4 we present basic statistics and our econometric estimates. We conclude in Section 5.

2. The Estimating Framework

Theories of search and matching do not imply a particular functional form of the matching function and most studies use the simple Cobb-Douglas form.[4] As mentioned above, we minimize the potential bias stemming from this restrictive specification by using a translog form, which provides a second-order approximation to any matching technology. We also include separately the new (current period) entrants into unemployment and those already unemployed in order to allow for the finding by Coles and Smith (1994) that new

[4] A notable exception is Warren (1996) in the U.S. context.

entrants interact with posted vacancies in a different manner than those already unemployed.

We start with the following dynamic specification

$$\ln O_{i,t} = \alpha_0 + \gamma_o \ln O_{i,t-1} + \sum_{k=u,s,v} \beta_k \ln X_{k,i,t} + .5 \sum_{k=u,s,v} \sum_{l=u,s,v} \gamma_{kl} \ln X_{k,i,t} \ln X_{l,i,t} + \delta' W_i + \varepsilon_{i,t} \quad (2)$$

where the second term on the right hand side is the lagged value of the dependent variable (the number of individuals flowing out from unemployment) and $X_{u,i,t} \equiv U_{i,t}$, $X_{s,i,t} \equiv S_{i,t}$, $X_{v,i,t} \equiv V_{i,t}$ represent those already unemployed, those flowing into unemployment in the current period and the number of lagged vacancies, respectively. W_i is a $K_I \times 1$ vector of time invariant, district-specific variables and δ is a $K_I \times 1$ vector of associated parameters. The term $\varepsilon_{i,t}$ represents the unexplained stochastic part of the matching process.

The term γlnO_{t-1} with $\gamma < 1$ allows for partial adjustment in the matching process which may be brought about by the fact that (a) the acceptance of a job offer precedes by several days or weeks the reported date of a match (start of work), with no search taking place during the interim period, and (b) information about newly posted vacancies is not diffused instantly to all job searchers.

Unlike other studies, we also analyze the nature and magnitude of differences in matching efficiency across districts. We use a two-stage procedure to analyze these inter-district differences and estimate the coefficients of a set of regressors characterizing the district-specific conditions that may influence matching.

We start with a general specification of the error as $\varepsilon_{i,t} \equiv \omega_i + \nu_{i,t}$, where ω_i stands for district-specific matching efficiency (district heterogeneity) that varies across districts but remains constant over time and $\nu_{i,t}$ is the effect of unobservable factors that vary across districts and over time. In order to make the model identifiable, we make the usual assumptions that $E\omega_i=Ev_{i,t}=0$, $E\omega_i v_{it}=0$, $E\omega_i W_i=0$, and $E\omega_i\omega_j=\sigma_\omega^2$ if $i=j$ and zero otherwise, $Ev_{it}v_{js}=\sigma_u^2$ if $i=j$ & $t=s$ and zero otherwise.

2.1 The Estimation Strategy

We estimate equation (2) using the two-stage Anderson-Hsiao Instrumental Variable (AHIV) procedure (Hsiao, 1986). In the first stage, we estimate β's and γ's of the matching function by exploiting the monthly within district variation in these time varying variables. We then use the consistent parameter estimates from this first stage and the mean log values of the associated variables to calculate district-specific efficiency residuals as the unexplained part of the matching process. In the second stage, we regress these efficiency residuals on explanatory variables W_i that are time invariant or have annual frequency and we obtain estimates of parameters δ.

The First Stage Estimation

The dependent variable and the error term in equation (2) are likely to contain a district-specific effect. As a result, in the absence of a transformation of the data that removes this effect, the error term and the explanatory variables would be correlated. The most widely used transformations are mean deviations and first differences. The mean deviation transformation is appropriate if the explanatory variables are strongly exogenous with respect to their lagged values. However, variables O_{t-1}, U_{t-1} and V_{t-1} are predetermined by the previous matching process and hence are not strictly exogenous. In particular, the lagged outflow is predetermined by definition, while unemployment and vacancies are partially predetermined by all preceding outflows. To see these links, note that the number of unemployed and vacancies at time $t\text{-}1$ is then given by identities[5]

$$U_{i,t-1} = U_{i,t-2} + S_{i,t-1} - O_{i,t-1} \tag{3}$$

$$V_{i,t-1} = V_{i,t-2} + Z_{i,t-1} - O_{i,t-1} \tag{4}$$

where $Z_{i,t-1}$ stands for the net inflow of newly posted vacancies. Since both $U_{i,t-1}$ and $V_{i,t-1}$ depend on $O_{i,t-1}$, which in turn depends on $\nu_{i,t-1}$, $U_{i,t-1}$ and $V_{i,t-1}$ are obviously only weakly exogenous. In this case, the mean deviation transformation leads to biased parameter estimates because of the correlation of the transformed values of weakly exogenous variables with the transformed error term $\varepsilon_{i,t}$ (see Appendix I).

The first difference transformation is often viewed as being superior to mean deviations in overcoming the problem of weak exogeneity and measurement errors. Although first differencing does not fully eliminate the problem of endogeneity, it permits one to identify the parameters using the AHIV or generalized method of moments methods. In particular, transforming U_{t-1} and V_{t-1} into first differences yields:

$$\Delta U_{i,t-1} = U_{i,t-1} - U_{i,t-2} = S_{i,t-1} - O_{i,t-1} \tag{5}$$

$$\Delta V_{i,t-1} = V_{i,t-1} - V_{i,t-2} = Z_{i,t-1} - O_{i,t-1} \tag{6}$$

Both first differences are obviously still affected by $O_{i,t-1}$ and are therefore correlated with $\Delta\nu_{i,t}=\nu_{i,t}\text{-}\nu_{it-1}$. But further lags of ΔU and ΔV are uncorrelated with the error term $\nu_{i,t-1}$ and therefore can be used as valid instruments to identify parameters of interest. The issue is, of course, whether the instruments have adequate explanatory power. Moreover, first differencing decreases significantly the variance of the explanatory variables while doubling

[5] We assume that all matches are brought about by the reported unemployed and vacancies (there being no out-of-register matching).

the variance of the error term and of the measurement error term, if present. This in turn leads to higher standard errors of estimated parameters.

In view of these tradeoffs, we use a specific form of mean deviations to secure a smaller variance in the error term, while at the same time avoiding the problem of weak exogeneity. In particular, we transform equation (2) into deviations from *forward* means and compute instruments as deviations from *backward* means. This transformation guarantees that the instruments are independent of the transformed error term.[6] In our empirical work we find that the explanatory power of these instruments is adequate.[7]

Adjusting for Varying Size of Districts

The variation in the size of the unit of observation (in our case the district[8]) has not been taken into account in the literature on matching. Yet, this neglect may mean that many of the existing studies have generated biased estimates. The reason for the bias, explained in detail in Appendix IV, relates to the fact that the size of a district, measured for instance by its labor force L, tends to be positively correlated with the values of O, U, S and V. In this situation, when district-level variables are not adjusted for the size of the district labor force, the intercorrelations among O, U, S and V 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-section data then provides biased estimates of coefficients unless there are constant returns to scale or U_i and V_i are uncorrelated with the

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[6] A detailed description of the transformation is provided in Appendix II.

[7] 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%).

[8] 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.

labor force L_i. The direction of the bias of β_u (β_v) 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 negative correlation (but not both) lead to a positive bias. 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, and therefore the likelihood of an incorrect acceptance of the constant returns hypothesis, 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 labour force. As we show in Appendix IV, the bias is very similar to that stemming from an omitted variable problem. In what follows we call this phenomenon the *spurious scale effect*.[9]

It can be shown that the spurious scale effect is avoided if one uses panel data and estimates a Cobb-Douglas function by fixed effects. In this case, the fixed effects transformation removes the spurious scale effects together with the district-specific time invariant effects. Unfortunately, the spurious scale effect problem remains even under fixed effects if the matching function is more complex than Cobb-Douglas. As we show in Appendix IV, in the case of our translog specification with three factors, the adjusted model should be estimated as

$$\begin{aligned} \Delta \ln O_{i,t} = {} & [\beta_u - L_i(\gamma_{uu} + \gamma_{us} + \gamma_{uv})] \Delta \ln X_{u,i,t} + \\ & + [\beta_s - L_i(\gamma_{ss} + \gamma_{us} + \gamma_{sv})] \Delta \ln X_{s,i,t} + \\ & + [\beta_v - L_i(\gamma_{vv} + \gamma_{uv} + \gamma_{sv})] \Delta \ln X_{v,i,t} + \\ & + .5 \sum_{k=u,s,v} \sum_{l=u,s,v} \gamma_{kl} \ln X_{k,i,t} \ln X_{l,i,t} \end{aligned} \tag{7}$$

The estimated coefficients of direct effects $\hat{\beta}$*'s* are hence identical with the coefficients of the unadjusted model only if the sums of γ coefficients in the brackets are equal to zero.

[9] 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.

As we show presently, this is also related to the necessary but not sufficient conditions for constant returns. In general, with bars denoting means of variables, the formula for short-term elasticities computed at the means of explanatory variables is:

$$\hat{\eta}_k = \frac{\partial \ln \frac{O}{L}}{\partial \ln \frac{X_k}{L}} = \hat{\beta}_k + \sum_{l=u,s,v} \hat{\gamma}_{kl} \overline{\ln \frac{X_l}{L}} = \hat{\beta}_k + \sum_{l=u,s,v} \hat{\gamma}_{kl} \overline{\ln X_l} - \overline{\ln L} \sum_{l=u,s,v} \hat{\gamma}_{lk} \tag{8}$$

It is obvious that the last term on the right hand side disappears if the sums of γ are zero. In this case expression (8) becomes identical to the formula for elasticities in the unadjusted model as presented in Appendix III.

The Second Stage Estimation

In the second stage we estimate $\delta s'$ which reflect the impact of district-specific variables W_i on the efficiency of matching, given unemployment, inflow into unemployment and vacancies. Using the consistent estimates from the first stage, we compute the unexplained first stage residuals and use them on the left hand side of the following cross section OLS regression:

$$\overline{\ln O_i} - \hat{\gamma}_o \overline{\ln O_{i,-1}} - \sum_{k=u,s,v} \hat{\beta}_k \overline{\ln X_{ki}} - .5 \sum_{k=u,s,v} \sum_{l=u,s,v} \hat{\gamma}_{kl} \overline{\ln X_k \ln X}_{li} = \alpha_0 + \delta' W_i + \omega_i \tag{9}$$

where the variables O and X are adjusted for district size.

Selecting the Functional Form

As mentioned earlier, we start with the translog matching function, which nests a number of simpler functional forms, including the widely used Cobb-Douglas. Moreover, we allow the new entrants into unemployment to be a separate factor from those already unemployed and test for the functional relationship between them. We therefore start with a three-input translog function and test the following four hypotheses about the appropriate form of the matching function and the separability of the already unemployed and the new entrants:[10]

1. Weak separability of X_u and X_s from X_v: This hypothesis presumes that the matching function has the form

$$\ln O = f[g(\ln X_u, \ln X_s), \ln X_v] \tag{10}$$

where $g(X_u, X_s)$ is a translog combination of factors X_u and X_s and f is a translog function of input aggregates. A test of this hypotheses represents testing whether $\beta_u/\beta_s - \gamma_{uv}/\gamma_{sv} = 0.$

2. Strong separability of X_u and X_s from X_v: This hypothesis presumes that the matching function has the form

$$\ln O = f[g(\ln X_u, \ln X_s), \ln X_v] \tag{11}$$

where $g(X_u, X_s)$ is a translog combination of factors X_u and X_s and f is a Cobb-Douglas function of input aggregates. A test of this hypothesis amounts to testing whether $\gamma_{uv} = \gamma_{sv} = 0.$

3. Cobb-Douglas functional form: This hypothesis presumes that the Cobb-Douglas functional form appropriately represents the matching process. A test of this hypothesis requires testing whether $\gamma_{uu} = \gamma_{ss} = \gamma_{vv} = \gamma_{us} = \gamma_{uv} = \gamma_{sv} = 0.$

[10] See Denny and Fuss (1977).

4. Constant Returns to Scale: This hypothesis presumes that the matching process displays constant returns to scale. A test of this hypothesis consists of testing the following four linear restrictions:

$$\beta_u + \beta_s + \beta_v = 1 \tag{12}$$

$$\gamma_{uu} + \gamma_{us} + \gamma_{uv} = 0 \tag{13}$$

$$\gamma_{ss} + \gamma_{us} + \gamma_{sv} = 0 \tag{14}$$

$$\gamma_{vv} + \gamma_{uv} + \gamma_{sv} = 0 \tag{15}$$

3. The Data and Variable Definition

In order to produce the best possible parameter estimates, we have assembled an extensive panel of district-level data on all 76 Czech and 38 Slovak districts. The Czech data cover the period January 1991 - September 1996, while the Slovak data cover the period January 1991 - December 1994. Both data sets contain monthly observations for the following variables, which are used in the first stage of estimation:

$O_{i,t}$ = the number of individuals flowing from unemployment in district *i* at time *t*; In the Czech Republic we have data on total outflow as well as outflow into jobs; In Slovakia the variable measures total outflow only;[11]

$U_{i,t}$ = the number of unemployed in district *i* at time *t-1*;

[11]The Slovak data contain information on outflow to jobs intermediated by the district labor offices (see e.g., Burda and Lubyova, 1995). However, an examination of these data indicates that this outflow is relatively low compared to the total expected outflow to jobs.

