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After, Before and During: Returns to Education in the Hungarian Transition

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After, Before and During:
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Abstract: How valuable are the education and skills acquired under socialism in a market economy? This paper uses data for about 3 million Hungarian wage earners, from 1986 to 1998, to throw light on this question. We find that returns to schooling reach 10 percent early on and remain at this high level. These estimates are larger than for other transition economies, but similar to those for middle-income developing countries. With the gap in average years of schooling unremitting, we argue that the Hungarian stock of human capital is considerably less than the existing figures have led us to believe.

Keywords: Human Capital, Labor Markets, Transition Economies, Hungary.
JEL classification: I20, J20, J24, J31, O15, O52, P20.

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1. Introduction

This paper estimates returns to years of schooling, to types of education and to experience. Over the years, economists have paid a great deal of attention to these matters and professional interest does not seem to have subsided. Mincer, Schultz and Becker laid down the theoretical foundations almost forty years ago. A flurry of empirical research followed taking advantage of the elegant specification that quickly became known as the "Mincer equation," which in its barest form stipulates that (log) earnings are a function of years of schooling, experience and experience squared. Estimates are now available for basically every country in the planet. Psacharopoulos (1994) summarizes this body of research: returns to education typically range from about 6-7 percent for high-income countries to about 11-12 percent for low-income countries. More recently, the literature has turned to issues such as the possible endogeneity of the schooling decision and the role of ability.[1]

This paper estimates returns to years of schooling, to types of education and to experience before and during the transition from centrally planned to market economy. An objective of the socialist experiment was to provide an egalitarian distribution of income. This goal was pursued by a deliberately compressed wage structure. Returns to education and experience were determined centrally and were set low. One hypothesis is, thus, that the liberalisation of labour markets at the outset of the transition would entail increasing returns to education. A competing hypothesis is based on the notion that skills acquired under socialism were excessively specialized and useful only for out-dated technology. Those skills should thus not be rewarded in a market economy and returns to education and to experience should decline during the transition. The empirical literature

[1] For reviews of this literature, see Card (1999) and Harmon, Ooesterbeek and Walker (2000).

addressing these issues is very large. Typically, it estimates returns to one year of education to be around 7 percent and the coefficient on years of potential experience to hover around 3 percent. Those studies with data for more than one year find that returns to education typically increase but that returns to experience are fairly constant.[2]

Despite the voluminous empirical literature on returns to education in transition economies, the number of studies that present results for pre- and post- 1989 is small. The availability of data is the binding constraint as, for example, Labor Force Surveys in transition economies tend to start only after 1991. Data collected before 1989 exist in large quantities. Their shortcoming is that most variables need to be re-coded (consider the differences in the classification of firm ownership). For this paper, primary data was re-coded to standard international classifications to allow us to investigate changes in returns to education and experience before and during the transition from plan to market.

This paper estimates returns to years of schooling, to types of education and to experience from 1986 to 1998for a transition country that has received surprising little attention in the literature: Hungary. Although among transition economies Hungary stands out for its gradualist approach to reform, it is in many ways representative of the economies chosen for the "first wave" of future members of the European Union. The size of the Hungarian population is between that of Poland and Estonia, and almost identical to that of the Czech Republic. It also has a per capita GDP higher than those of Poland and Estonia, but lower than the Czech and Slovenian figures. Labour market developments also place Hungary far from the Estonian aggressive laissez-faire and equally far from the more guarded Slovenian approach.

[2] See Svejnar (1999) and Boeri and Terrell (2002) for reviews of this literature.

The objective of this paper is to throw light on the question "how valuable are the education, qualifications and skills acquired under socialism in a market economy?" Our main conclusions are: (1) that returns to a year of schooling increased by 75 percent from 6.4 percent in 1986 to 11.2 percent in 1998; (2) that primary and vocational education show the smallest, and general secondary education and university show the largest changes in returns from 1986 to 1998; (3) that returns to experience decline early in the transition but show no overall trend as they recover to usual low levels (lower than for Western Europe and developing countries[3]); (4) that returns to education become higher for private sector workers late in the transition (only in 1998); and (5) that returns to education for older workers (age greater than median) are higher throughout.

Our core result shows that returns to one year of schooling are above 10 percent early in the Hungarian transition and stay at this high level throughout. This is in sharp contrast to the rest of the literature, which finds much lower rates of return to schooling, usually of the same order of magnitude of those for Western European economies. Why is that? Consider Portugal and Greece, two of the European Union's lowest per capita GDPs, with populations and demographic structures similar to Hungary. For 1995, estimates of returns to years of schooling are about 6.3 and 9.7 percent for Greek and Portuguese males, and about 8.6 and 9.7 percent for Greek and Portuguese females, respectively.[4] Note, however, how different their stocks of human capital are: Barro and Lee (2001) estimate that, for the population aged 15 and over in 1995, average years of schooling are 5.5 in Portugal and 8.1 in Greece. The same figure for Hungary in 1995 is about 8.8. A puzzle emerges when we contrast the high returns to education to such high

[3] See Harmon, Ooesterbeek and Walker (2000).
[4] The estimates, from specifications comparable to ours, for Greece are from Tsakloglou and Cholezas (1999, 2001) and the estimates for Portugal are from Pereira and Martins (1999, 2001).

stocks of human capital in Hungary. One possible answer, we offer, is that although Hungarian labour market rewards education in the same way as the Greek and Portuguese markets, the size of the economically meaningful stock of human capital in Hungary is considerably below what the existing figures show.

