SOCIOLÓGIA
SLOVAK SOCIOLOGICAL REVIEW
Časopis Sociologického ústavu
Slovenskej akadémie vied
Ročník 38, 2006, číslo 3

Socialist Egalitarian Policies and Education Inequality in Central Europe after World War II[1]

Martin Kreidl[2]
Department of Sociology, Faculty of Philosophy and Arts, University of West Bohemia, Plzeň

Socialist Egalitarian Policies and Education Inequality in Central Europe after World War II This paper investigates the effect of 'Communist Affirmative Action' on inequality in access to secondary and post-secondary education in five former socialist countries of Central and Eastern Europe between 1948 and 1989. I argue that earlier research failed to identify any periods of reduced inequality in former socialist countries because it employed inadequate definitions of both the dependent and independent variables. I correct these inaccuracies and I investigate data from the 'Social Stratification in Eastern Europe after 1989' survey. I am indeed able to document that inequality in access to education declined during the periods of the most extreme Communism in the early 1950s and, in some countries, also during the early 1970s.
Sociológia 2006, Vol. 38 (No 3: 199-220)

Key words: *education inequality; equal opportunity; socialism; affirmative action; Central and Eastern Europe*

A Persistent Question: Did 'Communist Affirmative Action' Reduce Socioeconomic Inequality in Access to Schooling?

Affirmative action, and the bearing it has on educational opportunity, has become the focus of an ongoing and heated debate among both policy-makers and the general public. During the 20th century, socialist states addressed this thorny issue and claimed that their egalitarian policies reduced socioeconomic inequality in access to education. Many radical egalitarian reforms were introduced during the Communist period to promote equal access to education. Fees were abolished and the two-tiered system of private and public schools was eliminated. 'Communist Affirmative Action' was introduced in order to promote educational opportunities among lower-class children, and negative selection criteria were applied to eliminate the advantages of the pre-communist intelligentsia, elites, petty-bourgeoisie, and private farmers as part of a *déclassement* campaign (Hanley, E., 2001; Wheeler, G. S., 1973; Connelly 1997, 2000; Kreidl, M., 2004; Hanley, E. – McKeever, M., 1997; Simkus, A. A., 1981; Simkus, A. – Andorka, R., 1982; Róbert, P., 1991). After the Communists had seized power, school admission practices began to include a complex screening of a student's „talent, interest in the chosen field, class origin, civic and moral considerations, social and political activism of the parents, and the result of the admission examination" (Ulč, O., 1978, 429).

Yet, despite these massive interventions into the system of social stratification, sociologists have so far found little evidence that socioeconomic inequalities in access to secondary and post-secondary education declined in the Soviet Bloc countries as a result of those policies (see e.g. Hanley, E., 2001; Hanley, E. – McKeever, M., 1997; Heyns, B. – Bialecki, I., 1993; Matějů, P., 1993; Nieuwbeerta, P. – Rijken, S., 1996; Róbert, P., 1991; Simkus, A. – Andorka, R., 1982; Szelényi, S. – Aschaffenburg, K., 1993)[3]. The literature on this topic offers a number of substantive explanations why this is so. For instance Szelényi and Aschaffenburg (1993) maintain that bribery and informal contacts with school administrators, bureaucrats, and Communist Party representatives helped create a 'culture of subversion' that inevitably led to the failure of the egalitarian policies (cf. Fiszman, J. R., 1972). The political ideology was, the argument goes, insufficiently enforced and too many exceptions were permitted so that the situation during Communism did not markedly differ from the pre-communist period.

Other authors rely mostly on the concept of cultural capital to explain why the upper classes of the pre-socialist period successfully maintained their advantages even during the Communist rule so that we observe no decline in the estimated effects of measures of socioeconomic background on the odds of success in educational transitions (see e.g. Hanley, E. – McKeever, M., 1997). Yet another stream of theoretical argument elaborates the concept of a 'new class' and maintains that the new Communist elites wanted to secure educational privileges for their offspring regardless of the official policies and ideology and thus effectively acted to forego or dismantle them (Konrad, G. – Szelenyi, I., 1979; Matějů, P., 1993; Parkin, F., 1971). As a result we again do not observe any

[1] An earlier draft of this paper was presented at the Spencer Foundation Spring Forum in Montréal, Canada in April 2006 and benefited greatly from the comments of the workshop participants. This paper was written with the support of the grant 'Changes in the educational stratification over the life course and the impact of political institutions on education inequality in Central and Eastern Europe after 1945" awarded to the author by the Czech Ministry of Education, Youth, and Sports (grant number 1K05010). The author would like to thank Donald J. Treiman and Robert D. Mare for their help while working on this research and the two anonymous reviewers of Sociológia/Slovak Sociological Review for helpful comments and suggestions on an earlier draft of this paper.
[2] Address: PhDr. Martin Kreidl, M.A., Ph.D., Chair, Department of Sociology, University of West Bohemia, Sedláčkova 15, Room 212, 306 14 Plzeň, Czech Republic. Tel.: +420-37-7635656, fax: +420-37-7635652, e-mail: kreidlm@kss.zcu.cz

[3] One exception to the rule is Deng and Treiman's (1997) finding that inequality in access to education declined significantly in China during the Cultural Revolution. Yet, it occurred in such a peculiar historical and political period that it was exceptional even among the socialist countries.

decline in education inequality between social classes in years of the Communist regime.

It seems, therefore, that there are good reasons to believe that the 'Communist Affirmative Action' in education had to be unsuccessful and, indeed, most authors accepted that conclusion as definitive. Nonetheless, I argue that this conjecture may be premature, because previous research has suffered from three major conceptual and measurement flaws that might have concealed existing policy effects. First, I propose that previous research paid insufficient attention to the specification of the dependent variable. Most notably it partially or completely ignored horizontal stratification of school types within levels of schooling, which was particularly pronounced at the secondary level in the countries of the former Soviet Bloc. Second, it focused too narrowly on the highest level of schooling completed and ignored the path leading to and timing of educational attainment. Third, prior research paid only limited attention to the timing of policy changes in former socialist countries, and thus applied inaccurate definitions of one of the key explanatory variables – namely of cohort. Below I develop these arguments in more detail. I conclude that it is worth revisiting the issue once more, correcting for some of the measurement and conceptual shortcomings and re-examining the results, before we accept the conclusion that 'Communist Affirmative Action' in education really failed to achieve greater equality.

Proposed Modifications of Previous Research

All previous studies of educational stratification in former socialist nations employed simplified views of the system of educational institutions and *failed to differentiate between types of secondary schools*[4]. Socialist education systems offered three major branches of study at the secondary level: vocational, professional, and academic. *Vocational secondary schools* lasted between two and three and a half years, trained students for a particular semi-skilled and skilled manual profession, and offered little general training. *Professional secondary schools* lasted four years and offered a curriculum with a balanced representation of general education and vocational preparation in areas such as electrical and civil engineering, agronomy, nursing or accounting. Professional schools were intended to prepare students for immediate labour market entry upon graduation, but provided the necessary formal certificates to apply to a university. Lastly, *academic secondary schools* put traditionally greater emphasis on education in humanities and general subjects and were meant to be preparing students for tertiary education.