$S_{i,t}$ = the normalized number of individuals flowing into unemployment in district i at t;[12]

$V_{i,t}$ = the number of vacancies in district i at time t-1;

Since a) outflow to jobs is a theoretically preferred variable to total outflows in the matching context and b) only total outflow is available for Slovakia, we carry out the estimation for the Czech Republic using both measures and compare them to the Slovak findings based on total outflow. As we show below, the Czech estimates based on total outflow and outflow to jobs are similar. As a result, the lack of data on outflow to jobs in Slovakia does not appear to affect dramatically the validity of the Czech-Slovak comparison that we carry out.

In the second stage of estimation, we use as regressors structural and policy variables that are time invariant or are measured with annual frequency. While there are numerous conjectures and empirical evidence about the effects on matching of variables such as unemployment duration, educational structure of the unemployed and extent of government policies, no formal theory exists in this area. We have therefore collected what we consider to be important district-level variables that might affect the efficiency of matching and carry out an exploratory analysis based on these variables. With τ denoting data that are available with annual frequency, the district-level variables that we use are defined as follows:

$PUED2_{i,\tau}$ = the proportion of unemployed with secondary education;

$PUED3_{i,\tau}$ = the proportion of unemployed with university education;

$PLTU_{i,\tau}$ = the proportion of unemployed with spells longer than nine months;

[12] Although the individuals flow into unemployment in the same calendar month, they enter on different days within the month. This means that they face different probabilities of finding vacancies during the calendar month. Assuming, that the inflow is approximately uniform over the month, we multiply the total monthly inflow by .5.

$Q_{i,\tau}$ = industrial output in million crowns per member of the district labor force,

$E_{i,\tau}$ = the number of self-employed persons and firms with fewer than 25 employees per member of the district labor force;

$IA90_i$ = the ratio of the value of agricultural production to industrial production in 1990 (the agricultural vs. industrial makeup of the district at the start of the transition);

$LDIS_i$ = the natural logarithm of the average distance (in kilometers) from the district capital to the Austrian or West German border (Austrian border for Slovakia);

$DENS_i$ = the density of population in the district (individuals per km^2);

$ALMP_{i,\tau}$ = the natural logarithm of expenditures on active labor market programs per members of the district labor force;

SUD_i = a dummy variable coded 1 if the district falls into the Sudeten lands and 0 otherwise.

The education variables are included in order to test the hypothesis that more educated individuals are more able to find jobs than their less educated counterparts. The hypothesis has grounding on both sides of the market: demand for individuals with human capital is hypothesized to have increased during the transition and more educated individuals are expected to be more adept at searching for jobs.

The share of unemployed with spells longer than nine months is included because individuals with long-term duration of unemployment may be expected to have different unobservable (to the econometrician) characteristics than the short term unemployed. In particular, the long term unemployed may suffer from discouragement, stigma, loss of skills, and skill mismatch.

The ratio of industrial output to the labor force controls for the differences in the level of economic activity in the various districts. It is expected to have a positive effect on matching unless vacancies fully capture this effect in the first stage of estimation.

The number of small-scale entrepreneurs proxies for the presence of new firms that are an important source of dynamism and jobs in the transition economies. It is hypothesized that this variable will have a positive effect on matching.

The ratio of the value of agricultural production to industrial production in 1990 proxies for the sectoral diversity of districts at the start of the transition. This variable proxies for the importance of different initial conditions on the functioning of the district labor market.

The distance of the district capital from the West German or Austrian border captures the ease with which the population can move to work in the high income western economies. By including this variable, we are able to check the widely held belief that the low unemployment in the Czech Republic is to a significant extent brought about by the ease with which Czech workers can participate in the German and Austrian labor markets.

Population density of the district proxies for the average geographic proximity of agents in the labor market. It was included to assess if this geographic proximity affects matching.

Within the Czech Republic, we were able to include two additional variables: the amount of government funds spent on active labor market policy programs (ALMPs) and whether the district belongs to the Sudeten lands. ALMP spending ideally captures the policies of district labor offices that are aimed at improving matching and hence reducing unemployment.[13] The ALMP variable should have a positive effect on job matching if

[13] Like other governments in the region, the Czech and Slovak governments introduced active labour market programs shortly after the Velvet Revolution. If these programs are successful in their objective of improving job matches and creating jobs, they should increase outflows from unemployment. For a more detailed description of these programs, see e.g., Münich and Terrell (1996) and Münich (1996).

higher expenditures yield more job matches and higher outflows.[14] Indeed, Svejnar et al. (1995) find the effect to be positive in the Czech Republic but insignificant in Slovakia, Svejnar et al. (1994) find the results to be sensitive to model specification and Burda and Lubyova (1995) find the effect to be positive in both republics. The econometric problem is that ALMP spending may be influenced by the levels of unemployment and outflow, thus possibly causing a spurious correlation between ALMP spending and outflows. Boeri and Burda (1995) therefore instrumented ALMP, using as instruments district-specific ALMP spending aggregated over the four quarters of the current year, the other regressors in the matching function, and in Slovakia also an explicit allocation rule for ALMP. As we have shown above, the unemployment and vacancy data are not strongly exogenous and the contemporaneous aggregation of ALMP may hence not be a valid instrument since it contains the district-specific effect. As a result, we instrument current ALMP spending in a district by its lagged values and, as we discuss below, we also test the sensitivity of estimates to the choice of other instruments.

By distinguishing Czech districts that fall into the Sudeten lands from others, we control for the possibly different structure and institutional nature of this part of the Czech Republic. The Sudeten lands were significantly depopulated after World War II as the significant German population was moved from these regions to West Germany. Property rights were less clearly established and more sluggish restitutions during the transition may have caused the labor market adjustments to differ in these regions from the rest of the country.

[14] The direct effect of ALMP on outflow counted by the share of total outflow flowing into ALMP programs cannot be used, since the spending say nothing about substitution effects (regular employees within an enterprise are substituted by subsidised employees), displacement effects (job loss in other enterprises as a consequence of changes on product market due to subsidised jobs). There is also an operational problem since some outflow to short term public works is not consistently reported as a part of outflow and also the better prospect to flow out due to the previous retraining cannot be distinguished directly

4. Basic Statistics and Econometric Estimates

4.1 Basic Statistics

Before discussing the estimates of our model, it is useful to note the basic patterns and interrelationships among the principal variables. As may be seen from Figure 2, between 1991 and 1994, the Czech and Slovak republics experienced a similar pattern of unemployment growth, but the rate of increase was much steeper in Slovakia than in the Czech lands. Indeed, with the Slovak labor force being about one-half of the Czech labor force, one observes the number of unemployed in Slovakia rising to about twice the absolute level observed in the Czech Republic (over 350,000 vs. less than 200,000). Moreover, while the total number of vacancies remained at a low level in range 8,000-13,000 in Slovakia, it gradually rose to about 100,000 in the Czech Republic. The average district-level statistics, reported in Appendix Table A1, provide a similar picture of a high unemployment and low vacancy economy in Slovakia versus a more balanced unemployment-vacancy situation in the Czech lands. An analysis of the underlying district-level mismatch indicates that while in the Czech Republic one finds a number of districts where the number of vacancies is similar or even exceeds the number of unemployed, in most Slovak districts the number of unemployed greatly exceeds the number of vacancies (Münich, Svejnar and Terrell, 1996).

In Figure 3 we plot the evolution of the overall relationship between the number of unemployed and number of vacancies in the two republics (the values for Slovakia are scaled by the size of the labor force to be comparable with the Czech republic and the scaling is also used for the Slovak flows on Figure 4 and 5). The vertical parts of the two plots relate to 1991 and depict the rapid rise of unemployment in the early phase of the transition. Since 1992, one observes negative Beveridge-type relationships between unemployment and vacancies in both republics. However, the Beveridge curve is steep in Slovakia, capturing the fact that vacancies varied little during significant unemployment fluctuations, while in the Czech Republic it is flat, reflecting the fact that vacancies fluctuated dramatically with relatively less pronounced variations in unemployment. As may be seen from Appendix Table A2, the negative correlation between adjusted district-level vacancies and unemployment is stronger for each year in the Czech Republic than in Slovakia.

4.2 Econometric Estimates

The First Stage

Since the transition period has been characterized by a rapid and profound regime change, we have collected monthly data in order to estimate the model separately for each year and test for changes in parameter estimates over time. The earliest data are from 1991 and these are used as instruments for the 1992 regressions.

The tests of the form of the matching function are summarized in Table 1. The tests are based on outflow to jobs and total outflow in the Czech Republic, and on total outflow in Slovakia. As may be seen from Table 1, the results are similar in the two countries in that only the hypothesis of weak separability between U and S cannot be rejected in any year. Importantly, the usual Cobb-Douglas specification is rejected in the Czech Republic in all years and in Slovakia in two out of the three years. Our results hence indicate that it is important to allow for a more flexible functional form than Cobb-Douglas.

We have also tested the frequently accepted hypothesis of constant returns to scale. As may be seen from Table 1, like the Cobb-Douglas specification, this hypothesis too is rejected in the Czech Republic in all years and in Slovakia, depending on specification, in one or two of the three years.

Finally, we have checked if the estimates of returns to scale differ when one uses adjusted and unadjusted data. We have found the two sets of estimates to differ somewhat, but not excessively. An examination of the estimated γ coefficients indicates that their sum is often not significantly different from zero, thus suggesting that this may be one of the reasons for the observed similarities of the two set of coefficients in the present study.

On the basis of the tests reported in Table 1, we have used the translog form with weak separability of U and S in further analysis.[15] The estimates of matching function elasticities and returns to scale, evaluated at the geometric means of variables, are reported in Table 2.[16] The estimates are again based on outflow to jobs and total outflow in the Czech Republic, and on total outflow in Slovakia. As may be seen from Table 2, the Czech estimates based on outflow to jobs and total outflows are similar, with the former showing somewhat higher point estimates of the returns to scale than the latter, especially in the first two years. The estimates are precisely estimated and they fluctuate without a trend in the region of increasing returns, ranging from 2.9 to 3.5 for outflows to jobs and 2.2 to 2.8 for total outflows. Contrary to many earlier studies, our estimates hence indicate that the matching process in the Czech Republic displays strongly increasing returns to scale. The corresponding estimates of returns to scale in Slovakia show an increasing trend with the point estimate rising from 0.5 in 1992 to 1.5 in 1993 and 2.8 in 1994. The increasing returns phenomenon may be brought about by a number of factors mentioned in the introduction. As was pointed to us by Dale Mortensen (1997), the one appealing interpretation of the Czech-Slovak differences in returns to scale may be attributed to "animal spirits". In particular, with the Czech and Slovak fundamentals being very similar, Mortensen (1997)

[15] The complete sets of estimated parameters for the general and weakly separable functional forms are reported for both countries in Appendix Table A4.

[16] The corresponding estimates based on the unconstrained translog model are analogous and are reported in Appendix Table A5.

shows that in an extended Pissarides model one can attribute the higher returns to scale in the Czech Republic to greater optimism in the Czech Republic.

In the Czech Republic and Slovakia the chi-square test does not allows us to reject the hypothesis of equality of returns to scale in consecutive years at conventional test levels (i.e., 1, 5%). Indeed, while the data do not allow us to reject strictly the hypothesis that the returns to scale were identical during the entire period in the Czech Republic, the acceptance of equality of returns to scale in consecutive years is not as strong in Slovakia. Finally, using the Chow test we are able to reject the hypothesis that the matching function coefficients were identical in any two contiguous years (Table 1, 6th hypothesis).

A comparison of the individual elasticity estimates in Table 2 indicates that vacancies and the newly unemployed play a much more important part in the Czech Republic than in Slovakia. The Czech estimates based on outflow to jobs and total outflow are again similar and they yield vacancy elasticities that are all significantly different from zero and range from 0.7 to 1.2. In contrast, the estimated elasticities in Slovakia range from 0.2 to 0.3 and in 1992 and 1994 they are not significantly different from zero.

The relative part played by the newly unemployed is analogous to that played by vacancies. The estimated elasticities in the Czech Republic are all positive, significantly different from zero and ranging from 0.4 to 0.9. In Slovakia, the estimates range from zero to 0.2 and, with the possible exception of the 1993 estimate, they are not significantly different from zero at conventional test levels.[17]

Finally, the estimated elasticities of the existing unemployed are high and statistically significant in the Czech Republic, ranging from 1.0 to 1.9. The corresponding Slovak estimates rise from 0.4 in 1992 to 1.0 in 1993 and 2.5 in 1993, with the 1991 estimate not

[17] The 1993 estimate is significant at the 10% test level.

being significantly different from zero at conventional test levels. The existing unemployed thus contribute in an important way to the outflow in the Czech Republic and increasingly so also in Slovakia.