The paper is organized as follows. The next section gives the theoretical background and reviews the literature. Section 3 describes the data and methodology. Section 4 presents our estimates of returns to education and experience in Hungary between 1986 and 1998. Section 5 concludes.

2. Theoretical background and related empirical literature

There are at least two competing hypotheses about the returns to education during transition from centrally planned to market economy. One is based upon the assertion that central planners purposefully undervalued education so as to achieve an egalitarian distribution of income through a compressed wage structure. Returns to education and experience were determined centrally and wages were set below equilibrium. The socialist system was characterized by excess demand for labor. Therefore, one hypothesis is that labor markets reforms at the onset would entail immediate increases in returns during transition.

A competing hypothesis is that the skills acquired from schooling under communism are not well suited for the market economy, and one would expect to see a temporary decline in returns to schooling during transition. Further, because these skills are "excessively specialized" (Laporte and Schweitzer, 1994; Boeri, 2000), they can not be easily transferred across occupations, jobs and sectors. Only after school curricula adjusted to the changes and successful re-training programs are put in place, one would expect increasing returns.

These two hypotheses represent extreme views. Although socialism deliberately compressed the distribution of earnings, it may have been able to reward different skills and experiences in non-pecuniary terms. The anecdotal evidence points to vacations and access to consumer goods as examples of these non-pecuniary rewards. It should be said that with our data we can not explicitly test this possibility (we are also unaware of research that does this), but this doesn't diminish its importance. On the other hand, the notion that skills and experience acquired under socialism are worthless in a market economy overlooks the fact that different skills are transferable to different degrees. Contrast computer skills to a good grasp of relativity theory. Moreover, given the speed of reduction of enrolment rates in vocational secondary education, one wonders what would be the time frame needed to capture these phenomena in a meaningful way. With these caveats in mind, let us turn to the relevant empirical evidence. Given the volume of this literature, this review is unavoidably selective. It gives priority to studies that examine pre and post 1989 years.

We start with the available estimates for rates of return to years of schooling during transition. Flanagan (1998) studies the Czech labor market and reports that the returns to a year of schooling increased from 4.3 percent in 1988 to 5.7 percent in 1996. For males, the change in returns to schooling was from 3.7 percent in 1988 to 4.5 percent in 1996, while for females this change was from 5.1 percent in 1988 to 7 percent in 1996.

Chase (1998) finds that the returns to schooling for men double between 1984 and 1993 in the Czech Republic (from 2.4 percent in 1984 to 5.8 percent in 1993). For Czech women, he finds that the returns to one year of schooling increased from 4.2 percent in

1984 to 7 percent in 1993. Chase also reports estimates for the Slovak Republic. For males, the change was from 2.8 percent in 1984 to 4.9 percent in 1993, while for females it was from 4.4 percent in 1984 to 5.4 percent in 1993.

Rutkowski (1996, 1997) also finds that returns increased in Poland between 1987 and 1992, though by a smaller extent (40 percent increase). The estimate for 1987 is relatively high, at 5 percent per year of schooling increasing to a rate of return of about 7 percent in 1993.

The evidence for the Russian Federation is somewhat contradictory. Brainerd (1998) finds that returns to education increased from 3.1 percent in 1991 to 6.7 percent in 1994 for males and, for females, from 5.4 percent in 1991 to 9.6 percent in 1994. Sheidvasser and Benitez-Silva (1999) find lower and fairly constant estimates for the period 1992 to 1999, at around 4 percent.

In contrast, Krueger and Pischke (1995) find that returns to schooling declined in the former Republic of East Germany between 1988 and 1991. Though, it should be noted that their results are based on data from different survey instruments over the two years. These authors estimate that the returns to one year of schooling decline from 7.7 percent in 1988 to 6.2 percent in 1991. Bird et al. (1994) also find evidence of this decline in the former Republic of East Germany between 1988 and 1991, from 4.4 percent to 4.1 percent.

In summary, most studies find evidence of increasing returns to schooling during the transition from plan to market (Newell and Reilly, 1999). These studies tend to present estimates of the order of 7 percent per year of schooling which is below the average return for middle-income developing countries (Psacharopolous, 1994).

The empirical transition literature also reports returns to different types of education. This is of particular interest in this context because "from the perspective of market economies, overinvestments in vocational-apprenticeship training and underinvestment in university education appear to be the major distortions in human capital formation under central planning" (Flanagan, 1998, p. 300). Therefore, if skills are "excessively specialized" one should observe a distinctive behaviour for returns to vocational-apprenticeship secondary education. The available evidence supports this distinctiveness. For instance, Orazem and Vodopivec (1997) report for Slovenian males that the premium (over incomplete primary) for elementary school increases from 4.4 percent in 1987 to 10.7 percent in 1991, for high school it increases from 31.9 percent to 40.6 percent, for university it increases from 71.5 percent to 94.3 percent, while for vocational it only increases from 16.3 percent to 20.1 percent. These changes are similar for Slovenian females, except that the premium to vocational education shows a smaller increase, from 16.2 to 18.3 percent. Kertesi and Kollo report that, for Hungary, the returns to vocational education also grow much slower than for other types of education. Keane and Prasad (2001) and Norkooiv, Orazem, Puur and Vodopivec (1998) present qualitatively similar evidence for Poland between 1986 and 1996 and for Estonia between 1989 and 1995, respectively. These two papers report that returns to vocational education increase much slower then for other types of education. In contrast, Flanagan (1998) shows that the premium to vocational education more than doubles in the Czech Republic between 1988 and 1996, while that for university increases only by a quarter.

