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THE MATCHING PROCESS IN LABOUR MARKETS IN TRANSITION

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First draft

Abstract

This paper uses the flow approach to investigate the way in which the transition of the Czech and Slovak economies influences the regional labour markets in the two countries. From estimates of matching functions we conclude that in most regions the labour markets are still in transition. We also examine the role of structural shocks as sources of shifts in unemployment-vacancy relationship in the Czech Republic. The evidence in support of structural hypothesis appears to be sensitive to the level of aggregation. The results are reversed when proceed to finer levels of disaggregation.

JEL-code: J63, J64
Keywords: matching function, unemployment, vacancies, transition

Part of the research of M. Lubyova was realized while staying at the Institute for Advanced Studies (IHS) in Vienna.

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Introduction.

The unemployment which has emerged in the Eastern Europe only recently and which has risen dramatically over two years horizon to two-digit figures is one of the most urgent problems of transitional economies. Especially annoying is the growth of long-term unemployment fraction and persistently high unemployment rates. Despite the common economic background with the other Eastern European countries, the Czech unemployment rate of 2-3 percent is much lower than that of many developed Western states. The puzzle of extremely low unemployment in the Czech Republic still remains unexplained.

Analysis of the unemployment data should help to explain the nature of transitional unemployment; mainly, to what extent it is structural phenomenon (due to the regional disparities, weak labour mobility, housing constraints, structural mismatch brought about by changes in the industrial structure, the transition of labour force from state to private sector, etc.), and how much of it is "real" unemployment (brought about by the decline in aggregate demand, breakdown of CMEA markets, etc.). These conclusions would have interesting policy implications for design of unemployment reducing strategies.

This paper is organized as follows:

In the first section we give a brief survey of theoretical approaches and recent work in the field of matching process. In the second section we provide in a descriptive way stylized facts on development and dynamics of unemployment in the Czech and Slovak Republics since the beginning of transition. The description is documented by set of charts included in Appendix 3. Section 3 contains description of the data set used in our estimation. In the fourth section we propose a theoretical specification of the matching function applicable to the Czech and Slovak labour markets and we estimate it. Section 5 is dedicated to the so-called Beveridge curve. We examine the role of structural changes in the outward shift of empirical UV relationship in the Czech Republic. In Section 6 we briefly conclude.

1. Recent Theoretical Approaches - A Brief Survey.

Many conceptual models of labour markets make heavy use of a matching function, the function that relates the flow of new hires to the stocks of vacancies and unemployment. Like the aggregate production function, the aggregate matching function is an abstract but convenient device, which partially captures a complex reality. The reality is one of workers looking for the right job, and of jobs looking for the right worker, with varying degrees of intensity and success. Changes in the nature of new jobs, in the location of job creation and job destruction, and in the search behaviour of the unemployed will all shift the matching function. The issue of matching function, its nature and characteristics have attracted the attention of many authors in the recent economic research.

Some attempts have been made to incorporate matching processes into more general macroeconomic framework. Attention of some authors has been focused on the so-called Beveridge curve which is closely related to the Phillips curve. As suggested by Blanchard and Diamond (1989), there may be a way of combining information obtained from the two that could be helpful in understanding labour market processes in macroeconomic relations.

In the recent literature labour market performance is investigated usually in terms of matching unemployed and vacancies via the so-called matching technology. It is possible to obtain direct evidence on the effectiveness of job search by examining the movement of unemployment relative to the level of vacancies. This is because there is a "hiring" (or "matching") function which explains the flow of unemployed people into work. This flow depends positively on the number of vacancies and also on the number of effective job-seekers (which can be obtained by multiplying the unemployment figure by effectiveness of job search. In many studies the matching function is assumed to be homogeneous of degree one in both inputs (vacancies and effective unemployment), which is equivalent to the constant returns-to-scale. Using the fact that in equilibrium the flows into and from unemployment are equal, we can infer the relationship between the number of vacancies and effective unemployed that delivers the famous Beveridge curve, i.e. a hyperbolical relationship in the unemployment-vacancies space. The empirical work on matching technology mostly builds on the simple Cobb-Douglas specification in which the number of matches which originate in a short time period depends on the numbers of job seekers and vacancies at the beginning of that period:

$$M_t=\lambda U_t^{\alpha} V_t^{\beta};0<\alpha<1;0<\beta<1 \quad (1)$$

where: M_t = flow of matches in period t

U_t = number of unemployed at the beginning of period t

V_t = number of vacancies at the beginning of period t

α,β,λ = parameters of the matching function

The matching function itself is a black box which incorporates information aspects, search behaviour, structural mismatch, etc. Parameter λ captures efficiency of labour market: the larger λ, the larger the flow of matches from given stocks of unemployed and vacancies.

Some empirical facts about matching and unemployment-vacancy relationships in Western Europe over the last twenty years are provided by Layard, Nickell, and Jackman (1991). The empirical relationships for sixteen developed Western countries indeed exhibit hyperbolical shape. However, there was an obvious upward shift in the Beveridge curve over the last twenty years in almost all investigated countries. The explanation would have to be either a fall in search effectiveness among the unemployed or an increase in "mismatch" between the pattern of unemployment and vacancies across sectors (i.e. regions, industries, or skill-groups). However, if the relevant mismatch indices are computed, it turns out that they have not risen at all since the early 1970s in most European countries. Therefore we come back to search effectiveness. This reflects not only how hard the workers look for work, but also how willing the employers are to consider them.

There are two directly measurable factors that affect search effectiveness. The first is the benefit-income (replacement) ratio. The empirical studies typically suggest that the elasticity of exit rates from unemployment with respect to the replacement ratio are of the order of 0.2 - 0.9. The increase in the replacement ratios in most European countries in the 1960s or 1970s may have had some lagged effect on unemployment but can explain only a fraction of the increase in unemployment in the 1970s, and none of that thereafter. It is possible that the absolute real value of benefits also has an effect (e.g. that the relevant replacement ratio relates to incomes above some subsistence level).

The second factor is how long people have been unemployed. Empirical results from Britain suggest that exit rates are much lower for the long-term unemployed. One possible explanation for this must be that the more effective job-seekers find jobs first, so that the long-term unemployed include a higher proportion of less effective job-seekers. Therefore the search effectiveness does not necessarily have to decline for a given individual over time, it can be just the change in the composition of the sample which delivers the above mentioned result. Layard, Nickell, and Jackman argue that the time-series evidence makes it clear that there is direct effect of unemployment duration on given individual. Long-term unemployment both demoralizes the individual and is also used by employers as a (biased) screening device.

A particular approach to the theory of unemployment was developed by Pissarides (1990). The central theme is the interaction between unemployment transitions and macroeconomic equilibrium. The equilibrium unemployment stock is derived from the transitions in and out of unemployment, which can be influenced both by aggregate events and by individual decisions. The transition out of unemployment is modeled as a trading process, with unemployed workers and firms with vacancies wanting to trade labour services. The theory is an "equilibrium" theory in the sense that firms and workers maximize their payoffs under rational expectations, given the stochastic process that breaks up jobs and the one that leads to the formation of new jobs. Furthermore, the wages are determined so as to exploit all the private gains from trade available to those who set them. The latter requirement is critical in the efficiency analysis of equilibrium. I general, the assumptions that give rise to unemployment in the model imply that the interests of the employed and the unemployed diverge. Therefore, if wages are determined at the level of the firm, where only employed workers are present, equilibrium will not in general be efficient. The model is a "macroeconomic" one in the sense that it is in terms of two representative agents, firms and workers, who move between two states, employment and unemployment, and aims to derive an equilibrium unemployment rate for the economy as a whole. The assumptions which ensure that the equilibrium unemployment rate is unique are natural extensions of those underlying the neoclassical model of growth.

The emphasis on the neoclassical model of balanced growth is driven by the belief that if a model of unemployment conforms to the long-run equilibrium requirements of balanced growth, it is the most suitable starting point for the extensions that will eventually explain unemployment in real economies. The basic model is shown to imply that in steady-state equilibrium there is a downward-sloping and convex-to-the-origin curve in vacancy-unemployment space - the Beveridge curve. When the steady state is disturbed, then stable adjustment path traces an anticlockwise loop around the curve. The model also implies the existence of a "long-run Phillips curve" in inflation-unemployment space, with a negative slope if there is an inflation-tax effect that affects the equilibrium real interest rate.

The extended model introduces plausible shift variables in the Beveridge curve. The two extensions are endogenous search intensity and job-specific variations in productivities that lead to the possibility of job rejection.

The implications of exogenous structural change and costly exchange in the model reach beyond the trivial ones of the existence of some "frictional" unemployment in otherwise conventional models. Since Walrasian and traditional Keynesian theory ignores trade, not much can be learned from them about its applications for macroeconomic equilibrium. The approach taken by Pissarides leads to the view that the decomposition of unemployment into frictional, cyclical, voluntary, involuntary, and so on is unhelpful in the theoretical and empirical analysis of unemployment.

A considerable amount of research on the Beveridge curve and matching processes in the U.S. was done by Blanchard and Diamond (1989). They believe that the model of labour market with large numbers of jobs being constantly created and destroyed is more appropriate for macroeconomic modelling than a representative agent model, or a centralized bargaining model with immobile labour.

Although jobs and workers are efficiently matched in the model, unemployment and vacancies coexist because of the sheer volume of jobs being created and destroyed. Blanchard and Diamond found that flows of newly hired workers depend both on unemployment and vacancies. The co-movement of unemployment and vacancies during the postwar period shows that short- and medium-term fluctuations in unemployment have been due mainly to aggregate activity shocks, that lead to both more (less) job creation and less (more) job destruction, rather than to changes in the degree in reallocation intensity, which lead to parallel movements in job creation and job destruction.

The issues of "equilibrium unemployment" are not dealt with in the model, since as a formalization strategy, the potential level of activity is taken as given. This depends in turn on wages. Thus, the next step in building a model of equilibrium unemployment is to draw out the implications of the approach for wage determination. The model so far has an implicit theory of wage determination. Wages are set to allow match to happen if it is mutually profitable. The surplus from the match between a worker and a firm is shared in some proportion (Nash bargaining assumption). This assumption implies that wages depend on the ratio of vacancies to unemployment, which reflects the relative bargaining strengths of workers and firms. The higher are vacancies given unemployment, the higher the wage; the higher is unemployment given vacancies, the lower the wage. The model suggests an integrated way of thinking about the Phillips curve and the Beveridge curve and the way to learn about the workings of the labour market and the source of the shocks from a joint examination of unemployment, vacancies, and wages. This is a approach explored by Solow in his 1964 Wicksell lectures, or, more recently, by Layard and Nickell in their analysis of unemployment in the United Kingdom.

A more ambitious task would be to consider wage-setting mechanism that sometimes prevents beneficial matching from taking place (such as present wages or union-negotiated wages). Blanchard and Diamond are challenged by the task of combining centralized wage setting and wage drift with description of the labour market which allows for the large flows of workers.

There is some previous research on unemployment and vacancies in the Czech and Slovak Republics. Burda (1993) estimates cross-sectional, time series, and pooled matching functions for Slovak and Czech districts and Czech regions. After testing for regime changes he concludes the matching function is relatively stable around changes in the benefit regime. The results indicate positive dependence of matching on both unemployment and vacancies, and plausibility of the constant returns-to-scale technology assumption in the matching process. Svejnar and Terrell (1994) estimate equations on cross-sectional data for inflows and outflows from unemployment using district data. They conclude there is an obvious difference between the years 1991 and 1992, indicating the labour markets are still in a transition period.

2. Unemployment in the Czech and Slovak Republics - Stylized Facts.

Let us remind that the two republics were originally united in the Czech and Slovak Federative Republic. Both the economic conditions and administration systems were therefore highly comparable at the beginning of the examined period and started to diverge more significantly only after the split was pre-announced in late 1992 and accomplished on January 1, 1993. The Czech to Slovak ratios in area and population size are roughly 2:1. There is 76 districts in the Czech Republic and 38 in Slovakia.

The history of unemployment in the two republics is rather short and dramatic. From virtually zero level at the end of 1989, unemployment rates rose sharply in both republics. At the beginning of 1992 the level was at 4 % in the Czech Republic and as high as 13 % in Slovakia. Development of unemployment is illustrated at Figures 1A and 2A. We prefer using absolute numbers to rates throughout the paper. The reason for this is high unreliability of the data on employment which are reported mostly on annual basis with two-years lag and in many cases they do not capture small firms (under 25 employees) and the newly emerging private sector. Also, we should note that the data used correspond to the official, i.e. registered figures and are not at this point adjusted using an additional information from microdata. Considering that the eligibility for unemployment benefits was conditioned by registration at labour offices and having in mind the poor economic situation in both republics, we can reasonably assume that the registered data on unemployment do not deviate too far from the real ones.

As can be seen from Figures 1A and 2A, the development of unemployment in the two republics keeps similar track. The slow growth of unemployment during 1990 can simply be attributed to some initial inertia in the firms' behaviour and labour hoarding during the period of uncertainty and rapid changes. This must have finished soon, as the firms were unable to maintain the excessive labour force[1]. The steep linear increase of unemployment during 1991 reflected the necessary reduction of the excessive stock of low-productive labour. This, together with the overall recession in the economy, can be considered as the main cause of the huge unemployment increase in 1991. There were two institutional changes during the period that should be noted. First, the administration of unemployment benefits was revised at the beginning of 1992. Since January 1, the eligibility period was cut by one half from 12 to 6 months, and the replacement ratios were slightly reduced. The resulting decrease in unemployment is obvious at both figures. Second, the increase in the unemployment at the beginning of 1993 can be partially explained by the introduction of the new social insurance scheme under which the social insurance contributions are paid out of monthly gross income at the rate of 46 %. The social insurance contributions for the unemployed are paid by the state. Therefore part of small entrepreneurs was forced to shut down and register themselves as unemployed, as well as part of the out of labour force population. The effect of these newcomers could only be temporary, since they were exiting the registered pool as their eligibility period elapsed.

In order to monitor trends in the unemployment behaviour we deseasonalised data using 12-months moving averages. The resulting series are plotted at Figures 3A and 4A. Now the two paths diverge significantly. Especially remarkable is the pronounced decline in the Czech case.

The development of numbers of registered vacancies and matches pre- and after deseasonalization is plotted at Figures 5A through 12A.

Interesting is also to look at the deseasonalized development of the labor market tightness indicator V/U and the "normalized" number of matches M/U (Fig. 13A to 16A). The two indicators exhibit the same dynamics in both republics, but surprisingly enough, the matches do not appear to react at the labor market conditions as captured by V/U. Quite opposite, the downturns and peaks of M/U precede the ones of V/U

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[1]If we compare the dynamics of decrease of industrial production (production of goods in the first half of the year 1991 decreased in comparison with the first half of 1990 in the Slovak Republic by 17.4 %) with the dynamics of decrease in the number of employees (9.1 % in the same period), we can see that the decrease in production is much greater than the one in employment. Before 1990, the level of employment relative to the level of production can be according to unofficial estimates viewed like 10 % overemployment. This means that the decline in employment almost did not offset the over-employment level and enterprises were still retainining considerable amount of excessive labor.

by approximately one month. When we look at the proportions between the numbers of unemployed and vacancies and especially at the amplitude of fluctuations in the series, we can see that the development of the V/U indicator is governed by the behaviour of unemployment rather than vacancies.

