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**SOCIALIST EGALITARIAN POLICIES AND INEQUALITY IN ACCESS TO SECONDARY
AND POST-SECONDARY EDUCATION: NEW EVIDENCE USING INFORMATION ON
DETAILED EDUCATIONAL CAREERS AND THEIR TIMING**

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ABSTRACT

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 definition of both dependent and independent variables. After correcting these inaccuracies, I analytically explore data from the Social stratification in Eastern Europe after 1989 data file and am indeed able to document that inequality in access to education indeed declined during periods of most extreme Communism in the early 1950s and, in some countries, also during the early 1970s.

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1 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 2001; Wheeler 1973; Connelly 1997, 2000; Kreidl 2004; Hanley, McKeever 1997; Simkus 1981; Simkus, Andorka 1982; Róbert 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 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 2001; Hanley, McKeever 1997; Heyns, Bialecki 1993; Matějů 1993; Nieuwbeerta, Rijken 1996; Róbert 1991; Simkus, Andorka 1982; Szelényi, Aschaffenburg 1993).¹ The literature on this topic offers a number of substantive explanations why this is so. For instance Szelényi, 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 (see also Fiszman 1976). 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 today observe no decline in the estimated effects of measures of socioeconomic background on the odds of success in educational transitions (see e.g. Hanley,

¹ 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.

McKeever 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, Szelenyi 1979; Matějů 1993; Parkin 1971). As a result we again do not observe any 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 conclusion 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 has 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 issues and reexamining the results, before we accept the conclusion that “Communist Affirmative Action” in education really failed to achieve greater equality in former socialist nation of Central and Eastern Europe.

2 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*². 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 almost an equal representation of general education and vocational preparation in areas such as

² There are often very good reasons to simplify and ignore some country-specific institutions. Those reasons can be data-related or design-related. For instance, Shavit and Blossfeld (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 such as for the purpose of this particular analysis.

electrical and civil engineering, agronomy, or accounting. Professional schools were intended to prepare students for immediate labor 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.

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 modeled “entry” into university (defined as university degree vs. no degree). Heyns and Bialecki (1993), Nieuwbeerta and Rijken (1996), and Szelényi and Aschaffenburg (1993) modeled 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 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 modeling 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 model 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, Jonsson 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 2001; Hanley, McKeever 1997; Matějů 1993; Simkus, Andorka 1982), they paid little attention to them in their actual statistical modeling 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 is not warranted.

There has been *insufficient attention paid 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 2001; Hanley, McKeever 1997; Heyns, Bialecki 1993; Matějů 1993; Nieuwbeerta, Rijken 1996; Róbert 1991; Simkus, Andorka 1982; Szelényi, Aschaffenburg 1993) – to determine any historical variations in the educational attainment process. Unfortunately, those cohorts were defined without any explicit reference to institutional and historical development and were therefore not well suited to identify any policy effects, while there is already some evidence suggesting that there was a lot of historical variation in the processes and consequences of Communist egalitarian policies (Kreidl 2004).

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, McKeever 1997; Heyns, Bialecki 1993; Matějů 1993; Nieuwbeerta, Rijken 1996; Róbert 1991; Simkus, Andorka 1982; Szelényi, Aschaffenburg 1993). 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. 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 were 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.

3 DATA, VARIABLES, AND THE MODELING 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, Czech Republic, Hungary, Poland, Russia, and Slovakia (see Treiman and Szelényi 1994 for details). However, due to a small yet important deviation in the Russian questionnaire the comparability of the Russian educational data with that of other countries is questionable, and I decided not to use in this paper.

The educational roster of this survey contains all the information required to define the dependent variables for this analysis. 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 respondent completed primary education and made a decision about secondary school attendance during the Communist period, and who can therefore be considered in the analysis of secondary school entry. Similarly, 7,882 respondent obtained complete secondary education during Communism and were at risk of progressing to the tertiary level³. 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.

³ 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 2003, 2005c). Moreover, non-standard education careers during Communism were studied elsewhere (Kreidl 2005c).

3.1 The dependent variables

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 never attended 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 enrollment 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

enrollment 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% 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 enrollments at both secondary and tertiary levels here, as they were reported in Kreidl (2005e) and the interested reader is referred there for more details.

3.2 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, De Graaf, and Treiman 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 siblings⁴ to minimize the influence of extreme values in the analysis.⁵ 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. Owing to the large number of responses missing for the question on the political status of the parents, three

⁴ 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.

