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EDUCATIONAL ACHIEVEMENT IN COMPARATIVE PERSPECTIVE

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ABSTRACT

We carry out a cross-national and cross-temporal analysis to assess how societal factors affect the dependence of educational attainment on parental status and the gender gap in educational attainment. Combining data from 541 sample surveys conducted in 54 nations, we estimate a micro-level model of the determinants of years of school completed in each of the 654 "contexts" created by crossing five-year schooling cohorts by nation and then carry out a macro-level analysis of the determinants of variations in the micro-level coefficients across contexts. We develop various hypotheses regarding the effect of modernization, educational expansion, educational inequality, and communism on the micro-level coefficients. Our hypotheses are generally confirmed.

INTRODUCTION

It is by now a commonplace that education is the primary locus of job training in the modern world and, as a consequence, the primary engine of social mobility. Most people prepare for their life's work by going to school, and those who go furthest in school obtain the best jobs—those with the greatest prestige and the highest earnings. Success in school is thus the best way to overcome the limitations of one's social origins; and failure in school among the children of the advantaged is a fairly sure route to downward mobility. This has not always been so. Until quite recently in human history, most men followed their fathers into the fields and shops, learning their trade through direct apprenticeship; and most women did the same, learning the skills of household management and agricultural production step-by-step under their mothers' tutelage.

Although schools as separate social institutions are an ancient invention, universal public education is very new—at most a century and a half old and in most nations of the world far more recent. But as work shifted out of the fields and into factories, formal organizations, and bureaucracies, and as processes of production and management became more complex, the efficiency of organized schooling came to be widely understood and valued (Harbison and Myers 1964) . As a consequence, education has expanded in virtually every nation on the globe throughout the 20th century, albeit more rapidly and more continuously in some societies than in others. At the turn of the century it was uncommon for as much as 10 per cent of the male population, and an even smaller percentage of the female population, to have any secondary schooling, and large fractions had no schooling at all. By the end of the century, primary schooling had become virtually universal and a majority of children have at least some secondary education in all industrialized and many non-industrialized nations.

The claim has been advanced (e.g., Treiman 1970) that as the availability of education has expanded, equality of educational opportunity has concomitantly increased so that over time the educational system has come to play a central role in promoting social mobility, both upward and downward—providing opportunities to the bright and hardworking children of those of humble origins but at the same time preventing the lazy, dull, or troubled children of advantage from enjoying the fruits of their parents' success.[1] In general, the hypothesized trend towards more equality of educational opportunity has been attributed to changes exogenous to the educational attainment process, such as the rise of universalistic values, efficiency, public funding, and urbanization, all of which would lead to a greater influx of children of disadvantaged backgrounds into each level of schooling and consequently to a reduction of social selectivity.[2]

The empirical evidence in support of these claims is mixed. Most studies of educational expansion or, more generally, trends in educational achievement over time have been limited to single countries, although there now is an accumulation of semi-comparable studies for a fairly large number of nations, mainly though the efforts of Shavit and Blossfeld (1993) who organized parallel analyses in 13 countries (see Blossfeld and Shavit [1993, pp. 4-5] for a partial review of such

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[1] Others have made the same general claim under various labels, such as the rise of meritocracy (Young 1963; Halsey, Heath, and Ridge 1980), modernization (Inkeles and Smith 1974), and the decline of ascription (Blau and Duncan 1967; Parsons 1970), each of which points to somewhat different mechanisms.

[2] There is, to be sure, a competing claim. Some (e.g., Collins 1971; Bowles and Gintis 1976; Bourdieu and Passeron 1977) see education, particularly higher education, as primarily a vehicle for social reproduction. While it is acknowledged that equality of educational opportunity may be increasing at low levels of education, access to universities and other elite institutions is seen as monopolized by the children of the rich and powerful. Thus, the claim is that the effect of social origins on educational attainment is greatest at the high end of the educational distribution. From this, it follows, *ceteris paribus*, that as education expands the dependence of education on social origins should increase since a higher fraction of all children and young adults will be attempting transitions into educational institutions at the high end of the educational system where social origins matter more. Assessment of variations in the effect of social origins across different transitions is beyond the scope of this paper and is presented separately.

studies; see also a paper by Müller and his colleagues [1990] comparing educational systems in nine European nations and a paper by Wong [2003] comparing educational attainment in five Eastern European nations during the communist period). In a summary of the 13 studies they organized, Blossfeld and Shavit (1993, pp. 13-16) report that the average level of education attained increased over time in all 13 countries (their Table 1.1) but that changes in the effect of social origins were quite varied (their Table 1.2): the effect of father's education declined in five nations, remained essentially unchanged in seven nations, and first declined and then increased in Czechoslovakia. The effect of father's occupational status was even more varied—declining in three nations, remaining essentially unchanged in nine nations, and increasing in one nation, Italy. They also report that the gender gap declined over time in all 10 nations where this was studied.

Ganzeboom and Treiman (1993) analyzed data for men from 29 nations, pooling all data sets available for each nation and using fully standardized measures of occupational status and education, and found results generally similar to those reported by Blossfeld and Shavit: education expanded over time in all countries; for the 26 nations for which both father's education and father's occupational status were available, the effect of father's education declined significantly over time in nine nations, remained essentially unchanged in 15 nations, and significantly increased in two nations, while the effect of father's occupational status declined significantly in five nations, remained essentially unchanged in 20 nations, and increased in one nation. However, they went substantially beyond previous analysis by showing, via least squares dummy variable and pooled models, that educational expansion reduced the effect of social origins on educational attainment while educational inequality increased the effect of social origins.

Treiman, Ganzeboom, and Rijken (2003) utilized a different analytic procedure, hierarchical linear modeling, and data for a wider range of societies (31) and expanded the analysis in various ways: by including women as well as men and analyzing how societal modernization, a world-wide secular trend toward greater equality of opportunity, and communist educational policies, as well as educational expansion and educational inequality, affect the dependence of educational on parental status and the gender gap in educational attainment. They showed that modernization, educational expansion, and the secular trend toward greater equality of opportunity reduced the dependence of educational attainment on parental status and the gender gap in educational attainment, while educational inequality increased both. Both the gender gap and the dependence of education on social origins were reduced at the outset of communist regimes but the effect of communism declined over time. The present paper further expands their analysis, reconsidering the same issues but with the inclusion of data from more nations (54), consideration of a richer set of hypotheses with a larger set of macro-variables, and improved estimation procedures.

Note that the present paper is concerned with explaining societal and temporal variation in the effect of social origins and gender on the *final level of education* attained by individuals, without regard to the process of educational attainment as measured by influences on *successive transitions*. In 1980 Mare introduced a model of the educational career as a sequence of educational transitions (Mare 1980, 1981). The popularity of this model has led to some confusion among analysts about the value of the metric regression model as a way of studying educational attainment, and some have no longer bothered to present results from the metric regression model (e.g. Shavit and Kraus 1990). We strongly disagree with this practice. While Mare's transition model yields important insights about the mechanics of educational opportunities, this should not obscure the fact that the parameters of the metric regression model (or some other measure of the over-all association of educational attainment with social origins), as well as changes in these parameters over time, are the fundamental explanandum of educational mobility research, because these are what directly measure the degree of inequality of educational opportunities in a society.

HYPOTHESES REGARDING EDUCATIONAL ATTAINMENT

In the present paper our primary interest is in how the factors affecting final educational attainment vary across different social contexts, where contexts are defined by national populations age 10-14 at specific points in time. We first propose a simple model of the factors affecting status attainment in all contexts (that is, a model that we claim holds for all societies throughout at least the past century); this is our micro model. We then consider how variations in social structure across societies and over time might be expected to modify the expectations of our micro model; this is our macro model.

Micro-level Hypotheses

Here our expectations are straightforward and unproblematic: those from more advantaged social origins (measured by father's education[3] and father's occupational status) should obtain more schooling; and men should obtain more schooling than women.

There are two reasons for expecting a positive effect of social origins on educational attainment. The most important is the role of family cultural capital as an intervening mechanism.

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[3] We also would expect mother's education to affect educational attainment in a similar way. However, many of our data sets lack information on mother's education. Thus we include only father's education in our micro model.

Well-educated parents provide home environments that generate the skills, expectations, and motivations that lead children to do well in school and to wish to continue in school (Bourdieu and Passeron 1977; Evans et al. 2010).[4] The same is true of parents with high status jobs, since such jobs are cognitively demanding (Kohn and Schooler 1983). Second, high status jobs tend to pay well, and hence provide the material resources that make it possible to pay school fees and to forgo the need for children to leave school in order to help support themselves or their families. In sum, those from high status families tend to enjoy superior cultural and material resources that both encourage and facilitate extended education.

Our expectation that men will obtain more education than will women of comparable social origin status is based on the assumption that families act rationally in light of the universal propensity for men to get higher returns to education than women (e.g., Treiman and Roos 1983). If a choice needs to be made between educating one's sons or one's daughters, the rational choice—from the point of view of providing for one's own social security—is to give preference to the education of sons. This suggests that in societies with guaranteed retirement benefits (whether from employers or from public sources), the difference in the educational attainment of men and women should be reduced.[5] In addition, when the earnings gap between men and women narrows,

[4] Desmond (2007, pp. 169-172) illustrates in a vivid way the role family and childhood socialization plays in creating competence that predisposes individuals to success at specific endeavors. In his case, it is fighting forest fires, which is not so much learned on the job or via specific training as by growing up in a "country-masculine habitus" that creates familiarity with and competence at the kinds of actions demanded by the job. In the same way, success at school depends heavily upon predispositions and skills acquired prior to ever setting foot in a classroom.

[5] We attempted to code contexts with respect to the availability of publicly funded retirement benefits from information provided by the U.S. Social Security Administration's description of welfare programs throughout the world (Social Security Administration 2002-2003) but ultimately abandoned the attempt because the available information generally lacked sufficient detail.

as it has in the U.S. (Morris and Western 1999), the gender gap in educational attainment also should narrow.

Macro-level Hypotheses

We next consider whether and in what ways the process of status attainment might be expected to vary in different social environments, that is, whether there is reason to expect systematic variation across contexts in the magnitude of the four coefficients of the micro-model: those associated with father's occupational status, father's level of schooling, gender, and the intercept. We consider the effects of six contextual variables (educational expansion, educational inequality, the level of societal development or modernization, communism, cohort [the years defining each context], and the interaction between communism and cohort).

Educational expansion refers to the proportion of the population of a society with access to schooling at each point in time. **Educational inequality** refers to the variability in educational attainment at each point in time. **Development** refers to the level of societal development or modernization, which is strongly correlated with but extends beyond economic development. **Communism** is a dichotomous variable distinguishing communist from non-communist regimes. **Cohort** defines the temporal context, and is used to model secular trends over the course of the past century. Precise operational definitions of these variables are given below.

Hypotheses regarding the intercept of the micro equation.—Since, as we define it (see below), the intercept indicates the expected level of education for men with within-context average social origins, it reflects whatever factors affect the average level of education. With our data, we can assess the effect of four such factors: the level of societal development, the level of access to education, a propensity for educational attainment to increase over time, and a propensity for education levels to be higher at the beginning of communist regimes than would be expected from their level of modernization. We consider an additional factor, the level of educational inequality, but have no specific hypotheses about its effect on the average level of education attained.