3. Data Description.

The data used come from the registries of district labor offices. There are 76 districts in the Czech Republic and 38 districts in Slovakia. Registration with the labor office is necessary condition for unemployment benefit eligibility. Individuals who are still employed but look for another job can register with the labor office as well. Therefore the term used by labor offices is job seeker rather than unemployed. Employers are not required by law to register their vacancies. Some employers did have an incentive to report their vacancies since posting of certain types of vacancies was subsidized from the state budget via labor offices within the framework of active labor policies. The data contain monthly records on the end of month stocks of unemployed and vacancies and on the flows of matches during a month. Records used in our calculations begin in October 1990 for Slovakia and in January 1991 for the Czech Republic and the last data used are those for December 1993. Despite all the well-known deficiencies of the recorded data, we consider the panel structure to be rich enough to produce a statistically robust results.

4. Matching function.

In this section we examine applicability of the standard matching approach to the Czech and Slovak labour markets from two aspects.

First, we test the existence of a stable matching function and determine some of its properties.

Second, we use a flow approach to investigate the way in which the transitions of Czech and Slovak economies influence functioning of their labour markets.

In section 4.1 we briefly recapitulate the theoretical framework of matching technology. In section 4.2 we consider possible effects of transition on matching. In section 4.3 we briefly summarize the results of tests regarding the stability, functional form and other properties of the Czech and Slovak matching technologies. In section 4.4 we estimate the matching functions for separate regions and we try to distinguish different regimes as suggested in section 4.2.

4.1 Specification of the matching function.

We start with the simple Cobb-Douglas form of matching technology as specified in equation (1). Dividing both sides of the equation by the number of unemployed respecifies the matching function in terms of the probability θ_t that an unemployed worker leaves unemployment within period t:

$$\theta_t = \lambda U_t^{\alpha-1} V_t^{\beta} \quad (2)$$

When there is neither competition between vacancies nor unemployed one would expect $\alpha=\beta=1$. Then, the outflow probability from unemployment is linearly dependent only on the number of vacancies: doubling the number of vacancies doubles the outflow probability. However, if there is a competition between unemployed - congestion - then $\alpha<1$ implies that an increase in the total number of unemployed decreases the outflow probability for an individual unemployed worker. In the same way, if $\beta<1$, there is competition between vacancies. As long as $\alpha+\beta>1$, the matching technology exhibits increasing returns to scale: doubling the numbers of both unemployment and vacancies more than doubles the outflow probability for an individual worker. A special case is $\alpha+\beta=1$, when the matching function exhibits constant returns to scale.

Most studies in the field find a constant or mildly increasing returns to scale matching function with β in the range of 0.4-0.6. In that case, the probability that an unemployed worker will find a job in period t depends on the vacancy-unemployment ratio:

$$\theta_t = \lambda (V_t/U_t)^{\beta} \quad (3)$$

It is difficult to give an explanation for constant returns to scale. Pissarides (1990) argues only constant returns to scale lead to a stable unemployment rate. Other explanations focus on the inaccuracy of measuring number of vacancies (Burda and Wyplosz (1994)), or on inadequate discrete time intervals when estimating the continuous time function (Burdett et al. (1994)). Both latter explanations suggest that in reality the matching function exhibits increasing returns to scale.

In some matching models employed job seekers are introduced (Pissarides (1994)):

$$M_t = \eta (U_t + S_t)^{\alpha} V_t^{\beta} \quad (4)$$

where: S_t = number of employed job seekers
Now, unemployed workers have to compete with employed job seekers. In that case, under assumption of equal search intensities for employed and unemployed job seekers, the outflow from unemployment depends on the composition of job seekers:

$$M_{U,t} = \eta U_t (U_t + S_t)^{\alpha-1} V_t^{\beta} \quad (5)$$

where $M_{U,t}$ = outflow from unemployment
If we define the share of unemployed (U/(U+S)) as δ, then:
From the last equation it is obvious that the composition of the pool of job seekers

$$M_{U,t}=\eta\,\delta^{1-\alpha}U_t^{\alpha}V_t^{\beta}\ (6)$$

becomes a shift variable in the matching function. So, only if this composition does not change we can interpret a shift as a change in labour market efficiency.

4.2 Matching and transition.

It is not clear whether we can apply standard matching analysis to the emerging labour markets in the Czech and Slovak Republics. We intend to analyze matching functions, or more specifically, their shape to identify the way in which transitions of the Czech and Slovak economies influences their labour markets. Let us consider the possible effects of transition.

Before the transition there was no unemployment. All the matches were based on either job-to-job, or out of labour force to job transitions. The costs of posting a vacancy were low since having workers in excess of the ones necessary for current production was not harmful. Some of the vacancies did not refer to current unfilled jobs but to the future ones. The transition brought about unemployment. No longer were employees in excess kept within firms. And, some of the employed job seekers became unemployed job seekers. How does this effect the matching between unemployed and vacancies? Our hypothesis is that there are 2 subsequent regimes at work:

1. Transitory period: Unlike before the transition, now there are two kinds of job-seekers to be distinguished: workers searching on-the job (employed job-seekers - ES), and unemployed job seekers (U). The peculiarity of this period is that the total number of job seekers can be assumed to be constant. In terms of equation (5), (U+S) = constant. Then, the size of the outflow from unemployment is in one-to-one relationship to the number of unemployed ($\alpha=1$) and less than proportional to the number of vacancies ($0<\beta<1$). In combination this implies increasing returns to scale and the following specification of the matching function:

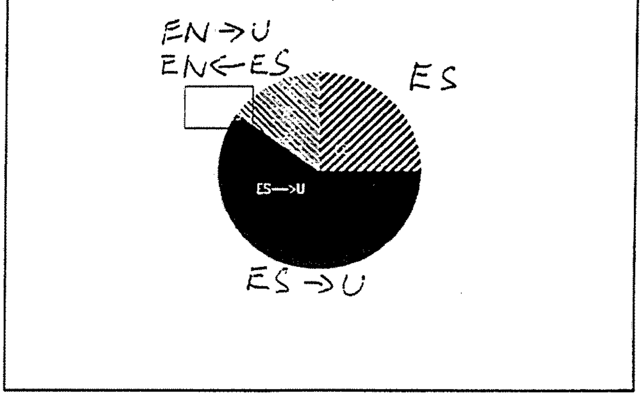
$$M_{u,t}=\eta^{*}U_tV_t^{\beta};0<\beta<1;\eta^{*}=\eta(U+S)^{\alpha-1};U+S=const.\ (7)$$

The constancy of the total number of job seekers can be grounded by the following reasoning:

At the initial stage of transition some of the employed job seekers (ES) became unemployed at the situation where they would otherwise stay on the job until they would find a new one. This group is depicted at Figure 1 like ES->U. There is a group of "consistent" employed job seekers, i.e. workers who were searching on the job before the transition and continued to do so afterwards, depicted like ES at Figure 1. So far, the total number of job-seekers remains constant.

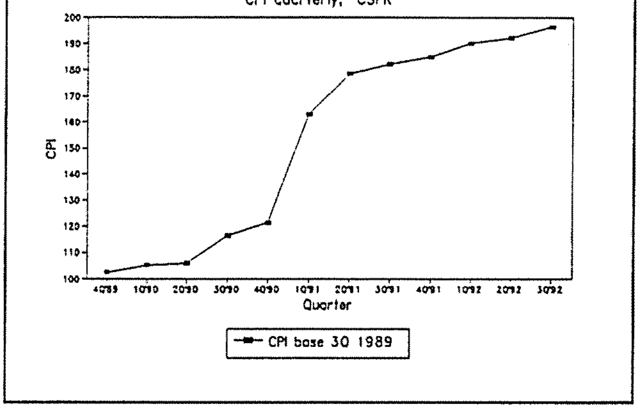
The third type of mobility took place in the following "exchange": some of the workers who were employed and not searching on-the-job before the transition (EN) became unemployed. This flow EN->U represents an addition to the total stock of job seekers. On the other side, the increased uncertainty and expectation of worsening labour market conditions led part of the employed job seekers before the transition to quit searching afterwards and stick to their current jobs. This flow ES->EN represents decrease in the total number of job seekers. As long as the two mentioned flows (EN->U and ES->EN) were balanced, the overall number of job seekers remained constant.

Figure 1.
Composition of job seekers



We want to show that taking into consideration the real economic development in the Czech and Slovak Republics, the assumption of balanced flows is plausible at the initial stage of transition.

Figure 2.
CPI quarterly, CSFR



The arguments supporting our hypothesis can be derived (at this stage in an informal way) from the standard job search model with incorporated signal extraction problem. The brief sketch of the situation is as follows: a job seeker samples offers from the ex-ante wage distribution characterized by the cumulative density function F(W). The offered wage corresponds to the expected real marginal product of a particular worker-job match: W=PX, where P is an i.i.d. random variable (sometimes thought of like the realized price level at the moment of sampling) and X denotes the expected productivity of the worker at the job, which captures the "quality of the match". The variables P and X influence the outcome of the search process in several ways. Among the most important characteristics are the degree of observability at the moment of sampling (single degenerate value versus "noisy" signal), the parameters of their distributions, and the evolution of these distributions in time (translations and spreads). Based on our observations of the underlying economic background, we would suggest the following sequence of events:

(1) A sudden upward shift of the P distribution. This corresponds to the lifting of price controls at the beginning of 1991 (Figure 2). Since the change is unanticipated, many job seekers are "tricked" into accepting jobs that they otherwise would not accept. This is because the high realization of P makes W=PX higher than the reservation wage even when the "quality of the match" as described by the productivity is low.

So far, we did not deal with information aspects of the story.[2] At the supply side of the market, the sudden increase of prices brings about temporarily higher profits, which enables firms to retain part of the excessive labour force that would otherwise have to be laid off.[3] Both these effects worked in the short run in the common direction of dampening the increase in the number of unemployed. In terms of our Figure 1, this means that at the initial stage of transition the flow EN->U is relatively small.

(2) Shortly after the translation of P distribution, the dispersion of X distribution increases. This is because under the new conditions the skills acquired in the previous regime may not be useful anymore while new types of jobs appear which require new (e.g. entrepreneurial) abilities. In other words, the quality of the match between a particular job-worker pair observed at the time of sampling the wage offer shifts from a "more inspection good" to a "more experience good". The screening from both sides becomes more noisy. Consequently, the distribution of ex-ante offered wages collapses, wage growth takes place more on-the-job than between jobs and workers search less once they are employed. In terms of Figure 1, this process results in the flow ES->EN which compensates the increase in the number of job seekers caused by the above mentioned flow EN->U.

Since both phenomena described by 1) and 2) were undoubtedly in effect at the initial stage of transition, we hypothesize that the two counterflows may balance each other so that the total number of job seekers remains roughly constant. Matching function can be characterized by (7), which implies $\alpha=1$, $0<\beta<1$, and increasing returns to scale.

2. Business-as-usual period. The upward drift in distribution of P slows down and becomes systematic. After recognizing and adjusting to this situation, the agents previously "tricked" into accepting offers quit. When the low quality of some of the realized matches is revealed, workers either quit or they are laid off (the "quit" terminology here requires more elaboration on the perceptions of labour market conditions by the workers). The abrupt jump in prices is not present anymore for some firms to override the real problems and firms are shedding the excessive labour. This brings about sharper increase in the number of unemployed, which is not anymore balanced by the decrease in the number of employed job seekers. The total number

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[2] So far, we did not deal with information aspects of the story. At this stage we would have to include an element of asymmetric information in order to avoid assuming some kind of "money illusion" at the demand side of the labour market.

[3] This argument hinges on an increase in the spread of price distribution. From the translation of the distribution itself, as documented at Fig. 2, we cannot infer anything about the changes in its variance. However, the underlying proces which brought about the shift is likely to having generated also increased dispersion. In further work we will try to make a quantitative assessment about the spread based on disaggregated price indeces.

of job seekers (U+S) does not remain constant. Agents are getting used to the new labour market environment. The functioning of the labour market approaches standard situation described by (4). The matching function exhibits constant or increasing returns to scale as specified in (6). Shifts may appear because of changes in labour market efficiency or because of changes in the composition of the pool of job seekers (δ).

4.3 Properties of the Czech and Slovak matching technologies.

At the preliminary stage of our analysis we used a set of twelve cross-sections over districts (six for every republic) to investigate the properties of matching technologies. We examined stability over time, spacial homogeneity, functional specification and returns-to-scale aspects.

We concluded that the matching functions in both republics were neither homogeneous in time, nor across districts. From the three basic functional forms tested - Cobb-Douglas, constant elasticity of substitution, and transcendental logarithmic form - the letter two were rejected in favour of the Cobb-Douglas one. The returns-to-scale aspects were tested for the Cobb-Douglas form. Constant returns to scale could not be rejected at 5 percent significance level in majority of cases. The exact functional forms and more details about estimation are contained in Appendix 1.

Based on the results of preliminary tests, we estimated the basic loglinear Cobb-Douglas functional form of matching function on pooled samples of all districts and regions and for districts within separate regions. Since according to the preliminary tests both the time and regional homogeneity were rejected, we repeated the estimation with adjustments for time and space heterogeneity. Detailed description of the method, results and discussion are contained in Appendix 1. Here we present the two representative results: Table 4.1 contains the OLS estimates of Cobb-Douglas loglinear form on the pooled sample of all districts, Table 4.2 contains the same estimates when an adjustment for geographical heterogeneity across districts was made by transforming the time series on all variables for every particular district into deviations from its own mean (the "means-substraction" procedure). An adjustment for time heterogeneity was done in both cases by adding 12 calendar dummy variables (reported in column 2), or the full sets of monthly dummy variables (column 3).

The main conclusion from comparing the two tables is the following: adjusting for geographical heterogeneity across the districts increases all the estiamted coefficients in their magnitude. However, adjusting for time heterogeneity within each of the two tables brings about substantial decrease in the magnitude of estimated coefficients. The results in the columns (3) of Table 4.2 where the adjustments are "finest" suggest decreasing returns to scale in both republics. This signals a need to elaborate in more details on the "time dimension" of the process. In particular, we need to distinguish between the shifts of the relationships over time due to the changes in efficiency of matching or composition of job-seekers, as captured by the constant term, and the changes in coefficients on lnU, lnV over time which capture the returns-to-scale aspects. There is a possibility of discrete regime breaks due to changes in labour market institutions or underlying economic conditions. We investigate these aspects in more details in Section 4.4. We switch from the level of districts to regions and we examine the time series behaviour in every separate region individually. This is because we believe that districts are spatially too small areas in the sense that there is too much commuting across the borders of districts. Therefore, the matches within a district can originate from unemployment registered in some other neighbouring district and in addition to the measurement problems mentioned in the Appendix 1, we have an additional "noise" in the relationship.

