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Education Delayed or Education Denied? Evidence on the Historically Variable Role of Delayed Educational Careers in Former Communist Countries

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EDUCATION DELAYED OR EDUCATION DENIED? EVIDENCE ON THE HISTORICALLY VARIABLE ROLE OF DELAYED EDUCATIONAL CAREERS IN FORMER COMMUNIST COUNTRIES

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ABSTRACT

This paper explores why previous research failed to find any empirical evidence confirming the success of "Communist affirmative action" in reducing inequality in access to secondary and tertiary education in Bulgaria, the Czech republic, Hungary, Poland, and Slovakia between 1948- 1989. I argue that scholars have too narrowly focused on ultimate educational attainment of each cohort and have thus overlooked important life-course and historical dynamics of educational stratification in former socialist countries. In this paper I study detailed information on educational careers from the Social Stratification in Eastern Europe after 1989 survey, distinguish the stratification of early and delayed school transitions and compare the differential degree of stratification of early and delayed transitions across cohorts. I show that delayed school transitions were usually stratified less on socioeconomic background than delayed transitions, yet this life-course differential was by no means stable over time. It turns out that delayed school transitions were stratified more strongly in cohorts, in which early transitions were stratified less as a result of the "Communist Affirmative Action". These two offsetting tendencies were overlooked by previous research and combined to produce and overall stable effect of SES on school transitions. I conclude that delayed education careers worked against the success of the egalitarian policies and offered a highly selective second chance for socioeconomically advantaged and politically disadvantaged students. This finding is statistically robust and is identified even in models that control for unmeasured individual-level heterogeneity. I argue that scholars should pay more attention to detailed educational careers and should not only study highest degree completed as otherwise their results may be biased and/or incomplete.

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1 SOCIAL INEQUALITY UNDER COMMUNISM AND INTERRUPTED EDUCATIONAL CAREERS

The transmission of social and economic status from one generation to the next and the crucial mediating role of education in this process are central issues of study in sociology. There is a remarkably uniform tendency in all modern societies for higher status children to obtain more education than their less advantaged peers (Blau, Duncan 1967; Featherman, Hauser 1978; Mare 1980, 1981; Shavit, Blossfeld 1993; Treiman, Ganzeboom, Rijken 1998). Also, those who go furthest in school obtain the best-paying jobs and enjoy the greater prestige associated with them (Blau, Duncan 1967; Featherman, Hauser 1978; Treiman, Yip 1989; Shavit, Müller 1997).

The persistent inequalities in access to education contrast sharply with both public and political concerns about the equality of educational opportunity. Many scholars, motivated by this concern, have looked at educational reforms worldwide, i.e. in advanced industrialized societies, in former Communist countries, and in other societies, to assess the potential impact of political and social institutional settings on socioeconomic inequality in access to education. With the sole exceptions of the Netherlands and Sweden in the second half of the 20th century, where there was a slight decrease in inequality (see e.g. Shavit, Blossfeld 1993; Erikson, Jonsson 1996; De Graaf, Ganzeboom 1993; Jonsson, Mills 1993), and China during the Cultural Revolution, where inequality was almost fully eliminated (Deng, Treiman 1997), sociologists have so far found little evidence that such attempts at reform can be successful (see e.g. Hanley 2001; Hanley, McKeever 1997; Heyns, Bialecki 1993; Matìjù 1993; Nieuwbeerta, Rijken 1996; Róbert 1991; Simkus, Andorka 1982; Szelényi, Aschaffenburg 1993). Only in a very recent study (Kreidl 2004) has some evidence been offered confirming the impact of reform, in the case of the short-lived yet substantively interesting success of policy intervention to equalize access to some types of secondary education in Communist Czechoslovakia.

This consistently negative finding is particularly remarkable in the countries of the former Soviet bloc, where many radical egalitarian reforms were introduced during the Communist period to promote equal access to education. Communist authorities declared equality to be the desired outcome of their policies, and numerous steps were taken in order to equalize 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” (Ulè 1978: 429).

Nevertheless, in most of the former Communist countries these efforts appear to have had little impact on educational stratification. Prior research has failed to identify period effects in educational stratification in the former state socialist countries and there are numerous reasons why this is so, ranging from model miss-specification in both the dependent and independent variables to inaccurate period specification (Kreidl 2004, 2005b). However, after correcting for these inaccuracies it is possible to empirically identify decreases in the effects of parental socioeconomic status (father's education, mother's education, main breadwinner's occupational status) on the odds of progressing from primary to secondary and from secondary to tertiary educational institutions during the periods of Communist "orthodoxy" (see Kreidl 2005b). The purpose of this paper is to revisit the issue and empirically evaluate some other explanations that literature on this topic has suggested and elucidate why researchers have so often failed to detect the effects of "Communist affirmative action" in their empirical data.

Some authors have claimed that the Communist egalitarian policies had little overall impact on educational inequality because students and their families adjusted their attainment strategies to the new institutional environment in an effort to maximize their chances of obtaining the level and type of schooling they aspired to. Following this line of argument an array of possible explanations is offered, from references to bribery and informal contacts with school administrators, bureaucrats and Communist Party representatives (Fiszman 1976; Szelenyi, Aschaffenburg 1993) to more abstract elaborations dealing with the role of cultural capital under Communism (Hanley, McKeever 1997), and theories of a "new class" that suggest 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; Matìjù 1993; Parkin 1971). These arguments usually maintain that because of the existence of a "culture of subversion", "Communist affirmative action" was bound to produce no change in the patterns of educational stratification.

1.1 The resource mobilization theory

Other theoretical streams propose that valuable insight into how Communist egalitarian interventions and people's responses to them worked could be gained by looking in detail at the exact timing of the educational transitions and incorporating an explicit life-course perspective into the conceptual models of educational attainment. This line of theorizing adopts the life-course theory of educational transitions and accepts that there is a systematic, predictable, and theoretically justified difference in socioeconomic background effects between educational transitions that occur early and later during one's life. The life-course theory suggests that the effects of measures of the socioeconomic background on the odds of progressing from one level of schooling to the next declines across subsequent educational transitions because the nature of the parent-child interaction changes (see e.g. Muller, Karle 1993; Mare 1980, 1993; Shavit, Blossfeld 1993).

When applied to former communist societies, this theory maintains that the life-course differential between early and later educational transitions plays out differently in various historical periods as people adjusted their educational plans and attainment strategies to the new institutional environment. Applicants who were denied education initially returned to school after an interruption. These life-course and historical dynamics have been overlooked in previous research in this area, which has focused too narrowly on the final educational attainment of each cohort and disregarded both the path leading to and the timing of educational attainment. It is therefore possible that Communist egalitarian policies did have an effect on the initial allocation of education in a cohort, which faded away as the cohort aged.

What could explain the disappearance of the effects of political intervention over the life course? Some authors have proposed a *resource mobilization theory* as an explanation, maintaining that socio-culturally advantaged but politically disadvantaged students were able to mobilize additional resources in response to being refused an education. It is argued that their *social networks* played a major role in compensating for the negative consequences of the initial policy intervention (cf. Kreidl 2003; Boguszak et al. 1990; Hanley, McKeever 1997; Matìjù 1993). There is some biographical and ethnographic evidence to support this claim. For instance, Kusá (1995) examined (auto-)biographical accounts of Slovak pre-1948 elites to assess the possible impact of parental social networks on the chances children had of overcoming their disqualifying class origin. She suggested that by means of their social capital parents could obtain additional information about schools that were less orthodox in applying Communist ideology, could possibly negotiate non-standard decision-making, and/or could virtually coerce someone into admitting the child to the desired school. My own personal communication with several witnesses suggests that parents were occasionally even able to exploit their social networks to obtain the admission examinations before the actual test date, which again increased the chances of the child being admitted.

A major critique of the resource mobilization argument is that families knew about the policies of positive and negative discrimination even before their children applied to a school, and therefore they should have been able to mobilize whatever resources they had immediately and without having to wait until they were faced with a negative outcome. However, I believe that the resource mobilization argument does make sense, because one of the chief characteristics of Communist egalitarian policies was that they varied greatly from school to school and even from one admission committee to the next within a single school (*cf*. Kusá 1995). It was therefore difficult for people to make a real assessment of the likelihood of the negative impact on them. Many families probably did not expect any discrimination, and those who had some suspicion of this could still harbor the belief that their kids would nonetheless get through. Another possibility is that some families simply took the risk on the theory that they should not mobilize their resources until it is absolutely certain that they are needed.

And finally, students – and parents – probably had a favored school they wanted to apply to but where admission was uncertain, while they were sure that they had enough resources to get into some other, less desirable school. In such a situation it is quite conceivable that they wanted to “try their luck” and applied first to their first-choice school, so they did not use their resources until they failed at least once. Kusá (1995), for instance, presents the story of one child of a non-Communist member of the pre-1948 Slovak government – clearly a student with politically a highly suspect background. This child was not admitted to his preferred secondary school in Bratislava, but he did eventually enroll in and graduate from an elite secondary school in Prague, because the school had quotas for Slovak students. These became more important in the admission process than class and political background. The family, as Kusá reports, learned about the school in Prague through their social networks. The above discussion suggests that families very often only dealt with rejections once they had become a reality and not just a possibility, which supports the resource mobilization argument.

1.2 Deliberately delayed educational careers

There is some biographical evidence that suggests yet another response to Communist school-admission policies by students and their families. In memoirs and biographies dissidents have recalled that an interruption in the educational career of their children proved to be a useful strategy for coping with the anticipated politically motivated discrimination. These students often deliberately forewent the application procedures for secondary and/or tertiary educational institutions in the year they graduated from the previous level and instead accepted a manual job in a factory. Then, once they were in the position to apply as “manual cadres”, they sent in an application to their desired school (see e.g. Vaculík 1983: XXX).