The different results for the Czech and Slovak republics raise the issue of whether the matching technology is different in high and low unemployment districts. In particular, while the Czech observations cover a wide variety of unemployment-vacancy combinations, with vacancies exceeding the number of unemployed in some districts, the Slovak districts reflect primarily low vacancy-high unemployment combinations. As a result, in order to assess whether high unemployment districts in the Czech Republic have similar matching technology as those in Slovakia, we have reestimated the matching function using only data from the half of the Czech districts with higher unemployment-vacancy ratio. In the context of theories of multiple equilibria, this excercise allows us to test Randall Filer's hypothesis that increasing returns in matching are not resulting from a higher intensity of search, as suggested by Dale Mortensen, but are brought about by higher intensity of hiring in the presence of a low local unemployment rate (high opportunity costs of not hiring). The estimates we obtained indicate that increasing returns to scale are observed even in high unemployment-low vacancy districts of the Czech Republic -- the estimated returns to scale (standard errors) are 3.06 (.358) for 1992, 3.74 (.423) for 1993, 3.14 (.463) for 1994 and 3.08 (.459) for 1995. The results hence suggest that the matching technology is similar in high and low unemployment districts of the Czech Republic and that the increasing returns are more consistent with multiple equilibria theories based on concepts such as high intensity of search than high intensity of hiring.

As these results indicate, the demand side of the matching process, as proxied by vacancies, has been much weaker in Slovakia than the Czech Republic during the transition.

Indeed, while vacancies have been an important component of the matching process in the Czech Republic, in Slovakia vacancies appear to have been an insignificant factor in the outflow of individuals from unemployment. Moreover, while the Czech matching process was relatively stable between 1992 and 1995, the Slovak process showed a major development from having virtually insignificant parameters in 1992, to becoming more structured by 1994.

(Second Stage

The estimated effects on matching of the structural and policy variables that are time invariant or measured only with annual frequency are reported in Table 3. We report estimates from four representative specifications since some of the estimates are somewhat sensitive to the inclusion of individual variables, we report estimates of four representative specifications. Since the number of estimates is large and the outflow to jobs and total outflow estimates for the Czech Republic are again very similar, we report only the estimates for outflow to jobs.

As may be seen from Table 3, education of the unemployed has a systematic effect in all years. In the Czech Republic the proportion of university educated individuals in the district unemployment pool increases the efficiency in matching in all years, while in Slovakia the effect only becomes positive in 1994, after being negative in 1992 and 1993. In contrast, the proportion of unemployed with high school education is found to increase matching efficiency in Slovakia in 1992, 1993 and in about one-half of the specifications in 1994, while it has an insignificant or a negative effect in the Czech Republic. Hence, while the two republics have a similar educational structure of the population, the matching process appears to favor more the university educated in the Czech Republic and those with high school education in Slovakia.

In both countries, the long-term (more than nine months) unemployed have a negative effect on the efficiency of matching in almost all specifications and years. This effect goes in the expected direction and has a much greater magnitude in Slovakia, where the problem of long-term unemployment has been more acute. Our estimates thus underscore the relevance of factors such as scaring and loss of skills in long-term unemployment.

The prevalence of small enterprises has a positive effect on matching in Slovakia in 1992, but the effect becomes insignificant in the later years. In the Czech Republic the coefficient is significant in the more parsimonious specifications of 1993-95, yielding a positive effect in 1994 and 1995, but a negative effect in 1993. Since small enterprises re sometime viewed as a source of needed job opportunities because they engage in labor-intensive activities, our results indicate that this effect, if present, does not work systematically through the efficiency of matching of the unemployed and vacancies.

In the Czech Republic, where ALMP expenditure data are available, we report estimates using the previous year's value of ALMP expenditure as an instrument. As can be seen from the table, the amount of ALMP spending is found to have had an insignificant effect on the efficiency of matching in 1992, a positive effect in 1993 and negative effects in 1994 and 1995. These estimates suggest that the effect of government's ALMP expenditures was insignificant in the early stages of the transition, became positive in 1993 and actually hindered the matching process in the more mature transition phase of the mid 1990s. The results raise the possibility that the one-year lagged ALMP expenditures, which we used as an instrument, did not avoid picking up the positive link between local unemployment and expenditures on ALMPs. Since our inquiries with officials at the district labor offices indicated that ALMP funds were indeed distributed with a view toward the unemployment and vacancy situation in the individual districts, we have also carried out our expenditure coefficient to be statistically insignificant. We have also performed the analysis using as instruments variables that capture the structure of district-level agencies (such as the per capita number of primary schools, teachers and government agencies) that could be correlated with the allocation of the budget among various categories including ALMP, and are likely to be independent of the local labor market conditions. These specifications generally also yielded insignificant effects of ALMP on matching efficiency. In view of (a) the inherent difficulty of separating the true underlying effect of ALMP expenditures on the efficiency of job matching and (b) the in-depth nature of our investigation, we conclude that the effect of ALMP expenditures has most likely been insignificant.

The remaining variables produce insignificant or only occasionally significant coefficients. The extent of agriculture in the district at the start of the transition process has no effect on matching in Slovakia and, with the exception of 1993 when it is negative, the effect is also insignificant in the Czech Republic. The level of industrial production, population density, distance to a western border, and the location of a district in the Sudeten lands have no effect on the efficiency of matching. The observed insignificance of these variables is important in that it suggests that the importance that some of them (e.g., distance from the border and Sudeten lands) have been attributed in popular policy discussions about the low unemployment in the Czech Republic is unfounded from the matching standpoint.

5. Concluding Observations

In the early to mid 1990s, all Central and East European countries except for the Czech Republic suffered double digit unemployment rates coupled with long unemployment durations. The Czech Republic experienced very low unemployment which was in large part brought about by a high outflow rate from unemployment. In our analysis, we have compared the nature and the low unemployment in the Czech Republic is unfounded from the matching standpoint.

5. Concluding Observations

In the early to mid 1990s, all Central and East European countries except for the Czech Republic suffered double digit unemployment rates coupled with long unemployment durations. The Czech Republic experienced very low unemployment which was in large part brought about by a high outflow rate from unemployment. In our analysis, we have compared the nature and determinants of Czech Republic's outflow from unemployment to that of its traditional counterpart republic, Slovakia. The reason for carrying out this comparison is that it brings us as close as possible to an interesting controlled experiment. Until 1993, the two republics shared the same currency, legal system and institutional framework. Moreover, except for unemployment, their aggregate economic indicators were very similar. Yet, the Slovak labor market indicators have resembled those of Poland, Hungary and other Central and East European economies rather than those of the Czech Republic. The Czech-Slovak comparison is hence useful per se as well as for understanding the difference between the Czech experience and that of other economies in the region. Our analysis, based on detailed monthly data from the Czech and Slovak districts points to the importance of the following factors.

In the early 1990s, the Czech and Slovak republics experienced a similar pattern of unemployment growth, but the rate of increase was much steeper in Slovakia than in the Czech lands. Moreover, while the total number of vacancies remained at a low level in Slovakia, it gradually rose to a high level in the Czech Republic. One hence observes a high unemployment-low vacancy economy in Slovakia versus a more balanced unemployment-vacancy situation in the Czech lands. At the district-level one finds a number of districts in the Czech Republic where the number of vacancies is similar or even exceeds the number of unemployed, while in most Slovak districts the number of unemployed greatly exceeds the number of vacancies.

In Slovakia, the number of vacancies varied little during relatively significant unemployment fluctuations since 1991, while in the Czech Republic vacancies fluctuated dramatically with relatively less pronounced variations in unemployment. The greater impact of demand on reducing unemployment in the Czech Republic than Slovakia is also visible from the fact that within each year the negative correlation between adjusted district-level vacancies and unemployment is stronger in the Czech Republic than in Slovakia.

In the context of the large literature on matching functions, we find that in many years the usual Cobb-Douglas specification and the hypothesis of constant returns to scale are rejected. A translog matching function with weak separability between the existing and newly unemployed is found to be the functional form best supported by the Czech and Slovak data. The Czech data generate very strong increasing returns to scale throughout the 1992-95 period, while the Slovak data point to initially very low but steadily increasing returns. An important part of our explanation of the Czech-Slovak differences hence lies in the much higher returns to scale in matching in the Czech Republic than in Slovakia in the crucial early period of the transition. In the context of recent macroeconomic literature, this outcome points to possible existence of multiple equilibria. In particular, Mortensen (1997) interprets our findings as indicative of greater optimism and search intensity in the Czech Republic than in the Slovak Republic.

Our theoretical analysis also indicates that by not adjusting data for the varying size of districts or regions, previous studies may have generated estimates of the returns to scale of the matching function that were biased toward unity. In the present study, the difference in estimated coefficients between the adjusted and unadjusted data is not very large. Our estimates also indicate that vacancies and the newly unemployed have played a much more important part in the matching process in the Czech Republic than in Slovakia. The demand side of the matching process, as proxied by vacancies, has thus been much weaker in Slovakia than the Czech Republic. In fact, while matching between the unemployed and vacancies was important in the Czech Republic, in Slovakia vacancies appear to have been an insignificant factor in the outflow of individuals from unemployment.

Our estimates indicate that the process of outflow from unemployment in the Czech Republic was relatively stable between 1992 and 1995, while in Slovakia the process developed from having virtually insignificant parameters in 1992 to being driven by the number of unemployed in 1994.

Only few of the structural-policy variables generate systematic effects on matching. The long-term unemployed have a negative effect on the efficiency of matching in both countries. The magnitude of this effect is much greater in Slovakia, where the problem of long-term unemployment has been more acute.

Education of the unemployed has a systematic effect but it varies across the republics. In the Czech Republic the proportion of university educated individuals in the district unemployment pool increases the efficiency in matching, while in Slovakia it is primarily the proportion of unemployed with high school education that is found to increase matching efficiency. Hence, while the two republics have a similar educational structure of the population, the matching process appears to favor more the highly educated in the Czech Republic and those with high school education in Slovakia.

We have carried out a careful instrumental variable analysis of the effects of expenditures on active labor market policies. We find that in most specifications the effect of ALMP on the efficiency of matching is statistically insignificant. Our estimates hence question the earlier findings of a positive effect of ALMP that were based on OLS estimates or weakly exogenous instrumetal variables.

Overall, our study suggests that the exceptionally low unemployment rate in the Czech Republic as compared to Slovakia and the other Central and East European economies has been brought about principally by (1) a rapid increase in vacancies along with unemployment, resulting in a balanced unemployment-vacancy situation at the aggregate as well as district level, (2) a major part played by vacancies and the newly unemployed in the outflow from unemployment, (3) a matching process with strongly increasing returns to scale throughout (rather than only in parts of) the transition period, and (4) ability to keep the long term unemployed at relatively low levels. Since the Czech economy has registered overall economic growth that is similar to that of the neighboring high unemployment economies (e.g., Slovakia or Poland) the interesting question is whether the favorable vacancy situation, coupled with the strong matching process, was brought about by a high level of initial economic activity (better initial conditions) or a delayed restructuring of firms.

Appendix I

Transforming equation (2) into mean deviations, we obtain the following form:

$$o_{i,t} - \bar{o}_i = \gamma\,(o_{i,t} - \bar{o}_{i,t-1}) + \beta'(x_{i,t} - \bar{x}_i) + \epsilon_{i,t} - \bar{\epsilon}_i, \quad i=1,\ldots,N, \quad t=2,\ldots,T$$

where

$$\bar{o}_i = \frac{1}{T-1}\sum_{t=2}^{T} o_{i,t}, \quad \bar{o}_{i,t-1} = \frac{1}{T-1}\sum_{t=1}^{T-1} o_{i,t}, \qquad (16)$$

$$\bar{x}_i = \frac{1}{T-1}\sum_{t=2}^{T} x_{i,t}, \quad \bar{\epsilon}_i = \frac{1}{T-1}\sum_{t=2}^{T} \epsilon_{i,t}\,.$$

where the letters denote logs of variables and input factors $X's$ are stacked in vector $x_{i,t}$. Each district-specific mean of outflow, $\bar{o}_{i,t-1}$, incorporates all $T-1$ error terms $\varepsilon_{i,t}$ that are specific to a given district i. The necessary assumption of independence of the error and explanatory variables is therefore explicitelly violated in partial adjustment models. Balestra and Nerlove (1967), Maddala (1971), Nickell (1981), and Arellano and Bond (1991) provide detailed evidence on the size and sign of biases of both $\hat{\gamma}$ and β's in experimental settings. For $\gamma > 0$, the bias in $\hat{\gamma}$ is negative and decreases with T, it is not approaching zero as $\gamma = 0$, and it is larger if other right hand side variables are include. Moreover, the bias in β is positive if other right hand side variables are positively related to lagged values o_{it-1}.