[4] I am well aware that there are often very good reasons to simplify and ignore some country-specific institutions. The reasons can be data-related or design-related. For instance, the Shavit and Blossfeld's (1993) volume required that authors follow a common analytic template to make results comparable across countries, and, in some cases, over time. While simplifications are clearly warranted in some situations, they may be sub-optimal for other purposes.

A review of previous research in this area reveals that most authors simplified the horizontal structure of education in former socialist nations to a significant degree. For instance, Matějů (1993) and Hanley (2001) alike limited their models of educational attainment in Czechoslovakia to merely two transitions. In the first transition they distinguished between primary and vocational secondary education on the one hand (failures), and complete secondary education at either professional or academic secondary schools on the other (successes). In the second transition they modelled 'entry' into university (defined as university degree vs. no degree). Heyns and Bialecki (1993), Nieuwbeerta and Rijken (1996), and Szelényi and Aschaffenburg (1993) modelled educational attainment at the secondary level using two 'transitions'. The first 'transition' distinguished between primary and any secondary school, and the second 'transition' differentiated between complete secondary education – i.e. a professional or an academic secondary school – and lower secondary (vocational) training. The common design employed in those studies fails to differentiate commonly recognized types of complete secondary education (academic and professional) and may thus suffer from model misspecification.

Simkus and Andorka (1982) and Hanley and McKeever (1997) achieved greater detail in modelling secondary education under socialism, yet their conceptualizations are unsatisfactory, too. Simkus and Andorka (1982) combined vocational and professional schools and distinguished them from no schooling or academic secondary schools. Hanley and McKeever (1997) employed a multinomial logistic regression to inspect the transition to professional and to academic secondary schools in contrast to not entering either of the two. Their comparison category merges vocational schooling or no schooling and thereby overlooks the considerable dissimilarity between no secondary schooling and vocational training.

Consequently, earlier research has, to a lesser or greater extent, overlooked the large dissimilarity between the students in the individual tracks with respect to their socioeconomic background (cf. Breen, R. – Jonsson, J. O., 2000). Moreover, while scholars verbally recognized possible variations in educational policies across different types of secondary schools and stressed – among other things – that quotas for working class youth were only established at academic secondary schools (Hanley, E., 2001; Hanley, E. – McKeever, M., 1997; Matějů, P., 1993; Simkus, A. – Andorka, R., 1982), they paid little attention to them in their actual statistical modelling work. As a consequence, schools with policy effects – such as academic secondary schools – and schools with no effects – such as professional secondary schools – might have been analyzed together producing an overall ‘no effect’ conclusion, where this was not warranted.

There has been *insufficient attention to the precise timing of policy changes* at both secondary and tertiary levels, leading to a potentially serious understatement of period differences. Previous research relied on a small number of birth cohorts – three to five in most studies (Hanley, E., 2001; Hanley, E., – McKeever, M., 1997; Heyns, B. – Bialecki, I., 1993; Matějů, P., 1993; Nieuwbeerta, P. – Rijken, S., 1996; Róbert, P., 1991; Simkus, A. – Andorka, R., 1982; Simonová, N., 2006; Szelényi, S. – Aschaffenburg, K., 1993) – to determine any historical variation in the educational attainment process. Unfortunately, cohorts were typically defined without any explicit reference to institutional and historical developments and were therefore not well suited to identify any policy effects. There is already some evidence suggesting that there was some significant historical variation in the processes and consequences of Communist egalitarian policies (Kreidl, M., 2004)[5].

Researchers have so far focused too narrowly on the ultimate educational attainment, and as a result they *have ignored not only the timing of educational transitions, but also the trajectory that led people to their degrees* (see e.g. Hanley, E., – McKeever, M., 1997; Heyns, B. – Bialecki, I., 1993; Matějů, P., 1993; Nieuwbeerta, P. – Rijken, S., 1996; Róbert, P., 1991; Simkus, A. – Andorka, R., 1982; Simonová, N., 2006; Szelényi, S. – Aschaffenburg, K., 1993, Wong, R. S-K., 1998). By doing so, scholars may have missed a subtle, but important, form of discrimination – the late start of children of the former bourgeoisie and intelligentsia at obtaining a higher education (cf. Kreidl, M., 2005a, 2005b). If this is the case, then our analyses should first focus on direct school to school transitions, which – as many authors implicitly agree – were potentially impacted by Communist preferential policies more strongly than other transitions. Only if we successfully demonstrate that inequality in direct school to school progressions was reduced during Communism, may we proceed to study why and how these effects did not persist over the entire life of school cohorts.

To summarize, I believe that after correcting for the above-mentioned conceptual and measurement shortcomings, I might still be able to show that socioeconomic inequality in access to secondary and post-secondary schooling indeed declined in former socialist nations. This effect might only exist during the years of the most vigorous Communist egalitarianism, though.

[5] Alternatively, one could claim that while the official policy per se remained rather unaltered, it was emphasized and followed to a different degree as the nature of the Communist regime changed. Obviously, these two competing interpretations result in the same implications for the analyses presented below and would not require that the models be re-specified and cohorts redefined. Nonetheless, interpretations should be modified accordingly.

Data, Variables, and the Modelling Strategy

The data used in these analyses were drawn from the ‘Social Stratification in Eastern Europe after 1989’ survey, which was conducted in 1993 in six post-Communist countries: Bulgaria, the Czech Republic, Hungary, Poland, Russia, and Slovakia (see Treiman, D. J. – Szelényi, I., 1994 for details). However, due to a small yet important deviation in the Russian questionnaire the comparability of the Russian educational data is questionable and I decided not to use in this paper.

The education roster of this survey contains all the information required to define the dependent variables. It included a list of all schools that the respondent ever attended, the year attendance began and ended, and information on whether the course of study was completed successfully, i.e. the relevant certificate was obtained. The data set from the selected countries contains a total of 23,957 completed interviews, of which only 17,942 respondents completed primary education and made the decision about secondary school attendance during the Communist period (i.e. between 1948-1989), and who can therefore be considered in the analysis of secondary school entry. Similarly, 7,882 respondents obtained complete secondary education during Communism and were at risk of progressing to the tertiary level[6]. However, some respondents were lost because the survey did not record their gender. Therefore, only 17,935 primary school graduates and 7,882 secondary school graduates are used in the analyses.

I limit the analysis to the Communist period mostly for methodological reasons. Because the education systems were heavily reformed in all countries under considerations immediately after the Communists had seized power in the late 1940s, it would be extremely difficult – if not impossible – to standardize the dependent variables across the old and new education systems (cf. Kreidl, M., 2004, 2005b). Hence the comparison of the Communist and pre-Communist periods seems unfeasible. The comparison of the Communist and the post-Communist is hampered by the lack of sufficiently rich, yet comparable high-quality data. The survey data I am using here were collected in 1993 and are quite thin in the post-1989 period as too prevent any serious and robust analyses[7].