There are reasons to believe that this strategy was indeed potentially successful. First, it is conceivable that the children's own status outweighed that of their parents once they had taken up gainful employment. Second, negative political discrimination applied against the pre-Communist elites was part of a larger *déclassement* campaign launched at the onset of Communism (see e.g. Simkus 1981; XXX). Its aim was, among other things, to re-socialize former “class enemies”. Working as a manual laborer in a factory could certainly have been viewed as a sign of successful re-socialization. As such it may have been sufficient for overriding whatever inauspicious socioeconomic and/or political background a person had. Even one year as a manual laborer often proved to be a sufficient qualification to gain college admission (Vaculík 1983: XXX). Third, all Communist countries had special policies in place to recruit and train working class youth to promote them to become future Communist cadres, managers, and supervisors (see e.g. Hanley, Treiman 2003; Kreidl 2003; Li, Walder 2001; Walder, Li, and Treiman 2000), which gave applicants from the industrial sector a leg up in the admissions process. This additional factor in the admissions process could easily have been the decisive element in converting a potential rejection into an admission.

Both the resource mobilization theory and the theory of deliberately delayed school progressions imply that *delayed educational transitions played a variable historical role in educational stratification in Communist societies*. Then, in the context of regression-type analyses, the estimated slope for the effects of socioeconomic background on delayed transitions should not only be flatter than for direct school-to-school progressions – as predicted by the life-course theory of educational stratification, but also the difference in slopes should vary by historical period. The difference in the level of stratification between direct and delayed progressions should be steeper for those cohorts that were in their early years impacted by communist egalitarian policies than for the cohorts who grew up during years of relatively less stringent educational policies. This is because during periods of Communist orthodoxy many advantaged students were initially denied access to schooling and were disproportionately highly represented among the late entrants. To phrase this hypothesis using the statistical terminology, these theories predict that there is a three-way interaction between individual measures of respondent's socioeconomic background, historical period and the variable differentiating direct and delayed enrollments.

Previous investigations have neglected the distinction between denied and delayed education. Researchers have so far focused too narrowly on 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; Matìjù 1993; Nieuwbeerta, Rijken 1996; Róbert 1991; Simkus, Andorka

1982; Szelényi, Aschaffenburg 1993). By doing so, scholars might have missed a subtle, but important, form of discrimination – the late entry into higher education of children of the former bourgeoisie and intelligentsia.

The primary aim of this paper is to investigate the role that life-course differences in socioeconomic background effects played in maintaining, suppressing, or increasing the effects of political intervention in the process of educational stratification and to test empirically the idea that delayed careers compensated for egalitarian intervention of the Communist government into the educational attainment process. Was it the case that while the "Communist affirmative action" reduced socioeconomic inequality in access to education, its effect was only temporary as delayed educational careers became more selective on socioeconomic background and this brought the overall level of inequality back to the level present before the intervention began? There are survey data (such a the Treiman and Szelényi's (1993) Social Stratification in Eastern Europe 1993 survey – see below for more details) that contain detailed information about each respondent's education career and thus enable to carry out direct comparisons between the initial assignment to school (school vs. no school, types of school etc.) and the ultimate educational attainment of each cohort. Moreover, it is also possible to compare the magnitude of socioeconomic background effects across the "direct" and "delayed" transitions to a given level of schooling, assess the stability of this difference across cohorts, and test the theoretical claim that delayed careers were a particularly frequent "second chance" for higher status children in cohorts whose educational stratification had been initially affected by a policy intervention[1].

[1] This effect could be offset by more intense Communist sponsoring of further education among the "red" working class youth in the "orthodox" periods. As long as communists sponsored early party entrants to obtain

2 THINKING ABOUT EDUCATIONAL ATTAINMENT: DO PEOPLE HAVE A "MASTER PLAN" FOR THEIR EDUCATION?

In order to test the hypotheses outlined above it is necessary to choose a conceptual and statistical model that will adequately represent the relevant social processes of interest. Therefore, this section provides an overview of the development in conceptual thinking on educational stratification and its relationship to life-course research. It highlights some issues that have surfaced in the estimation and interpretation of the variation in socioeconomic background effects on educational progressions across people's life course. Based on this discussion I choose a variety of statistical models, each of which is consistent with a different line of conceptual thinking about educational attainment. Moreover, each of the models is consistent with some biographical and ethnographic evidence describing the thinking and educational attainment strategies of the populations in former state socialist countries.[2]

The early generations of educational stratification researchers employed a rather simple measure of educational attainment, relying mostly on the number of years of completed schooling (see e.g. Blau and Duncan 1967; Featherman, Hauser 1978; Sewell, Hauser 1975). Later, Mare (1980, 1981) revolutionized the field by proposing that educational attainment be best viewed as a sequence of decisions that an individual faces throughout

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more schooling, these young communists were disproportionately from working class backgrounds, and sponsoring was more frequent in "red" periods, delayed educational careers might have worked to further reduce inequality in education. This criticism is not too serious, as long as we control for the political status of respondents – i.e. for their likelihood of being selected for educational sponsorship – and allow for the value of political capital to vary across periods.

the schooling process and introduced a logistic response model of school continuation. The model restricts the base population at risk for a successive transition to those students who have successfully completed all previous transitions, and it models the odds that they will continue to the next level. In a more recent paper Mare (1993) explained that the logistic continuation decision model is a more adequate representation of a person's educational career than years of completed schooling, because it best corresponds to the way in which people accumulate a formal education, "namely in a sequence of irreversible steps" (p. 353).

The logistic response model of school continuation decisions proved to be a more powerful analytical tool than earlier models in a number of ways. First, Mare (1981) showed that linear regression models of highest grade completed on measures of socioeconomic background showed little variation in background effects across cohorts because they were confounding two basic and offsetting trends – the upgrading of educational distribution and the growing effects of socioeconomic background on the process of allocation of schooling. It is only the logistic response model that offers estimates of background effects that are free from the ceiling and flooring effects that result from differences in marginal distributions of educational attainment across cohorts and/or societies.

Second, Mare (1980) argued that it may be useful to dis-aggregate formal school attainment into a series of grade progressions and to analyze the variation in the effects of social background on school continuation decisions, suggesting that not all transitions necessarily require "the same amount of familial resources and structural advantage" (p. 295).

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[2] Some of those claims may be less easily generalized to other societies, but it is obviously an issue worthy of empirical investigation.

Indeed, comparative research based on data from a variety of countries with rather different social, cultural, political, and economic institutions and at different levels of development later accumulated evidence in support of this hypothesis – see e.g. individual chapters in the Shavit and Blossfeld (1993) comparative volume. Many scholars have accepted the explanation that this variation is produced by changes in the parent-child relationship as the child grows up, and that it reflects the growing influence on the child of persons other than his/her parents (see e.g. Muller, Karle 1993). Others claim that this pattern likely results from selective attrition in subsequent transitions on unmeasured individual or family-level factors, including ability, motivation, ambitions, permanent family income, wealth, or even neighborhood characteristics (see e.g. Mare 1980; Shavit, Blossfeld 1993). Methods correcting for family-level sources of unobserved heterogeneity suggest that while both theories may have some explanatory power, they may also apply in varying degrees to different transitions (see e.g. Mare 1993). Nevertheless, the *need to control for all potentially confounding sources of unmeasured heterogeneity* remains a major challenge for scholars who wish to use the school continuation decision model – whether for comparisons of background effects across transitions within countries and cohorts or across nations or cohorts within transitions.

Third, the development of the school continuation model contributed significantly to further advances in comparative studies of educational stratification, because it offered scholars a conceptual tool that standardized the educational experiences of students in many diverse school systems. As a result, students of social stratification were able to estimate the effects of social background on the odds of success in selected key educational transitions – quantities that seem to be most directly comparable across nations. While approaches that pay more attention to qualitative differences within education systems – such as secondary school-tracking – may be more appropriate for some research questions (see e.g. Breen, Jonsson 2000; Gambetta 1996; Kreidl 2004; Lucas 2001), partly and/or entirely suppressing qualitative differences is a strategy that has been used to great effect for comparative purposes for instance by Shavit, Blossfeld and their colleagues (Shavit, Blossfeld 1993).

The arguments summarized above led sociologists to accept the school continuation model as the primary tool for the analysis of equality of educational opportunity. Nonetheless, even the OLS models of highest grade completed continue to be utilized. As De Graaf and Ganzeboom (1993) point out, linear models of the highest grade completed continue to provide valid, parsimonious, and useful summary statistics describing the educational experience of entire cohorts, and as such they are of intrinsic sociological interest. Indeed, there are many scholarly papers that – for reasons outlined by De Graaf and Ganzeboom – have employed linear models to conduct large-scale comparison of educational inequality both over time and across nations (see e.g. Ganzeboom, Treiman 1993; Treiman, Yip 1989; Treiman, Ganzeboom, Rijken 2003 along with chapters in the Shavit & Blossfeld's (1993) edited volume).

However, the school continuation decision model has recently been subjected to criticism from both sociologists and economists. Sociologists have proposed that the model should be extended to encompass qualitative differences within levels of schooling. Breen and Jonsson (2000) were perhaps the first to explicitly argue in favor of a multinomial transition model. They showed that for a population of Swedish students the effects of socioeconomic origin varied between school types within levels of schooling, a fact that could not be incorporated into previous versions of the model. The standard model – unlike their multinomial response alternative – employed binary response variables only. Breen and Jonsson (2000) also demonstrated that not only transition probabilities at later stages depended on the educational pathway the students had followed to reach that juncture, but also socioeconomic background effects varied in size by the student's previous educational trajectory, being stronger for students who followed less common attainment strategies. Similarly, Lucas (2001) argued that there are significant qualitative differences even within the US education system – such as high school tracks - that are not collinear with grade level and therefore cannot be distinguished using the simple version of the school continuation model. Lucas proposed that the sequential nature of Mare's model be maintained, but offered an ordered-probit model to represent the choices students face after completing each grade.