Appendix II

Equation (2) transformed into deviations from forward means has following form:

$$o^*_{i,t} = \gamma\, o^*_{i,t-1} + \beta' X^*_{i,t} + \epsilon^*_{i,t} \qquad (17)$$

where

$$o^*_{i,t} = o_{i,t} - \frac{1}{1+T_F}(o_{i,t} + o_{i,t+1} + \ldots + o_{i,t-1+T_F}) \qquad (18)$$

$$o^*_{i,t-1} = o_{i,t-1} - \frac{1}{1+T_F}(o_{i,t-1} + o_{i,t} + \ldots + o_{i,t-1+T_F}) \qquad (19)$$

$$x^*_{i,t} = x_{i,t} - \frac{1}{1+T_F}(x_{i,t} + x_{i,t+1} + \ldots + x_{i,t+T_F}) \qquad (20)$$

$$\epsilon^*_{i,t} = \epsilon_{i,t} - \frac{1}{1+T_F}(\epsilon_{i,t} + \epsilon_{i,t+1} + \ldots + \epsilon_{i,t+T_F}) \qquad (21)$$

The appropriate instruments are computed as deviations from backward means as

$$o^{IV}_{i,t-1} = o_{i,t-1} - \frac{1}{T_B}(o_{i,t-1} + o_{i,t-2} + \ldots + o_{i,t-T_B}) \tag{22}$$

$$o^{IV}_{i,t-2} = o_{i,t-2} - \frac{1}{T_B}(o_{i,t-2} + o_{i,t-3} + \ldots + o_{i,t-1-T_B}) \tag{23}$$

$$\vdots$$

$$x^{IV}_{i,t} = x_{i,t} - \frac{1}{1+T_B}(x_{i,t} + x_{i,t-1} + \ldots + x_{i,t-T_B}) \tag{24}$$

$$x^{IV}_{i,t-1} = x_{i,t-1} - \frac{1}{1+T_B}(x_{i,t-1} + x_{i,t-2} + \ldots + x_{i,t-T_B}) \tag{25}$$

$$\vdots$$

The transformed error term $\varepsilon^*_{i,t}$ incorporates all terms $\varepsilon_{i,t}$, $\varepsilon_{i,t+1},\ldots,\varepsilon_{i,t+TF}$ and the transformed right hand side variables $o^*_{i,t-1}$ and $x^*_{i,t-1}$ incorporate terms $\varepsilon_{i,t},\varepsilon_{i,t+1},\ldots,\varepsilon_{i,t+TF}$. Obviously $COR(\varepsilon^*_{i,t}, o^*_{i,t-1}) \neq 0$ and $COR(\varepsilon^*_{i,t}, x^*_{i,t}) \neq 0$. Instruments $o_{i,t-1}{}^{IV}$, $o^{IV}{}_{i,t-2}$ and $x^{IV}{}_{i,t-2}{}^{IV}$ incorporate terms $\varepsilon_{i,t-1}$, $\varepsilon_{i,t-2},\ldots,\varepsilon_{i,t-TB}$. The instruments are not correlated with forward errors $\varepsilon_{i,t}$, $\varepsilon_{i,t+1},\ldots,\varepsilon_{i,t+TF}$ and therefore $COR(\varepsilon^*_{i,t}, o^{IV}{}_{i,t-1})=0$ and $COR(\varepsilon^*_{i,t}, x^{IV}{}_{i,t})=0$. Having T cross-sections we could in principle increase the set of instruments with t (Holtz-Eakin 1988, Arrelano and Bond (1991). Since the explanatory power of instruments of farther lags is small, we limit the number of lags used as instruments to three.

Appendix III

Partial derivatives of equation (3) give us elasticities of matching with respect to the number of unemployed, inflow into unemployment and vacancies. The elasticities are given as:

$$\eta_u = \beta_u + \gamma_{uu} \ln X_u + \gamma_{us} \ln X_s + \gamma_{uv} \ln X_v \tag{26}$$

$$\eta_s = \beta_s + \gamma_{ss} \ln X_s + \gamma_{us} \ln X_u + \gamma_{sv} \ln X_v \tag{27}$$

$$\eta_v = \beta_v + \gamma_{vv} \ln X_v + \gamma_{uv} \ln X_u + \gamma_{sv} \ln X_s \tag{28}$$

and they vary with explanatory variables rather than being constants as in the case of the Cobb-Douglas specification. The elasticity of scale η is given by:

$$\eta = \eta_u + \eta_s + \eta_v$$

Substituting the expressions for elasticities into the formula for returns to scale and rearranging yields

$$\eta = (\beta_u + \beta_s + \beta_v) + (\gamma_{uu} + \gamma_{us} + \gamma_{uv}) \ln X_u + (\gamma_{ss} + \gamma_{us} + \gamma_{sv}) \ln X_s + (\gamma_{vv} + \gamma_{sv} + \gamma_{uv}) \ln X_v \tag{30}$$

Since constant returns to scale must hold independently of the values of the explanatory variables, the following linear parameter restrictions have to be satisfied:

$$\beta_u + \beta_s + \beta_v = 1 \tag{31}$$

$$\gamma_{uu} + \gamma_{us} + \gamma_{uv} = 0 \tag{32}$$

$$\gamma_{ss} + \gamma_{us} + \gamma_{sv} = 0 \tag{33}$$

$$\gamma_{vv} + \gamma_{uv} + \gamma_{sv} = 0 \tag{34}$$

Appendix IV

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 $(\beta_u+\beta_v>1)$. As a result of homogeneity, the outflow, unemployment and vacancies in each district, O_i U_i *and* V_i, are proportional to national aggregates O, U, *and* V,

$$\begin{aligned} O_i &= l_i O \\ U_i &= l_i U \\ V_i &= l_i V \end{aligned} \tag{35}$$

where l_i is the share of district i in the national labor force, defined as L_i/L.[18] Not taking the district size into account and estimating the matching function on unadjusted cross-sectional data amounts to estimating the model

$$\ln O_i = \alpha + \beta_u \ln U_i + \beta_v \ln V_i + \epsilon_i \tag{36}$$

Substituting (35) into (36), we get

$$\ln(l_i O) = \alpha + \beta_u \ln(l_i U) + \beta_v \ln(l_i V) \tag{37}$$

which in turn yields

$$\ln l_i = (\beta_u + \beta_v)\ln l_i + (\alpha - \ln O + \beta_u \ln U + \beta_v \ln V) + \epsilon_i \tag{38}$$

It is obvious that the estimation of (38) is identical to estimating (36). However, (38) represents a regression of l_i on itself plus a constant term. It will hence generate constant returns to scale $(\hat{\beta}_u+\hat{\beta}_v=1)$ and a zero constant term $(\hat{\alpha}=lnO-\hat{\beta}_u lnU-\hat{\beta}_v lnV)$. The estimation of model (36) therefore yields biased estimates since we have postulated increasing returns to scale.

A remedy for the problem is to adjust the variables by the district size in order to obtain the following model:

$$\ln\left(\frac{O_i}{L_i}\right) = \alpha + \beta_u \ln\left(\frac{U_i}{L_i}\right) + \beta_v \ln\left(\frac{V_i}{L_i}\right) + \epsilon_i \tag{39}$$

[18] 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.

which can be rearranged as

$$\ln O_i = \alpha + \beta_u \ln U_i + \beta_v \ln V_i + (\beta_u + \beta_v - 1)\ln L_i + \epsilon_i \tag{40}$$

A comparison of the adjusted model (40) to the unadjusted model (36) indicates that they are equivalent if and only if at least one of the two following conditions is satisfied:

(i) if $\beta_u + \beta_v - 1 = 0$ (the underlying matching displays constant returns to scale)

or

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

In our example, neither condition is satisfied because (i) we are assuming increasing returns to scale $(\beta_u + \beta_v - 1 > 0)$ and (ii) $U_i = UL_i/L$ *and* $V_i = VL_i/L$, resulting in $Cor(lnL_i, lnU_i) = Cor(lnL_i, lnV_i) = 1$.

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 intercorrelations among the unadjusted variables can of course be checked in advance. Judging from the Czech and Slovak data at our disposal, these intercorrelations are positive and significant.

b) Data Transformations, Functional Form and the Spurious Scale Effect

When the district size does not change over time, a fixed effects transformation eliminates the spurious scale effect in a Cobb-Douglas but not translog functional forms of the matching function.

A fixed effects (first difference) transformation of the adjusted equation (40) leads to

$$\Delta \ln O_{i,t} - \Delta \ln L_{i,t} = \beta_u \Delta \ln U_{i,t-1} + \beta_v \Delta \ln V_{i,t-1} + (\beta_u + \beta_v)\Delta \ln L_{i,t-1} \tag{41}$$

If the size variable L_i is constant over time, (41) can be simplified to yield

$$\Delta \ln O_{i,t} = \beta_u \Delta \ln U_{i,t-1} + \beta_v \Delta \ln V_{i,t-1} \tag{42}$$

Equation (42) is identical to a fixed effects transformation of the unadjusted model (36). There is hence no need to adjust variables if L_i is constant over time and a fixed effects transformation is used together with a Cobb-Douglas specification of the matching function.

The mis-specification caused by the spurious scale effect remains even with fixed effects and constant L_i if the true form of the matching function is more general than Cobb-Douglas, as in our case of a translog specification with three inputs. Such a model may be written in an adjusted form as

$$\ln(\frac{O_i}{L_i}) = \sum_{k=u,s,v} \beta_k \ln(\frac{X_{k,i}}{L_i}) + .5 \sum_{k=u,s,v} \sum_{l=u,s,v} \gamma_{kl} \ln(\frac{X_{k,i}}{L_i})\ln(\frac{X_{l,i}}{L_i}) \tag{43}$$

Decomposing the logarithms of adjusted variables and rearranging individual terms, one obtains

$$\begin{aligned}
&\ln O_i - \ln L_i = \\
&\sum_{k=u,s,v} \beta_k \ln X_{k,i} + .5 \sum_{k=u,s,v} \sum_{l=u,s,v} \gamma_{kl} \ln X_{k,i} \ln X_{l,i} + \\
&\ln L_i \sum_{k=u,s,v} \beta_k + \ln L_i \sum_{k=u,s,v} \ln X_{k,i} \sum_{l=u,s,v} \gamma_{kl} + \\
&.5 (\ln L_i)^2 \sum_{k=u,s,v} \sum_{l=u,s,v} \gamma_{kl}
\end{aligned} \tag{44}$$

and applying the fixed effects transformation by first differencing yields

$$\begin{aligned}
&\Delta \ln O_i - \Delta \ln L_i = \\
&\sum_{k=u,s,v} \beta_k \Delta \ln X_{k,i} + .5 \sum_{k=u,s,v} \sum_{l=u,s,v} \gamma_{kl} \, \Delta \, (\ln X_{k,i} \ln X_{l,i}) + \\
&\Delta \ln L_i \sum_{k=u,s,v} \beta_k + \ln L_i \sum_{k=u,s,v} \Delta \ln X_{k,i} \sum_{l=u,s,v} \gamma_{kl} + \\
&.5 \Delta (\ln L_i)^2 \sum_{k=u,s,v} \sum_{l=u,s,v} \gamma_{kl}
\end{aligned} \tag{45}$$

Since the district size is constant over time, first differencing cancels out several terms and (45) becomes

$$\Delta \ln O_i = \sum_{k=u,s,v} \beta_k \Delta \ln X_{k,i} + .5 \sum_{k=u,s,v} \sum_{l=u,s,v} \gamma_{kl} \, \Delta \, (\ln X_{k,i} \ln X_{l,i}) + \ln L_i \sum_{k=u,s,v} \Delta \ln X_{k,i} \sum_{l=u,s,v} \gamma_{lk} \tag{46}$$

Expanding the sum operator and grouping the terms by input factors yields

$$\begin{aligned}
&\Delta \ln O_i = \\
&[\beta_u - L_i(\gamma_{uu} + \gamma_{us} + \gamma_{uv})] \Delta \ln X_{u,i} + \\
&[\beta_s - L_i(\gamma_{ss} + \gamma_{us} + \gamma_{sv})] \Delta \ln X_{s,i} + \\
&[\beta_v - L_i(\gamma_{vv} + \gamma_{uv} + \gamma_{vs})] \Delta \ln X_{v,i} + \\
&.5 \sum_{k=u,s,v} \sum_{l=u,s,v} \gamma_{kl} \Delta \ln X_{k,i} \ln X_{l,i}
\end{aligned} \tag{47}$$

Equation (47) resembles the unadjusted specification. The direct effect parameters enter as a sum of the original coefficients βs' minus the product of L_i and the sum of cross-effect parameters γs'. Moreover, the sum of the cross-effect parameters is identical with left hand side of constraints (32)-(34). It follows that the spurious size effect disappears and the fixed effects transformation of the unadjusted specification is identical with the adjusted one only when these constraints are satisfied.[19] Since one knows nothing about the fulfilment of these constraints in advance, any test of returns to scale based on coefficient estimates from an unadjusted specification of the matching function is likely to yield biased estimates.

[19]Constraint (31) is not binding here, since it was removed by the fixed effects transformation as in the case of the simple Cobb-Douglas specification.

