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Returns to Mobility in the Transition to a Market Economy

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Part I
RETURNS TO MOBILITY IN THE TRANSITION TO A MARKET ECONOMY

(first draft, please do not circulate without the consent of the authors)

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Abstract

In spite of the ongoing dramatic changes in the structure of employment, transitional economies display rather low mobility across sectors, occupations and firms of different ownership. We characterise this low mobility of central and eastern European labour markets by computing mobility measures for transition matrices, the latter estimated on the basis of matched records across LFS waves. As low mobility can be explained by high costs of shifting jobs compared to the vbenefits one can get from the job change as well as by segmentation in the allocation of job offers, we develop an econometric model enabling to characterise intertemporal changes in probabilities of dismissal, remunerations and offer arrival rates on the basis of information only on observed transitions. This model is seminally implemented on matched data across various waves of the Polish LFS. Although our results are highly preòliminary, they point to significant 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. Were these findings supported in further work, they could support explanations for low mobility in transitional ecnomies, 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 on the fact that changing jobs and moving from public to private enterprises is costly especially for those with relatively long tenures and work records.

Non-Technical summary

In spite of ongoing dramatic changes in labor market structure, in transitional economies display rather low mobility of workers across sectors and occupations. In this paper we present evidence, drawn from the Polish Labour Force Survey, suggesting that such a low mobility can be attributed to the costs of moving to a new job in terms of forgone returns to tenure in the previous job as well as by market segmentation in the allocation of job offers. Job offers do not seem to reach those who are most prone to take up jobs, and 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.

1. Introduction

Transition, almost by definition, involves reallocation of workers across jobs, occupations and industries. Given the artificial full employment conditions inherited from the previous regime, transition in central and eastern Europe also involved the appearance of open unemployment and significant flows to inactivity.

Compared to other countries undergoing fast structural change – e.g., the Latin American economies – central and eastern Europe offers a wealth of data on labour market flows. Not only are administrative data, e.g., individual records from the registration of jobseekers at labour offices, often made available to researchers, but also most countries have introduced and are currently undertaking at regular intervals household surveys involving rotating panels, and hence allowing for longitudinal analyses.

Such data have been so far used mainly to describe the magnitude and characteristics of labour market flows or to carry out non-parametric analyses, mainly of hazard rates from unemployment. This has been of great help in characterising the specific features of labour market adjustment during economic transformation (e.g., the stsgnancy of transitional unemployment) and in assessing those personal characteristics that are most relevant in determining different labour market outcomes of those having been displaced in the course of the transition process. Preliminary evaluations of the effectiveness of active labour market policies have also been carried out, estimating the impact of participation in active programmes [Jones and Kato, 1993; Micklewright and Nagy, 1994; Steiner and Kwiatowski, 1995; Puhani and Steiner, 1996; Kotzeva et al., 1996], as well as of unemployment benefit levels and duration on outflows to jobs [Svejnar and Terrell, 1994; Vodopivec, 1995; Micklewright and Nagy,1996; van Ours and Lubyova, 1997; Boeri and Steiner, 1997].

These studies have contributed to our understanding of labour market adjustment under fast structural change and have provided relevant material for evaluations of the impact of labour market programmes, which does not need being confined to the East[1]. Yet, there are as number of issues that still need to be investigated and, more importantly, which *can* be addressed with the help of available data.

[1]The book recently published by OECD – *Lessons from Labour Market Policies in in the Transition Countries* (1996) – 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.

First and foremost, very little is known as to the allocation of job offers across individuals occupying different labour market states. How segmented are labour markets of transitional economies in conveying information on vacancies? Is the probability of being offered a post in the emerging private sector dependent on being employed in a state vs. a private firm and being unemployed or outside the labour force?

Second, what is the impact on returns to education and age and on wage-seniority profiles of the shift between public and private firms? How much one can loose in terms of tenure rating and can gain from better rewards to education and experience while moving from one sector to another?

Third, to which extent does the risk of dismissal vary between public and private firms? While differences in separation rates of firms of different ownership type 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 may originate from dismissals.