Two prominent labor economists have recently criticized the school continuation model for, among other things, being too loosely behaviorally motivated (Cameron, Heckman 1998). They maintain that the model assumes unnecessarily myopic agents and as such it is not an attractive interpretive tool for economists. Moreover, they argue that the pattern of the declining background effects across transitions relies on arbitrary distributional assumptions imposed on the data and would not appear if other quantities were used to represent the underlying processes of interest. Instead they have devised a simple low dimensional model that, in their view, has stronger behavioral foundations and enables corrections for dynamic selection bias. This model describes the data as well as the traditional model does, but it uses far fewer parameters. Therefore, Cameron and Heckman strongly favor it over previous representations of the schooling process.

Cameron and Heckman's preferred model is an ordered-discrete choice model (Cameron, Heckman 1998: xxx). It assumes that people know their endowments, know the cost of investment, and know the payoffs for certain levels of education. Based on the evaluation of these facts, according to Cameron and Heckman, people obtain just enough schooling to maximize their net benefits. Conceptually this means that people have an “educational master plan” that guides them throughout their lives. The dependent variable in their analysis is the years of school completed, treated as an ordinal variable, i.e. it allows for unequal distances between subsequent years. Effectively this implies that Cameron and Heckman are calling for a return to years of school completed as the focus of analysis, abandoning school transitions, and suggesting that a different functional form would resolve the problems attending the use of OLS regression.

There are a number of reasons why sociologists would find it hard to accept the model Cameron and Heckman have proposed. First, as suggested above, their model is not – unlike the school continuation decision model – suitable for the study of tracked education systems and other qualitative differences between educational credentials. Second, Lucas (2001) suggests that it is possible to maintain the school continuation model and get around some of the critical points raised by Cameron and Heckman if, for instance, a richer data set can be used, one which would provide, in addition to time-invariant explanatory variables, some time-varying explanatory variables such as measures of school achievement (school grades, test scores etc.) and high school track placement in the previous year.

Third, the ordered discrete-choice model presupposes that actors make educational choices in stable environments, which obviously ignores the many choices made by admissions committees, principals, and school administrators. These choices obviously limit the choices that students can make and, moreover, they are not – and in the recent past never have been – blind to students' statuses (see e.g. Walster, Cleary, Clifford 1971). Fourth, while most teenagers and/or their families express a motivation to go to college, only some of them in reality do enroll at the post-secondary level. For example, Dominitz and Manski (1996) present results from a sample of Wisconsin high school students in 1992 showing that an overwhelming majority of them expected to obtain a Bachelor's degree by the age of 30. For instance, 50% of the female students believed that their chances of graduating from college by the age of 30 were 95% or better, while 50% of male students believed that their odds of graduating from college by the age of 30 were 90% or better. Obviously, these estimates would imply far higher graduation rates than are observed in reality.

3 THE EDUCATION SYSTEMS IN THE FORMER COMMUNIST COUNTRIES

The countries in Central and Eastern Europe reformed and unified their education systems immediately after the Communists seized power in the late 1940s. The new system offered compulsory primary education for all pupils, which took between seven and nine years, depending on the country and the period (see e.g. Szelenyi, Aschaffenburg 1993; Heyns, Bialecki 1993; Simkus, Andorka 1982; Kreidl 2004). Students left primary school approximately at the age of 14 and could choose from several different options, which for analysis can be clustered into four basic categories. Obviously students could terminate their education and (1) not attend any secondary school,[3] or they could

[3] In the late 1970s and early 1980s most state socialist countries in Europe enacted ten-year compulsory education and thus most students had to enroll at the secondary level for at least one but usually for two years. The only exception were pupils who were "held back" during elementary education and had to repeat a grade or two.

go to one of the following institutions: They could attend (2) *a lower secondary vocational school* (sometimes referred to simply as vocational schools, or as lower secondary schools), which lasted from two to three years, after which they were not able to progress to the university level. Or, students could attend (3) one of a variety of *four-year professional secondary schools*, including various types of vocational schools and professional schools, which trained students in such diverse fields as civic and electrical engineering, accounting, nursing, agriculture, and administration. Instruction in these schools included both vocational and academic courses, but, in comparison to the fourth education option (see below), they were more explicitly oriented towards immediate labor market entry. Nonetheless, students at these schools attained a complete secondary education diploma and were formally eligible to attend university. The last option students had was to attend (4) *an academic secondary school,* which was primarily designed to prepare students for tertiary education and represented the most natural, though not the exclusive, steppingstone to university.

The tertiary sector offered a varied set of curricula, which differed in length and focus – ranging from short-track schools to six-year programs at medical schools. However, it is more difficult to grasp the qualitative differences within the tertiary sector using a set of clear-cut and internationally comparable categories and - instead of trying to create such classifications - the analysis relies primary on the distinction between individuals who did enter university and those who never studied at the tertiary level. There are a number of articles that give extensive reviews of the state socialist education systems in different countries (see e.g. Matìjù (1993), Szelényi, Aschaffenburg (1993), Kreidl (2004), Róbert (1991), or Heyns, Bialecki (1993)), so the interested reader is referred there for more detail.

4 AN ANALYTIC STRATEGY FOR MODELING DIFFERENCES BETWEEN STANDARD AND DELAYED PROGRESSION WITHIN EDUCATIONAL TRANSITIONS

Despite all the criticism of the idea that people have a “master plan” regarding their education and the problem of the compatibility of the econometric model's assumptions with some existing research, it is nonetheless an intriguing possibility. As demonstrated above, there is some biographical evidence pointing to the fact that at least some people during the Communist period did indeed have a “master plan” for their education, which guided them to delay the attempt to make a particular school transition, whether this was the transition to secondary school or the transition to university. In this section the analytical possibilities implied in such a strategy will be explored, along with more traditional models that strictly adhere to the logic of the school continuation model.

The traditional model fails to distinguish between early and delayed progressions from one level of schooling to the next. However, if behavioral theory leads us to believe that such a distinction is crucial, then it is necessary to devise models that accommodate it. If, for instance, the transition to secondary education is to be modeled using the traditional model, a multinomial logistic regression with four available outcomes can be used (no secondary education, vocational secondary education, professional secondary education, and academic secondary education). If the intention is to extend the traditional model to accommodate the differences in background effects between direct and delayed entry into secondary school, then a *discrete-time survival model with competing risks* can be used. Conceptually this is equivalent to assuming that everybody wants to continue their schooling at the secondary level immediately after completing primary education but that only some of the applicants succeed. Among the rest everybody keeps trying and some indeed succeed later, while some never succeed. Provided that there is a single variable distinguishing direct and delayed transitions ("DELAYED" in equations below), the multinomial logistic survival model can be written as:

$$\ln\left(\frac{P_V|X}{P_N|X}\right) = a_1 + \sum_{i=1}^{k} b_{i1} X_i * + c_1 * DELAYED + \\ + \sum_{i=1}^{j} d_{i1} * X_i * DELAYED \qquad \text{(Eq. 1)}$$

$$\ln\left(\frac{P_P|X}{P_N|X}\right) = a_2 + \sum_{i=1}^{k} b_{i2} X_i * + c_2 * DELAYED + \\ + \sum_{i=1}^{j} d_{i2} * X_i * DELAYED \qquad \text{(Eq. 2)}$$

$$\ln\left(\frac{P_A|X}{P_N|X}\right) = a_3 + \sum_{i=1}^{k} b_{i3} X_i * + c_3 * DELAYED + \\ + \sum_{i=1}^{j} d_{i3} * X_i * DELAYED \qquad \text{(Eq. 3)}$$

where P_N is the probability of not enrolling in any secondary education, P_V is the probability of enrolling in vocational training, P_P is the probability of enrolling in a professional secondary school, and P_A is the probability of entering an academic secondary school.

By definition $P_N + P_V + P_P + P_A = 1$ within each risk set. Further, Xi is a vector of k explanatory variables including father's education and mother's education, measured in years of schooling, and the occupational status of the main earner in the family, and some control variables (in this case the respondent's sex, number of siblings, parental political status, cohort, and country). “DELAYED” is a dichotomous variable, distinguishing between early and delayed entry, i.e. a measure of the underlying time dimension in the survival model. Finally, the model contains interactions between j of the k explanatory variables and the “DELAYED” dummy. Coefficients d_{i1}, d_{i2}, d_{i3} (the second subscript indexes the equations in the multinomial model) associated with those interactions would then be the crucial test for the stability of background effects across early and delayed progressions in this particular transition. Obviously, as previous research has shown, the estimates of these quantities may suffer from omitted variable bias.

This model can also be used to evaluate the hypothesis that the difference between early and delayed background effects varied systematically across historical periods, being greater in periods of Communist orthodoxy and smaller in other periods. This evaluation can be performed simply by adding a three-way interaction between each measure of SES (father's education, mother' education, ISEI), time (“DELAYED”) and cohort (“red” vs. “other”) in addition the necessary lower-level terms. All the above models can be simplified to a binary outcome and are thus also a suitable instrument for modeling entry into post-secondary education. These models are obvious extensions of the traditional school continuation decision model and are conceptually fully compatible with it.

However, if we believe is that people intentionally delayed the entry into the next level of schooling, it may be preferable to adopt a conceptually different model. While this doesn’t imply that people have a “master plan” for their entire education career, it leads to a less ambitious assumption that people have a master plan for each particular educational transition. The plan incorporates the explicit decision of whether to apply to the next school immediately upon graduation or after a delay. Under this model, people facing the transition to secondary school would not decide between the four options outlined above (no school, vocational secondary school, professional secondary school, and academic secondary school), but would instead choose from the following set of alternatives: direct entry into a vocational school, direct entry into a professional school, direct entry into an academic school, delayed entry into a vocational school, delayed entry into a professional school, and delayed entry into an academic secondary school (and an implicit choice of never progressing further).

