Returns to Mobility in the Transition to a Market Economy*

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Abstract

In spite of ongoing dramatic changes in labor market structure, we present statistical evidence that transitional economies display rather low worker flows across sectors and occupations. Such low mobility can be explained by low returns to job changes as well as by market segmentation in the allocation of job offers. We develop an econometric model which enables us to characterize intertemporal changes in probabilities of dismissal, remuneration, and offer arrival rates on the basis of information on observed transitions and wage payments. The model is estimated using data from the Polish Labor Force Survey. Our results indicate a significant degree of segmentation in the allocation of job offers, more stability in public sector versus private sector jobs, and little, if any, rewards to tenure and age in the private sector. These findings support explanations for low mobility in transitional economies, which are based on informational failures, notably that fact that job offers do not reach those who are most prone to take up jobs, and that moving from public to private enterprises is costly, especially for those with high levels of job tenure and labor market experience in the public sector.

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1. Introduction

Transition, almost by definition, involves the reallocation of workers across jobs, occupations, and industries. Given the artificial "full employment" conditions inherited from centrally-planned market structures, transition in central and eastern Europe also involved the appearance of open unemployment and significant flows into inactive status. Compared to other countries undergoing rapid structural change – e.g., the Latin American economies – central and eastern Europe offers a wealth of data on labor market flows. Not only are administrative data, e.g., individual records from the registration of jobseekers at labor offices, often made available to researchers, but also most countries have introduced and are currently undertaking household surveys involving rotating panels, which makes longitudinal analysis possible. Such data mainly have been used to describe the magnitude and characteristics of labor market flows or to carry out non-parametric analyses, mainly of hazard rates from unemployment. Such analyses have been useful in characterizing specific features of the labor market adjustment process during economic transformation (e.g., the stagnancy of transitional unemployment) and in identifying those personal characteristics which are most relevant in determining labor market outcomes in the course of the transition process. Preliminary evaluations of the effectiveness of active labor market policies have also been carried out, estimating the impact of participation in active programs [5], [10], [14], [**?**], [6], as well as of unemployment benefit levels and duration on outflows to jobs [8], [16], [10], [9], [4].

These studies have contributed to our understanding of labor market adjustment under rapid structural change and have provided relevant material for evaluations of the impact of labor market programs, which does not necessarily only apply to the East[1]. However, there are a number of issues that still need to be investigated and, more importantly, which *can* be addressed using available data.

First and foremost, very little is known concerning the allocation of job offers across individuals occupying different labor market states. How segmented are the labor markets of transitional economies in conveying information on employment opportunities? Is the probability of being offered a post in the emerging private sector dependent on one's current labor market state? Second, what is the impact on the relationship between wages and education, age, and tenure of the shift

[1] The book recently published by OECD [12] is an attempt to exploit the policy experiments carried out in these countries (e.g., radical changes in the generosity of unemployment benefit systems) in order to draw lessons which can be valuable also for the OECD countries.

between public and private sector employment? Third, to what extent does the risk of dismissal vary between public and private firms? While differences in separation rates of public and private firms do not seem to be that marked, it is possible that the composition of separations varies significantly across firms. For instance, separations from public firms may be mainly related to voluntary quits, whilst the bulk of separations from private units could originate from dismissals.

All these issues are very important in the light of the findings of the empirical literature on labor markets in transitional economies. It has frequently been suggested that low outflows from unemployment to jobs (and a negligible impact of the tigthening of benefits on flows from unemployment to employment) are the byproduct of an aggregate lack of vacancies [3]. However, an aggregate lack of vacancies is unlikely to persist under the current sustained economic recovery, and there are indeed indications that vacancy rates are on the rise in countries like Poland. Thus, the question arises as to whether it is an issue of a lack of vacancies overall or simply of a mis-allocation of job offers not reaching those who are looking for jobs and willing to take up them.

It has also been pointed out that labor markets in these countries display relatively low *churning* rates [2]. Aside from the carry-over of some labor market institutions from the previous regime which rewarded attachment to firms with a battery of social benefits (commodity subsidies, subsidized housing, recreational services, etc.), returns to mobility under the economic transformation may be too low to motivate people to abandon the most protected and unionised jobs in the public sector for more risky jobs in the private sector.

What appears to be lacking in the literature then is an evaluation of the private costs and benefits of job mobility. Such an assessment would greatly improve our understanding of the specific features of labor markets undergoing major structural change. One of the reasons why the large empirical literature on these countries has not yet addressed this issue is that raw data, by themelselves, are often uninformative in this respect. The assessment of the returns to mobility requires an empirical framework which encompasses both labour mobility across states (employment in the public and private sectors and non-employment) and the values associated with occupancy of these states.

This paper marks a first attempt to fill this gap. An econometric model is developed which enables us to characterize intertemporal changes in probabilities of dismissal, remuneration, and offer arrival rates on the basis of information only on observed transitions and wages. This model is implemented using matched data across various waves of the Polish LFS, which is not only the longest (it was begun in 1992), but is also the most complete survey (insofar as it contains wage and some retrospective labor market information).

The results we obtain are plausible and consistent across various LFS waves. To date we have estimated the model only over six quarters of the Polish LFS (three subsamples[2] each linking observations across two consecutive quarters), so that we certainly cannot claim that our empirical analysis has been exhaustive. With the above caveats in mind, our estimates point to significant segmentation in the allocation of job offers, more stability in public sector versus private sector jobs, and little, if any, association between, on the one hand, wages and, on the other hand, tenure and age in the private sector. Were these findings supported in further work, they would be consistent with explanations for low mobility in transitional economies based on informational failures, notably that fact that job offers do not reach those who are most prone to take up jobs, and that changing jobs and moving from public to private enterprises is costly, especially for those with relatively long tenures and work records in the public sector.

The plan of the paper is as follows. Section 2 documents low mobility in transitional economies. Some standard measures of mobility for transition matrices are produced, allowing us to summarize evidence on the extent of mobility across ownership types, sectors, and occupations in transitional economies and in one of the OECD countries with the least mobile labor market, Italy. The extent to which mobility is related to flows from and to non-employment (as opposed to shifts from one job to another) is also discussed. Section 3 develops an econometric model of labor market dynamics tailored for the data available from the Polish LFS. Section 4 presents our results and provides some suggestions for further research. Section 5 concludes by raising a number of issues concerning extensions of our empirical framework enabling to assess, inter alia, the impact of policy changes (e.g., the tightening of unemployment benefit systems) on the way in which labor markets operate.

[2] As explained in Section 2.1., 80 per cent of the LFS sample is retained across two consecutive quarters. We estimated the model over all eligible cases for the period Q3-Q4 1995, while we worked on a random subsample (consisting of 1/6 of the eligible cases) in the other two panel estimations. The parameter estimates are broadly stable across periods, and we believe that the fact of having worked on sub-samples (which greatly reduced data processing efforts) has not affected our results in any substantial way.

2. Low Mobility in Transitional Economies

This section produces and discusses some summary measures of mobility across labor market states, sectors, ownership types, and occupations. In addition to complementing the results from the econometric model developed in Section 3 (which employs smaller samples and covers only Poland), these measures help us to address three important empirical issues raised by the literature on transitional economies.

First, empirical work on labor market transitions in central and eastern Europe has documented a relatively low level of turnover in the unemployment pools in these countries [3]. This stands in sharp contrast to the deep falls in employment observed in the first three to four years of transition and with the scope of ongoing changes in the distribution of employment across sectors, ownership types, and occupations. More recent work has pointed to significant flows from employment to inactivity.[3] This could explain why large employment declines coexist in these countries with low unemployment inflows and outflows. Yet, one still has to explain how radical changes in the structure of employment by sector, occupation and ownership type are being achieved. Are those who are already in employment moving across jobs? Or is it those coming from out-of-the labour force who are taking most of the jobs in the expanding sectors of the economy? We hope to shed some light on this issue on the basis of measures of mobility which alternatively include and exclude non-employment.

