Human Capital in Transition

Czech Students and Workers Adapting to the Market

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CERGE-EI

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To my parents

CONTENTS

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The responsibility for the content of this book rests ultimately with the author.

Introduction

For a significant part of the twentieth century, over one-third of the world's population lived under the communist system. A great portion of the population had built their careers and lived out their lives in a centrally planned economy. The end of communism was abrupt, expected by few. The change of regime changed everybody's life. The economic transition constituted a huge transfer of decision-making powers and responsibilities from the state back to its citizens. The transfer was driven by the general wisdom that a market environment - relying on the free choices of entrepreneurs, workers, employers, managers, but also parents - is a much more efficient arrangement than any form of central planning. The end of communism, not being incorporated into individual expectations, provides the social scientist a unique opportunity to study *homo economicus* in an experiment-like setup with sizeable exogenous variation in deterministic factors. Moreover, the transition experience of Central European countries also provides invaluable experience to those reform economies still at transition's door.

This book provides a partial but very detailed analysis of the transition experience of the Czech Republic. It reviews the scale and scope of human actions in those spheres of life where people a spend great deal of their available time – education and work. The neoclassical economic paradigm suggests that transition from a centrally planned to a market economy will introduce intensive interactions between supply and demand forces. On the demand side, individuals will start clamoring for more education as the returns to education start to increase in the emerging private sector. Parents and pupils will change their schooling preferences according to the changing environment. Their effective choice will be, to some extent, supply constrained by incumbent schools but the choice will, on the other hand, encourage entry of new schools and competition among schools. Growing returns to education will be an important incentive and signal guiding the invisible hand towards efficient allocation of human and other resources. Interactions of this supply-demand nature will be repetitive ones. These and other theoretical predictions are the subject of the empirical analysis presented in this book.

The first chapter provides a detailed cross-sectional comparative analysis of the communist and transition economy wage structures and returns to human capital. Communist ideology assigned a great and historical role to physical capital, while human capital was attributed a small role and was even damned in the 1950s, the early decade of the communist regime. The regime started to recognize the importance of education only later when the development gap between the West and the East became obvious. Nevertheless, a large proportion of those who were in the labor force had their wages set according to a centrally-determined wage grid which effectively enforced ideological egalitarianism. Existing wage differentials were heavily biased by loyalty to the communist regime, working class origin and personal networks. While the effects of the grid *per se* have never been formally analyzed, there has been some evidence that earnings structures in centrally planned economies were very compressed, weakly rewarding education, and that there was decompression during the transition to a market system.

The second chapter extends the cross-sectional analysis of the previous chapter by focusing in greater detail on the role of labor market mobility and self-selection of individuals. This analysis is unique in that it is based on a sample of workers observed both during central planning and in mature transition.

The third chapter sheds light on the profound changes in the schooling system following the collapse of communism. Although the changes did not have a direct impact on the existing labor force, the changes were necessarily driven by changing labor market conditions described in detail in previous chapters. For generations schools had served not only as a means of training workers, but also as a vehicle of indoctrination designed to create a "new socialist man." Education was, by law, a state monopoly designed to respond to the dictates of the plan rather than the signals of the market. Very detailed curricula were prescribed by central authorities. Parental and student preferences played little, if any, role in determining how much or what type of training was provided. Entry into coveted disciplines, while certainly influenced by ability, was also heavily determined by political or other considerations. There are numerous examples of students including the author of this book with an interest in and aptitude for study in particular subjects being forced into entirely unrelated fields because they or their parents were considered politically unreliable or because central planners had to stick to widespread quotas. In such an environment it is not surprising that the transition process included a reform to overhaul the educational system to provide greater flexibility and give far more substantial decision-making power to students and parents. One key reform involved allowing non-state schools to challenge the state education monopoly.

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Prague Daniel Münich

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Chapter I

Returns to Human Capital

1. Introduction

During a significant part of the twentieth century, over one-third of the world's population lived under the communist system. A large proportion of those who were in the labor force had their wages set according to a centrally-determined wage grid. While the effects of the grid *per se* have never been formally analyzed, there has been some evidence that the earnings structures in centrally planned economies were very compressed and that there was decompression during the transition to a market system. In this chapter, we (a) analyze returns to human capital under the communist wage grid and (b) examine how wages and returns to human capital changed in the emerging market economy as the grid was supplanted by free wage setting in the sector composed of newly created private (*de novo*) firms and a modified wage grid in the public sector.

In analyzing the shift from the Communist wage grid, the Czech Republic because it is an excellent prototype of a sudden change of regimes among the leading transition economies. In the other transition countries, such as Poland and Hungary, central planners started losing control well before the 1989 revolutions and their adherence to the wage grid diminished as bargaining between firms and planners gained in importance (see e.g., Rutkowski, 1994). In the Czech Republic, the system remained intact until the very end of the communist regime and evidence from large firm-level data sets indicates that there was no significant rent sharing by workers (Basu et al., 1999). Moreover, while the Polish and Hungarian economies had significant private sectors already before the transition, the Czech economy was almost 100 percent state owned until 1990 and then it underwent one of the most rapid and extensive privatizations in the former Soviet bloc.[1]

The human capital studies carried out on the transition economies to date have examined returns in a cross-sectional setting, using one set of individuals at an early point in time during transition and in some cases also another set of individuals (sometimes from a different survey design) at a point in time during communism.[2] We complement these studies in several ways:

(a) We estimate the determinants of wages and returns to human capital using data on the *same* individuals during a large part of the communist period and the first six years of transition.

(b) We make use of the panel data to develop and asses if some individuals had high or low wage premiums related to unobservable characteristics and whether these premiums carried over into the transition period. In particular, we develop and apply a method decomposing the variance of worker-specific wages into components due to observable determinants and unobservable determinants in the old versus new regime.

(c) We use actual years of schooling as a measure of education rather than imputed years based on the highest degree obtained. We use the information on actual years of education and highest level attained for each individual to test for the bias created by using imputed measures of schooling and to measure *sheepskin effects* (jumps in wages when degrees are received, controlling for years of education).

(d) We test directly whether education and experience gained in the communist versus post-communist periods generate the same rate of return during the transition period.

(e) We examine the impact of firm ownership on returns to human capital during the transition. Privatization and the creation of new firms are key aspects of the transition process and understanding their impact on the wage structure is of great importance.

(f) We estimate changes in the structure of wages by industry and field of study (given attained education) to assess the impact of changes in the structure of the economy on wages.

(g) Finally, existing studies by Krueger and Pischke (1995), Chase (1998), Flanagan (1998), and Vecerník (2001) provide somewhat contradictory estimates of the returns to education and experience during the communist and post-communist regimes in a similar context. We provide additional evidence and ideas about how one might reconcile the differences in the various findings.[3]

In order to carry out our analysis, we collected data on the work histories of 2,284 men from a stratified random sample of households in the Czech Republic. Most of the men worked under communism, all worked during at least part of the 1990-96 transition period, and many worked in December 1996, the date of our survey. Using these data, we analyze the evolution of the returns to education and experience in various parts of the 1948-89 communist era and during the 1991-96 period of transition from plan to market. To our knowledge, no other data set provides information on individuals for such long periods of communism and transition.[4]

We demonstrate that the communist system used the wage grid to set and maintain an extremely low rate of return to education. We also show that the transition resulted in a major increase in the rates of return to education, which reached West European levels by 1996. Unlike Flanagan (1998), we find this increase in all ownership categories of firms.[5] Hence, as the economy opened to world competition, returns to education in the public sector (SOES and public administration) and privatized state-owned enterprises did not deviate from the market-driven, *de novo* firms.

We run regressions with different specifications of the education variable, using highest level (degree) attained vs. years of education, testing for sheepskin effects and estimating returns to fields of study. We find that those who have obtained (vocational) high school and university degrees experienced more rapid rates of increase in their returns than individuals with basic education (junior high school or apprentices). The sheepskin effect is prevalent and the effect is especially detectable in transition and for higher levels of education in both regimes. Certain fields of study have experienced tremendous increases in their returns (e.g., law), while others have not gained in the new market economy (e.g., health and education). We also show that the earlier studies may overestimate the rate of return to education by using years of education imputed from the highest degree obtained rather than actual years of schooling as an explanatory variable.

Our estimates of the effects of experience on earnings indicate that men's wage-experience profile was concave in both regimes and on average it did not change from the communist to the transition period. This finding differs from Chase (1998), Flanagan (1998) and to a lesser extent Krueger and Pischke (1995) who find wage-experience profiles becoming flatter during the new regime. When we estimate these profiles for workers in firms with different ownership types during the transition, we find that the *de novo* firms display a steeper and more concave profile than SOEs and public administration, hence paying a higher return to recent entrants' short experience than SOEs and public administration. We also find that private firms tend to pay higher wages than the SOEs and public administration, *ceteris paribus*.

Contrary to earlier conjectures, we find that the education and work experience gained during the transition have similar returns as the education and experience gained under communism. Unlike pre-transition studies that found the inter-industry wage structure to be similar in the market and centrally planned economies, we show that the inter-industry wage structure changed substantially as the transition unfolded between 1989 and 1996. In particular, men working in mining and quarrying lost much of their former wage premium, while those in trade, transport and telecommunications, and light manufacturing gained significantly.

Contrary to earlier studies that found the inter-industry wage structure to be stable and similar in the market and centrally planned economies, we show that men's inter-industry wage structure changed substantially as the economy switched from central planning to a nascent market system. The changes are in large part attributable to the *de novo* firms as they tend to pay a higher wage premium, irrespective of a worker's human capital, in trade, transport & telecommunications and other sectors of the economy.

Finally, we develop and apply a new methodology for decomposing the variance of worker-specific wages into components due to observable and unobservable determinants in communism vs. transition. We find the variance in wages due to unobserved effects dominates the variance due to observable determinants. Moreover, while over one-half of total variance is brought about by new unobservable characteristics introduced by the transition, there is considerable persistence of unobservable, individual-specific wage effects (e.g., skill premiums) from communism into the transition.

This chapter is organized as follows: In Section 2 we provide a brief institutional background, while in Section 3 we describe our data and methodology. Section 4 contains our empirical findings on returns to education under the communist grid and during the transition, while in Section 5 we present the corresponding returns to experience. In Section 6 we analyze the returns in transition to human capital obtained under communism. The shift in inter-industry wage differentials from the communist to the transition period is analyzed in Section 7, while in Section 8 we present and apply a new method for decomposing the variance of worker-specific wages. We conclude in Section 9.

2. The Wage Grids

As in other centrally planned economies, after the 1948 communist takeover of Czechoslovakia the government introduced the wage grid in an attempt to leave little discretion for managers or unions to set wages at the enterprise level. However, some discretion remained as managers could award “personal evaluation bonuses” that varied across workers with the same observable characteristics and could represent as much as 30% of the base wage. While in principle the trade unions and government jointly determined the grid and the level of wages within the grid, in practice the union and government officials by and large implemented the communist party policies as set out in the central plan.[6]

In Panel A of Table 1 we present the 1985 wage grid that was used for white collar workers in the last five years of communism. The columns represent wage levels by industry. Most workers were placed into wage tariff (class) categories I-Ib, while workers in heavy and construction industries were allocated into wage tariff categories II-Vb. Within each wage tariff category, workers were placed into salary classes 1-21 on the basis of their education, experience, occupation, and the number of employees that they supervised. The grid was accompanied by a detailed handbook that permits one to determine the relationship between education or experience and wages.

The system underlying the grid evolved over time. For example, the earlier grids (e.g., in the 1950s) were sector-specific, while the later ones were economy-wide.[7] As is evident from the 1985 grid in Table 1, planners favored workers in heavy industries and construction over those in other sectors.[8] Adjustments were also made for the number of hours worked per week and, as

Table 1
Wage Grids for the Communist and Post-Communist Period

A: 1985-1989 Wage Grid for White-Collar Workers in Czechoslovakia[a]

Salary Class	(A) All Industries except those in (B)				(B) Heavy Industry and Construction			
	I	**Ia**	**Ib**	**bonus**	**II**	**IIa**	**…**	**Vb**
1	1000	-	-	300	-	-	…	-
2	1100	-	-	300	-	-	…	-
3	1200	-	-	350	-	-	…	-
4	1300	-	-	400	-	-	…	-
5	1450	-	-	450	-	-	…	-
6	1600	1750	-	500	1700	1850	…	-
7	1750	1950	-	550	1850	2050	…	-
8	1950	2150	2350	600	2050	2250	…	3100
9	2150	2350	2600	650	2250	2450	…	3400
10	2350	2600	2850	700	2450	2700	…	3750
⋮	⋮	⋮	⋮	⋮	⋮	⋮	⋮	⋮
20	6300	-	-	1800	6500	-	…	-
21	7100	-	-	1900	7200	-	…	-

B: 1998 Wage Grid for the Public Sector in the Czech Republic[a]

Salary Class	Years of experience							
	< 1 yr.	**1-2**	**3-4**	**5-6**	**…**	**24-27**	**28-32**	**>32**
1	3,250	3390	3550	3700	⋮	4,660	4,820	4,980
2	3560	3720	3880	4050	⋮	5080	5250	5430
⋮	⋮	⋮	⋮	⋮	⋮	⋮	⋮	⋮
11	8800	9250	9710	10170	⋮	12910	13370	13840
12	10,000	10520	11030	11560	⋮	14710	15230	15760

[a]See text for description.

Sources:
Ministry of Labor and Social Affairs, (1985, 1986, 1998)

mentioned earlier, managers could at their discretion award workers significant bonuses.

The wage dispersion across the various categories in the grid was modest, given that unskilled workers were the pillar of the regime and the communist ideology dictated that wage differentials between the skilled and unskilled be kept small.[9] Correspondingly, during the communist period wages were compressed and income distribution in Czechoslovakia and the other Central and East European (CEE) countries was one of the most egalitarian in the world (see e.g., Atkinson and Micklewright, 1992).

Since the collapse of communism at the end of 1989, market forces have increasingly determined wages and employment in the *de novo* firms but the public sector and the privatized SOEs have continued to use a modified wage grid.[10] In Panel B of Table 1 we present the wage grid used in the public sector in 1998. In comparison to its communist predecessor, this grid was substantially simplified by eliminating the industry dimension and creating 12 experience-related categories (columns), together with 12 salary classes (rows) based primarily on education. The question that naturally arises is whether the rate of return on human capital under the transition grid matched or fell short of the market return provided by the new private firms.

3. Data and Methodology

3.1 Data

We use data from a retrospective questionnaire that was administered in December 1996 to 3,157 randomly selected households in all 76 districts of the Czech Republic.[11] For all working members of the household, we have information on the characteristics of the job held during the last year of socialism, in January 1989, and the current job held in December 1996.[12] The questionnaire first asks for the wage and other characteristics of the jobs held in

January 1989, the first month of the last year of the communist regime. Since the "big bang" of liberalization started January 1, 1991in Czechoslovakia, the questionnaire then traces the characteristics of all the jobs held by the surveyed individuals between January 1991 and December 1996. As a result, we have continuous labor market histories for each individual during the entire 1991-96 period. In particular, for each job we have the start wage and average hours of work, as well as the industry and ownership of the worker's firm. For the individuals employed in January 1991, we have also obtained information on wages and other characteristics at the start of the job held in January 1991. The starting dates of the jobs held in January 1991 span the entire 1948-89 communist period and we have used data from 1955 onward, while checking the robustness of our estimates by taking later starting points as well.[13] In particular, in order to test if our results are sensitive to the inclusion of observations from the 1950s, 1960s and 1970s, we have re-estimated our models with sub-samples that dropped observations on jobs that started before the1980s, 1970s and 1960s, respectively. As we report later, we found only negligible differences in the various estimates. Finally, for the 1991-96 period, we have collected information on each person's household and demographic characteristics, including changes in education.

The sample is representative of the 1996 population in terms of major demographic characteristics. It yields employment histories of 2,284 men who were employed for a minimum of two weeks during the period between January 1, 1991 and December 31, 1996. For the "mature" communist period of 1955-89, we use data on (a) the starting wages of 1,285 men who also held a job in January 1991 and (b) the cross section of wages of 1,955 men who were working during January 1989 (the first month of the last year of communism). For the transition period, we use cross section observations on wages and job characteristics of the 1,639 men who worked in December 1996, as well as the job start information on 2,107 men during the 1991-96 period. The data hence permit us to estimate (a) cross-sectional earnings functions using data from ongoing jobs at one point in time near the end of communism (January 1989) and one point in time in mature transition (December 1996), and (b) earnings functions using a long (1955-96) period of job start data under both regimes. The former estimates may be compared to Krueger and Pischke's (1995), Chase's (1998), Flanagan's (1998), and Vecerník (2001) cross-sectional estimates, while the latter ones provide a new longitudinal analysis during the communist and transition periods.

Different types of data sets have, by the nature of their design, different strengths and weaknesses. A potential weakness of retrospective data is the possibility of recall error. In our case, the potential problem is that individuals may not accurately remember their past wages. We expect this error to be relatively small, however, since wages set in the communist grid were clearly defined and did not change much through time. Moreover, the wages that we use from the relatively distant past are starting wages on the very last job held under communism, which we expect to be more readily recalled than wages during an arbitrary past job. With respect to wages during the transition period (1991-1996), we expect them to be remembered fairly accurately since there were few job changes: the average individual only held 1.6 jobs during this period.

Since we use the self-reported wage as a dependent variable rather than as a regressor, we avoid the usual problem of "errors in variables" with respect to the right hand side variables. Nevertheless, we check the magnitude of the recall error by performing two tests. First, we estimate the rate of return to education by using different starting points in the past and find the estimates to be invariant to whether we start in the 1950s, 1960s, 1970s, or 1980s. Second, we compare our basic estimates of rates of return to education with: (1)

Chase's (1998) estimates based on a 1984 and 1993 Czech household surveys, (2) Flanagan's (1998) and Vecerník (2001) estimates based on the 1988 Czech Microcensus and the 1996 Czech Survey of Economic Expectations and Attitudes, and (3) our estimates using a 1984 Czech firm-level survey. We find that these rates of return are similar to analogously calculated rates of return from our retrospective data.

Finally, there are two potential concerns related to the design of our retrospective data set. First, the sample is not fully representative of the communist era in that it is less likely to include individuals who were old men during the communist regime. In particular, we include in our sample those who were alive in 1996 and were not fully retired (i.e., worked at least two weeks) between 1991 and 1996. We hence exclude men who worked under communism and either fully retired before 1991 or died before 1996.[14] While this exclusion could be a problem if the individuals who retired/died had systematically different (e.g., lower) wages than others, there is no evidence that this was the case. Second, the communist era starting wage goes back further for individuals with long job tenure than for those with short job tenure. To the extent that these two types of individuals have systematically different unobserved characteristics that are correlated with some of the explanatory variables, the resulting time varying coefficients have a "duration bias." This concern is alleviated by our finding that parameter estimates are not affected in a material way by whether we make the starting point of the data be in the 1950s (when the sample is arguably the least representative of the population of starting wages), 1960s, 1970s, or 1980s (when the sample is the most representative of starting wages for the labor force in the 1980s).

In appendix Table A.1, we present the 1989 and 1996 means and standard deviations of the variables that we use in estimating the cross-sectional earnings functions. In appendix Table A.2, we report the corresponding information for the job start data during communism and the transition. As may be seen from the tables, the variables display sensible values and considerable variation both cross-sectionally and over time. Since manufacturing was the key part of the communist economy, over one-half of the men have apprenticeship education.

3.2 Estimation Strategy

In order to obtain estimates of the wage structure and returns to human capital at the end of communism (1989) and during the transition (1996), we first estimate the following augmented human capital earnings function with our 1989 and 1996 cross-sectional data:

$$\ln W_i = a_0 + a_1 E_i + a_2 X_i + a_3 X_i^2 + a_4 P_i + A_i' b + e_i \quad , \qquad (1)$$

where ln W_i, the natural logarithm of the monthly earnings of individual i, is taken to be a function of the individual's educational attainment (E_i), number of years of his potential labor market experience (X_i), a dummy variable for whether the individual worked in Prague (P_i), and a set of ten industry dummy variables for the industry location of the individual's job ($\boldsymbol{A_i}$).[15] The variables $\boldsymbol{A}$ and P control for industry wage effects, compensating differentials, and agglomeration effects of the central city. We have also estimated the traditional Mincer (1974) equation by omitting $\boldsymbol{A}$ and P from equation (1), but the coefficients on education and experience were virtually the same. In what follows we report estimates of equation (1).[16] We limit our analysis to workers with full-time jobs. In addition to examining all workers in 1989 and 1996, we estimate the regression separately for workers in three different ownership types: public administration and SOEs (henceforth "state"), privatized enterprises, and *de novo* firms.

An important stylized fact from the human capital literature is that the effect of education on wages often depends on how the education variable E is measured. Unlike Krueger and Pischke (1995), Chase (1998) and Flanagan

(1998), who have to impute *E* from the highest educational degree completed, we are able to use and test the relative merit of three different specifications of *E*: i) the actual self-reported number of years of education (net of grade repetition), ii) the highest level of attained schooling, and iii) a combination of i) and ii) above.[17]

The "number of years of education" specification yields an estimate of a constant marginal rate of return on an additional year of schooling and reflects the approach advocated by Layard and Psacharopoulos (1974). The "highest level of educational attainment" by type of degree obtained allows the rate of return to vary across types of completed education and reflects the criticism of the assumption of a constant rate of return to each year of education (Heckman, Layne-Farrar and Todd, 1996).[18] By including both of these variables, we are able to test between the competing specifications and see which one is better supported by the data in the communist and transitional contexts. Moreover, since we have data on actual years of schooling reported by the respondent,[19] rather than years imputed by the researchers from the reported school attainment, we can test the "sheepskin" hypothesis that "wages rise faster with extra years of education when the extra year also conveys a certificate" (Hungerford and Solon, 1987).[20]

As in most studies, our potential labor force experience variable *X* is calculated as the individual's age minus the sum of the individual's years of schooling and basic school enrollment age of six years.[21] In order to provide a good sense of the nature of the experience-earnings profile, we use two alternative specifications of experience: the traditional quadratic one and a spline function that fits the profile to three categories of years of experience.

Equation (1) enables us to compare cross-sectional estimates for late communism (1989) and mature transition (1996). For estimations covering the 1991-1996 period, we are able to include additional variables that capture important aspects of the transition and which are not relevant for the communist period. In particular, using our 1996 cross-section data, we estimate an equation that includes ownership dummy variables that capture whether the individual works in the state sector, privatized firm, or *de novo* firm. Finally, since we have data on wages at the start of jobs, we are also able to estimate continuous changes in the returns to human capital during the communist and transition periods. In order to capture these changes in a simple way, we extend equation (1) by estimating a time-varying-coefficient model by interacting the education (E) and experience (X and X^2) variables with an annual time trend τ, such that

$$a_k = a_k^t + t\ a_k^t \quad for\ k = 1-3 , \tag{2}$$

where subscripts k = 1-3 denote the coefficients on E, X, and X^2, respectively, and superscript t denotes the time invariant and superscript t the time varying portion of the coefficient. We stratify the data by the pre- and post-January 1991 periods and estimate separate equations for the communist and transition periods, allowing intercepts to vary across the regressions.[22]

It has become customary in the literature on earnings functions to correct for coefficient bias that may be brought about by the self-selection of a segment of non-representative individuals (usually women) into the labor market. Since labor force participation rates of both women and men declined after the fall of communism, we have tested for the presence of a selectivity bias in our sample but found it not to affect the coefficients of interest.[23]

4. Empirical Findings on Returns to Education

We divide our discussion of the returns to education into four parts: In Section 4.1 we present the returns to a year of education; in Section 4.2 the returns to an educational level; in Section 4.3 the returns from a model that encompasses both years and levels to test for sheepskin effects; and in Section 4.4 the returns to the field of study within each level of schooling. All estimates control for heteroskedasticity using the White (1980) method.

4.1 Returns to a Year of Education

In Table 2, we present our overall 1989 and 1996 cross-sectional estimates of the rates of return to a year of education based on equation (1).[24] For comparative purposes, we also report estimates from other studies in the Czech Republic and other selected countries. Our estimates suggest that in the last year of communism (1989), men's rate of return to a year of education was 2.7% and that it rose to 5.8% by 1996. The difference between the two coefficients is significant at 1% significance test level. Our findings are in line with the cross-sectional estimates of 2.4% for 1984 and 5.2% for 1993 obtained for the Czech Republic by Chase (1998), indicating that the return on education was low under the communist wage grid and that it rose substantially during the transition. Since both studies depict a lower starting level and a more pronounced increase in the return on education than the increase from 3.7% in 1988 to 4.5% in 1996 found by Flanagan (1998), we have gone back to Flanagan's data to re-estimate his equations and check for possible source of the discrepancy between his and our results. In replicating Flanagan's (1998) results we noticed two important facts. First, Flanagan's 1988 data set (*Microcensus* 1988) uses only data on heads of households. This may over-represent older and more able individuals, and hence account for the relatively higher rate of return on education reported by Flanagan for the communist period. Second, Flanagan's 1996 data set (the relatively small *Survey of Economic Expectations and Attitudes*) defines earnings as the sum of earned income and various social security benefits. Since the contribution of social security benefits to total income is more important for less educated workers,

Table 2
Estimated Returns to a Year of Education, Cross-Sectional Evidence for the Czech Republic and Other Countries

Country	Reference Years	Communism Men	Communism Men & Women	Transition Men	Transition Men & Women
CEE					
Czech Republic (1)	1989, 1996	0.027		0.058	
Czech Republic (2)	1984, 1993	0.024		0.052	
Czech Republic (3)	1989, 1996	0.037		0.045	
Czech Republic (4)	1988, 1996	0.040	0.044	0.083	0.088
East Germany (5)	1989, 1991		0.044		0.041
East Germany (6)	1988, 1991		0.077		0.062
Poland (7)	1987, 1992		0.050		0.070
Slovakia (2)	1984, 1993	0.028		0.049	
CIS					
Russia (8)	1991, 1994	0.031		0.067	
Latin America					
Argentina (9)	1989				0.103
Chile (9)	1989				0.120
Mexico (9)	1984				0.141
Venezuela (9)	1989				0.084
Europe					
West Germany (9)	1987				0.049
West Germany (6)	1988			0.075	0.077
Great Britain (9)	1984				0.068
Switzerland (9)	1987				0.079
United States (5)	1989			0.085	0.093

Note: Figures are reported coefficients from human capital (Mincer, 1976) earnings functions. All coefficients are statistically significant. CEE= Central and East Europe. CIS = Commonwealth of Independent States.

Sources:
(1)Authors' estimates, Table A.3
(2) Chase, 1998
(3) Flanagan, 1998
(4) Vecerník, 2001
(5) Bird et al., 1994
(6) Krueger and Pischke, 1995
(7) Rutkowski, 1997
(8) Brainerd, 1998
(9) Psacharopoulos, 1994

the construction of this dependent variable may explain the relatively low returns to education found in Flanagan's 1996 estimates.

The pattern of increased return on education is similar to that found by cross-sectional studies in other CEE countries, except for East Germany, in the early transition. As may be seen from Table 2, within a few years after the start of the transition, the rates of return on a year of education in CEE and Russia became similar to the rates in Western Europe, but not as high as the rates in the United States and Latin America.