If this is the right behavioral model, then its statistical equivalent is a discrete-choice model, which can be estimated using McFadden's (1973) *conditional logit model*. This model assumes that each of i individuals faces a set of j=1, 2,…J_i options (here J_i=6 for all individuals). Let y_{ij}=1 if individual i chooses option j and 0 otherwise. Let also x_{ij} be the vector of characteristics describing option j for person i. The explanatory variables may include characteristics of the options (here type of secondary school chosen, direct vs. delayed entry) and interactions between the options' characteristics and characteristics of individuals (e.g. interactions between measures of respondent's socioeconomic background and the “direct” vs. “delayed” dummy).[4] The model does not include an intercept and does not enable an estimation of the main effects of person-level characteristic on the choices (Allison 1999: xxx). The model can be formally written as follows:

$$\Pr(y_{ij}=1)=\frac{e^{bx_{ij}}}{e^{bx_{i1}}+e^{bx_{i2}}+\ldots+e^{bx_{iJi}}} \qquad \text{(Eq. 4)}$$

[4] The model would also be easily simplified to accommodate the situation with fewer choices, such as the direct vs. delayed university entry problem. The only difference between these models would be that a conditional logit model for university entry would contain only one option-specific characteristic, namely the “direct vs. delayed entry” dummy. Otherwise, the models are identical.

The standard multinomial logit can be shown to be a special case of the conditional logit model (Allison 1999). It is also true that the conditional logit model is identical to the fixed-effect multi-level model. One of the most powerful features of the conditional logit/fixed-effect model is that it can effectively control for the potentially confounding effects of all measured and unmeasured characteristics of respondents (Allison 1999: xxx). This is an extremely powerful feature of the model, as it is, unlike for instance the traditional school continuation model, immune against bias due to selection on unmeasured respondent-level characteristics. Consequently it offers statistically more robust evidence for empirical assessment of some hypotheses. This model, however, has some undesirable features, such as the inability to estimate the effects of stable individual-level variables on the choice between options (Allison 1999: xxx), which prevents it from becoming the analyst's primary choice. Fortunately, this is not a serious problem in this particular analysis, because my here aim is to determine whether there are interactions between individual-level and option-level characteristics. Because the choice between the two behavioral models of school choice is - as indicated above - unclear and complicated, both will be used and the results compared in order to gain a deeper insight into the role of delayed school progressions for educational inequality during the Communist period.

5 DATA AND VARIABLES

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 the decision was made not to include it 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 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, 8,997 completed secondary education during Communism and were at risk of progressing to the tertiary level. However, some respondents were lost because the survey did not record their gender. Therefore, only 17,935 primary school graduates and 8,989 secondary school graduates are use in the analyses.

The explanatory variables include the timing of attainment (direct vs. delayed entry), as well as 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 siblings[5] to minimize the influence of extreme values in the analysis.[6] A dichotomous variable was used to distinguish between men and women, and 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 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 during the year of extreme Communist orthodoxy and other respondents. I adopt the definition of the orthodox or "red" periods proposed elsewhere (see Kreidl 2005b) and define "red" years as 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. The descriptive statistics (means and standard deviations) for all the independent variables in the analysis of either the transition to secondary or tertiary educational institutions are presented in Table 1.

[5] 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.

[6] 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.

5.1 Direct and Delayed Transitions to Secondary and Tertiary Educational Institutions: A Look at the Distribution of the Dependent Variables

It was more common for students to progress from one level of schooling to the next immediately after graduating than it was to postpone enrollment. Nevertheless, delayed progressions were not uncommon. According to survey data utilized in this paper only 24% of all primary school graduates in Bulgaria, the Czech Republic, Hungary, Poland, and Slovakia between 1948 and 1989 did not continue on to secondary school directly (see column 6 in Table 2).[7] Of those who did not continue directly, however, only 72% never did so, while the remaining 28% did, but only after a delay of a year or more.

Among direct entrants into secondary schools, vocational secondary schools were the most frequent option, followed by professional secondary schools, and then academic secondary schools. Among delayed entrants the choices were ranked similarly, but a larger share of students chose vocational schools (see column 7 in Table 2). Table 2 also offers a comparison of the distribution of students with respect to direct and delayed school entry in "red" and "other" cohorts. This comparison shows that "no secondary education" was a more frequent option during the period of Communist orthodoxy, while during this same period professional and academic schools were chosen less often than in the less oppressive times. This should be no surprise, as the "red" cohorts are mostly concentrated in the early years of Communism, and the difference necessarily results from the expansion of education that all the former Communist countries experienced.

[7] Throughout this paper direct transitions are considered to be enrollment at the next level that occurred during the same calendar year as graduation from the antecedent level. The vast majority of people graduated during/after the spring semester, and schools accepted new students in September.

Table 3 presents the distribution of students by type of secondary school and enrollment timing for individual countries. It is evident that, across all the cohorts, the percentage of primary school graduates who did not continue directly to the secondary level varied somewhat between countries, ranging from 32% in Hungary and Bulgaria to19% in Slovakia and 17% in the Czech Republic. Of those who did not continue directly, the share re-entering school later was again approximately the same in all the countries observed – 24% in Bulgaria, 29% in the Czech Republic, 34% in Hungary, 25% in Poland, and 28% in Slovakia (see Table 2).

The ratio of direct and delayed entries was somewhat more balanced at the tertiary level. Table 4 shows the percentage of students in the transition to tertiary education displayed by country and enrollment timing. Overall 17% of secondary school graduates entered college directly, and out of the rest 16% experienced a delayed entry. Direct progressions were more common in the Czech Republic (23%) and Slovakia (22%) than in Hungary (13%), Bulgaria (13%), or Poland (13%). Delayed progressions, on the other hand, occurred more often in Hungary (25%) and Bulgaria (19%) than in the Czech Republic (11%) and Slovakia (9%), with Poland (17%) ranking in between (see Table 4). Interestingly, however, there is little difference between the “red” and other cohorts in the incidence of direct and/or delayed university enrollments.

6 RESULTS

The first results reported here refer to the traditional model of school progressions for both the entry into secondary school and the entry into university, and these are followed by the results from the discrete-choice conditional logit model. I rely mostly on the criteria of classical inference in evaluating the model fit in each section, yet I also show the Bayesian Information Criterion (BIC) for the interested reader. Finally, results based on the school continuation decision model and on the discrete choice model are compared and controlled.

6.1 Modeling secondary school entry using the school continuation model

The modeling of entry into secondary school begins with a simple additive model of all covariates (Model 1 in Table 5), which serves merely as a benchmark, against which the other models are evaluated. The estimated coefficients of Model 1 are presented in Table 6. Model 2 adds three two-way interactions into Model 1: father's education * “DELAYED” dummy, mother's education * “DELAYED” dummy, and ISEI * “DELAYED” dummy. These three variables appear to greatly improve the fit according to classical inference criteria (L^2= 66.7 with 9 d.f., see Table 5). Owing to concerns about the possible collinearity of these interaction effects, they were also added one by one to Model 1 to create Models 3, 4, and 5. Each of these additions is clearly a statistically significant improvement on the model fit (see Table 5). It seems clear, therefore, that all the three examined background variables do interact with the delayed entry variable, which suggests that their effect on the odds of progressing from the primary to the secondary school indeed differs between direct and delayed entries.

The estimated coefficients for Model 2 reported in Table 6 confirm expectations: the interaction between father's education and time is statistically significant at standard significance levels (p=0.05) in all three equations, and the interaction between mother's education and time is significant in two out of the three equations in the model.

On the other hand, the interaction between ISEI and time fails to reach the standard significance levels in every equation, and where it approaches the 0.05 level it tends to be positive, and not negative as predicted in the theory (see Table 6). However, when this interaction is estimated separately in a model that doesn't contain any of the other two-way interactions as in Model 5, it turns out to be significant and negative too (estimated coefficients for this particular model are not shown, but are available from the author upon request). In sum, there is rather convincing evidence that the effects of the education of both parents and the effect of the main breadwinner's occupational status decline between direct and delayed transitions to secondary school.

Models 6, 7, 8, and 9 extend the previously estimated models by adding interactions between all three measures of socioeconomic background and cohort using a simple dichotomous variable to contrast "red" and "other" cohorts. They also add interactions between the dichotomous variable for "red" cohorts and a dichotomous variable for "delayed" entry. These models were created to serve purely as a benchmark, against which the more elaborated containing also three-way interactions models can be compared. The estimated coefficients of Model 6 are shown in Table 7.

The crucial tests for my hypotheses are the three-way interactions between measures of socioeconomic background, cohort, and time. These are added in Models 10–13; Model 10 adds it to Model 6, and Models 11, 12, and 13 to Models 7, 8, and 9, respectively. Overall, these interactions show only marginal statistical significance. For instance, the contrast between Model 6 and Model 10 yields $L^2 = 16.3$ with 9 degrees of freedom ($p=0.061$), which is only marginally significant according to classical inference (see Table 5). Because it is likely that, given that the three background measures are correlated, the three-way interactions will also be correlated, it is worthy to add them to the model one by one in order to check for multicollinearity. This strategy is represented in Models 11, 12, and 13, and it reveals more clearly which of the interactions are statistically significant and interpretable. It turns out that whereas the contrast between Models 12 and 8 is strongly statistically significant (L^2=12.7 with 3 d.f.), the contrast between Models 11 and 7 and the contrast between Models 13 and 9 are far from significant. Clearly, only the mother's education interacts significantly with cohort and time in a three-way interaction, while father’s education and ISEI do not. The estimated coefficients for the complete model (Model 10) are reported in Table 7.