Table 4.1
Matching function, OLS estimation
District level: 76 districts for CR, 38 districts for SR
Monthly data: Jan 91-Oct 93 for CR, Jan 91-Sept 93 for SR

Dependent variable: m	Czech Republic			Slovak Republic		
	(1)	(2)	(3)	(1)	(2)	(3)
Independent variables:						
u	0.86	0.85	0.77	0.95	0.57	0.69
	(73.62)	(76.63)	(71.32)	(27.24)	(30.86)	(12.91)
v	0.23	0.23	0.17	0.22	0.20	0.22
	(21.72)	(22.61)	(18.14)	(8.81)	(7.90)	(8.72)
Summary statistic						
Obs.	2508	2508	2508	1254	1254	1254
Adj.R sq.	0.726	0.994	0.995	0.471	0.974	0.978

t-statistic in parentheses

Table 4.2
Matching function
Adjustment for geographical heterogeneity by means-substraction.

Dependent variable: m	Czech Republic			Slovak Republic		
	(1)	(2)	(3)	(1)	(2)	(3)
Independent variables:						
u	1.15	1.15	0.80	1.06	1.15	0.45
	(62.44)	(64.58)	(26.74)	(30.52)	(64.58)	(3.78)
v	0.36	0.35	0.12	0.21	0.35	0.11
	(23.22)	(24.25)	(6.16)	(6.70)	(24.25)	(3.72)
Summary statistic						
Obs.	2508	2508	2508	1254	1254	1254
Adj.R sq.	0.644	0.697	0.749	0.469	0.496	0.550

t-statistic in parentheses
Notes:
Column (1): estimated equation:

$$m_t=\beta_1+\beta_2 u_t+\beta_3 v_t+\epsilon_t$$

Here lowercase letters denote natural logarithms
Column (2): 12 calendar dummy variables included
Column (3): monthly dummy variables included (CR 34, SR 33)
Since the complete sets of dummy variables were used, constant in case (1) is not reported

4.4 Different matching regimes at the regional labour markets.

We want to investigate whether it is possible to distinguish between the two time periods described in Section 3. Since different regions may be in a different stage of transition, cross-sectional estimates at a specific moment in time or pooled time series estimates may be misleading. Furthermore, if the spatial areas are too small, there is too much commuting to and from the area. Then flows out of unemployment do not originate only from the matching process within the area, but also from the neighbouring areas. Therefore we do not use information on districts but on regions.

We estimate separate matching functions for each of the 8 Czech and 4 Slovak regions. When estimating a matching function, we use the current monthly outflow from unemployment and the stocks of unemployed and vacancies as measured at the end of the previous month.

We have the following estimation strategy. First, we estimate a matching model over the whole period Jan 1990 - Dec 1993. Then we investigate whether there is a structural break in the relationship within the period by allowing the matching function to have different specifications before and after a specific point in time. This was done at quarterly intervals. We analyzed whether there was a structural break between every subsequent March/April, June/July, September/October and December/January within the examined period. [4]If we find a structural break we investigate whether it refers to a shift in labour market efficiency or to a change in the shape of the matching function (parameters α and/or β). In all cases it appears that if a break occurs, it is connected with change in the constant of the matching function, which suggest a change in labour market efficiency. Our main interest is in the characterization of labour markets in terms of period in or after the transition. Essentially, there are two ways of testing whether $\alpha < 1$. The first one considers the estimated coefficient directly. The second one investigates whether the matching function exhibits constant returns to scale. If one of the two applies, we consider the labour market to be in what we call business-as-usual period. So, we investigated whether $\alpha < 1$ and whether $\alpha + \beta = 1$. When an unrestricted estimate reveals $\alpha > 1$, we tested whether it was allowed to restrict α to be equal to 1. In all cases it appeared that such a restriction was allowed.[5]

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[4] We also tried a more sensitive procedure of detecting the time of structural break: we examined the rlationship on monthly basis to decide where the point of break should be located such that the two separate regimes are as different as possible in terms of Chow-test. This method appears to be more sensitive and accurate in locating the break in the relationship but given the short time period under investigation, we could detect a "forced" break where in reality the regoinal labour market may still remain in the "transitory" period.

[5] In the sake of a more robust estimate of the break we also experimented with the transcendental logarithmic specification of the matching function which is more complex and therefore more

Table 4.3 gives an overview of our estimation results. The overall impression is that we achieved sensible results for each of the 12 regions. For 7 regions the coefficient of the unemployment variable is larger than 1. For Northern Bohemia (at 5 % significance level) and Western Bohemia (at 10 % significance level) we find the unemployment coefficient smaller than 1. From this we conclude that these two regions may be in the business-as-usual period. With the exception of Prague, all vacancy coefficients are in the range between 0 and 1. For Bratislava (Slovak capital) we find constant returns to scale matching function. Therefore we conclude Bratislava also overcame the transitory period. All other regional labour markets matching function exhibit increasing returns to scale and appear to be still in the transitory period.

For Western and Northern Bohemia we find a negative shift of the matching function possibly indicating a decrease in labour market efficiency starting in the third quarter of 1992. For Northern Moravia, Bratislava and Western Slovakia we find a positive shift of the matching function in various quarters of 1992. Only for Western Slovakia this shift coincides with the change in the social security system (which took place on Jan 1, 1992).

In the further work we are refining the analysis in the following ways:

We investigate a possibility of including second-order lags of independent variables into the regressions. This technique is also superior to the Cochrane-Orcutt transformation in case of eliminating autocorrelation.

We examine changes in returns-to scale aspects at a finer level by letting monthly time dummies interact with the independent variables. However, there may be a trade-off between these refinements and an increased noise in the relationship, as we would have to switch back to the district level which would give us more degrees of freedom.

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stable over time than the Cobb-Douglas one. As we expected, the procedure yielded dates of the regime breaks for separate regions. to be more clustered in time. However, the test of joint significance of coefficients over the whole period only favored the translog form over the Cobb-Douglas one for three Slovak regions (except the capital Bratislava). This confirms the more "chaotic" situation at the Slovak side.

Table 4.3 Estimation results[a)]

Czech Republic

Region	log(U)	log(V)	R^2	shift	Ftest	period
Prague[b)]	1.14 (0.13)	1.03 (0.23)	.785	no		1
Central[b)] Bohemia	1.15 (0.18)	0.33 (0.13)	.768	no		1
Southern Bohemia	1.03 (0.08)	0.51 (0.06)	.863	no		1
Western Bohemia	0.84 (0.09)	0.47 (0.13)	.775	3'92 (-)	7.41	2[c)]
Northern Bohemia	0.81 (0.07)	0.58 (0.11)	.821	3'92 (-)	7.02	2[d)]
Eastern Bohemia	1.17 (0.09)	0.38 (0.08)	.852	no		1
Southern Moravia	1.00 (0.08)	0.75 (0.11)	.878	no		1
Northern Moravia	1.10 (0.08)	0.67 (0.12)	.898	2'92 (+)	4.56	1

Slovak Republic

Region	log(U)	log(V)	R^2	shift	Ftest	period
Brati-slava	0.84 (0.15)	0.23 (0.09)	.775	3'92 (+)	4.52	2[e)]
Western Slovakia	1.09 (0.12)	0.23 (0.13)	.937	1'92 (+)	6.94	1
Central[b)] Slovakia	1.27 (0.16)	0.79 (0.24)	.908	no		1
Eastern Slovakia	0.87 (0.19)	0.69 (0.16)	.632	no		1

a) standard errors in parentheses, R^2 is corrected for degrees of freedom, Shift refers to the first quarter after the shift of the matching function where (-) means decrease in efficiency and (+) means increase in efficiency; F-test refers to the F-value when applying Chow-test for a structural break; period refers to the period of transition where 1 = transitory period and 2= business-as-usual period

b) estimation results after correcting for autocorrelation by Cochrane-Orcutt transformation

c) 10%-level $\alpha<1$

d) 5%-level $\alpha<1$

e) constant returns to scale cannot be rejected at 5% significance level

5. The Beveridge Curve.

In the 1960s and 1970s the Beveridge curve - an inverse relationship between unemployment and vacancies - was a popular device to study labour markets. Recently, there is a renewed interest in this type of analysis.

Most of analyses explains joint fluctuations in unemployment and vacancies and tries to distinguish between aggregate shocks and sectoral shifts as the main determinants. A noteworthy recent contribution to the latter type of analysis by Hosios (1994) identifies assumptions under which the unemployment-vacancy data represent a capable tool for distinguishing between the two mentioned factors.

The concept of Beveridge curve assumes a hyperbolical relationship between unemployment and vacancies in UV space of a form $U*V = k$, where k is a positive constant. The use of Beveridge curve as a diagnostic tool has been widely discussed in the literature (see for example Abraham (1987), Blanchard and Diamond (1989), Boersch-Supan (1991)). The idea of disentangling the shifts of the curve into movements along rays through an origin brought about by structural changes and movements along the curve brought about by aggregate shocks to an economy attracted attention of many. Boersch-Supan (1991) tested the hypothesis that a rightward shift of Beveridge curve indicates an increase in structural unemployment using a panel annual data of nine regions in West Germany for the period 1963 - 1988. The results did not support structural interpretation of the shifts of Beveridge curve. Cyclical variables performed better in explaining the outward shifts than variables associated with structural unemployment. Moreover, an evidence was found for part of the shift being caused by aggregation bias (Abraham, 1987).

The structure of our panel data appears to be richer (76 districts and 34 months for the Czech Republic, 38 districts and 36 months for Slovakia) and therefore suitable to test the structural hypothesis with reasonable statistical stability. We basically replicate the test used by Boersch-Supan, the main differences being in using monthly rather than annual data and slightly different variables capturing the composition of unemployment pool.

The UV relationships for the Czech and Slovak Republics are depicted at Fig. 18A and 19A, respectively. The similarity of the two curves is obvious. It appears there were three different regimes in place as captured by the three parts of UV curve. First, a downward-sloping quasi-hyperbolical part since the beginning of period until March 1991. Second, the upward sloping shift to the right since April till December 1991 and third, movements along the downward-sloping quasi-hyperbolical branch again, with an upward swing culminating in August 1992 for the Czech Republic and in September 1992 for Slovakia.

Our preliminary estimations were performed for the Czech Republic at three different levels of disaggregation: the republic as a whole, regional level (8 regions) and district level (76 districts). In this way we were able to observe how did the further levels of disaggregation influence our results.

Table 5.1 presents the results of estimation. The first two columns represent the level of republic, the second and third pair of columns - regional and district level, respectively. Within every pair of columns the first one contains the results of estimation including two dummy variables: SHIFT1 to control for the rightward shift during the period April - December 1991 and SHIFT2 for the period of the years 1992 and 1993. The second column presents the results of estimation including also variables related to the composition of unemployment pool: proportion of job seekers receiving unemployment benefits (COMP), proportion of unemployed blue collar workers (BLUE), proportion of unemployed technicians (TECH), and proportion of unemployed with completed university education (UNIV). If the hypothesis about the link between shifts of the UV curve and changes in structural unemployment is true, then these structural variables should have high explanatory power and substitute for the explanatory power of simple shift variables. As we can see from Table 5.2, for the level of republic this was really the case. Both shift variables are significant. After adding the "structural" variables into estimation, the originally highly significant coefficients on both shift variables dramatically decreased in their magnitudes and became insignificant. At this stage, the results would indicate that the outward shift was really related to structural changes. Among the structural variables only proportion of job seekers receiving unemployment benefits (COMP) and proportion of job seeking technicians (TECH) appear to be significant. As could be expected, COMP shifts the relationship outwards (since the unemployed with compensation are more "choosy" than the ones without). Surprisingly, the TECH variable shifts the relationship substantially inwards, as if technicians were matched extremely effectively (the term technicians denotes skilled workers with completed secondary education).

In order to verify somewhat arbitrary choice of the shift variables, we included set of monthly dummies into our estimation. This was possible only after increasing number of observations by going to the more disaggregated level of regions. The results are contained in Table 5.2.

Table 5.1

U-V relationship, Czech Republic
Structural variables included in columns (2)
Regional fixed effects

	Dependent variable: U					
	Republic		Regions		Districts	
Independent variable	(1)	(2)	(1)	(2)	(1)	(2)
Const.	6.36	16.72	0.03	4.36	0.12	0.30
	(0.99)	(0.75)	(0.14)	(3.22)	(10.39)	(4.60)
1/V	13804.74	251627.8	1898.74	2652.08	-7.58	-6.05
	(0.05)	(0.9)	(2.67)	(3.68)	(-5-98)	(-4.74)
SHIFT1	10.71	0.20	1.35	0.83	0.14	0.15
	(4.86)	(0.06)	(10.24)	(4.79)	(11.84)	(11.86)
SHIFT2	9.03	-1.01	1.33	0.89	0.10	0.09
	(2.78)	(-.197)	(9.12)	(4.94)	(8.62)	(7.14)
COMP		18.56		0.68		-0.15
		(1.73)		(1.5)		(-6.04)
BLUE		8.59		-4.55		-0.15
		(0.33)		(-3.08)		(-2.31)
TECH		-98.39		-5.55		-0.07
		(-2.77)		(-4.02)		(-1.13)
UNIV		134.9		-22.09		1.02
		(0.65)		(-2.76)		(2.15)
PRG			-0.87	-0.05		
			(-5.92)	(-0.18)		
BNORTH			0.38	0.42		
			(2.78)	(2.64)		
BWEST			-0.63	-0.71		
			(-4.58)	(-5.19)		
BSOUTH			-0.91	-1.07		
			(-4.98)	(-5.68)		
BEAST			0.13	0.20		
			(0.96)	(1.49)		
MNORTH			2.95	2.98		
			(21.3)	(20.95)		
MSOUTH			1.85	1.98		
			(13.46)	(14.07)		
Summary statistic:						
Obs.	32		262	2505	2505	
Adj. R sq.	0.41	0.66	0.84	0.85	0.07	0.09

Notes:
All models were fit using OLS
t-statistic in parentheses
Coefficients * 10 000
For identification of regional variables see Appendix 2.

Table 5.2

U-V relationship, Czech Republic
Regional and district level, monthly dummies

Independent variable	Dependent variable: U Regions		Districts	
Constant	-0.01	(-0.06)	0.11	(5.97)
1/V	1982.53	(3.25)	-8.03	(-6.42)
PRG	-0.86	(-7.58)		
BNORTH	0.38	(3.64)		
BWEST	-0.64	(-6.00)		
BSOUTH	-0.92	(-6.27)		
BEAST	0.13	(1.25)		
MNORTH	2.95	(27.76)		
MSOUTH	1.85	(17.54)		
Summary statistic				
Observations	262		2505	
Adjusted R sq.	0.91		0.12	

Notes:
All models were fit using OLS
t-statistic in parentheses
Coefficients * 10 000
The full set of coefficients on 33 monthly dummies is contained in Appendix 2.