Whereas this may be the first place where the rates of return to education for all of these transition countries are presented together, the stylized fact drawn from Table 2 is known. What is not yet known, however, is whether the rates of return to education vary with ownership. In the tables that follow, we report the rates of return by three important ownership categories: SOEs and public administration (State), privatized firms (Privatized) and private *de novo* firms (DeNovo). We are thus assess whether the new private entrepreneurs deviate from the communist era wage grid and reward human capital differently than their privatized and non-privatized SOE counterparts. This is an important question since post-communist adjustments in the wage grid, reductions in government subsidies to the state sector, and the opening up of the economy to international competition induced important changes in the pay policies of the SOEs and privatized firms as well. Whether the returns to human capital are higher in the *de novo*, privatized or public sector firms depends on the relative magnitudes of these effects.

In panels A and B of Table 3, we present estimated returns to a year of education using the cross-sectional and longitudinal data, respectively. In panel A, the 1996 cross-sectional estimates by ownership suggest that the privatized firms provide the highest rate of return to a year of education (6.5%), followed by the *de novo* firms (6.1%) and the state (5.6%).[25] However, these results --

Table 3
Estimated Returns to a Year of Education

	Reference	Communism	Transition			
	Years	**All**	**All**	**State**	**Privatized**	**DeNovo**
A: Cross-section data[a]	1989, 1996	0.0270 ***	0.0580 ***	0.0560 ***	0.0650 ***	0.0610 ***
		(0.004)	(0.005)	(0.009)	(0.007)	(0.010)
B: Time-Varying-coefficients[b]	1955-1989	-0.0004	0.0093 ***	0.0098 *	0.0104 ***	0.0077 ***
	(Annual change)	*(0.001)*	*(0.002)*	*(0.005)*	*(0.004)*	*(0.003)*
	Extrapolated to 1991	0.0170	0.0220 ***	0.0280 **	0.0270 **	0.0310 **
		(0.010)	*(0.007)*	*(0.012)*	*(0.012)*	*(0.012)*

[a]Taken from Table A.3.

[b]Taken from Table A.6. Based on job-starts

*, **, *** Statistically signigficant at the 10%, 5%, 1% level. Standard errors in parentheses.

based on 384 observations for state enterprises, 504 for privatized firms and 604 for *de novo* firms -- are not statistically different from one another, indicating no systematic difference in the education-based wage differentials across principal ownership forms.[26]

In panel B of Table 3, the time-varying-coefficients are presented as the 1991 base and the annual rate of change. The coefficient on the annual change (interaction term) is miniscule and insignificant during the communist period, indicating that under the communist grid the rate of return to a year of schooling remained constant over time at a mere 1.7%. Moreover, a test of the difference between the point estimates from the longitudinal (1955-89) and cross-sectional (1989) data indicates that there was no statistically significant difference. In order to check if our estimates are sensitive to the starting date, we have also estimated the time-varying-coefficients model with observations going back to the 1960s, 1970s and 1980s, respectively. We find that all three estimated coefficients on the interaction terms are insignificant and the base coefficients on education are in the 0.15 to 0.21 range and within one standard error of each other. Our results hence suggest that wage differentials based on education were low and stagnant under the decades of central planning, a finding that has not been documented before with micro data.

In contrast, our time-varying-coefficient estimates for 1991-96 show that the estimated rate of return to a year of education increased by almost 1% a year during the transition. While privatized firms recorded the fastest rate of annual increase (1.04%), followed by the state sector (0.98%) and *de novo* firms (0.77%), the differences across ownership categories are not statistically significant. This finding hence complements the cross-sectional estimates by showing that the rate of return rose steadily during the transition period and that on average firms with different ownership remained competitive in terms of education-based wage differentials.

4.2 Estimates Based on Attained Levels of Education

In panel A of Table 4, we report 1989 and 1996 cross-sectional estimates for several different levels of schooling, relative to the mandatory junior high school. (The full set of parameters is presented in Table A.4.) We use these estimates to calculate the annual returns to a year of education within each completed category of schooling (panel B).[27] The time-varying coefficients are presented in Table 5 and the full set of parameters is reported in Table A.7.

As may be seen from the first column of Table 4, at the end of the communist regime the earnings differentials between different types of schooling were small. For example, a university educated man earned just about 28% more than an otherwise identical man with a junior high school education. Similarly, men with a vocational high school degree earned 13% more than their counterparts with a junior high school education. Finally, the earnings of individuals with a two-year apprenticeship and junior high school were about the same.

By 1996 the returns to higher levels of education increased dramatically (column 2 of Table 4). University educated man earned 72% more (coefficient of .544) than his counterpart with junior high school education.[28] The difference between the 1989 and 1996 coefficients on university education is significantly different at the 0.01 confidence level. We also find that the difference between 1996 and 1989 in the returns to a vocational high school education is highly significant and that the percentage increase in this return is the largest among all the education levels. On the other hand, the return to an apprenticeship did not change significantly over time.

Examining the 1996 returns in Table 4 by firm ownership, one observes that privatized firms are the only ones valuing apprenticeship over junior high school education and that academic high school education is significantly valued only in the state sector. However, all firm types pay more to individuals

Table 4

Estimated Returns by Level of Educational Attainment, Cross Section Data[a]

	Communism (1989)	Transition (1996)			
	All	**All**	**State**	**Privatized**	**DeNovo**
A. Level of attainment					
-apprentices (2 years)	0.063	0.094	0.129	0.114 *	0.101
	(0.051)	*(0.057)*	*(0.121)*	*(0.065)*	*(0.137)*
-apprentices (3 years)	0.077 **	0.112 **	0.097	0.156 ***	0.065
	(0.037)	*(0.049)*	*(0.105)*	*(0.058)*	*(0.115)*
-vocational H.S. (4 years)	0.127 ***	0.294 ***	0.323 ***	0.327 ***	0.249 **
	(0.040)	*(0.050)*	*(0.105)*	*(0.058)*	*(0.118)*
-academic H.S. (4 years)	0.135 *	0.351 ***	0.401 ***	0.266	0.342
	(0.081)	*(0.107)*	*(0.142)*	*(0.164)*	*(0.309)*
-university	0.283 ***	0.544 ***	0.476 ***	0.673 ***	0.599 ***
	(0.045)	*(0.059)*	*(0.115)*	*(0.072)*	*(0.133)*
B: Calculated annual returns within attainment level[b]					
-apprentices (2 years)	0.032	0.048	0.067	0.059	0.052
-apprentices (3 years)	0.026	0.038	0.033	0.053	0.022
-vocational H.S. (4 years)	0.032	0.076	0.084	0.085	0.064
-academic H.S. (4 years)	0.034	0.092	0.105	0.069	0.089
-university	0.044	0.076	0.040	0.127	0.102

[a] Taken fom Table A.4, education in levels.

[b] Using the estimated coefficients ß on attainment in panel A and the years of education, annual returns are computed as exp(ß)-1.

*, **, *** Statistically significant at the 10%, 5%, 1% level. Standard errors in parentheses.

with vocational or university degrees. The estimated coefficient on university education is highest in privatized firms (0.673), followed by *de novo* firms (0.599) and state enterprises and public administration (0.476). The difference between the university coefficients for privatized firms and state enterprises approaches statistical significance (p-value of 0.14), but in all other pair-wise comparisons across ownership categories, one cannot reject the hypothesis of equality of returns. Our estimates hence indicate that firms with different ownership display tendencies to remunerate different types of human capital differently but, as in the case of returns to a year of schooling, these differences are not statistically significant.

As may be seen from Panel B of Table 4, in late communism the calculated return to a year of education was almost the same in all levels of schooling, except possibly the university. Yet, by 1996 the return to a year of academic or vocational high school education rose above the return to a year of apprenticeship, thus providing support for the hypothesis of uneven returns across educational categories. The estimates by ownership appear to amplify this finding.

When we estimate the time-varying-coefficient model on 1955-90 data, we find no change in the returns to educational attainment over time (Table 5). The small differences in returns among the various levels of education are also analogous to those based on the 1989 cross section data.[29] The 1991-96 estimates for all workers indicate that during the transition the rate of return on education rose significantly in all categories except for academic high school. The ownership-specific, time-varying estimates complement the cross-sectional estimates in Table 4 by showing that the increase in the rate of return on two-year apprenticeship has been driven by privatized firms. Moreover, while privatized and *de novo* firms provided a significant rate of return on vocational training already in 1991, the state sector registers a faster rate of increase in this

Table 5
Estimated Returns by Level of Educational Attainment[a]
(Time-Varying-Coefficients Model)

	Communism (1955-1989)	Transition (1991-1996)			
	All	**All**	**State**	**Privatized**	**DeNovo**
Apprentice (2 years)	0.057	-0.078	0.153	0.154	-0.066
	(0.101)	*(0.106)*	*(0.167)*	*(0.156)*	*(0.164)*
Apprentice (2 years)·t	n.a.	0.079 **	0.024	0.061 *	0.062
	n.a.	*(0.031)*	*(0.051)*	*(0.034)*	*(0.062)*
Apprentice (3 years)	0.069	0.049	0.095	0.118	0.087
	(0.075)	*(0.069)*	*(0.112)*	*(0.103)*	*(0.078)*
Apprentice (3 years)·t	0.000	0.053 **	0.065 ***	0.042 *	0.032 **
	(0.005)	*(0.021)*	*(0.022)*	*(0.022)*	*(0.015)*
Vocational H.S. (4 years)	0.056	0.051	0.059	0.203 *	0.183 **
	(0.082)	*(0.074)*	*(0.124)*	*(0.117)*	*(0.091)*
Vocational H.S.(4 years)·t	-0.001	0.077 ***	0.102 ***	0.047 **	0.032 *
	(0.006)	*(0.022)*	*(0.032)*	*(0.022)*	*(0.019)*
Academic H.S.(4 years)	0.338 *	0.090	0.299	0.059	0.013
	(0.178)	*(0.113)*	*(0.186)*	*(0.240)*	*(0.186)*
Academic H.S.(4 years)·t	0.010	0.033	0.037	0.104 *	0.032
	(0.011)	*(0.034)*	*(0.056)*	*(0.056)*	*(0.053)*
University	0.179 **	0.268 ***	0.330 **	0.405 ***	0.316 ***
	(0.089)	*(0.082)*	*(0.133)*	*(0.127)*	*(0.112)*
University·t	-0.005	0.100 ***	0.117 ***	0.076 ***	0.099 ***
	(0.007)	*(0.025)*	*(0.041)*	*(0.026)*	*(0.025)*

[a]Taken from Table A.6.
*, **, *** Statistically significant at the 10%, 5%, 1% level. Standard errors in parentheses.

return during the 1991-96 period, especially when compared to *de novo* firms. Finally, the rate of growth of returns for university was also growing most rapidly among the state sector.

Overall, our findings from Tables 4 and 5 indicate that education-related wage differentials were small and stagnant under communism. Market forces have increased wages for those with vocational high school and university education, but the gains were nil for those with lower education. The results based on firm ownership indicate that university education appears to be valued by all firm types, but most by the privatized firm and least by the state enterprises.

4.3 Regressions with Years and Levels of Education

Screening theories of education suggest that diplomas serve as a signal of higher productivity and one should therefore expect diplomas to be rewarded in the labor market. Various studies using US data test for sheepskin effects by estimating the difference in wages of individuals with and without a diploma, conditional on years of schooling (see e.g., Hungerford and Solon, 1987, Card and Krueger, 1992 and Jaeger and Page, 1996.) Except for the Jaeger and Page (1996) study, however, the US estimates are based on data that do not have information on the highest degree attained by an individual and therefore have to impute the level attained from the "usual number of years" it takes to complete a given degree. In contrast, researchers of transition economies usually have only information on highest degree attained and must impute the number of years of schooling of an individual by using the usual number of years it takes to complete a degree. To the extent that individuals obtain a diploma with more or fewer years of study, estimates of sheepskin effects in the US and returns to a year of education in the transition countries are biased. We are fortunate to have information on both the individual's reported years of education (net of any repeated grades) and the highest degree attained. We can thus obtain unbiased estimates of the sheepskin effect and also test for the bias using imputed vs. actual years of schooling. We also show a new way of testing for the sheepskin effect by estimating returns to years of study that lead to a degree and those that do not.

In Table 6 we present the coefficients for a specification that includes years of education (net of grade repetition) and dummy variables for highest degree attained, estimated from the 1989 and 1996 cross sectional data and controlling for the variables listed in Equation 1. In both years, we find sheepskin effects for higher levels of education -- vocational high school and university degrees in 1989 and these two degrees plus the academic high school diploma in 1996. We also find an overall effect associated with completing degrees in that we reject (at 1% in 1996 and 11% in 1989) the hypothesis that the coefficients on the five educational levels are jointly zero. The estimated coefficients on higher education also become greater over time but F tests on pair-wise differences of the coefficients between 1989 and 1996 do not find any of them to be statistically significant. Examining the sheepskin effect by firm ownership during the transition, we find that the privatized and *de novo* firms place more importance on diplomas than the state sector and that the state sector is the only owner that values years of education.[30]

Since many other studies, including Krueger and Pischke (1995), Flanagan (1998) and Chase (1998), had to impute the information on years of education from data on attainment, we have taken advantage of the dual reporting in our data and re-estimated our regressions with the imputed years of education in order to assess the magnitude of the errors-in-variables bias of this indirect, but commonly used, measure. Normally, the imputed years of education would generate a downward bias that is associated with errors-in-variables. However, in our case the imputed years of education may generate an upward bias because the measure underestimates the number of years of

Table 6
Sheep Skin Effects

	Communism (1989)	Transition (1996)				
	All	**All**	**Difference**	**State**	**Privatized**	**DeNovo**
Years of Education	0.006	0.020 **	0.014	0.042 ***	0.002	0.000
	(0.007)	*(0.009)*	*(0.239)*	*(0.016)*	*(0.013)*	*(0.018)*
Apprentices (1-2 years)	0.052	0.058	0.006	0.057	0.110	0.101
	(0.054)	*(0.061)*	*(0.942)*	*(0.119)*	*(0.071)*	*(0.142)*
Apprentices (3-4 years)	0.060	0.056	-0.004	0.001	0.150 **	0.066
	(0.043)	*(0.055)*	*(0.954)*	*(0.106)*	*(0.067)*	*(0.129)*
Vocational H.S. (4 years)	0.100 *	0.209 ***	0.109	0.176	0.318 ***	0.250 *
	(0.052)	*(0.062)*	*(0.178)*	*(0.113)*	*(0.077)*	*(0.143)*
Academic H.S. (4 years)	0.108	0.271 **	0.164	0.247 *	0.257	0.343
	(0.088)	*(0.112)*	*(0.251)*	*(0.137)*	*(0.170)*	*(0.316)*
University	0.229 ***	0.367 ***	0.137	0.144	0.655 ***	0.601 ***
	(0.078)	*(0.093)*	*(0.258)*	*(0.155)*	*(0.132)*	*(0.199)*

Note:
The regressions also include control dummies for child benefits, taxes and nine industries.
*, **, *** Statistically significant at the 10%, 5%, 1% level. Standard errors in parentheses.

schooling for people that study for additional years without obtaining a degree. Indeed, the coefficient on imputed years of schooling (the first row of Table 7) is higher than the coefficient on actual years attended (Table 3) for both communism (0.033 vs. 0.027) and transition (0.066 vs. 0.058). The associated standard errors are sufficiently large, however, not to permit us to reject the hypothesis that in both periods the coefficients on imputed and reported year of education are not statistically different from one another.[31] The downward and upward biases hence just about cancel one another out.

We also test for the sheepskin effect using data on years of education and attained degree. In particular, we test the hypothesis that years of education that lead to a degree have a higher payoff than those that do not result in a degree. To implement the test, we use our information on the total number of reported years of education and the highest degree obtained, together with the knowledge of the usual number of years needed to obtain a given degree. Using this information, we impute the number of years of education used for (a) obtaining the most advanced degree and (b) additional study not resulting in a degree. In Columns 2 and 4 of Table 7, we show the coefficients from a specification that enters these two measures as explanatory variables in the standard regression of Equation (1). In both 1989 and 1996 the coefficients on the additional years of study are significantly different from zero but smaller than the coefficients on the imputed years leading to a degree. The F tests indicate that the difference in the coefficients on imputed vs. additional year is significantly different from zero in 1996 but not in 1989.[32]

Overall, our results point to the presence of a sheepskin effect and the effect is more pronounced at higher educational levels and during the transition than under communism. They also caution that studies that impute years of education from educational attainment and do not control for the drop-out or repeater phenomena overestimate the rate of return on education.[33]

Table 7
Estimated Returns for "Imputed Years"and "Additional Years"of Schooling[a]

	Communism (1989)		Transition (1996)	
Imputed Years of Education	0.033 ***	0.032 ***	0.066 ***	0.065 ***
	(0.004)	*(0.004)*	*(0.006)*	*(0.006)*
An Additional Year of Schooling Above any Degree[a]	-	0.020 **	-	0.034 ***
		(0.008)		*(0.012)*

Note:

[a] "Imputed Years" denote the number of years of education imputed from the individual's highest level of educational attainment and the usual number of years it takes to attain that level/degree. "Additional Years" denote the number of years above the highest level of attainment and which do not lead to a degree. All the regressions also include control dummies for child benefits, taxes and nine industries.

*, **, *** Statistically significant at the 10%, 5%, 1% level. Standard errors in parentheses.

4.4 Returns to a Field of Study

Our data also permit us to estimate the returns to a field of study for a given level of education and assess whether there was a shift in these returns from the communist to the market system. As we show in Table A.5, there is no statistically significant change in the returns to the different fields of study from 1989 to 1996 for men who only attained an apprentice education. For men whose highest level of education was vocational high school, most of the coefficients on the fields of study rose between 15 and 25 percentage points from 1989 to 1996. Men trained in business and trade services gained relatively more over this period, as did men in manufacturing and the electro-technical area. Those trained in law, teaching and "other social branches" saw no change in their returns. For the university educated men all the coefficients basically doubled in size between 1989 and 1996. The high outlier is law where returns rose by a factor of almost three. On the low end, the returns of those trained in health, teaching and "other social branches," financed from the state budget, did not change over time. Our data hence reveal important shifts in the returns to fields of study. As expected, education in business and trade services has become more highly rewarded. Similarly, the higher rate of return for university educated lawyers is consistent with the increase in demand for legal services during the process of privatization and increased reliance on legal institutions.

5. Returns to Experience

We explore the returns to experience in the two regimes with our cross-section and longitudinal data and with the information provided by the wage grids. In Table 8, we present the coefficients and standard errors of the experience and experience squared terms estimated with the 1989 and 1996 cross-sectional data.[34] Focusing on the first two columns for "all workers," we find the coefficients are statistically significant, and within the standard range.

Table 8
Cross-Sectional Estimates of Returns to a Year of Labor Market Experience[a]

	Communism (1989)	Transition (1996)			
	All	**All**	**State**	**Privatized**	**DeNovo**
Experience	0.021 ***	0.021 ***	0.015 **	0.022 ***	0.030 ***
	(0.003)	*(0.005)*	*(0.006)*	*(0.007)*	*(0.004)*
Experience^2	-0.0004 ***	-0.0004 ***	-0.0003 ***	-0.0004 **	-0.0007 ***
	(0.0001)	*(0.0001)*	*(0.0001)*	*(0.0002)*	*(0.0001)*

[a] Taken from Table A.3, years of education.

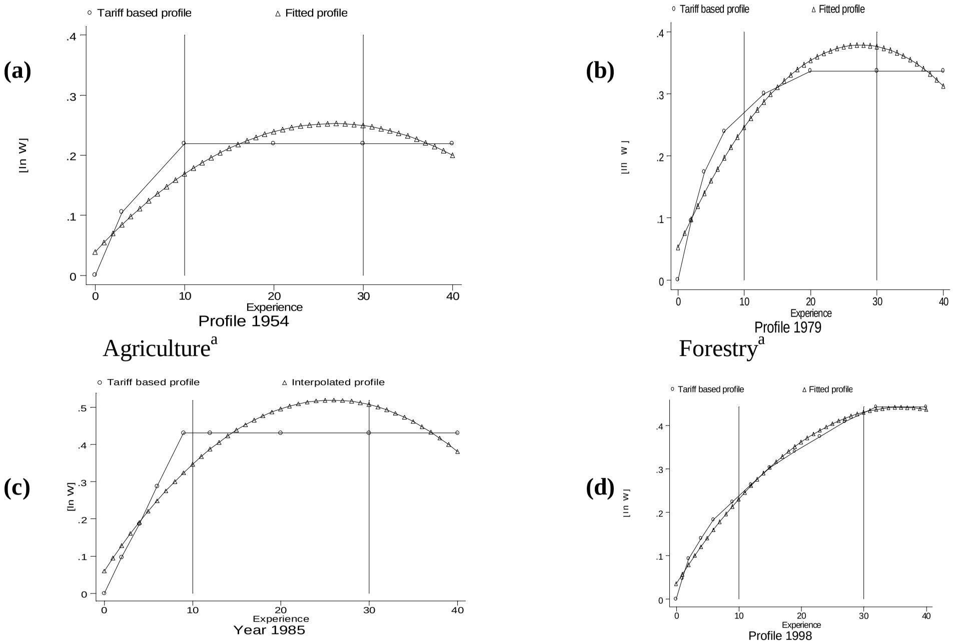
*, **, *** Statistically significant at the 10%, 5%, 1% level. Standard errors in parentheses.

We test the differences in these coefficients from 1989 to 1996 and find that the experience-earnings profile did not change from communism to the transition, peaking around 26 years in both years.

On the other hand, our estimates by ownership categories (columns 3-5 of Table 8) show marked differences in experience-based wage setting across the three types of ownership. The wage experience profile is flattest in the state sector, more concave in privatized firms and most concave in the *de novo* firms. The coefficients on the experience terms for the *de novo* firms are statistically different from those for the state and close to being statistically different from those of privatized firms in both specifications. Men's wage-experience profiles begin steeper in *de novo* firms than in the state sector, but they are also more concave and have an earlier turning point. *De novo* firms hence pay higher returns on a year of experience to employees with low experience (recent entrants into the labor market) and lower returns to men near retirement age.[35]

The similarity of the estimated wage-experience profile under communism, in the transition and in market economies has led us to collect data on wage grids in a number of periods of the communist regime, as well as transition, and analyze them more systematically. The search was surprisingly laborious, but we were able to obtain various wage grids, from 1954 to 1998. The wage-experience profiles given by these grids are presented for 1954, 1979, 1985-1989, and 1998 in Figure 1 (a)-(d) as the "grid based profiles." We note that we could not find grids pertaining to the same reference group over time, and the grids hence should not be compared longitudinally. For example, the 1954 grid is for agricultural worker, the 1985-89 grid is for white collar workers and the 1998 grid is for all workers in SOEs, public administration and privatized SOEs. As we noted in Section 1, the method for structuring the grid also changed over time; for example, in some years it had an industrial component and in some years it did not. Nevertheless, the grids permit us to

Figure 1
Wage Experience Profiles from the 1954, 1979, 1985 and 1998 Wage Grids
(Actual Grid Data Points and Curve Fitted with a Quadratic Wage-Experience Function)



Agriculture[a] Forestry[a]

White Collar Workers[b] All Workers in Public Admin. & SOEs[b]

Sources: [a]Ministry of Agriculture (1952, 1979), [b]Ministry of Labor and Social Affairs (1985, 1998) discern that in all years the wage-experience profiles are piece-wise linearly concave and have a flat region in the latter part of the profile. Hence, while ideology led the planners to impose narrow education-related wage differentials and cap the experience-earnings profile, they built into the grid enough wage progression in the early years of experience to generate a Mincerian-type concave profile.

Given the nature of all the grids, we fit the quadratic Mincerian earnings-experience function to the data of the five grids dating from 1954 to 1998. These coefficients are reported in Table 9 and also plotted in Figure 1(a)-(d). We see in Figure 1 that the quadratic function fits the wage grids fairly well and better in some years (e.g. 1998) than others (e.g., 1985). The goodness of fit is particularly high in the 1998 grid because of its fine gradation of earnings with seniority. The plots and the coefficients also show that the slope and concavity of the wage experience profile in agriculture was fairly flat whereas it was much steeper for all workers in 1998. We note that the coefficients in Table 9 for the 1998 wage grid are very similar to the coefficients from our data for all workers in 1996. It hence appears that the slope of the experience-earnings profile in the grid became steeper over time, but since the grids in the earlier years apply to different types of narrowly defined workers, we cannot formally draw this conclusion. Rather, we turn to our own data to test whether the experience-earnings profile changed over the communist period.

The time-varying estimates of the coefficients on experience (Table 10) permit us to provide the first assessment of the extent to which the concavity of the experience-earnings profiles change over time within the communist and transition periods. Although the coefficients on experience interacted with time are all positive and those on experience squared interacted with time are all negative, suggesting that the profile is becoming steeper and more concave over time, only the coefficient on experience interacted with time for the

Table 9

Parameters from Fitting the Wage-Grid with a Quadratic Wage-Experience Function.

	1954[a] (Agricul.)	1973[b] (Industry)	1979[c] (Forestry)	1982[d] (Manual Workers)	1985-1989[e] (White Collar)	1998[f] (SOEs/Pub Admin.)
Experience	0.016	0.017	0.024	0.032	0.039	0.023
Experience^2	-0.0030	-0.0004	-0.0004	-0.0006	-0.0006	-0.0003

Sources:

[a]Ministry of Agriculture (1952)

[b]Ministry of Industry (1973)

[c]Ministry of Agriculture (1979)

[d]Ministry of Defense (1982)

[e]Ministry of Labor and Social Affairs (1985, 1986)

[f]Ministry of Labor and Social Affairs (1998)

communist period in specification based on education levels (panel B) is marginally statistically significant (at 10% test level). In this latter specification, an F test on the joint significance of experience interacted with time and experience squared interacted with time also indicates that at 5% significance test level one cannot reject the hypothesis that the slope of the profile was changing during the 1955-89 period. In contrast, joint F tests performed on the overall estimates in panel A for 1955-90 and panels A and B for 1991-96 suggest that the profile was not changing significantly over time. Moreover, tests of equality of experience-related coefficients between the 1955-90 and 1991-96 periods indicate that one cannot reject the hypothesis of equality of the evolution of the experience profile during the two periods.[36] We hence conclude that the experience-earnings profile for all workers under communism approximated the Mincerian human capital earnings function; there is weak evidence that the profile was altered during communism but its evolution was not altered during the first six years of the transition.

The time-varying estimates based on firm ownership (columns 3-5 of Table 10) confirm that during 1991-96 the wage-experience profile is concave in all three types of ownership categories and that it does not change significantly over time. However, unlike in Table 8 (using cross-section data), the ownership-specific estimates in Table 10 suggest that the return to experience is highest in the state, followed by the *de novo* and privatized firms. The difference between the cross-sectional and the longitudinal estimates based on starting wages is brought about almost entirely by a change in the coefficients of the state sector. Unlike in *de novo* and privatized firms, new jobs in the state sector have a steeper and more concave profile than existing (cross-sectional) state jobs. The profile of the new state jobs also peaks earlier (23 years) than that of existing jobs (26 years). The asymmetry in compensating

Table 10
Time-Varying Estimats of Returns to a Year of Labor Market Experience

	Communism (1955-1989)	Transition (1991-1996)			
	All	All	State	Privatized	De Novo
A. Education in years[a]					
Experience	0.0236***	0.0285***	0.0349***	0.0256***	0.0283***
	(0.005)	*(0.005)*	*(0.007)*	*(0.009)*	*(0.006)*
Experience·t	0.0007	0.0017	0.0018	0.0012	0.0012
	(0.0005)	*(0.0014)*	*(0.003)*	*(0.003)*	*(0.002)*
Experience^2	-0.0005***	-0.0006***	-0.0008***	-0.0006**	-0.0007***
	(0.0002)	*(0.0001)*	*(0.0001)*	*(0.0002)*	*(0.0001)*
Experience^2·t	-0.000004	-0.00004	0.00003	-0.00002	0.00000
	(0.00002)	*(0.00004)*	*(0.00006)*	*(0.00007)*	*(0.00004)*
B. Education Levels[b]					
Experience	0.0244***	0.0291***	0.0411***	0.0252***	0.0303
	(0.006)	*(0.005)*	*(0.006)*	*(0.008)*	*(0.006)*
Experience·t	0.0009*	0.0022	0.0026	0.0033	0.0016***
	(0.0005)	*(0.0015)*	*(0.003)*	*(0.003)*	*(0.002)*
Experience^2	-0.0005***	-0.0006***	-0.0009***	-0.0006***	-0.0007
	(0.0002)	*(0.0001)*	*(0.0001)*	*(0.0002)*	*(0.0001)*
Experience^2·t	-0.00001	-0.00005	0.00006	-0.00006	-0.00002***
	(0.00002)	*(0.00004)*	*(0.00006)*	*(0.00007)*	*(0.00005)*

[a] Taken from Table A.6. [b] Taken from Table A.7.
*, **, *** Statistically significant at the 10%, 5%, 1% level. Standard errors in parentheses.

new versus existing workers in the state sector during the transition should be explored in future research.