⁵ All interval variables were centered on their means before entering analysis. Descriptive statistics reported below, however, refer to scales before centering for the ease of interpretation. A dummy replacement of missing data was also used for all interval explanatory variables: missing data was replaced with the mean and a dichotomous identification variable was added to the right-hand side of each model.

groups of respondents were distinguished: (1) those 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 Slovakia⁶.

The descriptive statistics (means and standard deviations) for all the independent

⁶ These years usually include the early years of the Communist rule in each country, when the fiercest Stalinism dominated politics and society alike. The end of this period is marked by Stalin’s death, yet the post-Stalinism did not spread across the entire region immediately: it usually started in the political centers 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 a 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 center 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.

variables in the analysis of either the transition to secondary or tertiary educational institutions are presented in Table 3.

3.3 Modeling 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:

$$\ln\left(\frac{P_V|X}{P_N|X}\right) = a_1 + b_1 * FE + c_1 * ME + d_1 * SEI + \sum_{i=1}^j e_{i1} * X_{i1} + f_1 * C_1 + g_1 C_1 * FE + h_1 C_1 * ME + k_1 C_1 * SEI \quad [\text{Eq. 1}]$$

$$\ln\left(\frac{P_P|X}{P_N|X}\right) = a_2 + b_2 * FE + c_2 * ME + d_2 * SEI + \sum_{i=1}^j e_{i2} * X_{i2} + f_2 * C_2 + g_2 C_2 * FE + h_2 C_2 * ME + k_2 C_2 * SEI \quad [\text{Eq. 2}]$$

$$\ln\left(\frac{P_A|X}{P_N|X}\right) = a_3 + b_3 * FE + c_3 * ME + d_3 * SEI + \sum_{i=1}^j e_{i3} * X_{i3} + f_3 * C_3 + g_3 C_3 * FE + h_3 C_3 * ME + k_3 C_3 * SEI \quad [\text{Eq. 3}],$$

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 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 identifiers 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 ($FE*C$, $ME*C$, $SEI*C$).

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 ($ISEI$). 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 and McKeever 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.

3.4 Modeling 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:

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

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.

4 RESULTS

4.1 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 enrollment 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, 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,

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. 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.

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.

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 neither 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. In periods of the most orthodox Communism, the advantages associated with higher family status declined significantly. 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, McKeever 1997; Matějů 1993) that the “Communist Affirmative Action” indeed focused most prominently on these most prestigious types of secondary education.

4.2 Entry into university

I modeled entry into university using simple binomial logistic regression models. In the context of 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 these interactions collectively. It 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).

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, 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 enrollment (see Table 8). Interaction effects reported in the lower panel of Table 8 document that indeed the effects of parental statuses were 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) in "Red" cohorts. Similarly, Model 12 also 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.

5 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 orthodox 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 be an artifact of the sample size rather than a result of 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 on the topic 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 bias 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 might have been less decisive in schools of sciences, engineering, medical schools, and other presumably 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 move to a multinomial logistic regression instead. 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 study may create more trouble than benefit. Nonetheless, it is an interesting issue to pursue in future research.

6 CONCLUDING DISCUSSION

This paper shows that we can indeed identify some significant declines in the socioeconomic 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 documents that previous research on this topic has utilized such specification 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. (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 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 brief periods of destratification during the Stalinist period in the early 1950s, and the Czech Republic also during the early 1970s during the so-called post-1968 “political normalization”. How does one adjudicate this finding and the consistent finding 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.

For instance, I suggested earlier that delayed educational careers might have worked to offset the effect of an egalitarian intervention in “red” cohorts. In other words, I proposed 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 school transitions.

Because I studied only direct school progression in this paper, and other scholars before 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 outlined above.

7 REFERENCES

Blau, Peter and Otis Dudley Duncan. 1967. *The American Occupational Structure*.
New York: Wiley.

Blossfeld, Hans-Peter and Yossi Shavit. 1993. "Persisting Barriers: Changes in
Educational Opportunities in Thirteen Countries." Pp. 1- 23 in: *Persistent
Inequality. Changing Educational Attainment in Thirteen Countries*. Edited
by Yossi Shavit and Hans-Peter Blossfeld. Boulder: Westview Press.