In the regional regressions we included set of constants for each region to control for the "fixed effects". Description of the variables is in Appendix 2. Central Bohemia served like reference region in the estimation. At the level of districts, the fixed effects are not included yet. The difference is obvious from comparing the adjusted coefficients of determination. the low level of 0.12 for the district level indicates a need to apply a panel technique of estimation in order to take into account heterogeneity over districts. As we expected, April 1991 and January 1992 appear to be breakpoints in the sense that a remarkable discrete jump in the magnitude of coefficients appeared in April (at the regional level only) and at both levels, the values of coefficients increase monotonically till January 1992 when they culminate and start to decrease afterwards (the full set of coefficients on monthly dummy variables can be found in Appendix 2). Therefore we continued using the two shift variables in our estimations.

We repeated the regressions at the regional level (see the second two columns of Table 5.1). As we moved from the level of republic to the more disaggregated level of regions, the evidence in favour of structural causes of shift became weaker. After adding the structural components, coefficients on the shift variables decreased slightly in their magnitudes and dropped about a half in their significance, but they still remained significant (with t-values of 4.79 for SHIFT1 and 4.94 for SHIFT2). Also the behaviour of structural variables changed considerably. COMP became insignificant while the other three variables were significant and shifting the relationship downwards. These results are in line with the findings of Boersch-Supan for Western Germany. The assessments about structural causes of shifts in UV relationship based on examination of annual time series data for a country may be misleading because the highly aggregated data do not contain enough information to discover the nature of the process.

In the next step we proceeded to the highly disaggregated level of districts (the last two columns of Table 5.1). Like in the regional case, the shift variables did not disappear after adding the structural components. Furthermore, they retained their magnitudes and high significance (the significance of coefficient on SHIFT1 even slightly increased). The behaviour of structural variables was quite inconsistent with the previous cases, COMP becoming significant again but negative, the same was true for BLUE, UNIV became positive and TECH insignificant. We are cautious with the interpretation of the district results. As we already mentioned, the panel of district data requires a panel technique of estimation. This is confirmed by the low values of adjusted coefficients of determination as compared to the regional case (0.07 and 0.09 versus 0.84 and 0.85).

In addition to the "structural" explanation of shifts in the UV relationship there is a "geometrical" one. As pointed out by Abraham (1987), linear aggregation of convex Beveridge curves over regional sub-samples leads to aggregation bias. In order to examine to what extent were the shifts observed in our data caused by this bias, we estimated separate Beveridge curves for each region. If an outward shift is detected in each region with magnitude smaller than the one of the shift at the republic level, then part of the outward shift can be explained by aggregation bias rather than a structural change. Estimation of separate regional Beveridge curves is presented in Table 5.3. For all the regions coefficients on both shift variables were significant. In all the regions the magnitude of corresponding coefficients was smaller than the one at the level of republic.

We constructed an average of coefficients on the separate regional shifts weighted by the labour force in each region as end of 1991. The average values are still smaller for the both shift variables than the overall shift at the republic level (15.9 as compared to 107.1 for SHIFT1 and 14.6 compared to 90.3 for SHIFT2, here all the coefficients are for simplicity reduced by factor of 1000). Therefore we can assume that part of the detected outward shift was caused by aggregation bias rather that a structural change.

Disaggregation by branches combined with the spatial one would yield more reliable results. Also, it is desirable to include structural variables capturing the supply side and repeat the tests for aggregate variables. However, given our short-run limitations, we have to work with monthly data which are not available for many of the desired indicators.

Table 5.3
U-V relationship for separate regions
Czech Republic

Dependent variable: U

Indep. variable	PRG	BC	BN	BW	BS	BE	MN	MS
Const.	0.39	0.74	0.68	0.10	-0.03	0.84	0.23	-0.64
	(2.46)	(1.47)	(1.14)	(0.33)	(-0.07)	(1.18)	(0.21)	(-0.39)
1/V	3077.8	-651.5	669.5	1406.4	794.5	-622.5	10220.5	9740.02
	(2.42)	(-0.26)	(0.20)	(1.28)	(1.08)	(-0.21)	(1.59)	(1.12)
SHIFT1	0.21	1.08	1.15	0.59	0.86	1.31	3.10	2.47
	(2.50)	(5.17)	(4.50)	(5.62)	(4.46)	(4.12)	(5.23)	(4.77)
SHIFT2	-0.28	0.92	1.26	0.68	0.75	0.90	3.23	2.88
	(-2.72)	(3.53)	(4.42)	(4.30)	(2.31)	(2.02)	(5.69)	(4.02)
Summary statistic								
Obs.	31	31	31	31	31	31	31	31
Adj. Rsq.	0.87	0.49	0.48	0.50	0.39	0.31	0.49	0.50

Notes:
All models were fit using OLS
t-statistic in parentheses
Coefficients * 10 000

PRG = Prague
BC = Central Bohemia
BN = Northern Bohemia
BW = Western Bohemia
BS = Southern Bohemia
BE = Eastern Bohemia
MN = Northern Moravia
MS = Southern Moravia

6. Conclusions.

We conclude that the matching function may be a valuable tool for analysing transition at labour markets. Although the available information refers to a short time period, we find plausible estimates of regional matching functions.

In general, preliminary tests on cross-sectional data at discrete points in time validated the Cobb-Douglas functional specification. The matching relationships exhibited considerable heterogeneity in both time and geographical dimensions. The geographical heterogeneity was reduced by a fixed-effects approach (means-substraction). An additional "noise" in the relationship due to commuting among relatively small spatial areas of districts was reduced by switching to higher level of spatial aggregation - regions.

The main field of our interest remains the evolution of matching technology over time. We attempt to distinguish between periods with different regimes. Our main conclusion is that most regional labour markets are still in the transitory period. This conclusion is in line with previous research but adds the new insight in that in some regions the labour markets are in the past-transition business-as-usual period.[6] Also, the matching functions (apart from one region) exhibit increasing returns to scale. It is not clear whether this is a real phenomenon or whether it was caused by exclusion of employed job seekers from the estimation.

In an empirical investigation of the shifts in unemployment-vacancy relationship in the Czech Republic, we find inconclusive results regarding the hypothesis about structural causes of these shifts. The structural variables are loosing explanatory power as we proceed to finer levels of geographical disaggregation. Our results suggest the aggregation bias can be of an importance when judging the case of "structural versus aggregate" causes of the shift.

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[6] We are aware of the fact that the "transitory" period itself may include several different regimes. At the present stage the the short time period does not allow us to investigate these aspects more rigorously.

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APPENDIX 1.

The Cobb-Douglas form of the matching function was specified in the familiar form:

$$M_t=\alpha U_t^{\beta_2} V_t^{\beta_3} e^{\epsilon_t} \tag{A.1}$$

Equation (A.1) can be rearranged to obtain the linear relationship

$$\ln M_t=\ln\alpha+\beta_2\ln U_t+\beta_3\ln V_t+\epsilon_t \tag{A.2}$$

The notation is simplified by denoting the logarithms by lowercase letters:

$$m_t=\beta_1+\beta_2 u_t+\beta_3 v_t+\epsilon_t \tag{A.3}$$

Equation (A.3) was estimated by the ordinary least squares procedure.

The nonlinear functional form of production function with constant elasticity of substitution (CES) was specified in the following form:

$$M_t=\gamma[\delta U_t^{\rho}+(1-\delta)V_t^{\rho}]^{-\nu/\rho} e^{\epsilon_t} \tag{A.4}$$

where $\gamma>0$ stands for the efficiency parameter, $0<\delta<1$ for the distribution parameter, $\rho>0$ for the returns-to-scale parameter, and $\nu>1$ for the substitution parameter. By taking the natural logarithms we obtain the linear form

$$\ln M_t=\ln\gamma-\frac{\nu}{\rho}\ln[\delta U_t^{\rho}+(1-\delta)V_t^{\rho}]+\epsilon_t \tag{A.5}$$

Equation (A.5) is transformed to obtain its approximation which is linear with respect to the parameter ρ[7]:

$$m_t=\ln\gamma+\nu\delta u_t+\nu(1-\delta)v_t-\frac{1}{2}\rho\nu\delta(1-\delta)[u_t-v_t]+\epsilon_t \tag{A.6}$$

The model then can be estimated like the one with the nonlinear restrictions on coefficients under exact identification. The "unrestricted" version of (A.6) is the

[7]The approximation was done by taking the Taylor series expansion of ln M_t around $\rho=0$, dropping the terms which contain the powers of ρ higher than one, and denoting the logarithms by the lowercase letters.

following one:

$$m_t=\beta_1+\beta_2u_t+\beta_3v_t+\beta_4[u_t-v_t]^2+\epsilon_t \tag{A.7}$$

The coefficients in equation (A.7) are related to the parameters in equation (A.4) in the following way:

$$\beta_1=\ln\gamma \tag{A.8}$$

$$\beta_2/(\beta_2+\beta_3)=\delta \tag{A.9}$$

$$\beta_2+\beta_3=\nu \tag{A.10}$$

$$-2\beta_4(\beta_2+\beta_3)/\beta_2\beta_3=\rho \tag{A.11}$$

The relevance of the CES specification can be verified by testing the value of β_4.[8]

The transcendental logarithmic form of the matching function was obtained from the equation (A.6) by allowing the terms in $[u-v]^2$ to differ in their coefficients:

$$m_t=\beta_1+\beta_2u_t+\beta_3v_t+\beta_4u_t^2+\beta_5v_t^2+\beta_6u_tv_t+\epsilon_t \tag{A.12}$$

The three types of production function were estimated by the OLS method using six cross-sections: February, July and October 1991 and 1992 for Slovakia and July and October 1991 and January, April, July and October 1992 for the Czech Republic.

The relevance of particular functional forms was tested by F-test on the significance of coefficient. The constant returns-to-scale hypothesis was tested for the Cobb-Douglas form. Also, the stability of the relationship over time and across the districts was examined using the Chow test within the six cross-sections groups for the both republics.

Generally, the CES and translog forms were rejected in favour of the Cobb-Douglas form. Constant returns-to-scale were accepted in vast majority of cases. The time homogeneity was generally rejected. The test of homogeneity over districts were

[8] When β_4 is not significantly different from zero, the CES specification is rejected in favor of the Cobb-Douglas form.

only applicable in the Cobb-Douglas case and homogeneity was rejected as well.

Since the preliminary tests rejected the time and regional homogeneity assumptions in our samples, an attempt to account for the heterogeneity in a simple way was done by combining various techniques in the following way:
the basic Cobb-Douglas equation was estimated by OLS for each republic at three levels:
A) pooled sample of all districts
B) pooled sample of all regions
C) districts within separate regions

A) and B) For the pooled samples an attempt to account for the geographical heterogeneity was done by:
- means substraction
- fixed effects (for regions only)
Within each of these methods an account for time heterogeneity was made by adding into equation:
- 12 calendar dummies
- 33 monthly dummies
within each of these, an attempt to detect a structural break was made by estimating the relationship for:
- the whole period Jan 1991 - Sep (Oct) 1993
- subperiod 1991
- subperiod 1992/93.
C) For the districts within separate regions, the basic equation was estimated by the following methods:
- OLS
- Instrumental variables (first-order lags of lnu, lnv)
- Cochrane-Orcutt transformation

After all, the above mentioned extensions did not change the results substantially. General patterns remain the same in all kinds of estimations. In the main text we presented the results of OLS estimation of the basic Cobb-Douglas functional form (1) in the pooled sample of districts with time-heterogeneity adjustments (Table 4.1) and time and geographical heterogeneity adjustments (Table 4.2).

After the adjustment by means substraction and including monthly dummies into regression, the coefficients on vacancies are decreased considerably both in magnitude and in significance in both republics (the lower entry in the columns (3) of Table 2A).

There are more possible reasons for this peculiar behaviour of vacancies. The most obvious one would be the low cooperation between employers and labour offices. Employers are not obliged to report their vacancies to the labour offices.

The second likely explanation is the methodological problem of estimating discrete time series data. Observations on monthly basis may be insufficient if the average duration of a vacancy or a job-search period is shorter than one month. The peculiar behaviour of vacancies was already obvious from our preliminary estimations on the monthly cross-sections. In many cases the coefficients on vacancies were insignificant and sometimes even negative. In our next work on vacancy durations we will examine the grounds for this explanation.

The third possible reason for poor performance of vacancies in our estimation is a possibility of simultaneity bias since the realized matches are depleting stocks of unemployed and vacancies at the same time. Tentative estimations including lagged values of independent variables did not increase the magnitude of coefficients substantially.

There is also a possibility of structural break in the relationship during the investigated period. We plotted the patterns of the monthly dummies at Fig. 17A. It is obvious that the year 1991 was somehow special in both republics. While the coefficients for corresponding months keep approximately the same path in the years 1992 and 1993, the values for 1991 are subject to wild fluctuations. This must be a result of already mentioned "big bang" of the year 1991, when the whole system started to react on the changed conditions and the increase of unemployment was steep and almost linear over the whole year. Apparently, an equilibrium approach is not relevant for the period. Repeated estimations with two subperiods 1991 and 1992/93 seem to support the idea that during 1991 the system was subject to disequilibrated stochastic dynamics while during 1992/93 the estimated relationship tended to be more stable with all the estimated coefficients significant and positive. Further investigation of this topic is contained in the main text (Section 4.2 and 4.3).

APPENDIX 2.

List of regions and corresponding fixed-effects variables used in the estimation of UV relationship in the Czech Republic

Czech Republic (8):

Prague (PRG)
Central Bohemia (Reference region)
Northern Bohemia (BNORTH)
Western Bohemia (BWEST)
Southern Bohemia (BSOUTH)
Eastern Bohemia (BEAST)
Northern Moravia (MNORTH)
Southern Moravia (MSOUTH)

The full set of coefficients on monthly dummy variables from the estimation reported in Table 5.2:
U-V relationship
Regional and district level, time dummies

	Dependent variable: U			
	Regions		Districts	
Independent variables:				
Constant	-0.01	(-0.06)	0.11	(5.97)
1/V	1982.53	(3.25)	-8.03	(-6.42)
PRG	-0.86	(-7.58)		
BNORTH	0.38	(3.64)		
BWEST	-0.64	(-6.00)		
BSOUTH	-0.92	(-6.27)		
BEAST	0.13	(1.25)		
MNORTH	2.95	(27.76)		
MSOUTH	1.85	(17.54)		
Feb91	0.03	(0.16)	0.02	(0.63)
Mar91	0.04	(0.17)	0.04	(1.46)
Apr91	0.48	(2.21)	0.07	(3.02)
May91	0.64	(2.94)	0.09	(3.58)
Jun91	0.87	(3.97)	0.12	(4.76)
Jul91	1.24	(5.74)	0.15	(6.09)
Aug91	1.50	(6.99)	0.17	(6.89)
Sep91	1.76	(8.21)	0.19	(7.83)
Oct91	1.88	(8.77)	0.20	(8.34)
Nov91	1.96	(9.12)	0.21	(8.67)
Dec91	2.03	(9.50)	0.21	(8.77)
Jan92	2.19	(10.21)	0.22	(9.09)
Feb92	2.07	(9.58)	0.20	(8.24)
Mar92	1.82	(8.38)	0.17	(6.87)
Apr92	1.51	(6.94)	0.13	(5.41)
May92	1.30	(5.89)	0.11	(4.30)
Jun92	1.25	(5.43)	0.09	(3.79)
Jul92	1.21	(5.43)	0.09	(3.70)
Aug92	1.20	(5.40)	0.09	(3.64)
Sep92	1.16	(5.24)	0.09	(3.54)
Oct92	1.08	(4.90)	0.08	(3.25)
Nov92	1.04	(4.73)	0.08	(3.17)
Dec92	1.11	(5.05)	0.09	(3.47)
Jan93	1.39	(6.33)	0.12	(4.83)
Feb93	1.36	(6.22)	0.11	(4.67)
Mar93	1.30	(5.96)	0.11	(4.37)
Apr93	1.19	(5.44)	0.10	(3.91)
May93	1.12	(5.09)	0.09	(3.56)
Jun93	1.14	(5.21)	0.09	(3.71)
Jul93	1.26	(5.76)	0.10	(4.28)
Aug93	1.38	(6.33)	0.12	(4.83)
Sep93	1.46	(6.73)	0.13	(5.31)
Summary statistic				
Observations	262		2505	
Adjusted R sq.	0.91		0.12	

Notes: All models were fit using OLS
t-statistic in parentheses
Coefficients * 10 000

APPENDIX 3.