The increase in the average level of educational attainment as societies develop is well known and the explanation is straightforward: outside of agriculture and small shops children generally cannot work along side their parents, learning vocational skills as they go. Thus, they are sent to school and schools become the primary locus of vocational training. In addition, as nations industrialize the cost of education declines because the state assumes greater responsibility, establishing public schools and subsidizing or completely covering the costs of schooling (Lindert 2004).

What is not as obvious is the possibility that there is a secular increase in average levels of education independent of shifts in the distribution of the labor force or the level of development. However, we think there are good grounds for expecting this to be the case. In virtually all societies, there is great pressure to expand the education system because parents believe (correctly) that education is the primary route to success for their children (see, e.g., Abernethy 1969).[6] Where schools exist at all, the benefits of schooling are visible for all to see: the more schooling people attain, the better their life chances. Since governments tend to be responsive to such public pressure, over time the availability of schooling and the level of educational attainment increase faster than

[6] As Schofer and Meyer (2005, p. 900) point out, "... as education becomes important in the attainment of social status, groups and individuals compete more intensively for success in education, producing inflationary credential expansion far beyond any original functional requirements (Collins 1971, 1979; Bourdieu and Passeron 1977)."

would be expected simply from the shift in the character of the labor force[7] or the level of development.

The 40 year experiment with communism in Eastern Europe, and the still lengthier regimes in the Soviet Union and China, produced a social system distinctive in many respects, among the most important of which was the imposition of centralized control on many aspects of life. Schools are instruments of social control *par excellence*, serving both socializing and gate-keeping functions. For this reason, we would expect communist regimes to have a special interest in ensuring that all of its citizens are exposed to at least a minimum level of standardized schooling, and thus posit a positive effect of communism on educational expansion net of other factors. (See Rijken [1997] for a similar argument.) However, we expect the "communist bonus" on educational attainment to diminish over time as communist regimes mature and ideological fervor wanes.[8]

Hypotheses regarding the effect of social origins.—We consider father's education and father's occupation together since we expect the effects of contextual factors to be similar for the two micro variables. Our main hypotheses are about how educational expansion, educational inequality, and development mediate the effect of social origins on educational attainment. We also consider competing hypotheses regarding the effect of communism.

The main effect of educational expansion should be to reduce the role of social origins on educational attainment. The argument for this expectation is as follows. The impact of social

[7] Of course, educational expansion itself is a force promoting a shift away from agriculture. Those who obtain schooling are generally reluctant to return to their farms and villages and tend to flood into cities to seek non-agricultural work.

[8] There is some evidence that educational expansion in communist nations occurred mainly at the lower end and that, especially after 1970, communist regimes resisted the expansion of higher education in order to retain party control (Schofer and Meyer 2005; see also Lenhardt and Stock 2000; and Baker, Köhler, and Stock 2004).

origins on the likelihood of moving from one level of education to the next tends to decline for successive transitions. Mare (1980, pp. 298-299) has suggested that this necessarily follows from differential selectivity (or, the same thing, differential attrition) at each transition. Since family cultural capital affects not only measured traits such as school performance but also generally unmeasured traits such as motivations, expectations, and abilities, at each step in the educational career children from lower status origins will tend to be more highly selected with respect to both measured and unmeasured traits than will children from higher status origins. When transition rates are high, only the very culturally deprived (who tend to be disproportionately from low status origins) fail to continue; when transition rates are low, only the most highly motivated and most talented students from low status origins continue, compared to higher proportions of those from high status origins for whom continued schooling is a normative expectation. The consequence is that the correlations between social origins and (typically unmeasured) intervening variables affecting educational success tend to decline for successive transitions, and thus the reduced-form effects of social origins on the odds of making successive transitions also decline. It then follows that as more students move into the ranges of education for which the effects of social origins are relatively small, the overall effect of social origins on educational attainment will decline,[9] all else equal (that is, specifically, if the variance in years of schooling and the effect of social origins on the odds of making each transition do not change).[10] The relationship between educational expansion

[9] Mare (1981, pp. 77-78) has shown this formally, by demonstrating that the linear effects of social origins on the highest level of education completed are weighted sums of the logistic regression parameters for the effects of social origins on the odds of making successive educational transitions, where the weights are functions of the proportion of persons making each transition. Clearly, if the logistic regression parameters for advanced transitions are small and if the proportions making such transitions are high, the linear effects of social origins will be small.

[10] Note that it is only if the *ceteris paribus* assumption holds—that is, only if the variance in years of schooling and the effect of social origins on the odds of making each transition do not change—that the relationship between

and the dependence of final educational attainment on social origins (inequality of educational opportunity, IEO) then turns on the relative strength of two competing mechanisms—the increase in selectivity over successive transitions with respect to unmeasured traits[11] and the reduction in overall selectivity as education expands—leaving as an empirical question how these two forces balance out.

An additional argument for why the effect of social origins declines with educational expansion has been advanced by Blossfeld and Shavit (1993, pp. 9-10): as more children stay in school, the crucial schooling decisions occur at successively higher ages, when students are less dependent upon the economic circumstances or preferences of their families, and become more able and willing to make their own decisions, thus reducing the dependence of educational transitions on social origins. In addition, parents are less likely to be able to help their children prepare for entrance examinations at higher levels of schooling.

All else equal, we expect educational inequality to *increase* the dependency of educational attainment on social origins. The argument follows that advanced by Treiman and Yip (1989, pp. 376-77; see also Tyree, Semyonov, and Hodge [1979]; and Kelley and Klein [1981, pp. 18-19]), who suggest that educational inequality may be taken as a proxy for social inequality more generally. In societies where social inequality is large, differences in both measured and

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educational expansion and the dependence of total years of schooling on social origins becomes a statistical tautology. But whether the *ceteris paribus* assumption holds is an empirical question and thus worth exploring. Also, we avoid any possibility of a statistical tautology by using a measure of educational expansion not drawn from our data—see below.

[11] The "Mare model" has been the subject of lively debate as to whether the decline in the effect of social origins over successive transitions should be regarded simply as an artifact of changes in the extent of unmeasured heterogeneity across transitions (see, in particular, Cameron and Heckman 1998) or as substantively meaningful (Treiman and Yamaguchi 1993; Lucas 2001) and to attempts to correct for unobserved heterogeneity (Mare 1993; and the papers in a special 2011 issue of *Research in Social Stratification and Mobility* edited by Buis). The details of this debate are beyond the scope of the present paper.

unmeasured social resources will be large. Thus, the difference in cultural and material capital between those from low and high status origins will tend to be larger in societies with a high degree of social inequality, increasing the impact on the educational achievements of their children. Consider two societies, one in which professionals average 14 years of schooling on average while laborers average four years of schooling and another society in which professionals average 16 years of schooling while laborers average 12 years of schooling. In the second society, the social distance between the children of professionals and laborers is much smaller than in the first society. They will be more likely to attend the same schools and to be exposed to the same social opportunities. Thus, the chances are enhanced that the amount of schooling they attain will be depend on their own talents rather than on their father's status.

We expect development to increase equality of opportunity—that is, to decrease the effect of father's education and father's occupational status. As we have noted earlier, formal schooling gains importance as the distribution of jobs shifts from agriculture and craft production to large scale manufacturing and distribution and to complex bureaucratic management.[12] The state thus increasingly takes responsibility for organizing and paying for education, creating new schools in rural and remote areas and reducing or eliminating school fees (Lindert 2004). The availability of low cost schooling increases opportunities for those of modest means, thus reducing the dependence of schooling on socioeconomic origins (Treiman 1970, p. 221; Treiman and Yip 1989, pp. 375-377).

There are two competing arguments regarding the impact of communism on equality of educational opportunity. One claim (e.g., Simkus and Andorka 1982) is that the net impact of

[12] To be sure, the causality may go the other way—educational expansion may promote economic development (for a thoughtful review of both the causal order and the weight of the evidence see Hannum and Buchmann 2003). But the causal order of the education-development relationship need not concern us here. It is sufficient to note the positive correlation between the two.

educational reproduction was reduced in communist regimes, particularly in their early years, as a result of a kind of "communist affirmative action"—social policies that favored the children of the proletariat and peasantry at the expense of the children of the former bourgeoisie and intelligentsia as well as policies that reduced the cost of education through reductions in fees and the provision of stipends to students to enable them to forgo employment. An additional feature of communist regimes, at least in Central and Eastern European, was the relative material advantage of skilled manual workers relative to routine nonmanual workers and, in consequence, their relatively higher prestige (Treiman 1977, pp. 144-148); this may have partly undercut the net advantage of non-manual origins generally found. Alternatively, new class theorists (e.g., Djilas 1957; Konrád and Szelényi 1979) argue that communist regimes are dominated by well-educated party intelligentsia who act to promote the educational opportunities of their children and thereby increase intergenerational educational reproduction. In a system in which the accumulation of material capital is very difficult and the ability to transmit it to one's children is almost non-existent, cultural capital becomes a necessary condition for social ascendancy. Thus, membership in the communist party is more open to the educated than to the uneducated (Marks 2004)—rhetoric about the worker's paradise notwithstanding—and access to high political positions depends heavily upon educational attainment (Szelényi, Wnuk-Lipinski, and Treiman 1995). Matějů (1993, p. 257; see also Blossfeld and Shavit [1993, p. 9]) suggests that both processes are at work. At the outset of communist regimes, "communist affirmative action" dominates. But then the emerging "new class" becomes entrenched and figures out how to exploit the new policies and bureaucratic procedures to create advantages for their children. This leads to the expectation that in communist nations, the effect of social origins declines at the outset of communist regimes but then increases over time.

Hypotheses regarding the effect of gender (male = 0; female = 1).—Here we consider a set of hypotheses regarding the factors that mediate the gender gap in status attainment. Although we cannot make a firm prediction regarding the intercept of the macro equation since it falls outside the range of our predictor variables, we expect women to obtain less schooling on average than men in non-developed societies in 1900 which, from the way the variable is specified, implies a negative coefficient. This follows from the argument we made above regarding the relative value of educating one's sons and daughters. In traditional societies at the beginning of the century, the value of education for sons should be substantially larger than that for daughters.

As societies modernize, the gender gap in education should be reduced; that is, we expect a positive effect of modernization on the "female" coefficient. There are two reasons for expecting the gender gap in educational attainment to narrow as societies modernize. First, as women increasingly enter the paid labor force their need for schooling increases. Whereas high rates of male labor force participation are universal, women are more likely to engage in paid economic activity outside their families as jobs outside of agriculture and the small-shop economy become more prevalent. Second, societal modernization generally brings an increase in publicly financed welfare, particularly old-age pensions. Hence, elderly parents become less dependent upon their children for their own social security. Under these circumstances, education becomes less of an investment by parents in their own future and more of an investment in the future of their children, which reduces the incentive to devote all of their resources to the improvement of their sons' earning capacities at the expense of their daughters.

For much the same reason that we expect education to expand over time independently of the degree of societal modernization, we also expect the gender gap in educational attainment to narrow. Over the course of the 20^{th} century, the welfare state has come into its own as a major societal type (Esping-Andersen 1990). Much of what used to be regarded as the private concern of individual families is now regarded as the responsibility of the state. While public welfare benefits tend to be more generous in more developed societies, we think it likely that a commitment to public welfare has increased more rapidly than would be expected from increasing societal development.[13] As suggested in the previous paragraph, this should have the consequence of equalizing educational opportunities for men and women.

We make no prediction regarding the effect of educational inequality on the gender gap. It is possible to have strong gender differences in otherwise relatively egalitarian societies and the converse—relative gender equality in otherwise highly stratified societies—is also possible.