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Table 1: Tests of the Functional Form of the Matching Function and Stability of Parameters over Years

Czech Republic - Outflows to Jobs

Year	1992		1993		1994		1995	
Test	χ^2	*P-value*	χ^2	*P-value*	χ^2	*P-value*	χ^2	*P-value*
1.Hypothesis[1)] - $\chi^2(1)$	2.45	*0.1173*	0.16	*0.6852*	1.62	*0.2030*	0.25	*0.6196*
2.Hypothesis[2)] - $\chi^2(2)$	10.23	*0.0060*	5.64	*0.0595*	9.47	*0.0088*	2.84	*0.2423*
3.Hypothesis[3)] - $\chi^2(5)$	30.26	*0.0000*	22.00	*0.0005*	19.89	*0.0013*	33.20	*0.0000*
4.Hypothesis[4)] - $\chi^2(6)$	32.09	*0.0000*	37.30	*0.0000*	20.29	*0.0025*	34.94	*0.0000*
5.Hypothesis[5)] - $\chi^2(4)$	28.32	*0.0000*	58.04	*0.0000*	20.00	*0.0005*	25.37	*0.0000*
6.Hypothesis[6)] - $\chi^2(10)$	n.a.	*n.a.*	36.90	*0.0000*	55.74	*0.0000*	23.76	*0.0082*

Czech Republic - Total Outflows

Year	1992		1993		1994		1995	
Test	χ^2	*P-value*	χ^2	*P-value*	χ^2	*P-value*	χ^2	*P-value*
1.Hypothesis[1)] - $\chi^2(1)$	2.15	*0.1427*	0.01	*0.9301*	0.05	*0.8266*	1.31	*0.2527*
2.Hypothesis[2)] - $\chi^2(2)$	12.33	*0.0021*	2.15	*0.3415*	8.06	*0.0178*	4.49	*0.1059*
3.Hypothesis[3)] - $\chi^2(5)$	31.64	*0.0000*	14.98	*0.0105*	23.58	*0.0003*	31.71	*0.0000*
4.Hypothesis[4)] - $\chi^2(6)$	34.74	*0.0000*	27.95	*0.0001*	24.66	*0.0004*	33.35	*0.0000*
5.Hypothesis[5)] - $\chi^2(4)$	43.48	*0.0000*	57.76	*0.0000*	23.37	*0.0001*	39.64	*0.0000*
6.Hypothesis[6)] - $\chi^2(10)$	n.a.	*n.a.*	44.49	*0.0000*	46.70	*0.0000*	47.16	*0.0082*

Slovak Republic - Total Outflows

Year	1992		1993		1994		1995	
Test	χ^2	*P-value*	χ^2	*P-value*	χ^2	*P-value*	χ^2	*P-value*
1.Hypothesis[1)] - $\chi^2(1)$	0.16	*0.6860*	0.01	*0.9370*	0.03	*0.8586*	n.a.	*n.a.*
2.Hypothesis[2)] - $\chi^2(2)$	1.26	*0.5319*	15.96	*0.0003*	2.91	*0.2329*	n.a.	*n.a.*
3.Hypothesis[3)] - $\chi^2(5)$	12.04	*0.0343*	32.08	*0.0000*	3.30	*0.6539*	n.a.	*n.a.*
4.Hypothesis[4)] - $\chi^2(6)$	12.06	*0.0607*	32.12	*0.0000*	5.28	*0.5080*	n.a.	*n.a.*
5.Hypothesis[5)] - $\chi^2(4)$	10.18	*0.0375*	28.26	*0.0000*	3.95	*0.4128*	n.a.	*n.a.*
6.Hypothesis[6)] - $\chi^2(10)$	n.a.	*n.a.*	31.66	*0.0009*	27.89	*0.0034*	n.a.	*n.a.*

1) Weak separability of unemployment and inflow into unemployment

2) Strong separability of unemployment and inflow into unemployment

3) Cobb-Douglas specification

4) Cobb-Douglas specification with constant returns to scale

5) Translog specification with constant returns to scale

6) Stability of coefficients between neighboring years

Table 2: Estimated Elasticities and Returns to Scale of the Weakly Separable Translog Matching Functions

Czech Republic - Outflows to Jobs

Year	1992	1993	1994	1995
Elasticity	Estimate *(Std.Error)*	Estimate *(Std.Error)*	Estimate *(Std.Error)*	Estimate *(Std.Error)*
η_U	1.659	1.309	1.620	1.928
	(0.191)	*(0.156)*	*(0.165)*	*(0.145)*
η_S	0.886	0.756	0.535	0.769
	(0.136)	*(0.106)*	*(0.090)*	*(0.092)*
η_V	0.912	1.189	0.776	0.679
	(0.176)	*(0.107)*	*(0.108)*	*(0.114)*
RTS	3.457	3.255	2.930	3.376
	(0.353)	*(0.276)*	*(0.271)*	*(0.220)*

Czech Republic - Total Outflows

Year	1992	1993	1994	1995
Elasticity	Estimate *(Std.Error)*	Estimate *(Std.Error)*	Estimate *(Std.Error)*	Estimate *(Std.Error)*
η_U	1.226	0.973	1.505	1.755
	(0.150)	*(0.135)*	*(0.153)*	*(0.125)*
η_S	0.745	0.569	0.425	0.754
	(0.114)	*(0.085)*	*(0.084)*	*(0.081)*
η_V	0.651	0.997	0.833	0.669
	(0.139)	*(0.092)*	*(0.102)*	*(0.106)*
RTS	2.623	2.539	2.763	2.177
	(0.275)	*(0.237)*	*(0.253)*	*(0.195)*

Slovak Republic - Total Outflows

Year	1992	1993	1994	1995
Elasticity	Estimate *(Std.Error)*	Estimate *(Std.Error)*	Estimate *(Std.Error)*	Estimate *(Std.Error)*
η_U	0.358	1.047	2.503	n.a.
	(0.285)	*(0.533)*	*(0.877)*	*n.a.*
η_S	-0.014	0.212	0.115	n.a.
	(0.113)	*(0.122)*	*(0.193)*	*n.a.*
η_V	0.170	0.253	0.186	n.a.
	(0.078)	*(0.083)*	*(0.129)*	*n.a.*
RTS	0.514	1.512	2.803	n.a.
	(0.317)	*(0.635)*	*(1.029)*	*n.a.*

η_U *elasticity of the outflow with respect to the unemployment*

η_S *elasticity of the outflow with respect to the inflow into unemployment*

η_V *elasticity of the outflow with respect to the vacancies*

RTS *returns to scale*

Note: The elasticities and returns to scale are evaluated at the geometric means of the relevant variables.

Table 3: Second Stage AHIV Estimates of the Effects of Structural and Policy Variables on Matching

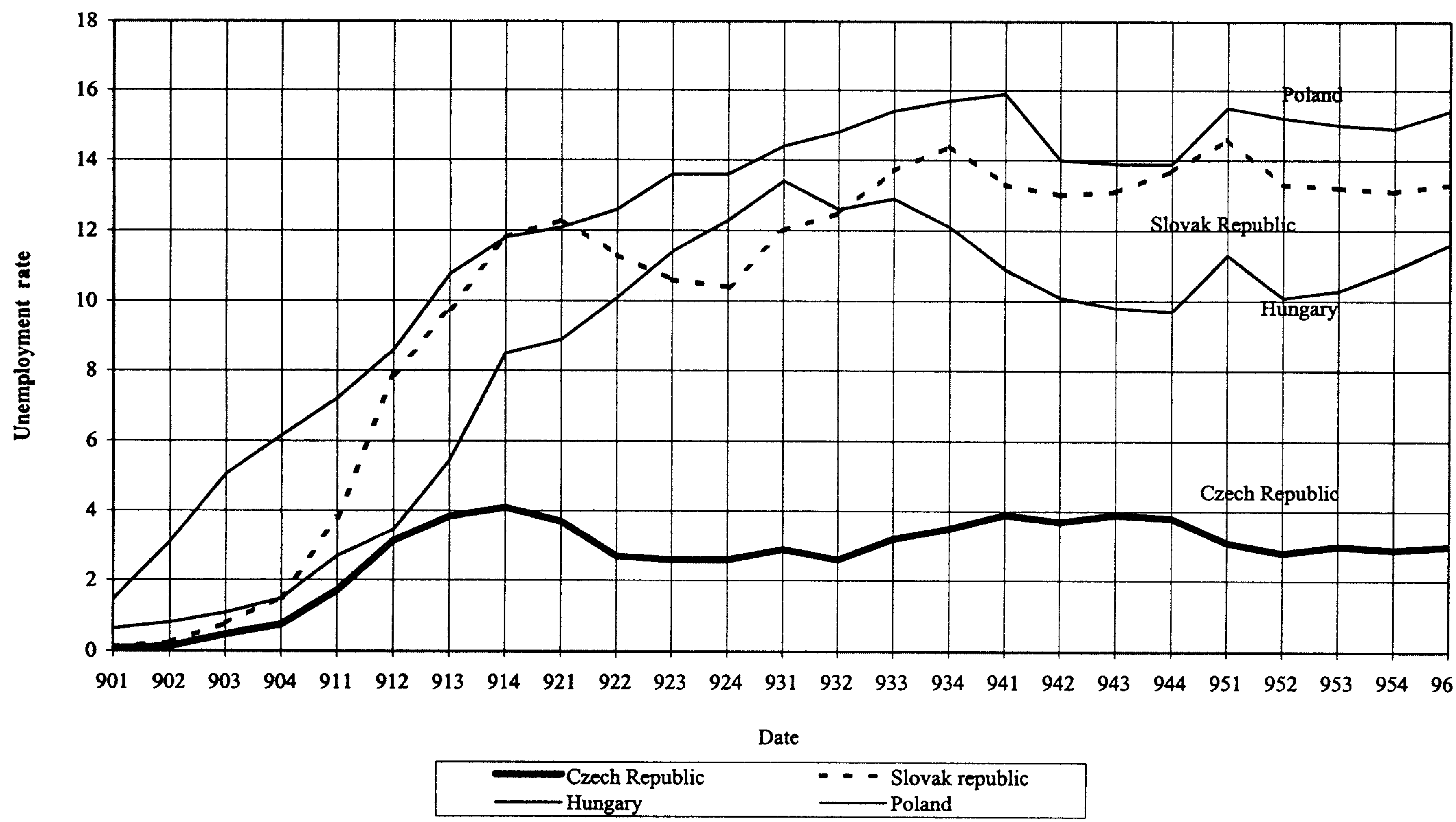
	Czech Republic				Slovak Republic			
1992	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*
C	103.98	109.14	109.11	109.44	25.75	18.92	19.24	19.76
	(3.56)	*(1.02)*	*(1.01)*	*(1.06)*	*(3.97)*	*(2.78)*	*(2.75)*	*(2.88)*
IA90	0.02	-0.12	-0.12	-0.07	2.35	-0.06	-0.45	-0.45
	(0.21)	*(0.11)*	*(0.11)*	*(0.11)*	*(1.55)*	*-(0.05)*	*-(0.36)*	*-(0.36)*
Q	-0.02	-	-	-	2.24	-	-	-
	(0.29)	-	-	-	*(2.36)*	-	-	-
E	-1.25	-0.71	-0.63	-0.06	1.64	1.97	0.75	0.92
	(0.85)	*(0.53)*	*(0.54)*	*(0.53)*	*(0.87)*	*(0.95)*	*(0.39)*	*(0.49)*
ALMP	-0.79	-	-	-	n.a.	n.a.	n.a.	n.a.
	(0.66)	-	-	-	*n.a.*	n.a.	n.a.	n.a.
PUED2	-4.16	-2.83	-1.00	2.23	2.64	4.59	2.20	1.01
	(3.80)	*(2.64)*	*(2.47)*	*(2.32)*	*(0.31)*	*(0.49)*	*(0.23)*	*(0.11)*
PUED3	19.59	28.03	27.11	6.02	-1.29	-7.31	-10.51	-1.21
	(14.75)	*(9.07)*	*(9.25)*	*(5.82)*	*-(0.05)*	*-(0.24)*	*-(0.34)*	*-(0.06)*
PLTU	-4.47	-7.59	-6.74	-4.76	-57.81	-60.83	-57.26	-57.12
	(4.29)	*(2.50)*	*(2.50)*	*(2.54)*	*-(7.61)*	*-(7.40)*	*-(7.08)*	*-(7.05)*
DENS	0.00	0.00	0.00	-	0.00	0.00	0.00	-
	(0.00)	*(0.00)*	*(0.00)*	-	*-(0.41)*	*(0.32)*	*(0.38)*	-
SUD	0.79	-0.42	-	-	n.a.	n.a.	n.a.	n.a.
	(0.47)	*(0.25)*	-	-	*n.a.*	*n.a.*	*n.a.*	*n.a.*
LDIS	-0.07	-	-	-	0.95	0.88	-	-
	(0.20)	-	-	-	*(1.65)*	*1.39*	-	-
adj.R^2	0.57	0.14	0.10	0.09	0.75	0.70	0.67	0.67