Second, the claim has often been made (mainly in an attempt to explain the low vacancy rates observed in these countries) that transitional labor markets tend to display much less churning than their western counterparts [2]. If confirmed by empirical evidence, this claim could explain the coexistence of rapid structural change and low gross worker flows. Low churning would also imply that structural change is being achieved by mobilizing a rather small segment of the working age population, which is relevant information for the design of labor market and social policies in these countries. Measures of mobility for transition matrices (across sectors, occupations, and ownership types) computed from labor market survey data offer a better basis than labor turnover or job turnover data – which are based on administrative records, and hence are affected by incomplete coverage of the small business sector – to assess the magnitude of job churning across countries.

Third, economic transformation, notably the growth of the private sector, is

[3] See the various articles of the (forthcoming) special issue of *Empirical Economics* on "Long-term unemployment and Social Assistance in Central and Eastern Europe".

supposed to significantly modify the way in which labor markets operate. Ongoing reforms of labor market institutions, notably the tightening of the fairly generous unemployment benefit systems introduced at the onset of the transition process and a partial relaxation of employment security schemes, are also likely to involve an increase over time in labor mobility in these countries. This can be readily assessed by analysing how mobility measures behave over time. By computing mobility indexes over as many years as possible, we also hope to remedy a limitation in the empirical analysis which follows, namely its coverage only of the 1994-5 period in Poland. This was a a period of sustained economic recovery, with a buyoant private sector and a much less generous unemployment benefit system than existed at the beginning of the transition period. Our mobility measures may speak to the representitiveness of this particular period vis-a-vis earlier periods in Poland or any period in the other transitional economies.

2.1. Data Issues

Indexes of mobility were computed from transition matrices for a number of countries. In order to make our measures comparable, we used a common procedure in estimating gross flows across states. In particular, we decided to draw on matched records across different LFS waves in all countries rather than using information gathered from the questions which requested retrospective information at a single point in time.

All surveys have a panel component so that the same individuals are interviewed at different points in time. The rotation scheme varies from country to country. In most countries, about 20 per cent of the sample is renewed at each survey date, so that the panel component is 80 per cent of the full sample across two consecutive quarters.

The linking of records was made easier in the Czech Republic, Poland and Slovak Republic by the assignment of unique identifiers to each sampled person. In Hungary identifiers were only provided for the household or the dwelling or there were no identifiers at all, and hence we could only match records across LFS waves on the basis of a battery of reported characteristics, such as the date of birth, the place of residence, and other personal or household characteristics.

The main problem with matched records is that sample attrition, non-response and errors in the classification of the labor market states of individuals at different points in time tend to bias results in a direction which is not predictable *a priori*. In most surveys, households moving from a sampled dwelling are not retained in the sample, but replaced by the new household moving into the originally sampled dwelling. Moreover, non-response is generally higher among movers than stayers and this generates a downward bias in estimated flows . Finally, errors in the classification of the labor market status of individuals tend to be (negatively) serially correlated (that is, response errors at a given survey date are not independent of errors in the previous LFS wave) and this creates many spurious changes in states, leading to over-estimating mobility rates . Another issue related to the use of matched records is that, unlike many retrospective questions, they cannot capture flows occurring between survey dates. In particular, matched records do not capture "roundtripping" across labor market states, which is sometimes significant in these countries. This period-censoring is, clearly, more of an issue when the interval across two subsequent LFS reference weeks is relatively long, say one year. All the measures displayed in Table 1 (and the transitions which offer the basis for the ensuing econometric analysis) are computed on the basis of quarterly transitions.

2.2. Results

Table 1 displays the following standard scalar measure of mobility for transition matrices:

$$I = \frac{s - trace(M)}{s - 1}$$

where s denotes the number of states (the number of rows of the transition matrix, M). As shown by Shorrock (1978), when matrices have a maximal diagonal – that is, stayer coefficients are larger than any individual mover coefficient — this index satisfies a number of desirable properties. In particular, the index is bounded between 0 and 1, is monotonically increasing in mobility, attaches value zero only to identity matrices, and is equal to one for matrices with identical rows (hence probabilities of moving independent of the state originally occupied). All the computed matrices had maximal diagonal, hence in our case the index satisfies the four properties listed above.[4]

States were defined on the basis of internationally agreed (ILO-OECD) definitions of employment, and comparable industry (ISIC) and occupational (ISCO88) classifications, as reported in the notes at the bottom of Table 1.

[4] A problem with this index is that it disregards information on the size of individual "mover" coefficients. We tried other indexes, but our results (and ranking of countries in terms of mobility) did not change.

Four facts, highlighted by Table 1, are particularly relevant. First, there seem to be significant differences in the extent of mobility across central and eastern European countries, with Poland always displaying greater mobility then the other countries. This visual impression is confirmed by tests of homogeneity of transition matrices across countries, reported in the last column on the right-hand-side of Table 1[5]. This may be explained by the fact that Poland is the fastest growing economy in the region. However, mobility measures in Poland are not increasing over time in step with increases in the pace of economic recovery.

Second, and quite strikingly, all transitional economies exhibit less mobility than Italy, a country typically pointed out as having all the ingredients of a rigid labor market with many "jobs for the life" [7].

Third, including or excluding those outside employment – that is, using transition matrices with (open) or without (closed) unemployment and "out of the labor force" status – does not greatly affect the measures of mobility. If most of the shifts were occurring for those individuals who were already employed, we would have expected mobility conditional on being employed to be larger than mobility unconditional on the initial labor market status of the individual. In the case of shifts across the public and private sectors, there are actually indications that mobility -especially in recent years- is significantly larger when non-employed individuals are included.

Fourth, there is no clear trend in mobility measures. Formal tests of stationarity[6] led us to reject the null hypothesis in most cases. But the direction of the change is ambiguous. If anything, mobility would seem to be slightly declining in recent years.

All this confirms the view of transitional labor markets as persistently characterised by relatively low churning levels. Why is churning so low? Is it costly to change firms, sectors or occupations compared with the benefits one can get from

[5] The tests conducted are likelihood ratio tests. Country-specific transition matrices are first computed for each time period. The restricted transition matrix is estimated pooling information *across countries* at each point in time. The likelihood ratio test statistic is twice the difference in log likelihoods associated with the unrestricted and restricted models, which under the null of homogeneity is distributed as a χ^2 random variable with $(K-1)(S)(S-1)$ degrees of freedom[1], where K is the number of countries and S is the number of states.

[6] Once again, these tests are based on likelihood ratio statistics. For each country, a transition matrix is estimated based on data pooled *over time*. The unrestricted model consists of the period-specific transition matrices. Let there be T time periods. Then twice the difference in the log likelihoods associated with the unrestricted and restricted models is distributed as a χ^2 random variable with $(T-1)(S)(S-1)$ degrees of freedom under the null of time invariance of the transition matrix.

the move? Are offer arrival rates disproportionately allocated to those who are less prone to move? The next two sections will try to provide some provisional answers to such questions.

3. An Econometric Model of Labor Market Dynamics in the Transition

The econometric model we develop in this section has been specifically designed with the data available in the Labor Force Surveys of central and eastern European countries. As we plan to initially implement the model on Polish data, some features of the model particularly fit the design of the Polish LFS, but can be amended and adapted for use with the data available in other transitional economies.

As discussed above, an advantage of the Polish LFS with respect to other household surveys used in economies in transition is that it contains information on the wages of individuals. Hence we can obtain more precise estimates of some of the key parameters, notably those characterizing returns to tenure and the standard deviation of the distribution of wage offers in the public and private sectors.

Our objective is to characterize intertemporal changes not only in remuneration, but also in employment opportunities, across two sectors of the economy, broadly classified as the state (g) and the private (p) sectors. Based on this characterization it is then possible to assess the costs and benefits associated with shifts from one sector to another.