Finally, we expect the gender gap in education to be smaller in communist regimes than in non-communist regimes at comparable levels of development. Communist regimes all have strong public policies promoting gender equality. While many such policies are honored in the breach, the chronic labor shortages faced by many post-war Eastern European societies had the consequence of encouraging high levels of female labor force participation, which created strong pressure to promote the schooling of women in order to prepare them for productive work.

Table 1 provides a summary of our expectations. Educational attainment should increase with development, educational expansion, and progression through the 20^{th} century, and should get a boost at the beginning of communist regimes but the communist advantage should decline over time. We have no expectations regarding the effect of educational inequality on the average level of

[13] As noted earlier, we have been unable to identify data adequate to directly test the effect of differing welfare regimes.

educational attainment. Development, educational expansion, progression through the 20th century, and early communism should increase equality of opportunity, manifest in reduced effects of father's education and father's occupational status and the gender gap penalizing women, while educational inequality should reduce equality of opportunity except that we have no clear expectation regarding the effect of educational inequality on the gender gap. The increase in equality of opportunity expected at the outset of communist regimes should fade as these regimes mature.

Having specified how we expect the process of educational attainment to be affected by variations in the social environment within which people made their major educational decisions, we now turn to our empirical analysis.

DATA

The data used in this analysis are from 541 sample surveys conducted in 54 nations[14] throughout the world. These surveys are drawn from the *International Stratification and Mobility File* (*ISMF*) [http://home.fsw.vu.nl/hbg.ganzeboom/ismf/ismf.htm]. Appendix A summarizes the surveys, giving the nation surveyed, the date each survey was conducted (the numerical part of the acronym for the survey), the number of male respondents, the number of female respondents, and the total number of respondents. Additional details regarding each survey are provided on the web

[14] In three cases we analyze separate "nations" within single "states": we divide Belgium into Flemish vs. French-speaking sectors; we divide Canada into Quebec vs. English Canada; and we divide the United Kingdom into England and Wales, Northern Ireland, and Scotland. We also divide Germany into the former East Germany and the pre-unification West Germany in order to be able to take account of the fact that the former East Germany had a communist regime from the mid-1940s through 1989. In two other cases nations have split during the 20th century: Taiwan separated from the remainder of China in 1950 and Czechoslovakia split into the Czech Republic and Slovakia in 1991. We treat these as four distinct nations. We also treat as distinct nations those formed when the Austro-Hungarian Empire disbanded after WWI and those formed after the Soviet Union collapsed in 1991.

page of the *ISMF*.[15] Two criteria govern inclusion of files in the *ISMF*: they must be based on a probability sample of a national (or sub-national) population (or labor force) and they must include detailed information on father's and respondent's occupation.

Although industrialized nations are over-represented, our coverage of world societies is quite broad. The 54 nations analyzed here include 14 former or current Communist nations (Bulgaria, China, Croatia, the Czech Republic, Estonia, former East Germany, Hungary, Latvia, Poland, Romania, Russia, Slovenia, Slovakia, and the Ukraine); nine Asian nations (China, Cyprus, India, Israel, Japan, Malaysia, the Philippines, Taiwan, and Turkey); four Latin/South American nations (Brazil, Chile, Mexico, and Suriname), and two African nations (Nigeria and South Africa). Most of the surveys utilized here were conducted from the 1970's onward, but they range from a 1947 U.S. survey to a number of surveys carried out in 2008.

We restrict our analysis to people age 25-74 for whom we have complete information on all variables included in the micro analysis: gender, educational attainment, father's (or mother's) educational attainment,[16] and father's occupational status. The lower age cutoff was chosen on the assumption that by age 25 nearly everyone has completed his/her education, even in highly

[15] Access to the catalogued surveys must be arranged with Ganzeboom (HBG.Ganzeboom@fsw.vu.nl) since use of some of the data sets requires the permission of the original investigators.

[16] Father's and mother's education are known to have similar effects on respondent's educational attainment. Thus, in cases where father's education was missing and mother's education was available and the difference in the mean father's and mother's years of schooling within a context was less than two years, we substituted mother's years of schooling, which reduced the amount of missing data on father's years of schooling from about 19% to about 16%. (We declined to make the substitution for the 17 contexts for which the difference was greater than 2.0 on the ground that larger discrepancies were likely to introduce bias in the measurement of father's education, but of an unknown sort.) In our judgment, making the substitution was preferable to averaging mother's and father's years of schooling since many studies did not include information on mother's years of schooling. Still, the relatively large amount of missing data on parental education is troublesome, especially given that people from low status origins are especially unlikely to know about the education of their parents.

industrialized nations.[17] The upper age restriction was imposed to reduce the extent of sample selection bias due to differential mortality[18] and to minimize reporting error due to cognitive deficits among the very elderly. These restrictions reduced our sample size from about 1.8 million cases to about 1.1 million cases.[19]

To prepare our data for analysis, we first combined all surveys conducted in each nation and then for each nation defined "schooling cohorts" by adding 10 years to the year of birth and dividing the sample into five-year intervals. The advantage of pooling data from many surveys is

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[17] In our data, among those age 25-29 5.6% were in school and among those age 30-34 1.4% were in school. Moreover, of those in school between ages 25-29 90.2% already had completed upper middle schooling and the same was true of 79.9% of those in school between ages 30-34. Thus, even for those in school at age 25 or older, the error in years of schooling is mainly restricted to those with at least some tertiary schooling.

[18] Regardless of the exact age cutoff there is bound to be sample selection bias with respect to social origins since in many nations differential mortality by education and occupational status begins at relatively young ages. For example, Xie (1996, p. 41) found that in China (using data from the 1990 census) "the standardized crude death rate for the male illiterate population is, respectively, 3.5, 2.2, 1.6, and 1.3 times that of college educated, senior high educated, junior high educated, and primary educated persons. Female illiterate crude death rates are, respectively, 3.1, 1.9, 1.6, and 1.2 times higher." And Banister and Hill (2004, p. 64) showed that for both men and women Chinese age-specific mortality rates increased in a nearly linear way from before age 30 until after age 90 (Banister and Hill 2004, p. 64). Given the tradeoff between reducing the variability in and upwardly biasing educational attainment among elderly respondents and increasing the range of cohorts for which we have data, we have opted for the latter. The resulting bias, if any, is conservative in that it reduces the range of variability in educational attainment across cohorts.

[19] Specifically, we began with 1,814,893 cases. Restricting respondents to people age 25-74 reduced the sample to 1,483,423. Omitting cases with missing data on any of the variables in the micro model further reduced the sample to 1,128,023, spread over 794 contexts. Finally, dropping contexts containing fewer the 100 cases, and also omitting for Israel the seven schooling cohorts prior to 1949, on the ground that for these cohorts the overwhelming majority had completed their schooling outside Israel, reduced the sample size to 1,115,615 and the number of contexts to 654. This is our analytic sample. The age restriction is, of course, an analytic decision. However, the 24% reduction in the sample size due to missing data is troublesome since missing data on socioeconomic variables often are not "missing completely at random" (Rubin 1987; Little and Rubin 2002). One possibility would have been to carry out a multiple imputation of these variables for each of the 541 surveys. But this is a daunting task, given the diversity of predictor variables available in each survey and the variability of coding specifications for specific variables across surveys. Moreover, it is unclear how to properly combine imputed data when we reorganize our data by county*cohort (see below). Thus, we have forgone any attempt at multiple imputation at the micro level; but see below for a description of multiple imputation carried out at the macro level. The final source of data loss is the omission of older cohorts for Israel on the ground that most members achieved their education before arriving in Israel, and restriction of the analysis to contexts including at least 100 people, which we did to achieve reliable estimates of the micro-level coefficients but also of the macro-level variables constructed for each context from the context-specific cases. But these restrictions had hardly any effect on the total sample size, although they did result in the loss of 140 contexts.

that doing so increases statistical power and also smooths out any idiosyncracies of individual surveys; using multiple surveys amounts to a multiple measurement perspective at the macro level, which is an appealing way to deal with measurement and comparability problems.[20] By contrast, using single surveys for each nation makes it impossible to distinguish between country-specific effects and survey-specific effects produced by variations in sampling procedures, question wording, coding procedures, etc. For these reasons, we consider our strategy superior to that of analysts such as Erikson and Goldthorpe (1992, pp. 49-53) who based their comparison of intergenerational occupational mobility in 12 nations on the (undefined) "highest quality" survey available for each nation.

Crossing schooling cohorts by nations produces a set of "contexts" within each of which we estimate a micro model of the process of educational attainment, described further below. We then study how the micro process varies depending on the institutional and demographic features of the context, that is, in each nation at the time respondents were age 10-14.[21] While the assumption that

[20] This position is quite foreign to the stance of most current comparative macro-sociological research, which tends to involve comparisons of two or at most a handful of countries. We find it a bit odd that comparisons of very limited numbers of countries are quite acceptable; certainly, no one would believe the results of similar designs at the individual level. Moreover, while the standard approach at the macro level is to go for the "best data" rather than all available data, this would be a quite unacceptable strategy at the individual level. However, the logic of macro- and micro-level comparisons is fundamentally similar: one tries to assess the relative influence of a number of different variables on one or more outcomes. The main difference is that usually there are data for many individuals available to model micro processes, but for only a few societies to model macro processes. It is therefore particularly ill advised to restrict the data available for comparative analysis. Exactly the same argument holds for multilevel models such as ours: first we use all the available data to analyze micro processes within each context, and then we use all the available macro data to analyze variability in the micro process across contexts.

[21] In some of our preliminary analysis we treated men and women as forming different contexts. However, separating men and women makes it awkward to assess societal determinants of the gender gap. Second, it reduces the number of cases per context and therefore the reliability of coefficients derived from the micro data. Third, it ignores the fact that in most nations males and females attend school together and are subject to many of the same influences. Its only advantage would be if there were important interactions between gender and the effect of social origins; but these are more readily handled by introducing interactions between gender and the social origin variables into the micro equations. We explored doing this. However, the interaction of gender and father's years of schooling was significant at the .05 level (2-tailed) in only 119 contexts and the interaction of gender and father's ISEI was significant in only 120 contexts (out of 643 contexts that included at least 10 men and at least 10 women).

nations define distinctive contexts is relatively unproblematic,[22] the appropriate choice for when in the course of school careers the social environment has the greatest impact on pupils is much less clear. We have somewhat arbitrarily settled upon age 10-14 on the ground that most people are still in school at that age but face their first major decision point—whether to continue on to lower middle school. Since our schooling cohorts range from 1904 to 1994, we could have as many as 1,026 (=19*54) contexts. In actuality, we have 654 contexts since we lack data for many cohorts in many nations (the number of cohorts per nation ranges from two for Croatia to 18 for five nations—France, Hungary, the Netherlands, Poland, and Sweden) and, as noted, we have restricted our analysis to contexts with at least 100 cases after excluding missing data, which eliminates Iceland and Suriname, neither of which has a minimum of 100 cases in any cohort.

Schooling cohorts alternatively may be thought of as people age 10 during a five year period, say 1915-1919, or as people age 10-14 in a specific year, say 1919. While a few contexts include small numbers of respondents, many are very large; indeed, the size of contexts ranges from the imposed minium of 100 to 25,591, with an unweighted mean of 1,706 and a weighted mean of 14,331. We corrected for the greater precision of estimates based on more cases by weighting each context by a variable constructed from the standard errors of the coefficients in the micro equations (details below).

