JOB AND WAGE CHANGES DURING THE TRANSITION:

Evidence from Czech Retrospective Data

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TITLE VIII PROGRAM

Project Information*

Sponsoring Institution:	**University of Michigan**
Principal Investigator:	**Katherine Terrell**
Council Contract Number:	**814-13g**
Date:	**October 4, 2000**

* The work leading to this report was supported in part by contract or grant funds provided by the National Council for Eurasian and East European Research, funds which were made available by the U.S. Department of State under Title VIII (The Soviet-East European Research and Training Act of 1983, as amended). The analysis and interpretations contained herein are those of the author.

Executive summary

The transition from central planning to a market economy provides one of the most interesting laboratories to study the simultaneous emergence of labor markets and the private sector. In this paper, we analyze a key issue in transition economies that has so far received little attention, namely the link between job mobility, individual wage changes and the dynamics of firm characteristics over time. Specifically, we examine the effect of job mobility on individuals' wage changes from 1989 to 1996, considering how workers left the old job and the characteristics of the new job, including whether it is a new private sector firm or not. The lack of attention to this issue appears surprising since the movement of labor from old and declining sectors of the economy to the new ones, coinciding with the exit, entry and restructuring of firms, is the very essence of the transition process. The scant literature stems largely from the lack of data for most of the transition economies that link individuals' wage and job characteristics to their firms' characteristics. In this paper we report results based on a retrospective survey for the 1989-1996 period that we carried out in the Czech Republic.

Section 1: Introduction

The transition from central planning to a market economy provides one of the most interesting laboratories to study the simultaneous emergence of labor markets and the private sector. In this paper, we analyze a key issue in transition economies that has so far received little attention, namely the link between job mobility, individual wage changes and the dynamics of firm characteristics over time. Specifically, we examine the effect of job mobility on individuals' wage changes from 1989 to 1996, considering how workers left the old job and the characteristics of the new job, including whether it is a new private sector firm or not. The lack of attention to this issue appears surprising since the movement of labor from old and declining sectors of the economy to the new ones, coinciding with the exit, entry and restructuring of firms, is the very essence of the transition process.[1] The scant literature stems largely from the lack of data for most of the transition economies that link individuals' wage and job characteristics to their firms' characteristics.[2] In this paper we report results based on a retrospective survey for the 1989-1996 period that we carried out in the Czech Republic.

Important related studies in this area are by Boeri and Flinn (1997), Burda and Metens (1998), Hunt (1998), and Noorkôiv, et al. (1998). Hunt's paper is the most similar to ours in that she uses panel data to investigate the determinants of the tremendous wage growth in former East Germany over the 1990-1996 period. Hunt identifies the demographic characteristics of the biggest gainers and estimates the returns to job mobility. Boeri and Flinn (1997) use Polish data on the same individuals over six consecutive quarters to estimate a structural econometric model characterizing inter-temporal changes in the probabilities of dismissal, remuneration and offer arrival rates. Their estimates of costs and benefits

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We would like to thank Jan Kmenta and Stepan Jurajda for their useful comments. We take responsibility for any remaining errors in the paper.

[1]This is not to say that the literature on labor markets in transition economies is thin, as a number of important areas have been carefully researched. For instance, the literature has examined the evolution of labor demand by firms, wage setting at the firm or industry level, determinants of wages in the context of the human capital model, and the flows of individuals among the three labor market states of employment, unemployment and out-of-the-labor force. (See Svejnar, 1999, for a recent survey of these studies.)

[2] For example, whereas the Russian Longitudinal Monitoring Survey (RLMS) has very good panel data on individuals, it has relatively few characteristics of the firm in which the individual works.

of job mobility are derived at a point in time when Poland was in a fairly mature stage in its transition. We are able to capture wage changes at the beginning of the transition process. Burda and Metens (1998) focus only on the wage changes of displaced workers in Germany, while we examine wages changes of both voluntary quits and involuntary layoffs. Finally, Noorkôiv et al. (1998) use a retrospective data set from Estonia that is very similar to our Czech data. However, in this paper they do not make full use of the panel nature of the data. They examine changes in employment by industry and estimate cross-sectional wage regression (similar to our first paper using the Czech retrospective data, Munich, et al. 1999). However, none of these studies examines the interactions between job and wage changes and the new emerging private sector.

The relationship between job mobility and wages has of course been one of the main areas of interest in western labor economics (see e.g., Light and McGarry, 1998, and Farber, 1999, for recent surveys). A number of models have been developed, starting with the Blumen, Kogan and McCarthy (1995) "stayer-mover" model which predicts that mobility is negatively related to wages because "good" (high-productivity) workers avoid turnover, while "bad" (low-productivity) workers move. If the transition is a process during which employers start behaving in a profit maximizing way and lay off low productivity workers, then one could expect to find support for the mover-stayer model, at least in the case of layoffs.

An opposite prediction is obtained from search models (e.g., Burdett, 1978, and Jovanovich, 1979), where mobility represents a means of accomplishing better (higher productivity) matches between workers and employers and thus leads to higher wages. If the transition is a process in which the creation and restructuring of firms creates major matching opportunities, then one should observe job mobility to be positively correlated with wage gains as workers and employers realize new productive matches.[3]

[3] There is also an extensive literature that examines the relationship between firm specific human capital, tenure, and wage changes that we will not address in this paper. See for e.g., Flinn, (1986), Neumark and Taubman (1995), Topel (1991).

In this paper we test these competing predictions about the relationship between job and wage mobility, taking into account the decision of workers to quit or stay and to move to firms in the new versus old sector of the economy. The outline of the paper is as follows. We describe our data set in Section 2 and our methodology in Section 3. We then discuss the results from the various specifications of the wage change equations in Section 4 and conclude the paper in Section 5.

Section 2: Data

The data we use in this paper are taken from a survey we conducted of 3,157 randomly selected households living throughout the 76 districts of the Czech Republic in December 1996.[4] Any member of the sampled household who had worked for a minimum of two weeks between January 1, 1991 and December 31, 1996 was asked a series of retrospective questions about the characteristics of each employment and non-employment spell during this period. We have monthly data on the labor market histories of 4,700 individuals (2,284 men and 2,416 women). Moreover, we have information on the characteristics of the jobs held in January 1989 and December 1996.[5]

In this paper we analyze how those individuals who had worked under communism fared in the transition. Hence, given the structure of our data, we select all individuals who held a job in both 1989 and 1996, as this enabled us to create a panel and compare wages at the same point in time for everyone.[6] This selection yields a sample of 3,072 individuals. After further cleaning the data to exclude individuals who held only part-time jobs, had missing wage data in 1989 or 1996, held more than two full-time jobs

[4] A comparison of the means and distributions of the major characteristics of the sample in our survey with those from the Labor Force Survey reveals they are very similar. Hence we believe our sample is representative of the population in 1996. See Munich et al. (1997) for a detailed description of the sample design and characteristics of the data.

[5] January 1989 was selected as a point in the last year of communism with the belief that people were likely to remember their labor market characteristics at the beginning of this year. We then used January 1, 1991 as the starting point for the detailed labor market histories as this is the year when the government effectively began to engineer the transition to a market economy.

[6] We therefore exclude all individuals who entered the labor market in 1991-1996 as this requires a different methodology. We plan to analyze their experience in future research.

simultaneously, had missing data for the question on how they left their job, we are left with a sample of 2,343 individuals. The characteristics of this sample are described in Section 4 of the paper.

Section 3: Methodology

The goal of the paper is to analyze the 1989-1996 wage changes of individuals with labor market experience in the communist regime, assess whether or not these workers benefited from job mobility, and evaluate whether the gain/loss from mobility varied with the type of firm to which they moved.