The three-way interaction between mother's education, cohort and enrollment timing is of enormous substantive interest. Its nature is best illustrated in Table 8, which presents the net effects of mother’s education in each equation of the multinomial logit for each combination of cohort and attainment timing. For instance, in the “non-red” cohorts each additional year of the mother's education increases, net of other factors, the log odds of a direct entry into a vocational school by 0.092, into a professional school by 0.181, and into an academic school by 0.202 (see Table 8) and are thus much steeper than in the “non-red” cohorts. In the “red” cohorts, the respective slopes are 0.04, 0.08, and 0.113 for direct entry into vocational, professional, and academic school – a clear decline in comparison with other cohorts. In the “non-red” cohorts, the estimated effect of an additional year of mother's education on the log odds of delayed entry into a vocational school is 0.01, into a professional school 0.09, and into an academic school 0.138, i.e. the effects decline in all cases to the level of the effects for direct entry in the “red” cohorts. Finally, the slopes for delayed entry in the “red” cohorts are 0.09, 0.204, and 0.216 for entry into vocational, professional, and academic schools, respectively (see Table 8). In other words, the slopes for delayed entry during “red” periods are as steep as the slopes for direct entry in the normal cohorts. Moreover, the slopes for late entry in the “red” cohorts are steeper than the slopes for direct entry in the same cohorts. This is exactly the pattern that the theory predicted!

Apparently, the difference between early and late transitions to a secondary school exists for all three of the background variables, whereas it is for mother's education that this difference plays a historically variable role. This should come as no surprise, as we know from previous research that it was mainly the effect of maternal education that was reduced during the years of Communist orthodoxy (see Kreidl 2005b). Given that “Communist affirmative action” policies did not mitigate the effects of the other background variables, it could hardly be expected that their effects on the later transitions would compensate for the equalization of socio-economic inequality. Hence, I conclude when Communist egalitarian policies indeed reduced inequality in the allocation of schooling among direct entrants, this success was of limited duration as it also meant that delayed school entries became more stratified on socioeconomic background. These two offsetting tendencies then combined to produce overall stable background effects that previous research has found.

6.2 Modeling university entry using the school continuation model

The modeling strategy to investigate the predictors of entry into university is the same as in the previous section. First, a simple additive model of all covariates is estimated (Model 14, see Table 9) and then two- and three-way interactions are added to it. Models 15–18 contain a two-way interaction between parental statuses and delayed entry, first added collectively (Model 15), and then one by one (Models 16, 17, and 18) to check for possible multicollinearity. These models reveal that the degree of stratification on socioeconomic background differs between delayed and direct entry into the university (L^2 for the comparison of Models 14 and 15 is 9.8 with 3 d.f., which implies a p-level of 0.02, see the lower panel in Table 9). All estimated coefficients of Model 15 are presented in Table 10. Their detailed inspection confirms that while father's education and mother's education do not seem to interact with attainment timing, parental ISEI does and the interaction is negative, suggesting that the odds of a delayed entry into university are less affected by parental occupational standing than direct entries are.

In the next step the Model 15 (and Models 16, 17, and 18) is extended to include also a full set of two-way interactions between measures of socio-economic background and cohort and in addition it also has an interaction between “red” cohorts and “delayed” entry into post-secondary schooling. This creates Models 19, 20, 21, and 22. As before, these models serve mainly for comparison with the more complex models reported below. The estimated coefficients from Model 19, which contains also interactions between all three measures of family socioeconomic background with the other variables, are presented in Table 11. The fundamental test of my hypotheses is the addition of three-way interactions, as performed in Model 23. Models 23, nonetheless, isn’t a statistically superior to Model 19 - a comparison of Model 23 and Model 19 produces L^2= 4.5 with 3 degrees of freedom (p=0.213). Thus I conclude that there is no evidence that measures of socioeconomic background, time and cohort would interact in a three-way fashion. The situation changes somewhat when these interactions are considered one at a time to reduce collinearity between the interaction effects as in Models 24, 25, and 26, which offer a slightly stronger evidence that individual measures of parental socioeconomic background interact with attainment timing, and cohort. Nonetheless, the p-value associated with the test for each interaction never drops below 0.1. The evidence from the school continuation decision model indicates that there is no evidence of a three-way interaction between respondent's socioeconomic background, attainment timing, and cohort. Hence, it appears that delayed university enrollments never played the compensatory role that was predicted in the theory and I conclude that the difference in socioeconomic background effects on direct and delayed entry into the university was stable across cohorts.

6.3 Modeling secondary school entry using the conditional logit model

As indicated in the theoretical sections of this paper, the school continuation decision model may be severely impaired by the presence of unobserved heterogeneity – a common problem in all non-experimental research, which, unfortunately, is highlighted in survival models. Nevertheless, a fortunate feature of the behavioral model that assumes that people plan their education and its timing and employ strategies that include a voluntary, yet temporary withdrawal from school is that it leads to a statistical model that explicitly controls for all potentially individual-level confounding factors, i.e. the conditional logit model. In this and the subsequent sections the conditional logit model is estimated for the entry into secondary education and for entry into university and the results are compared to the survival models estimated earlier.[8]

[8] The conditional logit model is a model of school choice and therefore people who never attended any school at the given modeled level cannot contribute anything to the estimation and are dropped from the data file.

The simplest conditional logit model of the choice of secondary school contains only two dummy variables that contrast the type of school chosen (professional vs. vocational, academic vs. vocational) – this is Model 27 reported in Table 12. The next model (Model 28) adds yet another option characteristic – the time dummy that contrasts the direct and delayed entry. I use this model purely as a benchmark and comparison model and its estimated coefficients are presented in Table 13. They conform to the general expectations: both professional secondary schools and academic secondary schools were chosen less frequently than vocational schools. Moreover, the odds of a delayed enrollment were also significantly lower than the odds of a direct school continuation (see Model 28 in Table 13).

Then I estimate a series of models, which include both option-level covariates and the interactions between individual-level covariates and option-level covariates. The variation in the size of the background effects across early and delayed transitions is examined first by adding three interactions to Model 28 – the interaction between time and father's education, between time and mother's education, and between time and parental ISEI. When all three interactions are added at once, a significantly better fit is yielded (L^2 comparing Models 29 and Model 28 is 291.1 with just 3 degrees of freedom – see Table 12). When these interactions are added one at a time I also obtain statistically highly significant differences, which is an indication that the results are satisfactorily robust (see Models 30, 31, and 32 as contrasted with Model 28 in Table 12). Clearly, even the conditional logit model, i.e. a model that is free from all potentially confounding effects of unobserved individual-level factors, confirms the finding that delayed school enrollments are less stratified on socioeconomic background than direct entries. Table 13 then reports estimated coefficients of Model 29 – the model with all three-way interactions. We can see that, for instance, the effect of father's education on the odds of experiencing a successful transition declines by 0.086 between direct and delayed enrollments, the effect of mother's education declines by 0.103, and the effect of ISEI drops by 0.005 (see Table 13). The declining effect of ISEI, however, seems to be only marginally statistically significant.

In the next step, the interaction between cohort and late enrollment are added to Model 29 to create Model 33 – a model that serves as a comparison model only and I do not aim to interpret its coefficients here. Then the essential three-way interactions visible in Model 34 are added to Model 33. This step improves the model fit significantly (L^2=10.7 with 3 d.f., see Table 12). Estimated coefficients for Model 34 are displayed in Table 13. There is a statistically significant three-way interaction that involves mother's education (with cohort and time), but no three-way interaction involving parental ISEI and cohort and time or father’s education and cohort and time (see Model 34 in Table 13). However, when each of the three-way interactions is added separately, there appears to be also a statistically significant three-way interaction between father’s education cohort and time[9]. The Table 13 also shows the estimated coefficients of this more restricted model (Model 35).We have fairly robust and convincing evidence that the life-course differential in socioeconomic background effect wasn’t stable over time and changed in response to transformations of the larger political and social environment. Hence, delayed educational careers served

[9] Statistics of fit for this particular model are not displayed here, because their evaluation would require that many antecedent, intermediate models be displayed as well, which would take up too much space, wouldn’t be necessarily very informative. Nonetheless, all these intermediate models and all the statistics of fit may be obtained from the author upon request.

as an avenue for politically discriminated individuals to overcome their disqualifying origin and thus helped make the “Communist affirmative action” an inefficacious enterprise.

In general, the conditional logit model of secondary school choice confirms the findings from the school continuation decision models: later progressions to secondary schools were less stratified on socio-economic background than early transitions. This difference, however, was not stable in time and played a distinct and historically variable role during the Communist period. More specifically, the difference in background effects on early and late transitions partially or completely disappears in “red” cohorts, when delayed school transitions compensated for the political interventions that reduced socioeconomic inequality in the direct entry to secondary schools (see Models 34 and 35 in Table 13). Again, this confirms the initial theory presented in this paper.

6.4 Modeling university entry using the conditional logit model

The conditional logit model of college entry contains only one option-level characteristic - the dichotomous variable contrasting direct vs. delayed progressions. A simple model containing only this explanatory variable is Model 36 presented in Table 14. The model is obviously statistically superior to the null model and its coefficient indicates that delayed university matriculations were less frequent that direct ones during the Communism (see Table 15). Model 37 then confirms that the background effects differ between direct and delayed transitions and an inspection of estimated interaction effects in Table 15 verifies that the interactions have the anticipated direction and the effects of parental statuses are indeed weaker for later transitions than for direct ones.

An interaction between time and cohort is then added to create the benchmark model for the subsequent models that will also contain the three-way interactions of interest. This intermediate model is Models 38 presented in Table 14. However, when is this model extended with the addition of the three-way interactions no evidence surfaces that these interactions are statistically significant (see Table 14 for the individual contrasts between models). Moreover, even if these interactions were added one at a time (details of this part of the analysis are not shown) they still appear to be statistically insignificant. Nonetheless, the estimated coefficients and standard errors for Model 39 are presented in Table 15. In sum, the conditional logit model gives no evidence to support the hypothesis of the variable historical role of delayed progression to university and leads to the same conclusion we obtained using the traditional hazard model of university entry.