The findings with respect to potential experience diverge less than it is the case with returns to schooling. There is broad consensus across studies that returns to experience are low and declining in transition economies. For instance, Flanagan (1998) estimates that the coefficient on (linear) potential experience in the Czech Republic decreases from 3.6 percent in 1988 to 2.2 percent in 1996 (the same pattern obtains for males and females). Chase (1998) reports that these coefficient estimates for Czech males decrease from 3.7 percent in 1984 to 3 percent in 1993. He notes a similar pattern for Slovakia (3.6 percent to 2.3 percent). For Slovakian females, Chase reports that these coefficient estimates stay constant at 2.5 percent between 1984 and 1993, while it declined from 1.4 percent to 1 percent for Czech females in the same period. Munich, Terrell and Svejnar (1999) find that returns to experience do not increase during the transition in the Czech Republic. For Poland, Keane and Prasad (2001) estimate that these coefficient estimates are 3.5 percent in 1986 and 3 percent in 1996. Rutkowski (1996) estimates that these coefficients on (linear) potential experience were about 3.1 percent in 1987 in Poland but decline to 2.1 percent in 1993. Orazem and Vodopivec (1997) report that for Slovenian males, these coefficients did not change between 1987 and 1991 (1.9 percent). Yet they declined for Slovenian females from 1.8 percent in 1987 to 1.1 percent in 1991. Sheidvasser and Benitez-Silva (1999) also report low values for Russia between 1992 and 1999.

As noted, one of the recent advances in this empirical literature is the use of instrumental variables techniques to address the potential endogeneity of the schooling decision. For transition economies, data availability critically constrains the set of suitable instruments. We are aware of two papers that present IV estimates; Filer, Jurajda and Plánovský (1999) and Sheidvasser and Benitez-Silva (1999). It should be noted that in these studies, although IV returns to education are higher then the values commonly found, they *decline* in the Czech Republic between 1995 and 1997 and in Russia between 1992 and 1999.

In summary, the empirical literature estimating returns to education and to experience in transition economies typically finds

(a) that the returns to one year of schooling in transition economies is about 7 percent, which is below the average for middle-income developing countries;
(b) that returns to one year of schooling tend to be higher for females than for males;
(c) that returns to vocational-apprenticeship secondary education have increased at a slower pace than for most other types of education; and
(d) that returns to experience during the transition stay at about 3 percent, which is low in international perspective.

3. Data and the Empirical Specification

The data used in our analysis come from the Wage and Earnings Survey (WES) of the National Labor Center in Hungary, and it contains information on wages, education, type of employment, and other demographic data. Considerable effort was spent in assuring that variables are coded consistently over time. This was carried out with assistance from the National Labor Center and the Hungarian Central Statistical Office and involved substantial recoding of the data on industrial sector, legal form (ownership) and occupation.[5]

We use data from the five years of 1986, 1989, 1992, 1995, and 1998 to cover the pre- and transition periods. The sample sizes range from a low of 136,829 in 1992 to a high of 933,282 wage earners in 1989. The sample sizes for the three other years are all above 450,000 observations resulting in a pool of more than 2.9 million wage earners.

[5] For details on how consistent definitions of industry, ownership, and occupation codes were obtained for the thirteen-year time frame of our analysis, see Campos and Žlábková (2001).

3.1 Sample Design

For all five years the samples are stratified on the characteristic of whether the wage earner is designated as a manual or non-manual laborer. The sampling units are wage earners and these are selected following a systematic, random selection procedure. The details of the procedures varied across the years in ways that affect the sample weights, but across all years the design is random and the estimates are representative of the sample frame. The sample frame for the years 1986, 1989, and 1992 includes all wage earners in private and public firms with more than 20 employees. The frame changed to include all wage earners in firms with ten or more employees for the 1995 and 1998 samples.

In addition to this change in the sample frame, the procedure for sample selection changed on three occasions. In 1986 and 1989 wage earners were selected following a systematic, random design with a fixed interval of selection. From the list of all manual wage earners in each firm, one observation was randomly drawn and then every seventh employee thereafter was drawn. The same random selection procedure was used for all non-manual wage earners, except that the interval of selection was set to every fifth employee. Kish (1965) notes that systematic sampling is "perhaps the most widely known selection procedure" (p. 113) and suggests that the simplicity of this design reduces the potential for introducing error in the selection procedure.

For the years 1992, 1995, and 1998, the selection procedure changed from a fixed-interval selection procedure to a systematic selection procedure based on date of birth. Again the frame was stratified on manual and non-manual laborers, and then single-stage random draws were made for each labor type. In the case of manual laborers, all workers born on the 5th or 15th day of each month were selected; and in the case of non-manual laborers, all were selected who were born on the 5^{th}, 15^{th}, or 25^{th} day of each month.[6] In this case, rather than a random starting point and fixed interval of selection, the design is based on the fact that date of birth is randomly and approximately uniformly distributed across days of the month.

For the years 1995 and 1998, there is one further element to the sample design that is associated with the changing frame. In these two years, the frame includes smaller firms with total number of employees between 10 and 20. There is an important difference in the sample design associated with this supplementation of the sample frame in that selection occurred in two stages. In the first stage, 20 percent of these small firms (10 to 20 employees) were selected and then in the second stage all wage earners in these firms were selected into the sample.