3.1 Simple Wage Regression Model

We begin by modeling the logarithm of wages in any given period t (W_t) as a function of time-invariant characteristics X and time-varying characteristics Y_t, respectively. Suppressing individual subscripts, we write the relevant wage equation for 1996 as:

$$\ln W_{96} = X'\alpha_{96} + Y'_{96}\beta_{96} + \varepsilon_{96}, \qquad (1)$$

where *ln* is the natural logarithm, α_{96} and β_{96} are the coefficient giving the 1996 payoffs to the values of the explanatory variables and ε_{96} is the 1996 individual specific error term. The corresponding wage equation for 1989 may be written as:

$$\ln W_{89} = X' (\alpha_{96} - \Delta\alpha) + Y'_{89} (\beta_{96} - \Delta\beta) + (\varepsilon_{96} - \Delta\varepsilon), \qquad (2)$$

where $\Delta\alpha = \alpha_{96} - \alpha_{89}$ and $\Delta\beta = \beta_{96} - \beta_{89}$ are the differences in payoffs to the explanatory variables and $\Delta\varepsilon = (\varepsilon_{96} - \varepsilon_{89})$ are the differences in the error terms between the transition (1996) and communism (1989). Subtracting equation (2) from (1) yields the equation for the percentage change in wages between 1989 and 1996:

$$\ln W_{96} - \ln W_{89} = X'\Delta\alpha + Y'_{89}\Delta\beta + (Y'_{96} - Y'_{89})\beta_{96} + \Delta\varepsilon. \qquad (3)$$

For the time-invariant explanatory variable X we hence obtain estimates of the changes in payoffs ($\Delta\alpha$) between 1989 and 1996, while for the time-varying characteristics Y we generate estimates of the coefficients for 1996 (β_{96}) and of the changes in payoffs ($\Delta\beta$) between 1989 and 1996.

Vectors X and Y_t contain variables relating to the individual's human capital (HC), local labor market conditions (D) and job characteristics (J):

$$X = (HC, D, J)$$

$$Y_t = (HC_t, D_t, J_t) \qquad (4)$$

Specifically, we model the wage change equation as a function of the following human capital characteristics: gender, education (time invariant), potential experience in 1989, the change in experience between 1989 and 1996 (netting out unemployment spells). The coefficients on gender, education and experience in 1989 hence capture the (relative) change in the return (i.e., 1989-1996 wage change) to each of these factors. The coefficient on the change in experience estimates the 1996 return to experience.

The local demand conditions are captured with a change in the district unemployment rates from 1989 to 1996 and a time invariant dummy for Prague.[7] The coefficient on the former tells us the relative effect of changes in local demand conditions on wages in 1996 (capturing the wage curve hypothesis of Blanchflower and Oswald, 1995), while the coefficient on the dummy for Prague indicates whether wages of people residing in the capital city grew faster or slower than wages of those living in other areas of the country.

We estimate the wage effect of job characteristics and job changes in several ways. We begin by estimating the effect of changing a job vs. staying on the job with a dummy variable. This corresponds to the basic "stayer-mover" model of Blumen, Kogan and McCarthy (1955). We then extend this framework by examining whether wage changes differed by the number of job changes over the period. We also test whether the wage effect depends on whether a person changed her industrial sector when she changed her job (we account for eight sectors).

[7] Since unemployment rates were zero in 1989, this variable is effectively the unemployment rate in 1996.

In line with the importance of industrial restructuring of firms and the growth of new firms in the transition context, we next estimate specifications that show whether moving from a job in the old sector to a job in the new vs another job in the old sector had an impact on wages. Since in pre-tests we did not find major differences in the estimates for state-owned enterprises (SOEs) and privatized SOEs, we present results in which we define the old sector as comprising SOEs, privatized SOEs, SOEs in the process of privatization, coops, and public administration. The new sector includes both entrepreneurial self-employment and jobs in new private firms.

We recognize that patterns of wage change may differ for individuals who changed jobs by voluntarily quitting their job versus being involuntarily laid-off. Herein also lies an econometric problem with estimating equation (3). Job termination is only exogenous to the laid off workers. Those who change jobs by quitting do so on the basis of an evaluation of the net expected gain from quitting vs. staying on the job.

3.2 Correcting for Selectivity Bias

The fact that a person chooses to change jobs creates a problem of selectivity bias in estimating the impact of quitting on the net wage change. Briefly, the problem is one where we want to estimate the net benefit (wage change) of quitting, however we only observe wage changes for those who quit and wage changes for those who stay. One component of the net benefit is the market wage individuals could obtain if they quit and change jobs compared with what they could obtain if they stay in the job. Therefore, among the determinants of the net benefit are factors that also affect the income received in either place. An analysis of income in a sample of quitters must account for the incidental truncation (non-randomness) of the quitter's income on a positive net benefit. Likewise, the income of the stayer is incidentally truncated on a non-positive net benefit. The model implies an income after moving for all observations, but we observe it only for those who actually do move.

Hence, we use Heckman's (1979) two-stage procedure to produce corrected versions of the wage regressions. We estimate the following selection equation using a standard probit model:

$$s_i^* = Z_i' a + v_i, \quad (5)$$

where s_i^* is the latent value of quitting a job for worker i; we observe a quit if s_i^* >0. Vector Z_i captures variables thought to determine quits; v_i captures unobserved individual specific costs of quitting and $v_i \sim N(0,1)$.

The decision to quit can be modeled in one of several ways. Following the human capital and job search models (e.g., Jovanovic, 1979 and Burdett, 1978) a simple illustrative model of the worker decision developed by Farber (1999) starts as follows. Let W_a represent the best alternative wage available to the worker in the market. This is the value of the general skills that the worker brings to the labor market. The work is also rewarded for specific capital inherent in the match between the worker and the firm (Y). Hence the wage paid to the worker by the current firm is:

$$W = W_a + \lambda Y, \quad (6)$$

where Λ is the worker's share of the value of specific capital. In the simplest world, where there is complete information about worker productivity and no costs of mobility, the worker will not quit as long as the firm pays the worker even a small amount more than her alternative wage ($W > W_a$), which implies $\lambda>0$. In order to generate quits in this model, some randomness in the alternative wage needs to be introduced. Burdett's (1978) model of quits relies on job search of employed workers; a simplified version of this model has a wage offer W_o arriving each period drawn from some wage offer distribution with mean W_a. Hence, the wage offer can be expressed as $W_o = W_a + \varphi$ where φ is a random variable with mean zero. A worker quits if the wage offer exceeds the current wage, which implies the condition for quitting of $W_a + \varphi > W_a + \lambda Y$, or $\varphi > \lambda Y$.

Clearly, the probability that the wage offer exceeds the current wage is greater the lower the person's wage is in the distribution of wage offers. Hence, a worker is more likely to quit the lower his wage is relative to the mean alternative wage. Search theory would also predict that the arrival rate of wage offers would be higher in tighter labor markets, so workers would be more likely to quit when demand is high.

As a corollary, workers might be more likely to quit if they are working in a job that is in a rapidly declining industry – the value of quitting and taking a job offer would exceed the value of staying, with a high probability of being laid-off in the next period. Moreover, it could be argued that the arrival rate of wage offers would be higher in large urban centers, where the structure of the labor market is more diversified and there are more different types of jobs to choose from.

Yet, since mobility is not costless and information about jobs is not perfect, a model of quitting behavior should incorporate these costs/risks that the individual is considering in the decision to quit. One way to incorporate uncertainty in the decision making is to consider the individual's decision as part of a household maximization decision, assuming the household members have common preferences and pool all sources of income. Modeling the interaction between individual household members is important for understanding how individuals are able to reduce their exposure to uncertainty.

Just as individuals can reduce their exposure to various investment risks by holding a diversified portfolio of assets, a household that shares earnings can ex-ante reduce its exposure to labor market uncertainty by holding a diversified portfolio of jobs (Stillman, 2000). Because of data limitations, we are not able to incorporate the jobs of other household members in the estimation strategy. However, we incorporate variables reflecting whether a person is married or not, the number of children in the household and the level of per-capita household income to proxy the wealth of the household. The assumption is that married people, especially those with more children, may be less likely to take risk given their family responsibilities. We also assume that wealthier households are in a position to take bigger risks and hence a worker in such a household might be more likely to quit for a new (uncertain) job.