[6] I only count students who finished complete secondary education following a standard, ‘normative” path as defined within the system. I recognize that a substantial fraction of students in each cohort proceeded through the education system following a rather non-standard trajectory. One such instance is students who returned to professional secondary schools after graduating from vocational secondary schools. I do not include students from the non-standard trajectory into the risk set because the choice of a non-standard educational trajectory itself might be reflective of a politically motivated intervention (see e g. Kreidl, M., 2005b, 2005c). Moreover, non-standard education careers and their stratification during Communism were studied elsewhere (Kreidl, M., 2005c).

[7] Other scholars have carried out analyses including also the post-Communist and pre-Communist period in the Czech Republic recently (see e.g. Simonová, N., 2003; Matějů, P. – Řeháková, B. – Simonová, N., 2004), but these analyses required significant compromises on the side of data quality and comparability.

The dependent variables

I adopt Mare's transition model of educational stratification in my analyses (Mare, R. D., 1981) and extend it to encompass also horizontally stratified education systems (cf. Breen, R. – Jonsson, J. O., 2000; Kreidl, M., 2004). I distinguish four possible outcomes for the analysis of secondary school entry, namely no secondary education, vocational secondary education, professional secondary education, and academic secondary education. Because it is much harder to develop a clear-cut and internationally comparable classification of the horizontal structure of tertiary education, I use a simple dichotomous dependent variable distinguishing respondents who entered university and respondents who did not attend college. For reasons outlined above I only consider direct transitions from primary to secondary school and from secondary to tertiary education as successes. Direct transitions mean transitions in which graduation from the previous level and enrolment at the subsequent level occurred during the same calendar year. Typically, students would complete primary education in June and start secondary schooling in September and would graduate from secondary education after the spring semester and enter the university in the fall of the same year.

Table 1 displays the distribution of respondents in types of secondary schools overall as well as by country. Altogether, 24% of primary school graduates did not continue their education at the secondary level in the year of their primary school graduation, 34% chose a vocational school, 24% entered a professional secondary school, and 18% enrolled in an academic secondary school (see the last column in Table 1). There are some differences in direct secondary school enrolment patterns between countries, though. No direct entry into secondary education was most common in Bulgaria (32%) and Hungary (32%), and least common in the Czech Republic (17%) and Slovakia (19%). Among direct school entrants vocational secondary education was the most frequent choice in all countries but Bulgaria, being most widespread in the Czech Republic (47%) and Slovakia (41%). On the other hand, academic secondary schools were the least common option among students at risk, enrolling 14% of students in the Czech Republic and Slovakia and 18% of students in Hungary and Poland. Bulgaria differs again from the other four nations with 29% of students choosing the academic track (see Table 1).

When we turn our attention to the tertiary level, we see that overall 19% of students at risk entered university immediately after they graduated from secondary education (see Table 2). The percentage of each secondary education graduation cohort directly continuing their schooling at the tertiary level was somewhat lower in Bulgaria (13%), Hungary (15%), and Poland (16%) than in Slovakia (24%), or the Czech Republic (26%, see Table 2). I do not report descriptive statistics regarding delayed enrolments at both secondary and tertiary levels here, as they were reported in Kreidl (2005a) and the interested reader is referred there for more details.

Table 1: **Percentage distribution of first secondary school attended by country, Central and Eastern Europe, 1948- 1989. Number of cases in parentheses. Total N= 17,935.**

	Bulgaria	Czech Republic	Hungary	Poland	Slovakia	All Countries
None	32%	17%	32%	25%	19%	24%
Vocational secondary	5%	47%	32%	40%	41%	34%
Professional secondary	35%	23%	17%	17%	26%	24%
Academic secondary	29%	14%	18%	18%	14%	18%
Total	101% (3,274)	101% (4,339)	99% (3,434)	100% (3,048)	100% (3,840)	100% (17,935)

Note: Only direct school to school transitions count, delayed transitions counted as failures.

Table 2: **Percentage distribution of university enrolment by country, Central and Eastern Europe, 1948- 1989. Total N= 7,882.**

Country	Percent entering university	Number of cases at risk
Bulgaria	13%	2,441
Czech Republic	26%	1,675
Hungary	15%	1,168
Poland	16%	973
Slovakia	24%	1,625
All countries	19%	7,882

Note: Only direct school to school transitions count, delayed transitions counted as failures.

Explanatory variables

The explanatory variables include the father's and the mother's education (measured in years of school attendance) as measures of the family's cultural capital. In addition, the socio-economic status of family origin was measured using the 'International Socio-Economic Index of Occupational Status' (ISEI, see Ganzeboom, H. B. G. –De Graaf, P. M. – Treiman, D. J., 1992) of the household head at the time the respondent was 14 years old. If the father was employed and his occupation was known, the father's occupation was used, otherwise the mother's occupation was used as a substitute. Family size was measured by the number of siblings a respondent had; the measure was topcoded at 4 siblings[8] to minimize the influence of extreme values in the analysis[9]. A dichotomous variable was used to distinguish between men and women. The parents' membership in the Communist Party was taken as a measure of the political status of the family.

Table 3: **Descriptive statistics of independent variables in the analysis of secondary and tertiary school entry among primary and secondary school graduates in Central and Eastern Europe, 1948-1989.**

	Primary school graduates		Secondary school graduates	
	Mean	**Standard deviation**	**Mean**	**Standard deviation**
Male	0.49	- -	0.44	- -
Father's education	9.2	3.5	10.1	3.6
Father's education missing	0.05	- -	0.03	- -
Mother's education	8.3	3.2	9.0	3.3
Mother's education missing	0.02	- -	0.02	- -
No. of siblings	2.0	1.3	1.6	1.1
No. of siblings missing	0.02	- -	0.01	- -
Main earner's ISEI	35.0	14.4	39.0	16.2
Main earner's ISEI missing	0.06	- -	0.07	- -
Parents CP members	0.22	- -	0.28	- -
Parents CP membership- missing	0.08	- -	0.06	- -
'Red' cohort	0.13	- -	0.09	- -
Country				
Bulgaria	0.18	- -	0.31	- -
Czech Republic	0.24	- -	0.21	- -
Hungary	0.19	- -	0.15	- -
Poland	0.17	- -	0.12	- -
Slovakia	0.21	- -	0.21	- -
Number of cases at risk	17,935		7,882	

Note: See text for individual variables' value coding; education, occupation, and sibling scales before centring (see text for details). Standard deviations are not shown for dichotomous variables as they are simple a function of the mean.

Owing to the large number of responses missing for the question on the political status of the parents, three groups of respondents were distinguished: (1) those

[8] Only very small fractions of respondents in each country – below 9% among primary school graduates and below 6% among secondary school graduates – had 5 or more siblings.