	Czech Republic				Slovak Republic			
1993	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*	**Coef.** *(Std.Error)*
C	4.47	13.33	13.11	13.09	-166.73	-163.94	-163.92	-162.82
	(1.22)	*(0.60)*	*(0.56)*	*(0.56)*	*-(35.25)*	*-(38.23)*	*-(39.30)*	*-(37.60)*
IA90	-0.12	-0.16	-0.16	-0.15	-1.95	-1.29	-1.30	-1.31
	(0.09)	*(0.07)*	*(0.08)*	*(0.08)*	*-(1.56)*	*-(1.25)*	*-(1.30)*	*-(1.24)*
Q	-0.12	-	-	-	-0.85	-	-	-
	(0.15)	-	-	-	*-(0.91)*	-	-	-
E	-0.46	-1.01	-1.07	-1.02	-0.49	-0.41	-0.42	-0.03
	(0.31)	*(0.28)*	*(0.29)*	*(0.29)*	*-(0.36)*	*-(0.29)*	*-(0.33)*	*-(0.02)*
ALMP	0.23	-	-	-	n.a.	n.a.	n.a.	n.a.
	(0.12)	-	-	-	*n.a.*	n.a.	n.a.	n.a.
PUED2	-2.51	-3.97	-2.62	-1.85	1.70	1.79	1.77	-2.47
	(1.88)	*(1.83)*	*(1.78)*	*(1.56)*	*(0.20)*	*(0.20)*	*(0.20)*	*-(0.28)*
PUED3	14.34	25.63	26.65	22.07	-121.77	-121.31	-121.54	-73.92
	(7.13)	*(6.56)*	*(6.76)*	*(4.37)*	*-(3.41)*	*-(3.35)*	*-(3.41)*	*-(3.24)*
PLTU	-2.52	-3.01	-2.51	-2.42	38.27	39.52	39.57	40.30
	(1.22)	*(1.21)*	*(1.23)*	*(1.23)*	*(6.56)*	*(6.86)*	*(7.14)*	*(6.93)*
DENS	0.00	0.00	0.00	-	0.01	0.00	0.00	-
	(0.00)	*(0.00)*	*(0.00)*	-	*(1.90)*	*(1.68)*	*(1.69)*	-
SUD	0.23	-0.35	-	-	n.a.	n.a.	n.a.	n.a.
	(0.18)	*(0.17)*	-	-	*n.a.*	*n.a.*	*n.a.*	*n.a.*
LDIS	-0.02	-	-	-	-0.03	0.02	-	-
	(0.09)	-	-	-	*-(0.07)*	*(0.03)*	-	-
adj.R^2	0.67	0.36	0.32	0.31	0.80	0.79	0.79	0.77

1994	Czech Republic				Slovak Republic			
	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)
C	45.97	49.72	49.73	49.77	-20.42	-20.53	-21.40	-19.53
	(1.09)	(0.76)	(0.73)	(0.72)	-(3.59)	-(3.60)	-(3.60)	-(3.82)
IA90	0.04	-0.04	-0.04	-0.04	0.11	0.05	0.17	0.20
	(0.06)	(0.06)	(0.06)	(0.06)	(0.21)	(0.11)	(0.38)	(0.43)
Q	0.03	-	-	-	0.09	-	-	-
	(0.11)	-	-	-	(0.23)	-	-	-
E	-0.17	0.70	0.70	0.72	0.22	0.21	0.64	0.52
	(0.37)	(0.36)	(0.36)	(0.35)	(0.37)	(0.36)	(1.14)	(0.99)
ALMP	-0.33	-	-	-	n.a.	n.a.	n.a.	n.a.
	(0.07)	-	-	-	n.a.	n.a.	n.a.	n.a.
PUED2	-1.05	-0.44	-0.46	-0.26	-0.67	-0.68	1.13	1.60
	(1.34)	(1.52)	(1.36)	(1.24)	-(0.07)	-(0.07)	(0.12)	(0.17)
PUED3	22.52	22.91	22.89	21.22	74.54	73.73	83.92	67.08
	(5.84)	(6.35)	(6.32)	(4.14)	(2.12)	(2.10)	(2.32)	(2.86)
PLTU	-2.79	-3.93	-3.93	-3.97	-32.07	-32.28	-33.61	-34.66
	(0.87)	(0.95)	(0.93)	(0.92)	-(6.37)	-(6.53)	-(6.56)	-(7.14)
DENS	0.00	0.00	0.00	-	0.00	0.00	0.00	-
	(0.00)	(0.00)	(0.00)	-	-(0.50)	-(0.44)	-(0.61)	-
SUD	0.11	0.00	-	-	n.a.	n.a.	n.a.	n.a.
	(0.14)	(0.14)	-	-	n.a.	n.a.	n.a.	n.a.
LDIS	0.06	-	-	-	-0.32	-0.32	-	-
	(0.07)	-	-	-	-(1.62)	-(1.66)	-	-
adj.R^2	0.42	0.56	0.56	0.56	0.83	0.83	0.82	0.81

1995	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)	Coef. (Std.Error)
C	30.35	31.52	31.50	31.73	n.a.	n.a.	n.a.	n.a.
	(1.19)	(0.75)	(0.73)	(0.74)	n.a.	n.a.	n.a.	n.a.
IA90	0.01	-0.10	-0.10	-0.10	n.a.	n.a.	n.a.	n.a.
	(0.07)	(0.07)	(0.07)	(0.08)	n.a.	n.a.	n.a.	n.a.
Q	0.07	-	-	-	n.a.	n.a.	n.a.	n.a.
	(0.12)	-	-	-	n.a.	n.a.	n.a.	n.a.
E	0.25	1.18	1.19	1.37	n.a.	n.a.	n.a.	n.a.
	(0.37)	(0.38)	(0.38)	(0.38)	n.a.	n.a.	n.a.	n.a.
ALMP	-0.38	-	-	-	n.a.	n.a.	n.a.	n.a.
	(0.07)	-	-	-	n.a.	n.a.	n.a.	n.a.
PUED2	-3.17	-0.58	-0.14	1.47	n.a.	n.a.	n.a.	n.a.
	(1.54)	(1.96)	(1.81)	(1.70)	n.a.	n.a.	n.a.	n.a.
PUED3	30.63	27.95	28.20	16.32	n.a.	n.a.	n.a.	n.a.
	(6.32)	(7.98)	(7.98)	(5.82)	n.a.	n.a.	n.a.	n.a.
PLTU	-3.40	-4.62	-4.54	-4.77	n.a.	n.a.	n.a.	n.a.
	(0.80)	(1.02)	(1.01)	(1.04)	n.a.	n.a.	n.a.	n.a.
DENS	0.00	0.00	0.00	-	n.a.	n.a.	n.a.	n.a.
	(0.00)	(0.00)	(0.00)	-	n.a.	n.a.	n.a.	n.a.
SUD	0.36	-0.10	-	-	n.a.	n.a.	n.a.	n.a.
	(0.15)	(0.17)	-	-	n.a.	n.a.	n.a.	n.a.
LDIS	0.04	-	-	-	n.a.	n.a.	n.a.	n.a.
	(0.08)	-	-	-	n.a.	n.a.	n.a.	n.a.
adj.R^2	0.77	0.61	0.61	0.58	n.a.	n.a.	n.a.	n.a.

Figure 1
Unemployment Rates in Central Europe



Source: 1990-1991 Unemployment in Transition Countries: Transient or Persistent?, OECD 1994, pg.29
1992-1996 Statistical Buletins Of the Czech Statistical Office, '93, '93,'95,'96/1

Figure 2
Unemployment and Vacancies in the Czech and Slovak Republic

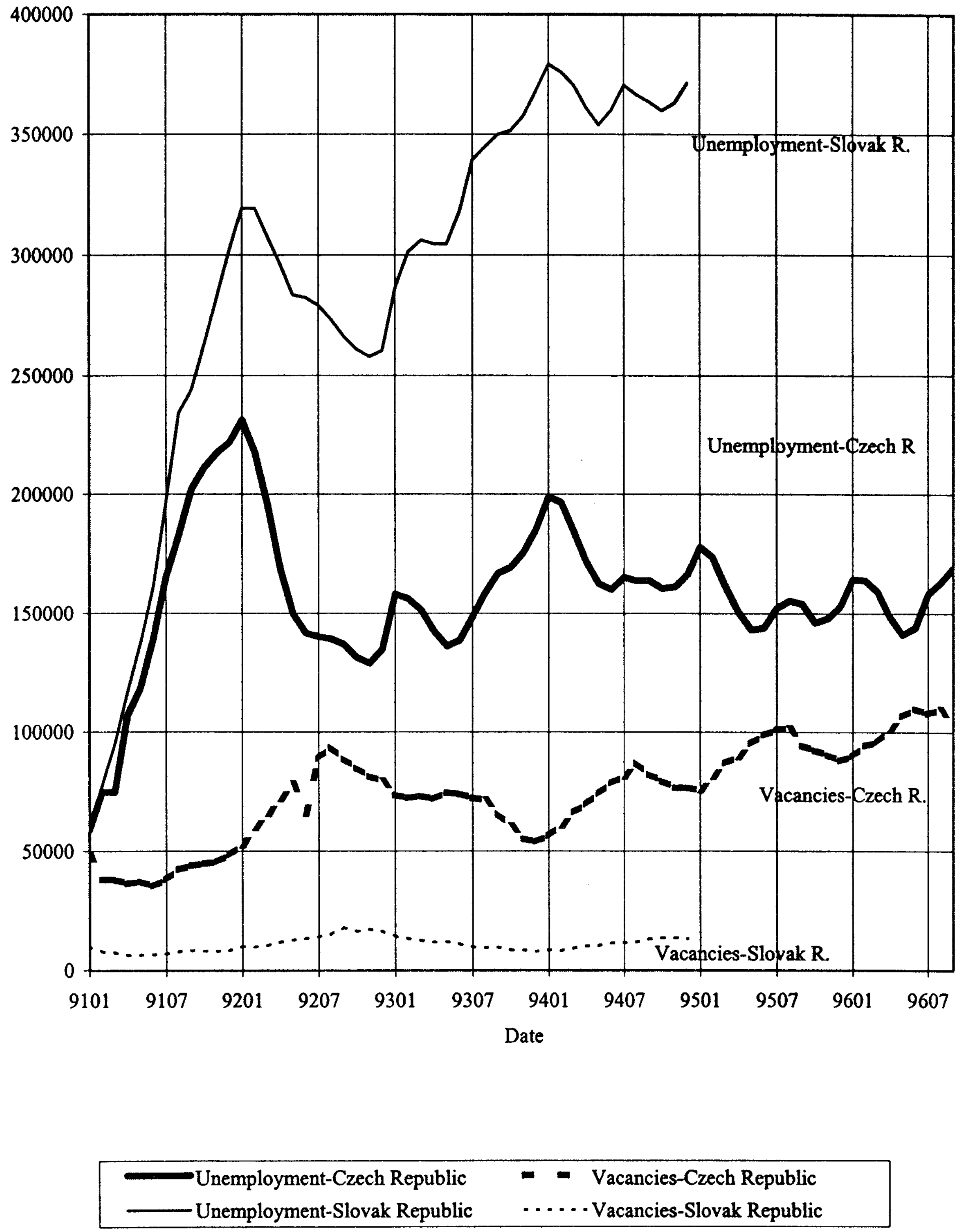
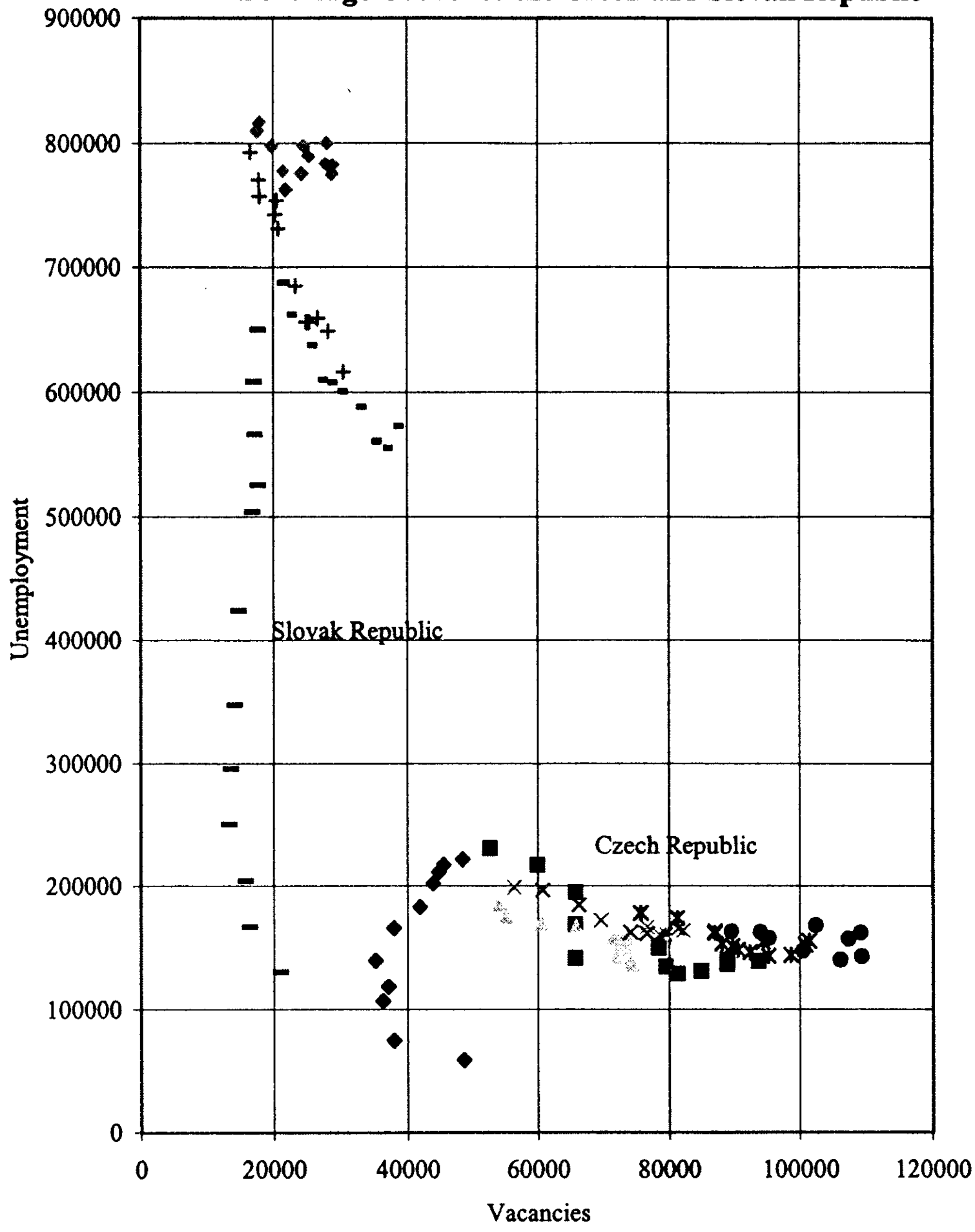