Hence, the Z_i vector contains the following variables: dummy variables for gender, marital status, one child, and two or more children; the log of per capita household income (in 1996), education, age, the log of the wage of the person relative to the mean wage of all individuals with the same level of schooling

(in 1996)[8], dummy variables for Prague and for the three sectors where employment declined dramatically (agriculture, mining and utilities and heavy manufacturing). We estimate the probit model with these variables and obtain the Mill's ratio to correct for selectivity bias in the wage regressions for quitters and stayers. For satisfactory identification, it is important that Z_i excludes at least one element that is in the wage regressions, or else the identification of the model hinges completely on functional form assumptions. Given we have several variables that affect the value of quitting but do not directly affect wages, we are confident that we have satisfactory identification.

3.3. The Decision to take a Job in the Old vs. New Sector

Workers who leave their job to take another one (either by quitting or being laid-off) also choose how/where to search and which job offer to accept as part of their optimization decision. In modeling the individual's decision, we assume that those who stay and quit are simultaneously deciding on whether to stay in the current job, quit and take a job offer in the new sector, and quit and take a job offer in the old sector.[9] On the other hand, those who are laid-off receive an exogenous shock and we assume that they only decide on which job to take *after* receiving the shock. Hence, we use the following models: a) a standard multinomial probit (rather than an ordered probit) for stay vs. quit and take a job in the new sector vs. quit and take a job in the old sector; b) a standard probit, conditional on lay-off, for the decision to take a job in the new vs. the old sector.

Using the same Z_i vector for these probit models, we derive the Mill's ratios to estimate selection corrected regressions of changes in log wages for these five groups of people.

[8] The relative wage variable is ln(w_i/average wage of the individual's education group), where education groups are defined as: apprentice, high school education without the CGE exam, high school education with the CGE exam and university and higher.

[9] Given that 90% of those individuals in our sample who quit did not experience a spell of non-employment, the assumption of simultaneity of the decision to accept an offer from one of the sectors and quit is consistent with the data. (We define a spell of non-employment as not working for more than one month. We assume non-employment spells of one month or less are being used for leisure.)

Section 4: Empirical Results

4.1. Summary Statistics

We start with a discussion of the means and standard deviations of the main variables. As may be seen from the values of real wage change in row 2 of Table 1 (at the end of this paper), the 1989-96 change in real full-time wages (nominal wages deflated by the consumer price index) was negative for all the principal groups of individuals.[10] For the entire sample, the real earnings fell by 16.5 percent between January 1989 and December 1996. The decline was deeper (22.8 percent) for those who stayed in their original jobs – "stayers." Those who quit their job experienced a smaller decline of about 8.5 percent, while those that were laid off lost 20.3 percent.

Therefore, on average the laid off individuals fared similarly as those who stayed in their original jobs. When one considers the 1996 destination of the movers, irrespective of whether they quit or were laid off since 1989, one observes that those who moved to the "old state sector" (public administration, state-owned enterprises and privatized firms) lost 21.6 percent, as compared to a smaller loss of 5.1 percent for those who moved to the newly formed private firms. Hence, individuals who quit and those who moved to the new private firms gained more in wages than those who stayed, were laid off, or moved into the old state sector.

In terms of other variables, we see that those who quit and those who move into new private firms are on average 2-4 years younger (and have 2-4 years less work experience) than those who stay in their original jobs, are laid off or move to the old sector. Women constitute 43.7 percent of the sample and they are found disproportionately among laid off workers (47.3 percent) and those moving into the old state sector (46.4 percent). They are under-represented among those moving to the sector of new private firms (37.1 percent). The mean of the variable 'change in experience between January 1989 and December 1996' (constructed as eight years minus the duration of unemployment spells during that time) is 7.9 years. This reflects the fact that only 7.6 percent of the sample experienced spells of unemployment

[10] Overall, real consumer wages did not reach their 1989 level until 1998.

and that on average these spells were rather short. Note that by construction, individuals who did not change jobs accumulated eight years of new (post-communist) work experience. Indeed, the extent of non-employment is greatest among individuals who suffered from layoffs, as their change in experience averages 7.7 years.

In view of the booming economy in the capital city of Prague, it is not surprising to find that individuals living in Prague suffer less from layoffs and move more frequently to the new private sector. Married individuals tend to be stayers and suffer less from layoffs, while individuals with children tend to be disproportionately located among those who quit and those who move to the new sector. Interestingly, individuals with two or more children suffer less from layoffs than individuals with fewer or no children, a finding that may signal the presence of social consideration in the allocation of layoffs. As might be expected, layoffs are observed in the declining industries.

4.2. A Simple Model of the Determinants of the 1989-96 Wage Change

We start our analytical discussion by providing a simple overall perspective on the determinants of the change in full-time wages between December1989 (the end of communism) and December 1996 (mature transition). We do so by presenting in Table 2 the estimates of several ordinary least squares (OLS) regression equations that relate the 1989-96 logarithmic change in wage to five sets of explanatory variables.

In line with the model presented in Section 3.1, the five specifications overlap in that they all include as common explanatory variables education, experience in 1989, change in experience between 1989 and 1996, a gender dummy variable coded 0 if the individual is a man and 1 if woman, a dummy variable for Prague, 1992 district unemployment rate, and eight dummy variables to capture (together with a constant term) any effects on wage changes from a job change that involves a change from one sector to another in the nine principal sectors of economic activity.

Beginning with the human capital variables, education is positively related to wage changes, with each year of education yielding about a 3 percent wage gain over the 1989-96 period. As may be seen from Table 2, this result is robust to differences in specification. Experience in 1989 captures the effect

of work experience gained under communism on wage changes during the 1989-96 period. This effect is found to be negative and highly significant, although the absolute size of the coefficient is small. It suggests that each year of work experience obtained under communism lowers the wage change during the 1989-96 transition period by 0.4 to 0.5 percent. The coefficient on '1989-96 change in experience' captures the returns to new experience gained in the post-communist period.

We would expect *a priori* the significance of this coefficient to be low since the variable has small variance (see Table 1). Interestingly, the coefficient is positive, ranging from 2.4 to 4.3 percent, and in some specifications it is or comes close to being statistically significant at the 10 percent test level. These estimates suggest that post-communist work experience is more valuable during the transition than experience obtained under communism. Finally, we note that the difference between men's and women's 1989-96 wage change is not statistically significant.

The variables capturing the effect of local demand conditions had the expected signs. Prague residence yields a 10-13 percent wage gain in all specifications. People living in districts with higher unemployment rates have lower wage gains than people living in districts with lower unemployment, providing an indirect support for the wage curve hypothesis (Blanchflower and Oswald, 1995). Finally, the net change of sector dummy variables indicates that individuals leaving (entering) mining and utilities lost (gained), while those leaving (entering) heavy manufacturing gained (lost). Other intersectoral moves are not associated with significant wage gains or losses during the 1989-96 period.

To capture the wage effects of job mobility, we include in column 1 of Table 2 as a regressor a dummy variable that is coded 1 if the individual changed jobs and 0 otherwise. The estimated coefficient on this variable is 0.09 and it is highly significant, indicating that individuals who moved to a new job over this period gained 9 percent relative to those who did not move, *ceteris paribus.*[11] In column 2 we present the effects of moving into a new private firm and moving into an old sector firm, relative to

[11] These findings are consistent with those from Hunt's (1998) study of former East Germany and both counter the predictions of Blumen et al. (1955) "mover-stayer" model.

staying in one's 1989 old sector job. The wage effect of moving into the new sector is 9.4 percent and it is statistically significant at the one percent confidence level.