[9] All interval variables were centered on their means before entering analysis. Descriptive statistics reported below, however, refer to scales before centring for the ease of interpretation. A dummy replacement of missing data was also used for all interval explanatory variables: missing data were replaced with the mean and a dichotomous identification variable was added to the right-hand side of each model. Overall mean computed over all respondents in all countries was used for replacements.

respondents who had at least one parent who was at some point a Communist party member, (2) those respondents whose both parents were never Communist party members, and finally (3) those respondents who did not declare the political status of their parents. Two dummy variables were used to contrast the first and the third group with the second. Finally, a dichotomous variable differentiated between respondents who came at risk of completing the particular transition during the year of extreme Communist orthodoxy and other respondents. I define the orthodox or 'red' periods as years 1949-1953 in Bulgaria, 1949-1953 and 1970-1973 in the Czech Republic, 1949-1953 in Hungary, 1949-1953 in Poland, and 1949-1954 in Slovakia[10]. The descriptive statistics (means and standard deviations) for all independent variables in the analysis of either the transition to secondary or tertiary educational institutions are presented in Table 3.

Modelling access to secondary education

The first step in the analysis is to estimate the following multinomial logistic regression model using a pooled data set for all countries:

$$\ln\left(\frac{P_V \mid X}{P_N \mid X}\right) = a_1 + b_1 * FE + c_1 * ME + d_1 * SEI + \sum_{i=1}^{j} e_{1i} * X_i + f_1 * C + + g_1 * C * FE + h_1 * C * ME + k_1 * C * SEI \qquad \text{[Eq. 1]}$$

$$\ln\left(\frac{P_P \mid X}{P_N \mid X}\right) = a_2 + b_2 * FE + c_2 * ME + d_2 * SEI + \sum_{i=1}^{j} e_{2i} * X_i + f_2 * C + + g_2 * C * FE + h_2 * C * ME + k_2 * C * SEI \qquad \text{[Eq. 2]}$$

$$\ln\left(\frac{P_A \mid X}{P_N \mid X}\right) = a_3 + b_3 * FE + c_3 * ME + d_3 * SEI + \sum_{i=1}^{j} e_{3i} * X_i + f_3 * C + + g_3 * C * FE + h_3 * C * ME + k_3 * C * SEI \qquad \text{[Eq. 3]}$$

where P_N is the probability of not attending any secondary school, P_V is the probability of enrolling in a vocational school, P_P is the probability of starting

[10] These years usually include the early years of the Communist rule when the fiercest Stalinism dominated politics and society alike. The end of this period is marked by Stalin's death. Yet post-Stalinism did not spread across the entire region immediately: it usually started in the political centres and spread to other regions with a delay. Hence, the orthodox period ends in 1953 in the Czech Republic, while it extends to 1954 in Slovakia despite the fact that these two nations were at that time members of one state. I also include years 1970-1973 among the orthodox cohorts in the Czech Republic, but I do not include these years in Slovakia. After the Prague Spring of 1968 a period of the so-called 'political normalization' of society followed, which was a period of political purges and increased ideological awareness. Yet these were mostly felt in Prague, where the centre of the liberal movement was, and touched other parts of the Czechoslovak federation and particularly Slovakia much less. Hence, I count these years as 'orthodox' only in the Czech Republic.

a professional school, and P_A is the probability of entering an academic school. By definition, $P_N + P_V + P_P + P_A = 1$. Then, *FE* is a measure of father's education, *ME* is a measure of mother's education, *SEI* is occupational status of the main earner in the family measured by the ISEI index and *C* is a dichotomous measure distinguishing 'red' (coded 1) and 'other' cohorts (coded 0). Finally, X_i is a set of *j* other explanatory variables used in the analysis (respondent's sex, country, parental membership in the Communist Party, number of siblings, and dummy identificators of mean-replaced missing values). Finally, the right-hand side of all three equations of the multinomial logit may contain interactions between cohort and measures of socioeconomic background (*FE*C, ME*C, SEI*C*). Because the parameters of the multinomial logistic regression model are not sensitive to changes in the marginal distribution of the dependent variable, modelling results presented below are by no means affected by the secular expansion of education each of the countries experienced during Communism (see e.g. Kreidl, M., 2002, 2004, 2005b).

The key statistical test of my hypotheses is going to be the statistical significance of the interaction between cohort (*C*) and father's education (*FE* in [Eq. 1], [Eq. 2], [Eq. 3]), cohort (*C*) and mother's education (*ME*), and cohort (*C*) and main earner's socioeconomic status (*SEI*). Because there is some ambiguity in the literature regarding whether the 'Communist Affirmative Action' applied to all types of secondary schools equally, or whether it was mostly limited to the most prestigious academic secondary schools (see e.g. claims made by Hanley, E., – McKeever, M., 1997) I will test for the presence of interactions between measures of SES and cohort in all three equations and then as the next step I will include it in one of the three equations only. I expect to see that the effect of socioeconomic background is weaker in the 'orthodox red' cohorts than in the other cohorts, i.e. I expect the interactions to be negative. If none of those interactions turns out to be statistically significant, I will reject the hypothesis that orthodox socialism reduced inequality in access to secondary education.

Modelling entry into university

I will model success in the transition from secondary to tertiary education using binomial logistic regression. The model will have the following form:

$$\ln\left(\frac{P_U|X}{1-P_U|X}\right) = a + b*FE + c*ME + d*SEI + \sum_{i=1}^{j} e_i * X_i + f*C + \quad \text{[Eq. 4]}$$
$$+ g*C*FE + h*C*ME + k*C*SEI$$

where P_U is the probability of enrolling in a university in the year of high school graduation, *FE* is father's education, *ME* is mother's education, *SEI* is

occupational status of the main earner in the family. X_i is a set of *j* other explanatory variables. The last explanatory variable will again be a dichotomous measure of cohort (*C*). Finally, the right-hand side of the equation contains interactions between cohort and measures of socioeconomic background (*FE*C, ME*C, SEI*C*). These, again, are the key test of the proposition that 'Communist Affirmative Action' had an effect on socioeconomic inequality in access to schooling and I expect them to be negative[11].

Results

Entry into secondary schools

When one adds interactions between all three measures of parental socioeconomic status (father's education, mother's education, head of household's ISEI) and cohort into all three equations of the multinomial logit, one obtains a significant improvement of the model fit (see model comparisons in Table 4). The maximum likelihood chi-square test returns L^2= 25.3 with 9 degrees of freedom, which is obviously statistically significant (p-value=0.003). A more restricted test would test for the presence of the interactions only in the equation predicting enrolment in an academic secondary school. Even this approach improves the model fit significantly (L^2=19.3 with 3 d.f., p-value=0.000). These results point to the fact that the two-way interactions in question are statistically significant and interpretable.