7 CONCLUSIONS AND DISCUSSION

This paper examines the explanation advanced by some scholars as to why previous research failed to empirically identify any effect of "Communist affirmative action" on socioeconomic inequality in access to secondary and tertiary schooling in the countries of the former Soviet bloc. The argument was that Communist egalitarian policies were successful in reducing inequality in direct school to school transitions, but the initial success of the intervention was later overridden by the more selective school re-entry among the previously rejected applicants. The theory proposes that the stratification of delayed school enrollments varied by historical period. In periods of communist orthodoxy direct school transitions were stratified less than in other cohorts as a result of the egalitarian policies and delayed cohorts were stratified more strongly as a result of people's adjusted attainment strategies. In periods of relative liberalization the relationship was reversed: early transitions were stratified more strongly than in orthodox periods and delayed ones were stratified less than in "red" cohorts. It seems that people indeed adjusted their education plans to their institutional and political context, and socio-economically advantaged and politically disadvantaged students were particularly likely to utilize delayed educational careers to obtain education they had been denied initially.

This pattern produced an apparent lack of changes over time in the overall effects of parental SES on the odds of school transition observed in previous studies, as prior research ignored the time dimension of attainment and focused too closely on the ultimate educational attainment of each cohort. By doing so it overlooked a subtle, yet important distinction, namely the delayed start of the children of the advantaged classes. Moreover, it also neglected the historically variable role played by the life-course differentiation of the stratification processes in former socialist countries and thus also overlooked the dynamic interplay between social and political institutions of a nation and individual actors' life strategies. Given the long-standing interest of sociology in the relationship between micro- and macro- level processes, this omission is most regrettable.

The initial hypotheses of this research were confirmed in the analysis for the transition to secondary schools, while the analyses of the transition to university were much less conclusive. This distinction could be the result of an insufficient sample size for the transition to the university level, but it could also stem from substantive differences in the process of educational attainment at the secondary and tertiary levels. Yet, we lack any theories that would offer a substantive and sustainable interpretation of the difference between transitions to secondary and tertiary educational institutions during Communism.

Obviously, a re-analysis of the issues raised in this paper based on a different and preferably larger data set would increase our confidence in the results reported in this study.

This paper proposes that there is considerable scholarly benefit from explicitly incorporating the time dimension into investigations of the educational attainment process. Moreover, it may not only be adequate, but also beneficial to adopt the idea that people have an educational "master plan", which may at some point lead them to a voluntary and planned withdrawal from school. This paper also maintains that such a temporary interruption in the educational career may be rational. This claim is less ambitious than the idea of some labor economists that people have a life-long educational master plan (cf. Cameron, Heckman 1998), as here the "master plan" is limited to being a strategy for how to best progress from one level of schooling to another. Nevertheless, this conceptual assertion makes it possible to estimate a series of models, which are robust in the presence of unmeasured variables that may bias standard school continuation models. Yet, given some conceptual ambiguities, I adhered both to the school continuation decision model and to the discrete-choice model and compared the results using both.

The conditional logit model, however, confirms the results yielded using the traditional school continuation model. The application of the conditional logit model – i.e. a multi-level fixed-effects model – along with the more traditional school continuation model has produced three important contributions to the educational stratification literature. First, it confirmed that in most circumstances delayed school transitions from one level of schooling to the next are stratified less on socioeconomic background than direct school to school progressions are. This was true for the transition to secondary as well as tertiary educational institutions. This result was obtained even when using models that by definition control for all measured and unmeasured respondents' aspects and thus do not suffer from selection bias. This finding supports the life-course theory of educational stratification more strongly than previous research did.

Second, it showed that delayed transitions might in fact be stratified more strongly than direct transitions under some circumstances. Hence, the universal generalizability of the life-course theory of educational stratification is clearly rejected by this finding.
This occurred, for instance, during the years of severe Communist orthodoxy, and it reflects the effort of upper status students and families to compensate in later life for disadvantages inflicted by political actors early on.

Thirdly, because the role of the life-course in educational stratification is not necessarily stable over time, we should take its contextual variability into account in future comparative research on inequality in access to education. Most notably, it should be subjected to strict scrutiny whenever, for instance, there is a suspicion that a change in the incidence of interrupted educational careers is occurring in populations. Similarly, it ought to be explicitly incorporated into comparative studies of educational stratification. It seems that the greater is the differences in the incidence of interrupted careers between countries or historical periods, the greater may be the need to explicitly incorporate it into the research design. Clearly, there seems to be an increasingly important need to look into delayed educational progressions in more detail, as an increasing incidence of temporary dropouts in many societies is being observed (XXX). Sometimes, these trends are even augmented by the institutional transformations of the education system, which are often motivated by governments and educational administrators with the aim of opening up the school system for easy re-entry and thus reducing inequality in access to schooling. Research reported in this paper, however, highlights that the potential of such efforts is heavily contingent upon the larger social, political, and cultural context, and their success is not obvious.

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9 TABLES AND FIGURES

Table 1: 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	- -	0.46	- -
Father's education	9.23	3.50	10.00	3.57
Father's education missing	0.05	- -	0.03	- -
Mother's education	8.26	3.22	8.90	3.24
Mother's education missing	0.02	- -	0.02	- -
No. of siblings	1.96	1.28	1.65	1.17
No. of siblings missing	0.02	- -	0.02	- -
Main earner's ISEI	35.0	14.4	38.6	15.9
Main earner's ISEI missing	0.06	- -	0.06	- -
Parents CP members	0.22	- -	0.27	- -
Parents CP membership-missing	0.08	- -	0.06	- -
"Red" cohort	0.13	- -	0.08	- -
Country				
Bulgaria	0.18	- -	0.28	- -
Czech Republic	0.24	- -	0.22	- -
Hungary	0.19	- -	0.16	- -
Poland	0.17	- -	0.14	- -
Slovakia	0.21	- -	0.21	- -
Number of cases at risk	17,935		8,989	

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 2: Percentage distribution of first secondary school attended by enrollment timing and cohort of primary school graduation, Central and Eastern Europe, 1948- 1989. Number of cases at risk for each cohort/enrollment timing combination in parentheses. Total N= 17,935.

	"Red" cohorts		Other cohorts		All cohorts	
	Direct entry	Delayed entry	Direct entry	Delayed entry	Direct entry	Delayed entry
None	39%	73%	22%	72%	24%	72%
Vocational secondary	32%	17%	34%	16%	34%	16%
Professional secondary	15%	7%	25%	7%	24%	7%
Academic secondary	14%	3%	19%	5%	18%	4%
Total	100% (2,259)	100% (873)	100% (15,675)	100% (3,518)	100% (17,935)	99% (4,391)

Table 3: Percentage distribution of first secondary school attended by country, enrollment timing, and cohort of primary school graduation, Central and Eastern Europe, 1948- 1989. Number of cases in parentheses. Total N= 17,935.

A. Bulgaria	"Red" cohorts		Other cohorts		All cohorts	
	Direct entry	Delayed entry	Direct entry	Delayed entry	Direct entry	Delayed entry
None	50%	83%	30%	75%	32%	76%
Vocational secondary	6%	9%	4%	8%	5%	8%
Professional secondary	15%	5%	37%	12%	35%	10%
Academic secondary	29%	4%	29%	6%	29%	6%
Total	100% (351)	101% (175)	100% (2,923)	101% (867)	101% (3,274)	100% (1,042)
B. Czech Republic	"Red" cohorts		Other cohorts		All cohorts	
	Direct entry	Delayed entry	Direct entry	Delayed entry	Direct entry	Delayed entry
None	21%	66%	15%	72%	17%	71%
Vocational secondary	52%	22%	46%	18%	47%	19%
Professional secondary	16%	8%	24%	8%	23%	8%
Complete secondary	10%	4%	15%	2%	14%	2%
Total	99% (971)	100% (206)	100% (3,368)	100% (514)	101% (4,339)	100% (720)
C. Hungary	"Red" cohorts		Other cohorts		All cohorts	
	Direct entry	Delayed entry	Direct entry	Delayed entry	Direct entry	Delayed entry
None	62%	72%	30%	65%	32%	66%
Vocational secondary	12%	20%	34%	24%	32%	23%
Professional secondary	13%	6%	18%	5%	17%	5%
Complete secondary	13%	2%	18%	6%	18%	6%
Total	100% (262)	100% (162)	100% (3,172)	100% (950)	99% (3,434)	100% (1,112)

Continued on next page.

Table 3- continued

D. Poland	"Red" cohorts		Other cohorts		All cohorts	
	Direct entry	Delayed entry	Direct entry	Delayed entry	Direct entry	Delayed entry
None	53%	73%	22%	76%	25%	75%
Vocational secondary	18%	14%	43%	14%	40%	14%
Professional secondary	16%	9%	17%	6%	17%	6%
Complete secondary	14%	4%	18%	5%	18%	5%
Total	101% (307)	100% (162)	100% (2,741)	101% (612)	100% (3,048)	100% (774)

E. Slovakia	"Red" cohorts		Other cohorts		All cohorts	
	Direct entry	Delayed entry	Direct entry	Delayed entry	Direct entry	Delayed entry
None	46%	71%	17%	72%	19%	72%
Vocational secondary	27%	19%	42%	19%	41%	19%
Professional secondary	15%	7%	27%	7%	26%	7%
Complete secondary	12%	2%	14%	2%	14%	2%
Total	100% (368)	99% (168)	100% (3,472)	100% (575)	100% (3,840)	100% (743)

Table 4: Percentage distribution of university enrollment by enrollment timing and cohort of secondary school graduation, Central and Eastern Europe, 1948- 1989. Number of cases at risk for each cohort/enrollment timing combination in parentheses. Total N= 8,989.