Figure 3
Beveridge Curve for the Czech and Slovak Republic



Czech Republic (CR), Slovak Republic (SR)

◆ U(CR)'91 ■ U(CR)'92 U(CR)'93 × U(CR)'94 ✱ U(CR)'95
● U(CR)'96 − U(SR)'91 - U(SR)'92 + U(SR)'93 U(SR)'94

Figure 4
Inflows into Unemployment in the Czech and Slovak Republic

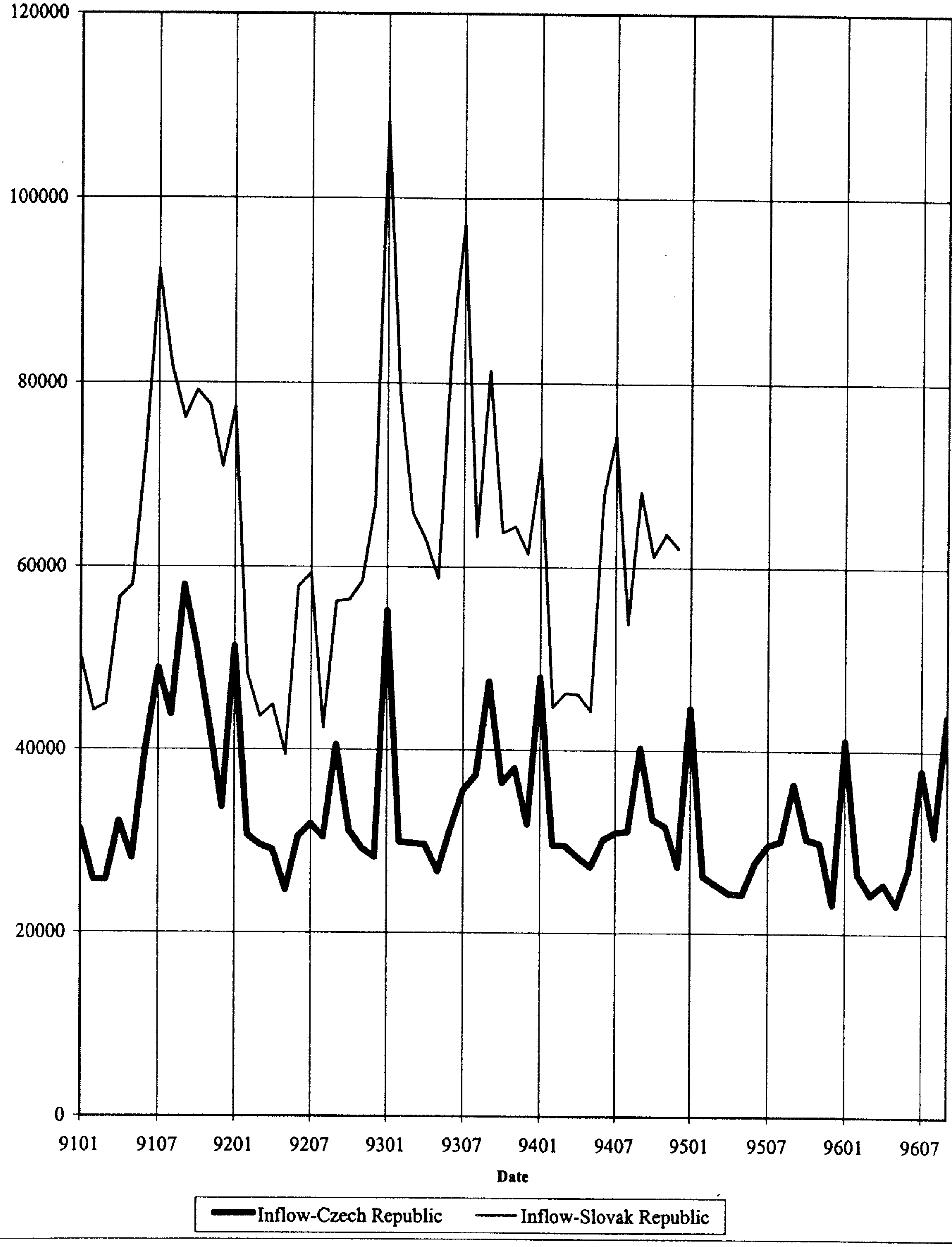


Figure 5
Outflows from Unemployment in the Czech and Slovak Republic

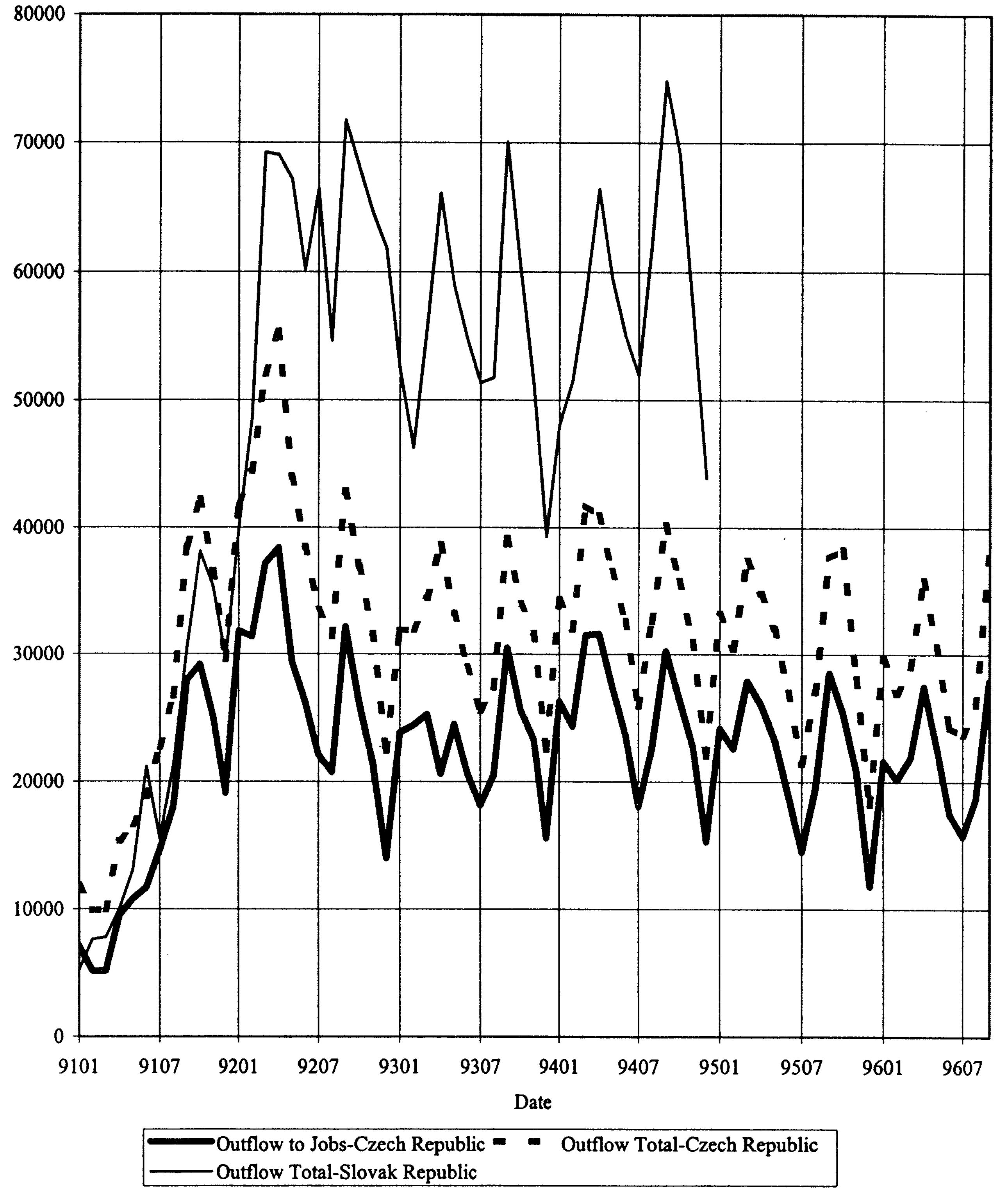


Table A1: Summary Statistics

Czech Republic

Year	Variable	Mean[a)]	Std. Dev.[a)]	Mean[b)]	Std. Dev.[b)]
1991	Inflow	520.2	224.8	506.2	413.7
	Outflow	203.6	151.3	201.1	228.8
	Unemployment	1988.2	1136.0	1945.2	1759.6
	Vacancies	462.9	291.1	543.4	1046.8
1992	Inflow	460.8	199.6	425.2	314.4
	Outflow	397.0	204.5	362.9	279.0
	Unemployment	2340.6	1158.2	2100.4	1506.2
	Vacancies	944.7	510.8	993.3	1894.1
1993	Inflow	513.6	235.1	471.3	353.7
	Outflow	343.4	161.7	308.2	211.2
	Unemployment	2292.4	1157.7	2069.9	1590.5
	Vacancies	866.5	460.1	898.6	1920.8
1994	Inflow	454.6	207.9	424.3	306.5
	Outflow	362.5	182.9	329.8	228.2
	Unemployment	2424.5	1283.8	2253.7	1815.7
	Vacancies	946.4	401.9	974.6	1704.4
1995	Inflow	409.9	202.4	386.2	286.7
	Outflow	313.0	166.2	288.9	210.9
	Unemployment	2175.5	1198.2	2039.6	1660.3
	Vacancies	1213.7	443.9	1200.6	1598.6
1996	Inflow	431.9	216.7	406.1	305.4
	Outflow	306.5	153.3	280.2	189.7
	Unemployment	2200.4	1214.5	2046.5	1610.5
	Vacancies	1380.0	519.8	1335.4	1651.4

Slovak Republic

Year	Variable	Mean[a)]	Std. Dev.[a)]	Mean[b)]	Std. Dev.[b)]
1991	Inflow	848.5	352.0	819.9	466.1
	Outflow	241.4	214.8	239.4	278.3
	Unemployment	4935.0	2760.4	4761.4	3218.8
	Vacancies	186.9	154.0	198.1	258.0
1992	Inflow	692.8	250.0	663.4	341.2
	Outflow	757.8	296.4	754.8	516.1
	Unemployment	7892.5	1941.0	7467.8	3040.0
	Vacancies	305.4	249.5	358.9	604.7
1993	Inflow	952.5	303.8	906.6	416.1
	Outflow	692.0	217.6	670.1	329.7
	Unemployment	9314.2	2618.1	8623.9	3058.7
	Vacancies	250.1	179.7	277.9	357.8
1994	Inflow	760.9	254.3	717.8	313.1
	Outflow	740.5	219.9	710.4	327.9
	Unemployment	10567.2	3110.0	9639.7	3194.3
	Vacancies	275.8	205.2	291.6	319.4

a) Data adjusted for district size
b) Unadjusted data

Table A2: Correlation Coefficients for the Czech Republic

Variables Unadjusted for District Size

Year	Variable	Inflow	Outflow	Unempl.
1991	Outflow	0.83		
	Unemployment	0.87	0.81	
	Vacancies	0.70	0.68	0.61
1992	Outflow	0.79		
	Unemployment	0.82	0.77	
	Vacancies	0.34	0.37	0.14
1993	Outflow	0.83		
	Unemployment	0.86	0.78	
	Vacancies	0.25	0.29	0.03
1994	Outflow	0.86		
	Unemployment	0.90	0.85	
	Vacancies	0.25	0.19	0.03
1995	Outflow	0.85		
	Unemployment	0.89	0.84	
	Vacancies	0.29	0.19	0.11
1996	Outflow	0.84		
	Unemployment	0.88	0.85	
	Vacancies	0.28	0.22	0.12