With the exception of the supplementation to the sample frame in 1995 and 1998, the sample design is a stratified, single-stage, systematic random draw that results in estimates which are representative of the sample frame population.[7] The relevant issue is that the comparison across time suffers the slight problem that the frame changed for the last two years, 1995 and 1998. For the years 1986, 1989, and 1992, estimates are representative of all wage laborers who work in firms that have at least 20 employees. For the years 1995 and 1998, the estimates are representative of all wage laborers who work in firms that have at least 10 employees. If it were the case that firms with 10 to 20 employees reward education in a way that is systematically different from firms with 20

[6] This sampling method gives a representative sample of 7-8% of the workers in such firms (6.4% of manual workers and 9.7% of non-manual workers).

[7] For more details on the properties of a stratified sample see Chapter 3 of Kish (1965) or Chapter 5 of Cochran (1977). For more details on single-stage, systematic random sampling, see Chapter 4 of Kish or Chapter 8 of Cochran.

or more employees, then it is possible that observed changes in returns to education over time result in part from changing the sample frame in 1995.

One way to correct for this change in the sample frame is to exclude from our analysis those firms with 10 to 20 employees in years 1995 and 1998. This would result in estimates for all years that are representative of a population of laborers employed in firms with 20 or more employees. The analysis loses some information on small firms, but the gain is that all estimates are based on a similar population. Unfortunately, our data does not identify these firms and we can only restrict our analysis by excluding firms with less than 50 employees from all years. Our preferred results are based on this restricted sample since they are representative of the same population across all years, but this analysis then ignores all small firms. In order to check the robustness of our results to this constraint on firm size, we also examine our results for the full samples across all years.

3.2 The Empirical Specification

Wage equations are estimated using a standard Mincer equation, taking the form:

$$\ln(w_i) \quad = \quad ß_1 S_i + ß_2 E_i + ß_3 E_i^2 + \mathbf{ß_4 X_i} + e_i \qquad (1)$$

where the i subscript denotes the individual, w is wages, S is years of schooling,[8] E is potential experience, and **X** contains a set of variables to control for institutional and demographic characteristics as well as spatial price differences. Each of these variables is described in more detail below.

The monthly value of wages used in our analysis is the sum of the official base wage received and other payments that the employee receives monthly (rewards given in the reference month, provisions, overtime work, shift work, other special payments, e.g. in mining). In addition, the value of wages includes a pro-rated estimate of irregular payments (1/12 of irregular payments in the previous year).

As a means of assessing the quality of the WES wage data used in this paper, Table 1 compares the monthly value of wages as measured by the WES and data from the International Labour Organization (ILO) and the Hungarian Central Statistical Office. Standard errors are only available for the WES data, so it is not possible to estimate p-values for the test of whether the differences in means are significant. From examining the ratio of the difference to the standard error for the WES data, though, it appears that the differences in mean values are likely to be statistically significant for all years. Due to the large sample sizes however, even qualitatively small differences in the mean values will be statistically significantly different. It is perhaps more useful to note that for all years except 1989, the difference between the two estimates for mean wages is less than (or equal to) one percent of the mean wage. This suggests that the WES wage data is very similar to other estimates of wages for Hungary, and we assert, of reasonably high quality.

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[8] In some specifications S will be a vector of dummy variables representing graduation from different school types, such as primary, secondary, or university.

Table 1: Comparison of Wage Estimates from WES and ILO Data

	1986	1989	1992	1995	1998
Mean Wage, ILO[1]	6,260	10,461	22,294	39,854	68,718
Mean Wage, WES[2]	6,312	9,761	22,466	40,190	69,415
Standard Error of Mean Wages, WES	4	8	41	77	162
Difference in Wages, ILO -WES	52	-700	172	336	697
Difference/Standard Error	13.9	-86.0	4.2	4.4	4.3
Difference as a percent of ILO Mean	0.8%	-6.7%	0.8%	0.8%	1.0%

[1]LABORSTA data, International Labour Organization (ILO) and the Hungarian Central Statistics Office.
[2] Wage and Employment Survey (WES), National Labor Center, Hungary.
Notes: Wages are monthly figures in Forint (HUF). ILO data on mean wages for 1986 and 1989 are based on ISIC (revision 2) codes 2-9, and the WES estimates are restricted to these same industry codes.

Two measures of schooling are examined in this paper. The first measure is a vector of six dummy variables that denote the highest type of completed schooling. The school types include primary, three types of secondary (vocational, technical, and gymnasium or general), college, and university.[9] In 1998, 22 percent of wage earners had only primary schooling or less, while 19 percent had college or university education. Of the remaining 59 percent who completed some form of secondary schooling, slightly less than half these wage earners attended vocational school (28 percent of the total).

The second measure of schooling is an estimate of years of school attainment, which is created by converting the data on highest school type completed into years of schooling. The average value of this variable increased from 9.7 in 1986 to 11.3 in 1998.[10] This measure of school attainment is also used to construct a measure of potential experience, which is the wage earner's age less six years and less years of schooling.

[9] The omitted category is those individuals with less than primary schooling.

[10] The difference between these two means is statistically significant with a p-value of less than 0.0001. School types are mapped into years of school attainment as follows: 5 completed years are given if the worker has "incomplete primary education;" 8 if the worker has "completed primary school," 10 if the worker has "secondary vocational school;" 12 if the worker has "technical secondary school;" 12 if the

The variables designated by **X** include a set of eight dummy variables to control for potential differences across industries.[11] The set of controls also includes dummy variables for the size of the firm. For the sample restricted to only those firms with 50 or more employees, the specification includes a dummy variable for those firms with more than 300 employees. For the full sample, an additional dummy variable is included in the regression that marks those firms with less than 50 employees. To control for the large differences in wages across gender, **X** includes a dummy variable for males.