The dependent variables of the model are period-to-period transition rates between discrete labor market states and observed wage outcomes, which are themselves the outcome of simple maximizing decisions made by labor market participants. The nature of the decision rules is discussed in some detail below. The advantages of this model are twofold. First, it enables us to combine labor market transition data with wage data in a natural way. Second, by positing a simple selection mechanism, it allows for consistent estimates of population parameters which are not otherwise obtainable from observations on wage outcomes associated with *chosen* alternatives. Of course, the consistency of our model estimates hinges critically on the validity of the choice mechanism we use in formulating the model. Unfortunately, given the data at hand, it does not appear possible to test the validity of the choice mechanism we posit.[7]

In defining the model the following notation is used:

> ξ denotes the sector of the economy from which the wage offer is made, where $\xi = g$ or p.
>
> X is a $(1 \times K)$ vector of time-invariant characteristics of the labor market participant.
>
> τ denotes tenure in the job, that is, the number of periods the individual has worked at his/her current employer.
>
> t denotes the time period

We specify that an individual who has worked at a specific firm in sector σ for a total of τ periods and who has characteristics X is offered a $\ln(wage)$ in period t of:

$$\ln(w(\xi, X, \tau, t)) = X\beta(\xi, t) + \delta(\tau, \xi, t) + \varepsilon(\xi, t),$$

where

> $\beta(\xi, t)$ is a $(K \times 1)$ vector of regression coefficients which vary by sector and time period
>
> $\delta(\tau, \xi, t)$ is a function of the agent's tenure level at the firm from sector ξ in period t making the offer. If the firm making the offer is not the agent's current employer, then $\tau = 0$ and $\delta(0, \xi, t) = 0$ all ξ and t.
>
> $\varepsilon(\xi, t)$ is an independently [over time] distributed, mean zero shock which is normally distributed in each period t. The shock is independently distributed across all firms in the market. The variance of the shock among private-sector firms in period t is $\sigma_{pp}(t)$ and the variance of the shock among state-sector firms is $\sigma_{ss}(t)$.

Time is discrete throughout our model. In each period, a labor market participant occupies one of three mutually exclusive and exhaustive states. Either s/he is nonemployed, employed by a firm in the state sector, or employed by a firm in the private sector. Let $l(t)$ denote the individual's labor market state in period t, with $l(t) = n, g$, or p. At the beginning of the next period, $t + 1$, the

[7] Thus our assumptions serve to "just identify" the model. This is not an uncommon situation when the analyst has access to rewards associated with states nonrandomly chosen by agents while the values of the alternatives not chosen are unobserved.

individual may either face dismissal from his period t job, if he was employed in period t, or may receive other offers of employment which s/he will then accept if they are preferable to the wage offered. Let $\pi(l(t), j, t)$ denote the probability that an individual in state $l(t)$ at time t experiences event j, where j is one of the following events:

1. dismissal occurs
2. no new job offer received
3. a job offer from a state-sector firm is received
4. a job offer from a private-sector firm is received

Certain combinations of events are logically impossible. For example, it is meaningless to speak of a nonemployed individual being "dismissed" from his state, so that $\pi(n, 1, t) = 0, \ \forall t$. In terms of other events, note that any of a number of situations are allowed in this structure. Let us consider the case of an individual employed by a private-sector firm in period t. Then $\pi(p, 1, t)$ is the probability that this individual loses his job at the end of period t and thus enters the nonemployment state in period $t+1$. The probability that an individual is not dismissed and receives no new wage offers during t is given by $\pi(p, 2, t)$. In this situation, we will assume that the individual stays at his period t private-sector job through period $t+1$.[8] $\pi(g, 3, t)$ is the probability that the agent receives a job offer from a public-sector firm at the beginning of period $t+1$ in addition to retaining his option of continuing to work in his (private-sector) period t job. Below we will characterize the decision rule s/he is assumed to utilize in making the choice between the two employment opportunities. Analogously, $\pi(p, 4, t)$ is the probability that the individual receives an offer from another private-sector firm. In period $t+1$ s/he then chooses either to continue to do her work at her (period t) private-sector employer or to switch to the other private-sector firm.

3.1. Decision Rules

As we mentioned in the introduction to this section, the decision rules are very much "static" in nature, in the sense that options are chosen based on current

[8] Given our assumptions concerning the value of the nonemployment state and the estimated parameter values, the choice never to quit into nonemployment is a utility-maximizing one. In future we plan to allow for voluntary quits into nonemployment.

period returns, and not on the anticipated future values associated with the options. There are three reasons we have chosen to work with this specification of the decision-making process.

First, a case can be made for myopic decision rules from the perspective of "realism." In a rapidly changing environment, individuals may find it impossible to assess the probabilities of future states of nature, since the possible states of nature which may appear in the future cannot be anticipated. Thus, in attempting to assess the future evolution of the labor market, Polish workers, especially in the early 1990s, may have found themselves in a situation in which there exists "Knightian uncertainty." When remuneration rates and dismissal and offer probabilities are moving in "unpredictable" ways, static decision rules may be the only ones which can be implemented.

Second, the modelling of static decision rules is motivated by data availability, notably the manner in which the sample is drawn. At most, we are able to observe the labor market movements of sample members and their compensation rates for an 18 month period. Then estimating a life-cycle model of labor market dynamics would require us to make very strong assumptions concerning the behavior of individuals and the environment they are likely to face over periods of substantial length.

A third rationale for the use of myopic rules in this particular case is the low asset position of most of the labor market participants and the very high inflation rates they face, particularly in the early stages of transition. In such a situation agents may effectively discount future rewards associated with any given option at a very high rate. Thus, a truly forward-looking rule, should one be feasible, may be well-approximated by a myopic one.

We are now ready to turn to the actual specification of the rules. First, we consider the decision rules utilized by an agent who was employed at a firm in sector ξ firm at time t and who receives an offer from another firm in sector ξ' in period $t+1$. Say that in period t the agent had τ units of tenure at his or her firm. Then in period $t+1$ the log wage offered by his period t employer is

$$\ln(w(\xi, X, \tau+1, t+1)) = X\beta(\xi, t+1) + \delta(\tau+1, \xi, t+1) + \varepsilon(\xi, t+1).$$

The log wage offer of the other firm is given by

$$\ln(w(\xi', X, 0, t+1)) = X\beta(\xi', t+1) + \varepsilon(\xi', t+1),$$

where the reader should note that the tenure level at the firm making the new offer is by definition equal to 0. For individuals with two wage offers, we assume the one offering the higher period $t+1$ log wage is the one selected. Then the individual *changes employer* in period $t+1$ if and only if

$$\ln(w(\xi', X, 0, t+1)) > \ln(w(\xi, X, \tau+1, t),$$

or

$$X[\beta(\xi', t+1) - \beta(\xi, t+1)] - \delta(\tau+1, \xi, t+1) > \varepsilon(\xi, t+1) - \varepsilon(\xi', t+1). \quad (3.1)$$

For notational simplicity, we will define the random variable

$$\eta(\xi, \xi', t) \equiv \varepsilon(\xi, t) - \varepsilon(\xi', t).$$

Then $\eta(\xi, \xi', t)$ is a normally distributed mean zero random variable in each period t. The variance of this random variable is given by $\sigma_{\xi\xi}(t) + \sigma_{\xi'\xi'}(t)$ under our independence assumptions. We will denote the *standard deviation* of $\eta(\xi, \xi', t)$ by $\psi(\xi, \xi', t)$.

Given that an agent can stay at his or her period t employer or move to a new firm in sector ξ' at time $t+1$, the probability that s/he will change is

$$\Phi\left(\frac{X[\beta(\xi', t+1) - \beta(\xi, t+1)] - \delta(\tau+1, \xi, t+1)}{\psi(\xi, \xi', t+1)}\right).$$

Note that if the alternative job offer is from a firm in the same sector as the agent's period t employer, than the probability of turnover is only a function of the agent's tenure level at their period t employer, or

$$\Phi\left(\frac{-\delta(\tau+1, \xi, t+1)}{\psi(\xi, \xi', t+1)}\right).$$

If the function δ is increasing in τ, the probability of turnover will be a decreasing function of tenure. We will estimate the function δ parametrically in what follows.