Further, inspection of the significant coefficients revealed no systematic pattern or association with macro factors in either case. Given the relative paucity of significant effects, and their unsystematic character, we decided to forgo additional exploration of the macro conditions under which gender interacts with social origins in affecting educational attainment. Finally, when we did separate contexts by gender in earlier analysis it turned out that there was little difference in results for male and female contexts.

[22] (except insofar as there is unobserved heterogeneity within nations because of regional and similar differences in educational systems)

VARIABLES

We utilize two kinds of variables: micro-level variables that distinguish the characteristics of individuals, and contextual, or macro-level, variables that distinguish the characteristics of social contexts: schooling cohorts within countries.

Micro Variables

Educational attainment.—As noted earlier, in this paper we restrict our analysis to the determinants of total educational attainment—that is, "virtual" years of school completed. We refer to this variable as "virtual" years rather than actual years because we engaged in extensive recoding of the data in order to render a wide variety of initial educational classifications comparable within and between nations. Our basic strategy was to recode educational categories to years of school completed where the correspondence was unambiguous and then to interpolate the remaining categories in such as way as to preserve a monotonic relationship between any rank ordering of categories claimed in the original data and our new educational measure. We then validated our preliminary assignments in two ways: by consulting experts on each educational system and by assessing the linearity of the relationships between our new educational measure and various criterion variables (father's occupational status, recoded spouse's education, and respondent's occupational status), and made adjustments as necessary. We think our resulting scales of "virtual" years of school completed for each country is cross-nationally valid with respect to both the assignment of scores to categories and their distributional properties. The latter is an important claim since we rely heavily upon comparisons of the means and standard deviations of educational attainment across cohorts and nations.

Father's education.—We measure father's education using the same coding of "virtual" years of schooling as we used for respondents. We center this variable within cohorts and nations by subtracting the within-context mean from each observation. Thus, what we are measuring is advantage or disadvantage in social origins relative to those attending school in the same nation at a specified point in time. This, of course, is not the same as measuring the schooling of fathers of a given age since there is non-trivial variation in age at paternity.

Father's occupation.—We measure father's occupational status when the respondent was growing up (typically the variable refers to when the respondent was age 14) by the *International Socio-Economic Index of Occupations*, the *ISEI* (Ganzeboom, De Graaf, and Treiman 1992; Ganzeboom and Treiman 1996, 2003). All data were recoded from national occupational classifications to both the 1968 and 1988 versions of the *International Standard Classification of Occupations, ISCO* (International Labour Office 1969, 1990), for each of which a standard recode module to assign *ISEI* scores exists (Ganzeboom et al. 1992; Ganzeboom and Treiman 1996). We assigned both 1968 and 1988 *ISEI* scores and averaged them. To facilitate interpretation, this variable was centered on the within-context mean and divided by 10.

Gender.—This variable is coded 1 for females and 0 for males.

Micro weight.—Many of the surveys used in the micro analysis included weight variables to adjust for differential sampling and/or response rates. Moreover, some of our data are from multiple waves of panel studies with partly overlapping respondents, which means that such cases must be down-weighted so that they are appropriately representative of the population. We take account of variations in effective sampling rates by utilizing Stata 13's "importance" weights (iweights) which, uniquely among Stata's weight options, do not normalize the weights to the number of observations

(StataCorp 2013). Doing this yields for each context a weighted sample size that correctly reflects the actual number of independent observations and gives surveys that contribute more such cases to a context more weight, which is exactly what we want on the assumption that all surveys contributing cases to a context are probability samples drawn from a single population—a given schooling cohort in a given nation.

Macro Variables

Where possible, we constructed our indicators of the social environment from external sources in order to avoid any possibility of introducing artifactual statistical relationships. However, where adequate external data were not available, we relied on measures calculated directly from our micro data. These instances are indicated below. Finally, where it was possible to construct several measures of the same concept we did so and then compared the results to assess the robustness of our estimates. Because some of the macro indicators are missing values for some contexts, we imputed missing values for such indicators using Stata 13's `-mi-` commands, carrying out 50 imputations.[23]

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[23] Specifically, we imputed missing values for our Development variable, GDP per capita, the percentage of the labor force engaged in agriculture, the percentage of the population enrolled in primary school, and the percentage enrolled in secondary school. The right-hand side variables used in the imputation were the ending year of the interval defining the context, the within-context mean and standard deviation of the respondent's years of schooling, the father's years of schooling, the father's occupational status (ISEI), the proportion female, the percentage with at least lower middle schooling, the percentage with at least upper middle schooling, whether the context had a communist regime, the percentage of fathers in agriculture or traditional sales occupations, each of the micro coefficients and their standard errors, the weight variable constructed from the standard errors, an inequality measure constructed by averaging the standardized standard deviations of father's ISEI and father's years of schooling, an inequality measure constructed by averaging the standardized standard deviations of father's ISEI and respondent's years of schooling, this measure lagged one cohort, the standard deviation of years of schooling lagged one cohort, and several welfare measures: the year health insurance was first enacted in a nation and similar measures for unemployment insurance, family and child subsidies, and when women were granted the vote, whether each of the four measures was introduced prior to the beginning year of the schooling cohort, the same measures for non-communist contexts with communist contexts coded zero, and the number of welfare measures among the four introduced prior to the beginning year of the cohort for all nations and for non-communist nations.

Educational expansion.—Our measure for each context is the proportion of the population enrolled in primary or secondary school. While it would have been preferable to have as a measure the proportion of the school age population enrolled, this information was not available for any time series that extended backwards much beyond the late 20th century. Although the measure we use is vulnerable to bias due to variations in the age distribution of the population—all else equal we would expect younger populations to have higher enrollment rates—we doubt that the amount of bias is great. However, because of the vulnerability of this measure to bias, we carry out sensitivity analysis using an alternative measure of educational expansion drawn from our data—the percentage of the cohort attaining at least some lower middle schooling, which ranges from 1% to 100%. The two measures are only moderately correlated ($r = .66$).[24]

Educational inequality.—This is the standard deviation of the years of schooling achieved by the members of the previous cohort. We lag the variable by one cohort to avoid any possibility of a statistical tautology resulting from the fact that the standard deviation is computed from the dependent variable in the micro equation. An alternative measure, the standard deviation of father's years of schooling, is unsatisfactory because the fathers of those in each cohort vary in age and thus vary in the year they completed their education, which means that such a variable would pertain to no fixed point in time. Although it would have been preferable to construct a measure of educational inequality external to our data, we were unable to identify such a measure with any historical depth and, indeed, cross-nationally consistent measurements of educational distributions are hard to come by for even a single point in time.

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[24] All the correlations in this section are based on data weighted by our standard error weight measure (see below). All correlations based on macro variables missing any observations, which is the case here, are calculated across all 50 imputations.

Societal development.—There are many ways to measure societal development/ industrialization/modernization. However, most such measures are highly correlated. Building on the work of Rijken (1999, pp. 169-178), Ganzeboom (2007) created a World Development Index (WDI) for 50 of our 54 nations, using a set of indicators from the *1815-1973 Political and Social Indicators* compiled by Banks (1976) combined with a newer set of such indicators for the 1965-1994 period provided by the World Bank (1997). The WDI indicator combines data on urbanization, economic indicators, and the welfare situation as well educational opportunities as measured by enrollment indicators. Detailed information is provided on the *ISMF* website cited above. Here it suffices to say that the final result, which we employ here, was a 0-1 measure that, as initially constructed, mapped contexts between the extremes of 0 (India in 1930) and 1 (United States in 1980). Since we subsequently have added four nations and additional contexts, we impute values for missing contexts as noted above. We also explore three alternative measures:

- The percentage of fathers of each cohort engaged in occupations that were neither in agriculture nor in the traditional sales, service, and craft sector, computed for each context from our data. As noted above, the ability to perform such occupations is usually directly transmitted from parents to offspring and does not require formal schooling. Moreover, a large agricultural sector usually indicates under-development. A less conventional measure is the proportion of the male labor force engaged in small retail, service, or craft enterprises, enterprises that tend to be run by families and to be passed on from generation to generation. Unfortunately, we have no good measure of such employment. We approximate it by counting the proportion of (employed) fathers who were classified as “working proprietors in wholesale or retail trade,” “salesmen, shop assistants, and related workers,” or “sales workers n.e.c.” in the *ISCO 68* classification.

We used the *ISCO 68* classification because it distinguishes small shop keepers better than does the *ISCO 88* classification.[25] Exploration of the combined measure (the percentage engaged neither in agriculture nor in traditional sales occupations) suggested that it better captures development than does the simple proportion not engaged in agriculture.

- A second measure of the proportion of the labor force not engaged in agriculture, drawn from the *Cross-national Time-series Archive* (CNTS Archive 2002).
- Gross domestic product (GDP) per capita. This is the conventional measure of how economically developed a nation is at a given point in time. We use the measure created by Haber and Menaldo (2011), who worked hard to generate a consistent time series over a long period of time.[26]

Cohort.—This variable, used to measure secular trends, was defined simply by taking the last two digits of the ending year of each birth cohort. Thus, for example, all members of the 1910-1914 cohort were coded 14 on this variable.

Communism.—This is a dummy variable, scored 1 for schooling cohorts corresponding to the communist period of sometime-communist nations and scored 0 otherwise. Specifically, we specified as communist the period between 1945 and 1989 for the nine non-Baltic Eastern European nations in our sample (Bulgaria, Croatia, the Czech Republic, East Germany, Hungary, Poland,

[25] Specifically, we defined traditional sales occupations as including *ISCO 68* categories 4100-4109 and 4500-4900 (see Ganzeboom et al. 1992, Appendix B). However, in some cases, the *ISCO 68* codes were highly aggregated. In cases where there were 20 or fewer distinct categories in the *ISCO 68* classification and more than 20 categories in the *ISCO 88* classification, we used the *ISCO 88* codes, defining traditional sales occupations as codes 5230 and 9100-9112. In both classifications, agricultural occupations have codes 6000-6999. Where neither the 1968 or 1988 version of *ISCO* had more than 20 codes, we used the proportion in agriculture as our measure of traditional occupations.

[26] Prof. Haber has kindly made available to us the data set and documentation he and Victor Menaldo constructed.

Romania, Slovakia, and Slovenia);[27] the period between 1940 and 1989 for the two Baltic nations (Estonia and Latvia); the period between 1920 and 1989 for the two formerly Soviet nations (Russia and the Ukraine);[28] and the period 1950 through 1994 for China. Of course, the remaining 40 nations were scored 0 for all cohorts.

Macro weight.—To adjust for the greater chance variability in point estimates of coefficients based on small samples, we have weighted all computations in the macro analysis by the product of the reciprocal of the squared standard errors for the three predictor variables in the micro equation: father's years of schooling, father's occupational status (ISEI), and gender (Sanchez-Meca and Marin-Martinez 1998). As we can see from the summary coefficients in Table 2, weighting the data in this way substantially reduces the variability of the coefficients across contexts because extreme coefficients arising by chance from samples with large standard errors are given reduced weight. Doing this also increases the mean size of the intercept, from 10.1 to 10.5, which reflects the fact that our more precise estimates tend to come from developed nations, mainly because the sample sizes are larger. Finally, compared to the unweighted data, the weighted data reveal a smaller average effect of gender. This most probably reflects the especially pronounced imprecision of the gender coefficient in contexts with very few cases and even fewer females.