In order to gain a deeper insight into which of the interactions are significant and also to check for possible multicollinearity between measures of socioeconomic background, I also added to the model one two-way interaction at a time and tested whether this partial model extension would improve its fit significantly. Results of these tests are reported in Table 4 as well. They reveal that each of the interactions or sets of interactions enhances the model fit significantly. For instance, when the interaction between father's education and cohort is added to all three equations of the model, it leads to a significantly improved fit (L^2=15.7 with 3 d.f. – see Table 4). When this interaction is added to the last equation only, it also results in an improved fit (L^2=13.2 with 1 d.f.). Similarly, when mother's education is interacted with cohort in all three equations of the multinomial logit, it improves the model fit significantly (L^2=20.5 with 3 d.f.), as it does when added to the last equation only (L^2=16.4 with 1 d.f.). Finally, the interaction of ISEI with cohort also produces significant improvements in the model fit (L^2=12.8 with 3 d.f. when added to all 3 equations; L^2=9.0 with 1 d.f.

[11] The parameters of the binomial logistic regression model are – similarly to the multinomial logit mentioned earlier – not sensitive to changes in the marginal distribution of the dependent variable.

when added to the last equation only – see also Table 4). Overall, it seems that the data offer strong and robust evidence that all utilized measures of family socioeconomic background interact statistically significantly with cohort, which means that their effects on the odds of progressing to secondary educational institutions differ significantly across cohorts.

Table 4: **Parameters of the maximum likelihood chi-square test for the comparison of multinomial logistic regression models of secondary school entry with and without interactions between measures of socioeconomic background and cohort, Central and Eastern Europe, 1948- 1989. Number of cases at risk = 17, 935.**

	Statistics of the test		
Model # and description of tested interactions	**L^2**	**d.f.**	**p-value**
M1: Father's education * 'Red', Mother's education * 'Red', ISEI * 'Red' (all equations)	25.3	9	0.003
M2: Father's education * 'Red', Mother's education * 'Red', ISEI * 'Red' ('academic' equation only)	19.3	3	0.000
M3: Father's education * 'Red' (all equations)	15.7	3	0.001
M4: Father's education * 'Red' ('academic' equation only)	13.2	1	0.000
M5: Mother's education * 'Red' (all equations)	20.5	3	0.000
M6: Mother's education * 'Red' ('academic' equation only)	16.4	1	0.000
M7: ISEI * 'Red' (all equations)	12.8	3	0.005
M8: ISEI * 'Red' ('academic' equation only)	9.0	1	0.003

Notes:
1. The multinomial logit has a dependent variable with 4 possible outcomes, i.e. it has 3 equations. I test for the presence of interactions either in all three equations, or in one of them only- then it is the equation predicting entry into academic secondary school as contrasted to no secondary education.
Apart from the tested interactions, each model also contains the main effects of the following variables: respondent's sex, parent's membership in the Communist Party, number of siblings, country, cohort, and dummy identificators of mean-replaced missing values.

Estimated coefficients and standard errors of selected multinomial logistic regression models are shown in Table 5. For instance, Model 1 in Table 5 contains a full set of two-way interactions between measures of SES and cohort in all three equations of the multinomial logit. Its coefficients confirm general expectations. Men have an advantage over women when entering the vocational and professional secondary schools, while women have an advantage at academic secondary schools. All three measures of family socioeconomic standing - father's education, mother's education, and ISEI – have a positive net effect on the odds of entering either vocational, professional, or academic schools. Moreover, it seems that these effects are strongest on the transition to academic secondary schools, weaker on the transition to professional schools, and weakest on the transition to

vocational schools (see Table 5). Similarly, the number of siblings has a negative effect on all three transitions. Again, this effect seems to be stronger on the transition to academic secondary schools than on the other two transitions. Finally, there is a clear advantage for students whose parents were members of the Communist Party.

Table 5: **Estimated coefficients and standard errors (in parentheses) of selected multinomial logistic regression models of secondary school entry, Central and Eastern Europe, 1948- 1989. Number of cases at risk = 17,935.**

	Model 1			**Model 2**		
Main Effects	**Vocational**	**Professional**	**Academic**	**Vocational**	**Professional**	**Academic**
Male	0.993 (0.048)	0.357 (0.053)	-0.266 (0.058)	0.993 (0.048)	0.357 (0.053)	-0.266 (0.058)
Father's education	0.106 (0.012)	0.152 (0.013)	0.186 (0.014)	0.106 (0.011)	0.152 (0.012)	0.186 (0.013)
Mother's education	0.095 (0.012)	0.194 (0.013)	0.223 (0.014)	0.095 (0.011)	0.190 (0.012)	0.221 (0.013)
Occupational status	0.007 (0.003)	0.030 (0.003)	0.048 (0.003)	0.008 (0.003)	0.030 (0.003)	0.048 (0.003)
# of siblings	-0.170 (0.019)	-0.378 (0.022)	-0.383 (0.024)	-0.170 (0.019)	-0.379 (0.022)	-0.383 (0.024)
'Red' cohorts	-0.548 (0.079)	-0.677 (0.089)	-0.239 (0.093)	-0.549 (0.065)	-0.664 (0.076)	-0.226 (0.082)
Parents CP	0.215 (0.070)	0.399 (0.072)	0.386 (0.076)	0.215 (0.070)	0.400 (0.072)	0.386 (0.076)
Interactions:						
FEd * 'Red'	-0.002 (0.028)	-0.002 (0.031)	-0.023 (0.032)			-0.021 (0.027)
MEd * 'Red'	0.003 (0.029)	-0.031 (0.032)	-0.069 (0.032)			-0.058 (0.027)
ISEI * 'Red'	0.005 (0.007)	-0.002 (0.008)	-0.005 (0.008)			-0.005 (0.005)
Constant	1.030 (0.062)	0.884 (0.066)	0.282 (0.074)	1.028 (0.062)	0.886 (0.066)	0.273 (0.075)

Note: Each model also contains the following explanatory variables: identificators of mean-replaced missing values and four country dummies (Bulgaria, Czech Republic, Hungary, Poland; Slovakia is the comparison category). Their coefficients are not shown to save space, yet are available from the author upon request.

Model 1 displayed in Table 5 also has a set of two-way interactions of family background and cohort. We have seen above that these interactions do collectively improve the model fit significantly. In Table 5, we can study in more detail the exact nature of these interactions. We see that the interaction between father's education and cohort and ISEI and cohort isn't statistically significant in none of the three equations, while the interaction between mother's education and cohort is significant and negative in the equation predicting entry into academic secondary schools. We can observe that each additional year of mother's schooling increases – net of other factors – the log odds of entering academic secondary school as compared to no schooling by 0.223 in 'normal' cohorts, while its effect is only 0.154 (=0.223-0.069) in 'red' cohorts. Interaction effects in Model 2, in which interactions are included only in the equation predicting entry into academic secondary schools, lead to the same substantive conclusions.