Our overall results for the transition period are similar to those of Flanagan (1998) and Vecerník (2001), but they differ from those of Chase (1998) and Krueger and Pischke (1995), who find a much flatter wage experience profile.[37] Our estimates by ownership categories and an examination of the wage grid over time provide a possible explanation of this discrepancy. As may be seen from Table 8, the wage experience profile is flatter in the state and privatized enterprises than in the *de novo* firms. Since Chase's (1998) and Krueger and Pischke's (1995) estimates relate to an earlier phase of the transition (1991 and 1993, respectively) when *de novo* firms were less prominent, the difference in the estimated wage experience profiles is likely to come from the different composition of firm ownership in these studies.[38]

6. Returns to Communist Human Capital in the Transition

Earlier studies have hypothesized that human capital acquired under communism is less appropriate for a market economy and it should hence receive a lower rate of return during the transition period than post-communist human capital. To test this hypothesis, we have calculated for each man the number of years of education and experience obtained under communism vs. transition and tested for differences in the returns during the transition to the two types of education and experience, respectively.

Since 14 percent of the men in our 1996 sample concluded their education during the 1990-1996 period, we have a sufficiently large sample to test three specifications. We begin by entering for each man as separate regressors his number of years of communist (old) education and his number of years of post-communist (new) education. The estimated coefficients are 0.058 and 0.031 for old and new, respectively, and at the 5% test level they are significantly different from zero and from each other.[39] The finding that post-communist education has a lower return than education obtained under communism strongly contradicts the aforementioned hypothesis. We have checked whether the result is arising because a large proportion of school leavers in 1990-96 have lower levels of education (junior high school and apprentices) that we know command relatively low returns during transition, but we find that this is not the case. We have therefore also tested for differences in returns to higher levels of education (vocational, academic high school and university) obtained during vs. after communism and find that the respective coefficients of 0.055 and 0.058 are significantly different from zero but not significantly different from each other. This result indicates that at the secondary and tertiary levels communist and transition education is indistinguishable in terms of labor market return. Finally, we have estimated a model that allows the education coefficients to be different for younger (less than 30 years) and older men, proxying for two vintages of human capital that correlate with the communist and transition periods. The resulting estimates do not allow us to reject the hypothesis that the education of the younger and older men generates the same rate of return.[40] Overall, the results of the three estimations contradict the hypothesis that education acquired under communism is less appropriate for a market economy than education obtained in transition. Rather, the findings are consistent with two other hypotheses: (i) education obtained under communism is (at least) as appropriate for a market economy as education obtained in transition and (ii) reforms of the educational system have proceeded slowly during the transition.

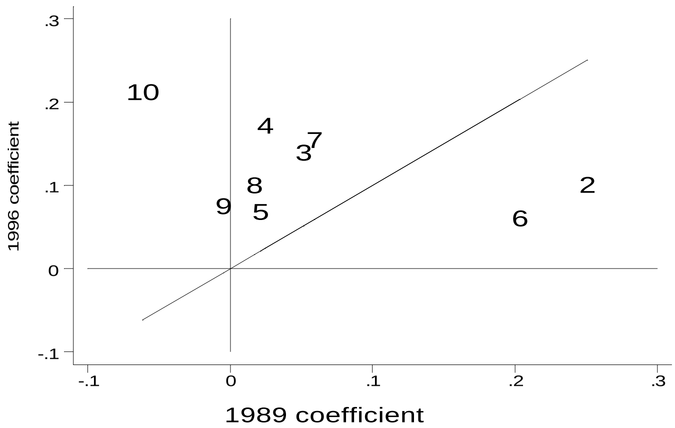
We have also tested the hypothesis that experience obtained after 1989 generates higher rates of return in the transitional than experience accumulated under communism. This is a conjecture made by policy makers and several authors, including Krueger and Pischke (1995) and Flanagan (1998). However, the cross-sectional data used in previous studies did not permit a direct test of this hypothesis because they do not have sufficient variation in the values of the post-communist experience variable. We can carry out the test on the 1991-96 job start data and we find that individually and jointly the coefficients on the pre- and post-communist experience and experience squared are not different from one another.[41] Our direct test hence suggests that the communist and transition experience command the same rate of return during the transition.

7. Shifts in Industry Wage Premiums Between 1989 and 1996

The literature on inter-industry wage differentials has found that these differentials are persistent and that the ranking of industries by their average wages was similar in market and planned economies (see e.g., Krueger and Summers, 1987 and Rutkowski, 1994). These findings were found to hold irrespective of whether one controlled for other factors and they pointed to a similar industry wage structure in western and communist economies. This is surprising given the differences in the institutions and degree of market forces governing wages in each economy.

In order to generate findings that are comparable to the existing literature, we analyze industry intercepts from the 1989 and 1996 regressions in which we control for years of education and experience. These intercepts are industry wage differentials relative to agriculture, holding constant the composition of workers' human capital characteristics.[42] Analogous to the approach adopted by Krueger and Summers (1987), we plot the industry intercepts (coefficients) for 1989 against those for 1996.[43] As seen in Figure 2, major changes have taken place in the structure of inter-sectoral wage premiums. Rather than fitting along the positively sloped 45 degree line, the coefficients fit more closely to a downward sloping line.[44] Between 1989 and 1996, relative wages in finance and mining and quarrying have decreased,

Figure 2
Scatter-plot of Estimated Coefficients on Industry-Specific Dummy Variables (1989 vs. 1996)



Legend

1	Agriculture=base (excluded)
2	Mining, Quarrying, Energy Production and Distribution
3	Construction
4	Wholesale, Retail, and Private Services
5	Public Administration, Education, Health, and Army
6	Finance, Insurance, Real Estate
7	Transportation and Telecommunication
8	Manufacturing-machinery
9	Manufacturing-Other
10	Other

while those in trade, transport and telecommunications, light manufacturing, and "other" activities gained. The long-term stability of the inter-industry wage differentials in these countries, documented in the earlier literature, has therefore been disrupted by the transition.

In order to verify the scatter diagram analysis in Figure 2, we report in Table 11 the industry intercepts and tests for the significance of their 1989 vs.1996 differences. An examination of the intercept coefficients indicates that while only miners enjoyed a significant positive wage premium relative to those in agriculture under communism in 1989, by 1996 seven of the nine sectors paid a premium. In analyzing pair-wise 1989-96 differences in the intercepts, we find that five are statistically significant. Men working in mining and quarrying indeed lost much of their former wage premium, with the decline occurring primarily in the privatized firms. Those in trade, transport and telecommunications, light manufacturing, and "other" activities gained significantly, with most of the gain brought about by higher wage premiums in the *de novo* firms and in the case of transport and telecoms also the privatized firm. However, the seemingly large decline in finance, insurance and real estate's wage differentials turns out not to be statistically significant. The interesting question is why we do not find a growing difference in intercepts in this expanding sector that has been hiring employees at very high wages? Our analysis indicates that the high wages of the employees in the finance sector reflect their relatively high levels of human capital and their concentrated location in the high premium city of Prague. Finally, a more detailed analysis of the differentials in Table 11 indicates that agriculture, the base sector whose share in total output and employment shrank dramatically, lost also in terms of its wage differential relative to the rest of the economy.

Table 11

Changes in Industry Wage Structure from 1989 to 1996[a]

	Communism (1989)	Transition (1996)				
	All	**All**	**Difference[b]**	**State**	**Privatized**	**DeNovo**
Mining & Quarrying	0.251 ***	0.092 **	-0.159 ***	0.245 **	0.063	-0.079
	(0.039)	*(0.044)*	*(0.007)*	*(0.099)*	*(0.058)*	*(0.159)*
Construction	0.051	0.131 ***	0.080	0.110	0.082	0.119
	(0.035)	*(0.040)*	*(0.134)*	*(0.120)*	*(0.058)*	*(0.091)*
Wholesale and Retail Trade	0.025	0.163 ***	0.139 **	-0.134	0.060	0.147 *
	(0.037)	*(0.041)*	*(0.012)*	*(0.138)*	*(0.062)*	*(0.087)*
Public Admin., Education, Health	0.021	0.059	0.038	0.080	-0.190	0.0850
	(0.035)	*(0.115)*	*(0.389)*	*(0.090)*	*(0.219)*	*(0.090)*
Finance, Insur. & Real Estate	0.203	0.052	-0.152	0.140	0.054	-0.017
	(0.139)	*(0.080)*	*(0.345)*	*(0.171)*	*(0.116)*	*(0.170)*
Transport & Telecom.	0.059	0.146 ***	0.087 *	0.096	0.122 *	0.275 ***
	(0.036)	*(0.040)*	*(0.100)*	*(0.095)*	*(0.062)*	*(0.095)*
Manufacturing-Food, Textile,	0.017	0.092 ***	0.075 *	0.045	0.063	0.118
	(0.028)	*(0.033)*	*(0.088)*	*(0.104)*	*(0.040)*	*(0.086)*
Manufacturing-Machinery	-0.005	0.066 *	0.071	0.152	0.036	0.111
	(0.030)	*(0.037)*	*(0.134)*	*(0.120)*	*(0.045)*	*(0.093)*
Not known	-0.062	0.200 ***	0.262 ***	-0.133	-0.021	0.520 ***
	(0.079)	*(0.038)*	*(0.001)*	*(0.137)*	*(0.226)*	*(0.170)*

Notes: Base = Agriculture.

[a]Source: Table A.3, education in years. [b]p-values from Chi Square test on differences in coefficients are in italics.

*, **, *** Statistically significant at the 10%, 5%, 1% level. Standard errors in parentheses.

8. Analysis of Unobserved Effects

Unlike other studies, we observe the same individuals before and after the regime change and can provide a superior analysis of the variation of wages of individual workers over time. In particular, since managers had discretion in awarding wage premiums under the communist wage grid, it is of interest to assess if individuals who had high or low wage premiums (residuals) related to unobservable characteristics such as skills during communism also enjoyed these premiums during the transition. Using our regression estimates, we decompose the variance of worker-specific wages into the components due to observable determinants and those due to unobservable determinants in the old vs. the new regime. This gives us an interesting insight into the persistence of unobserved components of worker's wages during the regime change.

8.1 The Model

Let observed logarithms of wages of individual *i* under communism

(t = 1) and during the transition ($t = 2$) be given by

$$\begin{aligned} w_{1i} &= x_{1i} b_1 + e_{1i} \\ w_{2i} &= x_{2i} b_2 + e_{2i}, \end{aligned} \tag{3}$$

where x_{1i} and x_{2i} are vectors of observed characteristics in each regime, b's are vectors of corresponding coefficients and e's reflect unobserved determinants of wages. The unobserved individual component of a person's wage in the first period, e_1, may have an effect on the unobserved component in the second period, so that

$$e_{2i} = q\, e_{1i} + u_{2i}. \tag{4}$$

The parameter q captures the persistence of the unobserved individual-specific wage component across regimes, while e_{2i} captures the unobserved component of the wage that is introduced by the transition and is orthogonal to e_{2i}. Hence

$$x_{1i} \perp e_{1i}, x_{2i} \perp e_{2i} \text{ and } e_{1i} \perp n_{2i}. \tag{5}$$

Using equations (4) and (5), the relationship between variances in the unobserved wage can be expressed as

$$\begin{aligned} V(e_{2i}) &= q^2 V(e_{1i}) + V(u_{2i}) \\ COV(e_{1i}, e_{2i}) &= q\, V(e_{1i}). \end{aligned} \tag{6}$$

Note that repeated cross-sectional data do not allow one to inspect the relationships in (6). Our panel data permit us to do so and also to analyze the variance of a worker-specific wage change, $V(w_{2i} - w_{1i})$. Substituting from (4) into (3) and taking into account (6) yields

$$\begin{aligned} V(w_{2i} - w_{1i}) &= V(x_{2i}b_2 + u_{2i} + qe_{1i} - x_{1i}b_1 - e_{1i}) \\ &= V[(x_{2i}b_2 - x_{1i}b_1) + e_{1i}(q-1) + u_{2i}] \\ &= V(\Delta B) + (q-1)^2 V(e_{1i}) + V(u_{2i}) \end{aligned} \tag{7}$$

where $\Delta B \equiv x_{2i}b_2 - x_{1i}b_1$.

Equation (7) decomposes the variance of a worker-specific wage change into three mutually exclusive components: (i) the variance due to changes in observable worker/job characteristics and coefficients of these characteristics, (ii) the variance due to workers' unobserved characteristics determining the wage in the first period, and (iii) the variance due to unobserved determinants of the wage that are introduced by the transition and are orthogonal to unobserved determinants in the first period.

The first component in (7) reflects changes in individual and job characteristics and the corresponding payoffs. For example, a rise in returns to education contributes positively to V(ΔB), while the effect of changing labor market experience depends on where the individual is on the concave wage-experience profile. The value of the second component depends on the persistence of the unobserved individual-specific effect. In the case of full persistence, *q=0,* the part played by unobserved characteristics in the unexplained variation of wages remains unchanged across the regimes and regime change does not affect unobserved wage component of a worker's wage (e.g., general ability is rewarded equally under planning through the wage premium and in the wage setting during the transition). With no persistence, *q=0,* the unobserved component under communism does not translate into the unobserved component during the transition (e.g., entrepreneurial skills are rewarded only by the market and did not appear as an unobserved component in communist wages). One can also expect negative sorting, *q<0,* where communist party membership is for instance rewarded by a wage premium during communism but is punished through negative wage discrimination during the transition. The value of the last term in (7) depends on the extent to which new unobserved components of wages, orthogonal to the unobserved wage component during communism, are introduced during the transition.

Applying the decomposition in (6) and (7) to our panel data, we are able to assess the extent to which wage changes experienced by individual workers stem from their observable characteristics versus unobservable time-invariant and regime-specific effects.

8.2 The Estimating Framework

From the estimated coefficients $\hat{b}_1$ and $\hat{b}_2$, we calculate the residuals for each individual *i* as

$$\begin{aligned} \hat{e}_{1i} &= w_{1i} - x_{1i}\hat{b}_1 \\ \hat{e}_{2i} &= w_{2i} - x_{2i}\hat{b}_2. \end{aligned} \tag{8}$$

The variances in (6) can be consistently estimated as

$$\hat{V}(e_{ti}) = V(\hat{e}_{ti}) = \tfrac{1}{N}\sum_{i=1}^{N}\hat{e}_{ti}^{2}, \text{ for } t = 1,2 \text{ and } i = 1,..,N, \tag{9}$$

where N is the number of individuals. The parameter q can be obtained as an OLS coefficient in (4) or identically as

$$\hat{q} = COV(\hat{e}_{1i}, \hat{e}_{2i}) / V(\hat{e}_{1i}). \tag{10}$$

The remaining variance in (7) is obtained by substituting estimates from (9) and (10) into (6):

$$\hat{V}(u_{2i}) = \hat{V}(e_{2i}) - \hat{q}^2 \hat{V}(e_{1i}). \tag{11}$$

The variances in (7) contribute to the overall variance in wages as follows:

$$\begin{aligned} V(w_{1i}) &= V(x_{1i}b_1) + V(e_{1i}) \\ V(w_{2i}) &= V(x_{2i}b_2) + V(e_{2i}) \end{aligned} \tag{12}$$

and the variance in the deterministic components in (12) can be estimated as

$$\begin{aligned} \hat{V}(x_{1i}b_1) &= V(x_{1i}\hat{b}_1) \\ \hat{V}(x_{2i}b_2) &= V(x_{2i}\hat{b}_2). \end{aligned} \tag{13}$$

Finally,

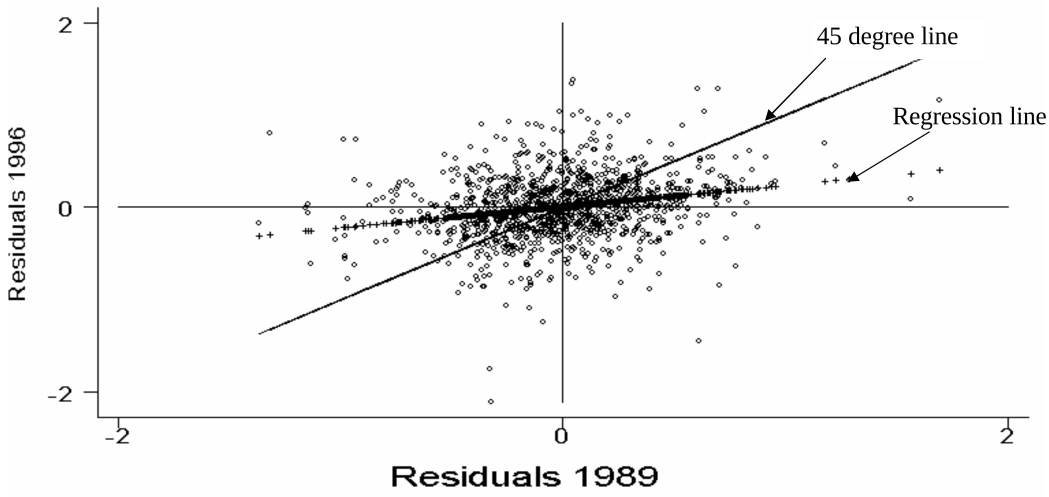
$$\hat{V}(\Delta B) = V(\Delta \hat{B}). \tag{14}$$

8.3 Empirical Estimates

As may be seen from Figure 3, there is a positive relationship in the scatter plot of the 1989 and 1996 residuals. The point estimate of parameter q, capturing this relationship in terms of equation (4), is 0.23 with a standard error of 0.027.

The estimates of the variance components of observed wages are presented in Table 12. The individual cells in the table correspond to the components in equations (6) and (7). Rows 1989 and 1996 refer to cross-sectional variances in these years, while the row titled “within” refers to the variance in worker-specific wage changes. Panel A presents actual variances, while panels B and C present variances as a percentage of the overall cross-

Figure 3
Scatter Plot of Residuals from Equation (8) Estimated with the 1989 vs. 1996 Data



sectional variance for each year and for 1989, respectively. As may be seen throughout Table 12, the variance in wages due to unobserved effects dominates the variance due to observable determinants. However, the variance due to observed determinants increases both absolutely (from 0.019 to 0.031) and in relative terms (from 13% to 20% of total variance) from communism in 1989 to the transition in 1996. The variation in wage changes experienced by individual workers is greater than the cross-sectional variance in both regimes (0.219 > 0.15 in panel A), implying that individual workers experience relatively substantial wage changes. Furthermore, panel B shows that 34% of the variance in wage changes experienced by individual workers is due to unobserved characteristics determining the wage already in 1989 (hence showing persistence over time), while 54% is due to new transition-specific unobserved determinants of wages that are orthogonal to those in 1989. Finally, 11% is due to changes in observed characteristics and their associated coefficients.

The fact that our decomposition estimates are by definition based on the sample of workers who worked in both periods raises the issue of whether our results are biased by excluding workers who worked in only one period. We recognize the problem but think that this exclusion does not substantially change our results for two reasons. First, between 1989 and 1996, labor force participation of Czech men has been very high and the unemployment rate extremely low. Second, we have estimated the cross-sectional variances including all workers in each of the two years and found the results to be very similar to those presented in Table 12.

9. Conclusions

In sum, we find more changes in the returns to education than in the returns to experience. The transition brought about a major increase in these

Table 12
Variance in Wages and Its Decomposition

	$V(w_i)$	$V(X_ib)$	$V(e_i)$	$V(u_i)$
A: Variance of Wages (Vw_i)				
1989	0.144	0.019	0.126	-
1996	0.156	0.031	0.007	0.118
within	0.219	0.024	0.075	0.118
B: Variance in % of $V(w_i)$				
1989	100	13	88	-
1996	100	20	4	76
within	100	11	34	54
C: Variance in % of $V(w_i)$ in 1989				
1989	100	13	88	-
1996	108	22	5	82
within	152	17	52	82

returns to a year of education and the magnitude of this increase is similar in private *de novo* firms, privatized state-owned enterprises (SOEs), and the state sector (SOEs and public administration). We find that those who have obtained (vocational) high school and university degrees experienced more rapid rates of increase in their returns than individuals with basic education (junior high school or apprentices). The sheepskin effect is prevalent and the effect is especially detectable in transition and for higher levels of education in both regimes. Certain fields of study have experienced tremendous increases in their return (e.g., law), while others have not gained in the new market economy (e.g., health and education). On the other hand, with respect to experience, our estimates indicate that men's wage-experience profile was concave in both regimes and did not change from the communist to the transition period. However, we find that the *de novo* firms have a more concave and steeper profile than the state sector, indicating that *de novo* firms pay a higher return to new entrants than the state. Contrary to earlier conjectures, we cannot reject the hypothesis that education and experience obtained during communism are rewarded identically to the education and experience obtained in transition.

The findings presented in this chapter provide two important insights into the functioning of the communist system and the transition economy. First, for decades the communist planners used the wage grid to maintain extremely low education-related wage differentials, but they also generated a significant amount of human capital that is as highly rewarded as post-communist human capital in the nascent market economy. The communist system was hence able to maintain an effective educational system while decoupling it from education-related pecuniary rewards. Moreover, a large part of unobservable, individual-specific wage effects (e.g., skill premiums) has carried over from communism to the market economy. Second, except for the greater concavity of the wage-experience profile in the *de novo* firms, firm ownership during the transition is found to be unrelated to wage differentials based on education and work experience. Hence, factors such as the reduction of state subsidies, opening up of the economy to the world and allowing competition in the labor market are sufficient to generate human capital-related wage differentials that on average do not vary with principal types of firm ownership in the economy.

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Notes for Chapter I

[1] See e.g., Dyba and Svejnar (1995).

[2] See for example Bird, et al. (1994), Chase (1998), Flanagan (1995, 1998), Jones and Illayperuma (1994), Krueger and Pischke (1995), Nesterova and Sabirianova (1999), Orazem and Vodopivec (1997) and Rutkowski (1996).

[3] Krueger and Pischke (1995) deal with East Germany and Chase (1998) and Flanagan (1998) with the Czech Republic. Unlike Hungary and Poland, East Germany and the Czech Republic both adhered to the wage grid until the very end of the communist regime and hence provide interesting laboratories.

[4] A potential weakness of the retrospective data set is recall error, as individuals may not accurately remember their past wages. As we discuss below, we check the magnitude and minimize the effect of this error in a number of ways.

[5] Chase (1998) does not have data on firm ownership.

[6] See e.g., Windmuller (1970), Svejnar (1974), Adam (1984), and Flanagan (1998). In addition to personal evaluation bonuses, the managers could influence total compensation and hence compete for workers by offering various social benefits, such as subsidized housing. However, they could not change the centrally set wage rates.

[7] See Ministry of Agriculture (1952) and Ministry of Labor and Social Affairs, (1981, 1985 and 1986).

[8] For many years, planners favored “productive” sectors (industry, construction and agriculture) over the “unproductive” sectors (trade and services) and wages in the productive sectors were boosted above the others. In some years, the location of the job within the government hierarchy (headquarters vs. branch office) also mattered.

[9] Discussions with officials who used to administer the wage grid indicate that the process was taken very seriously and that administrators from various Soviet bloc countries compared notes and experiences. In this respect, the wage grid was an integral part of the centrally planned system.

[10] We have examined the internal wage setting practices within several hundred firms with diverse ownership. Using the Trexima firm data over 1995 to 1998, we have found that as late as 1998, most state owned and privatized firms still used a modified wage grid that had been carried forward from the communist days. The privatized enterprises were not required to pay according to a grid and their adherence to a grid system reflected inertia in (transaction costs

related to changing) their compensation practices. In contrast, the *de novo* private firms have been found to operate outside this grid. Moreover, government intervention in private sector wage setting has been minimal, although loose wage controls were in effect intermittently from 1991 to 1995.

[11] The January 1989 date was selected as a point in time for which people were likely to remember their labor market characteristics since 1989 was the year of the revolution that toppled the communist regime. See Münich et al. (1997) for a description of the survey and sample design as well as the descriptive statistics of the sample relative to the Labor Force Survey data.

[12] In order to confirm that our sample is representative of the working population in 1996 and 1989, we compare the means and distributions of the demographic characteristics of the working age population in our survey with those from the last quarter of the 1996 Labor Force Survey and of the 1989 Social Attitudinal Survey. As seen in Appendix Table 1, we find they are strikingly similar.

[13] In fact, this question yields data on jobs that began as early as the 1940s: 0.3 percent of all the job starts reported occurred before 1951, 2.6 percent occurred during the 1951-60 period, 5.5 percent during 1961-70, 9.2 percent during 1971-80, 18.9 percent during 1981-90, and 63.5 percent during 1991-96. We concluded that the very early data points went too far back in time to be reliable and that they might be confounded with the systemic changes that accompanied the communist takeover of 1948. As a result, we restricted our observations on job starts to those that occurred from 1955 onward since by 1955 the revolutionary period, nationalization and currency reform that followed the communist *coup d'état* of 1948 were over and the centrally planned system was fully in place.

[14] The retirement age for men was 60 years of age, although many retirees continued to work on a full- or part-time basis.

[15] The monthly nominal earnings are meant to be net of payroll and income taxes. This is the most common way that the Czechs recall their salary, since both of these taxes are taken out before they receive their pay. However, about 25 percent of the respondents preferred to report their gross rather than net earnings. As a result, we have included as a regressor a dummy variable to control for this discrepancy in reporting. In addition, net earnings in some cases include benefits provided by the state, through the employer, for raising children. We

have therefore also included a dummy variable to control for the cases when the reported earnings include children benefits.

[16] We have also tested for the effect of marital status in equation (1) and found it to be insignificant.

[17] We would like to thank Orley Ashenfelter for suggesting the combined specification to us.

[18] Our data permit us to estimate a specification with six categorical variables reflecting the highest degree attained: 1) junior high school (mandatory education of 9 years), 2) apprentices in 2 year programs, 3) apprentices in 3 year programs, 4) technical high school graduates and apprentices in 4 year programs who received the technical high school diploma, 5) academic high school graduates, and 6) university graduates and above.

[19] The respondents were asked not to report any years of repeated grades.

[20] The "sheepskin effect" refers to the fact that wages may not increase steadily with years of education within a given level of schooling but may jump up when a degree is received (see also Heckman et al., 1996). Using U.S. data, Hungerford and Solon (1987) find significant discrete jumps in the return to education upon receiving a degree.