All these issues are very important in the light of the findings of the empirical literature on labour 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 [Boeri, 1994].

An aggregate lack of vacancies, however, is unlikely to persist under the current sustained economic recovery and there are indeed indication of a rise in vacancy rates 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 labour markets in these countries display relatively low *churning* rates [Blanchard, Commander and Corricelli, 1995]. Asides from legacies of the previous regime, which rewarded attachment to the firms with a battery of social benefits (commodity subsidies, subsidised housing, recreational services, etc.), returns to mobility under the economic transformation may have been 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.

In a nutshell, what is still lacking is an assessment of the individual costs and benefits of job mobility. Such an assessment would greatly improve our understanding of the specific features of labour markets undergoing major structural change. One of the reason 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 for the most complementing the data provided by the household surveys (which do not convey information on the alternatives rejected by the individuals and rarely provide information on wages) with assumptions concerning the stochastic processes governing labour market transitions.

The purpose of this paper is to contribute to fill this gap. An econometric model is developed which enable us to characterise intertemporal changes in probabilities of dismissal, remunerations and offer arrival rates on the basis of information only on observed transitions. This model is seminally implemented on matched data across various waves of the Polish LFS, which is not only the longest (was started as early as in 1992), but also the most complete survey (insofar as it contains wage and retrospective information).

The results we obtain are plausible and consistent across different LFS waves. As we have so far been estimating the model just over six consecutive quarters of the Polish LFS (and not always over the full sample!), we certainly cannot claim generality and robustness for our results. 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, rewards to tenure and age in the private sector. Were these findings supported in further work, they could support explanations for low mobility in transitional ecnomies, 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 on the fact that changing jobs and moving from public to private enterprises is costly especially for those with relatively long tenures and work records.

The plan is as follows. Section one documents low mobility in transitional economies. Some standard measures of mobility for transition matrices are produced allowing to summarise evidence on the extent of mobility across ownership types, sectors, and occupations in transitional economies and in what is commonly considered as the most rigid OECD labour 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 two develops an econometric model of labour market dynamics tailored on the data available from the Polish LFS. This household survey was started as early as in 1992 and provides information on wages unlike household surveys in many other transitional economies. Section 3 presents our preliminary results and provides some suggestions for further research. Section 4 concludes by raising a number of issues concerning the use of our empirical framework in assessing the impact of policy changes (e.g., the tigthening of unemployment benefit systems) on the way in which labour markets operate.

2. Low Mobility in Transitional Economies

This section produces and discusses some summary measures of mobility across labour market states, sectors, ownership types and occupations. In addition to complementing the results from the econometric model developed in section 2 (which draws on smaller samples and so far concern only Poland), these measures contribute to addressing three important empirical issues raised by the literature on transitional economies.

First, empirical work on labour market transitions in central and eastern Europe has documented a relatively low turnover of the unemployment pools in these countries [Boeri, 1994]. This stands in sharp contrast with deep falls in employment experienced 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 either occurring directly (particularly at early stages of transition) or indirectly, that is, involving often longlasting spells of unemployment and discouraged worker effects[2]. This could explain why large employment declines coexist in these countries with low unemployment inflows and outflows. Yet, one has still to explain how radical changes in the structure of employment by sector, occupation and ownership type are being achieved. Whether it is mainly those who are already in employment, as opposed to those who are out of it who move across jobs is still an open empirical issue. We hope to shed some light on this issue on the basis of measures of mobility involving or excluding non-employment.

Second, the claim has often been made (mainly in an attempt to explain the low vacancy rates generated by these countries) that transitional labour markets tend to display much less churning than their western counterparts [Blanchard, Commander and Corricelli, 1995]. If confirmed by empirical evidence, this claim could also explain the coexistence in these countries of rapid structural change and low gross worker flows. Low churning would also imply that deep structural change is being achieved by mobilising a rather small segment of the working

[2] See the various articles of the (forthcoming) special issue of The Journal of Empirical Economics on "Long-term unemployment and short-term benefits".

age population. Measures of mobility for transition matrices (across sectors, occupations and ownership types) offer a better basis than labour turnover or job turnover data – which are based on administrative records, and hence are affected by their incomplete coverage of the small business sector – to assess the magnitude of job churning and compare it across countries.