In addition to the controls for industry, size of firm, and gender; **X** contains a dummy for each of the 19 counties of Hungary and for the capital, Budapest. These spatial-specific, binary variables control for any variation that is specific to Budapest or some particular county. In particular, these dummy variables control for spatial variation in prices, which are likely to be significant with wages and prices in Budapest higher than other regions.

The county dummy variables will also control for region specific differences in labor markets, which are also potentially important given that unemployment rates are relatively lower in Budapest, the counties along the Budapest–Vienna highway, and the counties along the Hungarian–Austrian border. Similarly, the county dummies will control for the potential measurement issue that a year of schooling may result in different levels of human capital accumulation over different regions if there are differences in schooling quality across regions.

The controls for firm size and industry, as well the county fixed effects, greatly reduce the potential for omitted variable bias in our estimation of equation (1). Having

workers has "grammar school;" 15 if the worker has "college;" and 17 if the worker has completed "university." For details on how school types are mapped into years of school attainment, see Campos and Žlábková (2001).

data that has been collected using the same survey instrument over the years of 1986 to 1998 also significantly improves the credibility of measured changes. Frequently, comparisons over time come from different data sources and one is left with the question of whether the change over time is a consequence of actual changes in the population, or is simply the result of changing the survey instrument. These are important advantages to using the WES data.

The disadvantage of using the WES data is that the choice of variables is limited. The data contain no credible instruments for schooling (ie. there are no variables that are correlated with schooling that are also reasonably excluded from the wage equation). For this reason, we do not attempt to control for endogeneity bias when estimating equation (1). Another caveat is that the WES data contain no information on those who are not wage earners, and we therefore do not attempt to control for sample-selection bias. As evidence that this potential bias might not be severe, we note that Saget (1999) finds no evidence of sample-selection bias in determining female labor supply for Hungary in 1992.[12]

Nonetheless, in interpreting the results, one should keep these two potential sources of bias in mind. A somewhat mitigating factor is that our analysis is focused how returns have changed over the pre- and transition years. If the magnitude of these potential biases has not changed over time, then their impact on drawing inferences about change over time is limited.

[11] The eight classifications are: industry, construction, transportation and telecommunications, trade, water, services, health and social services, and public services. The excluded classification is agriculture.
[12] Saget notes that female labor force participation is much higher in Eastern European countries than OECD countries. She suggests that female labor is viewed more as an economic necessity and less as a choice, and this mitigates the severity of the selection bias.

4. Results

To examine how returns to schooling have changed from 1986 to 1998, county fixed-effects estimates of equation (1) are provided in Table 2. Panel A of Table 2 provides estimates for all firms with more than 50 employees, and Panel B lists estimates for the full sample of all firms.

The standard errors listed in all regression tables are corrected for heteroskedasticity of unknown form through use of the 'sandwich variance estimator'. This commonly used estimator was first introduced by Huber (1967) and White (1980) and has the advantage of providing consistent estimates of the variance-covariance matrix when errors are heteroskedastic or if the residuals exhibit some form of dependence.[13] The disadvantage, as noted by Kauermann and Carroll (forthcoming), is that the sandwich variance estimator is inefficient and often times results in estimated standard errors that are too conservative (large). Given the large sample sizes, the cost of the 'sandwich' or robust variance estimates are not qualitatively important, and the benefit of consistency is desirable.

4.1 Returns to Years of Schooling

The results from Panel A of Table 2 show that the returns to a year of schooling increased by 75 percent from a return of 6.4 percent in 1986 to 11.2 percent in 1998.[14]

[13] An important advantage of the WES design is that the sample was selected in a single stage, and there is therefore no need to correct estimates of the sampling variance for any sort of design-induced dependence. Scott and Holt (1982) show that this type of correction for dependence can be quite large for multi-stage sample designs.
[14] This increase in returns is statistically significant.

This dramatic increase in returns to schooling is particularly noteworthy given that the average level of schooling has also increased over this time period.

Recall that Panel A restricts our analysis to only those firms with 50 or more employees. While this restriction corrects for the change made to the sample frame in 1995, it comes with the disadvantages that a large portion of the sample has been dropped, and our analysis excludes wage earners working in smaller firms. By examining the full sample in Panel B of Table 2, it is apparent that the restriction on the sample does not affect the qualitative nature of the results.

Panel B shows that the return to schooling for wage earners from firms with 20 or more employees increased from 6.2 percent in 1986 to 11.2 percent for wage earners from firms with 10 or more employees in 1998. While there are small differences across the panels from year to year, the overall change from 1986 to 1998 is almost identical over the two panels. For the remainder of the analysis, only those estimates based on the samples of wage earners from firms with 50 or more employees are presented, but the results for the sample of all wage earners continue to be qualitatively very similar.