[21] The shortcoming of this variable is that it includes periods during which the individual may have been out of the labor market and acquired less labor force experience. This of course tends to be less of a problem in the case of men than women, who are likely to take long maternity leaves (Mincer and Polachek, 1974 and Mincer and Ofek, 1982).

[22] Since the dependent variable is in nominal terms, we include annual dummies to control for changes in prices in all the models with time-varying coefficients. We have also tested for the validity of a higher than linear time-varying-coefficient model but we have not found strong support for this higher order specification.

[23] Paukert (1995) finds that between 1989 and 1994 labor force participation rates of men and women (over 15 years of age) fell between six and eight percentage points in the Czech Republic, Hungary, Poland and Slovakia, and that the absolute decline was about the same for men and women in each country. Our survey provides us with a number of variables that can be used to impose exclusion restrictions in that they are likely to affect the respondent's labor force participation decision but not his wage. In particular, we derived Heckman's (1979) λ by estimating a probit equation with the 1996 cross-section data, using as explanatory variables a marital status dummy, a dummy variable for the presence of children under 15 years of age in the household, the per capita household income minus the income of the

respondent, a dummy variable for Prague, the district level vacancy rates (the number of vacancies per working age population), and the respondent's age, age^2, and education (in years). The estimation yields a positive and significant λ, but the estimated coefficients on education and experience remain unaffected by the correction procedure.

[24] The complete set of our estimates of equation (1) using the 1989 and 1996 cross-sectional data is presented in appendix Table A.3.

[25] The overall cross-sectional estimate for 1989 (2.7%) and 1996 (5.8%) are the same estimates presented in Table 3.

[26] The lowest p value is 0.43 for the difference between State and privatized firms. Flanagan (1998) found the returns to a year of education in 1996 to be lowest in the new private firms (5.8%), highest in the privatized firms (7.2%) and intermediate (6.2%) in the state sector. However, since Flanagan does not report standard errors and relative tests of significance for these estimates, it is not possible to know if they are statistically different from one another or from our estimates as well. We note that in Flanagan's data the years of education are imputed and include both men and women which may account for the possible difference in his and our estimates. Finally, Flanagan's and our data also reveal lower payoffs to vocational education in the newly created private firms, but the difference in our data is not statistically significant. Again, Flanagan (1998) does not report formal tests for differences of coefficients and we hence cannot establish if the two studies yield similar or dissimilar results.

[27] Each of the four schooling levels below university level represents a direct path from junior high school the mandatory level of education). Hence, the annual return to a year of education within these levels of schooling relative to junior highs school (r_s) is calculated as the nth root of the rate of return to the schooling level (R_s), where s represents the level of schooling and n represents the number of years of education in each level: $r_s = (R_s)^{1/n}$. However, the return to a year of university education represents a return above either academic or vocational high school, and hence it is calculated as $r_u = (R_u - R^*_{hs})^{1/n}$, where the star denotes the average value.

[28] The return is calculated as [exp(coefficient) – 1]; in this case [exp(0.544) – 1] = 72%.

[29] The 1955-90 results also indicate that men with academic high school and university degrees had higher starting wages than others and that the wages of high school and university graduates were not statistically different from each other (p-value of 0.96).

[30] In fact, the coefficient difference between State and privatized firms and State and *de novo* firms is found to be statistically significant at 5% and 9%, respectively.

[31] The P values for the F tests are 0.560 in 1989 and 0.558 for 1996.

[32] The F(1,1934) = 1.36, P value = 0.243 for 1989; and F(1,1610) = 5.72, P value = 0.017 for 1996.

[33] The actual coefficients reported from other studies in Table 2 are not necessarily higher than ours. As mentioned earlier, Flanagan's (1996) estimates come from data that, by construction of the earnings variable, produce a downward bias. The coefficients from the other studies refer to earlier years in the 1990s when the return on education was still low.

[34] These results are based on the regression where education is measured as "actual years of schooling" (Table A.3). We also estimated experience coefficients in a regression with education measured as "level of attainment" (Table A.4) and found that there was no statistically significant difference in the experience-earnings profiles estimated with years vs. attainment. The F test statistics are F(2, 3547) = 0.07 for the 1989 vs. 1996 comparison based on the specification with years of education and F(2, 3539) = 0.28 for the 1989 vs. 1996 comparison based on the specification with levels of education. To save space, we only report one set of experience coefficients.

[35] In order to check the robustness of these findings, we have also re-estimated the three ownership-specific equations with all coefficients, except those on education, experience and experience squared, being constrained to be equal. The resulting estimates are very similar to those reported in Table 8. We have also estimated spline experience-earnings profiles, where the splines capture three ten-year experience intervals from the start of one's career. Although the spline functions generate similar results to the coefficients on the quadratic experience profiles, in that they are similar in 1989 and 1996 for all workers, they highlight a greater decline in the returns to workers with more than 30 years experience in 1996 than in 1989; it is clear that it is the *de novo* firms that are driving this steeper slope for the 30+ segment. As we noted in the paragraph above, this corresponds to the greater concavity of the wage-experience profile in the *de novo* firms. And as with the quadratic experience estimates described above, the spline profile (at least for men with 30 or fewer years of experience) in *de novo* firms is clearly above that of the privatized firms and state sector, which are very similar. Estimating spline functions at other than ten year intervals did not fundamentally change the results.

[36] The relevant F statistic is F(4, 3266) = 0.29 for the model based on years of education and 0.28 for the model based on levels of educational attainment.

[37] Chase's coefficients on experience and experience squared are 0.014 and –0.0003, respectively. The corresponding coefficients for Krueger and Pischke are 0.014 and -0.0002, and coefficients for Vecerník (2001) are 0.037 and -.00077.

[38] The various data sets may also have different age compositions of workers. In particular, depending on the number of individuals that a sample contains from different age categories, one's estimates may reflect the concave or flatter parts of the wage-experience profile.

[39] The coefficient for "communist education" was 0.058 (S.E.=0.005) and the coefficient for "post-communist education" was 0.031 (S.E.= 0.013). F(1,1610) = 4.65, with Prob > F = 0.03.

[40] The coefficients on the education coefficient for individuals less than 30 and for those greater than or equal to 30 are 0.063 and 0.059, respectively, in 1996. The F test indicates that the hypothesis of zero difference cannot be rejected.

[41] The F test value on the joint significance is F(2, 2078) = 1.22.

[42] These coefficients are reported in full in Table A.3.

[43] The reported pattern is very similar to the one obtained when one does not control for workers' human capital characteristics.

[44] The nine-point scatter in fact generates a negative correlation coefficient of –0.41.

Appendix tables for Chapter I

Table A.1
Means and Standard Deviation of Variables in Cross-Sectional Data

	Communism 1989		Transition 1996	
	mean	**st.dev.**	**mean**	**st.dev.**
Log of monthly wage	8.227	(0.394)	8.961	(0.404)
Experience (years)	18.2	(11.458)	20.4	(11.992)
Experience^2	463.3	(490.445)	559.8	(545.45)
Education in years	12.776	(2.519)	12.626	(2.347)
% of Population with Given Education:				
Junior High School (reference group)	0.057	(0.394)	0.047	(0.212)
Apprentices w/2 years	0.048	(0.213)	0.035	(0.184)
Apprentices w/3 years	0.484	(0.500)	0.503	(0.500)
Vocational H.S. w/4 years	0.258	(0.438)	0.274	(0.446)
Academic H.S. w/4 years	0.022	(0.147)	0.023	(0.149)
University	0.131	(0.338)	0.119	(0.323)
Apprenticeship:				
Machine control	0.028	(0.164)	0.029	(0.168)
Manuf. Machinery and Metalurgy	0.199	(0.399)	0.200	(0.400)
Electrotechnics, transport, telecom.	0.069	(0.254)	0.073	(0.260)
Chemistry, Food processing	0.016	(0.125)	0.018	(0.132)
Textile, Clothing	0.007	(0.084)	0.004	(0.061)
Wood, Shoes manufacturing	0.025	(0.157)	0.031	(0.173)
Construction	0.089	(0.284)	0.089	(0.284)
Agriculture, Forestry	0.040	(0.197)	0.042	(0.202)
Trade, Services	0.029	(0.168)	0.022	(0.145)
Other	0.030	(0.170)	0.031	(0.173)
Academic High School	0.022	(0.147)	0.023	(0.149)
Vocational High School:				
Natural sciences	0.004	(0.060)	0.002	(0.050)
Manufacturing-Machinery	0.091	(0.288)	0.094	(0.292)
Electrotechnics	0.046	(0.209)	0.058	(0.235)
Construction	0.019	(0.136)	0.017	(0.130)
Other technical branches	0.016	(0.127)	0.018	(0.135)
Agriculture	0.023	(0.149)	0.022	(0.147)
Health	0.003	(0.055)	0.006	(0.074)
Business, Trade, Services	0.028	(0.164)	0.027	(0.162)
Law	0.001	(0.032)	0.001	(0.035)
Teaching	0.002	(0.045)	0.002	(0.050)
Other social branches	0.005	(0.071)	0.004	(0.065)
Other	0.021	(0.142)	0.020	(0.141)
University:				
Natural sciences	0.010	(0.098)	0.007	(0.082)
Manufacturing-Machinery	0.023	(0.150)	0.024	(0.153)
Electrotechnics	0.009	(0.096)	0.009	(0.096)
Construction	0.013	(0.112)	0.012	(0.107)

Table A.1 continued
Means and Standard Deviation of Variables in Cross-Sectional Data

	Communism 1989		Transition 1996	
	mean	**st.dev.**	**mean**	**st.dev.**
Other technical branches	0.010	(0.101)	0.008	(0.089)
Agriculture	0.013	(0.115)	0.012	(0.107)
Health	0.008	(0.087)	0.008	(0.089)
Business, Trade, Services	0.012	(0.110)	0.009	(0.096)
Law	0.006	(0.078)	0.005	(0.070)
Teaching	0.016	(0.125)	0.015	(0.123)
Other social branches	0.005	(0.068)	0.004	(0.061)
Other	0.006	(0.078)	0.006	(0.078)
Other variables:				
Prague	0.106	(0.307)	0.116	(0.320)
Child benefits included	0.197	(0.398)	0.110	(0.313)
Gross earnings reported	0.247	(0.431)	0.226	(0.418)
Industry:				
Mining & Quarrying	0.088	(0.283)	0.074	(0.261)
Construction	0.116	(0.320)	0.122	(0.327)
Wholesale, Retail,	0.099	(0.299)	0.138	(0.345)
Broad public	0.127	(0.333)	0.136	(0.343)
Finance, Insurance, Renting & Real Estate	0.005	(0.068)	0.015	(0.121)
Transport, Telecommunications	0.082	(0.274)	0.082	(0.274)
Manufacturing-Food, Textile,	0.241	(0.428)	0.252	(0.434)
Manufacturing-Machinery	0.118	(0.323)	0.112	(0.315)
Households + Exteritorial + Not known	0.010	(0.101)	0.009	(0.096)
Firm Size:				
1-25 employees			0.258	(0.438)
26-100 employees			0.211	(0.408)
101-500 employees			0.238	(0.426)
>500 employees			0.256	(0.437)
Not known			0.037	(0.214)
Ownership:				
Privatized			0.310	(0.445)
SOE & Public Administration			0.236	(0.341)
De Novo Private			0.371	(0.483)
Other & not known			0.083	(0.276)
Employment status:				
Employee			0.900	
Employer			0.025	(0.157)
Self-employed			0.067	(0.250)
HH Helper + Not known			0.008	(0.089)
Log of district level enemployment rate			0.035	(0.021)
No. of Obs.	1951		1627	

Table A.2
Means and Standard Deviation of Variables for Start Date Data

	Communism (1989)		Transition (1996)	
	mean	**st.dev.**	**mean**	**st.dev.**
Log of earnings	8.049	(0.549)	8.509	(0.484)
Experience	7.009	(9.178)	13.442	(12.653)
Exper. x time	-640	(1184.8)	381	(534.65)
Experience ^{2}	135	(302.504)	341	(511.855)
Exper. ^{2} time	-921	(2053.8)	786	(1598.7)
Years of education	12.843	(2.526)	12.428	(2.261)
Education · time	-151	(126.8)	32	(23.52)
Apprentice (2 years)	0.037	(0.190)	0.036	(0.185)
Apprentice (2) · time	-	-	0.088	(0.572)
Apprentice (3 years)	0.475	(0.500)	0.533	(0.499)
Apprentice (3) · time	-5.4	(8.482)	1.4	(1.836)
Vocational H.S.	0.268	(0.443)	0.243	(0.429)
Vocational H.S. · time	-3.3	(7.300)	0.6	(1.442)
Academic H.S.	0.022	(0.146)	0.036	(0.185)
Academic H.S. · time	-0.3	(2.501)	0.1	(0.671)
University	0.143	(0.350)	0.101	(0.302)
University · time	-1.4	(5.016)	0.2	(0.896)
Prague	0.111	(0.314)	0.121	(0.327)
Child ben. incl,	0.136	(0.343)	0.089	(0.284)
Gross earnings	0.258	(0.437)	0.226	(0.418)
Industry:				
Machine Control	0.093	(0.290)	0.049	(0.216)
Electro., trans., tele.m.	0.098	(0.298)	0.175	(0.380)
Chemistry, Food processir	0.096	(0.295)	0.187	(0.390)
Textile, Clothing	0.125	(0.331)	0.112	(0.315)
Wood, Shoes manufac.	0.007	(0.083)	0.012	(0.108)
Construction	0.075	(0.264)	0.062	(0.241)
Agriculture, Forestry	0.244	(0.429)	0.254	(0.435)
Trade, Services	0.134	(0.341)	0.080	(0.272)
Other	0.007	(0.083)	0.008	(0.089)

Table A.2 continued
Means and Standard Deviation of Variables for Start Date Data

	Communism (1989)		Transition (1996)	
	mean	**st.dev.**	**mean**	**st.dev.**
Firm Size:				
1-25 employees	-	-	0.336	(0.472)
26-100 employees	-	-	0.245	(0.430)
101-500 employees	-	-	0.209	(0.407)
>500 employees	-	-	0.172	(0.377)
Not known	-	-	0.038	(0.192)
Ownership:				
Privatized	-	-	0.196	(0.397)
SOE & Public administrati	-	-	0.229	(0.420)
De Novo Private	-	-	0.495	(0.500)
Other & not known	-	-	0.081	(0.272)
Employment status:				
Employee	-	-	0.911	(0.472)
Employer	-	-	0.018	(0.131)
Self-employed	-	-	0.061	(0.240)
HH helper + Not known	-	-	0.010	(0.102)
No. of Obs.	1285		2107	

Table A.3
Cross-sectional Earnings Functions, 1989 and 1996 (Education by years)

	Communism	Transition			
	All	**All**	**State**	**Privatized**	**DeNovo**
Education	0.027	0.058	0.056	0.065	0.061
	(0.004)	*(0.005)*	*(0.009)*	*(0.007)*	*(0.010)*
Experience	0.021	0.021	0.015	0.022	0.030
	(0.003)	*(0.005)*	*(0.006)*	*(0.007)*	*(0.004)*
Experience^2	-0.0004	-0.0004	0.000	0.000	-0.001
	(0.0001)	*(0.000)*	*(0.000)*	*(0.000)*	*(0.000)*
Prague	0.015	0.120	0.151	0.088	0.177
	(0.027)	(0.032)	*(0.047)*	-*(0.064)*	*(0.057)*
Child benefits included	0.061	0.064	0.051	0.112	0.054
	(0.022)	*(0.026)*	-*(0.042)*	*(0.052)*	-*(0.045)*
Gross earinings	0.122	0.069	0.082	0.045	0.091
	(0.020)	*(0.022)*	*(0.041)*	-*(0.031)*	*(0.040)*
Mining & Quarrying	0.251	0.092	0.245	0.063	-0.079
	(0.039)	*(0.044)*	*(0.099)*	*(0.058)*	*(0.159)*
Construction	0.051	0.131	0.110	0.082	0.119
	(0.035)	*(0.040)*	*(0.120)*	*(0.058)*	*(0.091)*
Wholesale, Retail	0.025	0.163	-0.134	0.060	0.147
	(0.037)	*(0.041)*	*(0.138)*	*(0.062)*	*(0.087)*
Public Adm., Educ, Health	0.021	0.059	0.080	-0.190	0.085
	(0.035)	*(0.115)*	*(0.090)*	*(0.219)*	*(0.090)*
Finance, Insur., Real Est.	0.203	0.052	0.140	0.054	-0.017
	(0.139)	*(0.080)*	*(0.171)*	*(0.116)*	*(0.170)*
Transport & Telecom.	0.059	0.146	0.096	0.122	0.275
	(0.036)	*(0.040)*	*(0.095)*	*(0.062)*	*(0.095)*
Manuf. - Food, Textile,	0.017	0.092	0.045	0.063	0.118
	(0.028)	*(0.033)*	*(0.104)*	*(0.040)*	*(0.086)*
Manuf.-Machinery	-0.005	0.066	0.152	0.036	0.111
	(0.030)	*(0.037)*	*(0.120)*	*(0.045)*	*(0.093)*
Not known	-0.062	0.200	-0.133	-0.021	0.520
	(0.079)	*(0.038)*	*(0.137)*	*(0.226)*	*(0.170)*
Constant	7.620	7.916	7.919	7.812	7.845
	(0.055)	*(0.071)*	*(0.143)*	*(0.099)*	*(0.157)*
adj.R ^{2}	***0.118***	***0.190***	***0.256***	***0.23***	***0.23***
nobs	***1951***	***1627***	***384***	***504***	***604***

Base = people working outside Prague, whose earnings are net of tax and child benefits, and who work in agriculture.

Table A.4

Cross-sectional Earnings Functions, 1989 and 1996 (Education by levels)

	Communism		Transition				
	(1)	**(2)**	**(1)**	**(2)**	**State**	**Privatized**	**DeNovo**
Apprentice (2 years)	0.0701	0.0635	0.1128	0.0939	0.1290	0.1143	0.1009
	(0.052)	*(0.051)*	*(0.058)*	*(0.057)*	*(0.121)*	*(0.065)*	*(0.137)*
Apprentice (3 years)	0.0923	0.0773	0.1434	0.1122	0.0968	0.1559	0.0652
	(0.038)	*(0.037)*	*(0.049)*	*(0.049)*	*(0.105)*	*(0.058)*	*(0.115)*
Vocational H.S.	0.1374	0.1265	0.3228	0.2943	0.3232	0.3266	0.2492
	(0.040)	*(0.040)*	*(0.050)*	*(0.050)*	*(0.105)*	*(0.058)*	*(0.118)*
Academic H.S.	0.1525	0.1346	0.3822	0.3508	0.4011	0.2656	0.3425
	(0.080)	*(0.081)*	*(0.102)*	*(0.107)*	*(0.142)*	*(0.164)*	*(0.309)*
University	0.2793	0.2826	0.5515	0.5439	0.4758	0.6734	0.5993
	(0.044)	*(0.045)*	*(0.058)*	*(0.059)*	*(0.115)*	*(0.072)*	*(0.133)*
Experience	0.022	0.021	0.024	0.024	0.021	0.027	0.033
	(0.003)	*(0.003)*	*(0.005)*	*(0.005)*	*(0.006)*	*(0.007)*	*(0.004)*
Experience^2	-0.00047	-0.00045	-0.00050	-0.00051	-0.00041	-0.00054	-0.00076
	(0.0001)	*(0.0001)*	*(0.0001)*	*(0.0001)*	*(0.0001)*	*(0.0002)*	*(0.0001)*
Prague	-	0.009	-	0.102	0.142	0.045	0.140
	-	*(0.027)*	-	(0.032)	(0.047)	(0.061)	(0.055)
Child benefits included	-	0.065	-	0.076	0.056	0.122	0.076
	-	*(0.021)*	-	*(0.026)*	*(0.042)*	*(0.051)*	*(0.044)*
Gross earnings	-	0.125	-	0.080	0.088	0.050	0.102
	-	*(0.020)*	-	*(0.021)*	*(0.041)*	*(0.031)*	*(0.038)*

Table A.4 continued: Cross-sectional Earnings Functions, 1989 and 1996 (Education by levels)

	Communism		Transition				
	(1)	**(2)**	**(1)**	**(2)**	**State**	**Privatized**	**DeNovo**
Mining & Quarrying	-	0.250	-	0.095	0.271	0.063	-0.045
	-	*(0.039)*	-	*(0.043)*	*(0.089)*	*(0.058)*	*(0.150)*
Construction	-	0.053	-	0.145	0.150	0.096	0.144
	-	*(0.035)*	-	*(0.040)*	*(0.114)*	*(0.060)*	*(0.091)*
Wholesale and Retail	-	0.020	-	0.150	-0.067	0.028	0.136
	-	*(0.036)*	-	*(0.040)*	*(0.120)*	*(0.058)*	*(0.086)*
Public Admin, Edu, Health	-	0.012	-	0.034	0.095	-0.039	0.068
	-	*(0.035)*	-	*(0.038)*	*(0.080)*	*(0.210)*	*(0.102)*
Finance, Insur. & Real Estate	-	0.210	-	0.024	0.091	0.046	-0.064
	-	*(0.131)*	-	*(0.076)*	*(0.157)*	*(0.119)*	*(0.185)*
Transport & Telecom.	-	0.057	-	0.149	0.115	0.144	0.305
	-	*(0.036)*	-	*(0.039)*	*(0.087)*	*(0.057)*	*(0.094)*
Manuf. - Food, Textile,	-	0.018	-	0.092	0.077	0.063	0.135
	-	*(0.028)*	-	*(0.032)*	*(0.097)*	*(0.039)*	*(0.085)*
Manuf. - Machinery	-	-0.010	-	0.066	0.172	0.026	0.137
	-	*(0.030)*	-	*(0.036)*	*(0.116)*	*(0.044)*	*(0.091)*
Not known	-	-0.064	-	0.180	-0.167	-0.016	0.525
	-	*(0.082)*	-	*(0.111)*	*(0.094)*	*(0.223)*	*(0.176)*
Constant	7.910	7.847	8.516	8.404	8.331	8.324	8.401
	(0.043)	*(0.046)*	*(0.054)*	*(0.059)*	*(0.136)*	*(0.077)*	*(0.143)*
R^2	***0.070***	***0.120***	***0.181***	***0.210***	***0.280***	***0.270***	***0.270***
nobs	***1955***	***1951***	***1639***	***1627***	***384***	***504***	***604***

Table A.5
Cross-sectional Earnings Functions, 1989 and 1996
(Education by Levels and Field of Study)

	Communism 1989	**Transition 1996**
Apprenticeship Fields of study **:**		
Machine control	0.123	0.084
	(0.053)	*(0.062)*
Manuf. Machinery and Metalurgy	0.113	0.139
	(0.040)	*(0.051)*
Electrotechnics, transport, telecom.	0.076	0.122
	(0.045)	*(0.056)*
Chemistry, Food processing	0.122	0.031
	(0.068)	*(0.085)*
Textile, Clothing	-0.056	-0.194
	(0.071)	*(0.133)*
Wood, Shoes manufacturing	0.071	0.073
	(0.056)	*(0.061)*
Construction	0.054	0.154
	(0.046)	*(0.060)*
Agriculture, Forestry	-0.040	-0.007
	(0.053)	*(0.064)*
Trade, Services	0.007	0.161
	(0.067)	*(0.071)*
Other	0.093	0.163
	(0.061)	*(0.067)*
Academic High School	0.138	0.352
	(0.081)	*(0.106)*
Fields within vocational high school:		
Natural sciences	0.185	0.745
	(0.127)	*(0.303)*
Manufacturing-Machinery	0.120	0.289
	(0.045)	*(0.052)*
Electrotechnics	0.120	0.361
	(0.052)	*(0.058)*
Construction	0.138	0.309
	(0.077)	*(0.079)*
Other technical branches	0.238	0.265
	(0.070)	*(0.073)*
Agriculture	0.011	0.163
	(0.065)	*(0.063)*

Table A.5 continued
Cross-sectional Earnings Functions, 1989 and 1996
(Education by Levels and Field of Study)

	Communism 1989	**Transition 1996**
Health	-0.011	0.084
	(0.118)	*(0.129)*
Business, Trade, Services	0.099	0.280
	(0.068)	*(0.069)*
Law	0.539	0.617
	(0.348)	*(0.119)*
Teaching	0.215	0.223
	(0.172)	*(0.154)*
Other social branches	0.198	0.240
	(0.101)	*(0.198)*
Other	0.210	0.354
	(0.071)	*(0.082)*
Fields within university education:		
Natural sciences	0.135	0.454
	(0.106)	*(0.157)*
Manufacturing-Machinery	0.274	0.571
	(0.074)	*(0.082)*
Electrotechnics	0.300	0.746
	(0.069)	*(0.130)*
Construction	0.275	0.569
	(0.076)	*(0.104)*
Other technical branches	0.488	0.753
	(0.079)	*(0.136)*
Agriculture	0.305	0.496
	(0.077)	*(0.080)*
Health	0.315	0.246
	(0.091)	*(0.166)*
Business, Trade, Services	0.350	0.643
	(0.117)	*(0.144)*
Law	0.394	1.054
	(0.112)	*(0.138)*
Teaching	0.266	0.314
	(0.083)	*(0.091)*
Other social branches	0.129	0.139
	(0.087)	*(0.101)*
Other	-0.007	0.548
	(0.129)	*(0.088)*

Table A.5 continued
Cross-sectional Earnings Functions, 1989 and 1996
(Education by Levels and Field of Study)

	Communism 1989	**Transition 1996**
Experience	0.021	0.025
	(0.003)	*(0.0049)*
Experience^2	-(0.00044)	-(0.00052)
	(0.00006)	*(0.0001)*
Prague	0.008	0.108
	(0.028)	*(0.031)*
Child Benefits	0.063	0.081
	(0.021)	*(0.026)*
Gross Earnings	0.130	0.085
	(0.020)	*(0.021)*
Industry:		
Mining & Quarrying	0.214	0.046
	(0.040)	*(0.045)*
Construction	0.027	0.086
	(0.039)	*(0.045)*
Wholesale and Retail Trade	-0.005	0.098
	(0.037)	*(0.041)*
Finance, Insur. & Real Estate	0.167	-0.014
	(0.132)	*(0.077)*
Transport & Telecom.	0.019	0.097
	(0.037)	*(0.042)*
Manuf.-Food, Textile,	-0.021	0.046
	(0.029)	*(0.034)*
Manuf.-Machinery	-0.051	0.013
	(0.033)	*(0.039)*
Public Admin., Edu, Health	-0.015	0.017
	(0.038)	*(0.041)*
Not known	-0.089	0.135
	(0.082)	*(0.112)*
Constant term	7.877	8.431
	(0.046)	*(0.060)*
adj. R^2	***0.129***	***0.240***
nobs	***1951***	***1627***

Base = Jr. H.S. graduates working outside Prague in agriculture, whose earnings net of tax and child benefits.