Third, economic transformation, notably the growth of the private sector, is supposed to significantly modify the way in which labour markets operate. Ongoing reforms of labour market institutions, notably the tightening of the fairly generous unemployment benefit systems introduced at the outset of transition and a partial relaxation of employment security schemes, are also likely to involve an increase over time of labour market mobility in these countries. This can be assessed by computing rather simple measures of mobility and looking at their behaviour over time. By computing mobility indexes over as many years as possible, we also hope to remedy to a limitation of the econometric part to follow, 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 at the outset of transition. Comparing mobility matrices across years can be informative as to the stability of observed patterns of mobility (it is just based on these transitions that the model is estimated) across labour market states, private and public firms, etc..

2.1. Data Issues

Indexes of mobility were computed from transition matrices in 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 retrospective questioning, which is also allowed in many questionnaires.

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

The linking of records was eased in Poland and Hungary by the assignment of unique identifiers to each sampled person. In other countries identifiers were only provided for the household or the dwelling 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 residence, and other personal or household characteristics, depending on the nature of the survey.

The main problem with matched records is that sample attrition, non-response and errors in the classification of the labour market statutes 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 labour market status of individuals tend to be 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 overestimating changes . Another issue related to the use of matched records is that unlike many retrospective questions – they cannot capture flows occurred within survey dates. In particular, matched records do not capture "roundtripping" across labour market states, which is sometimes significant in these countries [Lehmann and Gora, 1994]. 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 a common summary measure of mobility for transition matrices. In particular, the scalar measure is given by the index:

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

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 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, attach value zero only to identity matrices, and one to matrices with identical rows (hence probabilities of moving independent of the state originally occupied). All the computed matrices had maximum diagonal, hence in our case the index satisfies the four properties listed above[3].

States were defined on the basis of internationally agreed (ILO-OECD) definitions of employment, and comparable industry (ISIC) and occupational (ISCO88) classifications.

Three 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[4]. Strikingly enough, all transitional economies do not exhibit larger mobility than Italy, a country typically pointed out as having all the ingredients of a rigid labour market with many "jobs for the life" and much less mobile than the United States [Flinn, 1997].

Second, including or excluding those outside employment – that is, using transition matrices with or without unemployment and the inactive status – does not greatly affect the measures of mobility. If most of the shifts were occurring across those who are already employed, we would have expected mobility conditional to being employed to be larger than mobility unconditional to the initial labour market status of 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 the non-employed individuals are included.

Third, there is not a clear trend in mobility measures. Formal tests of sta-

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

[4] Under the null hypothesis, $\pi_{ij(k)=}\pi_{ij(\forall k'\neq k)}$. Under the alternate hypothesis, the estimates of the transition matrix for country k are given by:

$\widehat{\pi}_{ij}\ (k) = \frac{m_{ij(k)}}{\sum m_{ij(k)}}$

The likelihood function maximised under the null hypothesis is:

$\prod_k \prod_{ij} m_{ij}(k)$

whereas under the alternative is:

$\prod_{ij} m_{ij}$

The ratio is the likelihood ratio criterion:

$\lambda = \prod_k \prod_{ij} \frac{m_{ij}}{m_{ij(k)}}$

Extensions of the Cramer and Neyman theorem (cf. Anderson and Goodman, 1956) show that -2$\lambda\lg(\lambda)$ is distributed as a χ^2 with (k-1)[s(s-1)] degrees of freedom when the null hypothesis is true. We plan to undertake similar tests for individual mover coefficients.

tionarity of transition matrices[5] led us to reject in most cases the hypothesis of stationarity of transition matrices. But the direction of the change is ambiguous. If anything, mobility would seem to be slightly declining in recent years. While this suggests that the econometric results we present in Section 3 need to be confirmed over data covering other time periods, we do not have the impression from the analysis of mobility matrices that we have so far been taking observations which are out of range with respect to the low mobility characterising the Polish labour market.