Table 2: Returns to Years of Schooling, 1986-1998
Fixed-effects Estimation of Equation (1)

Panel A: Firms with 50 or more employees	**1986**	**1989**	**1992**	**1995**	**1998**
Years of Schooling	**0.064**	**0.085**	**0.100**	**0.112**	**0.112**
	(0.0003)	(0.0002)	(0.0007)	(0.0002)	(0.0002)
Dummy: Male=1	0.247	0.255	0.157	0.112	0.106
	(0.0016)	(0.0013)	(0.0040)	(0.0014)	(0.0016)
Potential Experience	0.034	0.032	0.031	0.035	0.033
	(0.0002)	(0.0002)	(0.0006)	(0.0002)	(0.0002)
Experience Squared / 100	-0.047	-0.044	-0.038	-0.044	-0.039
	(0.0005)	(0.0004)	(0.0015)	(0.0005)	(0.0006)
Firm Size: 300+ Employees	-0.008	-0.005	-0.019	-0.054	-0.061
	(0.0033)	(0.0013)	(0.0050)	(0.0012)	(0.0013)
Observations:	149,274	383,720	48,261	371,882	334,207
R-squared:	0.52	0.50	0.45	0.52	0.50
Panel B: All Firms					
Years of Schooling	**0.062**	**0.074**	**0.095**	**0.109**	**0.112**
	(0.0001)	(0.0002)	(0.0004)	(0.0002)	(0.0002)
Dummy: Male=1	0.267	0.286	0.175	0.123	0.118
	(0.0007)	(0.0009)	(0.0023)	(0.0012)	(0.0013)
Potential Experience	0.032	0.024	0.031	0.032	0.030
	(0.0001)	(0.0001)	(0.0004)	(0.0002)	(0.0002)
Experience Squared / 100	-0.044	-0.034	-0.038	-0.040	-0.035
	(0.0002)	(0.0003)	(0.0009)	(0.0004)	(0.0005)
Firm Size: 20-50 Employees	0.046	0.014	0.082	0.059	0.144
	(0.0009)	(0.0010)	(0.0026)	(0.0014)	(0.0016)
Firm Size: 300+ Employees	-0.003[a]	-0.013	-0.016	-0.054	-0.071
	(0.0030)	(0.0013)	(0.0050)	(0.0012)	(0.0013)
Observations:	831,407	933,282	136,829	529,928	487,160
R-squared:	0.48	0.39	0.43	0.52	0.50

[a] The only statistically insignificant point estimate—Panel B, 1986, Firm Size: 300+ Employees.

Notes: Dependent variable is the log of monthly wages. Standard errors, in parentheses, are robust to heteroscedasticity of unknown form. The eight industry dummy variables are jointly significant and are excluded from the table. County, fixed effects are also jointly significant. All listed point estimates are significant with a p-value of less than or equal to 0.001 except for two: The 1986 dummy variable for firm size greater than 300 employees has a p-value of 0.013 in Panel A and 0.28 in Panel B.

As discussed above, there are (at least) two competing hypotheses about the returns to education during transition. One assertion is that the skills acquired from schooling under communism are not well suited for the market economy, and one would expect to see a temporary decline in returns to schooling during transition. Once school curricula adjusted to the changes, one would then expect to see increasing returns. A competing hypothesis is that central planners undervalued education, and one would expect to see immediate increases in returns during transition. The results from Table 1 strongly support the second hypothesis.

Table 3 compares the change over time of returns to school years by various breakdowns. In Panel A, returns are separately estimated for wage earners in private and public firms. The code identifying the legal status of firms is first available for 1992, when returns were slightly higher in public firms. By 1998, this reversed and returns were greater in private firms. This result potentially indicates a growing competitiveness of private firms in rewarding more highly skilled workers.

Panel B examines returns by gender and shows that in 1986 women had significantly (both statistically and qualitatively) greater returns to schooling than men, but by 1998 this difference in returns was eliminated and men had slightly higher returns to schooling. This result is perhaps not so surprising given that in 1986 average female wages were equal to 74 percent of average male wages. Given that returns in Tables 2 – 4 are estimated in percentages of wages, an equivalent change in the level of returns would mean a higher percentage change in returns for the lower base wage. By 1998 the ratio of female to male wages increased to 86 percent, and by this time the gender gap between (percentage) returns to schooling was eliminated.

Panel C separates the sample into young and old wage earners, where old is defined as an age greater than the median in each year.[15] Perhaps one of the more striking results is that returns to schooling for the older wage earners are greater than the returns for younger workers.[16] This result is true over all the years and the difference between returns has been a fairly constant two-percentage points. It is noteworthy that over this time period, average wages for older workers were also greater than for younger wage earners.[17] While this result doesn't refute a hypothesis that younger workers adjust more quickly to the changes associated with transition, it certainly does argue against it. One explanation for the result could be that the quality of education declined during the 1970s and 1980s such that a year of schooling for younger workers attained during this period was of less value.

[15] Median age increased over the reference period with a value of 38 years in 1986, and growing to 41 years by 1998.
[16] Note that this is true when controlling for gender and potential experience.
[17] In 1986, young workers earned an average wage equal to 85 percent of the average wage earned by older workers. By 1998, this ratio dropped to 80 percent.

Table 3: Comparing Private-Public, Male-Female, Old-Young Returns, Fixed-effects Estimation of Equation (1)