In contrast, the effect of moving into the old sector is not significantly different from staying in the old job. In column 3 we assess the wage effect of moving across several jobs during this period while controlling whether the individual ended in a new sector job. We find that holding two and four or more (but not three) jobs results in a 7 percent wage gain, while the effect of moving into the new sector remains positive and significant at 7.3 percent. Multiple job holding is hence associated with wage gains but the effect is not uniform or monotonic.[12] The estimates in column 4 indicate that quits have a positive 6 percent wage effect, while layoffs have no significant effect on wages, when one controls for the effects of the basic set of variables and for whether the movers go into the new sector.

Finally, in column 5 we separate the wage effects of different job separations (quits and layoffs) and destinations (new vs. old sector jobs). The results indicate that the workers who quit or were laid off and moved into a new sector job obtain on average a similar wage gain (13.4 and 11.2 percent, respectively) relative to stayers. Workers who quit and move to the old sector register a 3.8 percent gain that is not statistically significant. However, workers who are laid off and end up in the old public sector suffer a 11.2 percent wage decline compared to the stayers. Overall, the findings in Table 2 demonstrate that movers on average gain relative to stayers and that the positive wage effect is associated with quitting and also with moving into the new sector. In contrast, the wage effect of being laid off is neutral, while the effect of moving into the old sector is negative.

4.3. Determinants of Quits and Selection of New Jobs

In the analysis of the preceding section, we have implicitly assumed that individuals are exogenously assigned to the categories of stayers, quitters and laid off individuals, as well as to the destination of their move (new vs. old sector firms). While the decision to lay off a worker is arguably a decision of the firm and as such is exogenous to the worker, the decision to quit or stay is probably not.

[12] This finding is similar to that of Keith and McWilliams (1995) using US data.

Similarly, once the worker is laid off, one can argue that rather than being randomly assigned to a firm, she makes a decision on whether to join a firm in the new sector or get a job in the traditional state owned sector. In this section, we present estimates that take this decision-making process into account.

We have estimated three probit models. The marginal (slope) coefficients of the explanatory variables and the associated standard errors are presented in Table3. The estimated (raw) probit coefficients and standard errors are presented in Appendix Table A1.

The first set of coefficients in Table 3 gives the estimated effects of marginal changes in the explanatory variables on the probability that a person quits his/her job rather than staying. As may be seen from the binary probit estimates in panel (a) of Table 3, the probability of quitting is negatively related to being married or older and having a relatively high wage in the original job. It is positively related to total household per capita income and the vacancy (job opening) rate in the district. Moreover, we also detect a positive and almost significant effect on quits of having children (especially two or more), being more educated and being in a declining sector. Finally, gender and Prague location have no systematic effect. With the possible exception of the children effect, these findings are intuitively plausible.

Considering the negative effects first, married individuals tend to be more risk averse and incur higher transaction costs in moving. Older individuals may also be more risk averse and incur higher transaction costs and they may have accumulated more job specific capital for which they will lose a return.[13] Moreover, older people have a shorter remaining working life in which to recoup the return on investment in moving. Relatively high income in the original job makes it less likely that a higher paying job will be found elsewhere. In contrast, higher household income relaxes the budget constraint and makes investment in moving more feasible, while a high number of vacancies signals high demand for labor and increases the expected rate of return on moving.

[13] We would have preferred to use the tenure variable here, but we did not obtain this information from people on their 1989 job.

In panel (b) of Table 3 we present estimates of a probit model that postulates that an individual makes the following two decisions simultaneously: a) to quit vs. stay and b) to take a new vs. old sector job. This multinomial model may be interpreted as assuming that an individual considers job offers from firms in the two sectors and then decides whether to stay or whether to accept one of the offers and quit his old job. The fact that only 10 percent of the quitters in our sample experienced an unemployment spell of more than one month is consistent with this conceptual framework.

In interpreting the coefficients in panel (b) of Table 3, it is important to remember that in a multinomial model the estimated coefficients reflect the relative risk (odds) ratios. Each marginal coefficient in panel (b) of Table 3 is hence expressed in the form of $\exp(\beta)$ rather than β itself. In terms of interpretation, this means that the marginal effect is negative ($\beta < 0$) when $\exp(\beta) < 1$ and that it is positive ($\beta > 0$) for $\exp(\beta) > 1$. As may be seen from column 1 of panel (b), the probability of quitting and moving to the new sector (rather than staying) is negatively related to being a woman (the coefficient is significantly smaller than unity). The probability is positively related to higher per capita household income, education and district vacancy rate (these three coefficients are significantly greater than unity).

On the other hand women and individuals living in districts with higher vacancy rates are more likely to quit and move to the old sector rather than stay. This probability is negatively related to being married, being older and having an above average wage at the original job. Being a woman hence decreases the probability that a person quits for the new sector but increases the probability that the person quits and moves to the old sector. This could be interpreted as indicating that women tend to go for less risky alternatives, as does the finding that high household income is conducive to quitting for the new sector and being married and older increases the probability of quitting for the old sector. The finding that a relatively high income at the original job reduces the probability that a person quits for an old sector job is plausible since old sector jobs do not offer higher average incomes.

In panel (c) of Table 3 we present the estimated effects of marginal changes in the explanatory variables on the probability that a laid off person enters the sector of new private firms rather than the

sector of the old state-owned and privatized firms and government agencies. The interpretation of these binomial probit coefficients is analogous to that in panel (a). As may be seen from panel (c), the only coefficient that is statistically significant is the positive coefficient on per capita household income, indicating that laid off individuals from households with greater income tend to go to the new rather than the old sector.

This is an intuitively acceptable finding since working in the new sector is more risky. The fact that the probability a laid off person enters the new vs. old sector is unrelated to his/her other demographic characteristics, local demand conditions and growth/decline in sector of previous job is interesting. It suggests that the allocation of laid off individuals to firms in the two sectors may be related more to firm than individual characteristics. If so, it would suggest that the re-employment process is more a hiring process of the firm than a search process of the individual. Of course, the statistically insignificant estimates could in part be also brought about by the fact that they are based on only 218 observations.

4.4. The Determinants of the 1989-96 Wage Change with Adjustment for Selectivity

We next present estimates of the determinants of the 1989-96 wage changes when we take into account the selectivity issues analyzed in the preceding section. In line with our conceptual framework, we first examine in Table 4 the wage changes of the following three groups: those who stayed in their 1989 job, those who quit and those who were laid off. We then present in Table 5 our analysis of the wage changes of five groups: those who quit for jobs in the new sector, quit for jobs in the old sector, were laid off and moved to a job in the new sector, and were laid off and moved to the old sector. The coefficients on Heckman's λ (the Mill's ratio) are significant in all equations, indicating that there is a correlation between the unobserved factors that determine the choices (of staying and not staying, etc) and the individual's wage changes in 1989-1996.

4.4.1. *Stayers, Quitters, Laid-off*

We begin with the determinants of wages changes for stayers and quitters. As may be seen from Table 4, several of the coefficients on the wage change regressions for stayers and quitters are affected by

the correction for selectivity bias. The most notable change is on the coefficient for experience in 1989. In the uncorrected regressions, it would appear that people with more labor market experience in 1989 who quit have a lower wage change than those with less experience who quit. However, after correcting for selectivity bias, the coefficient is positive and significant. Similarly, for stayers, the coefficient on this experience variable is not significantly different from zero in the uncorrected wage equation and it becomes positive and significant in the corrected equation.

Hence, once we take into account the fact that older people are less likely to move, we find that the wage gain is in fact positive for people with more experience at the start of the period. For quitters, the gain from finding a job in the new sector is reduced in the corrected equation (falling from 8.4% to 7.7%). Finally, the other noteworthy coefficient that is affected by the selection correction is the district unemployment rate in 1996. It is estimated at –5.7 (-5.8) for quitters (stayers) in the uncorrected equation and it becomes not significantly different from zero for both in the corrected equation. Hence demand conditions affect the decision to quit vs. stay but not the wages of quitters and stayers.