Centering.—To facilitate interpretation of the results, we center our measures of development, educational expansion, and educational inequality around their means and center our

[27] Although several Eastern European nations had social democratic governments for a few years after World War II, all except Yugoslavia (represented in our sample by Croatia and Slovenia) were occupied by the Soviet Army during these years, starting in 1945 or sometimes earlier, and the Soviet Union exerted heavy pressure to form communist governments (Snyder 2010).

[28] 1920 is, of course, slightly later than the establishment of the Soviet Union but given the chaos of the Revolution and its aftermath, which included a civil war and much fighting between Soviet and foreign forces in the Ukraine and other parts of the Soviet Union (Figes 1998; Snyder 2010), it would seem to be an appropriate beginning date for the impact of communist policies on education.

cohort measure around 50. The result of doing this is that in each macro equation the intercept is the expected value of the dependent variable (the micro coefficient being studied) for non-communist contexts in 1950 with average levels of development, educational expansion, and educational inequality, and the coefficient for communism gives the expected difference in 1950 between communist and non-communist contexts with average levels of development, educational expansion, and educational inequality.

ANALYSIS

As noted, we carry out our estimation in two steps. First we estimate a micro model for each of the 654 contexts, using OLS. We then treat the coefficients yielded by the micro model for each context as variables in a macro-level analysis, predicting the size of each coefficient from the characteristics of each context that we discussed above. The macro models are estimated using Stata 13's `-xtgls-` command, specified to permit nation-specific first-order autocorrelation (AR1), which is surely likely give the continuity of social structure, and a heteroskedastic error structure with no cross-sectional correlation (the specification of no cross-sectional correlation is necessary because the panels are unbalanced—that is, nations vary with respect to the contexts for which we have data).[29]

[29] An alternative would have been to carry out the estimation via conventional hierarchical linear modeling (Raudenbush and Bryk 2002). The main advantage of hierarchical linear models is that the micro and macro equations are estimated simultaneously, which results in optimal estimation of both the micro and macro errors. However, there are several disadvantages, at least as implemented using Stata 13's `-mixed-` command. First, it is very slow whereas `-xtgls-` is very fast, because `-mixed-` operates on micro data whereas `-xtgls-` operates on macro data—the difference between 1.1 million and 654 cases in the present analysis. This is particularly true given that, as noted, we employ multiple imputation with 50 imputations to deal with missing macro-level data. Secondly, `-mixed-` sometimes does not converge, at least in our analysis. We did some exploration of multilevel estimation using `-mixed-` and had difficulty achieving convergence, although simpler models than those estimated using `-xtgls-` and presented here did converge and produced similar results using the two commands. Finally, `-mixed-` proved to be rather inflexible given our analytic needs.

Micro Model

The micro-level equation, estimated for each of the 654 contexts, is

$$E = \beta_0 + \beta_1(E_F - \overline{E}_F) + \beta_2(S_F - \overline{S}_F) + \beta_3(F) + \varepsilon \tag{1}$$

where E = years of school completed by the respondent; E_F = years of schooling completed by the respondent's father; S_F = the status of the father's occupation (that is, the father's *ISEI* score) when the respondent was approximately age 14; F is a dummy variable, scored 1 for females and scored 0 for males; and ε is the error term. As indicated in our above discussion, we expect positive coefficients β_0, β_1, and β_2 in all contexts (that is, all schooling cohorts in each nation) and a negative coefficient β_3 in all contexts until near the end of the 20th century when the educational attainment of males and females converged or, in developed nations, even crossed (DiPrete and Buchmann 2013). Note that because we center the father's education and father's occupation variables around their within- context means, the coefficient β_0 gives the expected years of schooling of men whose fathers are at the average with respect to years of schooling and ISEI scores relative to all fathers of respondents in the same context and β_3 gives the difference between the expected years of schooling of females and males whose fathers were at the average with respect to years of schooling and ISEI scores relative to all fathers of respondents in the same context.

The (weighted) average values of the coefficients across the 654 estimated equations are:

$$E = 10.5 + .344(E_F - \overline{E}_F) + .470(S_F - \overline{S}_F) - .352(F) + \varepsilon \tag{2}$$

This result tells us that, on average, if we compare people living in a given country at a given point of time, those whose fathers differ by a year in schooling can be expected themselves to differ by about a third of a year of schooling, net of father's occupational status and gender. Similarly, those whose fathers' occupations differ by 10 points on the ISEI scale can be expected to differ by nearly half a year of schooling, net of father's education and gender. Women on average get about a third of a year less schooling than men of the same social status living in the same nation at the same point in time. Finally, the intercept tells us that, on average, men whose fathers have average (within-context) education and occupational status would be expected to achieve 10.5 years of schooling, whereas women would be expected to achieve 10.1 (= 10.5 - .352) years of schooling. These results are about what we would expect given the theoretical discussion earlier about the factors driving educational attainment at the individual level.

The standardized coefficients also are of interest. Here are the (weighted) average values across the 654 contexts:

$$e = .329(e_F - \overline{e}_F) + .198(s_F - \overline{s}_F) - .061(f) + \varepsilon \tag{3}$$

These coefficients tell us that of the three predictor variables, father's years of schooling is the most important determinant of educational outcomes, in the sense that a one standard deviation difference with respect to father's schooling predicts, on average, an expected 1/3 standard deviation in the schooling of our respondents, compared to a much smaller effect of father's occupational status and a very small effect of gender.

Of perhaps even greater interest than the mean values of the micro coefficients is their great variability. Leaving aside for the moment the most extreme values (shown in Table 3), let us focus on the (weighted) 1^{st} and 99^{th} percentiles shown in Table 2; these give us a good sense of the range of the variable uncontaminated by outliers. For the two social origin variables the the largest (99^{th} percentile) coefficients are about three times as large as the smallest (1^{st} percentile) coefficients. These are very large differences. For example, in the contexts with the greatest equality of opportunity (measured at the 1^{st} percentile), those whose fathers' schooling differed by a year would be expected to differ in their own attained schooling by less than a fifth of a year (precisely, .196) whereas in the contexts with the least equality of opportunity, those whose fathers' schooling differed by a year would be expect to differ in their attained schooling by nearly 3/5ths of a year (precisely, .577). The ratio of the 99th to 1st percentile coefficients is similar for father's occupational status (2.99 = .797/.267). Finally, the effect of gender ranges from handicap of nearly 1 ½ years for women at the 1st percentile (precisely, 1.43) to an advantage of more than half a year at the 99th percentile (precisely, .55). Clearly, there is a substantial amount of variability across social contexts in equality of educational opportunity.

Table 3 provides further detail, by showing the contexts with the five smallest and five largest coefficients for each of the micro-model predictor variables. These results are reassuring, since they are quite consistent with *a priori* expectations. The smallest intercepts—that is, the lowest levels of expected schooling for the sons of men with average socioeconomic status—are for Nigeria and India prior to the middle of the 20th century whereas the largest intercepts are for highly developed nations toward the end of the century. The effect of father's education is strongest in countries that were poorly developed prior to the Second World War and generally is weakest for communist and post-communist nations. A roughly similar, although less clear cut, pattern holds for the effect of father's occupational status. The female disadvantage is strongest in pre-war East and Southeast Asian nations and is reversed—there is a female advantage—in several communist and post-communist nations plus Sweden. Finally, the dependence of education on social origins and gender (indicated by the R^2's) is greatest in South Africa, reflecting the strong correlation of race and opportunity (Treiman, McKeever, and Fodor 1996; Treiman 2007, 2009, pp. 39-42).

Macro Models

We next turn to analysis of what factors promote variations in equality of opportunity. To do this, we estimate a set of four successively more complex models predicting each of our micro-level outcomes—the coefficients for the intercept, father's years of schooling, father's ISEI/10, and female gender; descriptive statistics for the variables in these models are shown in Table 4. The first model considers the effect of societal development, educational expansion, and educational inequality. The second model adds cohort to assess the effect of secular trends on equality of educational opportunity. The third model adds whether the regime was communist and the fourth model adds the interaction between cohort and the communist dummy variable to assess whether the effect of communism declines over time, as we predict. Although, as indicated in Table 1, in some cases we have no predictions regarding the signs of the macro-level predictors, we include all predictors in each model. The full macro model—Model 4 in Table 5—is given by

$$\begin{aligned}
\beta_0 &= \gamma_{00} + \gamma_{01}(D-\overline{D}) + \gamma_{02}(E-\overline{E}) + \gamma_{03}(I-\overline{I}) + \gamma_{04}(T-50) + \gamma_{05}(C) + \gamma_{06}((T-50)*C) + \mu_0 \\
\beta_1 &= \gamma_{10} + \gamma_{11}(D-\overline{D}) + \gamma_{12}(E-\overline{E}) + \gamma_{13}(I-\overline{I}) + \gamma_{14}(T-50) + \gamma_{15}(C) + \gamma_{16}((T-50)*C) + \mu_1 \\
\beta_2 &= \gamma_{20} + \gamma_{21}(D-\overline{D}) + \gamma_{22}(E-\overline{E}) + \gamma_{23}(I-\overline{I}) + \gamma_{24}(T-50) + \gamma_{25}(C) + \gamma_{26}((T-50)*C) + \mu_2 \\
\beta_3 &= \gamma_{30} + \gamma_{31}(D-\overline{D}) + \gamma_{32}(E-\overline{E}) + \gamma_{33}(I-\overline{I}) + \gamma_{34}(T-50) + \gamma_{35}(C) + \gamma_{36}((T-50)*C) + \mu_3
\end{aligned} \tag{4}$$

where D = the level of development of the context, E = the level of educational expansion, I = the level of educational inequality, T = time, measured by the final year defining each cohort, C is a dichotomous variable scored 1 for communist contexts and scored 0 otherwise, and the μ's are the macro-level error terms. The macro equations for the first through third models are identical to Eq. (4) except that various coefficients are set to zero. Having estimated the set of models just described, we then carry out sensitivity analysis by substituting alternative measurements of some of the macro variables.

Consider first the determinants of the micro-level intercept. Recall that because of the way we have defined the micro equation, the micro-level intercept represents the expected level of schooling for men whose father's education and occupational status are average for their cohort. What accounts for variation in this coefficient? First, we note that the macro intercept is consistently a bit larger than 10 across all models. This tells us that the expected level of schooling is just over 10 years for men who were in non-communist 1950 schooling cohorts with average levels of development, educational expansion, and educational inequality. In short, already by 1950 the average level of educational attainment was quite high, the equivalent of some senior high school.

Now consider the specific models. Model 1 indicates, not surprisingly, that the average level of schooling is strongly associated with the level of development. Also, as predicted, educational expansion increases the level of average schooling. That is, among equally developed contexts the contexts with higher fractions of the population enrolled in school exhibit higher levels of average educational attainment. Interestingly, educational inequality, about which we made no prediction, also increases the level of schooling; but, as we will see, this effect disappears in more elaborate models.

Model 2 permits us to assess the effect of any secular trend toward increasing education across the 20th century. This effect is very large. The coefficient of .033 implies an expected increase of about three years of schooling (precisely 2.97 = .033*90) between the beginning of the century and the 1990-1994 cohort, the most recent cohort for which we have data, net of the level of development, the extent of educational expansion, and the degree of educational inequality. That there is, as hypothesized, a significant effect of time net of the effect of development is consistent with the claim that at any level of development there is inexorable pressure to get as much education as possible, which translates into social and political pressure to expand educational opportunities. Because in most nations development has increased over the 20th century, part of the seeming effect of development estimated in Model 1 is taken up by the time variable introduced in Model 2, reducing the net effect of development. In a similar way, the introduction of time reduces the effects of economic expansion and educational inequality seen in Model 1 to non-significance.