Firms with 50 or **more employees**	**1986**	**1989**	**1992**	**1995**	**1998**
Panel A: Returns by Legal Status					
Private Firms	NA	NA	0.097	0.113	0.131
			(0.0012)	(0.0011)	(0.0013)
Public Firms	0.064	0.085	0.103	0.113	0.112
	(0.0003)	(0.0002)	(0.0007)	(0.0002)	(0.0002)
Difference (Private-Public)			*-0.006*	*0.000*[a]	*0.019*
R^2 (Private / Public)	0.52	0.50	0.37 / 0.59	0.34 / 0.58	0.34 / 0.59
Panel B: Returns by Gender					
Male Wage Earners	0.061	0.077	0.097	0.105	0.115
	(0.0004)	(0.0004)	(0.0011)	(0.0004)	(0.0005)
Female Wage Earners	0.071	0.091	0.103	0.115	0.111
	(0.0004)	(0.0002)	(0.0009)	(0.0002)	(0.0002)
Difference (Male-Female)	*-0.010*	*-0.014*	*-0.006*	*-0.010*	*0.004*
R^2 (Male / Female)	0.44 / 0.49	0.35 / 0.51	0.41 / 0.46	0.45 / 0.55	0.45 / 0.53
Panel C: Returns by Age					
Old (Greater than median)	0.080	0.098	0.116	0.122	0.123
	(0.0004)	(0.0003)	(0.0010)	(0.0003)	(0.0003)
Young	0.057	0.081	0.089	0.103	0.101
	(0.0004)	(0.0003)	(0.0011)	(0.0003)	(0.0004)
Difference (Old-Young)	*0.023*	*0.017*	*0.027*	*0.019*	*0.022*
R^2 (Old / Young)	0.46 / 0.34	0.47 / 0.40	0.46 / 0.35	0.54 / 0.44	0.52 / 0.41

[a] The only statistically insignificant point estimate—Panel A, 1995, Private-Public Difference in returns.

Notes: Dependent variable is the log of monthly wages. Standard errors, in parentheses, are robust to heteroscedasticity of unknown form. The remaining results from estimation of equation (1) are suppressed for the sake of brevity. All point estimates for the experience and gender variables are statistically significant. The firm-size and industry dummies are jointly significant as well as the county, fixed effects. All listed parameters are statistically significant with a p-value less than 0.001 except for one: The 1995 difference between returns private and public firms is not statistically significant. The smallest t-statistic for all other *difference* estimates is 4.2, while all t-statistics for the point estimates are greater than 80.

4.2 Returns by Type of School

By assigning years of schooling to persons who have attended different types of schooling, an implicit assumption is being imposed that a year of vocational schooling, for example, is the same as a year of general secondary schooling. This assumption made in section 4.1 simplifies interpretation of the results and helps to clearly demonstrate that returns to schooling have been increasing, but it may mask some important information in terms of what types of schooling are being more heavily rewarded in the labor market.

There are several potential explanations for the varying returns to types of schools during transition. One explanation is that different types of school produce different skill sets and these skills may be more or less well suited to the needs of the new market economy. Another related explanation is that the government traditionally steered students into certain types of schools and this planned aspect of the economy no longer provided the correct mix of skills.

Both of these explanations are based on the idea that the changing market environment produced changes in the market value of certain skills. These hypotheses ignore the fact that under the planned economy, returns to skills were set by planners and not determined by the market. Prior to transition, wage setting was used to favor certain industries and certain types of labor. To this end, labor that had been trained in technical and vocational schools and was involved in the production of certain goods tended to be more highly valued, while labor that had been more academically trained and less likely working in the physical production of goods was less highly valued. Presumably the market economy rewards the value added by labor and is indifferent as to whether the added value is in terms of some physical commodity or, for example, in terms of some service.

Table 4 lists the returns to each of the six types of schooling (ranging from primary to university) over five points in time from 1986 to 1998. The listed estimates are those based on the consistent sample of wage earners in firms with 50 or more employees.[18] One result to note is that the returns to primary and vocational schooling were virtually unchanged between 1986 and 1998. This result is striking given the dramatic increase in the return to a year of schooling over this time period.

Table 4: Returns to School Types, 1986-1998
Fixed-effects Estimation of Equation (1)

Firms with 50 or ***more employees***	**1986**	**1989**	**1992**	**1995**	**1998**	**% Change 1986-1998**
School Types						
Primary	0.109	-0.083	0.102	0.105	0.107	-2%
	(0.0038)	(0.0047)	(0.0096)	(0.0052)	(0.0065)	
Secondary, Vocational	0.247	0.077	0.305	0.301	0.259	5%
	(0.0041)	(0.0050)	(0.0138)	(0.0055)	(0.0068)	
Secondary, Technical	0.417	0.341	0.275	0.631	0.597	43%
	(0.0042)	(0.0050)	(0.0103)	(0.0054)	(0.0068)	
Secondary, General	0.339	0.239	0.541	0.569	0.541	60%
	(0.0044)	(0.0050)	(0.0100)	(0.0055)	(0.0069)	
College	0.638	0.544	0.848	0.910	0.885	39%
	(0.0049)	(0.0049)	(0.0103)	(0.0054)	(0.0067)	
University	0.752	0.743	1.059	1.097	1.110	48%
	(0.0045)	(0.0052)	(0.0118)	(0.0057)	(0.0071)	
Observations:	149,274	383,720	48,261	371,882	334,207	
R-squared:	0.52	0.51	0.48	0.53	0.51	

Notes: Dependent variable is the log of monthly wages. Standard errors, in parentheses, are robust to heteroscedasticity of unknown form. The remaining results from estimation of equation (1) are suppressed for the sake of brevity. All point estimates for the experience and gender variables are statistically significant. The firm-size and industry dummies are jointly significant as well as the county, fixed effects. All listed parameters are statistically significant with a p-value less than 0.001. The point estimate for primary schooling in 1992 has the smallest t-statistic with a value of 10.6.

[18] Results for the full samples are available from the authors upon request and are similar in that the percentage change to secondary general education is the greatest of the period from 1986 to 1998. Similarly, when using the full sample, the rate of change for each school type follows the same patterns.