Table A.6
Earnings Regressions with Time Varying Cofficients; Communism and Transition (Education in Years)

Period	Communism	Transition			
		All	State	Privatized	DeNovo
Education	0.0166	0.0219	0.0276	0.0273	0.0308
	(0.0099)	*(0.0072)*	*(0.0118)*	*(0.0121)*	*(0.0123)*
Education·t	-0.0003	0.0093	0.0098	0.0104	0.0077
	(0.0007)	*(0.0020)*	*(0.0050)*	*(0.0039)*	*(0.0030)*
Experience	0.0236	0.0285	0.0349	0.0256	0.0283
	(0.0053)	*(0.0053)*	*(0.0066)*	*(0.0086)*	*(0.0060)*
Experience·t	0.0007	0.0017	0.0018	0.0012	0.0012
	(0.0005)	*(0.0014)*	*(0.0030)*	*(0.0026)*	*(0.0018)*
$Experience^2$	-0.0005	-0.0006	0.0008	0.0006	0.0007
	(0.0002)	*(0.0001)*	*(0.0001)*	*(0.0002)*	*(0.0001)*
$Experience^2$·t	0.0000	0.0000	0.0000	0.0000	0.0000
	(0.0000)	*(0.0000)*	*(0.0001)*	*(0.0001)*	*(0.0000)*
Prague	-0.1257	0.1506	0.1911	0.1111	0.1856
	(0.0460)	*(0.0279)*	*(0.0561)*	*(0.0651)*	*(0.0369)*
Child benefits included	0.2282	0.1194	0.1817	0.0512	0.1136
	(0.0403)	*(0.0296)*	*(0.0616)*	*(0.0653)*	*(0.0392)*
Gross Earnings	0.1328	0.0420	0.0863	0.1133	0.0360
	(0.0514)	*(0.0441)*	*(0.0944)*	*(0.0780)*	*(0.0613)*
Gross Earnings·t	0.0899	0.0176	0.0628	0.0302	0.0216
	(0.1601)	*(0.0131)*	*(0.0326)*	*(0.0259)*	*(0.0176)*
Industry:					
Mining & Quarrying	0.2759	0.0448	0.1965	0.0382	0.2002
	(0.0553)	*(0.0548)*	*(0.1287)*	*(0.0792)*	*(0.1094)*
Construction	0.1337	0.1287	-0.0241	0.1627	0.0249
	(0.0520)	*(0.0430)*	*(0.1255)*	*(0.0707)*	*(0.0578)*
Wholesale and Retail	-0.0540	0.1186	-	0.1844	0.0110
	(0.0589)	*(0.0447)*	-	*(0.0945)*	*(0.0593)*
Public Adm., Edu, Health	0.0937	0.0650	0.1244	-	-
	(0.0513)	*(0.0470)*	*(0.1310)*	-	-
Finance, Insurance, R. Est.	0.1161	0.0047	0.1294	0.0406	0.0489
	(0.2079)	*(0.0818)*	*(0.1142)*	*(0.1378)*	*(0.0746)*
Transport & Telecom.	0.0963	0.1010	0.2232	0.0060	0.1819
	(0.0632)	*(0.0551)*	*(0.1959)*	*(0.1032)*	*(0.1379)*
Manuf.-Food, Textile	-0.0021	0.0253	0.1072	0.1721	0.0753
	(0.0441)	*(0.0414)*	*(0.1204)*	*(0.1008)*	*(0.0856)*
Manuf.-Machinery	-0.0162	0.0855	0.0018	0.0086	0.0079
	(0.0494)	*(0.0482)*	*(0.1171)*	*(0.0634)*	*(0.0569)*
Not known	0.0639	0.1963	0.0351	0.0778	0.1842
	(0.1303)	*(0.1032)*	*(0.1379)*	*(0.1119)*	*(0.1578)*
Constant	7.9297	7.7520	7.5578	7.6788	7.8586
	(0.1289)	*(0.0944)*	*(0.1799)*	*(0.1553)*	*(0.1707)*
adj.R^2	***0.172***	***0.285***	***0.384***	***0.269***	***0.356***
nobs	***1285***	***2107***	***483***	***1045***	***579***

Base= individuals working outside Prague in agriculture, whose earnings are net of tax, and child benefits.

Table A.7
Earnings Regressions; Time Varying Cofficients for Communism and Transition (Education in Levels)

Period	Communism	Transition			
		All	**State**	**Privatized**	**DeNovo**
Apprentice (2 years)	0.0566	0.0783	0.1532	0.1542	-0.0658
	(0.1007)	*(0.1062)*	*(0.1673)*	*(0.1562)*	*(0.1635)*
Apprentice (2 years)·t	n.a.	n.a	n.a	n.a	n.a
	n.a.	*n.a*	*n.a.*	*n.a.*	*n.a.*
Apprentice (3 years)	0.0690	0.0489	0.0950	0.1185	0.0865
	(0.0745)	*(0.0691)*	*(0.1117)*	*(0.1032)*	*(0.0775)*
Apprentice (3 years)·t	-0.0003	0.0528	0.0652	0.0417	0.0315
	(0.0051)	*(0.0206)*	*(0.0217)*	*(0.0216)*	*(0.0150)*
Vocational H.S.	0.056	0.051	0.0591	0.2034	0.1827
	(0.082)	*(0.074)*	*(0.1243)*	*(0.1169)*	*(0.0911)*
Vocational H.S.·t	-0.0014	0.0768	0.1022	0.0474	0.0322
	(0.0059)	*(0.022)*	*(0.0323)*	*(0.0217)*	*(0.0191)*
Academic H.S.	0.3378	0.0896	0.2993	0.0585	0.0133
	(0.1783)	*(0.1126)*	*(0.1857)*	*(0.2400)*	*(0.1862)*
Academic H.S.·t	0.0104	0.0335	0.0367	0.1037	0.0315
	(0.0106)	*(0.0338)*	*(0.0560)*	*(0.0559)*	*(0.0535)*
University	0.1789	0.2675	0.3302	0.4048	0.3160
	(0.0888)	*(0.0822)*	*(0.1332)*	*(0.1270)*	*(0.1120)*
University·t	-0.0047	0.0996	0.1168	0.0762	0.0987
	(0.0066)	*(0.0245)*	*(0.0409)*	*(0.0263)*	*(0.0253)*
Experience	0.0244	0.0291	0.0411	0.0252	0.0303
	(0.0054)	*(0.0053)*	*(0.0063)*	*(0.0077)*	*(0.0057)*
Experience·t	0.0009	0.0002	0.0026	0.0033	0.0016
	(0.00048)	*(0.0001)*	*(0.0027)*	*(0.0025)*	*(0.0018)*
Experience^2	-0.0006	-0.0006	0.0009	0.0006	-0.0007
	(0.0002)	*(0.0001)*	*(0.0001)*	*(0.0002)*	*(0.0001)*
Experience^2·t	-0.00001	-0.000004	0.0001	0.0001	0.0000
	(0.000020)	*(0.00000)*	*(0.0001)*	*(0.0001)*	*(0.0000)*
Prague	-0.130	0.140	0.1629	0.0794	0.1667
	(0.046)	*(0.028)*	*(0.0575)*	*(0.0736)*	*(0.0364)*
Child benefits included	0.228	0.122	0.2099	0.0493	0.1126
	(0.040)	*(0.029)*	*(0.0597)*	*(0.0579)*	*(0.0400)*
Gross earnings	0.134	0.048	-0.0325	0.0828	0.0609
	(0.051)	*(0.044)*	*(0.0904)*	*(0.0779)*	*(0.0631)*
Gross earnings·t	0.004	0.002	0.0441	0.0085	0.0185
	(0.004)	*(0.001)*	*(0.0305)*	*(0.0264)*	*(0.0180)*

Table A.7 continued
Earnings Regressions; Time Varying Cofficients for Communism and Transition (Education in Levels)

Period	Communism	Transition			
Industry		All	State	Privatized	DeNovo
Mining & Quarrying	0.272	0.046	0.202	0.061	-0.206
	(0.055)	*(0.055)*	*(0.129)*	*(0.078)*	*(0.111)*
Construction	0.132	0.130	-0.0001	0.200	-0.001
	(0.052)	*(0.042)*	*(0.123)*	*(0.071)*	*(0.059)*
Wholesale and Retail	-0.054	0.119	-	0.180	-0.014
	(0.059)	*(0.044)*	-	*(0.097)*	*(0.061)*
Public Adm., Edu, Health	0.083	0.055	0.129	-	-
	(0.053)	*(0.047)*	*(0.113)*	-	-
Finance, Insur., R. Estate	0.083	0.095	0.188	0.103	-0.252
	(0.053)	*(0.055)*	*(0.167)*	*(0.095)*	*(0.137)*
Transport, Telecom.	0.090	0.025	0.118	0.211	0.058
	(0.063)	*(0.041)*	*(0.123)*	*(0.098)*	*(0.087)*
Manuf.-Food, Textile,	-0.002	0.025	0.009	0.026	-0.037
	(0.044)	*(0.041)*	*(0.116)*	*(0.063)*	*(0.059)*
Manuf.-Machinery	-0.017	0.087	0.128	0.114	0.026
	(0.049)	*(0.048)*	*(0.130)*	*(0.071)*	*(0.072)*
Not known	0.068	0.182	0.030	0.091	0.136
	(0.131)	*(0.099)*	*(0.148)*	*(0.130)*	*(0.149)*
Constant	8.063	7.959	7.719	7.864	*(8.148)*
	(0.084)	*(0.078)*	*(0.151)*	*(0.138)*	*(0.078)*
adj.R^2	***0.172***	***0.296***	***0.344***	***0.339***	***0.27***
nobs	***1285***	***2107***	***483***	***413***	***1045***

Base= Jr. H.S. graduates working outside Prague in agriculture, whose earnings are net of tax and child benefits.

Chapter II

Returns to Job Mobility

1. Introduction

In Chapter I we presented results of cross-sectional analysis of returns to human capital. The advantage of cross-sectional results is their straightforward comparability with results found in many other studies from different countries and different years. However, the cross-sectional analysis told us little about the transition process leading to wage structures observed in mid 1990s. Newly emerging wage structures was not only a an outcome of changes in relative prices of labor and human capital but also an outcome of high labor market mobility experienced by individual workers. The relationship between job mobility and wages has been a long-standing area of interest in economics (see e.g., Light and McGarry, 1998, and Farber, 1999, for recent surveys). The underlying theoretical concepts and explanations range from human capital to job-matching and sectoral shift models. Whereas the literature is based almost entirely on U.S. data, the context of the sudden elimination of central planning and the emergence of new private firms during the transition to a market economy constitutes a much more dramatic laboratory in which to study this mobility-wage relationship. In particular, the extent of labor mobility during the transition has been high. Within four to five years after the start of transition over one-half of the labor force moved from the old state sector to the sector composed of newly created private firms.[1] This environment provides researchers with the opportunity to observe a large number and variety of moves in addition to workers who stay in their old job (stayers). Moreover, the magnitude of changes in the wages should be greater for a given period as compared to the U.S. since in transition the wage distribution expanded from a compressed, centrally determined wage to market wages determined by firms in new sectors competing for workers.[2]

The literature on wage changes of workers in transition economies is still thin, largely because there are few panel data sets that contain information on individuals' wages and job/firm characteristics.[3] In particular, to the best of our knowledge, no study examines the interactions between job and wage changes and the new emerging private sector. However, several studies examine other, related issues (Boeri and Flinn (1997), Burda and Mertens (1998), Hunt (1998), Noorkôiv, et al. (1998) and Sabirianova (2000).[4]

The goal of this hapter is to analyze the wage changes of individuals between the last year of the communist regime (1989) and seven years after the beginning of the transition to a market economy (1996). We use the same data as in Chapter I, a 1989-1996 panel of individuals working in the Czech Republic, to assess whether or not these workers benefited from job mobility. We also evaluate whether the gain/loss from mobility varied with the type of separation (voluntary and involuntary) and type of firm to which they moved, as characterized by industrial sector and whether the firm is in the new sector (private firms created after 1990) or the old sector (enterprises in existence before 1990, whether state owned enterprises or privatized, and the public sector).[5]

The outline of the chapter is as follows. In Section 2, we briefly review the principal theoretical models driving the relevant literature. We describe our methodology in Section 3 and our data in Section 4. We then discuss the results from the various specifications of the wage change equations in Section 5 and conclude the chapter in Section 6.

2. Theoretical Background

A number of models have been developed since the early Blumen, Kogan and McCarthy (1955) "stayer-mover" model which emphasized underlying personal characteristics of workers as driving mobility and wage differences associated with job changes. The model predicts that mobility is negatively related to wages because high-productivity workers avoid turnover, while low-productivity workers undergo persistent mobility. People are consistently high or low mobility individuals over their entire life cycle.

The human capital models stress the importance of skills and the firm's valuation of these skills in explaining job mobility (see e.g., Bartel and Borjas, 1981; Farber, 1994; Mincer, 1981; Neal, 1995; Topel, 1986, 1991). Whereas the earlier human capital literature focuses on explaining mobility over the life-cycle (i.e., why it is higher earlier in the life-cycle than later), the more recent literature is concerned with explaining why job separation is higher in the early part of a job and it declines as tenure increases. The central idea is that as long as workers are being paid for their accumulated firm specific human capital, they will not quit. Moreover, the more firm specific human capital the worker has obtained, the more costly it is for the worker to quit. Conversely, firms will not layoff a worker if his/her productivity (including firm specific human capital) is equal to or above the wage. Much of this empirical literature has been mired by the fact that firm specific capital is not observed and tenure is not necessarily a good proxy for it.[6] We focus instead on mobility over the life-cycle.

The job-matching models, started by Jovanovic (1979) in his seminal paper, focus on the fact that the worker-firm match varies across firms and the quality of this match is not known ex-ante. The quality of the match is revealed over time as tenure accumulates. In the earlier models, there is no randomness in the wage offer distribution or shocks to productivity or demand, therefore all turnover is generated simply by the revelation of the match. Separations occur when match quality is poor and hence those who stay are better matched and earn higher wages. Later, search models (e.g., Burdett, 1978), introduced the importance of the distribution of wages, and shifts in these distributions in explaining voluntary mobility.

The sectoral shifts models (e.g., Jovanovic and Moffit, 1990) focus on mobility arising from of shifts in sectoral demand as the economy is restructuring, important for the study of economies in transition. Although Jovanovic and Moffit (1990) found with U.S. data that wages of movers tend to be lower than wages of non-movers, and that most of the mobility and change in wages was derived from poor matching rather than sectoral shifts, we expect the opposite in transition economies. If the transition is a process in which the creation of new productive private sector firms generates 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. Methodology

We analyze the wage changes of individuals between the last year of the communist regime (1989) and seven years after the beginning of the transition to a market economy (1996). We assess whether or not these workers benefited from job mobility, and evaluate whether the gain/loss from mobility varied with the type of separation (voluntary and involuntary) and type of firm to which they moved, as characterized by industrial sector and whether a new private sector firm (new sector) or a state or privatized firm (old sector).

3.1 Basic 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:

$$\ln W_{96} = X' a_{96} + Y_{96}' b_{96} + e_{96,} \qquad (1)$$

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

$$\ln W_{89} = X^{'}(a_{96} - \Delta a) + Y_{89}^{'}(b_{96} - \Delta b) + (e_{96} - \Delta e), \tag{2}$$

where $Da = a_{96} - a_{89}$ and $Db = b_{96} - b_{89}$ are the changes in the payoffs to the explanatory variables over the period and $De = (e_{96} - e_{89})$ is the difference 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:

$$\ln W_{96} - \ln W_{89} = X^{'}\Delta a + Y_{89}^{'}\Delta b + (Y_{96}^{'} - Y_{89}^{'})b_{96} + \Delta e. \tag{3}$$

For the time-invariant explanatory variable *X* we hence obtain estimates of the changes in payoffs (*Da*) between 1989 and 1996, while for the time-varying characteristics Y we generate estimates of the coefficients for 1996 (b_{96}) and of the changes in payoffs (*Db*) 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*). 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 to each of these factors (i.e., *Da* or *Db* in equation 3). The coefficient on the change in experience estimates the 1996 return to experience (i.e., b_{96})

Local demand conditions are captured with a change in the district unemployment rates from 1989 to 1996 and a time invariant dummy for Prague. Since unemployment rates were zero in 1989, this variable is effectively the district unemployment rate in 1996. However, its coefficient indicates the relative effect of changes in local demand conditions on changes in wages (capturing the wage curve hypothesis of Blanchflower and Oswald, 1995). The coefficient on the dummy for Prague (Da) indicates whether wages of people residing in the capital city relative to the wages of those living in other areas of the country changed over this 1989-1996 period.

We estimate the wage effect of changes in job characteristics in several ways. We start by estimating the effect of simply changing a job vs. staying on the same job with a dummy variable. This corresponds to the basic "stayer-mover" model of Blumen et al. (1955). We extend this framework by examining whether job changes that also involved changing certain attributes of the job make a difference. We first test whether the wage effect depends on whether a person changes her industrial sector when she changes her job (we account for eight sectors). Derek Neal (1995) argues that workers receive compensation for skills that are specific to their industry, i.e., industry-specific human capital. Hence a worker who changes industry may not experience as much of a wage gain as a worker who changes job without changing industry, *ceteris paribus*. However, we also recognize that the industrial wage structure was changing dramatically over this period.[7] In line with the importance of the growth of new private economic activity in transition countries, we next estimate specifications that show the impact on wages of moving from a job in the old sector to a job in the new sector vs. changing jobs but staying in the old sector.

We recognize that patterns of wage change may differ for individuals who changed jobs by voluntarily quitting their job and those who were involuntarily laid-off. 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 the Endogeneity of the Decision to Quit

The fact that an individual chooses to leave the labor force or to quit and change jobs creates a problem of selectivity bias in estimating the overall wage change and impact of quitting on the wage change. The problem is one where we want to estimate the net benefit (wage change) of quitting or staying, however we only observe wage changes for those who quit and for those who stay, and have not left the labor force. The net benefit is the market wage an individual could obtain if he/she quit and change jobs compared with what he could obtain if he stayed in the job.

Formally, abstracting from the time varying vs. time non-varying nature of the variables, let the wage growth for an individual who quits be:

$$\Delta \ln W_q = m_q(X) + e_q\,, \qquad (4)$$

and the wage growth if the same individual stays

$$\Delta \ln W_s = m_s(X) + e_s\,. \qquad (5)$$

The gain from quiting for this individual is

$$\Delta \ln W_q = \Delta \ln W_s\,, \qquad (6)$$

and the average gain for individuals with X is

$$b(X) = m_q(X) - m_s(X)\,. \qquad (7)$$

Let q is the probability of quitting, then the observed wage growth is

$$\begin{aligned} \Delta \ln W &= q\Delta \ln W_q + (1-q)\Delta \ln W_s \\ &= m_s(X) + b(X)q + [e_s + (e_q - e_s)q]. \end{aligned} \qquad (8)$$

If $(e_q - e_s) \perp Z \equiv (X, X_e)$ for some instruments X_e, then

$$\begin{aligned} E[e_s + (e_q - e_s)q|Z] &= E(e_q - e_s|q = 1, Z)E(q|Z) \\ &= j\,(Z)E(q|Z) \end{aligned} \qquad (9)$$

so that

$$\Delta \ln W = m_s(X) + [b(X) + j\,(Z)]q + u\,, \qquad (10)$$

where $E(? \mid Z) = 0$. Note that $E[\mu_s(X)]$ is the average gain for stayers, $E[b(X)]$ is the overall average gain, and $E[b(X) + f(Z)]$ is the average gain for movers.

If given Z, q is independent of $(e_q - e_s)$ then $f(Z) = 0$ and $E(? \mid q, Z) = 0$ so that q is exogenous and $b(X)$ can be estimated by OLS. If the gain for individuals with X is homogeneous $(e_q - e_s) = 0$ and $f(Z) = 0$. In this case, subject to identification, $b(X)$ can e estimated by IV methods. If the gain is heterogeneous, but

$$\Pr(q = 1 | Z) = \Pr(Zp + u > 0 | Z) = \Pr(u > -pZ) \tag{11}$$

And u, e_q, $e_s \mid Z \sim N(0, S)$ then,

$$j(Z) = (s_{qu} - s_{su})\, l(Zp)\,. \tag{12}$$

Hence we will estimate and compare the returns using both IV and selectivity correction methods.

3.2.1 Correcting for Selectivity Bias

The decision to quit can be modeled several ways. Following the human capital and job search models (e.g., Jovanovic, 1979 and Burdett, 1978) a simple model of the worker's decision to quit is developed by Farber (1999) 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:

$$W = W_a + lY, \tag{13}$$

where l 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 $l > 0$.

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, dependent on worker's general ability, and with dispersion that reflects cross-job variation in the worker's job-specific ability. Hence, the wage offer can be expressed as $W_o = W_a + F$, where F 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 $W_a + F > W_a + lY$, or $F > lY$. 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 and hence, the more likely a worker will quit. Search theory also predicts that the arrival rate of wage offers is higher in tighter labor markets, so workers are more likely to quit in economic upswings. 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. 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 to capture risk: dummy variables for, marital status, one child, and two or more children; the log of per capita household income (in 1996). The vector includes the following variables to capture human capital and search characteristics: gender education, age, the log of the individual's wage relative to the mean wage of all individuals with the same level of schooling (in 1996)[8] and dummy variables for Prague and for the three sectors where employment declined dramatically (agriculture, mining and utilities, and heavy manufacturing).

We estimate a 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 determine wage levels (marital status, number of children and household per capita income), we are confident that we have satisfactory identification.

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 employed individuals are simultaneously deciding on whether to stay in the current job, quit and take a job offer in the new sector, or 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 two models: a) a standard multinomial probit (rather than an ordered probit) for stay vs. i) quit and take a job in the new sector or ii) 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 both of 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.

4. Data

The data we use are already described in Chapter I, Section 1.[10] Selecting those individuals who had a job in both periods, and who were not likely to have retried during the seven year period between 1989 and 1996, yielded a sample of 3,072 individuals for whom we could create panel data on their wages and other job characteristics. After further cleaning the data to exclude individuals who held only part-time jobs and more than one job or had missing wage, we are left with a sample of 2,343 individuals.

The means and standard deviations of the main variables are presented in Table 1. As may be seen from the values in row 2, the 1989-96 change in real full-time wages was negative for all the principal groups of individuals.[11] (This variable is defined as nominal monthly salaries of full-time workers, net of tax, deflated by the consumer price index.) For the entire sample, the real wages 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 laid off

Table 1: Summary Statistics

Variable Name	All Sample		Stayers		Quit Job		Laid Off	
	Mean	(St. Dev.)	Mean	(St. Dev.)	Mean	(St. Dev.)	Mean	(St. Dev.)
(W96-W89)/W89 in '89 Kcs	-0.265	0.484	-0.312	0.411	-0.213	0.540	-0.288	0.495
Log of relative wage[u]	0.003	0.429	0.042	0.418	-0.019	0.428	-0.045	0.444
Relative wage[u]	1.101	0.559	1.133	0.497	1.083	0.624	1.059	0.575
Log per cap HH inc.	8.463	0.438	8.465	0.396	8.473	0.473	8.416	0.473
Age	35.984	9.127	37.687	8.222	33.714	9.639	36.513	8.803
Women (dummy)	0.438	0.496	0.438	0.496	0.433	0.496	0.479	0.501
Education (years)	12.544	2.417	12.435	2.376	12.727	2.437	12.188	2.455
Experience (years)	16.399	9.415	18.185	8.734	13.983	9.740	17.270	8.907
Experience in 1989-1996	7.889	0.422	8.000	0.000	7.806	0.578	7.694	0.556
Prague	0.111	0.314	0.083	0.276	0.103	0.304	0.104	0.306
Married	0.806	0.395	0.832	0.374	0.787	0.410	0.771	0.421
One child in home	0.201	0.401	0.186	0.389	0.226	0.418	0.192	0.394
Two+ children in home	0.188	0.391	0.159	0.366	0.241	0.428	0.121	0.327
Sector of Job in 1989:								
Agriculture	0.089	0.285	0.063	0.243	0.111	0.315	0.129	0.336
Mining and Utilites	0.064	0.245	0.082	0.275	0.046	0.209	0.096	0.295
Construction	0.075	0.264	0.062	0.241	0.089	0.286	0.079	0.271
Light Manufacturing[a]	0.153	0.360	0.152	0.360	0.142	0.349	0.163	0.370
Heavy Manufacturing[υ]	0.201	0.401	0.234	0.423	0.182	0.386	0.175	0.381
Trade, Restaurants and Hotels	0.091	0.287	0.072	0.259	0.104	0.306	0.108	0.311
Fin., Real Est., Trans.&Comm.	0.120	0.326	0.114	0.318	0.130	0.336	0.092	0.289
Public Administration	0.206	0.404	0.220	0.415	0.195	0.396	0.158	0.366
Change in sector of Job:								
Agriculture*	-0.036	0.220	0.000	0.000	-0.065	0.305	-0.083	0.292
Mining and Utilites*	-0.014	0.171	0.000	0.000	-0.017	0.216	-0.067	0.310
Construction *	0.007	0.228	0.000	0.000	0.014	0.314	0.017	0.342
Light Manufacturing*[a]	-0.005	0.285	0.000	0.000	-0.013	0.385	0.025	0.408
Heavy Manufacturing*[υ]	-0.021	0.299	0.000	0.000	-0.042	0.431	-0.033	0.376
Trade, Restaurants and Hotels*	0.037	0.293	0.000	0.000	0.067	0.398	0.071	0.418
Fin., Real Est., Trans.&Comm.*	0.020	0.297	0.000	0.000	0.032	0.419	0.075	0.392
Public Administration*	-0.039	0.194	0.000	0.000	-0.073	0.261	-0.083	0.277

Table 1: Summary Statistics, continued

	All Sample		Stayers		Quit Job		Laid Off	
Variable Name	Mean	(St. Dev.)	Mean	(St. Dev.)	Mean	(St. Dev.)	Mean	(St. Dev.)
Decline Sector[c]	0.153	0.360	0.145	0.352	0.157	0.364	0.225	0.418
District unemp. rate in 1996 (%)	-3.712	0.735	-3.643	0.692	-3.753	0.703	-3.651	0.755
Avg. district vacancy rate (%)	-4.255	0.404	-4.287	0.415	-4.225	0.377	-4.253	0.374
Dummy for stayed	0.477	0.500	1.000	0.000	0.000	0.000	0.000	0.000
Dummy for changed jobs	0.523	0.500	0.000	0.000	1.000	0.000	1.000	0.000
Held two jobs	0.164	0.371	0.000	0.000	0.401	0.490	0.000	0.000
Held three jobs	0.292	0.455	0.000	0.000	0.460	0.499	0.896	0.306
Held four+ jobs	0.068	0.251	0.000	0.000	0.139	0.346	0.104	0.306
Dummy if unemployed	0.078	0.268	0.000	0.000	0.110	0.313	0.317	0.466
Unemp. Duration	1.332	5.059	0.000	0.000	2.325	6.932	3.667	6.676
Old Sector	0.684	0.465	1.000	0.000	0.413	0.493	0.379	0.486
New Sector	0.302	0.459	0.000	0.000	0.560	0.497	0.596	0.492
No. of Observations	2277		973		872		240	

Table 1: Summary Statistics, continued

Variable Name	Old State Sector Mean	(St. Dev.)	New Private Sector Mean	(St. Dev.)
(W96-W89)/W89 in '89 Kcs	-0.278	0.463	-0.187	0.571
Log of relative wage[u]	-0.068	0.402	0.007	0.443
Relative wage[u]	1.015	0.456	1.119	0.678
Log per cap HH inc.	8.438	0.449	8.487	0.491
Age	34.340	10.327	34.369	9.135
Women (dummy)	0.539	0.499	0.370	0.483
Education (years)	12.555	2.489	12.672	2.422
Experience (years)	14.809	10.459	14.658	9.238
Experience in 1989-1996	7.738	0.681	7.825	0.462
Prague	0.100	0.301	0.124	0.330
Married	0.756	0.430	0.808	0.394
One child in home	0.201	0.401	0.222	0.416
Two+ children in home	0.217	0.413	0.219	0.414
Sector of Job in 1989:				
Agriculture	0.101	0.301	0.121	0.327
Mining and Utilites	0.064	0.244	0.046	0.209
Construction	0.068	0.252	0.103	0.304
Light Manufacturing[α]	0.146	0.353	0.163	0.369
Heavy Manufacturing[υ]	0.187	0.390	0.174	0.379
Trade, Restaurants and Hotels	0.086	0.281	0.117	0.322
Fin., Real Est., Trans.&Comm.	0.117	0.322	0.128	0.335
Public Administration	0.232	0.423	0.148	0.356
Change in sector of Job:				
Agriculture*	-0.048	0.273	-0.081	0.316
Mining and Utilites*	-0.021	0.257	-0.029	0.213
Construction *	-0.029	0.256	0.040	0.346
Light Manufacturing*[α]	-0.012	0.375	-0.013	0.403
Heavy Manufacturing*[υ]	-0.033	0.426	-0.046	0.410
Trade, Restaurants and Hotels*	-0.048	0.280	0.157	0.439
Fin., Real Est., Trans.&Comm.*	0.070	0.401	0.017	0.412
Public Administration*	-0.043	0.204	-0.091	0.287

Table 1: Summary Statistics, continued

Variable Name	Old State Sector Mean	Old State Sector (St. Dev.)	New Private Sector Mean	New Private Sector (St. Dev.)
Decline Sector[c]	0.164	0.371	0.167	0.373
District unemp. rate in 1996 (%)	-3.675	0.721	-3.796	0.745
Avg. district vacancy rate (%)	-4.240	0.380	-4.225	0.379
Dummy for stayed	0.000	0.000	0.000	0.000
Dummy for changed jobs	1.000	0.000	1.000	0.000
Held two jobs	0.379	0.486	0.265	0.442
Held three jobs	0.512	0.500	0.591	0.492
Held four+ jobs	0.109	0.311	0.144	0.351
Dummy if unemployed	0.182	0.387	0.121	0.326
Unemp. Duration	3.143	8.167	2.094	5.549
Old Sector	1.000	0.000	0.000	0.000
New Sector	0.000	0.000	1.000	0.000
No. of Observations	488		702	

*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 = w_i/(average wage of the individual's education group), where education groups are defined as: apprentice, high school eduation without CGE exam; high school education with CGE exam, university and higher.

workers 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 enterprises) lost 21.6 percent, as compared to a smaller loss of 5.1 percent for those who moved to the newly formed private firms (including self-employed). 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 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 size and vigorous growth of the 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 (defined as loosing more than 10% of their workforce over the 1989-1996 period).