All this confirms the view of transitional labour markets as 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 the move? Are offer arrival rates disproportionately allocated to those who are less prone to move?

The next two sections will try and provide some provisional answers to such questions.

[5] For any country k, under the null hypothesis, $\pi_{ij(t)=}\pi_{ij(t=1...T)}$..Under the alternate hypothesis, the estimates of the transition matrix for time t are given by:

$\widehat{\pi}_{ij(t)}=\frac{m_{ij}}{\sum m_{ij(t-1)}}$

The likelihood function maximised under the null hypothesis is:

$\prod_t \prod_{ij} m_{ij}(t)$

whereas under the alternative is:

$\prod_{ij} m_{ij}$

The ratio is the likelihood ratio criterion:

$\lambda = \prod_t \prod_{ij} \frac{m_{ij}}{m_{ij(t)}}$

Extensions of the Cramer and Neyman theorem (cf. Anderson and Goodman, 1956) show that $-2\lambda \lg(\lambda)$ is distributed as a χ^2 with (T-1)[s(s-1)] degrees of freedom when the null hypothesis is true.

Table 1

MEASURES OF MOBILITY (a) (b)

BETWEEN PUBLIC AND PRIVATE FIRMS

open*

	92-93	93-94	94-95	95-96	96-97	stationarity
POLAND	0.13	0.17	0.17	0.15		rej.
HUNGARY				0.08	0.08	acc.
SLOVAK R.			0.08	0.08		acc.
homogeneity			rej.	rej.		

* including shifts to and from unemployment and inactivity

closed*

	92-93	93-94	94-95	95-96	96-97	stationarity
POLAND	0.07	0.07	0.07	0.05		rej.
HUNGARY				0.02	0.01	rej.
SLOVAK R.			0.04	0.02		rej.
homogeneity	-	-	rej.	rej.	-	

* excluding shifts to and from unemployment and inactivity

ACROSS 12 SECTORS

open*

	92-93	93-94	94-95	95-96	96-97	stationarity
POLAND	0.12	0.17	0.19	0.13		rej.
HUNGARY				0.04	0.04	acc.
homogeneity				rej.		
ITALY		-	0.18	0.16	0.16	rej.

* including shifts to and from unemployment and inactivity

closed*

	92-93	93-94	94-95	95-96	96-97	stationarity
POLAND	0.14	0.18	0.19	0.12		rej.
HUNGARY				0.04	0.03	rej.
homogeneity				rej.		
ITALY		-	0.18	0.17	0.16	rej.

* excluding shifts to and from unemployment and inactivity

ACROSS 9 OCCUPATIONS

open*

	94-95	95-96	96-97	stationarity
POLAND	0.17	0.13		rej.
HUNGARY		0.05	0.04	rej.
homegeneity		rej.		
ITALY	0.21	0.19	0.19	rej.

* including shifts to and from unemployment and inactivity

		closed*		
	94-95	95-96	96-97	stationarity
POLAND	0.20	0.12		rej.
HUNGARY		0.03	0.03	rej.
homegeneity		rej.		
ITALY	0.23	0.20	0.16	rej.

* excluding shifts to and from unemployment and inactivity

Notes:

(a) M(P)=(n-traceP)/(n-1)

(b) Yearly average of quarterly transition matrices

(c) See text for details on the mobility indexes being used.

The following sectoral classification was used:

1) Agriculture & Fishing
2) Energy & Water
3) Manufacturing
4) Construction
5) Trade & Repair
6) Hotels & Reastaurants
7) Transport & Communications
8) Financial Services
9) Real Estate
10) Public administration & Defence
11) Education & Health
12) Other public Services

The following classification of occupations was used:

1) Legislators, Head Directors & Entrepreneurs
2) Highly Specialised Professionals
3) Intermediate Professionals, Technicians
4) Operational Staff (white collars)
5) Marketing
6) Artisans & Specialised Workers and Agricultural Workers

7) Supervisors of Machinery
8) Unskilled Workers
9) Officials of Armed Forces

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 Labour Force Surveys of central and eastern European countries. As we plan to initially implement the model on Polish data, some features of the model fit particularly the design of the Polish LFS, but can be readily amended and adapted to the other countries.