We have described the turnover rules for employed labor market participants, and now turn our attention to the rules utilized by unemployed agents. In any period t, an unemployed agent may receive no offer of employment, may receive an offer from a private-sector firm, or may receive an offer from a state-sector firm. As before, let the sector of the firm making the job offer be given by ξ. Then we posit the existence of a vector, $\beta^*(t)$, which is used to determine the value of the unemployment state to a type X individual. The individual is assumed to compare the $\ln(w)$ offer with the function $X\beta^*(t)$ and accept the offer when

$$\ln(w(\xi, X, 0, t)) > X\beta^*(t)$$

or

$$\varepsilon(\xi, t) > X[\beta^*(t) - \beta(\xi, t)].$$

Then the probability that an individual of type X accepts a sector ξ wage offer in period t is

$$1 - \Phi\left(\frac{X[\beta^*(t) - \beta(\xi, t)]}{\sqrt{\sigma_{\xi\xi}(t)}}\right).$$

Note that we have assumed that there is no random component to the value of nonemployment for an individual. Furthermore, while the parameter vector $\beta^*(t)$ is in principle identified given the data available to us, we decided to normalize the value of $\beta^*(t)$ to 0 for all t so as to enhance interpretability of the estimates of the other parameters of the model.

3.2. Implications of the Model for Labor Market Transitions

Before turning to estimation issues, it will be useful to consider what the model implies in terms of observed patterns of labor market dynamics. Consider first the case of an agent who is unemployed at time t. Next period we may find him or her in the state of nonemployment, working for a public sector firm, or working for a private sector firm. The probability of finding him or her unemployed is the probability that no offer was received plus the probability that unacceptable public- or private-sector offers were made, thus

$$\begin{aligned}\Lambda(n, n, X, t) &= \pi(n, 2, t) + \pi(n, 3, t)\Phi\left(\frac{X[\beta^*(t+1) - \beta(g, t+1)]}{\sqrt{\sigma_{gg}(t+1)}}\right) \\ &\quad + \pi(n, 4, t)\Phi\left(\frac{X[\beta^*(t+1) - \beta(p, t+1)]}{\sqrt{\sigma_{pp}(t+1)}}\right).\end{aligned}$$

The probability that a nonemployed agent will find a job in the public sector by period $t+1$ is

$$\Lambda(n, g, X, t) = \pi(n, 3, t)\tilde{\Phi}\left(\frac{X[\beta^*(t+1) - \beta(g, t+1)]}{\sqrt{\sigma_{gg}(t+1)}}\right),$$

where $\tilde{\Phi}(z) \equiv 1 - \Phi(z)$ is the survivor function. Similarly, the probability of transition from nonemployment to employment in the private sector is

$$\Lambda(n,p,t) = \pi(n,4,t)\tilde{\Phi}\left(\frac{X[\beta^*(t+1) - \beta(p,t+1)]}{\sqrt{\sigma_{pp}(t+1)}}\right).$$

Next consider the transition probabilities for an individual employed in the public sector in period t. The probability of a transition into nonemployment is simply the probability of dismissal, or

$$\Lambda(g,n,t) = \pi(g,1,t).$$

The probability of remaining in a public sector job is equal to the sum of the probabilities of staying in the same public sector job and the probability of accepting a new public sector job. The probability of staying in the same public sector job is equal to the probability of not receiving any alternative offers (and not being dismissed) plus the probabilities of receiving unacceptable offers from other firms. Then we have

$$\begin{aligned}
\Lambda(g,g,X,\tau,t) &= \{\pi(g,2,t) + \pi(g,3,t)\tilde{\Phi}\left(\frac{-\delta(\tau+1,g,t+1)}{\psi(g,g,t+1)}\right) \\
&\quad +\pi(g,4,t)\tilde{\Phi}\left(\frac{X[\beta(p,t+1) - \beta(g,t+1)] - \delta(\tau+1,g,t+1)}{\psi(g,p,t+1)}\right)\} \\
&\quad +\pi(g,3,t)\Phi\left(\frac{-\delta(\tau+1,g,t+1)}{\psi(g,g,t+1)}\right) \\
&= \pi(g,2,t) + \pi(g,3,t) + \\
&\quad \pi(g,4,t)\tilde{\Phi}\left(\frac{X[\beta(p,t+1) - \beta(g,t+1)] - \delta(\tau+1,g,t+1)}{\psi(g,p,t+1)}\right),
\end{aligned} \tag{3.2}$$

where the term in brackets on the RHS of the first line of [3.2] is the probability of staying at the period t public sector job and the additional term is the probability of moving to another public sector firm.

Finally, the probability of a transition to a private-sector job from period t public sector employment is

$$\Lambda(g,p,X,\tau,t) = \pi(g,4,t)\Phi\left(\frac{X[\beta(p,t+1) - \beta(g,t+1)] - \delta(\tau+1,g,t+1)}{\psi(g,p,t+1)}\right).$$

The derivation of the transition probabilities given period t private sector employment is similarly performed, and for purposes of brevity is omitted.

It is important to note that *virtually all the parameters in our model are estimable using only data on transitions between observed labor market states.*[9] Information on wage payments from firms in the private and public sector are *not* strictly required for the identification of most parameters, though this type of information can be expected to greatly increase the precision of estimates of all parameters in the model, most especially $\beta(\xi, \cdot)$ and $\delta(\cdot, \xi, \cdot)$.

3.3. Identification and Estimation Issues

The data to which we have access consists of information on the labor market states and demographic characteristics of individuals at two points in time.[10] The sampling scheme for the Polish Labor Force Survey has selected households interviewed for two consecutive quarters, then omitted for two consecutive quarters, and then put back into the sample for two consecutive quarters. As there are four panels, the overlap between any two consecutive quarters (and years) is 50 per cent.

In order to consistently estimate the econometric model developed in this paper it is not necessary to use all four quarters of sample information for each household member participating in the labor market. For simplicity, we have constructed samples which contain information from two consecutive quarters only. We will refer to these quarters as t and $t+1$.

The key to estimating the model is recognizing that it is first-order Markov - that is, the probability of transiting from state m [his or her period t labor market state] to state m' [his or her period $t+1$ labor market state] is solely a function of his or her "type" (X) and, if employed at time t, his or her tenure at the job held at t.

Likelihood contributions are defined as follows. Consider first the case of individuals who were nonemployed in the first period [$m = n$]. If such an individual is also nonemployed in the second period [$m' = n$], the log likelihood contribution is simply $\ln(\Lambda(n, n, X, t))$. If a nonemployed individual accepts a job in the public sector, their log likelihood contribution is determined as follows. Given that the new job was chosen by the individual, we have that the [conditional] accepted

[9] The principal exceptions being the variance parameters, $\sigma_{pp}(\cdot)$ and $\sigma_{gg}(\cdot)$, which would have to be normalized.

[10] In this preliminary analysis only two adjacent time points are utilized. We are currently working on incorporating more observations per respondent and relaxing some of the distributional assumptions made in the present analysis.

wage density is given by

$$f(w_g|w_g > X\beta^*(t{+}1), t{+}1) = \sigma_{gg}(t{+}1)^{-.5}\frac{\phi((w_g - X\beta_g(t+1))/\sigma_{gg}(t+1)^{.5})}{\Phi(X[\beta_g(t+1) - \beta^*(t+1)]/\sigma_{gg}(t+1)^{.5})}.$$

Now the joint density/probability of the public sector wage offer and the event that it is acceptable is

$$\begin{aligned} f(w_g|w_g &> X\beta^*(t+1), t+1) \times \Pr(w_g > X\beta^*(t+1); t+1) \\ &= \sigma_{gg}(t+1)^{.5}\phi((w_g - X\beta_g(t+1))/\sigma_{gg}(t+1)^{.5}). \end{aligned}$$