Because the measures of father's education, mother's education, and occupational status are correlated, interaction terms involving these variables are correlated as well, and interaction effects reported in Model 1 may be affected by multicollinearity. Therefore, I also build a series of models that only contain one of these interactions at a time, which lessens the likelihood of a serious bias due to multicollinearity. Estimated coefficients and standard errors of these simpler models are displayed in Table 6. They generally indicate that all three measures of parental socioeconomic status do indeed interact significantly with cohort once these interactions are tested separately. For instance, we can see in Model 4 that one additional year of father's education increases – net of other factors – the log odds of entry into academic secondary education by 0.193 in normal cohorts and only by 0.125 (=0.193-0.068) in 'red' cohorts. Similarly, the effects of mother's education in Model 6 is 0.224 in normal cohorts and 0.140 (=0.224-0.084) in 'red' cohorts, and the effect of ISEI in Model 8 is 0.049 in normal cohorts and 0.036 (=0.049-0.013) in 'red' cohorts (see Table 6).

Altogether, we have seen that the effects of socioeconomic background on the odds of students' successful progression from primary to secondary school indeed varied by historical period. The advantages associated with higher family status declined significantly in periods of the most orthodox Communism. Yet, these effects were limited to one type of secondary education – namely to academic secondary schools – and were not present in other secondary school types. This finding confirm the claim made in the literature (see e.g. Hanley, E. – McKeever, M., 1997; Matějů, P., 1993) that the 'Communist Affirmative Action' indeed focused most prominently on this most prestigious type of secondary education.

Entry into university

I modelled entry into university using simple binomial logistic regression models. In these models I again tested for the statistical significance of two-way interactions between father's education and cohort, mother's education and cohort, and the household head's occupational status (ISEI) and cohort. Statistics for the maximum-likelihood chi-square test for the significance of these interactions are reported in Table 7. In Model 9, I test the statistical significance of these interactions collectively. The test yields L^2 of 8.3 with 3 degrees of freedom, which is only marginally statistically significant (p-value= 0.04). Because of possible multicollinearity between interactions, it is worth testing their significance also individually and I accomplish this in Models 10, 11, and 12 (see Table 7). It turns out from Model 10 that while the interaction between father's education and cohort isn't statistically significant at all (p-value for the test = 0.249), and the interaction between ISEI and cohort (Model 12) is only marginally statistically significant (L^2=3.7 with 1 degree of freedom, p-value= 0.055), the interaction between mother's education and cohort in Model 11 indeed is fairly highly statistically significant (L^2=6.4 with 1 degree of freedom, p-value= 0.012).

Table 7: **Parameters of the maximum likelihood chi-square test for the comparison of binomial logistic regression models of university entry with and without interactions between measures of socioeconomic background and cohort, Central and Eastern Europe, 1948- 1989. Number of cases at risk = 7,882.**

	Statistics of the test		
Model # and description of tested interactions	**L^2**	**d.f.**	**p-value**
M9: Father's education * 'Red' + Mother's education * 'Red' + ISEI * 'Red'	8.3	3	0.04
M10: Father's education * 'Red'	1.3	1	0.249
M11: Mother's education * 'Red'	6.4	1	0.012
M12: ISEI * 'Red'	3.7	1	0.055

Note: Apart from the tested interactions, each model also contains the main effects of the following variables: respondent's sex, parent's membership in the Communist Party, number of siblings, country, cohort, and dummy identificators of mean-replaced missing values.

Estimated coefficients and standard errors of Models 9, 10, 11, and 12 are shown in Table 8. They confirm our expectations: men are advantaged in comparison with women[12], and higher father's education, mother's education, and ISEI all independently increase the chances of enrolling in a university, while larger sibship sizes tend to reduce the odds of university enrolment (see Table 8). Interaction effects reported in the lower panel of Table 8 document that indeed the effects of parental statuses were indeed weaker in periods of elevated Communist orthodoxy, while they were relatively stronger during more liberal periods. For instance, we learn from Model 9 that whereas each additional year of mother's education increased on average the log odds of university entry by 0.079 in normal periods, its effect was slightly negative (-0.011=0.079-0.09) in 'red' cohorts. Similarly, Model 12 documents that the effect of parental occupational status differed across cohorts. One additional point on the ISEI scale increased – net of other factors – the log odds of university entry by 0.017 in normal cohorts, while its effect was 0.005 (=0.017-0.012) in orthodox periods. Father's education doesn't interact with cohort at all.

[12] We know from other research that male advantage in education was reduced and eventually disappeared during Communism. This paper doesn't focus on gender inequality and thus simplifies the model by not including interactions between gender and cohort.

Table 8: **Estimated coefficients and standard errors (in parentheses) of selected binomial logistic regression models of entry into university, Central and Eastern Europe, 1948- 1989. Number of cases at risk = 7,882.**

	Model 9	**Model 10**	**Model 11**	**Model 12**
Main Effects				
Male	0.368 (0.061)	0.371 (0.061)	0.369 (0.061)	0.370 (0.061)
Father's education	0.080 (0.014)	0.088 (0.013)	0.084 (0.013)	0.085 (0.013)
Mother's education	0.079 (0.013)	0.071 (0.013)	0.079 (0.013)	0.070 (0.013)
Occupational status	0.017 (0.002)	0.016 (0.002)	0.016 (0.002)	0.017 (0.002)
# of siblings	-0.078 (0.030)	-0.074 (0.030)	-0.076 (0.030)	-0.076 (0.030)
'Red' cohorts	0.190 (0.112)	0.222 (0.108)	0.183 (0.107)	0.242 (0.108)
Parents CP	0.050 (0.067)	0.053 (0.067)	0.051 (0.067)	0.053 (0.067)
Interactions:				
FEd * 'Red'	0.044 (0.042)	-0.033 (0.028)		
Med * 'Red'	-0.090 (0.042)		-0.084 (0.032)	
ISEI * 'Red'	-0.010 (0.008)			-0.012 (0.006)
Constant	-1.459 (0.073)	-1.457 (0.073)	-1.460 (0.073)	-1.458 (0.073)

Note: Each model also contains the following explanatory variables: identificators of mean-replaced missing values and four country dummies (Bulgaria, Czech Republic, Hungary, Poland; Slovakia is the comparison category). Their coefficients are not shown to save space, yet are available from the author upon request.

Summary of Results

In sum, we have seen that the effects of family socioeconomic background on the odds of progressing from primary to secondary education varied across cohorts being smaller in 'red' cohorts, and stronger in cohorts that graduated from primary schools during more liberal times. We have observed a similar pattern for the entry into university as well. However, results regarding entry into university were less conclusive and robust than results for secondary education. There is a variety of reasons why this could be so.