Model 3 introduces a distinction between communist and non-communist contexts and Model 4 includes an interaction between time and whether the context is communist. The optimal strategy for assessing these two models is first to consider Model 4, which posits a difference between communist and non-communist nations over time, and then, if the effect of the communist dummy and the interaction between time and the communist dummy are jointly significant, to consider whether the effect of communism changes over time or whether there is simply a time-constant communist effect (Treiman 2009, pp. 124-131). As we see from Model 4, net of other factors those in communist contexts in 1950 appear to average about 8/10ths of a year more schooling than those in non-communist contexts, but there is a substantial decline in the communist advantage over time. A more precise comparison between communist and non-communist contexts is deferred until we discuss Table 6.

The effects of macro-social characteristics on variations in the effect of father's schooling on respondent's schooling are generally strong and consistent with our expectations. We see first that in the "average" context, as we have defined it, each additional year of father's schooling yields about a third of a year of additional schooling for respondents, an outcome that holds across models. Model 1 shows that both development and educational expansion significantly reduce inequality of opportunity (that is, the size of the father's education effect) while educational inequality increases it. Although the effects of development and educational expansion change somewhat in Model 2, they remain significant when time—which has the predicted negative effect—is introduced. Models 3 and 4 reveal that the apparent secular trend does not hold up but is rather an artifact of the concentration of communist regimes in the second half of the 20th century. As predicted, communist regimes exhibit less inequality of educational opportunity—as measured by the size of the father's education effect—than do non-communist regimes; but there is no significant change in the effect of communism over time, which lends support to the "communist affirmative action" hypothesis as against the "new class" hypothesis.

A difference of 10 points on the ISEI scale yields an expected difference of about a half year of schooling in the average context in 1950. The macro-social determinants in variations in the effect of father's occupational status are, as predicted, generally similar to those for the effect of father's years of schooling, except that there is a consistent decline in this aspect of educational inequality over time.

Finally, the disadvantage in educational attainment experienced by women in the average context in 1950 ranged between a third and a half a year, depending on the model. As expected, this effect declined very substantially with development (a positive coefficient indicates an increase in the years of schooling of women relative to men) and over time and, as it turned out, increased with educational inequality (recall that we made no prediction regarding this coefficient), but, contrary to expectation, educational expansion has no significant effect. As expected, women did better relative to men in communist than in non-communist contexts and there appears to have been a decline in the communist advantage for women over time although the coefficient is only marginally significant.

We can further assess the relative size of these effects by contrasting expected net effects for various levels of each of our macro variables. The first three rows of Table 6 give the expected net difference between contexts near the extremes of each distribution—that is, contexts at the 99th and 1st percentiles. We chose the 1st and 99th percentiles to avoid distortion due to extreme outliers. The fourth row gives the contrast between our earliest (1900-1904) and latest (1990-1994) contexts. Consider first the effects on the intercept. Here we see that development and the secular trend over the 20th century have very strong effects. People in the most developed contexts would be expected to average 3.5 years more schooling than those in the least developed contexts, net of all other factors; and those in contexts near the end of the 20th century would be expected to average 2.5 years more than their counterparts at the beginning of the century. By contrast, the effects of educational expansion and educational inequality are quite small, with the difference between the most and least expansive contexts less than half a year in expected schooling (precisely, .39) and the difference between the most and least educationally unequal contexts only .10 year.

To evaluate the effect of communism and how this changes over time, we contrast communist and non-communist nations in 1950 and again in 1990. We choose 1950 as a good approximation for the early years of communist regimes (excepting Russia and the Ukraine). By 1990 communism had ended in all the nations in our sample except China (and Russia, where the complete collapse did not come until 1991). Thus, this a reasonable approximation for the effect of end-stage communism. Because of the way the variables were defined, the 1950 effect is just the coefficient associated with communism, borrowed from Table 5. As already noted, in 1950 the communist bonus in educational attainment was .78 years. However education increased more slowly in communist nations so that by 1990 there was a communist deficit of .26 years, as we see in column 1 of Table 6. In short, net of development, educational expansion, and educational inequality communist nations did not keep up with non-communist nations with respect to the increase in educational attainment over the 40 year communist experiment.

Now consider macro determinants of the effect of father's years of schooling. Here we see that the expected coefficient is .12 points lower in the most developed contexts than in the least developed contexts, net of other factors. This is a very large effect. The remaining effects are also about as expected. Educational expansion moderately reduced the effect of father's schooling. The effect of father's schooling is substantially larger in the most educationally unequal contexts than in the most equal contexts. Communism reduced the effect of father's schooling in 1950 and the communist effect remained largely intact in 1990. However, as would be expected from the non-significant coefficients in Table 5, net of other factors there is almost no reduction in the effect of father's schooling over time.

The effect of father's occupational status also is strongly dependent on macro-social factors. There is a very large reduction in this effect over time—more than 2.6 standard deviations between the earliest and most recent contexts net of other factors (precisely, -2.63 = -.0040*90/.137). Educational inequality sharply increased the effect of father's ISEI; and development reduced the effect. However, the effect of communism was at best modest and became even more modest over time. And the effect of educational expansion was extremely modest.

Finally, women benefitted substantially from development, from secular changes over the 20th century, and, in 1950, from communism, and were disadvantaged noticeably by educational inequality. However, the communist advantage dropped by more than half between 1950 and 1990. For women, educational expansion had little effect.

The bottom row of Table 6 gives the expected differences between the contexts most and least likely to experience equality of educational opportunity—that is, between contexts at the 99th percentile of development, the 99th percentile of educational expansion, and the first percentile of educational inequality in 1990-1994, on the one hand, and contexts at the 1st percentile of development, the 1st percentile of educational expansion, and the 99th percentile of educational inequality in 1900-1904, on the other. (Here we ignore the effect of communism since it varies over time and because there were no communist regimes in 1900-1904 nor, with the exception of China, in 1990-1994.) We see that the expected difference between these extremes is more than six years of schooling among men who are from social origins that are average for their contexts; that the expected difference in the effect of father's schooling is -.35 (nearly four standard deviations); that the expected difference in the effect of father's occupational status is -.74 (more than five standard deviations); and that the expected difference in the effect of gender is 2.2 years. These are extremely large effects.

Sensitivity Analysis

Development.—In our main analysis, we used a composite measure of development that we created by combining 14 distinct measures. Here we explore whether our conclusions would have been altered if we had utilized a different measure of development. There are three candidates: two measures of the composition of the labor force and one measure of gross domestic product (GDP) per capita. As it turns out, all these measures are highly correlated, which makes it likely that they are substitutable for one another. The correlations among our four measures are

	D	G	A_1	A_2
D: Development	1.00			
G: GDP/capita	.93	1.00		
A_1: % not in ag., sales	.90	.90	1.00	
A_2: % not in agric.	.93	.88	.93	1.00

Appendix B shows coefficients for Model 4 of Table 5 except that G, A_1, and A_2, respectively, are substituted for D. As we see, we would be led to essentially the same conclusions regardless of which measure we employed: 10 of the 12 coefficients based on alternative measures of development are highly significant as are all 12 development coefficients in Table 5.

Educational expansion.—The same is true of an alternative measure of educational expansion taken directly from our data—the mean years of schooling obtained by members of the previous cohort (a lag introduced for the same reason that we lagged our measure of educational inequality—to avoid any possibility of a statistical tautology). Here the correlation between the two measures is quite modest: r = .66. Still, substituting the measure drawn from the data yields conclusions qualitatively similar to those reported earlier: three of the four coefficients are comparable in their level of significance and the exception involves a marginally significant coefficient for our main measure and a highly significant coefficient for the substitute measure. In addition, the remaining coefficients in the model mainly have the same sign and significance regardless of what measure is used. Of 112 other coefficients in the models based on the alternative measures, 94 have the same sign and are in agreement regarding their significance level with the coefficients reported in Table 5; 10 have marginally different significance levels (less than .05 vs. between .05 and .20 or between .05 and .20 vs. > .20); and eight are inconsistent with respect to either sign or significance levels. Given these strong qualitative similarities, we have added confidence that our conclusions are not importantly dependent upon our choice of measures.

CONCLUSIONS

In this paper we have explored how the process of educational attainment—which we characterize as depending on social origins and on gender, men generally obtaining more schooling than women until very recently in advanced nations—varies across social settings. We have defined such settings by the combination of the society in which people live and the period at which they were making important schooling decisions, that is, the year they were age 10. Using data from 54 nations, we defined 654 such settings, or "contexts," for which we had full data on the characteristics of men and women and on relevant macro-social factors.

We have been able to show that both the level of schooling reached in a population and the size of the gender gap (the difference in the average schooling of men and women) depend heavily upon the level of societal development of the context. In addition, there is a secular trend toward increased education net of the level of societal development and there is also a secular trend toward increased gender equality in educational attainment.

We also have shown that the degree of equality of educational opportunity—as measured by the effect of social origins (father's education and father's occupational status) on educational attainment—increases with development and that net of development the effect of father's occupational status, but not the effect of father's education, declines over time; we have no clear explanation for the discrepancy in these two effects. Equality of opportunity also depends on two distributional properties of the educational system: the level of educational expansion and the level of educational inequality. Educational expansion promotes equality of educational opportunity because it moves the bulk of students into that segment of the educational system at which dependence of educational attainment on social origins is weak. Educational inequality promotes social closure through education because it widens the gap in cultural capital between those from low and high status origins. Both properties of educational systems affect equality of opportunity in the ways we predicted.

We explored two contrasting hypotheses regarding the effect of (mainly) Central- and East-European communism on equality of educational chances: the "communist affirmative action" hypothesis, which posits a reduction in the advantage of those from high status origins relative to non-communist societies, and the "new class" hypothesis, which posits an increasing advantage for the intelligentsia and *nomenklatura* as they consolidate their positions over time. The results favor the "communist affirmative action" hypothesis: the effects of social origins are reduced in communist contexts but there is no significant evidence of reductions over time in the size of these effects.

The societal variations we have identified are very substantial, as we have shown in Table 6. They suggest strongly that systems of social stratification are not all of a piece but vary dramatically depending on the nature of the social structures within which they are embedded.

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TABLE 1
Summary of Hypotheses Regarding Macro-level Effects

	Macro Variable					
	D	E	I	T	C	T*C
Intercept	+	+	0	+	+	-
Fr's years of school . .	-	-	+	-	-	+
Fr's ISEI	-	-	+	-	-	+
Female	+	+	0	+	+	-

Note: D = development, E = educational expansion, I = social inequality, T = time (ending year of schooling cohort), C = communist regime; + and - indicate the expected sign; 0 means we have no prediction regarding the sign.

TABLE 2
Distribution of the Micro-level Coefficients Across 654 Contexts

	Weighted?	β_0	β_1	β_2	β_3	R^2	*s.e.e.*
Mean	No	10.1	.345	.449	-.45	.249	2.68
	Yes	10.5	.344	.470	-.35	.253	2.78
Standard	No	2.6	.171	.223	.83	.099	.51
deviation	Yes	1.6	.093	.137	.55	.054	.36
Minimum	No Yes	1.1	-.124	-.413	-3.40	.025	1.16
1st percentile . . .	No	2.7	.059	-.056	-2.93	.073	1.58
	Yes	6.7	.196	.267	-1.43	.152	1.63
Median	No	10.7	.312	.436	-.40	.234	2.65
	Yes	10.9	.339	.395	-.35	.250	2.85
99th percentile . .	No	14.0	.856	1.008	1.10	.545	3.76
	Yes	13.7	.577	.797	.55	.380	3.44
Maximum	No Yes	15.1	.955	1.375	1.33	.662	4.07

Notes:

β_0 = the intercept, interpreted as the expected years of schooling in each context for men whose fathers had average education and occupational status.