Boguszak, Marek, Ivan Gabal and Petr Matějů. 1990. „Ke koncepcím vývoje sociální
struktury ČSSR. Příspěvek k reflexi československé sociologie.“
Sociologický časopis 26: 168- 186.

Breen, Richard and Jan O. Jonsson. 2000. “Analyzing Educational Careers: A
Multinomial Transition Model.” *American Sociological Review* 65: 754- 772.

Featherman, David L. and Robert M. Hauser. 1978. *Opportunity and Change*. New
York: Academic Press.

Ganzeboom, Harry B. G., Paul M. De Graaf, and Donald J. Treiman. 1992. “A
Standard International Socio-Economic Index of Occupational Status.” *Social
Science Research* 21: 1 - 56.

- Ganzeboom, Harry B. G., and Donald J. Treiman. 1993. "Preliminary Results on Educational Expansion and Educational Achievement in Comparative Perspective." Pp. 467-506 in *Solidarity of Generations: Demographic, Economic and Social Change, and Its Consequences*. Edited by Henk A. Becker and Piet L. J. Hermkens. Amsterdam: Thesis Publishers.
- Gerber, Theodore P. and Michael Hout. 1995. "Educational Stratification in Russia during the Soviet Period." *American Journal of Sociology* 101: 611- 60.
- Hanley, Eric. 2001. "Centrally Administered Mobility Reconsidered: The Political Dimension of Educational Stratification in State-Socialist Czechoslovakia." *Sociology of Education* 74: 25- 43.
- Heyns, Barbara and Ireneusz Bialecki. 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, George and Ivan Szelenyi. 1979. *Intellectuals on the Road to Class Power*. New York: Harcourt Brace Jovanovich.
- Kreidl, Martin. 2004. „Politics and Secondary School Tracking in Socialist Czechoslovakia, 1948- 1989.“ *European Sociological Review* 20: 123- 139.

Kreidl, Martin. 2005e. „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)

Mare, Robert D. 1981. “Change and Stability in Educational Stratification.” *American Sociological Review* 46: 295- 305.

Matějů, Petr. 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.

Nieuwbeerta, Paul and Susanne Rijken. 1996. “Educational Expansion and Educational Reproduction in Eastern Europe, 1940- 1979.” *Czech Sociological Review* 4: 187- 210.

Róbert, Peter 1991. “Educational Transition in Hungary from the Post-war Period to the End of the 1980’s.” *European Sociological Review* 7: 213- 236.

Shavit, Yossi and Hans Peter Blossfeld (eds.). 1993. *Persistent Inequality. Changing Educational Attainment in Thirteen Countries*. Boulder: Westview Press.

- Simkus, Albert A. 1981. "Historical Changes in Occupational Inheritance Under Socialism: Hungary 1930- 1973." *Research in Social Stratification and Mobility* 1: 171- 203.
- Szelényi, Sonja and Karen Aschaffenburg. 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, Donald J. and Ivan Szelényi. 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].
- Walder, Andrew. 1995. "Career Mobility and the Communist Political Order." *American Sociological Review* 60: 309- 328.
- Walder, Andrew, Bobai Li and Donald Treiman. 2000. „Politics and Life Chances in a State Socialist Regime: Dual Career Paths into Urban Chinese Elite, 1949- 1996." *American Sociological Review* 65: 191- 209.
- Wong, Raymond Sin-Kwok. 1996. "The Social Composition of the Czechoslovak and Hungarian Communist Parties in the 1980s." *Social Forces* 75: 61- 89.

Wong, Raymond Sin-Kwok. 1998. "Multidimensional Influences of Family Environment in Education: The Case of Socialist Czechoslovakia." *Sociology of Education* 71: 1- 22.

Xie, Yu. 1994. "The Log-Multiplicative Models for Discrete-Time, Discrete-Covariate Event-History Data." Pp. Xxx-xxx in: *Sociological Methodology*. Edited by Peter V. Marsden. Oxford: Blackwell Publishers.

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 enrollment 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.

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 centering (see text for details). Standard deviations are not shown for dichotomous variables as they are simple a function of the mean.

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.

Model # and description of tested interactions	Statistics of the test		
	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.*
2. *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 identifiers of mean-replaced missing values.*

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.

Main Effects	Model 1			Model 2		
	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: indicators 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.

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.

Main Effects	Model 4			Model 6			Model 8		
	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: identifiers 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.

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 identifiers of mean-replaced missing values.

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: indicators 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.