	"Red" cohorts		Other cohorts		All cohorts	
	Direct entry	Delayed entry	Direct entry	Delayed entry	Direct entry	Delayed entry
All countries	19% (737)	19% (596)	17% (8,252)	16% (6,857)	17% (8,989)	16% (7,453)
Bulgaria	15% (144)	22% (122)	13% (2,330)	19% (2,024)	13% (2,474)	19% (2,146)
Czech Republic	22% (328)	18% (257)	23% (1,666)	10% (1,286)	23% (1,994)	11% (1,543)
Hungary	19% (63)	20% (51)	13% (1,353)	25% (1,183)	13% (1,416)	25% (1,234)
Poland	13% (93)	25% (81)	13% (1,143)	16% (991)	13% (1,236)	17% (1,072)
Slovakia	22% (109)	11% (85)	22% (1,760)	9% (1,373)	22% (1,869)	9% (1,458)

Table 5: Goodness-of-fit statistics of selected multinomial logistic regression models of entry into secondary schools, Central and Eastern Europe, 1948- 1989. Total N= 17,935.

Model	L^2	d.f.	p-value	BIC
Additive effect only:				
M1: additive effects of all covariates[10]	11771.4	51	0.000	-11260.7
Two-way interactions of background and delayed entry:				
*M2: M1 + FEd * Late + MEd * Late + ISEI * Late*	*11838.1*	*60*	*0.000*	*-11237.3*
M3: M1 + FEd * Late	11822.9	54	0.000	-11282.2
M4: M1 + MEd * Late	11813.4	54	0.000	-11272.6
M5: M1 + ISEI * Late	11784.4	54	0.000	-11243.5
Contrasts				
M2-M1	66.7	9	0.000	23.4
M3-M1	51.5	3	0.000	-21.5
M4-M1	42.9	3	0.000	-11.9
M5-M1	12.9	3	0.005	17.2

[10] Male, father's education (FEd), mother's education (MEd), main breadwinner's ISEI (ISEI), number of siblings, parents' CP membership, "red" cohort (Red), plus identificators of mean-replaced missing values, four country dummies (Bulgaria, Czech Republic, Hungary, Poland; Slovakia is the comparison category), dummy for early vs. late entry (Late). Indicated abbreviations used throughout the paper.

Table 5 continued: Goodness-of-fit statistics of selected multinomial logistic regression models of entry into secondary schools, Central and Eastern Europe, 1948-1989. Total N= 17,935.

Model	L^2	d.f.	p-value	BIC
Combine models with two-way interactions:				
M6: M2 + Fed * Late + MEd * Late + ISEI * Late +Red * Late	11887.2	72	0.000	-11166.2
M7: M3 + FEd * Late +Red * Late	11866.6	60	0.000	-11265.8
M8: M4 + MEd * Late +Red * Late	11858.9	60	0.000	-11258.1
M9: M5 + ISEI * Late + Red * Late	11831.5	60	0.000	-11230.7
Three-way interactions:				
M10: M6 + (FEd * Red * Late) + (MEd * Red * Late) + (ISEI * Red + Late)	11903.5	81	0.000	-11092.4
M11: M7 + (FEd * Red * Late)	11870.9	63	0.000	-11240.1
M12: M8 + (MEd * Red * Late)	11871.7	63	0.000	-11240.8
M13: M9 + (ISEI * Red + Late)	11833.2	63	0.000	-11202.3
Contrasts				
M10-M6	16.3	9	0.061	73.8
M11-M7	4.3	3	0.234	25.7
M12-M8	12.7	3	0.005	17.3
M13-M9	1.7	3	0.641	28.4

Table 6: estimated coefficients and standard errors (in parentheses) of selected multinomial logistic regression of secondary school entry among primary school graduates, Central and Eastern Europe, 1948- 1989. Total N = 17,935.

	Model 1			Model 2		
	Vocational	Professional	Academic	Vocational	Professional	Academic
Male	0.857 (0.038)	0.214 (0.044)	-0.406 (0.050)	0.858 (0.038)	0.216 (0.044)	-0.404 (0.050)
Father's education	0.087 (0.009)	0.126 (0.010)	0.156 (0.011)	0.101 (0.010)	0.141 (0.011)	0.175 (0.012)
Mother's education	0.067 (0.009)	0.153 (0.010)	0.175 (0.011)	0.082 (0.010)	0.168 (0.011)	0.189 (0.012)
Occupational status	-0.001 (0.002)	0.021 (0.002)	0.038 (0.002)	-0.002 (0.002)	0.020 (0.002)	0.038 (0.002)
# of siblings	-0.125 (0.016)	-0.323 (0.019)	-0.326 (0.021)	-0.123 (0.016)	-0.322 (0.019)	-0.326 (0.021)
"Red" cohorts	-0.492 (0.054)	-0.557 (0.067)	-0.186 (0.074)	-0.485 (0.054)	-0.550 (0.067)	-0.178 (0.074)
Parents CP	0.106 (0.053)	0.290 (0.056)	0.277 (0.061)	0.105 (0.053)	0.291 (0.057)	0.279 (0.061)
"Delayed" entry	-1.514 (0.049)	-1.926 (0.067)	-2.143 (0.084)	-1.631 (0.057)	-2.040 (0.069)	-2.191 (0.084)
Interactions:						
FEd * Late				-0.051 (0.022)	-0.062 (0.029)	-0.125 (0.035)
MEd * Late				-0.066 (0.023)	-0.069 (0.028)	-0.044 (0.033)
ISEI * Late				0.009 (0.005)	0.001 (0.006)	-0.003 (0.007)
Constant	0.463 (0.048)	0.344 (0.054)	-0.202 (0.064)	0.487 (0.049)	0.364 (0.054)	-0.194 (0.064)

Note: coefficients associated with dummy variables identifying missing values replaced with means as well as country dummies - a total of 9 effects in each equation- are not shown in any model, but will be made available upon request from the author. All numbers are rounded to three decimal places.

Table 7: estimated coefficients and standard errors (in parentheses) of selected multinomial logistic regression of secondary school entry among primary school graduates, Central and Eastern Europe, 1948- 1989. Total N = 17,935.

	Model 6			Model 10		
	Vocational	Professional	Academic	Vocational	Professional	Academic
Male	0.866 (0.038)	0.224 (0.044)	-0,398 (0,050)	0,867 (0,038)	0,225 (0,044)	-0,397 (0,050)
Father's education	0.100 (0.011)	0.141 (0.011)	0,174 (0,013)	0,100 (0,011)	0,139 (0,012)	0,174 (0,013)
Mother's education	0.087 (0.011)	0.175 (0.012)	0,197 (0,012)	0,092 (0,011)	0,181 (0,012)	0,202 (0,013)
Occupational status	-0.003 (0.002)	0.020 (0.003)	0,038 (0,003)	-0,003 (0,003)	0,020 (0,003)	0,038 (0,003)
# of siblings	-0.122 (0.016)	-0.321 (0.019)	-0,325 (0,021)	-0,121 (0,016)	-0,321 (0,019)	-0,325 (0,021)
"Red" cohorts	-0.674 (0.070)	-0.793 (0.080)	-0,326 (0,085)	-0,730 (0,073)	-0,849 (0,082)	-0,370 (0,086)
Parents CP	0.104 (0.053)	0.288 (0.057)	0,275 (0,061)	0,104 (0,053)	0,289 (0,057)	0,276 (0,061)
"Delayed" entry	-1.728 (0.061)	-2.160 (0.075)	-2,226 (0,092)	-1,755 (0,063)	-2,176 (0,075)	-2,232 (0,092)
2-way interactions:						
FEd * Late	-0.050 (0.022)	-0.063 (0.029)	-0,125 (0,035)	-0,055 (0,024)	-0,042 (0,032)	-0,121 (0,038)
MEd * Late	-0.060 (0.023)	-0.057 (0.028)	-0,041 (0,033)	-0,082 (0,026)	-0,091 (0,031)	-0,064 (0,036)
ISEI * Late	0.010 (0.005)	0.001 (0.006)	-0,002 (0,007)	0,011 (0,006)	0,001 (0,007)	-0,004 (0,008)
FEd * Red	-0.007 (0.024)	-0.011 (0.029)	-0,004 (0,031)	-0,013 (0,027)	0,003 (0,032)	-0,002 (0,033)
MEd * Red	-0.014 (0.026)	-0.055 (0.030)	-0,052 (0,032)	-0,052 (0,030)	-0,101 (0,033)	-0,089 (0,034)
ISEI * Red	0.003 (0.005)	0.000 (0.006)	-0,008 (0,006)	0,004 (0,006)	0,000 (0,007)	-0,008 (0,007)
Late * Red	0.614 (0.123)	0.801 (0.173)	0,200 (0,225)	0,839 (0,156)	0,992 (0,193)	0,382 (0,238)
3-way interactions:						
FEd * Late * Red				0,036 (0,057)	-0,113 (0,079)	-0,023 (0,101)
MEd * Late * Red				0,132 (0,062)	0,215 (0,082)	0,167 (0,102)
ISEI * Late * Red				-0,009 (0,013)	0,004 (0,016)	0,014 (0,020)
Constant	0.511 (0.049)	0.387 (0.054)	-0,182 (0,064)	0,514 (0,049)	0,391 (0,055)	-0,179 (0,064)

Note: coefficients associated with dummy variables identifying missing values replaced with means as well as country dummies - a total of 9 effects in each equation- are not shown in any model, but will be made available upon request from the author. All numbers are rounded to three decimal places.

Table 8: estimated net slopes for the effect of mother's education on the log odds of the transition into vocational, professional, and academic secondary schools based on a multinomial logistic regression model of secondary school entry among primary school graduates, Central and Eastern Europe, 1948- 1989. Total N = 17,935.