Finally, the joint density/probability of a public sector wage, the event that it is acceptable, and the event that it is received, is

$$\begin{aligned} f(w_g|w_g &> X\beta^*(t+1), t+1) \times \Pr(w_g > X\beta^*(t+1); t+1) \times \pi(n,3,t) \\ &= \sigma_{gg}(t+1)^{-.5}\phi((w_g - X\beta_g(t+1))/\sigma_{gg}(t+1)^{.5}) \times \pi(n,3,t), \end{aligned}$$

where ϕ denotes the standard normal probability density function. The log of this expression is the contribution to the log likelihood function made by a nonemployed individual with characteristics X who transits from nonemployment to public-sector employment at a wage w_g at time $t+1$. Similarly, the log likelihood contribution of a nonemployed individual who enters private-sector employment at wage w_p is given by the log of

$$\sigma_{pp}(t+1)^{-.5}\phi((w_p - X\beta_p(t+1))/\sigma_{pp}(t+1)^{.5}) \times \pi(n,4,t).$$

The log likelihood contributions for individuals employed at the baseline date are only slightly more complicated. Consider the case of an individual who is employed at a public sector firm at date t and at that time has τ units of tenure. If s/he is terminated from his or her job by time $t+1$, the log likelihood contribution is simply $\ln(\Lambda(g,n,t)) = \ln(\pi(g,1,t))$. If the individual is working at his old firm in period $t+1$, we know that this could occur either because s/he was not offered any new job, s/he was offered an unacceptable job by a different public sector firm, or s/he was offered an unacceptable job by a private sector firm. If the individual was not offered another job in the period, then there are no selection issues with respect to the period $t+1$ wage at their old employer. This is not the case if an alternative offer was received. Say that the alternative offer came from another public sector firm. Then for it not to be accepted, it must be the case that $w'_g \leq w_g \Rightarrow \varepsilon'_g \leq \delta(\tau+1, g, t+1) + \varepsilon_g$. Now the joint density of ε'_g and ε_g can be written as $f(\varepsilon'_g|\varepsilon_g)f(\varepsilon_g) = f(\varepsilon'_g)f(\varepsilon_g)$ under our independence assumptions. Then the probability that the other public-sector offer was unacceptable can be written as

$$\Pr(\varepsilon'_g < w_g - X\beta_g(t+1)) = \Phi\left(\frac{(w_g - X\beta_g(t+1))}{\sigma_{gg}(t+1)^{.5}}\right).$$

Similarly, the agent may have received an unacceptable offer from a private-sector firm. The probability of this event is

$$\Pr(\varepsilon'_p < w_g - X\beta_p(t+1)) = \Phi\left(\frac{(w_g - X\beta_p(t+1))}{\sigma_{pp}(t+1)^{.5}}\right).$$

Thus the joint probability distribution/density of staying at the last period's public-sector employer at the new observed wage w_g is

$$\sigma_{gg}(t+1)^{-.5}\phi((w_g - X\beta_g(t+1) - \delta(\tau+1, g, t+1))/\sigma_{gg}(t+1)^{.5})$$
$$\times \{\pi(g,2,t) + \pi(g,3,t)\Phi\left(\frac{(w_g - X\beta_g(t+1))}{\sigma_{gg}(t+1)^{.5}}\right) + \pi(g,4,t)\Phi\left(\frac{(w_g - X\beta_p(t+1))}{\sigma_{pp}(t+1)^{.5}}\right).$$

The log of this expression is such an agent's log likelihood contribution.

For an individual who switches public sector firms between t and $t+1$, let the (observed) wage offer at the new firm be denoted w'_g. Then given w'_g, the probability that the old public-sector employer's offer was less than this amount is

$$\Phi\left(\frac{(w'_g - X\beta_g(t+1) - \delta(\tau+1, g, t+1))}{\sigma_{gg}(t+1)^{.5}}\right).$$

Then the log likelihood contribution of such an individual is the log of

$$\sigma_{gg}(t+1)^{-.5}\phi((w'_g - X\beta_g(t+1))/\sigma_{gg}(t+1)^{.5}) \times \pi(g,3,t)$$
$$\times \Phi\left(\frac{(w'_g - X\beta_g(t+1) - \delta(\tau+1, g, t+1))}{\sigma_{gg}(t+1)^{.5}}\right).$$

Finally, using similar arguments, the log likelihood contribution of an individual who switches from public-sector to private-sector employment at a wage w'_p is given by the log of

$$\sigma_{pp}(t+1)^{-.5}\phi((w'_p - X\beta_p(t+1))/\sigma_{pp}(t+1)^{.5}) \times \pi(g,4,t)$$
$$\times \Phi\left(\frac{(w'_p - X\beta_g(t+1) - \delta(\tau+1, g, t+1))}{\sigma_{gg}(t+1)^{.5}}\right).$$

The derivation of the log likelihood contributions for individuals who were private-sector employers in the baseline period is identical (after switching the appropriate subscripts).

The model was estimated using parametric maximum likelihood methods. The structural parameters are the vector of probabilities associated with random choice sets. Given the fact that these transition rates must sum to unity, there are eight free parameters to estimate here, namely $\pi(n,3), \pi(n,4), \pi(g,2), \pi(g,3), \pi(g,4)$, $\pi(p,2), \pi(p,3)$, $\pi(p,4)$. In addition, we estimate the standard deviations of the wage shocks in the public and private sectors [$\sigma_{gg}(t+1)^{.5}$ and $\sigma_{pp}(t+1)^{.5}$]. We parameterize the "returns to tenure" function δ as a quadratic, so that $\delta(\tau,\xi,t) = \delta_1(\xi,t)\tau + \delta_2(\xi,t)\tau^2$. When we estimate models which include tenure effects, this adds 4 parameters (two for each sector). We have chosen to normalize the value of nonemployment to 0 by setting $\beta^*(t) = 0$ for all t. While this parameter is theoretically identifiable under our model assumptions, in practice it is difficult to do so [i.e., the standard errors associated with estimates of the elements of the vector $\beta^*(t)$ are very large]. Interpretation of the parameter estimates should be made with this normalization in mind. In particular, given the other estimates of the model, the implication is that all offers received from private-sector or public-sector firms by nonemployed individuals will be accepted. Given that the sample is comprised of 30-50 year old males and an ample empirical literature pointing to an overall lack of vacancies as the main determinant of unemployment duration, this implication may not be grossly at odds with reality. The remaining parameters consist of the elements of $\beta_g(t)$ and $\beta_p(t)$. All of the elements of these vectors are identified given the sample information available to us.

4. Sample Description and Empirical Results

Below we present estimates of the econometric model developed in the previous section using subsamples of observations from the Polish Labor Force Surveys over an 18 month period beginning in the third quarter of 1994 and ending in the fourth quarter of 1995. While we have access to LFS data beginning in 1992, a number of variables required for the estimation of the model are not available until this later period.[11] We have estimated the model for three matched samples

[11] The design of the questionnaire, the coding of variables, and the rotation scheme (initially it was to be a full panel with no missing quarters) have been frequently changed, making it difficult to construct labor market histories and reducing the over-time comparability of observations on the occupational and sectoral affiliations.

of individuals: those matched for Q3 and Q4 of 1994, those matched for Q1 and Q2 of 1995, and those matched for Q3 and Q4 of 1995.

Prior to estimating the model we implemented a number of sample inclusion restrictions. So as to bypass consideration of the labor market participation decision, we have only utilized information for males between the ages of 30 and 50, inclusive.[12] A benefit of the Polish LFS is the availability of wage information; the problem is that there is a large amount of missing wage information for individuals reporting that they are currently employed [at the time of the interview]. From our experiences with the three subsamples utilized below, it seems that roughly 30% of otherwise eligible individuals did not report a wage rate. We choose simply to exclude such individuals from the analysis. Since only employed respondents were subject to this exclusion restriction, the proportion of employed individuals in the final sample is most probably smaller than it is in the population. While this poses problems for the interpretation of estimates of transitions between the nonemployment state and the employment states, these are not our primary focus of interest here. Since we are primarily interested in transitions between public-sector and the private-sector employment, the problem of missing wages may not be of great concern if the missing wage pattern by sector is approximately random. True shifts of workers between public and private sectors were disentangled from simple ownership changes by using job tenure information.[13]

Table 2 reports descriptive statistics for the three samples. As can be seen from the first three rows of the table, unemployment has been roughly stable from Q3 94 to Q4 95 (however, it is slightly increasing in the fourth quarter of each year due to seasonal factors), while the employment share of the private (public) sector has increased (decreased) over this period. This is consistent with aggregate data pointing to the increasing importance of private sector employment. Note that

[12] The current statutory retirement age in Poland is 60 for women and 65 for men, while the minimum statutory contribution rate (allowing one to work and get a pension) is 25 years. As a result of rather liberal rules concerning pre-retirement, an extensive use of early retirement schemes and no decrual rates for those retiring before reaching the retirement age, the actual age of retirement significantly dropped – relative to the statutory retirement age - at early stages of transition. In 1992, the actual average age of retirement was 57. The average age of disability pensioners was 46.