β_1 = the effect of father's years of schooling on respondent's years of schooling.

β_2 = the effect of father's occupational status (10 ISEI points) on respondent's years of schooling.

β_3 = the effect of female gender on respondent's years of schooling.

TABLE 3
Contexts with Minimum and Maximum Micro-level Coefficients

Coefficient	Minimum				Maximum			
	Nation	Cohrt	Coef	N	Nation	Cohrt	Coef	N
β_0 (intercept)	Nigeria	1934	1.1	134	Cyprus	1984	14.1	210
	Nigeria	1939	1.7	169	Norway	1994	14.2	111
	Nigeria	1944	2.0	216	USA	1984	14.2	942
	India	1924	2.4	474	Norway	1989	14.3	441
	Nigeria	1949	2.5	215	Can.-	1984	15.1	106
					English			
β_1 (effect of father's years of schooling) . . .	China	1994	-.124	161	No. Ireland	1924	.915	476
	Czech Rep.	1994	-.070	112	South Afr.	1934	.916	161
	Austria	1989	-.039	332	Nigeria	1939	.917	169
	Gr. Dem. R	1934	-.012	192	Philippines	1909	.946	201
	Ukraine	1979	.024	301	No. Ireland	1939	.955	645
β_2 (effect of father's ISEI*10)	Cyprus	1949	-.413	130	Portugal	1969	1.092	663
	Ukraine	1994	-.254	147	Can.-	1919	1.101	450
	Philippines	1909	-.186	201	Quebec	1949	1.148	906
	Latvia	1979	-.085	341	Chile	1959	1.215	161
	Croatia	1994	-.083	101	Turkey	1959	1.375	789
					Portugal			
β_3 (effect of female gender)	China	1944	-3.40	296	Latvia	1989	1.11	183
	Taiwan	1939	-3.25	1406	Latvia	1979	1.14	341
	China	1939	-3.19	145	Sweden	1994	1.18	138
	Malaysia	1944	-3.12	279	Estonia	1989	1.25	302
	Taiwan	1934	-3.08	463	Poland	1989	1.33	499
R^2	Croatia	1994	.025	101	Turkey	1949	.551	105
	Can.-	1984	.044	106	South Afr.	1964	.559	700
	English	1969	.052	104	South Afr.	1944	.603	282
	N. Zealand	1934	.058	134	South Afr.	1934	.646	161
	Nigeria	1974	.062	376	Turkey	1954	.662	132
	Latvia							
s.e.e.	Czech Rep.	1924	1.16	218	China	1969	3.85	1,084
	Austria	1919	1.21	416	Chile	1964	3.88	1,458
	Finland	1924	1.39	164	South Afr.	1939	3.90	262
	Finland	1919	1.46	100	Chile	1959	3.95	1,346
	Austria	1924	1.49	1,866	Israel	1949	4.07	1,533

TABLE 4
Descriptive Statistics (Correlations, Means, and Standard Deviations) for the Macro-level Analysis (654 contexts) [a,b]

	β_0	β_1	β_2	β_3	D	E	I	T	C
β_0: Micro intercept	1.00								
β_1: Micro fr's educ. coef. .	-.53	1.00							
β_2: Micro fr's ISEI coef. . .	-.67	.16	1.00						
β_3: Micro female coef.	.57	-.28	-.61	1.00					
D: Development[c]	.85	-.34	-.73	.75	1.00				
E: Educational expansion[c] .	.72	-.48	-.51	.44	.63	1.00			
I: Educational inequality . .	.05	.22	.22	-.04	.05	-.01	1.00		
T: Ending yr of cohort	.67	-.22	-.63	.67	.74	.42	.36	1.00	
C: Communist regime	-.13	-.20	-.00	.10	-.31	-.17	-.12	-.11	1.00
Mean	10.5	.34	.47	-.35	.58	.16	3.30	63	.12
Standard deviation	1.6	.09	.14	.55	.21	.03	.48	16	.32
Minimum	1.1	-.12	-.41	-3.40	-.39[d]	-.01[e]	1.32	4	0
Maximum	15.1	.96	1.38	1.33	1.03[d]	.29	5.56	94	1

Notes:

[a] These are weighted coefficients; see the text for a discussion of the level-2 weights.

[b] See text for details on the construction of these variables.

[c] Because D and E include imputed values and the Stata 13 `-mi-` commands do not permit estimation of correlation coefficients from imputed data, we estimated these coefficients by pooling data from all 50 imputations and executing Stata's `-corr-` command. That is, the correlations involving these variables and the means and standard deviations for these variables are based on 32,700 cases (= 654 observations * 50 imputations).

[d] Although this variable, as originally constructed, had a 0-1 range, the multiple imputation procedure produced values outside this range.

[e] Although this variable, as constructed, has a theoretical minimum of zero, the multiple imputation procedure produced negative values.

TABLE 5
Macro-level Coefficients for a Multi-level Model of Educational Attainment (p-values in Parentheses).

	Model 1	Model 2	Model 3	Model 4
Micro intercept:				
Macro intercept	10.68 (.000)	10.28 (.000)	10.27 (.000)	10.23 (.000)
(D) Development	5.81 (.000)	4.13 (.000)	4.80 (.000)	4.73 (.000)
(E) Educational expansion	4.12 (.019)	1.86 (.282)	2.62 (.104)	2.80 (.085)
(I) Educational inequality .	.333 (.000)	.066 (.506)	.128 (.139)	.046 (.586)
(T) Cohort		.033 (.000)	.023 (.000)	.028 (.000)
(C) Communist regime . . .			.558 (.000)	.781 (.000)
(C*T) Comm.*Cohort . . .				-.026 (.000)
Joint sig. of C & C*T				(.000)
Father's education:				
Macro intercept	.317 (.000)	.336 (.000)	.333 (.000)	.336 (.000)
(D) Development	-.160 (.000)	-.082 (.012)	-.178 (.000)	-.165 (.000)
(E) Educational expansion	-.370 (.001)	-.448 (.000)	-.463 (.000)	-.471 (.000)
(I) Educational inequality .	.059 (.000)	.069 (.000)	.052 (.000)	.056 (.000)
(T) Cohort		-.0015 (.001)	-.0001 (.677)	-.0004 (.230)
(C) Communist regime . . .			-.073 (.000)	-.082 (.000)
(C*T) Comm.*Cohort . . .				.0004 (.518)
Joint sig. of C & C*T				(.000)
Father's ISEI:				
Macro intercept	.455 (.000)	.515 (.000)	.517 (.000)	.520 (.000)
(D) Development	-.390 (.000)	-.161 (.000)	-.195 (.000)	-.191 (.000)
(E) Educational expansion	-.479 (.001)	-.261 (.055)	-.330 (.024)	-.344 (.019)
(I) Educational inequality .	.055 (.000)	.095 (.000)	.085 (.000)	.088 (.000)
(T) Cohort		-.0041 (.000)	-.0039 (.000)	-.0040 (.000)
(C) Communist regime . . .			-.049 (.001)	-. 060 (.004)

continued

TABLE 5. Continued

	Model 2	Model 3	Model 4	Model 5
(C*T) Comm.*cohort				.0002 (.830)
Joint sig. of C & C*T				(.000)
Female:				
Macro intercept	-.323 (.000)	-.481 (.000)	-.472 (.000)	-.473 (.000)
(D) Development	1.750 (.000)	1.942 (.000)	1.461 (.000)	1.442 (.000)
(E) Educational expansion	.710 (.201)	.480 (.358)	.800 (.133)	.621 (.235)
(I) Educational inequality .	-.068 (.020)	-.199 (.000)	-.129 (.000)	-.132 (.000)
(T) Cohort		.0140 (.000)	.0077 (.000)	.0084 (.000)
(C) Communist regime . . .			.526 (.000)	.537 (.000)
(C*T) Comm.*Cohort . . .				-.0077 (.059)
Joint sig. of C & C*T				(.000)

Note: The coefficients shown are estimated using the Stata 13 command, `-xtgls-`. The coefficients are the equivalent of OLS regression coefficients and are interpreted in the same way. For example, the coefficient associated with Development in Model 1 of the intercept equation indicates that the expected difference in years of schooling for those living in the theoretically most and least developed societies is 5.56 years, holding constant the level of educational expansion and educational inequality.

TABLE 6

Expected Differences in the Model 4 Micro-level Coefficients Between Contexts at the 99th and 1st Percentiles of the Macro variables and Expected Values for Selected Years

Variable (99th p-tile - 1st p-tile)	Intercept	Fr's Educ.	Fr's Occ.	Female
Development (.942 - .194)	3.54	-.123	-.143	1.08
Educ. Expansion (.234 - .096)	.39	-.065	-.047	.09
Educ. Inequality (4.10 - 1.92)	.10	.122	.192	-.29
Cohort (1994 - 1904)[a]	2.52	-.036	-.360	.76
Diff.: Com. vs. Non-Com. (1950) .	.78	-.082	-.060	.54
Communist cohort (1990)	.86	-.082	-.212	.56
Non-communist cohort (1990)	1.12	-.016	-.160	.34
Difference	-.26	-.066	-.052	.22
Expected diff. between extremes[c] . .	6.34	-.346	-.742	2.21

Notes:

[a] This is the entire range of the variable. The coefficients show the expected increase in educational attainment between our earliest and our most recent cohorts.

[b] Ignoring other factors, the expected value of the difference between communist and non-communist cohorts in 1950 is simply the coefficient for communism in the models in Table 5. For 1990, the intercept for communist cohorts is $\gamma_{04}(90-50)+\gamma_{05}(1)+\lambda_{06}((90-50)*1)$ while for non-communist cohorts it is $\gamma_{04}(90-50)$ (see Eq. 4); the corresponding expected values the remaining micro variables are computed in the same way.

[c] Between the least developed (1st percentile) non-communist contexts with the smallest educational expansion (1st percentile) and the highest educational inequality (99th percentile) in 1900-04 and the most developed (99th percentile) non-communist contexts with the greatest educational expansion (99th percentile) and the smallest educational inequality (1st percentile) in 1990-94. There were no communist regimes in 1900-1904 or, with the exception of China, in 1990-1994.