First, obviously the *sample size* for the latter part of the analysis was significantly smaller (n=7,882) than for the former (n=17,935) and so the difference in achieved significance levels might simply reflect the sample size rather than the difference in the substantive processes governing the transition from primary to secondary and from secondary to tertiary education. Second, there obviously might be *substantive reasons* why admission processes worked differently at the secondary and tertiary levels. Yet, a careful review of the existing literature doesn't reveal any strong theoretical explanations, and thus I leave this topic open for future investigations. Finally, it is possible that *admission procedures also differed within segments of tertiary education* and failing to distinguish them analytically might have biased the results. For instance, it is possible that political criteria were more salient in schools of social sciences, philosophy, humanities, law, and education, while they were less decisive in schools of sciences, engineering, medical schools, and other politically less sensitive fields. In order to test this assertion one would need to go beyond a simple dichotomous dependent variable measuring entry into university, employ a nominal dependent variable and a multinomial logistic regression. This should be possible with the data file used in this paper, as the survey also recorded the field of study at the university level. It is however going to be a little troublesome, because categories measuring field of study were nation-specific and using them in a cross-national comparative research may create more trouble than benefit. Nonetheless, it is an interesting issue to pursue in future research.

Concluding Discussion

This paper shows that we can identify some significant declines in the socio-economic inequality in access to schooling in former socialist countries in Central and Eastern Europe during Communism. Hence, it questions conclusions and claims made earlier by a number of scholars who believed that the so-called 'Communist Affirmative Action' was inefficient and who thus expressed doubts regarding to possibility that political actors can alter the patterns of educational stratification. This paper maintains that previous research has utilized such specifications of both the key dependent and independent variables that inevitably led it to miss some important aspects of the process of educational stratification in former socialist nations.

I have found three major shortcomings in previous research. (1) It did not distinguish direct and delayed school transitions and thus overlooked the fact that delayed transitions could offset the initial success of an egalitarian intervention

(cf. Kreidl, M., 2005a). (2) It failed to acknowledge the fact that secondary education wasn't internally homogenous and that according to historical sources only some types of secondary schools were supposed to be subject to political intervention. (3) Previous research utilized historically ambiguous definitions of cohorts; definitions that did not separate cohorts that went through the crucial school transitions during years of extreme Communist orthodoxy and cohorts that proceeded through the system in years of a relative political and ideological liberalization.

Once we correct the definitions of both the dependent and independent variables, we see clear, statistically significant, and fairly robust, yet very brief period effects. We see that Bulgaria, the Czech Republic, Hungary, Poland, and Slovakia experienced ephemeral periods of destratification during the early 1950s, and the Czech Republic also during the early 1970s. How does one adjudicate this finding and the consistent conjecture of all earlier research that there were no periods of reduced socioeconomic inequality in access to schooling in former socialist nations? I think the best answer is to look for a theoretical explanation that would link differences in the definitions of variables to substantively interesting social processes.

Delayed educational careers might have worked to offset the effect of an egalitarian intervention in 'red' cohorts. In other words, I propose a theory of the historically variable role of delayed school progressions that maintains that socio-economically advantaged and politically disadvantaged families utilized delayed transitions more often and more successfully in 'red' cohorts than in liberal cohorts. Thus, in 'red' cohorts we can see reduced inequality in direct school transitions and increased inequality in delayed school transitions. Consequently, these two processes combine to produce an overall stability in the effects of socioeconomic background on the odds of success in school transitions. Because I studied only direct school progression in this paper, and other scholars focused on ultimate attainment, the two conclusions may in fact be describing the same social reality and do not necessarily contradict each other. Rather, they seem to be compatible. Yet, without looking at direct school progressions more specifically, as I did in this paper, we would have missed a subtle point and would have been interpreting earlier research incorrectly. Clearly, we need to carry out an explicit empirical test of this substantive explanation.

Seeing that the Communist regimes indeed achieved at least partial success in reducing socioeconomic inequality in access to schooling may inspire new research questions that would look in more detail into what specific aspect of the egalitarian intervention produced the most salient effect? The opening discussion emphasized that the policy of equal opportunities targeted both the advantaged and the disadvantaged groups. Recently publicized research suggests that the negative discrimination of the formerly advantaged families was more efficient than the policy of positive discrimination of the working class youth in equalizing access to schooling (Simonová, N., 2006). Since this research suffers from the same flaws as other investigations reviewed in the opening sections of this paper, I maintain that it be scrutinized carefully before accepting its conclusions. Nonetheless, it seems to be a plausible hypothesis worth of further empirical tests.

Martin Kreidl *obtained his Ph.D. in sociology from the University of California, Los Angeles (UCLA). He is the Chair of the Department of Sociology, University of West Bohemia in Plzeň, the Czech Republic. He conducts research in social stratification, social demography and sociology of health. His recent publications include „Trends in Educational Assortative Mating in Central Europe: the Czech Republic, Slovakia, Poland, and Hungary between 1988 and 2000" (in European Sociological Review Vol. 22, with T. Katrňák and L. Fónadová) and "Politics and Secondary School Tracking in Socialist Czechoslovakia, 1948-1989." (in European Sociological Review Vol. 20).*

REFERENCES

BREEN, R. – JONSSON, J. O., 2000: Analyzing Educational Careers: A Multinomial Transition Model. American Sociological Review 65, pp. 754-772.

CONNELLY, J., 1997: Communist Higher Education Policies in Czechoslovakia, Poland, and East Germany. Pp. 191-216. In: The Establishment of Communist Regimes in Eastern Europe, 1944-1949. Edited by Norman Naimark and Leonid Gibianskii. Boulder: Westview Press.

CONNELLY, J., 2000: Captive University. The Sovietization of East German, Czech, and Polish Higher Education, 1945-1956. Chapel Hill: The University of North Carolina Press.

DENG, Z. – TREIMAN, D. J., 1997: The Impact of Cultural Revolution on Trends in Educational Attainment in the People's Republic of China. American Journal of Sociology 103, pp. 391-428.

FISZMAN, J. R., 1972: Revolution and tradition in people's Poland: education and socialization. Princeton: Princeton University Press.

GANZEBOOM, H. B. G. – DE GRAAF, P. M. – TREIMAN, D. J., 1992: A Standard International Socio-Economic Index of Occupational Status. Social Science Research 21, pp. 1-56.

HANLEY, E., 2001: Centrally Administered Mobility Reconsidered: The Political Dimension of Educational Stratification in State-Socialist Czechoslovakia. Sociology of Education 74, pp. 25-43.

HANLEY, E. – MCKEEVER, M., 1997: The Persistence of Educational Inequalities in Socialist Hungary: Trajectory Maintenance vs. Counterselection. Sociology of Education 70, pp. 1-18.

HEYNS, B. – BIALECKI, I., 1993: Educational Inequality in Post-War Poland. Pp. 303-335. In: Persistent Inequality. Changing Educational Attainment in Thirteen Countries. Edited by Yossi Shavit and Hans – Peter Blossfeld. Boulder, Westview Press.

KONRAD, G. – SZELENYI, I., 1979: Intellectuals on the Road to Class Power. New York, Harcourt Brace Jovanovich.

KREIDL, M., 2002: Trendy ve výběru střední školy v socialistickém Československu. Sociologický časopis 37, pp. 565-592.

KREIDL, M., 2004: Politics and Secondary School Tracking in Socialist Czechoslovakia, 1948-1989. European Sociological Review 20, pp. 123-139.