Variables Adjusted for District Size

Year	Variable	Inflow	Outflow	Unempl.
1991	Outflow	0.65		
	Unemployment	0.71	0.66	
	Vacancies	-0.09	0.03	-0.18
1992	Outflow	0.54		
	Unemployment	0.64	0.60	
	Vacancies	-0.30	-0.29	-0.52
1993	Outflow	0.66		
	Unemployment	0.80	0.70	
	Vacancies	-0.46	-0.40	-0.58
1994	Outflow	0.70		
	Unemployment	0.84	0.74	
	Vacancies	-0.42	-0.37	-0.50
1995	Outflow	0.71		
	Unemployment	0.83	0.73	
	Vacancies	-0.27	-0.22	-0.29
1996	Outflow	0.68		
	Unemployment	0.82	0.74	
	Vacancies	-0.20	-0.19	-0.21

Table A3: Correlation Coefficients for the Slovak Republic

Variables Unadjusted for District Size

Year	Variable	Inflow	Outflow	Unempl.
1991	Outflow	0.51		
	Unemployment	0.76	0.61	
	Vacancies	0.41	0.45	0.41
1992	Outflow	0.56		
	Unemployment	0.68	0.57	
	Vacancies	0.52	0.59	0.49
1993	Outflow	0.66		
	Unemployment	0.63	0.60	
	Vacancies	0.52	0.58	0.36
1994	Outflow	0.67		
	Unemployment	0.65	0.57	
	Vacancies	0.53	0.61	0.30

Variables Adjusted for District Size

Year	Variable	Inflow	Outflow	Unempl.
1991	Outflow	0.36		
	Unemployment	0.67	0.52	
	Vacancies	-0.14	0.08	-0.09
1992	Outflow	0.10		
	Unemployment	0.29	0.01	
	Vacancies	-0.07	0.13	-0.32
1993	Outflow	0.20		
	Unemployment	0.30	0.11	
	Vacancies	0.00	0.04	-0.24
1994	Outflow	0.18		
	Unemployment	0.41	0.23	
	Vacancies	0.02	0.08	-0.24

Table A4: Estimated Coefficients of Unconstrained Translog Matching Function (1) and with Weak Separability Between the Existing and Newly Unemployed Individuals (2)

Czech Republic - Outflows to Jobs

Year	1992				1993				1994				1995			
Model	(1)		(2)		(1)		(2)		(1)		(2)		(1)		(2)	
Variable	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*
Trend	0,061	*0,024*	0,057	*0,023*	0,052	*0,022*	0,056	*0,020*	0,117	*0,031*	0,107	*0,029*	0,108	*0,044*	0,074	*0,037*
γ_0	0,319	*0,052*	0,320	*0,051*	0,297	*0,056*	0,303	*0,053*	0,267	*0,045*	0,244	*0,041*	0,410	*0,050*	0,345	*0,040*
β_U	-10,004	*4,474*	0,588	*0,904*	11,350	*2,862*	3,430	*2,156*	8,259	*2,967*	7,110	*2,852*	10,035	*3,159*	8,850	*2,780*
β_S	18,851	*6,655*	22,376	*5,955*	-6,087	*3,909*	1,057	*3,031*	8,056	*6,495*	1,993	*2,289*	9,931	*4,271*	-1,452	*1,132*
β_V	-2,783	*4,288*	6,515	*1,729*	3,775	*3,697*	-0,198	*3,392*	12,325	*3,466*	11,444	*3,272*	4,246	*3,092*	7,202	*2,590*
γ_{US}	0,041	*0,009*	0,041	*0,009*	0,031	*0,008*	0,035	*0,008*	0,026	*0,009*	0,022	*0,008*	0,026	*0,010*	0,016	*0,009*
γ_{UV}	-1,248	*0,512*	0,011	*0,017*	1,031	*0,367*	0,067	*0,276*	0,854	*0,320*	0,795	*0,315*	0,109	*0,378*	0,484	*0,316*
γ_{SV}	0,301	*0,230*	0,411	*0,206*	-0,424	*0,254*	0,021	*0,101*	0,317	*0,305*	0,222	*0,259*	-0,128	*0,076*	-0,079	*0,066*
γ_{UU}	-1,732	*0,623*	-0,407	*0,250*	1,349	*0,363*	0,428	*0,294*	0,754	*0,476*	0,523	*0,448*	2,005	*0,582*	1,328	*0,469*
γ_{SS}	2,047	*0,719*	2,424	*0,651*	-0,623	*0,389*	0,010	*0,338*	0,740	*0,679*	0,051	*0,147*	1,171	*0,521*	-0,241	*0,110*
γ_{VV}	-0,362	*0,384*	0,504	*0,223*	0,483	*0,482*	-0,394	*0,417*	1,367	*0,399*	1,388	*0,392*	1,008	*0,616*	1,321	*0,539*
adj.R^2	0,172		0,184		0,148		0,152		0,150		0,166		0,200		0,260	
Nobs	912		912		912		912		912		912		912		912	

Slovak Republic - Total Outflows

Year	1992				1993				1994				1995			
Model	(1)		(2)		(1)		(2)		(1)		(2)		(1)		(2)	
Variable	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*
Trend	0,010	*0,070*	0,023	*0,065*	-0,264	*0,099*	-0,132	*0,074*	-0,202	*0,116*	-0,186	*0,075*	n.a.	*n.a.*	n.a.	*n.a.*
γ_0	0,380	*0,050*	0,377	*0,049*	0,173	*0,070*	0,163	*0,061*	0,309	*0,092*	0,308	*0,090*	n.a.	*n.a.*	n.a.	*n.a.*
β_U	-2,630	*6,805*	0,446	*2,680*	0,525	*6,871*	5,208	*5,816*	-12,186	*12,330*	-11,700	*11,780*	n.a.	*n.a.*	n.a.	*n.a.*
β_S	-10,392	*4,101*	-9,973	*4,005*	-21,169	*4,823*	-21,980	*4,225*	-4,368	*4,700*	-4,827	*3,856*	n.a.	*n.a.*	n.a.	*n.a.*
β_V	0,505	*2,133*	1,309	*1,367*	-13,141	*3,613*	-6,001	*2,033*	-5,980	*3,669*	-5,604	*2,944*	n.a.	*n.a.*	n.a.	*n.a.*
γ_{US}	0,000	*0,013*	-0,002	*0,012*	0,047	*0,018*	0,025	*0,013*	0,022	*0,019*	0,020	*0,011*	n.a.	*n.a.*	n.a.	*n.a.*
γ_{UV}	-0,225	*0,444*	-0,008	*0,047*	-0,734	*0,436*	0,180	*0,215*	-0,830	*0,688*	-0,839	*0,672*	n.a.	*n.a.*	n.a.	*n.a.*
γ_{SV}	0,149	*0,169*	0,168	*0,164*	-1,093	*0,275*	-0,791	*0,219*	-0,397	*0,353*	-0,346	*0,205*	n.a.	*n.a.*	n.a.	*n.a.*
γ_{UU}	-0,777	*2,154*	0,065	*1,307*	1,417	*2,183*	1,263	*1,915*	-5,045	*4,630*	-4,744	*4,222*	n.a.	*n.a.*	n.a.	*n.a.*
γ_{SS}	-1,431	*0,489*	-1,391	*0,481*	-1,979	*0,544*	-2,298	*0,466*	-0,275	*0,660*	-0,372	*0,368*	n.a.	*n.a.*	n.a.	*n.a.*
γ_{VV}	-0,061	*0,113*	-0,029	*0,092*	-0,591	*0,239*	-0,098	*0,124*	-0,239	*0,217*	-0,237	*0,212*	n.a.	*n.a.*	n.a.	*n.a.*
adj.R^2	0,212		0,213		0,035		0,055		0,019		0,030		n.a.		n.a.	
Nobs	456		456		456		456		190		190		n.a.		n.a.	

Czech Republic - Total Outflows

Year	1992				1993				1994				1995			
Model	(1)		(2)		(1)		(2)		(1)		(2)		(1)		(2)	
Variable	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*	Estimate	*Std.Error*
Trend	0.062	*0.019*	0.058	*0.018*	0.039	*0.017*	0.041	*0.017*	0.109	*0.028*	0.111	*0.027*	0.079	*0.039*	0.049	*0.034*
γ_O	0.407	*0.047*	0.398	*0.045*	0.291	*0.051*	0.295	*0.050*	0.196	*0.045*	0.200	*0.042*	0.448	*0.046*	0.390	*0.037*
β_U	4.395	*3.749*	0.354	*0.770*	5.711	*2.401*	4.220	*2.310*	5.549	*2.705*	5.660	*2.660*	9.132	*2.778*	8.539	*2.520*
β_S	15.693	*5.328*	19.134	*4.949*	-0.477	*2.982*	0.695	*2.803*	1.495	*5.664*	2.704	*1.814*	7.254	*3.591*	-1.711	*0.903*
β_V	-0.449	*3.516*	6.389	*1.352*	2.388	*2.968*	1.874	*2.965*	11.634	*3.201*	11.891	*2.993*	5.059	*2.582*	7.015	*2.253*
γ_{US}	0.032	*0.007*	0.030	*0.007*	0.023	*0.007*	0.024	*0.006*	0.022	*0.008*	0.023	*0.007*	0.017	*0.009*	0.008	*0.008*
γ_{UV}	-0.904	*0.432*	0.009	*0.020*	0.445	*0.308*	0.285	*0.300*	0.700	*0.289*	0.704	*0.289*	0.224	*0.330*	0.539	*0.280*
γ_{SV}	0.343	*0.182*	0.479	*0.166*	-0.023	*0.199*	0.047	*0.184*	0.305	*0.270*	0.336	*0.232*	-0.132	*0.064*	-0.108	*0.058*
γ_{UU}	-1.331	*0.521*	-0.324	*0.207*	0.640	*0.307*	0.490	*0.293*	0.228	*0.438*	0.252	*0.425*	1.711	*0.510*	1.233	*0.434*
γ_{SS}	1.658	*0.578*	2.006	*0.540*	-0.127	*0.299*	-0.022	*0.247*	-0.042	*0.595*	0.090	*0.108*	0.851	*0.436*	-0.253	*0.087*
γ_{VV}	-0.148	*0.321*	0.413	*0.176*	-0.008	*0.382*	-0.115	*0.379*	1.341	*0.356*	1.341	*0.356*	1.120	*0.525*	1.287	*0.475*
adj.R^2	0.209		0.221		0.161		0.139		0.165		0.165		0.242		0.298	
Nobs	912		912		912		912		912		912		912		912	

Table A5: Estimated Elasticities and Returns to Scale of the Unconstrained Translog Matching Functions

Czech Republic - Outflows to Jobs

Year	1992	1993	1994	1995
Elasticity	Estimate *(Std.Error)*	Estimate *(Std.Error)*	Estimate *(Std.Error)*	Estimate *(Std.Error)*
η_u	1.508	1.457	1.636	2.104
	(0.213)	*(0.163)*	*(0.170)*	*(0.174)*
η_s	0.791	0.728	0.602	0.872
	(0.151)	*(0.112)*	*(0.108)*	*(0.109)*
η_v	0.960	1.222	0.744	0.758
	(0.189)	*(0.111)*	*(0.114)*	*(0.131)*
RTS	3.260	3.407	2.982	3.734
	(0.386)	*(0.289)*	*(0.282)*	*(0.277)*

Slovak Republic - Total Outflows

Year	1992	1993	1994	1995
Elasticity	Estimate *(Std.Error)*	Estimate *(Std.Error)*	Estimate *(Std.Error)*	Estimate *(Std.Error)*
η_u	0.424	1.386	2.555	n.a.
	(0.424)	*(1.386)*	*(2.555)*	*n.a.*
η_s	-0.016	0.284	0.106	n.a.
	-(0.016)	*(0.284)*	*(0.106)*	*n.a.*
η_v	0.193	0.272	0.201	n.a.
	(0.193)	*(0.272)*	*(0.201)*	*n.a.*
RTS	0.600	1.942	2.862	n.a.
	(0.600)	*(1.942)*	*(2.862)*	*n.a.*

η_u *elasticity of the outflow with respect to the unemployment*

η_s *elasticity of the outflow with respect to the inflow into unemployment*

η_v *elasticity of the outflow with respect to the vacancies*

RTS *returns to scale*

Note: The elasticities and returns to scale are evaluated at the geometric means of the relevant variables.