[13] In particular, individuals reporting employment in a firm of ownership type ξ in quarter $t+1$ and employment in a firm of different ownership type ξ' at time t were dropped from the sample if they reported a job tenure greater than 4 months at time $t+1$. While the analysis of the wage effects of changes in firm ownership is potentially of significant interest, there were too few observations on this event in our samples to justify modifying the model to incorporate such an eventuality [less than 3 percent of sample members were excluded due to this criterion].

the standard deviation (reported in parenthesis) of private sector wages is larger than the standard deviation of wages paid on public-sector jobs. Average tenures are declining throughout this period, but the decline is more marked in the private than in the public sector.

Table 3 displays the estimated transition matrices across the various labor market states. The notation $p_{ss'}$ denotes the probability that an individual in state s in the first quarter is observed in state s' in the second. At the bottom of the table, the notation $p_{s \nrightarrow s}$ denotes the probability that an individual who stays employed in sector s in quarter one and two does not change firms, while $p_{s \rightarrow s}$ denotes the probability that an individual employed in the same sector at the two points in time is employed by a different firm at the two interview dates.

The probability of escaping from unemployment are relatively low in all samples. For example, the value for p_{nn} estimated for Q3-Q4 94 implies an average unemployment duration of about 14 quarters. For those managing to find employment, the proportion finding private-sector jobs exceeds the proportion finding public-sector jobs except in the last sample, where the split is even. For those employed in the first quarter, the probability of a dismissal is substantially higher if the individual originally held a private-sector job. Moreover, individuals are much more likely to transit from private-sector jobs into public-sector ones than vice versa. In the end, public-sector employment tends to be more persistent than public-sector employment in all samples. Also note that individuals in public-sector employment in both quarters are less likely to change firm than individuals employed in private-sector jobs at both points in time. All these estimates seem to lead to a conclusion that private-sector employment is less stable than is public-sector employment.

Before examining the estimates of the model, a few words of caution are in order. Due to the very compressed time period over which our three samples have been collected, it is more than a bit dangerous to attempt to discern systematic movements in the parameters characterizing the econometric model, though on occasion we may do so. We should also alert the reader to the fact that, while the sample for Q3 and Q4, 1995 contains all eligible respondents in the LFS for that period, the samples for the other two periods consist of (randomly-selected) subsamples of all eligible respondents in the LFS for those periods.[14] The final sample size for Q3-Q4, 1995 is 2,516, while the sample sizes for Q3-Q4, 1994 and Q1-Q2, 1995 are 367 and 465, respectively. The relatively small sizes of these two samples can be expected to result in parameter instability and relatively large

[14] We restricted the size of these samples for purposes of computational expediency only.

estimated [asymptotic] standard errors. This is particularly the case with respect to parameters associated with labor market transitions which occur relatively infrequently in the population. Nonetheless, the relative stability of the estimates across quarters leads us to believe that the estimates from all samples are credible.

Bearing the above caveats in mind, we can now turn to a summary of the estimates. The compensation measure used is the natural logarithm of the individual's wage rate. We will consider the estimates sample by sample in chronological order. Table 4 contains the estimates for the Q3-Q4, 1994 sample. For this sample, and for each of the others, we have estimated two specifications. The first specification includes no regressors in the wage offer function and does not allow for tenure effects in this function either. The second specification includes age as a regressor and allows for linear and quadratic tenure effects on current wage offers from the previous period employer. We have not included additional regressors, such as schooling levels, at this point due to some stability problems we had in solving the likelihood equations. In the future, we intend to include regressors such as schooling in both the wage offer and the arrival rate functions.

Considering first the estimates reported in the basic model [Specification 1], we see that the rate of arrival of offers from private-sector firms is substantially larger than the rate of arrival of offers from public-sector firms for individuals who were not employed at the baseline interview. Recall that, given our normalization of the value of nonemployment to zero, all these offers would be accepted, which implies that the estimated values of $\pi(n, 3)$ and $\pi(n, 4)$ come directly from the observed transition matrix. It is of interest to compare the probability of receiving wage offers from alternative firms conditional on the sectoral identification of the respondent's baseline employer. Individuals who were employed at a public-sector firm in the baseline period had a probability of receiving an employment offer from another public-sector firm of .289; an individual employed by a private-sector firm at the baseline had only a probability of .091 of receiving such an offer. Conversely, private-sector employees had a probability of .377 of receiving an offer from another private-sector firm, whereas public-sector employees had only a probability of .053 of receiving such an offer. This pattern suggests that the labor market may be segmented in the sense that search behavior and network formation [for employment contacts] tend to be located in one sector or the other.

We find that the dismissal rate from private-sector jobs [.118] is substantially greater than is the dismissal rate from public-sector firms [.021] (these estimates also come directly from the transition matrix estimates in Table 3 given our assumptions on the dismissal process). There is another sense in which public-sector jobs are less "risky" then private sector ones - the population standard deviation of the disturbance term in the log wage equation is .392 for public-sector jobs but is .502 for private-sector jobs. The mean log wage offer in the public-sector is significantly larger than the mean log wage offer in the private-sector. Of course, this could merely be an artifact of there being substantially different types of people in the two sectors. In the second specification of the model we allow for differences in the ages of participants and their tenure level at last period's employer [if remaining at that employer is a current period option] to shift the wage offer function. The standard deviation of the wage shocks continues to be significantly larger in the private than in the public sector.

The estimated event arrival rates from the "full" model are broadly similar to those found in the base model that doesn't include age or tenure effects. The only event probabilities substantially impacted by a respecification of the wage offer function are those connected with job-to-job changes. This is to be expected, since identification of the individual effects of event arrival rates and the wage offer distribution is only obtained through functional form assumptions. We now find that the "segmentation" of the markets in terms of offer arrival rates is even more noticeable than before. It is possible that this segmentation is partly a byproduct of a lack of circulation of information over vacancies across regions. There is ample evidence of marked differentials across regions in the allocation of vacancies – with most vacancies concentrated in urban areas where the private sector is concentrated – and plans to computerise the Polish Public Employment Service (allowing it to provide on-line vacancy registers covering the country as a whole) have not yet been completed.

We find that there are strong age effects in the wage offer function associated with the public sector, though this is not the case for the private sector. We also find strong tenure effects in log wage offers for public-sector jobs, but not for private-sector ones. Thus the log wage offer distribution associated with private-sector jobs is essentially time-invariant, while this is most definitely not the case with respect to public-sector jobs. A possible interpretation for this result is that, prior to 1989, private-sector non-agricultural employment accounted for barely 5 per cent of employment and family work in agriculture (by and large the dominant component of private sector employment at the beginning of the decade) typically does not reward tenure and age. Another explanation we cannot rule out is that contracts in the private sector – significantly less unionised than the public sector and still largely concentrated in services – do not reward seniority in the firm and actually do not value previous work experience (perhaps because the human capital inherited from the previous regime is deemed obsolete or most new jobs require low levels of skill). What is clear from our estimates is that the log wage-tenure profile is concave in public-sector firms.