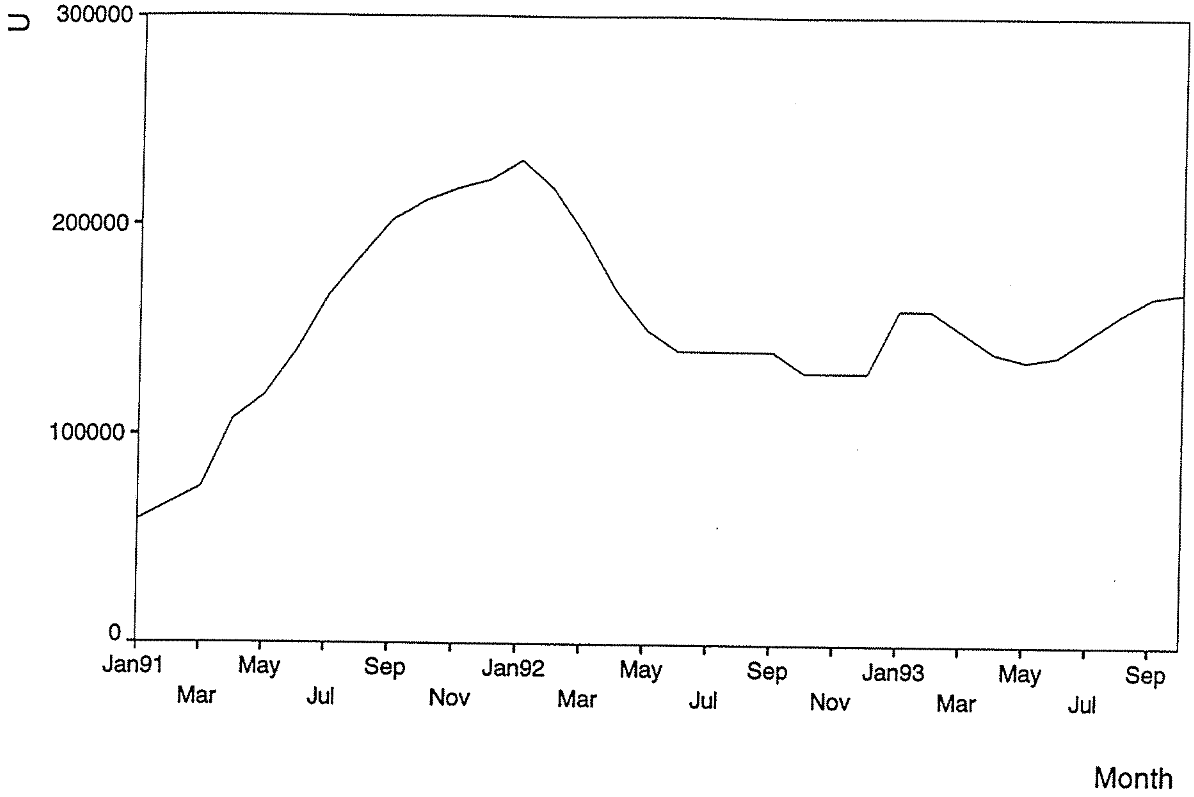
Figures 1A - 19A

Unemployment in the Czech and Slovak Republics, Stylized Facts

(Section 2)