In general, the coefficients from Table 4 indicate that the more educated individuals experienced a faster growth of wages than the less educated, irrespective of whether they stayed, quit or were laid off. This finding is consistent with Munich, Svejnar and Terrell (1999) finding on the rising rate of return to education during the transition.

Work experience accumulated as of 1989 has a positive effect on wages of stayers and quitters, while its effect is negative and almost significant for the laid off individuals. Hence, workers who succeed in keeping their existing jobs or voluntarily move to new jobs are able to secure a positive rate of return on their communist era work experience. Those who are laid off tend to find that the effect of this experience is nil or negative. (It is worth noting again that these results lead to different conclusions than those obtained when one does not correct for selectivity.)

The effect of a change in experience between 1989 and 1996 can only be measured for those who quit or were laid off since stayers by definition all accumulated eight years of new experience. The effect of a change in experience for quitters and laid off individuals is positive but it is statistically insignificant.

Hence, when one breaks down the data into the three categories, one cannot detect a significant wage effect of change in work experience.

The gender effect is insignificant for all three categories of workers, as it was in the more aggregated regressions of Table 2. Prague location has an insignificant effect for stayers and quitters, but the effect is large (32.4 percent) and statistically significant for the laid off individuals.

The effects of changes of industries are as follows: Being laid off from mining and utilities is associated with a major decline in earnings, while quitting from heavy manufacturing is found to have a significant positive effect on earnings.

Finally, both quitters and laid off workers who move into the new sector jobs experience wage gains compared to those who moved to the old sector jobs. The wage effect is estimated at about 8 percent for quitters and 21 percent for individuals who were laid off.

4.4.2. *Quit for New vs. Old Sector and Laid-off for New vs. Old Sector*

The estimates in Table 5 give the effects of explanatory variables on the 1989-96 wage change of four groups: those who quit for jobs in the new sector, quit for jobs in the old sector, were laid off and moved to the new sector, and were laid off and moved to the old sector.

As may be seen from the table, education has a positive effect on the wage change of the first three groups, but negative and insignificant effect for those who were laid off and moved to the old sector. The increase in the rate of return to human capital is hence related to job and sectoral mobility, being insignificant for those who suffer involuntary separation and move to the old sector. The coefficient is not affected by correction for selectivity bias.

The effect of 1989 experience is positive for quitters and not significantly different from zero for laid off individuals, irrespective of whether they moved to the new or old sector. Consistent with the findings in Table 4, the sign of this coefficient for each type of quitter changes from negative to positive when it is corrected for selectivity bias. We also learn, that the return to experience cumulated in 1989 is higher for quitters who moved into the new sector than for quitters who took a job in the old sector.

Clearly the new sector with its market determined wage is rewarding experienced people (who quit) more than the old sector. On the other hand, the effect of a change in experience between 1989 and 1996 is not statistically significant in all cases except for laid off individuals who moved to the old sector. These individuals register a positive and marginally significant coefficient of 14.5 percent. As mentioned earlier, there is very little variation in this wage for quitters so this result is not surprising.

New information revealed by this set of regressions is that there are significant differences between men and women's wage changes during 1989-1996, in certain categories. Moreover, this effect is not revealed until the regression is corrected for selectivity bias. Women who quit and accept a job in the new sector experience a sizable positive wage gain (40.8 percent) relative to men. On the other hand, women who quit and accept a job in the old sector experience a 21.7 percent lower wage gain than comparable men. Hence, we find that once we control for the fact that women are less likely to enter the new sector, those who do obtain a considerable wage premium over the men who quit and enter the new sector.

The interesting finding with respect to the "Prague effect" in Table 4 is magnified in Table 5. We learned earlier that among those laid off, being laid-off in Prague had its advantages over being laid off in the rest of the country. In Table 5, we find that those who gain the most are those who are laid-off and find a job in the old sector. Given that Prague is the seat of the government, this is not surprising.

Finally, the coefficients on the dummy variables denoting net sector change continue to be mostly insignificant, perhaps due to sample size. The only one that is significant is the coefficient for those who quit jobs in heavy manufacturing and found a new one in the public sector. They experienced a substantial wage gain.

Section 5: Conclusions

Our analysis of job and wage changes in the Czech Republic between 1989 (the last year of communism) and 1996 indicates that job mobility, the sector of destination of movers and wage changes are significantly related. In particular, we draw the following conclusions:

1. The Czech workers experienced a significant fall in real wages between 1989 and 1993. Wage growth occurred after 1993 but even by the end of 1996 the real wage of our entire sample was 16.5 percent below that of January 1989. The decline was deeper (22.8 percent) for those who stayed in their original jobs and those who were laid off (20.3 percent). It was less pronounced for those who quit (8.5 percent). Movers to the "old state sector" (public administration, state-owned enterprises and privatized firms) lost 21.6 percent, as compared to a smaller loss of 5.1 percent for those who moved to the newly formed private firms. Hence, individuals who quit and those who moved to the new private firms gained more in wages than those who stayed, were laid off, or moved into the old state sector. These findings are confirmed in a multivariate regression framework.

2. Those who quit and those who move into new private firms are on average 2-4 years younger than those who stay in their original jobs, are laid off or move to the old sector. Women are found disproportionately among laid off workers and those moving into the old state sector.

3. Our analysis of the probability that an individual quits rather than staying in her original job shows that the probability of quitting is negatively related to being married or older and having a relatively high wage in the original job. It is positively related to total household per capita income and the vacancy (job opening) rate in the district. Moreover, we also detect a positive and almost significant effect on quits of having children (especially two or more), being more educated and being in a declining sector. Finally, gender and Prague location have no systematic effect. With the possible exception of the children effect, these findings are intuitively plausible.

4. When we analyze a model in which an individual decides simultaneously on whether to a) quit and take a new job vs. stay and b) quit and take an old sector job vs. stay, we find that the probability of quitting and moving to the new sector (rather than staying) is negatively related to being a woman and positively related to higher per capita household income, education and district vacancy rate. On the other hand women and individuals living in districts with higher vacancy rates are more likely to quit and move to the old sector rather than stay. This probability is negatively related to being married,

being older and having an above average wage at the original job. Being a woman hence decreases the probability that a person quits for the new sector but increases the probability that the person quits and moves to the old sector. This could be interpreted as indicating that women tend to go for less risky alternatives, as does the finding that high household income is conducive to quitting for the new sector and being married and older increases the probability of quitting for the old sector. The finding that a relatively high income at the original job reduces the probability that a person quits for an old sector job is plausible since old sector jobs do not offer higher average incomes.

The probability that a laid off person enters the new rather than the old sector is affected positively by per capita household income, but no other variable has a significant effect. The finding that laid off individuals from households with greater income tend to go to the new rather than the old sector is an intuitively acceptable since working in the new sector is more risky. The fact that the probability a laid off person enters the new vs. old sector is unrelated to his/her other demographic characteristics, local demand conditions and growth/decline in sector of previous job suggests that the allocation of laid off individuals to firms in the two sectors may be related more to firm than individual characteristics. If so, it suggests that the re-employment process is more a hiring process of the firm than a search process of the individual. The statistically insignificant estimates could in part be also brought about by the fact that they are based on only 218 observations.

5. When we adjust the 1989-96 wage change equations for selectivity of individuals in terms of quitting vs. staying, we find that the more educated individuals experienced a faster growth of wages than the less educated, irrespective of whether they stayed, quit or were laid off. Work experience accumulated as of 1989 has a positive effect on wages of stayers and quitters, while its effect is negative and almost significant for the laid off individuals. Hence, workers who succeed in keeping their existing jobs or voluntarily move to new jobs are able to secure a positive rate of return on their communist era work experience. Those who are laid off tend to find that the effect of this experience is nil or negative. (It is worth noting that these results are different than those obtained when one

does not correct for selectivity.) The effect of a change in experience for quitters and laid off individuals is positive but it is statistically insignificant. The gender effect is insignificant for all three categories of workers, as it was in the more aggregated regressions of Table 2. Prague location has an insignificant effect for stayers and quitters, but the effect is large (32.4 percent) and statistically significant for the laid off individuals. Being laid off from mining and utilities is associated with a major decline in earnings, while quitting from heavy manufacturing is found to have a significant positive effect on earnings. Finally, quitters and laid off workers who move into the new sector jobs experience wage gains of 8 and 21 percent, respectively.