Many of the patterns we have discussed so far are also found in the estimates for the Q1-Q2, 1995 sample, though some appear not to be present. For example, there is no indication of a significant difference in flows from nonemployment to public-sector and private-sector firms, though the private-sector offer arrival rate is still slightly higher than the public-sector rate. We continue to find that dismissals occur more frequently from private-sector firms than from public-sector ones. We also continue to find evidence that the public and private sector markets are segmented in terms of offer arrivals, with a public-sector employee much more likely to receive an offer of a public-sector job than a private-sector job [.153 versus .010], with the converse being true for private-sector employees [.031 versus .183]. For this period, dispersion in the disturbance term of the log wage equation is essentially identical in the two sectors. In the model which includes age and tenure effects, we find once again that these terms have little influence on log wage offers in the private sector but do impact public-sector offers in much the same way as was found for the previous sample.

The results for the last period are reported in Table 6. As was true for the early 1995 sample, there is little difference in the probability of exit from unemployment into public-sector and private-sector jobs, though in this case exit into a public-sector job is slightly more likely in contrast to our estimates from the other periods. We continue to find the same concentration of employment opportunities that has been remarked upon previously. We also continue to find that the dismissal rate from private-sector jobs is much higher than the dismissal rate for public-sector jobs [.098 versus .028]. As was true for the Q3-Q4, 1994 sample, the estimated standard deviation of the disturbance in the log wage offer equation is significantly greater for private-sector jobs. When we estimate the model with regressors, we find once again that age and tenure effects are important determinants of public-sector wage offers, but play no significant role in determining private-sector wage offers.

While the brevity of the entire sample period does not allow us to draw inferences regarding trends in the Polish labor market, the consistency of results obtained across the various samples does lend some credibility to our interpretations of mobility patterns and the wage determination process in this market.

Our results hint at three possible explanation for low churning in the Polish labor market. First, segmentation in the allocation of job offers makes it very difficult for those outside the private sector to receive job offers in the emerging small business sector. Second, public sector employment is significantly more stable than private sector employment, and although its share in total employment has declined considerably, it still involves a large fraction of the active population.[15] Third, there is no evidence of steep tenure and age profiles in the private sector. Thus, shifting jobs and moving from public to private enterprises may have high costs relative to the returns one can get from it, especially for those with relatively long tenures and work experience in public sector jobs.

5. Conclusion

In spite of major changes occurring in the distribution of employment across the public and private sectors, industries and occupations, worker mobility in transitional economies is low. In this paper we compute some standard measures of mobility for transition countries showing that there is strikingly more interindustry and interoccupational mobility in countries with rigid labor markets, like Italy, than in central and eastern European transforming economies.

In order to explain why mobility is so low, we develop a simple model of labor market transitions across and within public and private firms as well as between employment and nonemployment. The model is sufficiently flexible so as to capture the complicated dynamics of transitional labor markets. It also offers a rather parsimonious representation of the stochastic process governing labor market transitions and one that can be readily interpreted from a behavioral perspective. The model enables us to make inferences concerning the distribution of job offers conditional on occupying different labor market states, the probabilities of moving across jobs and labor market states, and the distributions of the values associated with occupancy of those states.

Our results point to three possible explanations for low mobility in transitional

[15] The privatization process in Poland has been much slower than in other transition economies, such as the Czech Republic. By June 1996, privatization had been completed in 2,624 enterprises out of the 8,853 state enterprises existing in 1990, when the process was started. These numbers should also be interpreted with caution, as the state has in some cases retained substantial participation in privatized units, remaining the *de facto* controlling shareholder. The size of the residual state sector may provide a better measure of the extent of privatization than the number of privatized units per se: the total book value of the state shares in the business sector was estimated in November 1995 to amount to about 140 billion zlotys. This compares with a book value of about 8 billion zlotys for all firms quoted at the Warsaw Stock Exchange and revenues from the sale of state assets of the order of 2.6 billion in 1996.

economies. The first is segmentation in the allocation of job offers: those outside the private sector are less likely to be offered a job from private-sector firms. The second explanation relates to the fact that – in spite of its shrinking size (which is also related to the privatization of continuing firms) – the public sector still offers more stable jobs than private units. Finally, the low return to experience and age and the greater dispersion of private-sector offers makes employment in this sector less attractive, on average, than public-sector employment.

In future work we plan to estimate the model fully exploiting the panel structure of the Polish LFS. Having more observations on the same individual would enable us to relax some of our distributional assumptions on the error terms. Other planned extensions of the model concern the treatment of voluntary quits. Separations in the current model are treated as purely exogenous. Transitions to inactivity, especially at the beginning of economic transformation in Central and Eastern Europe, had a significant voluntary component, given that under the previous regime having a job was considered an obligation. Disentangling quits from layoffs would also allow us to better characterize asymmetries in the riskiness of public sector vs. private sector jobs.

Finally, we also plan to extend the model by estimating the value of non-employment and explicitely incorporating various aspects of the unemployment benefit system. This would allow us to capture the effects on mobility of the radical changes in the generosity of the unemployment benefits observed in Poland (and other transitional economies) during the fiscal crises of 1992-3.

LEGEND

$\pi(n,2)$	probability of not receiving a job offer conditional on being unemployed at t
$\pi(n,3)$	" of receiving an offer from a public firm conditional on being unemployed at t
$\pi(n,4)$	" of receiving an offer from a private firm conditional on being unemployed at t
$\pi(g,1)$	" of being laid-off conditional on being employed in a public firm at t
$\pi(g,2)$	" of not receiving a job offer conditional on being employed in a public firm at t
$\pi(g,3)$	" of receiving an offer from another public firm conditional on being in a public firm at t
$\pi(g,4)$	" of receiving an offer from a private firm conditional on being in a public firm at t
$\pi(p,1)$	" of being laid-off conditional on being employed in a private firm at t
$\pi(p,2)$	" of not receiving a job offer conditional on being employed in a private firm at t
$\pi(p,3)$	" of receiving an offer from a public firm conditional on being in a private firm at t
$\pi(p,4)$	" of receiving an offer from another private firm conditional on being in a private firm at t

Constant (g)	Intercept of wage offer equation for public sector workers
Age (g)	Age coefficient of wage offer function for public sector workers
Constant (p)	Intercept of wage offer equation for private sector workers
Age (p)	Age coefficient of wage offer function for private sector workers
σ_g	standard deviation of the wage shocks in the public sector
σ_p	standard deviation of the wage shocks in the private sector
$\tau_1(g)$	linear tenure term in the public sector
$\tau_2(g)$	quadratic tenure term in the public sector
$\tau_1(p)$	linear tenure term in the private sector
$\tau_2(p)$	quadratic tenure term in the private sector
L	log likelihood values

Table 1. Mobility Indexes[a,b]

	Between Public and Private Firms					
Open[c]	92-93	93-94	94-95	95-96	96-97	*Stationarity*
Hungary				0.08	0.08	*acc.*
Poland	0.13	0.17	0.17	0.15		*rej.*
Slovak R.			0.08	0.08		*acc.*
Homogeneity			*rej.*	*rej.*		
Closed[d]	92-93	93-94	94-95	95-96	96-97	*Stationarity*
Hungary				0.02	0.01	*rej.*
Poland	0.07	0.07	0.07	0.05		*rej.*
Slovak R.			0.04	0.02		*rej.*
Homogeneity			*rej.*	*rej.*		
			Across 12 sectors[e]			
Open[c]	92-93	93-94	94-95	95-96	96-97	*Stationarity*
Czech R.		0.08	0.09	0.07	0.06	*rej.*
Hungary				0.04	0.04	*acc.*
Poland	0.12	0.17	0.19	0.13		*rej.*
Homogeneity		*rej.*	*rej.*	*rej.*	*rej.*	
Italy			0.18	0.16	0.16	*rej.*
Closed[d]	92-93	93-94	94-95	95-96	96-97	Stationarity
Czech R.		0.03	0.02	0.02	0.02	*rej.*
Hungary				0.04	0.03	*rej.*
Poland	0.14	0.18	0.19	0.12		*rej.*
Homogeneity		rej.	rej.	rej.	rej.	
Italy			0.18	0.17	0.16	*rej.*

Table 1. Mobility Indexes[a,b] (continues)

			Across 9 occupations[f]			
Open[c]	92-93	93-94	94-95	95-96	96-97	*Stationarity*
Czech R.		0.08	0.10	0.08	0.07	*rej.*
Hungary				0.05	0.04	*rej.*
Poland			0.17	0.13		*rej.*
Homogeneity			*rej.*	*rej.*	*rej.*	
Italy			0.21	0.19	0.19	*rej.*
Closed[d]	92-93	93-94	94-95	95-96	96-97	*Stationarity*
Czech R.		0.02	0.04	0.04	0.02	*rej.*
Hungary				0.03	0.03	*acc.*
Poland			0.20	0.12		*rej.*
Homogeneity			*rej.*	*rej.*	*rej.*	
Italy			0.23	0.20	0.16	*rej.*

Notes:

(a) I=(s-tr(M))/(s-1))

(b) Yearly average of quarterly measures of mobility.