KREIDL, M., 2005a: Education Delayed or Education Denied? Evidence on the Historically Variable Role of Delayed Educational Careers in Former Communist Countries. California Center for Population Research On-line Working Paper Series 47/05. Los Angeles, California Center for Population Research. (available online at http://www.ccpr.ucla.edu/ccprwpseries/ccpr_047_05.pdf).

KREIDL, M., 2005b: Inequality of Access to Secondary and Post-Secondary Education in Countries of the Former Soviet Bloc after 1948. Ph.D. dissertation. Los Angeles, UCLA, Department of Sociology.

KREIDL, M., 2005c: Educating Communist Cadres: School Re-entry and Sponsored Educational Mobility in State Socialism. California Center for Population Research On-line Working Paper Series 55/05. Los Angeles, California Center for Population Research. (available at http://www.ccpr.ucla.edu/ccprwpseries/ccpr_055_05.pdf).

MARE, R. D., 1981: Change and Stability in Educational Stratification. American Sociological Review 46, pp. 295-305.

MATĚJŮ, P., 1993: Who Won and Who Lost in a Socialist Redistribution in Czechoslovakia? Pp. 251-271. In: Persistent Inequality. Changing Educational Attainment in Thirteen Countries. Edited by Yossi Shavit and Hans – Peter Blossfeld. Boulder, Westview Press.

MATĚJŮ, P. – ŘEHÁKOVÁ, B. – SIMONOVÁ, N., 2004: Kulturní a sociálně ekonomické zdroje nerovností v šancích na dosažení vysokoškolského vzdělání v České republice v letech 1948-1999. Sociológia 36, pp. 31-56.

NIEUWBEERTA, P. – RIJKEN, S., 1996: Educational Expansion and Educational Reproduction in Eastern Europe, 1940-1979. Czech Sociological Review 4, pp. 187-210.

PARKIN, F., 1971: Class Inequality and Political Order: Social Stratification in Capitalist and Communist Societies. New York, Praeger.

RÓBERT, P., 1991: Educational Transition in Hungary from the Post-war Period to the End of the 1980's. European Sociological Review 7, pp. 213-236.

SHAVIT, Y. – BLOSSFELD, H. P., (Eds.) 1993: Persistent Inequality. Changing Educational Attainment in Thirteen Countries. Boulder: Westview Press.

SIMKUS, A. A., 1981: Historical Changes in Occupational Inheritance Under Socialism: Hungary 1930-1973. Research in Social Stratification and Mobility 1, pp. 171-203.

SIMKUS, A. – ANDORKA, R., 1982: Inequalities in Educational Attainment in Hungary, 1923-1973. American Sociological Review 47, pp. 740-751.

SIMONOVÁ, N., 2003: The Evolution of Educational Inequalities in the Czech Republic after 1989. British Journal of Sociology of Education 24, pp. 469-483.

SIMONOVÁ, N., 2006: Vzdělanostní nerovnosti a vzdělanostní mobilita v období socialismu. Pp. 62-91. In: Nerovné šance na vzdělání. Vzdělanostní nerovnosti v České republice. Edited by Petr Matějů and Jana Straková. Praha, Academia.

SZELÉNYI, S. – ASCHAFFENBURG, K., 1993: Inequalities in Educational Opportunity in Hungary. Pp. 273-302. In: Persistent Inequality. Changing Educational Attainment in Thirteen Countries. Edited by Yossi Shavit and Hans – Peter Blossfeld. Boulder, Westview Press.

TREIMAN, D. J. – SZELÉNYI, I., 1994: Social Stratification in Eastern Europe after 1989. General Population Survey. Provisional Codebook. Treiman, D. J. and I. Szelényi [principal investigators]; Los Angeles, UCLA, Social Science Data Archive [distributor].

ULČ, O., 1978: Some aspects of Czechoslovak Society since 1968. Social Forces 57, pp. 419-435.

WHEELER, G. S., 1973: The human face of socialism: the political economy of change in Czechoslovakia. New York, L. Hill.

WONG, R. S-K., 1998: Multidimensional Influences of Family Environment in Education: The Case of Socialist Czechoslovakia. Sociology of Education 71, pp. 1-22.

Table 6: **Estimated coefficients and standard errors (in parentheses) of selected multinomial logistic regression models of secondary school entry, Central and Eastern Europe, 1948- 1989. Number of cases at risk = 17,935.**

	Model 4			**Model 6**			**Model 8**		
Main Effects	**Vocational**	**Professional**	**Academic**	**Vocational**	**Professional**	**Academic**	**Vocational**	**Professional**	**Academic**
Male	0.993 (0.048)	0.357 (0.053)	-0.265 (0.058)	0.993 (0.048)	0.357 (0.053)	-0.267 (0.058)	0.993 (0.048)	0.357 (0.053)	-0.264 (0.058)
Father's education	0.107 (0.011)	0.154 (0.012)	0.193 (0.013)	0.106 (0.011)	0.151 (0.012)	0.183 (0.012)	0.106 (0.011)	0.152 (0.012)	0.183 (0.012)
Mother's education	0.094 (0.011)	0.188 (0.012)	0.213 (0.013)	0.095 (0.011)	0.190 (0.012)	0.224 (0.013)	0.094 (0.011)	0.188 (0.012)	0.213 (0.013)
Occupational status	0.007 (0.003)	0.030 (0.003)	0.048 (0.003)	0.007 (0.003)	0.030 (0.003)	0.048 (0.003)	0.008 (0.003)	0.030 (0.003)	0.049 (0.003)
# of siblings	-0.170 (0.019)	-0.379 (0.022)	-0.383 (0.024)	-0.170 (0.019)	-0.379 (0.022)	-0.383 (0.024)	-0.171 (0.019)	-0.379 (0.022)	-0.384 (0.024)
'Red' cohorts	-0.542 (0.064)	-0.651 (0.076)	-0.200 (0.080)	-0.544 (0.064)	-0.653 (0.076)	-0.242 (0.082)	-0.528 (0.064)	-0.642 (0.076)	-0.175 (0.080)
Parents CP	0.215 (0.070)	0.400 (0.072)	0.385 (0.076)	0.215 (0.070)	0.399 (0.072)	0.387 (0.076)	0.215 (0.070)	0.400 (0.072)	0.387 (0.076)
Interactions:									
FEd * 'Red'			-0.068 (0.019)						
MEd * 'Red'						-0.084 (0.020)			
ISEI * 'Red'									-0.013 (0.004)
Constant	1.027 (0.062)	0.885 (0.066)	0.275 (0.074)	1.027 (0.062)	0.885 (0.066)	0.277 (0.074)	1.025 (0.062)	0.883 (0.066)	0.275 (0.074)

Note: Each model also contains the following explanatory variables: identificators of mean-replaced missing values and four country dummies (Bulgaria, Czech Republic, Hungary, Poland; Slovakia is the comparison category). Their coefficients are not shown to save space, yet are available from the author upon request.