6. In the final analysis, we have taken selectivity into account while analyzing the determinants of wage changes of those who quit for jobs in the new sector, quit for jobs in the old sector, were laid off and moved to a job in the new sector, and were laid off and moved to the old sector. We find that education has a positive effect on the wage change of the first three groups, but negative and insignificant effect for those who were laid off and moved to the old sector. The increase in the rate of return to human capital is hence related to job and sectoral mobility, being insignificant for those who suffer involuntary separation and move to the old sector.

   The effect of 1989 experience is positive for quitters and not significantly different from zero for laid off individuals, irrespective of whether they moved to the new or old sector. The return from 1989 experience is higher for those quitters who moved into the new sector than those who took a job in the old sector. The new sector with its market-determined wage is rewarding experienced people (who quit) more than the old sector. On the other hand, the effect of a change in experience between 1989 and 1996 is not statistically significant in all cases except for laid off individuals who moved to the old sector. These individuals register a positive and marginally significant coefficient of 14.5 percent. As mentioned earlier, there is very little variation in this wage for quitters so this result is not surprising.

There are significant differences between men and women's wage changes during 1989-1996, in certain categories. Women who quit and accept a job in the new sector experience a sizable positive wage gain (40.8 percent) relative to men. On the other hand, women who quit and accept a job in the old sector experience a 21.7 percent lower wage gain than comparable men. Hence, once we control for the fact that women are less likely to enter the new sector, those who enter obtain a considerable wage premium over the men who quit and enter the new sector.

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Table 1: Summary Statistics

Variable Name	All Sample		Stayers		Quit Job		Laid Off		Old State Sector		New Private Sector	
	Mean	(St. Dev.)	Mean	(St. Dev.)	Mean	(St. Dev.)	Mean	(St. Dev.)	Mean	(St. Dev.)	Mean	(St. Dev.)
(W96-W89)/W89 in '89 Kcs	-0.165	0.563	-0.228	0.553	-0.085	0.596	-0.203	0.415	-0.216	0.548	-0.051	0.588
Log of relative waged	0.001	0.428	0.030	0.426	-0.021	0.426	-0.045	0.444	0.000	0.421	0.007	0.444
Relative wage[d]	1.098	0.554	1.123	0.500	1.079	0.606	1.059	0.575	1.090	0.489	1.119	0.679
Log per cap HH inc.	8.461	0.437	8.461	0.395	8.473	0.472	8.412	0.471	8.453	0.413	8.486	0.491
Age	35.918	9.153	37.663	8.315	33.846	9.700	36.278	8.863	36.652	9.101	34.358	9.137
Women (dummy)	0.437	0.496	0.432	0.496	0.433	0.496	0.473	0.500	0.464	0.499	0.371	0.483
Education (years)	12.537	2.408	12.466	2.382	12.712	2.426	12.159	2.403	12.493	2.415	12.665	2.415
Experience (years)	16.341	9.425	18.128	8.765	14.127	9.805	17.071	8.953	17.118	9.434	14.654	9.244
Experience in 1989-1996	7.892	0.411	8.000	0.000	7.814	0.556	7.711	0.520	7.920	0.394	7.830	0.449
Prague	0.107	0.309	0.101	0.302	0.116	0.320	0.098	0.298	0.101	0.301	0.123	0.328
Married	0.808	0.394	0.831	0.375	0.791	0.407	0.771	0.421	0.808	0.394	0.807	0.395
One child	0.203	0.402	0.185	0.389	0.225	0.418	0.192	0.395	0.190	0.393	0.223	0.416
Two+ children	0.193	0.395	0.164	0.370	0.242	0.429	0.131	0.338	0.180	0.384	0.220	0.414
Sector of Job in 1989:												
Agriculture	0.090	0.286	0.064	0.244	0.108	0.311	0.135	0.343	0.075	0.263	0.121	0.327
Mining and Utilites	0.065	0.246	0.076	0.266	0.044	0.205	0.094	0.293	0.072	0.259	0.046	0.209
Construction	0.076	0.265	0.062	0.241	0.090	0.286	0.082	0.275	0.064	0.244	0.103	0.304
Light Manufacturing[a]	0.153	0.361	0.154	0.361	0.150	0.357	0.164	0.371	0.152	0.359	0.163	0.369
Heavy Manufacturing[b]	0.201	0.400	0.227	0.419	0.178	0.383	0.168	0.375	0.215	0.411	0.173	0.378
Trade, Restaurants and Hotels	0.091	0.288	0.074	0.261	0.106	0.308	0.111	0.314	0.077	0.267	0.117	0.322
Fin., Real Est., Trans.&Comm.	0.121	0.326	0.118	0.323	0.130	0.337	0.090	0.287	0.118	0.323	0.129	0.335
Public Administration	0.204	0.403	0.224	0.417	0.194	0.395	0.156	0.363	0.227	0.419	0.149	0.356
Change in sector of Job:												
Agriculture*	-0.035	0.220	0.000	0.000	-0.063	0.303	-0.082	0.289	-0.014	0.151	-0.081	0.317
Mining and Utilites*	-0.013	0.169	0.000	0.000	-0.015	0.210	-0.066	0.307	-0.006	0.141	-0.029	0.213
Construction *	0.007	0.225	0.000	0.000	0.012	0.305	0.016	0.339	-0.009	0.141	0.040	0.346
Light Manufacturing*[a]	-0.004	0.285	0.000	0.000	-0.015	0.392	0.025	0.405	-0.004	0.206	-0.013	0.404
Heavy Manufacturing*[b]	-0.021	0.300	0.000	0.000	-0.043	0.424	-0.033	0.373	-0.010	0.234	-0.046	0.410
Trade, Restaurants and Hotels*	0.036	0.289	0.000	0.000	0.069	0.394	0.070	0.415	-0.014	0.156	0.157	0.440
Fin., Real Est., Trans.&Comm.*	0.020	0.296	0.000	0.000	0.030	0.414	0.074	0.388	0.021	0.223	0.017	0.412
Public Administration*	0.010	0.294	0.000	0.000	0.026	0.410	-0.004	0.401	0.036	0.247	-0.046	0.365
Decline Sector[c]	0.154	0.361	0.140	0.347	0.152	0.359	0.229	0.421	0.147	0.354	0.167	0.373
District unemp. rate in 1996 (%)	-3.706	0.727	-3.673	0.724	-3.761	0.723	-3.640	0.743	-3.673	0.721	-3.790	0.737

Avg. district vacancy rate (%)	-4.255	0.399	-4.278	0.419	-4.229	0.381	-4.258	0.371	-4.266	0.408	-4.226	0.379
Dummy for stayed	0.476	0.500	1.000	0.000	0.000	0.000	0.000	0.000	0.696	0.460	0.000	0.000
Dummy for changed jobs	0.524	0.500	0.000	0.000	1.000	0.000	1.000	0.000	0.304	0.460	1.000	0.000
Held two jobs	0.164	0.371	0.000	0.000	0.392	0.488	0.000	0.000	0.115	0.320	0.265	0.442
Held three jobs	0.292	0.455	0.000	0.000	0.471	0.499	0.898	0.303	0.156	0.363	0.591	0.492
Held four+ jobs	0.068	0.252	0.000	0.000	0.136	0.343	0.102	0.303	0.033	0.179	0.144	0.351
Dummy if unemployed	0.076	0.266	0.000	0.000	0.105	0.307	0.310	0.464	0.055	0.229	0.120	0.325
Unemp. Duration	1.295	4.936	0.000	0.000	2.226	6.673	3.465	6.240	0.956	4.728	2.043	5.384
Old Sector	0.685	0.465	1.000	0.000	0.400	0.490	0.388	0.488	1.000	0.000	0.000	0.000
New Sector	0.299	0.458	0.000	0.000	0.567	0.496	0.588	0.493	0.000	0.000	1.000	0.000
No. of Observations	2343		1116		982		245		1604		701	

*The "change in industrial sector" variable was constructed as -1 if the person left the sector, 1 if the person moved into the sector, and 0 otherwise.
a Light manufacturing includes food, textiles, paper and chemicals, and other manufacturing.
b Heavy manufacturing includes machinery, metals and equipment.
c Declining sectors were agriculture, mining and utilities and heavy manufacturing.
d Relative wage = wi/(average wage of the individual's education group), where education groups are defined as: apprentice, high school education without CGE exam; high school education with CGE exam, university and higher.