5. Empirical Results

5.1. An OLS 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 1989-1996 logarithmic change in wage to five sets of explanatory variables.

In line with the model presented in Section 3, 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, 1996 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.[12]

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 1989-1996 period. As may be seen from Table 2, this result is robust to differences in specification. Experience in 1989 captures the effect

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 (1996-89)	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[a]	0.007	0.062	-0.050	0.062	-0.009	0.062	-0.006	0.062	-0.031	0.062
Mining and Utilites[a]	0.177 *	0.071	0.123	0.071	0.160	0.072	0.156	0.072	0.139	0.071
Construction[a]	0.027	0.059	-0.038	0.061	-0.001	0.060	0.001	0.060	-0.027	0.061
Light Manufacturing[a,b]	-0.015	0.055	-0.068	0.055	-0.034	0.055	-0.029	0.055	-0.048	0.055
Heavy Manufacturing[a,c]	-0.103 *	0.053	-0.156 *	0.053	-0.121 *	0.053	-0.119 *	0.053	-0.140 *	0.053
Trade, Restaurants and Hotels[a]	-0.001	0.054	-0.074	0.056	-0.035	0.056	-0.034	0.056	-0.070	0.056
Finance, R. Estate, Trans. Comm.[a]	0.084	0.053	0.042	0.053	0.071	0.053	0.074	0.053	0.057	0.053
Public Administration[a]	-0.004	0.072	-0.069	0.072	-0.014	0.072	-0.017	0.072	-0.040	0.072

Table 2: continued

	(1)		(2)		(3)		(4)		(5)	
Variable Name	Coeff.	St.Error	Coeff.	St.Error	Coeff.	St.Error	Coeff.	St.Error	Coeff.	St.Error
Dummy for Changed Job	0.089 *	0.022								
Dummy Changed Job; in New Sector			0.094 *	0.025						
Dummy for Changed Job and in Old			-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; in New Sector in '96									0.134 *	0.026
Quit; in Old Sector in '96									0.038	0.030
Laid-off; in New Sector in '96									0.112 *	0.043
Laid-off; in Old Sector in '96									-0.112 *	0.052
N	2276		2276		2276		2276		2276	
Adjusted R^2	0.091		0.091		0.093		0.093		0.096	

[a]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.

[b] Light manufacturing includes food, textiles, paper and chemicals, and other manufacturing

[c] Heavy manufacturing includes machinery, metals and equipment.

[*]Statistically Significant at 10% confidence level.

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 a priori expect 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 and women’s 1989-1996 wage change is not statistically significant *ceteris paribus*.

The variables capturing the effect of local demand conditions had the expected signs. Prague residence yields a 10-13 percent wage gain over those living outside Prague in all specifications. People living in a district with higher unemployment rate 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 indicate that individuals leaving (entering) mining and utilities lost (gained), while those leaving (entering) heavy manufacturing gained (lost). Other inter-sectoral 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*.[13] In column 2 we present the effects of moving into a new private firms and moving into an old sector firm, relative to 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.[14] 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. The wage change of workers who quit and move to the old sector is not statistically different from the wage change of stayers (in the old sector). However, workers who are laid off and end up in the old public sector suffer an 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.

5.2. Determinants of Quits

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. 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 Table 3. 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, being more educated and being in a declining sector. Finally, gender and Prague residence 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.[15] 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

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	Marginal	St. Error	Marginal	St. Error	Marginal	St. Error	Marginal	St. Error
Women	-0.015	0.026	-0.105 *	0.028	0.078 *	0.018	-0.074	0.079
Married	-0.067 *	0.031	-0.007	0.032	-0.049	0.027	0.061	0.085
1 child	0.052	0.032	0.048	0.031	-0.007	0.026	0.101	0.092
2+ children	0.066	0.036	0.028	0.035	0.031	0.028	0.119	0.115
Per capita HH Income	0.078 *	0.030	0.068 *	0.030	0.017	0.024	0.178	0.082
Education (yrs.)	0.008	0.005	0.008	0.005	0.001	0.004	-0.007	0.015
Age	-0.011 *	0.002	-0.009 *	0.002	-0.003 *	0.001	0.006	0.005
Relative wage[a]	-0.069 *	0.030	-0.016	0.030	-0.050 *	0.024	0.131	0.094
Prague	-0.003	0.045	-0.021	0.044	0.008	0.034	0.100	0.133
District Vacancy Rate	0.109 *	0.032	0.054	0.032	0.062 *	0.027	0.145	0.105
Declining Sector[b]	0.054	0.033	0.012	0.033	0.035	0.026	-0.068	0.087
Pseudo R2	0.048		0.0438				0.064	
N	1905		1876				218	

a Relative wage = w_i/(average wage of the individual's education group), where education groups are defined as: apprentice, high school eduation without CGE exam; high school education with CGE exam, university and higher.
b Declining sectors are agriculture, mining and utilities and heavy manufacturing.
* Statistically Significant at 10% confidence level.
NB: Marginal effects are evaluated for single men without children, living outside of Prague, not employed in a declining industry

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.

5.3. Determinants of Quitting and Selecting a Job in the New or Old Sector

In panel (b) of Table 3 we present estimates of a probit model that postulates that an individual makes the following decisions simultaneously: a) to quit and take a new sector job vs. stay and b) quit and take an old sector job vs. stay. 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.

As may be seen from Table 3, row 1 of panel (b), women are less likely than men to quit and move into a new sector job but are more likely than men to quit and move into an old sector job. Married individuals are indifferent between quitting for a new sector job and staying but are less likely to quit for an old sector job. Neither the presence of children in the household nor the worker's level of education seem to affect the decision making process. As expected, the probability of quitting for a new sector job is positively related to higher per capita household income, whereas this variable does not affect the probability of quitting for an old sector job. Older people are less likely than younger workers to quit for either sector. It is very interesting to see that individuals on the high end of the wage distribution are less likely to quit for an old sector job but are not deterred from quitting for a new sector job. Hence, this last finding combined with the findings on gender, per capita income, education and age, seems to point to that men, younger and more able individuals, living in households with relatively high income are those who are taking the decision/challenge of quitting for a new sector job. Finally, the results in panel (b) indicate that among the variables capturing demand conditions, only the coefficient on the district vacancy rate is significant, indicating that people who live in districts with tighter labor markets are more likely to quit for a job in either sector. Neither living in Prague nor having a job in a declining sector affect these probabilities.

5.4. Selection of Job in Old vs. New Sector for those who were Laid-Off

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 consistent with the finding in panel (b) and intuitively acceptable, since working in the new sector is riskier, especially for the self-employed. 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. Of course, the statistically insignificant estimates could in part be also brought about by the fact that they are based on only 218 observations.

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

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 four groups: those who quit for jobs in the new sector, quit for jobs in the old sector, were laid off and found a job in the new sector, and were laid off and found a job in 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.

5.5.1. *Stayers, Quitters, Laid-off*

We begin with the determinants of wages changes for stayers and quitters. As may be seen from Table 4, three 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 the rising rate of return to education during the transition we found in Chapter I.

Whereas work experience accumulated as of 1989 has a positive effect on wages of stayers and quitters, 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. The gender effect is also not significant for all three categories of workers, as it was in the more aggregated regressions of Table 2.

The results for the demand variables were mixed. 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. We noted above that changes in the local demand conditions (proxied by the 1996 district unemployment rates) did not affect wage changes for individuals in all three groups (if using the wage equations corrected for selectivity bias). The wage effects of changing of industry of the job are generally not significant except for the following: Being laid off from mining and utilities is associated with a major

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					
	Stayed		Quit		Stayed		Quit		Laid-off	
Variable Name	Coeff.	St. Error	Coeff.	St. Error	Coeff.	St. Error	Coeff.	St. Error	Coeff.	St. Error
Education	0.040 *	0.005	0.036 *	0.007	0.036 *	0.006	0.034 *	0.007	0.027 *	0.013
Experience in 1989	0.029 *	0.003	0.029 *	0.004	0.000	0.002	-0.007 *	0.002	-0.007 *	0.004
Change in Experience (1996-89	-	-	0.046	0.029	-	-	0.029	0.030	0.072	0.054
Women	0.020	0.024	0.023	0.035	0.010	0.026	0.027	0.036	0.070	0.062
Prague	0.055	0.061	0.092	0.080	-0.004	0.066	0.145	0.084	0.324 *	0.150
District Unemp. Rate in 1996	0.057 *	0.026	0.020	0.036	-0.057 *	0.026	-0.058	0.037	0.019	0.062
Agriculture[a]	-	-	0.010	0.077	-	-	-0.074	0.081	0.003	0.148
Mining and Utilites[a]	-	-	0.134	0.093	-	-	0.054	0.098	0.279 *	0.148
Construction[a]	-	-	-0.090	0.075	-	-	-0.080	0.079	0.050	0.138
Light Manufacturing[a,b]	-	-	-0.084	0.069	-	-	-0.079	0.073	-0.070	0.130
Heavy Manufacturing[a,c]	-	-	-0.167 *	0.064	-	-	-0.167 *	0.067	-0.162	0.136
Trade, Restaurants, Hotels[a]	-	-	-0.085	0.069	-	-	-0.100	0.072	-0.066	0.136
Finance, R.Estate, Trans. Comn	-	-	-0.002	0.065	-	-	0.004	0.068	0.086	0.137
Public Administration[a]	-	-	-0.090	0.090	-	-	-0.083	0.094	-0.018	0.167
New Sector Job	-	-	0.097 *	0.036	-	-	0.084 *	0.038	0.210 *	0.068
Lambda	-1.655 *	0.127	-1.802 *	0.184	-	-	-	-	-	-
Constant	-2.548 *	0.156	-0.180	0.296	-1.191 *	0.125	-1.264 *	0.289	-1.363 *	0.473
N	973		872		973		872		240	
Adjusted R2	0.1937		0.1781		0.053		0.086		0.149	

a 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.

b Light manufacturing includes food, textiles, paper and chemicals, and other manufacturing

c Heavy manufacturing includes machinery, metals and equipment.

* Statistically Significant at 10% confidence level.

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.

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

The estimates in Table 5a give the selectivity corrected effects of our 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. Table 5b presents the coefficients for the same variables without correcting for selectivity bias.

As may be seen from Table 5a, education continues to have a positive effect on the wage change of the first three groups, but it has no effect on the wage changes of those who were laid off and moved to the old sector. The coefficient is not affected by correction for selectivity bias.

The effect of 1989 experience continues to be positive for quitters but moves from being negative (Table 4) to not significantly different from zero for laid off individuals, irrespective of whether they moved to the new or old sector. 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.

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							
	Quit for New		Quit for Old		Laid Off for New		Laid Off for Old	
Variable Name	Coeff.	St. Error	Coeff.	St. Error	Coeff.	St. Error	Coeff.	St. Error
Education	0.031 *	0.011	0.029 *	0.009	0.045 *	0.017	-0.021	0.021
Experience in 1989	0.031 *	0.007	0.008 *	0.003	-0.001	0.005	-0.005	0.007
Change in Experience (1996-198	0.026	0.056	0.037	0.031	0.048	0.074	0.141	0.079
Women	0.408 *	0.092	-0.212 *	0.071	-0.189 *	0.099	0.183	0.108
Prague	0.169	0.120	0.144	0.112	0.306	0.200	0.808 *	0.275
District Unemp. Rate in 1996	-0.025	0.053	0.052	0.050	-0.052	0.085	0.028	0.094
Agriculture*	-0.148	0.140	0.109	0.095	0.065	0.291	-0.281	0.180
Mining and Utilites*	0.172	0.191	0.110	0.103	0.432	0.294	0.185	0.201
Construction *	-0.156	0.131	-0.017	0.100	0.168	0.275	-0.045	0.184
Light Manufacturing*[a]	-0.181	0.129	0.032	0.083	0.082	0.276	-0.208	0.156
Heavy Manufacturing*[b]	-0.195	0.126	-0.162 *	0.071	-0.163	0.284	0.035	0.161
Trade, Restaruants and Hotels*	-0.123	0.124	-0.094	0.095	0.050	0.275	-0.201	0.180
Finance, Real Estate, Trans.&Co	-0.018	0.124	-0.025	0.075	0.181	0.291	-0.002	0.151
Public Administration*	-0.173	0.151	-0.057	0.130	0.066	0.312	-0.151	0.237
Correction for Selectivity Bias	-2.017 *	0.343	-1.218 *	0.223	0.798 *	0.269	0.142	0.294
constant	0.742	0.622	0.742	0.444	-2.021 *	0.708	-1.355 *	0.681
N	488		360		123		90	
Adjusted R^2	0.136		0.167		0.241		0.166	

a 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.

b Light manufacturing includes food, textiles, paper and chemicals, and other manufacturing

c Heavy manufacturing includes machinery, metals and equipment.

* Statistically Significant at 10% confidence level.

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							
	Quit for New		Quit for Old		Laid Off for New		Laid Off for Old	
Variable Name	Coeff.	St. Error	Coeff.	St. Error	Coeff.	St. Error	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[a]	-0.217	0.145	0.083	0.099	0.043	0.301	-0.296 *	0.176
Mining and Utilites[a]	0.092	0.197	0.011	0.105	0.416	0.305	0.176	0.199
Construction[a]	-0.161	0.136	-0.055	0.104	0.193	0.285	-0.025	0.179
Light Manufacturing[a,b]	-0.203	0.133	0.028	0.086	0.091	0.286	-0.205	0.155
Heavy Manufacturing[a,c]	-0.213	0.131	-0.175 *	0.074	-0.184	0.294	0.037	0.160
Trade, Restaurants and Hotels[a]	-0.151	0.129	-0.149	0.099	0.109	0.284	-0.212	0.177
Finance, Real Estate, Trans.&Comm.[a]	-0.028	0.129	-0.033	0.078	0.212	0.301	-0.002	0.150
Public Administration[a]	-0.201	0.157	-0.006	0.135	0.008	0.322	-0.148	0.236
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 R^2	0.075		0.098		0.186		0.174	

a 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.

b Light manufacturing includes food, textiles, paper and chemicals, and other manufacturing

c Heavy manufacturing includes machinery, metals and equipment.

* Statistically Significant at 10% confidence level.

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 21.2 percent lower wage gain than comparable men. We find hence 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 5a. We learned earlier that among those laid off and finding a job in the old sector, being laid-off in Prague had its advantages over being laid off in the rest of the country. In Table 5a, we find that people who live in Prague and who are laid-off and find a job in the old sector gain 80.8 percent more than their counterparts living outside of Prague. Given that Prague is the seat of the government, this is not surprising.

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.

Finally, in Table 6 we compare the relative wage changes for quitters, laid-off and stayers that appear in the data as averages, with the predicted changes from the selectivity corrected regressions (Tables 4 and 5), and with those estimated in the simple wage equation (Table 1). We learn that stayers were worse off than the average person in the sample (by about 7%) but those who were laid-off and took a job in the old sector were the biggest loosers (with a 16% lower wage change than the average for the sample.) In comparing the results from Tables 4 and 5 with those in Table 1, we find that the findings are quite similar, with the predicted values from the selectivity equations being higher for some and lower for others. Those who quite are predicted to have 14.3% higher wages than those who stayed in Table 4 but 6.1% higher wages in Table 1. The wage changes of those who were laid off are not discernibly different from those who stayed in either estimation. People who quit and took a new sector job did the best of all people – predicted to have 13-19% higher wage growth than those who stayed. However those who quit and went to an old sector job had 4-9% higher wages than those who stayed. Finally, those where laid off and took a new sector job fare better (wage gains 8-11% higher than those who stayed) while those who were laid off and took an old sector job saw wage changes that were about 10-11% smaller than those who stayed.

6. Conclusions

Our analysis of job and wage changes of individuals working in the Czech Republic between 1989 (the last year of communism) and 1996 indicates that the relationship between job mobility and wage changes varies by type of separation and the sector of destination of movers. In particular, we draw the following conclusions:

The Czech workers experienced a significant fall in real wages between 1989 and 1993. Wage growth occurred in the economy 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. Our findings indicate that the decline was deepest for those who were laid off and took employment in the old sector followed by those who stayed in their original jobs. Those who quit and took a new sector job were experiencing the highest gains (smallest losses). 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.

We asked to what extent these findings were biased by selectivity and proceeded estimate wage equations which included Heckman's selectivity bias parameter based on estimated probits of the probability of quitting as well as the probabilities of quitting and taking an old sector job vs. a new sector job. We learned from the probit estimates that quitting behavior is found to be higher in tight local labor markets and among people who are younger, more able, in a better position to take risk (single, higher per capita household income) and who would gain the most from a move (lowest relative earnings in original job). Since layoff decisions are exogenous to the worker, we did not estimate a probability of layoff. On the other hand, both quitters and those who are laid off must decide what type of job to take. The estimates from modeling the choice of sector (new vs. old) indicates that individuals who quit and take a new sector job are more likely to be men, younger, single, with higher education and family income, and working in tight local labor markets. Those with higher relative earnings are not deterred from taking a new sector job. On the other hand those who quit and take a job in the old sector are more likely to be women, older, married and with lower relative wages. 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.

When we estimate 1989-96 wage change equations corrected for selectivity bias as individuals choose to quit or take a new vs. old sector job. We find that correction for selectivity bias strongly affects the coefficients on general human capital (i.e., experience) accumulated under communism. These coefficients become positive and significant when the OLS estimates were negative but not significant.

For quitters, laid off and stayers, wage changes were greater for the more educated and those with more experience. Relative changes in local demand conditions (district unemployment rate) do not matter. 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. As for changing industry we find that 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.

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Notes for Chapter II

[1] Jurajda and Terrell (2001).

[2] Svejnar (1999).

[3] The Labor Force Surveys, which only began after several years of transition had passed, do not collect wage data in most of these countries. Three exceptional data sets with good wage data are: Russian Longitudinal Monitoring Survey (RLMS), the Trexima data in the Czech and Slovak Republics and the Estonian Retrospective Employment History Survey (EREHS).

[4] Hunt's paper is the most similar one 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 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 from the beginning of the transition process. Burda and Mertens (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, they do not make use of the panel nature of the data and instead estimate cross-sectional wage regressions, of which there are several for the transition economies (see, e.g. Bird et al., 1994, Brainerd, 1998, Chase, 1998).

[5] We distinguish the new private sector from the privatized state owned enterprises because a growing literature indicates that this new sector appears to be the fountainhead of growth (see eg., Berkowitz and Cooper, 1997; Berkowitz and DeJong, 2001; Brixy and Kohaut, 1998, Jurajda and Terrell, 2002 and McMillan and Woodruff, 2001) while the privatized state firms continue to operate much like state-owned enterprises during the early years of transition (see eg., Roland, 2001).

[6] Furthermore, Farber (1999) argues that the theoretical link between wages and marginal product of labor exists for general human capital, as proxied by experience, but not for specific human capital, as proxied by tenure.

[7] See Münich et al. (1999) for an analysis of changes in industrial wage structure using this same data set.

[8] The relative wage variable is ln(w_i/average wage of the individual's education group), where education groups are define 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.)

[10] The questionnaire was designed by the authors and survey was carried out by a private firm, MEDIAN. See Münich et al. (1997) for a detailed description of the sample design and the characteristics of the sample.

[11] Overall, real consumer wages did not reach their 1989 level until 1998 (cite source).

[12] 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.

[13] 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.

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

[15] We would have preferred to use a tenure variable in the probit, but we did not obtain this information from the 1989 job.

Appendix tables for Chapter II

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 R^2	0.048		0.044				0.064	
N	1905		1876				218	

Chapter III

Responses of Schools to the Demand for Human Capital

1. Introduction

In previous chapters, we explored large scale and scope of changes in returns to human capital and education in particular experienced during first six transitional years. The changes were driven not only by changes in relative prices of skills but also by a large scope of job mobility realized by the workforce. The analysis focused only on the sub-population of individuals who had acquired formal education already during decades of the communist regime. The long-term changes in wage structures will be, however, necessarily driven by what have happened in the schooling sector during 1990s.

In this chapter, we analyze the pattern of changes at the upper secondary schooling level during the transitional period. We focus on the upper-secondary sector because this level is closely linked to the developments on the labor market. Moreover, the upper-secondary sector experienced by far the most profound institutional and curricular changes. Our analysis of the reform implemented in the Czech Republic tackles several issues whose importance reaches beyond simple transitional experience.

2. The Czech Educational System[1]

Prior to 1989 education in the Czech Republic was, by law, a state monopoly. Schools were both a means of training workers and a vehicle for political indoctrination. Very detailed curricula were prescribed by central authorities (Micklewright, 1999). Parental and student preferences played little, if any, role in determining how much or what type of training was provided. Entry into coveted disciplines, while certainly influenced by ability, was also heavily determined by political or other considerations.

In such an environment it is not surprising that one of the first reforms during the transition was to provide greater flexibility and give more substantial decision-making power to parents and individual schools. A key reform was to allow nonstate[2] schools to challenge the state education monopoly. Table 1 shows the extent of nonstate education in various Central European countries by the middle of the first decade after the collapse of communism. Several trends stand out. In most countries, nonstate education has achieved only limited market share. In the Czech Republic, Slovakia and Hungary, however, the share of students in nonstate schools is comparable to that in nearby EU countries such as Germany and Austria.[3] It is not a coincidence that these three countries provide the most generous level of state funding for private and religious schools. In each, funding may be as much as 100 per cent of that provided to government schools.

We focus on secondary education, which began after either eight or nine years of primary education until 1996, when nine years of primary education was made compulsory.[4] Students applied for various types of secondary school depending on their future career plans, with admission to over-subscribed programs rationed on the basis of exam performance and other considerations.

The lowest level of additional education available involves two years of vocational training.[5] Full high school education normally lasts for four years and is divided into three types: vocational education leading to a certification exam, specialized secondary (technical) education in professional fields such as nursing and engineering, and general secondary education in academic high schools known as gymnázia. Students from all three types of secondary education may continue on to university, although it is rare for those from vocational school to

Table 1
Fraction of Primary and Secondary School Students in Nonstate Schools, 1996/97

Country	% in Nonstate Schools
Central Europe	
Bulgaria	0.5%
Czech Republic	5.0%
Estonia	1.3%
Latvia	0.7%
Lithuania	0.2%
Hungary	4.6%
Poland	2.0%
Romania	0.9%
Slovenia	0.4%
Slovakia	4.6%
EU Comparison	
Austria	7.4%
Belgium	58.8%
France	20.6%
Germany	4.7%
Italy	5.7%
Netherlands	77.1%
United Kingdom	6.5%
EU Average	15.7%

Source: European Commission (1999)

do so and the majority of university students come from academic high schools.[6]

Although educational levels were on average relatively high, the structure of education is highly skewed towards vocational and away from general academic training. In 1989-90 fewer than one-quarter of secondary school students were enrolled in an academic, as opposed to a technical or vocational, program.[7] This percentage contrasts with slightly under half of secondary-level students in general academic programs in the average OECD country (OECD, 1997). By the end of decade, fewer than one in five upper secondary students in the Czech Republic were enrolled in general academic schools, the lowest proportion in any OECD member country (OECD, 2001). Furthermore, the technical and vocational education system is highly specialized. There are over 300 separate "tracks" in the Czech Republic, compared with 16 in Germany (Laporte and Schweitzer, 1994).

The legacy of the allocation system imposed by the planning authorities has resulted in substantial excess demand for various types of education (CEPR, 1998). In 1989 only 52 percent of those seeking university admission in the Czech Republic were offered at least one place. It is not possible to reconstruct from official data the success rate of students seeking admission to academic high schools.[8] It is widely understood, however, that many more students seek admission to these schools than there are places available. Similarly, places in popular fields in technical and vocational high schools, especially those required for the expanding service sector, were severely rationed during the 1990s. Thus, there should have been market niches that could be filled by entrepreneurial education providers.

In addition, public schools in the Czech Republic have substantial weaknesses that may encourage parents to seek alternatives. In particular, the public school systems are overly focused on memorization rather than creative thinking (Tomášek et. al., 1997). Finally, some parents regard public schools with distrust, given their role in indoctrination under communism, a situation paralleling the attitudes of groups such as fundamentalist Christians in the United States.

Development of the Czech educational system during the 1990s was driven by demographic trends as well as educational reforms. Table 2 shows the population at various ages in 1991 and 2000. It is clear that there were massive declines in birth rates during the final years of communism and early years of the transition. The number of children between 14 and 17 fell by over 200,000 to less than three-quarters of its 1991 level by the end of the decade. This demographic trend means that private schools were being created at a time of substantial excess capacity in the established public school system and makes their rapid increase in market share all the more remarkable.

Beyond establishing their legality, other major education reforms in the Czech Republic since the start of the transition have influenced the rise of nonstate schools. In particular, individual high schools, both public and private, were given legal status and decision-making authority over enrollment and curricula, enabling them to compete for students.