Another result shown in Table 4 is that the largest percentage change in returns is for those wage earners who completed secondary general education. Their returns to schooling increased by 60 percent between 1986 and 1998. The next largest change in returns to schooling is a 48 percent increase for those who completed university, followed by a 43 percent increase in returns for those completing secondary technical schooling. The result that the returns to general secondary schooling increased the most during transition is consistent with the belief that the planned economy under-valued (relative to the market economy) labor used in the production of non-physical goods and services. It also suggests that human capital resulting from general education is valuable in terms of adjusting to a changing environment, which adds to the discussion of the relative benefits of general education versus training in specific skills.

Figure 1 summarizes the results from Table 4 and focuses on the returns to four types of schooling that had both the highest level of returns and the most change over time. From this figure, the relative ordering becomes clear with the returns to university training being greater than college, college greater than secondary technical, and secondary technical greater secondary general education.[19] While the return to technical schooling is slightly greater than the return to secondary general schooling, the difference in these returns declined between 1986 and 1998.

The figure also clearly highlights that the returns to general education (secondary general, college, and university) increased more quickly in response to transition, even during the period of recession from 1990 to 1994. An explanation for the different pattern of change in returns over time could be that the value of training in specific skills (or the

[19] The one exception to this was in 1992, when the returns to completing secondary general education were greater than for secondary technical schooling.

training at vocational and technical schools) is more dependent on market fluctuations than general training. When skills training is well targeted to the specific demands of the market, then returns are high; when market conditions change, there will be a lag before the curricula can adjust to provide the correct mix of skills.

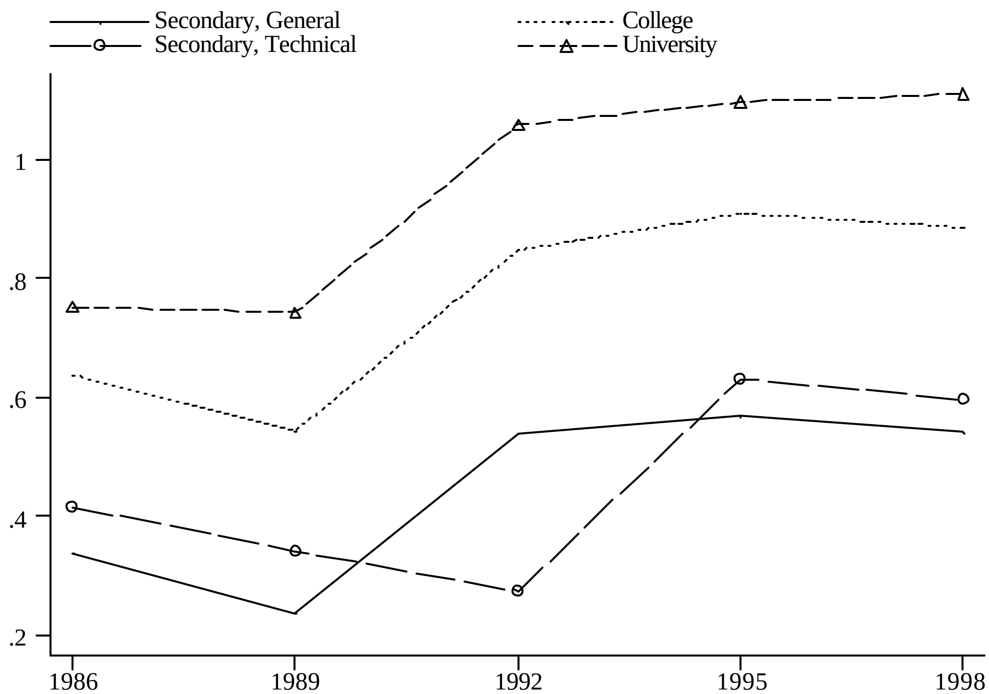

Figure 1: Returns to School Types, 1986-1998
(Summary Figure of Results from Table 4)

5. Conclusions

This paper uses data for about 2.9 million wage earners from 1986 to 1998 in Hungary to support the following five conclusions:

(1) returns to a year of schooling increased by 75 percent from 6.4 percent in 1986 to 11.2 percent in 1998;

(2) primary and vocational education show the smallest, and general secondary education and university show the largest changes in returns from 1986 to 1998;

(3) returns to experience decline early in the transition but show no overall trend as they recover to usual low levels (lower than for Western Europe and developing countries);

(4) returns to education become higher for private sector workers late in the transition (only in 1998); and

(5) returns to education for older workers (age greater than median) are higher throughout the transition.

Our core result is that returns to schooling are large throughout the Hungarian transition, at around 10 percent and above since 1992. These returns are much larger than those available for other transition economies and for Western Europe, and are in line with returns to education in middle-income developing countries. We offer three main reasons for this discrepancy: (a) our data was collected using the same survey instrument over the years of 1986 to 1998, covering pre- as well as transition years; (b) our data was painstakingly re-coded to current standard international classifications to minimize errors in comparisons over time; and (c) we favoured an econometric approach that is attentive to the limitations of the data and of the sampling procedures.

What are then the implications from our core result? We believe it lends support to the view that the stock of human capital, one of the few positive legacies from communism, is misleadingly large. It suggests that those figures are inflated. The puzzle that emerges by contrasting the high years of schooling for the average worker to these high rates of return to years of schooling throughout the transition may be disentangled by considering the possibility that the market is rewarding just a fraction of the total stock of human capital. In other words, the economically meaningful stock of human capital is much smaller then the existing figures have led us to believe. These stocks seem to be smaller than for the richer European countries and in line with some of the poorest members of the European Union (as well as with richer developing countries). In this light, the disappointing performance in terms of rates of economic growth and delays in catching-up with Western Europe should be much easier to comprehend.

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