(c) including shifts to and from unemployment and inactivity

(d) excluding shifts to and from unemployment and inactivity

(e) The following sectoral classification was used: 1) Agriculture & Fishing, 2) Energy & Water, 3) Manufacturing, 4) Construction, 5) Trade & Repair, 6) Hotels & Restaurants, 7) Transport & Communications, 8) Financial Services, 9) Real Estate, 10) Public Administration & Defence, 11) Education & Health, 12) Other public Services

(f) The following classification of occupations was used: 1) Legislators, Senior Officials and Managers, 2) Professionals, 3) Technicians and Associate Professionals, 4) Clerks, 5) Service Workers and Shop and Market Sales Workers, 6) Skilled Agricultural and Fishery Workers, 7) Craft and Related Trades Workers & , 8) Plant and Machine Operators and Assemblers, 9) Elementary Occupations.

Table 2. Descriptive Statistics

		Q3-Q4 94	Q1-Q2 95	Q3-Q4 95
		n = 367	n = 465	n = 2516
	n	0.223	0.265	0.230
	g	0.523	0.480	0.493
	p	0.253	0.256	0.277
	n	0.248	0.254	0.257
S1 (a)	g	0.515	0.486	0.489
	p	0.237	0.260	0.254
	w2g (σ_g)	5.925 (0.376)	6.033 (0.373)	6.316 (0.393)
S2 (b)	w2g (σ_p)	5.897 (0.453)	6.003 (0.385)	6.206 (0.440)
	$\tau(g)$	13.371 (8.227)	12.163 (8.808)	12.483 (8.632)
	$\tau(p)$	8.647 (8.581)	7.162 (8.544)	6.479 (7.771)
	age n	39.286	41.525	41.028
	age g	39.936	39.548	39.763
	age p	38.667	38.851	39.606

Notes:

(a) Variables referred to the first of the two quarters in the matched sample.

(b) Variables referred to the second of the two quarters in the matched sample.

Table 3. Transition Matrices

	Q3-Q4 94	Q1-Q2 95	Q3-Q4 95
=====	Across	Status	=======
p_{nn}	0.927	0.870	0.931
p_{ng}	0.012	0.057	0.035
p_{np}	0.061	0.073	0.035
p_{gn}	0.021	0.022	0.031
p_{gg}	0.958	0.973	0.964
p_{gp}	0.021	0.005	0.006
p_{pn}	0.118	0.050	0.100
p_{pg}	0.043	0.017	0.022
p_{pp}	0.839	0.933	0.878
=====	Across	Employers	=======
$p_{g\nrightarrow g}$	0.859	0.922	0.947
$p_{g\rightarrow g}$	0.141	0.078	0.053
$p_{p\nrightarrow p}$	0.821	0.901	0.931
$p_{p\rightarrow p}$	0.179	0.099	0.069

Table 4. Model Estimates: Quarters 3 and 4, 1994

Parameters	Specification 1	Specification 2
$\pi(n,2)$	.927	.927
	(.029)	(.029)
$\pi(n,3)$	.012	.012
	(.012)	(.012)
$\pi(n,4)$	.061	.061
	(.026)	(.026)
$\pi(g,1)$	.021	.021
	(.005)	(.005)
$\pi(g,2)$	.637	.518
	(.057)	(.089)
$\pi(g,3)$	.289	.405
	(.053)	(.086)
$\pi(g,4)$	.053	.057
	(.027)	(.029)
$\pi(p,1)$	.118	.118
	(.033)	(.034)
$\pi(p,2)$	.413	.346
	(.086)	(.107)
$\pi(p,3)$	.091	.131
	(.044)	(.064)
$\pi(p,4)$	.377	.405
	(.083)	(.101)

Constant (g)	5.841	5.043
	(.029)	(.207)
Age (g)		.153
		(.052)
Constant (p)	5.717	5.371
	(.053)	(.369)
Age (p)		.076
		(.093)
σ_g	.392	.381
	(.026)	(.026)
σ_p	.502	.496
	(.030)	(.030)
$\tau_1(g)$		.039
		(.010)
$\tau_2(g)$		-.133
		(.037)
$\tau_1(p)$		.007
		(.018)
$\tau_2(p)$		.007
		(.067)
L	-371.027	-356.993

Table 5. Model Estimates: Quarters 1 and 2, 1995

Parameters	Specification 1	Specification 2
$\pi(n,2)$	.870	.870
	(.030)	(.030)
$\pi(n,3)$	.057	.057
	(.021)	(.021)
$\pi(n,4)$	.073	.073
	(.023)	(.023)
$\pi(g,1)$	.022	.022
	(.002)	(.002)
$\pi(g,2)$	.815	.779
	(.038)	(.049)
$\pi(g,3)$	.153	.188
	(.036)	(.047)
$\pi(g,4)$	.010	.010
	(.010)	(.010)
$\pi(p,1)$	.050	.050
	(.020)	(.020)
$\pi(p,2)$	.735	.714
	(.058)	(.064)
$\pi(p,3)$	.031	.038
	(.022)	(.027)
$\pi(p,4)$	.183	.197
	(.053)	(.058)

Constant (g)	6.000 (.025)	5.432 (.184)
Age (g)		.112 (.046)
Constant (p)	5.955 (.035)	5.725 (.260)
Age (p)		.046 (.067)
σ_g	.378 (.018)	.368 (.018)
σ_p	.388 (.024)	.386 (.024)
$\tau_1(g)$		.019 (.009)
$\tau_2(g)$		-.045 (.032)
$\tau_1(p)$		.018 (.013)
$\tau_2(p)$		-.062 (.053)
L	-369.808	-359.146

Table 6. Model Estimates: Quarters 3 and 4, 1995

Parameters	Specification 1	Specification 2
$\pi(n,2)$	.928	.928
	(.015)	(.015)
$\pi(n,3)$	.039	.039
	(.011)	(.011)
$\pi(n,4)$	.032	.032
	(.010)	(.010)
$\pi(g,1)$	.028	.028
	(.001)	(.001)
$\pi(g,2)$	.849	.807
	(.016)	(.023)
$\pi(g,3)$	.107	.148
	(.014)	(.021)
$\pi(g,4)$	.016	.016
	(.006)	(.007)
$\pi(p,1)$	.098	.098
	(.013)	(.013)
$\pi(p,2)$	.728	.709
	(.026)	(..028)
$\pi(p,3)$	.040	.051
	(.010)	(.014)
$\pi(p,4)$	.134	.142
	(.021)	(.023)

Constant (g)	6.339 (.012)	6.109 (.113)
Age (g)		.010 (.027)
Constant (p)	6.190 (.020)	6.012 (.157)
Age (p)		.034 (.039)
σ_g	.379 (.010)	.378 (.010)
σ_p	.450 (.012)	.451 (.013)
$\tau_1(g)$		.031 (.005)
$\tau_2(g)$		-.086 (.017)
$\tau_1(p)$		.015 (.008)
$\tau_2(p)$		-.054 (.029)
L	-1544.322	-1514.742

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