Table 3 shows the development of public spending over the decade. During the 1990s there were significant variations in public spending on education both in amount per student and in share of the Gross Domestic Product (GDP). Spending rose both in constant dollars and as a share of GDP until the last years of the decade when it fell somewhat due to reduced cohort sizes and pressure on the state budget when economic growth slowed after 1997. Between 1991 and its peak in 1996, real spending per student increased by at least 37 percent. At the end of the decade it remained 16 percent higher than at the start of transition. The pattern of spending for secondary school students follows an exaggerated version of the same path, due to the fact that the peak spending year of 1996 coincided

Table 2
Age Structure of the Czech School-Aged Population
(thousands)

Age	**7**	**8**	**9**	**10**	**11**	**12**	**13**	**14**	**15**	**16**	**17**
# of Children 1991	133.9	134.5	138.6	141.1	150.4	168.0	174.0	177.2	182.0	187.0	188.6
# of Children 2000	120.2	120.9	128.7	127.9	126	129.9	128.3	130.7	133.5	134.2	134.7

Source: Czech Statistical Yearbooks, Various Years.

with an abnormally low number of secondary school students due to the addition of a ninth mandatory year of primary school that year. Filer and Hanousek (2000) have argued that inflation measures in transition economies contain substantial upward biases. If this is true, then real expenditures were substantially greater at the end of the decade than indicated in Table 3.

As we show in Chapter 1, one of the most profound changes during the transition from communism has been a rapid and sustained increase in the value of education.[9] Table 4 shows the additional income that accrued to workers with various degrees beyond what would be earned by primary school graduates in the Czech workforce for selected years between 1989 and 1997.[10] Clearly the value of all types of education has been increasing, with the greatest increase occurring for workers with general academic or specialized technical education. Filer, Jurajda and Plánovský (1998) show that both levels of additional earnings and the increase in these levels associated with various degrees are greater for younger workers, even though many of them were trained under the communist regime.

It would be surprising if individuals did not respond to such massive changes in private pecuniary returns. Indeed, Figure 1 shows that enrollment in secondary school as a fraction of the appropriate age cohort increased throughout the decade such that by 1999 enrollment was close to 100 percent among 14 to 17 year-old young men and women.[11] In addition, as can be seen in Figure 1, the increase in enrollments was greater in those types of schools where the increase in returns was greatest.

A similar pattern can be seen in the demand for university education. Figure 2 shows the fraction of each cohort applying to and enrolling in university, where the “cohort” is defined as those who turn 18 in a given year. Given the high rejection rate among applicants,[12] and the tendency for rejected applicants to reapply for several years, it is not appropriate, however, to infer that between 60

Figure 1
Share of Age Cohort in Various Types of Secondary Schools

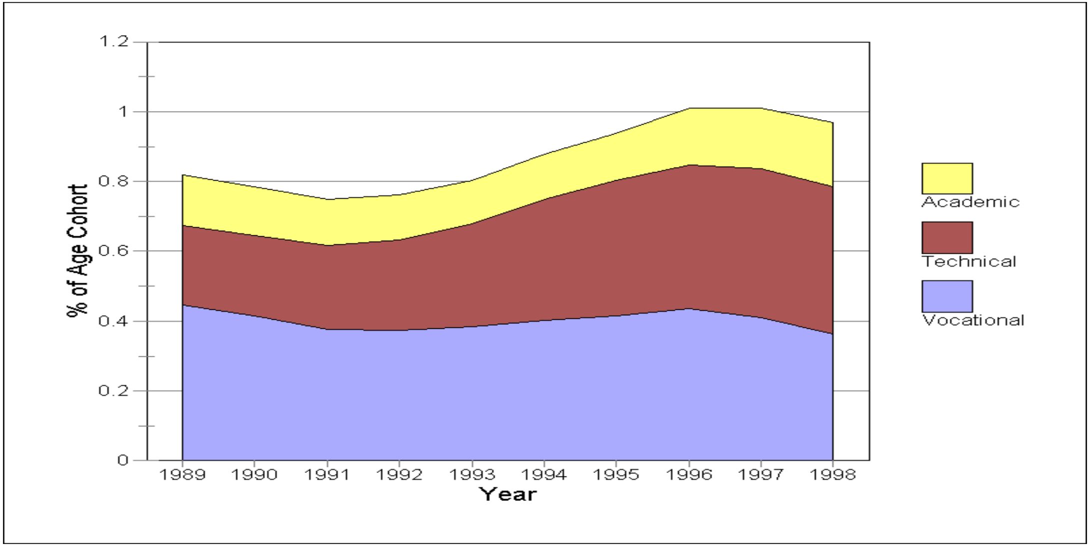

Figure 2
Applications and Admissions to University

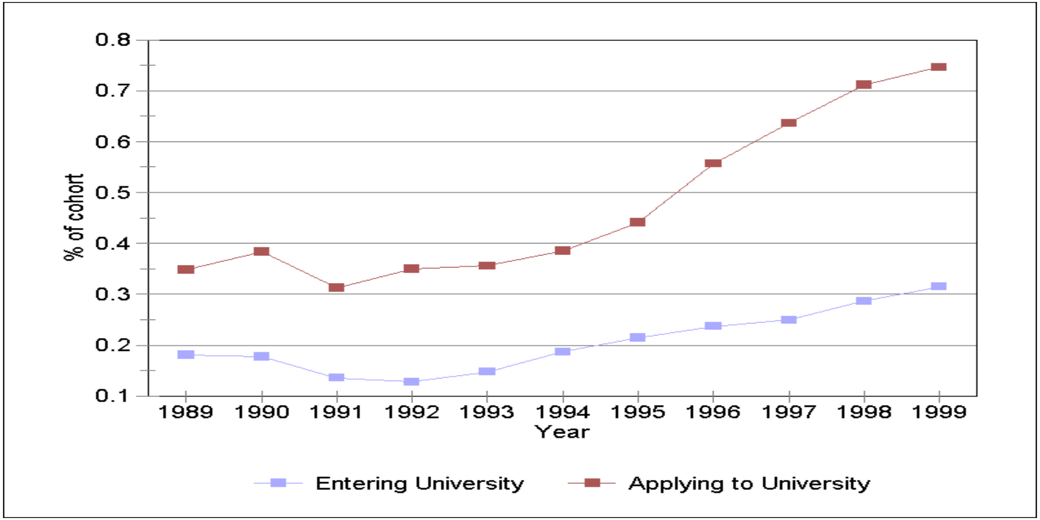

Table 3
Public Expenditures on Education

	1989	**1990**	**1991**	**1992**	**1993**	**1994**	**1995**	**1996**	**1997**	**1998**	**1999**
Billions of 1989 crowns*	21.7	21.4	18.1	20.1	23.3	24.9	26.0	27.2	24.2	22.2	22.8
1989 Crowns per Student	8870	9050	7960	8990	10490	11100	11590	12190	10970	10260	10940
As % of GDP	4.0	4.1	4.1	4.5	5.3	5.3	5.2	5.2	4.7	4.4	4.6
1989 Crowns per Secondary School Student			8150	9710	10580	12020	12780	16020	13460	12290	11700

Source: Authors' calculations based on Czech Statistical Yearbook, Various Years.

*The exchange rate has varied between 25 and 40 crowns to the dollar over the decade of the 1990s.

and 75 percent of eighteen-year-olds actually sought to go to university. It is also the case that the mean number of applications per applicant has been rising over time. In the Czech Republic the average number of applications per applicant increased from 1 (the limit allowed by the communists) in 1989 to 2.2 in 1992, after which it remained roughly constant at 2.45 or less for the remainder of the decade. Approximately 50 per cent of applicants seeking to go to university were admitted somewhere throughout the decade of the 1990s.

3. Growth of Nonstate Schools

The key reform of interest is the rise of private and church schools. Such schools were first legalized in the Czech Republic in 1990. Starting in 1992 they were funded from the state budget at a level that was at first equal to and then a substantial fraction of that provided to state schools through capitation grants based on enrolled students. In addition to public spending, those running private or church-sponsored schools generally obtain additional funding from other sources. Czech private schools generally charge tuition fees,[13] while church-sponsored schools, which are prohibited from charging such fees, are provided additional funds from congregational or diocesan resources for capital expenditures.

Despite their legality, there has been very little growth of nonstate primary schools in the Czech Republic. By the 1999/2000 school year there were only 51 private and church-related primary schools (1.3 percent of the total of 4,068 primary schools in the country), enrolling approximately 0.6 percent of all primary school pupils. Their role has been limited, frequently specializing in marginal students such as those needing special education or unable to adapt to normal school conditions. Experts suggest that the relatively minor role of nonstate primary schools may be due to their inability to attract a critical mass of students

Table 4
Increased Earnings Compared to Primary School Graduates Over Time in the Czech Republic

Level of Education	1989	1993	1996	1997
Academic HS	13.5%	27%	35.1%	52%
Technical HS	12.7%	28%	29.4%	57%
Vocational HS	7.7%	n.a.	11.2%	37%
University	28.3%	60%	54.4%	125%

Figures for 1993 calculated from Chase (1998)
Figures for 1997 from Filer, Jurajda and Plánovský (1999)
Figures for 1989, 1996 from Chapter I

since the difficulty of young children traveling long distances means that primary schools must be neighborhood-based.

At the secondary-school level the story is very different, with nonstate education playing a more important role. From a base of zero in 1990, nonstate secondary schools grew to approximately 25 percent of institutions by the middle of the 1990s. Since the average private or church-related school was significantly smaller than the average public school, however, around 13 percent of students were enrolled in nonstate secondary schools by 1995. Both the number of schools and the share of students enrolled in them appear to have leveled off by about the 1995/96 academic year and there has been little change since then.

Table 5 shows the number of state and nonstate secondary schools of various types between 1989/90 and 1999/2000.[14] It is clear that, despite declines in the number of students in the relevant age range, there has not been a commensurate decrease in the number of secondary schools since educational reform began in 1990. Indeed, the total number of secondary schools increased by 42 percent from 1246 to 1764, down from a peak of 2116 in the 1995/96 school year.[15] Two-thirds of this increase was accounted for by nonstate schools, which grew from none to 401 institutions by the end of the decade (again down from a peak of 544 institutions three years earlier). One implication of this increase, combined with the decline in the number of students in the relevant age range seen in Table 2, is that the average school size fell precipitously over the decade. Even allowing for the fact that a greater share of secondary students have enrolled in academic high schools than in the past, the average state academic high school in the Czech Republic in 1999 was 12 percent smaller than a decade ago, while enrollment in the average technical or vocational school shrank by over 40 percent.[16] This excess capacity, while creating a difficult environment in which to establish nonstate schools, also worked to their benefit. One key factor was to

Table 5
Czech Secondary Schools by Type, 1989-1999

	89/90	90/91	91/92	92/93	93/94	94/95	95/96	96/97	97/98	98/99	99/00
Academic											
State	225	227	234	244	262	276	282	283	284	277	276
Nonstate	0	2	24	41	62	72	79	84	82	79	67
Technical											
State	375	390	564	575	598	677	711	668	653	626	604
Nonstate	0	4	57	133	222	294	314	333	297	260	232
Vocational											
State	646	671	663	669	643	638	625	611	574	508	483
Nonstate	0	0	27	34	84	93	105	127	117	109	102
Total											
State	1246	1288	1461	1488	1503	1591	1618	1562	1511	1411	1363
Nonstate	0	6	108	208	368	459	498	544	496	448	401
% Nonstate	0	0.5	6.9	12.3	19.7	22.4	23.5	25.8	24.7	24.1	22.7

recognize that a school is not a building, but rather an organizational entity. Newly established private schools, therefore, were often able to rent space within buildings occupied by contracting public schools.[17]

Table 6 shows the total number of students in various types of schools over the decade. It is clear that both the fraction of teen-agers enrolled in school and the share of secondary school students in nonstate schools increased dramatically over the decade. Since the share of the cohort in state schools was approximately the same at the end of the decade as at its start, the increase in overall enrollment over the decade was almost entirely due to the rise of nonstate schools.

4. Factors Influencing the Establishment of Nonstate Schools

We now turn to the question of the distribution of nonstate schools across space, first examining academic high schools (gymnázia) and then technical high schools. We will rely on district level data for much of our analysis. Studies of labor markets have found that there is little commuting for employment across district boundaries, especially given the relatively small size of districts and the large differences in job opportunities (Erbenová, 1997). Mobility for employment, already low during communism, has declined further since 1990 (Andrle, 1998). Whether due to intense localism or poor transportation infrastructure, the lack of mobility suggests that there is also likely to be little commuting to attend schools that are in some way more attractive than those nearby.

4.1 Academic High Schools

Table 7 shows the distribution of the number of state and nonstate gymnázia by district for 1992, 1995 and 1999. There is obviously considerable variation in the presence of nonstate gymnázia. Many districts have no nonstate alternative to the state academic high school(s). We are interested in what factors determine whether a nonstate gymnázium is founded in a given district. In particular, is it the

Table 6
Enrollment in Czech Secondary Schools by Type, 1989-99
(thousands)

	89/90	90/91	91/92	92/93	93/94	94/95	95/96	96/97	97/98	98/99	99/00
Academic											
State	100.7	101.8	95.9	89.9	80.5	76.6	77.1	66.8	66.4	68.4	72.3
Nonstate	0	0.1.	0.9	3.5	5.8	8.4	9.2	8.3	7.9	7.4	7.6
Technical											
State	155.1	161.9	165	161.5	164.2	173.4	178.9	142	147.4	145.4	140.1
Nonstate	0	0.2	3.3	10.1	20.5	32.3	37.4	31.5	29.0	23.5	20.1
Vocational											
State	310.2	301.8	277.4	248.8	237.6	231.9	211.1	139.8	122.4	111	145.4
Nonstate	0	0	0	17.2	27.0	24.3	23.4	15.9	12.7	11.5	14.6
Total											
State	566	565.5	538.3	500.2	482.3	481.9	467.1	348.6	336.2	324.8	357.8
Nonstate	0	0.3	4.2	30.8	53.3	65.0	70.0	55.7	49.6	42.4	42.3
% Nonstate	0	0	0.7%	5.8%	10.0%	11.9%	13.0%	13.8%	12.9%	11.5%	10.6%

case that such schools arise when the state alternative is, in some sense, less attractive?

It should be expected that nonstate gymnázia would be more likely to be established in areas where the local public gymnázia are of lower quality. Fortunately, unlike in many other studies where the quality of public schools is measured by inputs such as spending per student, we have a direct output measure based on the success of their graduates in obtaining admission to university. Recall that the primary purpose of gymnázia is to prepare students for university admission and that the vast majority of gymnázium graduates seek to go on to tertiary education, although many applications for admission are not successful. For each school in the Czech Republic, data from the Ministry of Education enables us to calculate the success of applicants from that school in obtaining university admission measured as the ratio of *applications accepted* to *applications filed* for students from the school. We weight each observation by the ratio of total applications to total admissions for the university in question, thereby placing more weight to the better (and, therefore, more highly demanded) universities. Although there are approximately two dozen universities in the Czech Republic, there are widely perceived quality differences among them. By general consensus, three institutions (Charles University and the Czech Technical University in Prague and Masaryk University in Brno) are regarded as significantly better than other alternatives and attract substantial excess demand every year. The technical success measure is defined as:

(1) $$z_j = \sum_i (k_i * x_{i,j} / y_j)$$

where k_i = applications/admissions for university i; $x_{i,j}$ = number of applicants admitted to university i from gymnázium j; and y_j = total number of applications

Table 7
Distribution of the Number of Academic High Schools by District
Czech Republic

Number of Gymnázia in a District	Number of districts with given number of gymnázia					
	1992	1992	1995	1995	1999	1999
	State	Nonstate	State	Nonstate	State	Nonstate
0	2	59	1	44	1	45
1	5	12	5	16	6	20
2	29	2	23	11	24	9
3	22	2	21	2	23	0
4	11	0	13	1	8	1
5	2	0	6	0	7	0
6	2	1	2	0	4	0
7	1	0	2	0	1	0
8	2	0	0	0	0	0
9	0	0	1	1	1	0
>=10	1	1	2	1	2	2
Total	**77**	**77**	**77**	**77**	**77**	**77**

to all universities from gymnázium *j*. The pattern of results reported below are not affected if we do not weight the observations or if we restrict the analysis to just the three top universities.

It is not appropriate, however, to measure the quality of a school solely by the raw success of its graduates in obtaining university admission. Some schools will start with more advantaged students who would have a high success rate even if the school is actually performing poorly. Other schools may be exceptional performers in that they produce high value-added even if the overall success rate of their students is not particularly high due to poor inputs (especially the quality of entering students). We therefore measure quality by a school's performance *relative* to how well it would be expected to do based on its environment. In the first stage we estimate a school's success rate as a function of the share of the local population with at least a secondary school maturita-level education, share of the population living in towns (as opposed to villages), average class size in primary schools in the district (on the assumption that this will be reflected in the average quality of students arriving in the local gymnázia) and grade-point average of applicants to the gymnázium from primary school (again to reflect quality of incoming students). For each public gymnázium we calculate the residual between the predicted and actual success in university admissions. We then average these school-specific residuals for all public gymnázia in a district weighted by enrollment and use this district average residual as a measure of public school quality in each local market.[18] We use this public school quality measure together with other explanatory variables in the second stage model of nonstate school entry.

Table 8a reports descriptive statistics on explanatory variables considered in the first and second-stage regressions. The first stage OLS regression results are presented in Table 8b. Success in obtaining admission to university among students in a given district is positively affected by the average education of adults in that district and the elementary school performance of the students. It is negatively affected by average class size in the district's primary schools and the degree of urbanization of the district.

The results of second-stage probit regression are presented in Table 8c. The dependent variable takes the value of one if a nonstate gymnázium was established in a district by 1995.[19] Several factors are associated with whether a nonstate gymnázium was established including the education level of the population (which should reflect both the preference for and ability to pay for academic education among parents), changes in high-school age cohort size, and unemployment rate. The positive relationship between a district's unemployment rate and the presence of a nonstate gymnázium may reflect dissatisfaction with the performance of the public school system if recent graduates from that system are having a hard time finding jobs.[20] It may also reflect a shift in demand from technical or vocational education towards more general academic skills in the presence of economic uncertainty. Large declines in cohort size may be particularly relevant if they signal significant excess capacity in local public schools. We also include the number of students in nonacademic high schools and the population density of the district. These are highly collinear so that either, but not both, is significant when included, although they are jointly significant. Obviously both serve as a measure of potential demand.

In addition to the factors reported in Table 8c, we analyze several other factors that might be related to the establishment of nonstate schools. These include the share of votes for the ruling coalition parties (on the grounds that regions that supported the government might have received favorable capital investment treatment in the public system), average wages in the district, the share of employment in agriculture, the distance from a major urban area (where

Table 8a
Descriptive statistics

1st stage regression	**mean**	**std.dev.**	**min**	**max**
Edu23	0.30	0.080	0.21	0.47
Townpop	0.71	0.199	0.34	1.00
Pclass	23.02	1.054	20.09	24.02
Agrade	1.30	0.138	1.06	1.90
Prague	0.13	0.338	0	1
2nd stage regression				
Edu3	0.06	0.021	0.03	0.16
DelCohort	1.35	0.087	1.17	1.57
Dens	210.30	392.83	36.00	2451.10
Non-Gym	2037	1815	614	15984
Unemp	0.04	0.02	0.003	0.08
Q	-1.73	13.36	-30.96	33.40

1st STAGE: Estimating district level public schools' quality

Agrade	Grade average of applicants to Gymnazia from Primary school. A measure similar to elem. 1<=agrade<5; 1 is the best and 5 the worst.
Pclass	Pupils/class district ratio in Primary school in district
Edu23	Share of population with at least full-secondary education
Townpop	Share of population living in towns
Prague	Dummy if Prague

2nd STAGE: Estimates of nonstate school entry (probit)

Dens	District population density
Non-Gym	Number of students in technical and vocational high schools in 1991
DelCohort	Percentage decline in high school aged cohort between 1992 and 1994
Edu3	Share of district population with university education
Unemp	District unemployment rate
Q	Public schools' quality

Table 8b
1st stage estimates of school graduates' quality regression

	Coef.	Std. Err.	t-stat
Edu23	3.33	0.45	7.32
Townpop	-0.38	0.12	3.19
Pclass	-0.91	0.55	-1.66
Agrade	-0.42	0.13	-3.19
Prague	-0.63	0.08	8.03
Const	2.13	0.71	3
Nobs		***219***	
AdjR^2		***0.32***	

Table 8c
Probit Estimates of the Probability of a
Nonstate Gymnázium in a District by 1995

	dF/dx	Std.Err.	z
Edu3	21.18	8.74	2.46
DelCohort	-1.95	0.93	2.07
Dens	0.002	0.001	1.25
Non-Gym	0.0001	0.0002	0.76
Unemp	10.24	4.79	2.19
Q	-1.12	0.58	1.91
Nobs		***76***	
pseudo R^2		***0.29***	

universities are located), and the size of the largest town in the district without a public gymnázium. None had a significant effect, so they have not been included in the estimates reported.

As can be seen in Table 8c, all coefficients have reasonable signs. District education levels, potential student pool (jointly captured by population density and number of students enrolled in nonacademic secondary schools), and unemployment rate all increase the probability that a nonstate gymnázium is founded in a district, while a rapidly decreasing cohort size reduces that probability. Most importantly, the effect of public gymnázia quality on whether or not a nonstate gymnázium is established is significant and negative. If public schools in an area do better than expected in getting their graduates into university, it is less likely that nonstate competitors will emerge. Evaluated at mean levels, a one standard-deviation increase in the unexplained admission success of public gymnázia in a district results in approximately a one-third reduction in the probability of a nonstate gymnázium being established in that district.

We experimented with alternative specifications of quality. When we use the weighted success rate calculated in directly in Equation (1), not surprisingly, there was a significant negative relationship between state schools' success in obtaining university admission for their students and the likelihood that private competition would arise. If we include total admissions success, partitioned into explained and unexplained components, it is clear that it is the unexplained component (our residual) that matters.

There is an issue of simultaneity that might bias our results. The earliest data we have available for measuring admission success come from the 1995/96 academic year. Thus, we are using a measure of success *after* the establishment of nonstate schools to predict their establishment. While it would obviously be preferable to use admission success from 1992 or earlier, before there were a significant number of nonstate schools, it is likely that any biases introduced by the use of later data will work against the effect we see, thereby strengthening our results. We say this for three reasons.

As we will discuss below, there is evidence that state gymnázia respond to the challenge of nonstate competition by improving quality. To the extent that such improvement had already taken place by the time we measured public-school quality, the quality differences between districts where nonstate schools came into existence and those where they did not should have been even greater at the time the nonstate schools were established.

Furthermore, nonstate gymnázia typically draw from the lower end of the quality distribution among potential applicants. Such a finding contrasts strongly with those for Chile in Hseih and Urquiola (2002) and casts doubt on the assertion that private school vouchers would result in cream skimming. It is, however, consistent with maximizing behavior. Top students in state schools can expect admission to leading universities in any case and so have little reason to pay the additional charges for private schools. It is the marginal students in poorer state schools who have the most to gain from attending private institutions. As a rough measure, the grade-point average of elementary school students entering state gymnázia in 1995 was 1.30 (where 1 is the best grade and 5 is the worst), while the average for those entering nonstate gymnázia was 1.5. Thus, where nonstate gymnázia exist, there should be a selectivity effect increasing the average quality of students remaining in the public system, again serving to reduce cross-district differences in public school quality from the time when the nonstate schools were established.

Finally, the pattern of development of nonstate schools means that even if a nonstate gymnázium existed in a district in 1996, the fact that new schools rarely enrolled students in their last year of education meant that most graduates during that year would be from state schools.

4.2 Technical Schools

The role of secondary technical schools is quite different from that of gymnázia. Technical schools are expected to provide education that directly affects the labor market productivity of graduates, instead of preparing them for successful enrollment in university. The proportion of technical school graduates who apply to universities is much smaller than the proportion from gymnázia, while differences in the proportions admitted are even more pronounced. Thus, the admission rate to university is not a proper indicator of technical school quality. Moreover, the curricula provided by technical schools is heterogeneous by definition. Although there have been some attempts to shrink the number of fields of study after 1989, hundreds of different curricula remain at the end of the 1990s.[21]

As discussed above, inter-regional labor force mobility in the Czech Republic is extremely low, with what mobility that exists arising mainly because of marriages. This implies that regions are, to a large extent, local markets and there should be a close relationship between the education provided and the prevailing industrial/occupational structure at the regional level. On the other hand, when disruptions due to the transition resulted in a mismatch between the educational institutions in a given region and that region's current labor market needs, there will be little possibility of resolving this mismatch through mobility of workers trained in other regions.

We would expect private technical schools to be more responsive than public ones to local labor market conditions and to arise when there are obvious niches to be filled in the demand for education. To test this hypothesis we consider a simple linear model:

(2) $$S^{I}_{j,d} = D^{'}_{j,d} b + e_{j,d}$$

where S represents an education supply indicator, D represents a vector of determinants of local demand for education, *ß* is a vector of parameters, ***e*** is a stochastic error term, and the subscripts j and d identify vocational branch and region. As in the case of gymnázia, public-school quality should be an important determinant of whether private schools arise. Since, however, we cannot rely on admission rates to universities, we focus instead on returns to education as the major determinant of school choice and demand for individual fields of study. Our hypothesis is that the likelihood that a private school with a given occupational specialization is established in a region depends positively on the relative scarcity of labor of a particular type in the region.

To obtain a proxy for the unmet demands of the market economy, we look at the growth in unexplained earnings (the residual from a standard Mincerian wage equation) between 1989 and 1996. The intuition is that occupation/location cells with a high growth in unexplained earnings are those in disequilibrium because of labor shortages.[22] The indicator is:

$$\Delta w_{j,d} = w_{j,d}^{96} - w_{j,d}^{89} \tag{3}$$

where w represents the average residual from an economy-wide log wage equation for individuals with secondary-school technical education (and the maturita exam) in a given region/occupation cell. This equation was estimated using a representative survey of approximately 3500 workers who were asked retrospective wage histories for the period 1989-1996 (see Münich et al. (1999), for a description of this data).[23] Averages are taken across 17 occupations and 8 regions including Prague.[24]

We would also expect the likelihood of a nonstate school being established to be greater in regions where there is significant unemployment, thereby indicating that current schools are not providing appropriate training. The local unemployment rate is therefore included on the right-hand side of our estimating equations, as well as dummies for Prague and for the Finance and Business field where we know *ex ante* that there were significant labor shortages.

Although we have 136 possible observations (17 fields times 8 regions), we are constrained by the number of observations available in the retrospective survey. For some cells we do not have a sufficient number of individuals to estimate reliable mean residuals from the wage equations. We have excluded cells with fewer than 5 observations, leaving us 91 observation units. Increasing this threshold does not change the results.

We have examined two indicators of the responsiveness of private schools to labor market conditions. The first is the ratio of private school to public school enrollment in a given region and field of study, while the second is the ratio of nonstate school enrollment in a given region and field to total regional enrollment in technical schools, both private and public, across all fields. Finally, for comparison purposes, we examine the growth rate of public technical school enrollment in a given field/region cell between 1991 and 1996 to see if public schools also respond to labor market conditions.

Regression results for alternative supply of education indicators are presented in Table 9. Since many region/field cells do not have any private school enrollment, estimates for the extent of private school enrollment are of the truncated regression (Tobit) form, while the growth in public school enrollment is estimated using OLS.