Table 2: Determinants of 1989-1996 Wage Change
(OLS Regressions; Dependent Variable = ln(W96/W89))

	(1)		(2)		(3)		(4)		(5)	
Variable Name	Coeff.	St. Error	Coeff.	St. Error	Coeff.	St. Error	Coeff.	St. Error	Coeff.	St. Error
Constant	-1.085	0.204	-0.917	0.203	-1.046	0.214	-1.039	0.204	-1.263	0.212
Education	0.030	0.004	0.029	0.004	0.030	0.004	0.029	0.004	0.030	0.004
Experience in 1989	-0.005	0.001	-0.005	0.001	-0.005	0.001	-0.004	0.001	-0.004	0.001
Change in Experience (1996-1989)	0.043	0.024	0.024	0.024	0.037	0.025	0.036	0.024	0.029	0.024
Women	0.006	0.020	0.012	0.020	0.010	0.020	0.010	0.020	0.014	0.020
Prague	0.131	0.036	0.130	0.036	0.131	0.036	0.132	0.036	0.100	0.046
District Unemp. Rate in 1996	-2.590	0.671	-2.523	0.672	-2.531	0.672	-2.420	0.673	-0.057	0.020
Agriculture*	0.007	0.062	-0.050	0.062	-0.009	0.062	-0.006	0.062	-0.031	0.062
Mining and Utilites*	0.177	0.071	0.123	0.071	0.160	0.072	0.156	0.072	0.139	0.071
Construction *	0.027	0.059	-0.038	0.061	-0.001	0.060	0.001	0.060	-0.027	0.061
Light Manufacturing*[a]	-0.015	0.055	-0.068	0.055	-0.034	0.055	-0.029	0.055	-0.048	0.055
Heavy Manufacturing*[b]	-0.103	0.053	-0.156	0.053	-0.121	0.053	-0.119	0.053	-0.140	0.053
Trade, Restaurants and Hotels*	-0.001	0.054	-0.074	0.056	-0.035	0.056	-0.034	0.056	-0.070	0.056
Finance, Real Estate, Trans.&Comm.*	0.084	0.053	0.042	0.053	0.071	0.053	0.074	0.053	0.057	0.053
Public Administration*	-0.004	0.072	-0.069	0.072	-0.014	0.072	-0.017	0.072	-0.040	0.072
Dummy for Changed Job	0.089	0.022								
Dummy for Changed Job and in New Sector job in '96			0.094	0.025						
Dummy for Changed Job and in Old Sector in '96			-0.022	0.028						
Dummy for New sectorJob					0.073	0.029	0.072	0.028		
Dummy for held 2 jobs					0.066	0.032				
Dummy for held 3 jobs					0.032	0.031				
Dummy for held 4+ jobs					0.074	0.047				
Dummy for Quit							0.061	0.027		
Dummy for Laid-Off							0.001	0.039		
Quit and in New Sector job in 1996									0.134	0.026
Quit and in Old Sector job in 1996									0.038	0.030
Laid-off and in New Sector job in 1996									0.112	0.043
Laid-off and in Old Sector job in 1996									-0.112	0.052
N	2276		2276		2276		2276		2276	
Adjusted R2	0.091		0.091		0.093		0.093		0.096	

*The "change in industrial sector" variable was constructed as -1 if the person left the sector, 1 if the person moved into the sector, and 0 otherwise.
a Light manufacturing includes food, textiles, paper and chemicals, and other manufacturing.
b Heavy manufacturing includes machinery, metals and equipment.

Table 3: Marginals from Probits for Determinants of the Probability of:
(a) Quitting vs. Staying; (b) Quitting for New Sector Job vs. Staying and Quitting for the Old Sector Job vs. Staying; (c) Moving to New vs. Old Sector Conditional on Layoff

	(a)		(b)				(c)	
	Quit vs. Stay (base)		Quit/New Sector vs. Stay		Quit/Old Sector vs. Stay		Laid-off/New Sector vs.Laid-off/Old Sector	
Variable Name	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error
Women	-0.015	0.026	0.677	0.085	1.413	0.190	-0.074	0.079
Married	-0.067	0.031	0.871	0.135	0.694	0.111	0.061	0.085
1 child	0.052	0.032	1.266	0.197	1.050	0.185	0.101	0.092
2+ children	0.066	0.036	1.231	0.215	1.323	0.252	0.119	0.115
Per capita HH Income	0.078	0.030	1.478	0.213	1.300	0.205	0.178	0.082
Education (yrs.)	0.008	0.005	1.046	0.025	1.026	0.028	-0.007	0.015
Age	-0.011	0.002	0.950	0.007	0.962	0.008	0.006	0.005
Relative wage[a]	-0.069	0.030	0.832	0.122	0.679	0.106	0.131	0.094
Prague	-0.003	0.045	0.910	0.198	1.014	0.236	0.100	0.133
District Vacancy Rate	0.109	0.032	1.502	0.236	1.740	0.306	0.145	0.105
Declining Sector[b]	0.054	0.033	1.146	0.185	1.318	0.237	-0.068	0.087
Constant								
Pseudo R2	0.048		0.044				0.064	
N	1905		1876				218	

a Relative wage = wi/(average wage of the individual's education group), where education groups are defined as: apprentice, high school education without CGE exam; high school education with CGE exam, university and higher.
b Declining sectors were agriculture, mining and utilities and heavy manufacturing.

Table 4: Determinants of the 1989-97 Wage Change For Those Who Stayed, Quit and Were Laid-off

	Corrected for Selectivity Bias				Not Corrected for Selectivity Bias					
Variable Name	Stayed Coeff.	St. Error	Quit Coeff.	St. Error	Stayed Coeff.	St. Error	Quit Coeff.	St. Error	Laid-off Coeff.	St. Error
Education	0.035	0.006	0.033	0.007	0.036	0.006	0.034	0.007	0.027	0.013
Experience in 1989	0.022	0.002	0.010	0.003	-0.000	0.002	-0.007	0.002	-0.007	0.004
Change in Experience (1996-1989)	-	-	0.019	0.030	-	-	0.029	0.030	0.072	0.054
Women	0.004	0.024	0.012	0.036	0.010	0.026	0.027	0.036	0.070	0.062
Prague	0.046	0.063	0.117	0.083	-0.004	0.066	0.145	0.084	0.324	0.150
District Unemp. Rate in 1996	0.031	0.026	-0.028	0.037	-0.057	0.026	-0.058	0.037	0.019	0.062
Agriculture*	-	-	-0.039	0.079	-	-	-0.074	0.081	0.003	0.148
Mining and Utilites*	-	-	0.059	0.096	-	-	0.054	0.098	0.279	0.148
Construction *	-	-	-0.086	0.077	-	-	-0.080	0.079	0.050	0.138
Light Manufacturing*[a]	-	-	-0.085	0.071	-	-	-0.079	0.073	-0.070	0.130
Heavy Manufacturing*[b]	-	-	-0.176	0.066	-	-	-0.167	0.067	-0.162	0.136
Trade, Restaruants and Hotels*	-	-	-0.104	0.071	-	-	-0.100	0.072	-0.066	0.136
Finance, Real Estate, Trans.&Comm.*	-	-	-0.004	0.067	-	-	0.004	0.068	0.086	0.137
Public Administration*	-	-	-0.086	0.092	-	-	-0.083	0.094	-0.018	0.167
New Sector Job	-	-	0.077	0.037	-	-	0.084	0.038	0.210	0.068
Lambda	-0.824	0.073	-0.626	0.101	-	-	-	-	-	-
Constant	-0.524	0.132	-0.809	0.293	-1.191	0.125	-1.264	0.289	-1.363	0.473
N	973		872		973		872		240	
Adjusted R2	0.162		0.141		0.053		0.086		0.149	

*The "change in industrial sector" variable was constructed as -1 if the person left the sector, 1 if the person moved into the sector, and 0 otherwise.
a Light manufacturing includes food, textiles, paper and chemicals, and other manufacturing.
b Heavy manufacturing includes machinery, metals and equipment.