Fig. 1. Unemployment

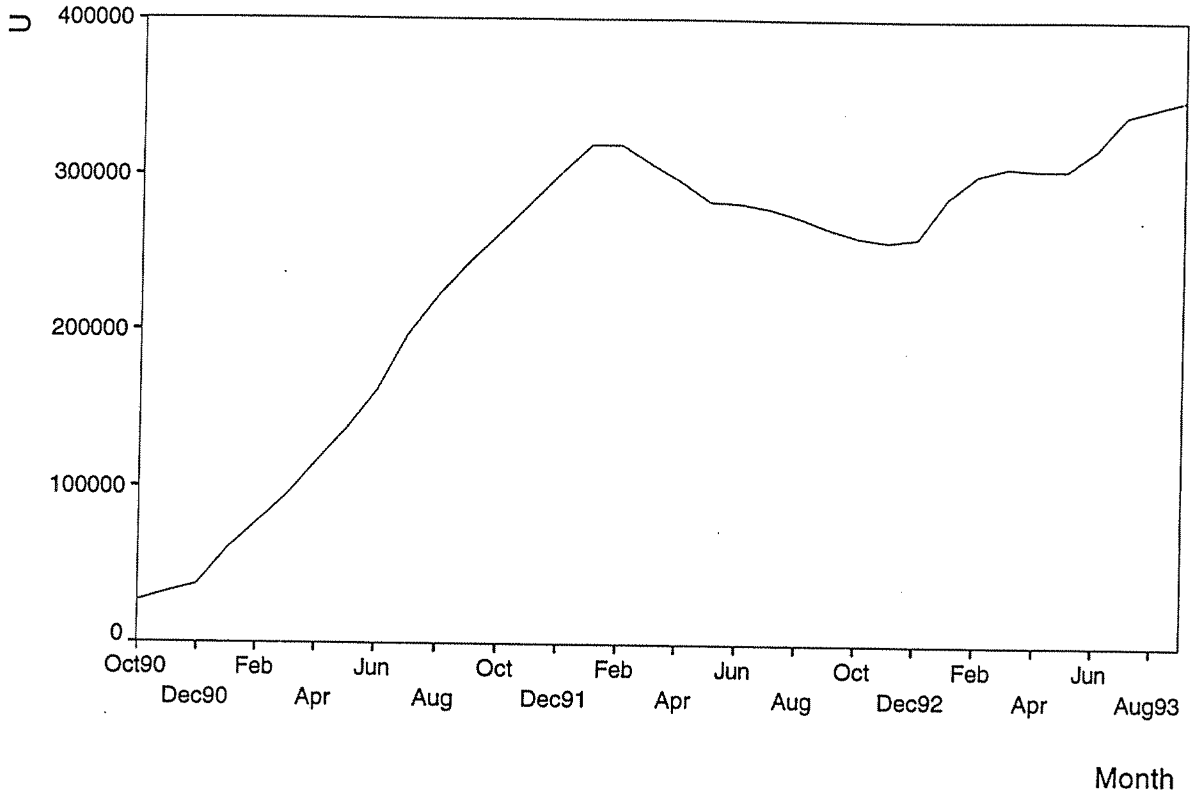
Czech Republic



End of month stocks

Fig. 2. Unemployment

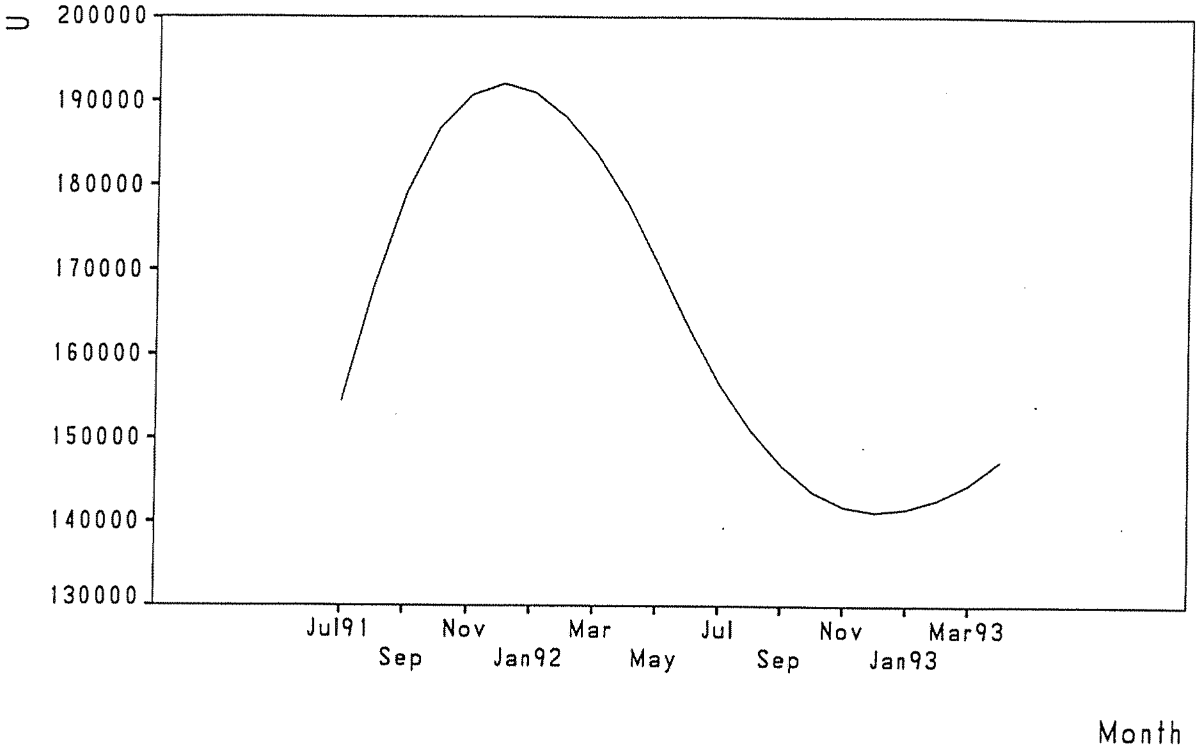
Slovak Republic



End of month stocks

Fig. 3. Unemployment

Czech Republic

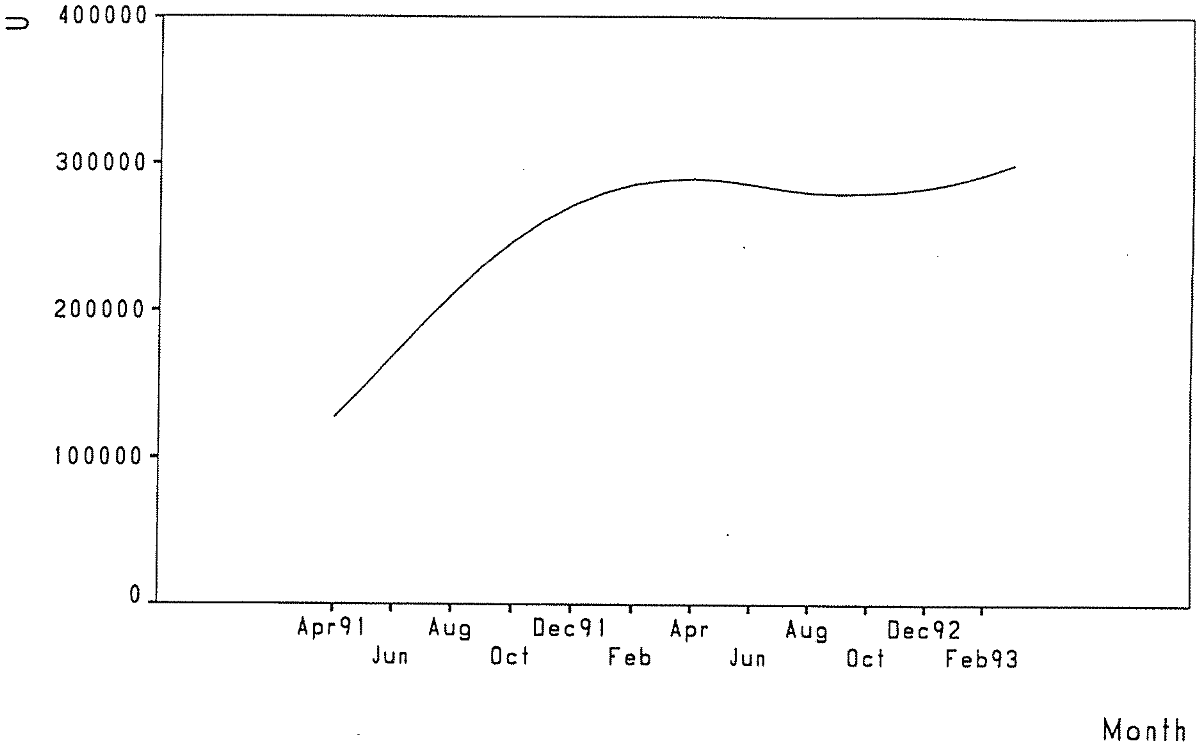


U = MA(12) using end of month stocks

for period Jan91 - Oct93

Fig. 4. Unemployment

Slovak Republic

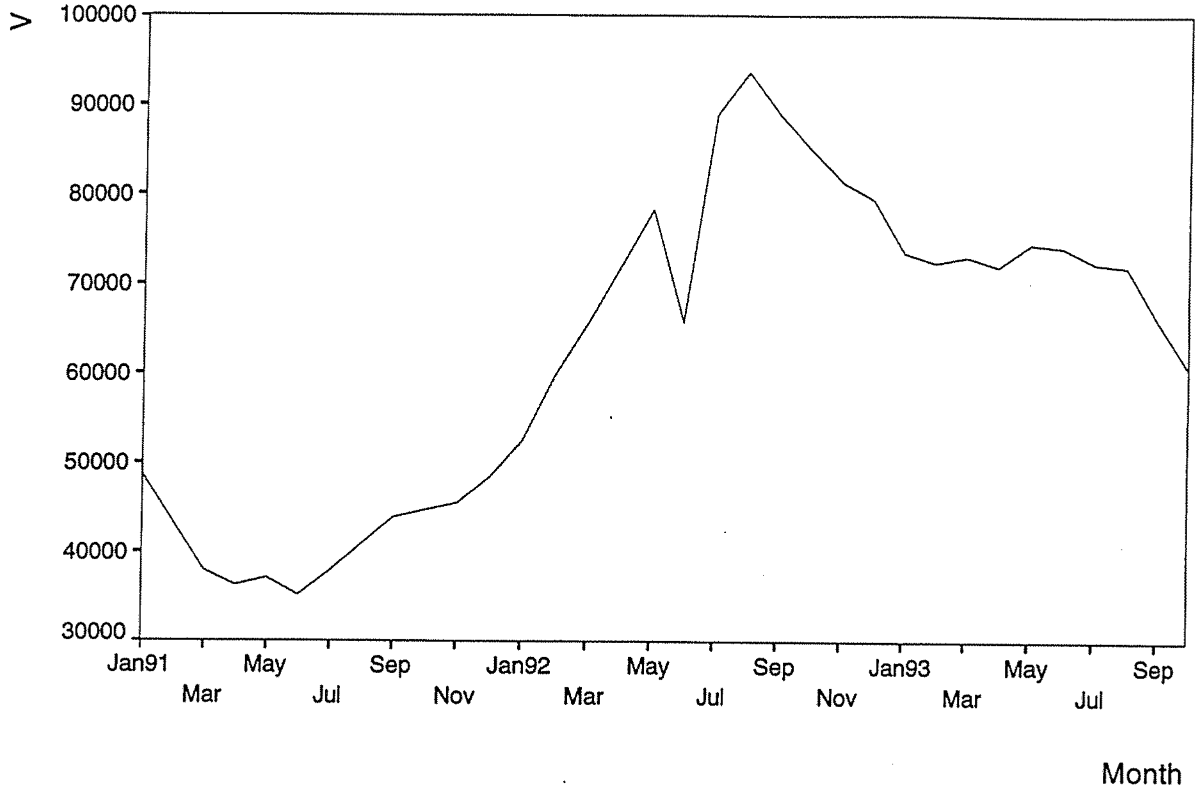


U = MA(12) using end of month stocks

for period Oct90 - Sept93

Fig. 5. Vacancies

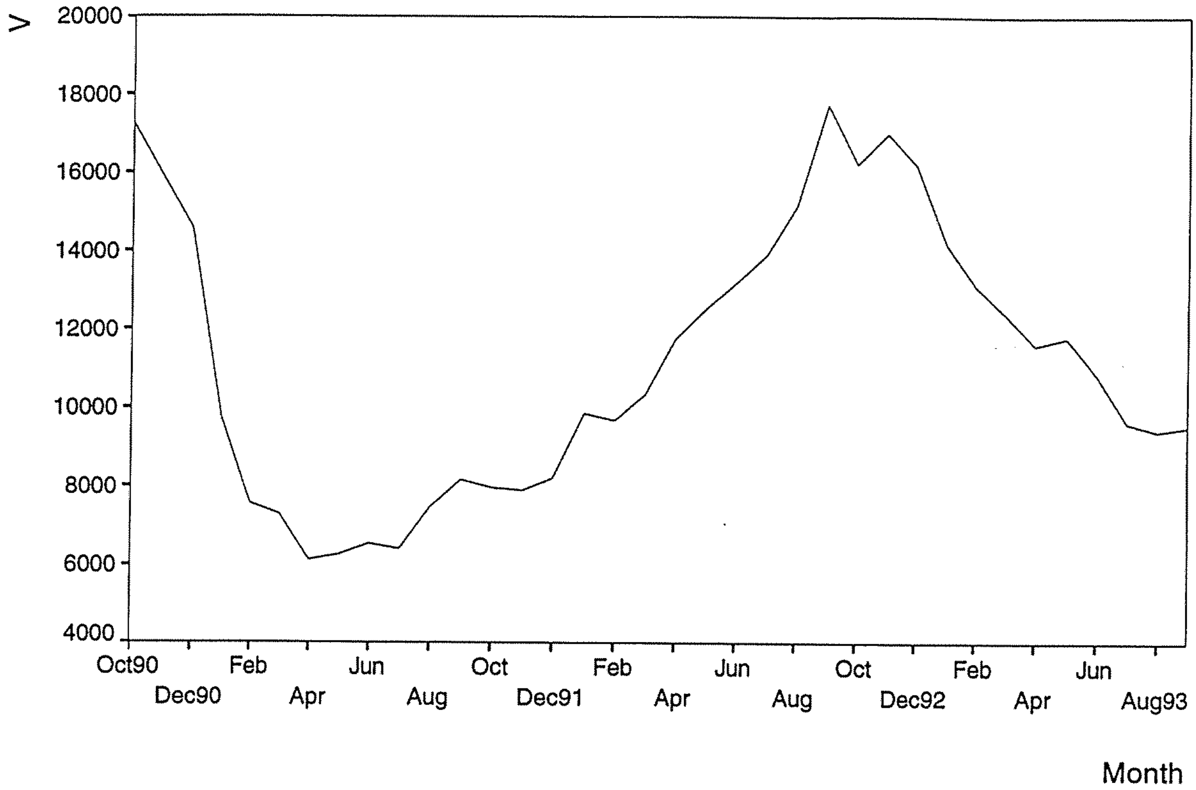
Czech Republic



End of month stocks

Fig. 6. Vacancies

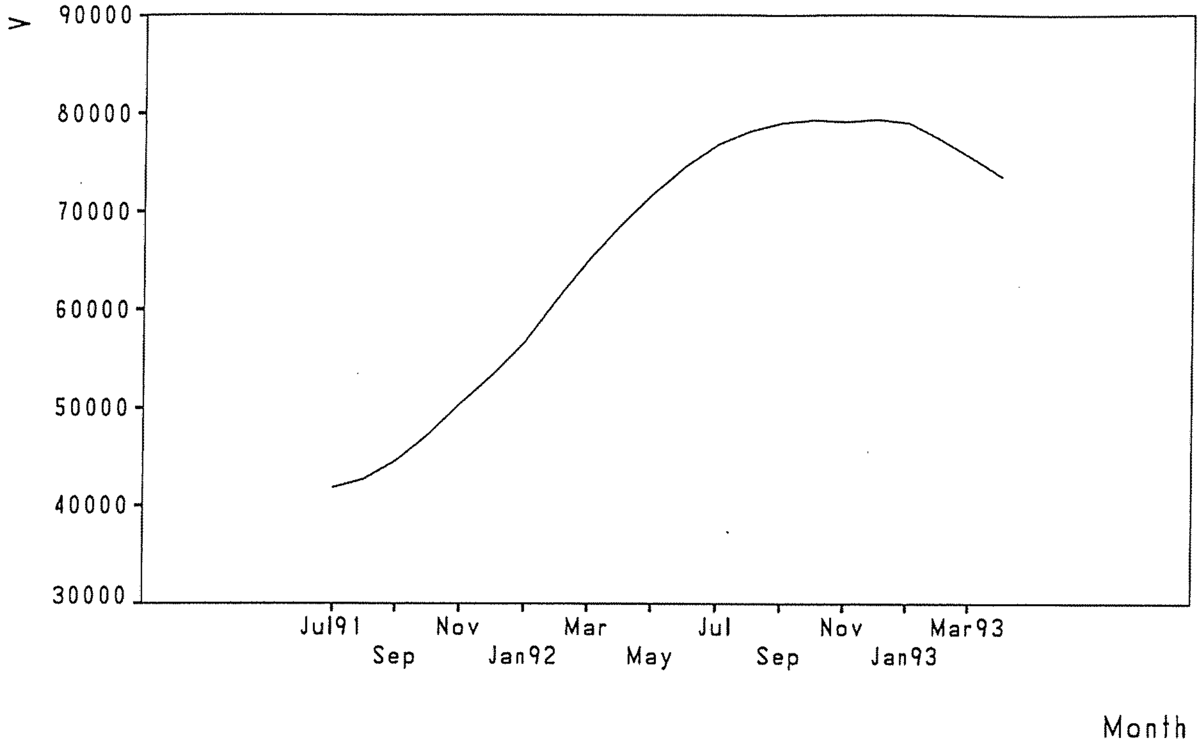
Slovak Republic



End of month stocks

Fig. 7. Vacancies

Czech Republic

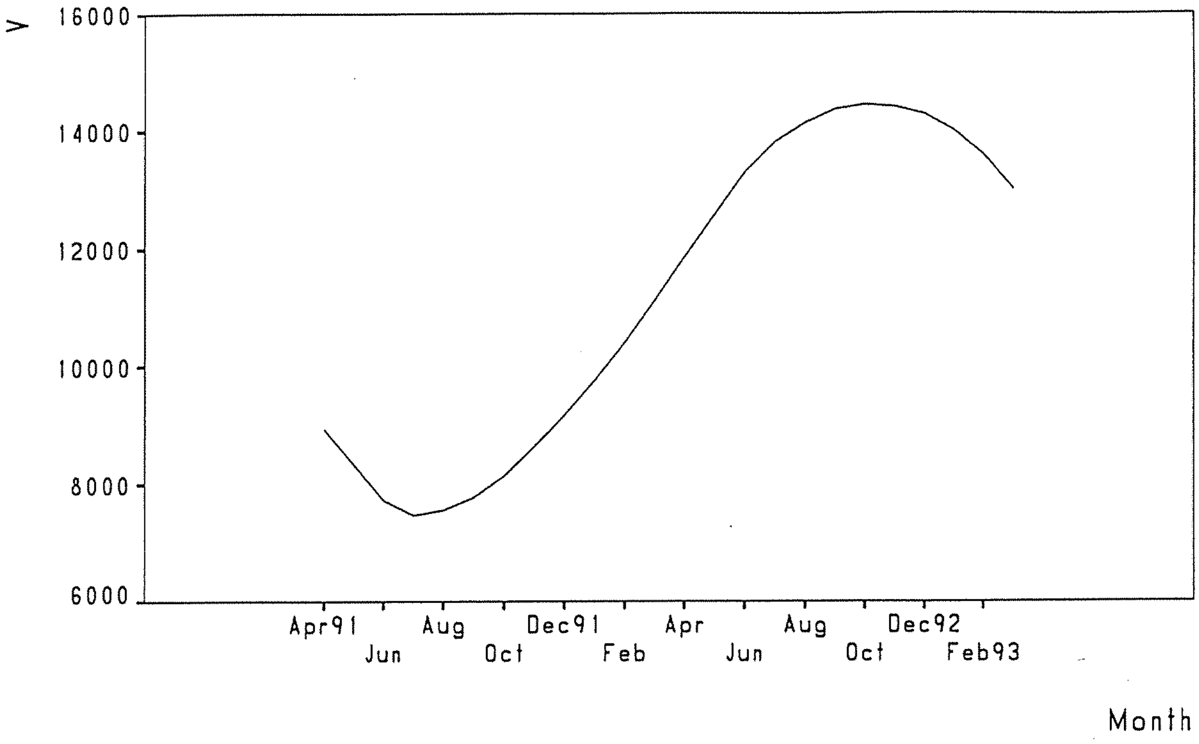


V = MA(12) using end of month stocks