	Directly entry		Delayed entry	
	Normal cohorts	Red cohorts	Normal cohorts	Red cohorts
Vocational schools	0.092	0.04	0.01	0.09
Professional secondary schools	0.181	0.08	0.09	0.204
Academic secondary schools	0.202	0.113	0.138	0.216

Note: estimates based on model 10. No secondary education is the comparison category.

Table 9: Goodness-of-fit statistics of selected binomial logistic regression models of entry into tertiary schools, Central and Eastern Europe, 1948- 1989. Total N= 8,989.

Model	L^2	d.f.	p-value	BIC
Additive effect only:				
M14: additive effects of all covariates[11]	931.1	17	0.000	-766.1
Two-way interactions of background and delayed entry:				
M15: M1 + FEd * Late + MEd * Late + ISEI * Late	941.0	20	0.000	-746.8
M16: M1 + FEd * Late	935.2	18	0.000	-760.4
M17: M1 + MEd * Late	937.0	18	0.000	-762.2
M18: M1 + ISEI * Late	939.2	18	0.000	-764.4
Contrasts				
M15-M14	9.8	3	0.020	19.7
M16-M14	4.0	1	0.045	5.7
M17-M14	5.8	1	0.016	3.9
M18-M14	8.0	1	0.005	1.7

[11] Male, father's education (FEd), mother's education (MEd), main breadwinner's ISEI (ISEI), number of siblings, parents' CP membership, "red" cohort (Red), plus identificators of mean-replaced missing values, four country dummies (Bulgaria, Czech Republic, Hungary, Poland; Slovakia is the comparison category), dummy for early vs. late entry (Late).

Table 9 continued: Goodness-of-fit statistics of selected binomial logistic regression models of entry into tertiary schools, Central and Eastern Europe, 1948- 1989. Total N= 8,989.

Model	L^2	d.f.	p-value	BIC
Combine models with two-way interactions:				
M19: M15 + FEd * Late + MEd * Late + ISEI * Late + Red * Late	948.6	24	0.000	-715.6
M20: M16 + FEd * Late + Red * Late	937.6	20	0.000	-743.4
M21: M17 + MEd * Late + Red * Late	943.3	20	0.000	-749.2
M22: M18 + ISEI * Late + Red * Late	943.1	20	0.000	-748.9
Three-way interactions:				
M23: M19 + (FEd * Red * Late) + (MEd * Red * Late) + (ISEI * Red + Late)	953.0	27	0.000	-690.9
M24: M20 + (FEd * Red * Late)	940.0	21	0.000	-736.1
M25: M21 + (MEd * Red * Late)	943.8	21	0.000	-734.0
M26: M2 + (ISEI * Red + Late)	943.1	21	0.000	-739.3
Contrasts				
M23-M19	4.5	3	0.213	24.7
M24-M20	2.4	1	0.121	7.3
M25-M21	0.5	1	0.485	15.2
M26-M22	0.0	1	0.836	9.6

Table 10: estimated coefficients and standard errors (in parentheses) of selected binomial logistic regression of tertiary school entry among secondary school graduates, Central and Eastern Europe, 1948- 1989. Total N = 8,989.

	Model 14	Model 15
Male	0.356 (0.043)	0.358 (0.043)
Father's education	0.074 (0.009)	0.072 (0.012)
Mother's education	0.063 (0.009)	0.073 (0.012)
Occupational status	0.016 (0.002)	0.018 (0.002)
Number of siblings	-0.092 (0.021)	-0.092 (0.021)
"Red" cohorts	0.362 (0.078)	0.362 (0.078)
Parents CP members	0.101 (0.048)	0.103 (0.048)
"Delayed" entry	0.024 (0.044)	0.056 (0.045)
Interactions:		
FEd * Late		0.004 (0.019)
MEd * Late		-0.023 (0.018)
ISEI * Late		-0.006 (0.003)
FEd * Red		
MEd * Red		
ISEI * Red		
Constant	-1.918 (0.059)	-1.939 (0.060)

Table 11: estimated coefficients and standard errors (in parentheses) of selected binomial logistic regression of tertiary school entry among secondary school graduates, Central and Eastern Europe, 1948- 1989. Total N = 8,989.

	Model 19	Model 23
Male	0.355 (0.043)	0.356 (0.043)
Father's education	0.071 (0.013)	0.066 (0.013)
Mother's education	0.078 (0.012)	0.078 (0.013)
Occupational status	0.019 (0.002)	0.019 (0.002)
Number of siblings	-0.094 (0.021)	-0.094 (0.021)
"Red" cohorts	0.336 (0.108)	0.346 (0.110)
Parents CP members	0.099 (0.048)	0.099 (0.048)
"Delayed" entry	0.061 (0.048)	0.057 (0.048)
2-way interactions:		
FEd * Late	0.004 (0.019)	0.015 (0.020)
MEd * Late	-0.023 (0.018)	-0.024 (0.019)
ISEI * Late	-0.006 (0.003)	-0.008 (0.003)
FEd * Red	0.010 (0.031)	0.061 (0.041)
MEd * Red	-0.060 (0.033)	-0.060 (0.043)
ISEI * Red	-0.007 (0.006)	-0.014 (0.008)
Late * Red	-0.046 (0.154)	-0.096 (0.167)
3-way interactions:		
FEd * Late * Red		-0.121 (0.063)
MEd * Late * Red		0.002 (0.065)
ISEI * Late * Red		0.016 (0.012)
Constant	-1.946 (0.060)	-1.944 (0.060)

Table 12: Goodness-of-fit statistics of selected discrete choice/conditional logistic regression models of entry into secondary schools, Central and Eastern Europe, 1948-1989. Total N= 17,935.

Model	L^2	d.f.	p-value	BIC
Additive effect only:				
M27: type of secondary school only	1183.3	2	0.000	-1160.5
M28: M27 + Late	13183.7	3	0.000	-13149.6
Two-way interactions of background and delayed entry:				
M29: M28 + FEd * Late + MEd * Late + ISEI * Late	13474.9	6	0.000	-13406.5
M30: M28 + FEd * Late	13414.47	4	0.000	-13368.9
M31: M28 + MEd * Late	13416.8	4	0.000	-13371.2
M32: M28 + ISEI * Late	13282.0	4	0.000	-13236.4
All two-way interactions:				
M33: M29 + Late * Red	13517.5	7	0.000	-13437.7
Add three-way interactions				
M34: M33 + (FEd * Red * Late) + (MEd * Red * Late) + (ISEI * Red + Late)	13528.1	10	0.000	-13414.2
Contrasts between models	**L2**	**d.f.**	**p-value**	**BIC**
M29-M28	291.1	3	0.000	-256.9
M30-M28	230.7	1	0.000	-219.3
M31-M28	233.0	1	0.000	-221.6
M32-M28	98.2	1	0.000	-86.8
M34-M33	10.7	3	0.014	23.5

Table 13: estimated coefficients and standard errors (in parentheses) of selected discrete choice/conditional logistic regression of secondary school entry among primary school graduates, Central and Eastern Europe, 1948- 1989. Total N = 8,989.

	Model 28	Model 29	Model 34	Model 35
Characteristics of options				
Type of secondary school (vocational omitted)				
Professional secondary	-0.406 (0.019)	-0.406 (0.019)	-0.406 (0.019)	-0.406 (0.019)
Academic secondary	-0.686 (0.021)	-0.686 (0.021)	-0.686 (0.021)	-0.686 (0.021)
"Late" entry	-2.396 (0.030)	-2.423 (0.031)	-2.505 (0.034)	-2.505 (0.034)
Cross-level interactions:				
Father's education * Late		-0.086 (0.014)	-0.083 (0.015)	-0.155 (0.011)
Mother's education * Late		-0.103 (0.014)	-0.111 (0.015)	
ISEI * Late		-0.005 (0.003)	-0.004 (0.003)	
Late * Red			0.650 (0.086)	0.659 (0.083)
Father's education * Late * Red			-0.004 (0.035)	0.053 (0.028)
Mother's education * Late * Red			0.110 (0.038)	
ISEI * Late * Red			-0.004 (0.007)	

Table 14: Goodness-of-fit statistics of selected discrete choice/conditional logistic regression models of entry into tertiary schools, Central and Eastern Europe, 1948-1989. Total N= 8,989.

Model	L^2	d.f.	p-value	BIC
Additive effect only:				
M36: Dummy for "Late" transitions only	39.30	1	0.000	-30.7
Two-way interactions of background and delayed entry:				
M37: M36 + FEd * Late + MEd * Late + ISEI * Late	79.4	4	0.000	-44.9
M38: M37 + Late * Red	80.0	5	0.000	-36.9
Add three-way interactions				
M39: M38 + (FEd * Red * Late) + (MEd * Red * Late) + (ISEI * Red + Late)	83.7	8	0.000	-14.9
Contrasts between models	**L2**	**d.f.**	**p-value**	**BIC**
M37-M36	40.1	3	0.000	-14.2
M39-M38	3.8	3	0.286	22

Table 15: estimated coefficients and standard errors (in parentheses) of selected discrete choice/conditional logistic regression models of tertiary school entry among secondary school graduates, Central and Eastern Europe, 1948- 1989. Total N = 8,989.

	Model 36	Model 37	Model 39
Characteristics of options			
"Late" entry	-0.240 (0.038)	-0.153 (0.041)	-0.144 (0.044)
Cross-level interactions:			
Father's education * Late		-0.010 (0.015)	-0.001 (0.016)
Mother's education * Late		-0.026 (0.015)	-0.029 (0.016)
ISEI * Late		-0.009 (0.003)	-0.011 (0.003)
Late * Red			-0.132 (0.146)
Father's education * Late * Red			-0.092 (0.053)
Mother's education * Late * Red			0.009 (0.055)
ISEI * Late * Red			0.017 (0.011)