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 characterising returns to tenure and wage differentials between the public and private sector.

Our objective is to characterize intertemporal changes not only in employment opportunities, but also in remuneration across two sectors of the economy, broadly classified as the state (g) and the private (p) sectors. Based on this characterisation 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 labour 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 labour 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 itself we posit.[6]

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.

[6] 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 whilst the values of the alternatives not chosen are not observed.

X is a $(1 \times K)$ vector of time-invariant characteristics of the labour market participant, which includes sex, education level (since all sample members will have completed schooling), etc.

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:

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

where

$\alpha(\xi, t)$ is the constant term in the regression function, which is indexed by sector and time period

$\beta(\xi, t)$ 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 $\xi$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 labour 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 labour market state in period t, with $l(t) = n, g$, or p. At the beginning of the next period, $t + 1$, the 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 $t+1$.[7] $\pi(g,3,t)$ 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 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 state-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 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

[7] 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.

nature which may appear in the future cannot be anticipated. Thus, in attempting to assess the future evolution of the labour market, Polish workers especially in the early 1990s may have found themselves in such a situation of "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 labour market movements of sample members and their compensation rates for an 18 month period[8]. Thus estimating a life-cycle model of labour market dynamics would require us to make very strong assumptions concerning the behaviour 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 modelling labour market decisions in this particular case are the low asset positions of most of the labour market participants and the very high inflation rates countersigning particularly the initial 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

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

[8]Longitudinal data from LFS, which are available for the other central and eastern countries cover even shorter a time period.

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

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

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 labour 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 Labour Market Transitions

Before turning to estimation issues, it will be useful to consider what the model implies in terms of observed patterns of labour 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 labour market states.*[9]

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

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 we have access to consists of information on the labour market states and demographic characteristics of individuals at 2 points in time.[10] The sampling scheme for the Polish labour Market Survey has selected households interviewed in two consecutive quarters, then skipped them for two consecutive quarters, and then back in 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 who is participating in the labour market. For simplicity, we will confine our estimates covering two *consecutive* quarters only. Without lack of generality, we shall refer to these quarters as t and $t+1$.

The key to estimating the model is recognizing that it is strictly Markovian - that is, the probability of transiting from state m [his or her period t labour market state] to state m' [his or her period $t+1$ labour 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 wage density is given by

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

[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 being made at present.

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

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

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

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

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

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

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 $\pi(n,3), \pi(n,4), \pi(g,2), \pi(g,3), \pi(g,4), \pi(p,2), \pi(p,3), \pi(p,$ 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 some preliminary estimates of the econometric model developed in the previous section using subsamples of observations from the Polish labour Force Surveys over an 18 month period extending from the third quarter of 1994 to 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 and the rotating scheme (initially it was a full panel) has been changed repeatedly making it more difficult to reconstruct individuals' histories and making observations on the occupation and sector not strictly comparable over time.

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 serious consideration of the labour 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 indicating 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 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. We have not as yet performed a systematic analysis of missing wage patterns in these data, but intend to do so in the near future.

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.[13] The final sample size for Q3-Q4, 1995 is 1830, while the sample sizes for Q3-Q4, 1994 and Q1-Q2, 1995 are 367 and 465, respectively. The relatively small sizes of these two

[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 these 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]We restricted the size of these samples for purposes of computational expediency only. In the revision of this paper we will report results based on the full samples available for all periods.

samples can be expected to result in parameter instability and relatively large estimated [asymptotic] standard errors. This is particularly the case with respect to parameters associated with labour market transitions which occur relatively infrequently in the population.