for period Jan91 - Oct93

Fig. 8. Vacancies

Slovak Republic

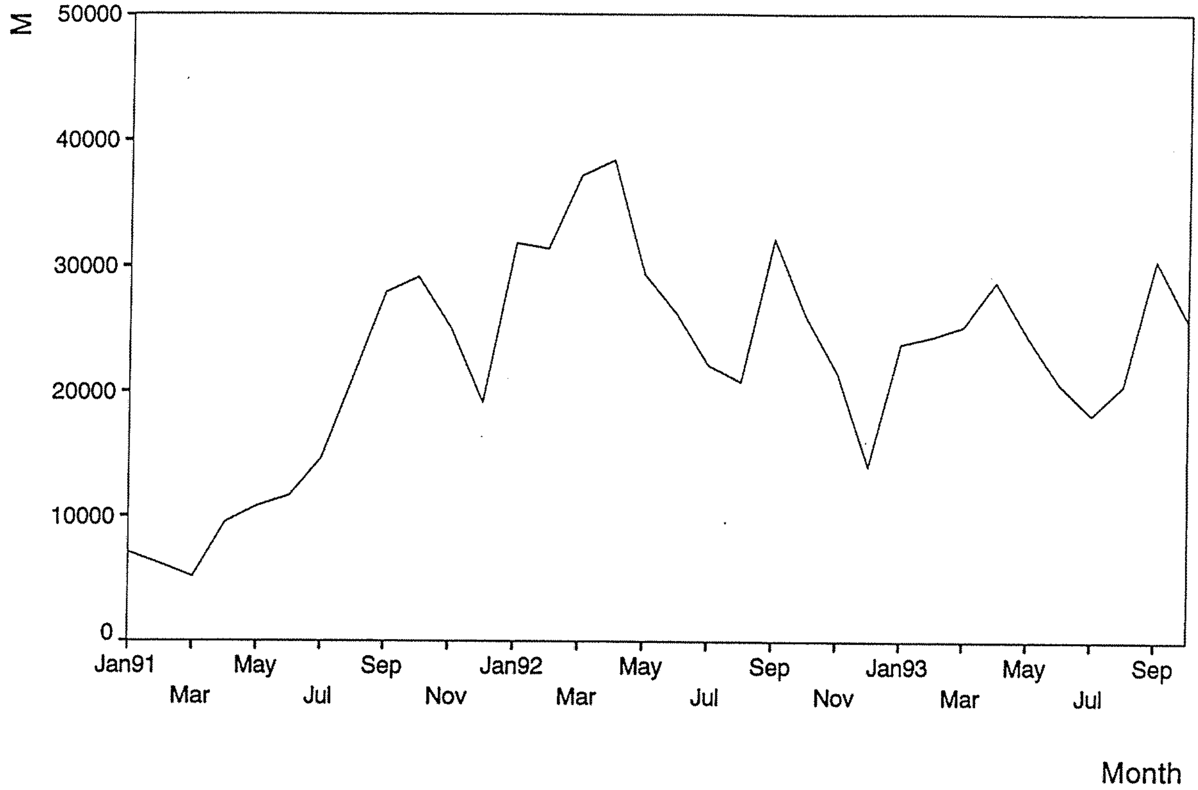


V = MA(12) using end of month stocks

for period Oct90 - Sept93

Fig. 9. Matches

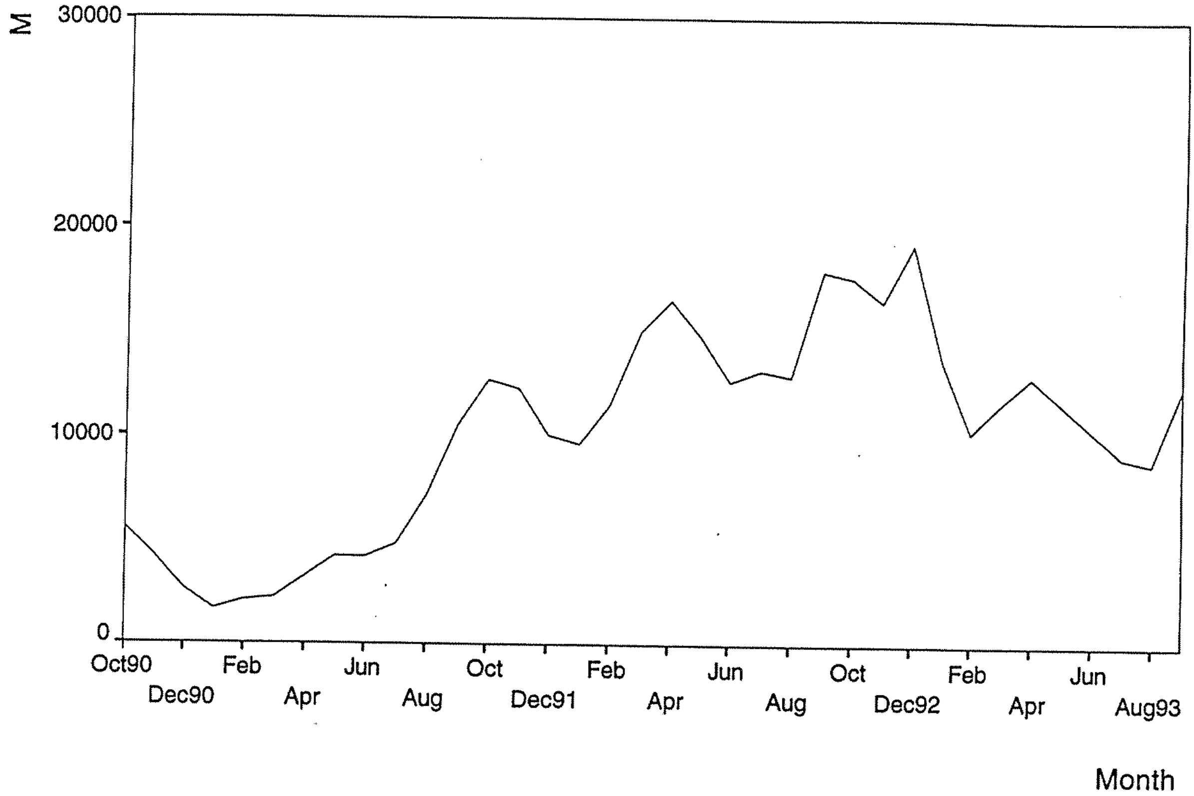
Czech Republic



Monthly flows

Fig. 10. Matches

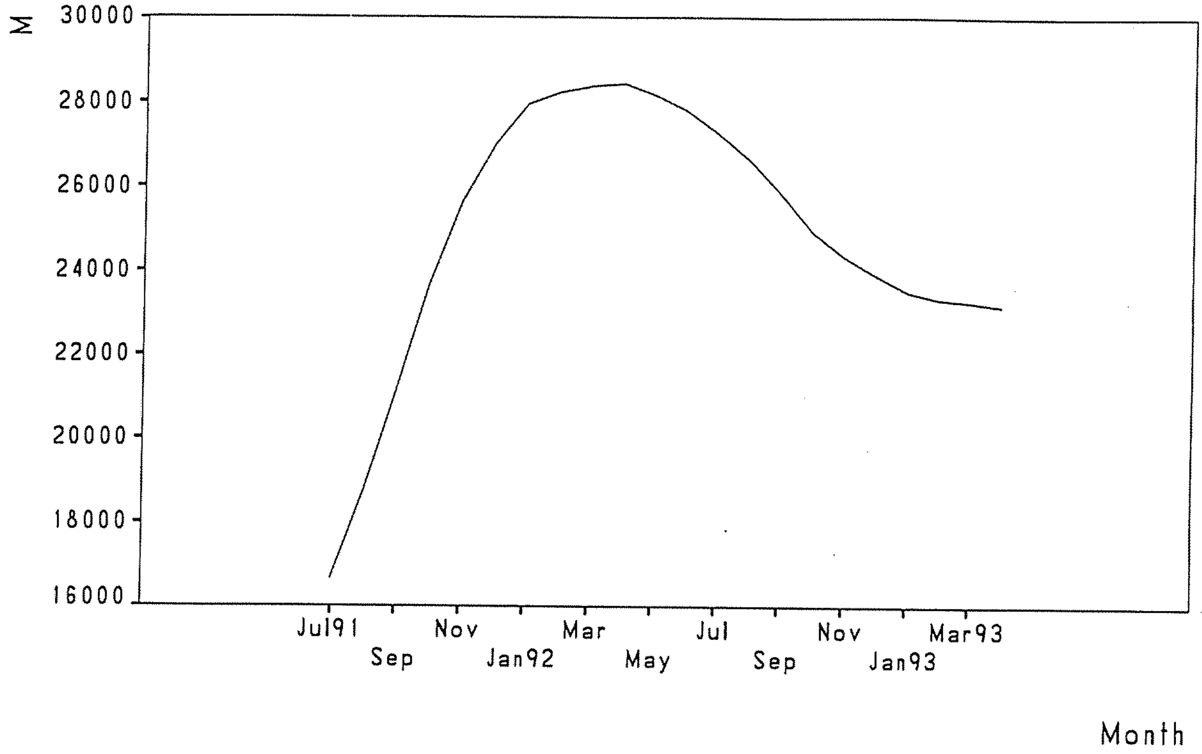
Slovak Republic



Monthly flows

Fig. 11. Matches

Czech Republic

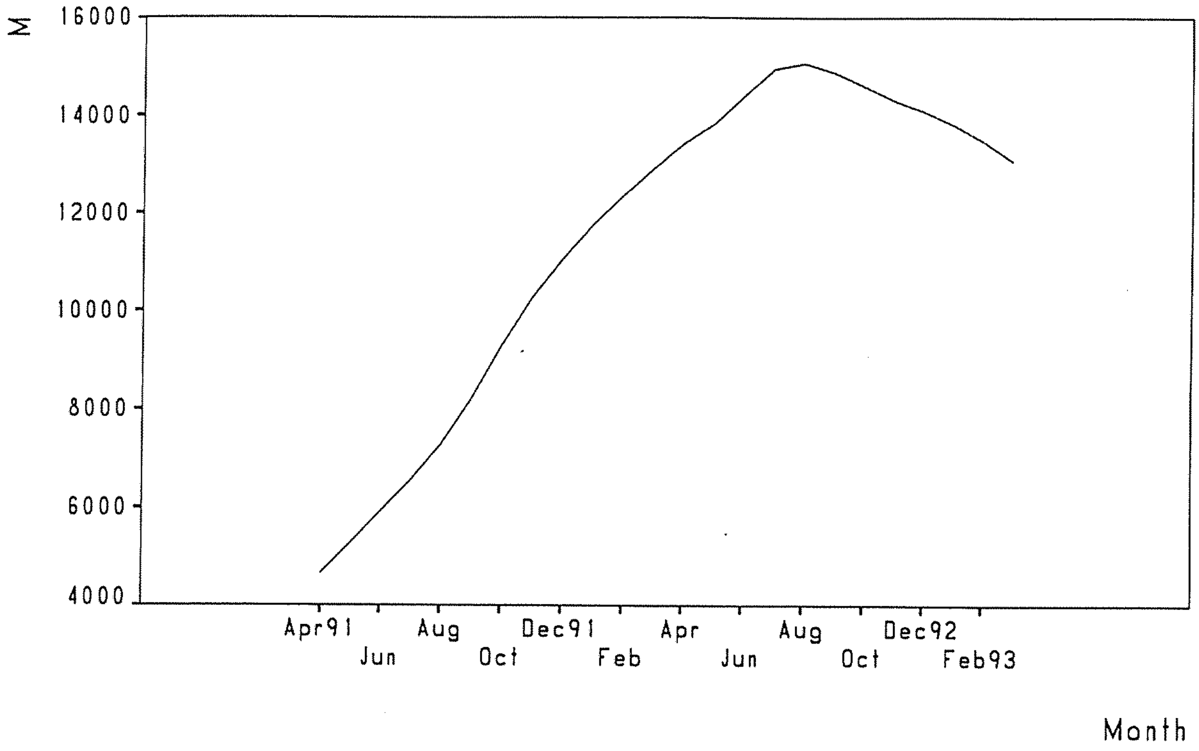


M = MA(12) using monthly flows of filled vacancies

for period Jan91 - Oct93

Fig. 12. Matches

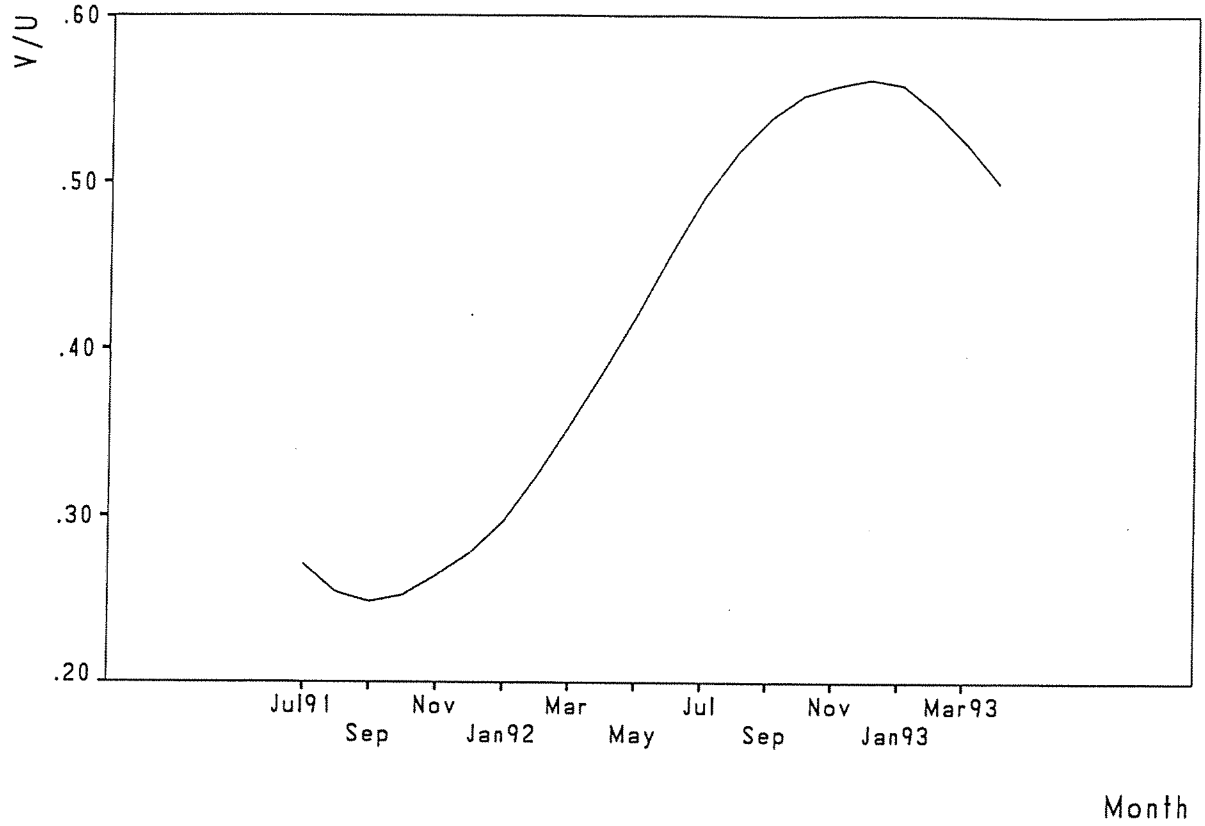
Slovak Republic



M = MA(12) using mothly flows of filled vacancies

for period Oct90 - Sept93

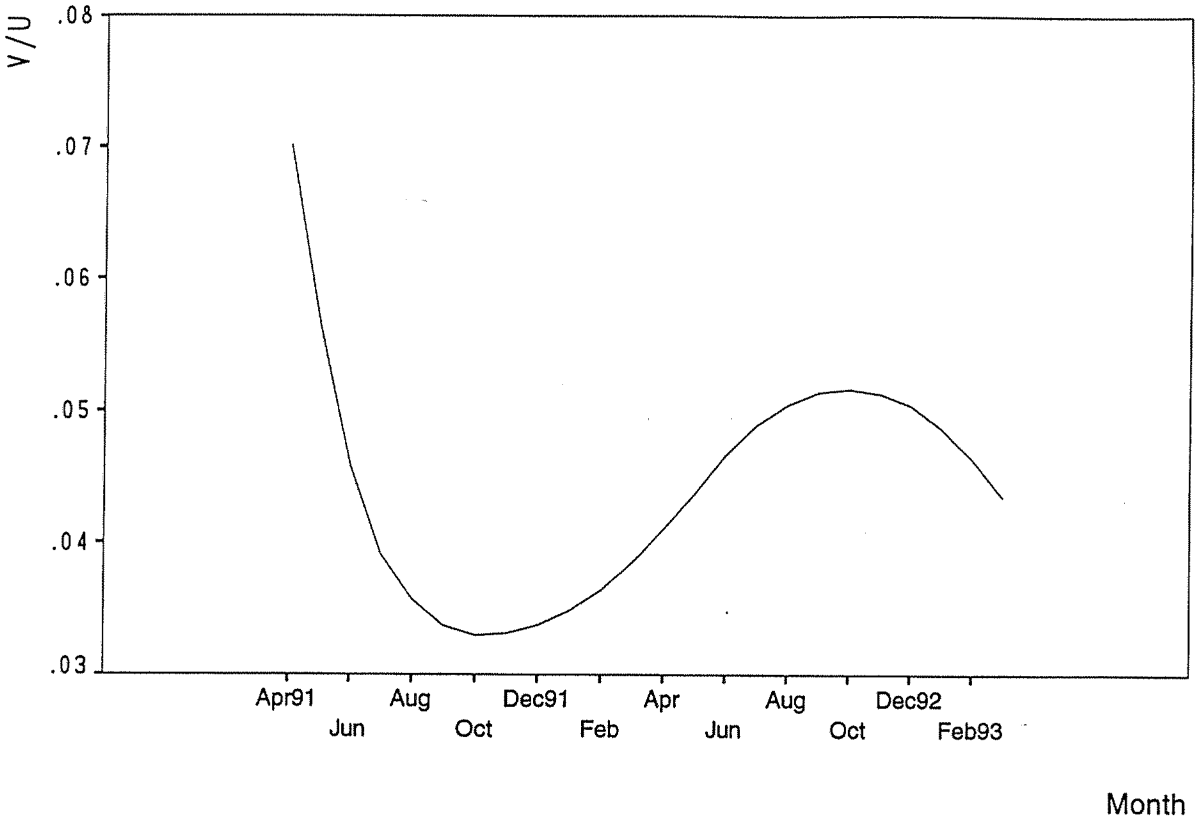
Fig. 13. Czech Republic



MA(12) using monthly data Jan91 - Oct93