Table 5a: Determinants of the 1989-96 Wage Change For Those Who Stayed, Quit For a New Sector Job, Quit For an Old Sector Job, Laid-off and Went To a New Sector Job, Laid-off and Went To a Old sector

	Corrected for Selectivity Bias							
Variable Name	Quit for New Coeff.	St. Error	Quit for Old Coeff.	St. Error	Laid Off for New Coeff.	St. Error	Laid Off for Old Coeff.	St. Error
Education	0.031	0.011	0.029	0.009	0.043	0.017	-0.020	0.020
Experience in 1989	0.031	0.007	0.008	0.003	-0.004	0.005	-0.004	0.006
Change in Experience (1996-1989)	0.026	0.056	0.037	0.031	0.047	0.076	0.145	0.078
Women	0.408	0.092	-0.212	0.071	-0.126	0.097	0.164	0.102
Prague	0.169	0.120	0.144	0.112	0.189	0.198	0.810	0.263
District Unemp. Rate in 1996	-0.025	0.053	0.052	0.050	-0.032	0.087	0.026	0.093
Agriculture*	-0.148	0.140	0.109	0.095	0.035	0.297	-0.277	0.177
Mining and Utilites*	0.172	0.191	0.110	0.103	0.436	0.301	0.186	0.199
Construction *	-0.156	0.131	-0.017	0.100	0.171	0.281	-0.063	0.182
Light Manufacturing*[a]	-0.181	0.129	0.032	0.083	0.081	0.282	-0.215	0.155
Heavy Manufacturing*[b]	-0.195	0.126	-0.162	0.071	-0.185	0.290	0.025	0.160
Trade, Restaruants and Hotels*	-0.123	0.124	-0.094	0.095	0.068	0.280	-0.188	0.179
Finance, Real Estate, Trans.&Comm.*	-0.018	0.124	-0.025	0.075	0.195	0.297	-0.017	0.151
Public Administration*	-0.173	0.151	-0.057	0.130	0.031	0.318	-0.139	0.236
Correction for Selectivity Bias	-2.017	0.343	-1.218	0.223	0.415	0.205	0.225	0.210
constant	0.742	0.622	0.742	0.444	-1.601	0.694	-1.694	0.727
N	488		360		123		90	
Adjusted R2	0.136		0.167		0.209		0.175	

*The "change in industrial sector" variable was constructed as -1 if the person left the sector, 1 if the person moved into the sector, and 0 otherwise.
a Light manufacturing includes food, textiles, paper and chemicals, and other manufacturing.
b Heavy manufacturing includes machinery, metals and equipment.

Table 5b: Determinants of the 1989-96 Wage Change For Those Who Stayed, Quit For a New Sector Job, Quit For an Old Sector Job, Laid-off and Went To a New Sector Job, Laid-off and Went To a Old sector

	Not Corrected for Selectivity Bias							
Variable Name	Quit for New Coeff.	St. Error	Quit for Old Coeff.	St. Error	Laid Off for New Coeff.	St. Error	Laid Off for Old Coeff.	St. Error
Education	0.039	0.011	0.026	0.010	0.043	0.017	-0.021	0.020
Experience in 1989	-0.006	0.003	-0.007	0.002	-0.008	0.005	-0.007	0.005
Change in Experience (1996-1989)	0.038	0.058	0.026	0.032	0.054	0.077	0.138	0.078
Women	-0.033	0.055	0.088	0.047	-0.053	0.091	0.208	0.094
Prague	0.114	0.122	0.204	0.115	0.149	0.200	0.768	0.261
District Unemp. Rate in 1996	-0.102	0.053	0.018	0.051	0.009	0.086	0.035	0.092
Agriculture*	-0.217	0.145	0.083	0.099	0.043	0.301	-0.296	0.176
Mining and Utilites*	0.092	0.197	0.011	0.105	0.416	0.305	0.176	0.199
Construction *	-0.161	0.136	-0.055	0.104	0.193	0.285	-0.025	0.179
Light Manufacturing*[a]	-0.203	0.133	0.028	0.086	0.091	0.286	-0.205	0.155
Heavy Manufacturing*[b]	-0.213	0.131	-0.175	0.074	-0.184	0.294	0.037	0.160
Trade, Restaruants and Hotels*	-0.151	0.129	-0.149	0.099	0.109	0.284	-0.212	0.177
Finance, Real Estate, Trans.&Comm.*	-0.028	0.129	-0.033	0.078	0.212	0.301	-0.002	0.150
Public Administration*	-0.201	0.157	-0.006	0.135	0.008	0.322	-0.148	0.236
Correction for Selectivity Bias								
constant	-1.470	0.512	-0.878	0.344	-1.237	0.680	-1.395	0.673
No. of Observations	488		360		123		90	
Adjusted R2	0.075		0.098		0.186		0.174	

*The "change in industrial sector" variable was constructed as -1 if the person left the sector, 1 if the person moved into the sector, and 0 otherwise.
a Light manufacturing includes food, textiles, paper and chemicals, and other manufacturing.
b Heavy manufacturing includes machinery, metals and equipment.

Table A1: Coefficents from Probits For Determinants of the Probability of:
(a) Quitting vs. Staying; (b) Quitting For New Sector Job vs. Quitting For the Old Sector Job vs. Staying; (c) Moving to New vs. Old sector Conditional on Layoff

	(a)		(b)				(c)	
	Quit vs. Stay (base)		Quit/New Sector vs. Stay		Quit/Old Sector vs. Stay		Laid-off/New vs. Laid-off/Old	
Variable Name	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error
Women	-0.037	0.064	-0.390	0.126	0.346	0.135	-0.313	0.326
Married	-0.168	0.078	-0.138	0.154	-0.365	0.160	0.239	0.349
1 child	0.131	0.081	0.236	0.155	0.049	0.176	0.432	0.399
2+ children	0.166	0.091	0.208	0.175	0.280	0.191	0.521	0.517
Per capita HH Income	0.196	0.075	0.391	0.144	0.262	0.158	0.765	0.367
Education	0.021	0.013	0.045	0.024	0.026	0.027	-0.031	0.065
Age	-0.027	0.004	-0.051	0.007	-0.038	0.008	0.027	0.019
Relative wage	-0.174	0.075	-0.184	0.147	-0.388	0.157	0.528	0.389
Prague	-0.008	0.113	-0.094	0.218	0.014	0.233	0.438	0.613
District Vacancy Rate	0.275	0.082	0.407	0.157	0.554	0.176	0.587	0.434
Declining Sector	0.136	0.084	0.136	0.161	0.276	0.180	-0.291	0.358
Constant	0.233	0.726	-0.857	1.399	0.203	1.546	-4.387	3.390
Pseudo R2	0.048		0.044				0.064	
N	1905		1876				218	