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 2 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 implying that the estimated values of $\pi(n, 3)$ and $\pi(n, 4)$ come directly from the observed transition matrix. It is interesting to compare estimates of the probability of dismissal from the two sectors. It is also interesting 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 at 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 labour market may be segmented in the sense that search behaviour and network formation [for employment contacts] tends 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]. 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 which 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 tenures in private firms are likely to be seriously affected by length bias [14] – given that prior to 1989 private sector non-agricultural employment accounted for barely 5 per cent of employment - while 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

[14] We will try and address this censoring problem in future work.

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 experirence (perhaps because the human capital inherited from the revious regime and organisation of production is deemed obsolete) 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 4. 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 labour 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 preliminary results hint at three possible explanation for low churning in the Polish labour 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 profile 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 records.

4.1.

[15]The privatisation process in Poland has been much slower than in countries like the Czech Republic. By June 1996, privatisation 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 privatised units, remaining de facto, the controlling shareholder. The size of the residual state sector may provide a better measure of the extent of privatisation than the number of privatised 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 on the Warsaw Stock Exchange and revenues from the sale of state assets of the order of 2.6 billion in 1996.

5. Final Remarks

The model proposed in this paper is sufficiently flexible so as to capture the complicated dynamics of transitional labour markets. It also offers a rather parsimonious representation of the stochastic process governing labour market transitions and one that can be readily interpreted from an economic perspective.

A main limitation of our model is that it cannot capture directly the effects of policy changes, as the policy environment is not explicitly modelled in choice rules. However, we can still make inferences on the effects of policy changes. This can be done by checking whether changes in the value of parameters of interest are consistent with changes in the policy variables. For instance, the decision rule of nonemployed individual does not explicitly model the value of the nonemployment benefits. However, the cofficients $\alpha * (., t)$ and $\beta * (., t)$ capture, respectively, effects of the unemployment benefits which are orthogonal to our vector of individual characteristics and effects which are associated with age, duration of the unemployment spell in progress, etc.. We would expect the tightening of unemployment benefits should result in reductions of the parameters $\alpha * (., t)$ and $\beta * (., t)$.

Comparisons of the estimated parameters across countries can alkso shed some light on the impact pof policies on labour market transitions. In particular, we expect to gain valuable insights by comparing results from countries like Poland and Hungary with estimated valkues of the parameters in the Czech Republic, the countries with the tigthest enforcement of work-tests, with the shiortest unemployment benefits in the arena of transitional economies and where much efforts have been made to implement on a wide scale active labour market policies.

Table 2. 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
	(.003)	(.005)
$\pi(g, 2)$	.637	.518
	(.057)	(.092)
$\pi(g, 3)$	.289	.404
	(.053)	(.089)
$\pi(g, 4)$	.053	.057
	(.027)	(.029)
$\pi(p, 1)$	.118	.118
	(.033)	(.033)
$\pi(p, 2)$	.413	.346
	(.116)	(.119)
$\pi(p, 3)$	.091	.131
	(.044)	(.064)
$\pi(p, 4)$	.377	.405
	(.110)	(.111)

Constant (g)	8.143	7.345
	(.029)	(.207)
Age (g)		.153
		(.052)
Constant (p)	8.020	7.674
	(.055)	(.362)
Age (p)		.076
		(.092)
σ_g	.392	.381
	(.026)	(.026)
σ_p	.502	.496
	(.030)	(.030)
$\tau_1(g)$		.040
		(.010)
$\tau_2(g)$		-.133
		(.037)
$\tau_1(p)$		.007
		(.018)
$\tau_2(p)$		.007
		(.066)
L	-371.026	-356.993

Table 3. Model Estimates: Quarters 1 and 2, 1995

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

Constant (g)	6.339	6.109
	(.012)	(.113)
Age (g)		.010
		(.027)
Constant (p)	6.190	6.012
	(.020)	(.157)
Age (p)		.034
		(.039)
σ_g	.379	.378
	(.010)	(.010)
σ_p	.450	.451
	(.012)	(.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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