on end of month stocks for V, monthly flows for M

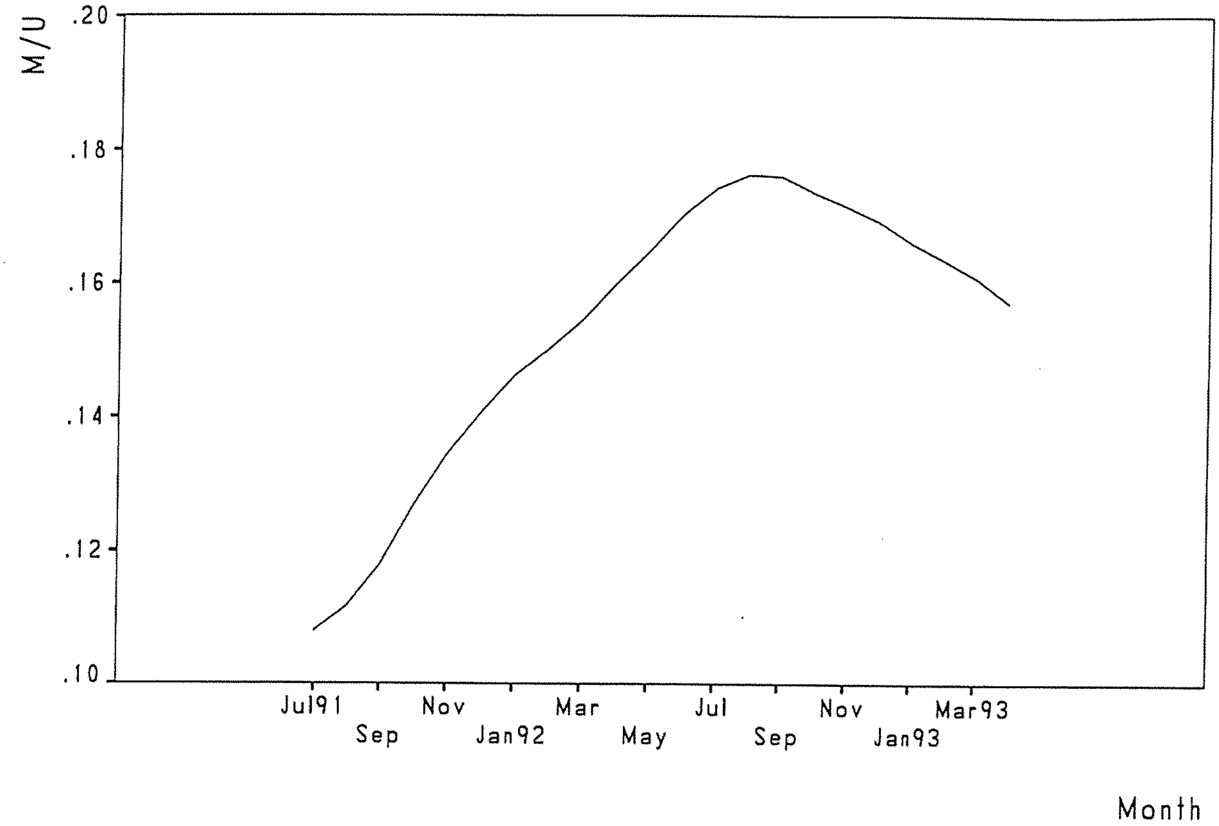
Fig. 14. Slovak Republic



MA(12) using monthly data Oct90 - Sept93

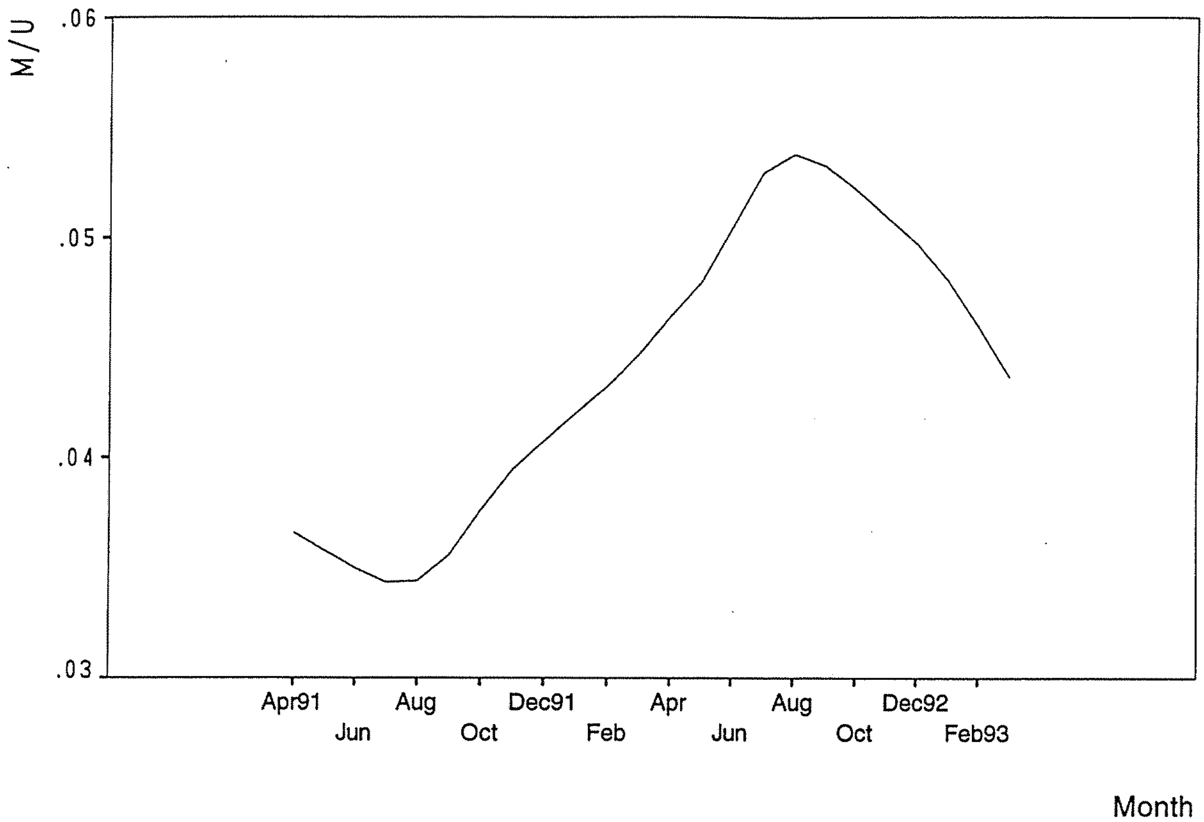
on end of month stocks for V and U

Fig. 15. Czech Republic



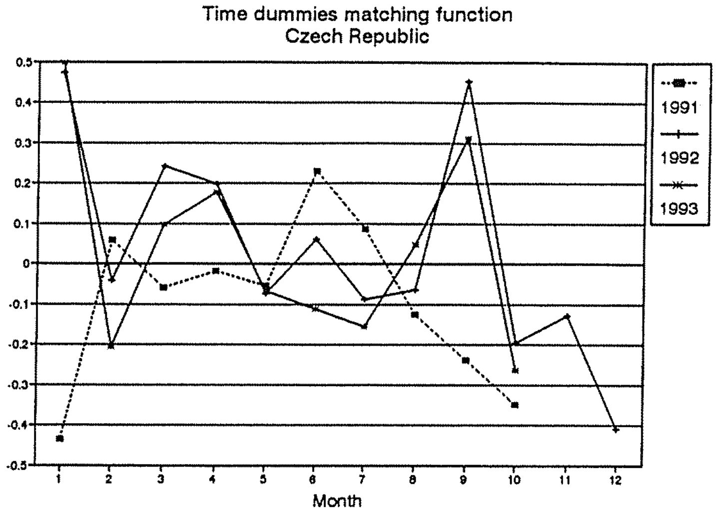
MA(12) using monthly data Jan91 - Oct93
on end of month stocks for U, monthly flows for M

Fig. 16. Slovak Republic



MA(12) using monthly data Oct90 - Sept93
on end of month stocks for U, mothly flows for M

Fig. 17. Czech Republic and Slovak Republic



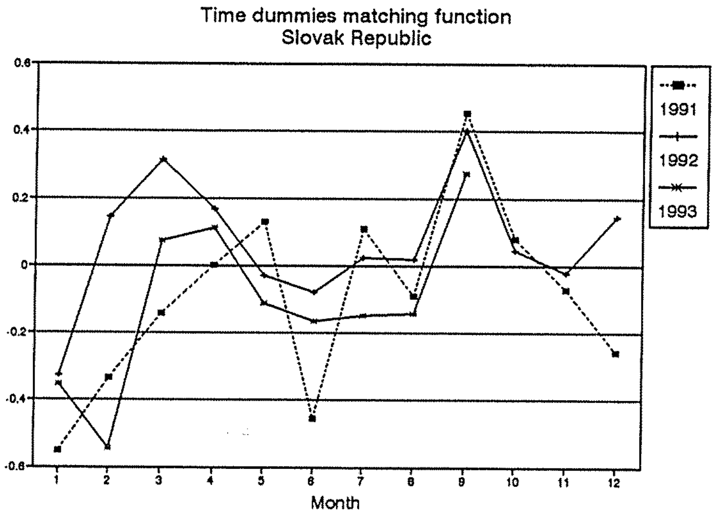


Fig. 18. UV curve

Czech Republic

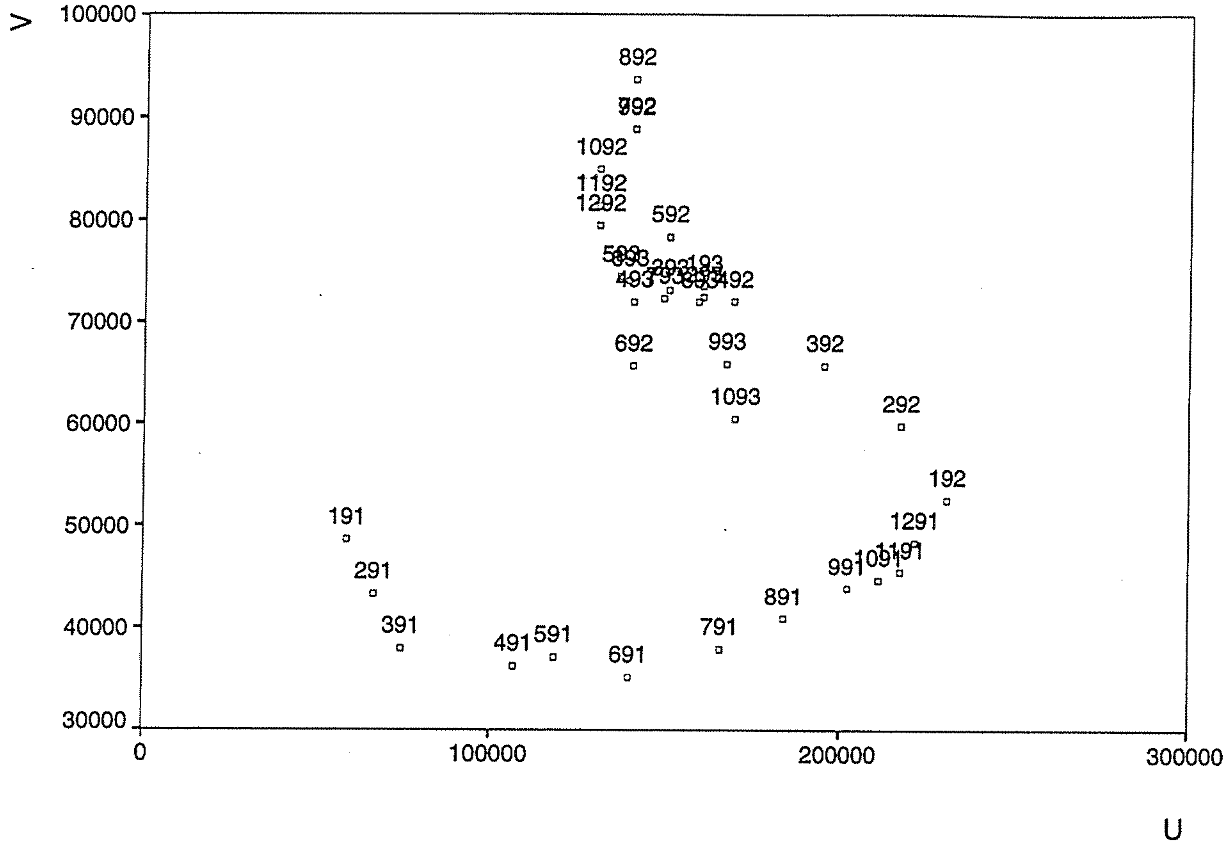
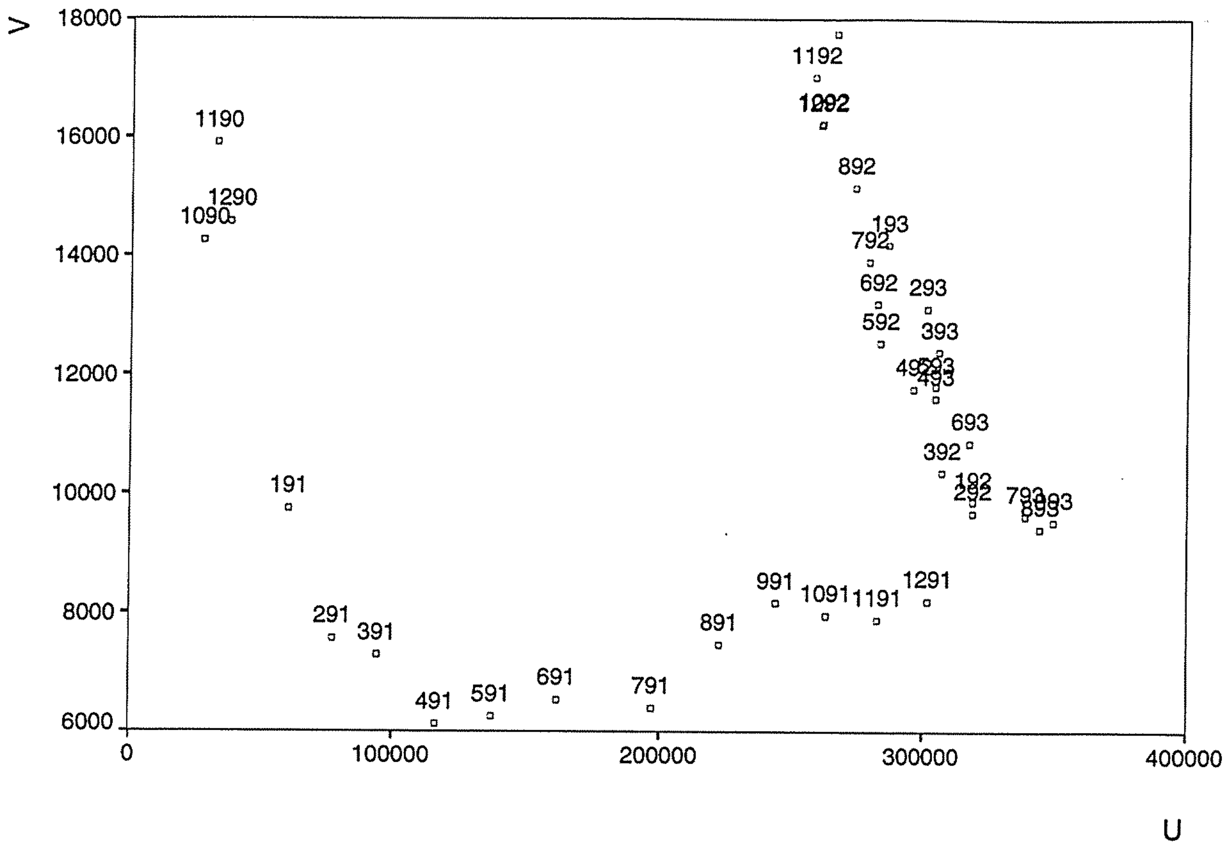


Fig. 19. UV curve

Slovak Republic



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