Results are quite clear. Both public and private technical schools have responded to market demands by increasing training in financial and business subjects. Here, however, the similarity ends. Nonstate schools have created opportunities for training in the areas and fields where wages have been growing rapidly, indicating increasing demand, and where unemployment rates are highest,

Table 9
Determinants of Vocational School Enrollment, 1996
(standard errors in parentheses)

	Private Enrollment/ Public Enrollment (Within Cell)	**Private Enrollment (Within Cell)/ Total Enrollment**	**Increase in Public Enrollment (1991-1996)**
Increase in Wage Residual	.193*** (.067)	.042*** (.015)	-.121 (.330)
Unemployment Rate	.104** (.045)	.015* (.009)	-.117 (.189)
Prague	.358*** (.105)	.055*** (.021)	-.489 (.444)
Finance & Business	.618*** (.050)	.202*** (.012)	.759** (.307)
Constant	-.212*** (.065)	-.041*** (.013)	.198 (.242)
Estimation Method	Tobit	Tobit	OLS

***Significant at 1% level
**Significant at 5% level
*Significant at 10% level

indicating a greater regional mismatch of workers and jobs. Public technical schools, on the other hand, exhibit no such response to market conditions. In fact, the signs on the rate of wage growth and local unemployment rates are negative, although insignificant.

This pattern makes sense. As we saw earlier, returns to technical education increased the most during the post-communist period. It is likely that pre-existing state technical schools provided training better suited to the old industrial structure, thereby leaving gaps in the curricula demanded by employers in the emerging market economy. These gaps were filled by newly created nonstate schools which, unconstrained by past investments in physical and human capital, have concentrated in high-demand areas such as commerce, economics and hotel management. State schools, on the other hand, have not been able to keep up with shifting market demands, perhaps because they have less incentive to do so than nonstate alternatives. Thus, our results are consistent with recent reports that the unemployment rate of graduates from nonstate secondary schools is substantially lower than that of graduates from public schools (UIV, 2000).

5. Perception of Nonstate Schooling and Parental School Choice

To gain an understanding of public awareness of nonstate school entry and competition we examined a representative sample of 1411 individuals surveyed in 1996.[25] The survey asked about family status, occupation, educational background, children, and schooling-related opinions and preferences. Using these data we examine perceptions of nonstate schools among parents as well as factors influencing parents' actual school choice.

People responded to several statements about nonstate schools using a 4-point rating scale ranging from (1) strong disagreement to (4) strong agreement. Although answers are categorical, they are ordered. We, therefore, estimate an ordered logit model to identify key determinants of individuals' opinions.

Table 10 presents results of factors influencing opinions regarding private schools.[26] Columns (2) deals with the statement: "Private schools serve as competitors to public schools, enhancing the quality of schooling." Overall, we find that public school teachers are more likely to disagree with the statement, reflecting their inherent biases and self-interest. Agreement is highest for individuals of about 50 years of age, by parents of school-aged children, and especially by parents of a child in a nonstate school who find local school choices satisfactory. Disagreement with the statement is more likely among parents who do not find local school choice satisfactory yet still have a child in such a school.[27] Higher-income women are less likely to agree with this statement, perhaps reflecting the concentration of such individuals in Prague where state schools are better in general. Interestingly, education and being a teacher in a private school do not affect the opinion.

Respondents were also asked their agreement with the assertion that "Private schools are mostly of better quality than public ones." Results in column (3) suggest that public-school teachers' opinions serve their self-interest by

Table 10
Determinants of Perceptions About Nonstate Schools

	Statement No.*					
	1 Total	**1 Males**	**1 Females**	**2 Females**	**3 Females**	**Definition of variables**
TEACHER	-0.399	-0.192	-0.505	0.021	-0.788	dummy=1 if teacher, 0 otherwise
	-0.072	*-0.603*	*-0.073*	*-0.075*	*-0.006*	
TEACH_PR	1.083	-	1.156	-2.324	1.513	dummy=1 if teacher in nonstate school, 0 otherwise
	-0.286	-	*-0.254*	*-0.007*	*-0.122*	
EXP_BAD	-0.295	0.051	-0.479	0.138	0.087	dummy=1 if child in state school and poor local school choice
	-0.163	*-0.886*	*-0.069*	*-0.069*	*-0.748*	
EXP_GOOD	1.899	33.102	1.501	-0.527	1.851	dummy=1 if child in nonstate school and good local school choice
	-0.009	*-0.999*	*-0.052*	*-0.506*	*-0.007*	
INFORMED	0.318	0.232	0.336	-0.302	-0.014	dummy=1 if school age child present
	-0.021	*-0.278*	*-0.065*	*-0.11*	*-0.943*	
EDU	-0.028	-0.041	-0.018	-0.047	-0.131	years of education
	-0.321	*-0.322*	*-0.641*	*-0.236*	*-0.001*	
AGE	-0.052	-0.043	-0.059	0.099	-0.012	years of age
	-0.05	*-0.285*	*-0.101*	*-0.011*	*-0.74*	
AGE2	0.512	0.444	0.553	-0.878	0.094	age squared /1000
	-0.078	*-0.309*	*-0.158*	*-0.04*	*-0.818*	

Table 10: continued

GENDER	2.022	-	-	-	-	dummy=1 if female
	-0.319	-	-	-	-	
LEARN_M	-0.026	-0.033	-	-	-	log(monthly earnings) of primary male in household
	-0.892	*-0.886*	-	-	-	
LEARN_W	-0.292	-	-0.283	-0.027	-0.387	log(monthly earnings) of primary woman in household
	-0.066	-	*-0.089*	*-0.873*	*-0.021*	
Pseudo R2	***0.0124***	***0.0085***	***0.0156***	***0.025***	***0.045***	
Nobs	***1142***	***483***	***659***	***673***	***653***	

p-values in parenthesis

*Statement No.1: “Private schools serve as competitors to public schools, enhancing quality of schooling.”
Statement No.2: “Private schools are accessible mainly to the rich”
Statement No.3: “Private schools are mostly of better quality than public ones.”
Answers: ranging from (1) strong disagreement to (4) strong agreement

disagreeing with this statement while, as might be expected, parents who have opted to send their child to a private school even though they believe the local state school is good are especially likely to believe in the quality of the private school. More educated and higher earning parents are less likely to agree, perhaps representing a Prague effect.

Overall, it appears that the perception of the role of nonstate schools depends strongly on the amount of information and experience an individual has with the school system. Incumbent state-school teachers express a negative attitude towards nonstate schools, reflecting their vested interests. Conversely, those who work in nonstate schools are generally supportive. Parents, or at least mothers, appear to have reasonable and reliable opinions regarding the quality of local schools, but may overestimate the costs of alternatives.

The second issue we look at is the decision to enroll a child in a nonstate school. The data possess some limitations because there are only 661 persons in our sample who have a child in either a grammar or upper secondary school.[28] To identify factors determining state/private school choice, we use a standard probit model, with results presented in column (4) of Table 10. We find that teachers in state schools are less likely to place their child into a nonstate school. Parents with higher earnings and families with more education are more likely to choose a nonstate school. As dicussed above, a primary market for nonstate schools are students who are not admitted to local gymnázia. Educated parents are likely to be more upset if their child does not go to a gymnazium, and will, therefore be more likely to enroll him in a nonstate gymnazium whereas less educated parents will simply send the child to a vocational/apprentice school. The likelihood of a nonstate school choice declines at a diminishing rate with parental age until the mid-40's after which it increases. At the same time, 28 percent of respondents reporting a school age child are older than 43 years.[29]

6. The Role of Nonstate Schools in Promoting Reform of State Schools

Finally, we turn to the issue of how state schools respond when confronted with competition from nonstate schools. Here our evidence comes only from the study of gymnázia due to a lack of data on other types of public schools. We divided the 77 districts into three groups according to the fraction of newly enrolled students entering nonstate gymnázia in 1995. Public gymnázia in 44 districts faced no competition from nonstate alternatives, while in another 7 districts newly established private gymnázia were of the extended format and enrolling students only in the lower grades in 1995. This left 26 districts where there was significant competition. These were further divided into two groups: those where less than an arbitrary 20 percent of new enrollees in gymnázia opted for nonstate schools and those where this percentage was greater than 20.[30]

Table 11 shows that there were substantial differences in how state schools behaved depending on the degree of competition they faced from private alternatives. Although the size of the cohort decline between 1992 and 1999 was similar across districts with differing degrees of nonstate competition, total enrollment in gymnásia remained almost constant when competition was extensive. More critically, despite a 16 per cent decline in enrollment, public gymnásia in districts where there was significant competition increased the number of classroom teachers by 22 percent and saw average class sizes fall by almost a third, a 50 percent greater drop than public gymnásia that did not face extensive competition, where teaching staff remain almost constant and cohort declines resulted in average class size falling by a fifth.[31] Since operating funding for schools is based on capitation grants and is not influenced by the degree of competition from local nonstate schools. Indeed, given that funds are a linear function of enrollment, the existence of fixed costs for administrators should mean that diversion of students to nonstate schools should result in *increases* in average

Table 11
Determinants of Choice of Nonstate Schooling
(dependent variable: Child in nonstate/state school ~1/0**)**

	dF/dx	**Definition of variables**
	(p-value)	
TEACHER	-0.027	Dummy=1 if teacher, 0 otherwise
	(0.063)	
TEACH_PR	0.225	Dummy=1 if teacher in nonstate school, 0 otherwise
	(0.080)	
AGE	-0.014	Years of age
	(0.069)	
AGE2*1000	0.164	age squared *1000
	(0.071)	
LEARN	0.030	log(monthly earnings)
	(0.024)	
EDUMAX	0.009	Maximum parental years of education
	(0.009)	
obs. P	***0.038***	
pred. P	***0.024***	
Pseudo R2	***0.136***	
Nobs	***474***	

p-values in parentheses, p>|z| are the test of the underlying coefficient being 0;
dF/dx is for discrete change of dummy variable from 0 to 1

class sizes. Since by far the largest component of costs is salaries (77 per cent of total costs in 1997), the marked reduction in class size suggests that schools facing competition must be reallocating resources away from administrative and other noninstructional personnel towards classroom teachers, real differences in behavior designed to make the public gymnázia more attractive.[32] Finally, between 1996 and 1998 while state gymnázia in districts facing the greatest competition increased the PC/pupil ratio by 683 percent, state gymnázia in districts where there was no private competition increased the number of personal computers per pupil by 533 percent.

Perhaps the greatest difference can be seen in what we have argued is the true test of the performance of gymnázia, success in gaining university admission for graduates. If we rank districts from 1 to 77 according to the success of graduates from their state gymnázia in obtaining university admission,[33] public gymnázia facing significant competition improved their relative rank by 4.5 positions between 1996 and 1998, while those facing moderate competition improved their ranking by an average of 0.6 positions. Given that there are a fixed number of districts, these improvements came at the expense of state gymnázia in districts where there was no competition from nonstate alternatives. State gymnázia in these districts saw their relative position deteriorate by an average of 1.4 positions. This result is not a statistical artifact created by selection effects with the worst students leaving public gymnázia for private ones. If we include admissions for *all* graduates in a district, whether from state or nonstate schools, the improvement in a district's performance when there is competition is even stronger. Combining state and nonstate schools, the typical district with strong competition improved its success in obtaining university admission for its graduates by 13.3 positions. Districts without the spur of competition saw their relative rank decline by an average of 1.6 positions.

Table 12
Changes in Public School Inputs and Quality
According to Degree of Competition by 1995

	Little or No Competition	**Moderate Competition**	**Extensive Competition**
Percentage decline in population aged 14 to 17			
1992-1999	-24.9%	-22.6%	-26.2%
1995-1999	-14.0%	-13.0%	-16.8%
Percentage change in students in all gymnásia			
1992-1999	-19.6%	-19.8%	-4.6%
1995 - 1999	-4.4%	-15.2%	-6.9%
Percentage change in students in public gymnásia			
1992-1999	-20.3%	-24.0%	-15.9%
1995-1999	-5.1%	-15.1%	-2.4%
Percentage change in teachers in public gymnásia			
1992-1999	+1.0%	-4.3%	+21.8%
1995-1999	-5.5%	-16.6%	-1.0%
Percentage change in public student/teacher ratio			
1992-1999	-6.4	1.3	-4.1 (p=.08)
1995-1999	2.5	8.7	4.7 (p=.22)
1992-1996	-3.9	-7.4	-9.9 (p=.02)
Public schools student/teacher ratio			
1992	11.71	11.79	12.3
1995	11.46	11.02	11.23
1999	11.53	11.93	11.76
Computers per pupil 1998	0.27	0.22	0.28
Percentage increase (PC/student, 1996-1998)	533%	453%	683% (p=.48)
Relative success in obtaining admission to university (1 = best, 77 = worst)			
Public Schools Only			
Average change in rank	+1.4	-0.6	-4.5
Average Rank 1996	42.0	40.9	27.26
Average Rank 1998	43.5	40.4	22.8
Public&Priv. Schools			
Average change in rank	+1.62	+10.5	-13.3
Average Rank 1996	40.2	32.27	39.8
Average Rank 1998	41.8	42.8	26.5
Number of districts	51	11	15

p-values correspond to significance level of a difference in mean values between extensive and no-competition group

Unfortunately, there is also some evidence of efforts on the part of state school bureaucrats to capture the funding mechanism and reestablish their favored position. In a personal letter to the Czech Prime Minister when the structure of the Czech funding mechanism was being designed, Milton Friedman pointed out that to ensure effective school choice it was “important to examine the details and not only the titles of the proposals” in order to ensure that entrenched interests would be forced to provide real choice. The fate of private schools in the Czech Republic underlines how critical it is in ensuring effective competition for funding of nonstate schools to be automatic and subject to the lowest possible level of interference from those with a vested interest in seeing alternatives to current arrangements fail.

The impact of the efforts of vested interests to thwart the growth of competition may be seen in Tables 5 and 6, where the growth of nonstate schools came to an abrupt halt around 1996, with their share of schools and of students actually shrinking in the past few years. In large part this can be attributed to road-blocks created by administrators in the Czech Ministry of Education under the influence of lobbying from state schools. In the 1992 reforms, the per student funding level for nonstate schools was equal to that for state schools of the same type in the same area. Under pressure from education authorities, the principle of “equal treatment” for nonstate schools was abandoned in 1995, and the level of support for nonstate schools was set equal to 60 to 90 percent of the subsidy provided to state schools, with the exact amount being set by the Ministry of Education and regional school authorities on the basis of poorly specified performance criteria. This is exactly when the growth in nonstate schools halted, perhaps due to the need of such schools to increase tuition in order to offset reduced public support.

The complexity introduced by these reforms resulted in a lack of transparency and created uncertainty for operators of nonstate schools. The budget outlining Ministry of Education support for schools grew from 65 pages in 1992, the first year of the capitation system, to over 350 pages by 1999. The problems for nonstate schools that were inherent in this arbitrary process created dissatisfaction and resulted in the law being amended again in late1998, to take effect with the 1999-2000 school year. Currently, public support for private schools is based on a two-part formula. Base support at the level of 50 percent of total support for state schools is now given according to the type of school and is independent of quality or ownership. There is then a supplement that varies according to quality as evaluated by district (county) level school offices (with final determination approved by a board at the Ministry of Education). Nonstate schools can obtain a maximum supplement equal to 90 percent of that available to state schools. In addition, the law now limits the discretion of the ministry and schools offices when evaluating quality to an explicit set of criteria. This policy was adopted in order to protect nonstate schools from arbitrary denial of funds by public officials.[34] Interestingly, 1999 was the first year for quite some time when the average tuition charged by private schools fell. It remains to be seen whether this will reverse the decline in the number and enrollment of these schools during the previous three years.

7. Contribution to the debate about education vouchers

Our analysis contributes to the ongoing debate about so called educational vouchers also called the movement for school choice. Given reluctance to adopt major reforms without evidence that they would be successful, there have been few opportunities to test the theoretical assertions of advocates for educational vouchers (see Friedman and Friedman, 1981 or Chubb and Moe, 1990). Key among these assertions are that, if universal vouchers were adopted, (1) nonstate schools would arise to offer options to students currently trapped in poorly performing public schools; (2) parents would be aware of, and make use of, the options available to them; and (3) public school authorities would respond by reforming and improving school quality rather than using bureaucratic regulation to stifle effective challenge to their monopoly position.

Without actually adopting a universal voucher system it is impossible to determine how the market would respond were one to be implemented.[35] The limited evidence that does exist suggests that public schools respond to increased competition by increasing quality[36] and that parents make use of choice to enhance their children's opportunities.[37]

The fundamental difficulty in testing the assertions of voucher advocates, mostly in the United States, has led to interest in establishing their impact in other countries. As Hanushek (2002, p. 79) has observed: "International evidence currently offers the best chances for understanding the impacts of voucher systems, particularly the longer run implications."[38] Here the transition experience of the Czech Republic offered us an ideal experiment. Although the legacies of communism create some unusual initial conditions, they also offered a unique opportunity to examine the responses of both public and private schools when an entire system is opened up to significant possibilities for competition at a single stroke.

Results are generally encouraging for voucher advocates. We find that private schools did arise, even in a period of substantial excess capacity among state schools. Moreover, they were more likely to arise where the state schools were doing a worse job in meeting their mission. Parents were aware of and responded to school quality. Finally, competition from private schools led to reforms in the public schools that increased their quality. On the negative side, however, there is evidence that the public school bureaucracy at least partially captured and redesigned the voucher system in order to preserve their privileged status.

8. Summary and Conclusions

Post-communist Central Europe provides an interesting laboratory in which to investigate possible responses were a large state to adopt universal education vouchers. Although public schools in the Czech Republic were relatively good by objective standards, there was an initial surge in demand for private alternatives that eventually reached between 10 and 15 percent of the secondary school population. Private schools appear to have arisen in response to distinct market incentives. They are more common in fields where public school inertia has resulted in an under-supply of available slots. They are also more common where the state schools appear to be doing a worse job in their primary educational mission, as demonstrated by the success rate of academic high schools in obtaining admission to the top universities for their graduates.

There is also evidence that state schools facing private competition do improve their performance. State schools facing competition spend a larger fraction of their resources on classroom instruction and significantly reduce class sizes. Furthermore, Czech public academic high schools facing significant private competition in 1995 substantially improved their relative success in obtaining university admissions for their graduates between 1996 and 1998.

While generalizations must be made with caution due to the unique nature of the times during which private schools arose in the Czech Republic and to the limited time over which to observe the responses of public schools, evidence from the adoption of the Czech nation-wide voucher scheme supports the claims of advocates for such systems. Private schools supported by vouchers increased educational opportunity and spurred public schools to improve performance. They also spurred public schools to engage in bureaucratic manoeuvring designed to preserve their entrenched position. Overall, the experience of education funding reform in the Czech Republic since the fall of communism provides support for the theoretical underpinnings of the case for school vouchers. It does, however, point out how important it is that a funding system be simple and leave as little opportunity as possible for discretionary actions on the part of implementing officials if it is to avoid capture by the current school bureaucracy and enable private schools to provide effective competition to state schools' monopoly status.

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Notes for Chapter III

1.For more detail on the Czech educational system and reforms since 1989 see Filer and Münich (2002).

2.We will use the term nonstate to refer to all types of education administered by nongovernment entities such as churches, foundations, profit-making corporations and individuals.

3.The share of students in nonstate schools in the Czech Republic lags considerably behind the EU average of almost 16 percent. This average is heavily influenced by very high nonstate enrollments in countries such as France, Belgium and the Netherlands where the tradition is for each of several antagonistic linguistic or religious groups to operate independent school systems with state funding.

4.Prior to 1996, students were allowed to apply for secondary education after eight years of primary school but could, particularly if they did not obtain their desired placement, remain in primary school for a ninth year. In the 1995/96 school year only about 5 percent remained in primary school for the ninth grade, down from about 20 per cent a few years earlier.

5.Except for students who studied only for one more year before reaching the legal school-leaving age if they remained in primary schools for the full nine years possible.

6.In order to enroll in university students must leave secondary school with an exam credential known as a maturita. Whether or not a student receives this credential, and can therefore continue on to university studies, depends on their program or course of study. In the Czech Republic, all gymnázia and 96 percent of technical school programs, but only 14 percent of vocational school programs, lead to a maturita and the possibility of university admission. In fact, many vocational schools are three years or less in length and cannot provide the maturita required for university admission. A market niche has arisen, therefore, for schools providing what is known as “addendum” programs to allow such students to qualify for university.

7.Prior to approximately 1992, all academic high schools were four-year programs. With the freedom allowed after 1989, a number of gymnázia began admitting students after the fourth, fifth or sixth year of primary school. By the 1997/98 school year, these “extended gymnázia” accounted for over 40 percent of gymnázia students in the secondary-school (above grade 9) years. There has been considerable discussion about the impact of this reform on primary schools. It is generally assumed that the more talented and academically motivated students leave primary school for the extended gymnázia, resulting in less classroom stimulation and lower probability of academic success for those left behind. If this is true, then the trend will be self-reinforcing and the share of extended gymnázia should continue to grow over time. It should be noted, however, that this reform developed independently of the rise of nonstate schools. Indeed, the division between extended and conventional gymnázia is approximately the same in the state and nonstate sectors. Below we focus only on students in the secondary school part of extended gymnásia.

8.Data is reported on the number of applications and the number of acceptances but not on the number of applicants.

9.Chase (1998), Filer, Jurajda and Plánovský (1999, 1998) provide alternative discussions of trends in returns to education in the Czech Republic.

10.Although often called such in the literature, the results presented are not technically "returns to education" since they show only the private benefit of a given degree and ignore both social returns and the costs associated with that degree.

11.Determination of precise enrollment rates is complicated by the extension of primary school that occurred in 1996. We have omitted 14 year-olds from the relevant population in 1996 and later. The fact that there is no discontinuity in the trend line in Figure 1 at this point suggests that this is approximately the correct adjustment.

12.In the mid-1990s roughly 80 percent of Czech gymnázium graduates, 37 percent of technical secondary school graduates and 22 percent of eligible vocational school graduates (i.e., the 8 to 10 percent of vocational school graduates who were enrolled in courses leading to the maturita) were successful in enrolling in university within two years of their graduation from secondary school. Obviously some graduates elect not to apply to university but overall places are still severely rationed.

13.In 1998 the mean annual tuition charged by nonstate gymnázia was approximately 15,000 Czech crowns ($450) with a range of from 1,500 crowns to 29,000 crowns. By way of reference, the mean annual wage during this year was approximately 150,000 crowns per worker while most households had at least two workers.

14.Approximately 85 percent of nonstate schools were private while 15 percent were church-related. The number of church-related schools is too small to enable independent analysis of various types of nonstate schools.

15.These figures exclude a small number of highly specialized schools such as dance and music academies.

16.This obviously raises questions of over-capacity and excess spending on fixed plant. Although there have been attempts to close unneeded public schools, given entrenched bureaucracies and reluctance to commute long distances, these attempts have met with only limited success.

17.This suggests an additional reason why nonstate schools were less likely to be established at the primary level. Schools at this level have much less autonomy than secondary schools, so individual school officials were not able to sign contracts to lease space and any revenues generated would have had to be returned to the state budget.

18.Although derived independently, the methodology used to assess school quality is similar to that devised by NORC at the University of Chicago for a recent survey of high-school quality in the US (*U.S. News*, 1999). There is one potential complication with our measure of university admission success. The number of applications per student, the quality of the schools to which a given student applies, or the fraction of students in a cohort who apply to university at all could be correlated with school quality if better school management motivates students to a greater extent. Application rates are so close to 1 among gymnázium graduates that we doubt the last of these is a problem. The first two, if a factor, should bias our results towards zero.

19.As we saw earlier, there has been little change in the extent of nonstate schools since 1995 in the Czech Republic.

20.In the early years of the decade the overall unemployment rate in the Czech Republic was quite low, especially for a transition economy, averaging under 4 percent. Some districts, however, had unemployment rates that approached 10 percent. In recent years the overall rate has risen to slightly under 10 percent with the rate in some districts exceeding 20 percent.

21.Personal conversations have revealed that there are approximately fifteen people at the Czech Ministry of Education, each of whom supervises a different set of vocational fields. The number of fields within a given general branch depends on the belief of the individual in charge of that branch, with new fields proposed by schools being approved or rejected based on personal opinion. This has resulted, for example, in many highly-specialized fields in the Electronics branch but few in the Engineering branch.

22.We use residuals, rather than levels, in order to account for the likelihood that even under communism unmeasured job characteristics resulted in higher wages for industries such as mining.

23.The Czech Labor Force Survey does not collect wage information and there is no other data source such as household budget surveys or micro-censuses that provides information on wages for both the pre- and post-transition periods by branches.

24.Educational data distinguish 31 occupational fields of study. These fields are not identical with occupational data from the retrospective survey. Therefore, for each two-digit occupational code we identified the closest matching educational field. We merged several fields if they fell within a single occupational group (examples include Architecture-Urban Planning and Construction-Geodesy-Cartography, Health and Veterinary, Chemistry and Technical Chemistry, Trade and Law). We also excluded a few educational fields from our analysis when they were too divorced from market operations (such as theology) to estimate reliable wage equations.

25.The survey was conducted by Analysis Marketing Data (AMD) for the Institute for Information in Education (UIV). The survey does not provide detailed information on each child if there is more than one in a family. We exclude observations with missing values (mostly income). Dropping income from the estimating equations so that these observations can be included does not change the results.

26.We initially analyzed responses for men and women together and independently. No coefficient was ever significant in the men’s equation, however. This suggests that Czech fathers do not extensively participate in their children’s education and schooling, leaving these decisions, as well as active involvement, to the mothers. Given this difference, we present only results for women

27.Although we do not know the district of residence, it is likely that these parents live in districts where there is no effective private competition.

28.We also defined a smaller group of parents with children at secondary school only, but the results are not substantially different.

29.Other variables that were not significant include a dummy if a person finds the local choice of schools poor, municipality-size dummies, marital status.

30.Nationwide approximately 12 percent of new entrants to gymnázia in 1995 enrolled in nonstate schools. Because there were no nonstate schools in the majority of districts, the percentage opting for such schools where they were available was significantly higher.

31.This decline offset a substantially higher student/teacher ratio in 1992 in districts that later attracted competition. Since as we have already seen, these districts were ones that performed worse than expected in obtaining university admission for their graduates, this result strongly suggests that class size is an important input in educational success. Estimation of education production functions for the Czech Republic using this school-level data is a promising area for future research.

32.There is also some evidence that state schools facing declining enrollment may rent their unused space to nonstate schools. This may provide additional revenue that could be used to reduce class sizes.

33.As discussed above, success is measured as acceptances divided by applications weighted by the attractiveness of the universities to which students applied.

34.The difference in support is somewhat larger than these formulae would suggest since public schools are also eligible for capital funds for construction and maintenance from state sources. During the past decade such investment funds added about 10 percent to the level of support for state schools that was not available to nonstate institutions.

35.Downes and Greenstein (1996), Goldhaber (1999), and Figlio and Stone (2001) find that the location of and enrollment in private schools in the United States are heavily influenced by the quality of public schools in an area.

36.Borland and Howsen (1992), Couch, Shughart and Williams (1993), Hoxby (1995), Arum (1996), Borland and Howsen (1996), Dee (1998), McMillan (1999), Brasington (2000), Hoxby (2000), Marlow (2000), Greene (2001), and Grosskopf et. al. (2001) all find such an effect. Contrary results have been reported by Newmark (1995), Simon and Lovrich (1996) and Sander (1999).

37.See Kleitz, et. al (2000) and Teske and Schneider (2001).

38.For recent studies that investigate the impact of vouchers or voucher-like systems in countries other than the United States, see, for example, Toma (1996), West (1997), Carnoy (1998), King, Orazem and Wohlgemuth (1999), McEwan and Carnoy (2000), Mizala and Romaguera (2000), Angrist, et. al. (2001), Chandler (2001), McEwan (2001) and Hsieh and Urquiola (2002).

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