APPENDIX A - SURVEYS USED

Country Study	Male Total	Female Total	Total
1. Australia:			
aus1965	1283	0	1283
aus1965l	642	0	642
aus1967	915	0	915
aus1967l	0	768	768
aus1973	3166	1773	4939
aus1973l	1708	2998	4706
aus1984	1458	1551	3009
aus1987	885	945	1830
aus1987i	824	839	1663
aus1989i	2216	2297	4513
aus1990	811	793	1604
aus1992i	1171	1026	2197
aus1999i	791	861	1652
2. Austria:			
aut1969	835	934	1769
aut1974p	649	936	1585
aut1982	21291	24898	46189
aut1986i	454	573	1027
aut1987i	431	541	972
aut1989i	886	1111	1997
aut1991i	430	554	984
aut1992i	438	589	1027
aut1999i	422	594	1016
aut2002e	1042	1215	2257
aut2004e	1041	1215	2256
aut2006e	1118	1287	2405
3. Belgium - Flemish:			
bef1971c	428	444	872
bef1975d	340	344	684
bef1976d	287	276	563
bef1991d	1421	1363	2784
bef2002e	664	569	1233
bef2004e	502	526	1028
bef2006e	557	571	1128
bef2008e	521	540	1061
4. Belgium - Walloon:			
bew1971c	274	313	587
bew1975f	387	436	823
bew1975v	892	502	1394
bew1976f	145	284	429
bew1976m	845	503	1348
bew1991f	886	841	1727
bew2002e	299	338	637
bew2004e	372	378	750
bew2006e	283	387	670
bew2008e	343	356	699
5. Brazil:			
bra1972	572	742	1314
bra1973	13622	0	13622
bra1982	13422	13628	27050
bra1988	15644	16597	32241
bra1999i	985	1015	2000
6. Bulgaria:			
bul1991j	667	737	1404
bul1993	2360	2543	4903
bul1999i	531	571	1102
bul2006e	546	854	1400
7. Canada - English:			
can1965m	564	0	564
can1965w	0	567	567
can1973	17383	18471	35854
can1982w	1334	1243	2577
can1984	1467	1910	3377
can1986	4220	5147	9367
can1994	4205	5037	9242
can1994l	1689	2262	3951
can1999i	494	298	792
8. Chile:			
chl1998l	3204	3962	7166
chl1999i	1314	1692	3006
9. China:			
chl2001	7088	0	7088
chn1996	3088	3002	6090
chn2008	1186	1814	3000
10. Croatia:			
cro2008e	635	838	1473
11. Cyprus:			
cyp1999i	501	499	1000
cyp2006e	474	521	995
12. Czech Republic:			
czr1984	1879	2226	4105
czr1991	1237	1609	2846
czr1991j	386	423	809
czr1992i	332	342	674
czr1993	2531	3090	5621
czr1998l	1302	1830	3132
czr1999i	831	1003	1834
czr2002e	644	707	1351
czr2004e	1414	1612	3026
czr2008e	984	1034	2018
13. Denmark:			
den1972	502	498	1000
den1972l	369	373	742
den1976	2565	2601	5166
den2002e	762	740	1502
den2004e	722	765	1487
den2006e	738	767	1505
den2008e	799	811	1610
14. England and Wales:			
eng1963e	817	998	1815
eng1964e	759	891	1650
eng1966e	906	988	1894

Country Study	Male Total	Female Total	Total
eng1969e	478	529	1007
eng1970e	454	531	985
eng1970ex	574	670	1244
eng1972	10309	0	10309
eng1974p	673	810	1483
eng1984w	899	871	1770
eng1987i	513	588	1101
eng2002e	838	966	1804
eng2004e	758	904	1662
eng2006e	942	1147	2089
eng2008e	970	1129	2099
15. Estonia:			
est1991j	406	594	1000
est2004e	818	1171	1989
est2006e	660	857	1517
est2008e	704	957	1661
16. Finland:			
fin1972	477	517	994
fin1972l	323	296	619
fin1975p	612	611	1223
fin1994k	891	812	1703
fin1998l	1493	1435	2928
fin2002e	960	1040	2000
fin2004e	948	1074	2022
fin2006e	919	977	1896
fin2008e	1077	1118	2195
17. France:			
fra1967e	947	1061	2008
fra1971c	999	1096	2095
fra1978	2170	2337	4507
fra1995e	1937	2141	4078
fra1997i	458	553	1011
fra1999i	1104	785	1889
fra2002e	679	824	1503
fra2004e	841	965	1806
fra2006e	930	1056	1986
fra2008e	941	1132	2073
fra1958e	469	548	1017
18. German Democratic Republic:			
gdr1991a	719	825	1544
gdr1991j	467	552	1019
gdr1992a	531	617	1148
gdr1994a	533	575	1108
gdr1996a	534	582	1116
gdr1998a	448	574	1022
gdr1999i	248	263	511
gdr2002e	534	564	1098
gdr2004e	521	498	1019
gdr2006e	509	531	1040
gdr2008e	486	481	967
19. German Federal Republic:			
ger1969e	899	1046	1945
ger1969f	536	622	1158
ger1972	485	573	1058

Country Study	Male Total	Female Total	Total
ger1975p	1057	1250	2307
ger1976z	972	1064	2036
ger1977z	903	1099	2002
ger1978c	917	1113	2030
ger1978x	821	1191	2012
ger1979x	866	1141	2007
ger1979z	913	1099	2012
ger1980a	1359	1596	2955
ger1980c	890	1049	1939
ger1980p	991	1104	2095
ger1980z	915	1082	1997
ger1982a	1340	1651	2991
ger1984a	1423	1581	3004
ger1985w	1154	680	1834
ger1986a	1445	1650	3095
ger1987i	614	783	1397
ger1988a	1356	1696	3052
ger1990a	1476	1575	3051
ger1991a	739	775	1514
ger1991j	923	914	1837
ger1992a	1132	1268	2400
ger1994a	1208	1134	2342
ger1996a	1204	1198	2402
ger1998a	1050	1162	2212
ger1999i	447	474	921
ger2002e	873	948	1821
ger2004e	860	991	1851
ger2006e	928	948	1876
ger2008e	964	820	1784
20. Greece:			
grc2002e	1113	1453	2566
grc2004e	1051	1355	2406
grc2008e	941	1131	2072
21. Hungary:			
hun1973	19208	21218	40426
hun1982	7271	8568	15839
hun1983	15352	16949	32301
hun1986	2686	3313	5999
hun1987i	1144	1462	2606
hun1989i	447	553	1000
hun1990e	424	557	981
hun1991i	431	569	1000
hun1991j	477	523	1000
hun1992	13500	15506	29006
hun1992i	578	672	1250
hun1993	2320	2657	4977
hun1998l	1192	1401	2593
hun1999i	520	688	1208
hun2002e	809	876	1685
hun2004e	646	852	1498
hun2006e	627	891	1518
hun2008e	702	842	1544
22. Iceland:			
ice2004e	271	299	570

Country Study	Male Total	Female Total	Total
23. India:			
ind1967e	1973	314	2287
ind1971	2100	537	2637
ind1971e	2748	2174	4922
24. Ireland:			
ire1973	2291	0	2291
ire1989	459	513	972
ire1991i	508	497	1005
ire2002e	944	1102	2046
ire2004e	986	1300	2286
ire2006e	790	944	1734
ire2008e	810	954	1764
25. Israel:			
isr1974j	5793	6116	11909
isr1974n	1661	1490	3151
isr1991	3833	4325	8158
isr1997j	496	541	1037
isr1997n	297	199	496
isr1999i	485	574	1059
isr1999n	83	63	146
isr2002e	1143	1351	2494
isr2008e	1140	1350	2490
26. Italy:			
ita1963	1568	0	1568
ita1968	1242	1258	2500
ita1972	888	953	1841
ita1975p	886	893	1779
ita1985	2465	2551	5016
ita1985x	2465	2551	5016
ita1987i	499	528	1027
ita1990e	753	747	1500
ita1992i	480	516	996
ita1993b	8916	9408	18324
ita1995b	9057	9472	18529
ita1997	5238	5671	10909
ita1997x	4631	5139	9770
ita1998b	7966	8241	16207
ita1998l	1352	1622	2974
ita1998s	22402	23728	46130
ita1999	3892	3555	7447
ita2000b	8627	8859	17486
ita2001	3281	3097	6378
ita2001e	1614	1595	3209
ita2002b	8207	8759	16966
ita2002e	548	659	1207
ita2003	18704	20656	39360
ita2004b	8052	8532	16584
ita2004e	751	778	1529
ita2005	1530	1366	2896
ita2005c	1197	761	1958
ita2006	16825	20688	37513
ita2006b	7485	8020	15505
ita2006e	1995	2021	4016
ita2008i	549	529	1078

Country Study	Male Total	Female Total	Total
ita2009i	581	489	1070
27. Japan:			
jap1955	2014	0	2014
jap1965	2077	0	2077
jap1967	942	1031	1973
jap1971	1300	1357	2657
jap1975	2724	0	2724
jap1991j	362	415	777
jap1999i	634	691	1325
jap2000	1318	1575	2893
28. Latvia:			
lat1999i	494	606	1100
lat2006e	781	1176	1957
lat2008e	747	1233	1980
29. Luxembourg:			
lux2002e	734	815	1549
lux2004e	846	789	1635
30. Malaysia:			
mal1967l	5457	0	5457
mal1976	1134	0	1134
mal1976l	0	1250	1250
31. Mexico:			
mex2006	6322	966	7288
mex2006l	102	5429	5531
32. Netherlands:			
net1958	543	664	1207
net1967m	620	0	620
net1967t	327	382	709
net1967w	0	620	620
net1970	984	854	1838
net1971	949	797	1746
net1971c	975	698	1673
net1974p	596	605	1201
net1975g	656	658	1314
net1976j	661	94	755
net1977	2093	2066	4159
net1977e	883	973	1856
net1979p	800	786	1586
net1981e	1073	1232	2305
net1982e	724	817	1541
net1982n	1306	1318	2624
net1982u	497	256	753
net1985o	2004	2016	4020
net1986e	764	865	1629
net1986l	1842	2198	4040
net1987c	934	1056	1990
net1987i	777	861	1638
net1987j	397	398	795
net1987s	445	457	902
net1988o	2237	2227	4464
net1990o	2226	2212	4438
net1990s	1238	1146	2384
net1991j	950	833	1783
net1992f	902	898	1800

Country Study	Male Total	Female Total	Total
net1992o	2250	2286	4536
net1992t	1695	1540	3235
net1994e	887	925	1812
net1994h	440	595	1035
net1994o	2278	2258	4536
net1995h	1019	1014	2033
net1995s	990	1029	2019
net1995y	680	641	1321
net1996	413	377	790
net1996c	813	1065	1878
net1996o	2293	2266	4559
net1996y	355	435	790
net1998	535	390	925
net1998e	1022	1079	2101
net1998f	1000	1029	2029
net1998o	2380	2400	4780
net1999	1431	1080	2511
net1999a	5565	6024	11589
net1999i	851	767	1618
net2000f	779	782	1561
net2000s	502	506	1008
net2002e	1042	1322	2364
net2003	3420	4741	8161
net2003f	1063	1111	2174
net2004e	783	1098	1881
net2005i	995	828	1823
net2006e	868	1021	1889
net2006i	984	935	1919
net2008e	818	960	1778
net2008i	1344	1499	2843
net2008v	701	853	1554
33. New Zealand:			
nze1976	1575	1658	3233
nze1990	1041	1061	2102
nze1992i	612	614	1226
nze1999i	512	560	1072
34. Nigeria:			
nig1971	1402	767	2169
nig1973	510	510	1020
35. Northern Ireland:			
nir1968	597	694	1291
nir1971p	5416	0	5416
nir1973	2415	0	2415
nir2002e	350	430	780
nir2004e	32	34	66
nir2006e	26	44	70
nir2008e	35	44	79
36. Norway:			
nor1957e	762	784	1546
nor1972	496	509	1005
nor1972l	389	367	756
nor1982w	1272	1260	2532
nor1992i	780	758	1538
nor1999i	631	637	1268
nor2002e	1103	933	2036
nor2004e	914	846	1760
nor2006e	891	859	1750
nor2008e	807	742	1549
37. Philippines:			
phi1968	9922	10627	20549
phi1973	8141	8251	16392
phi1999i	600	600	1200
38. Poland:			
pol1972	39317	35234	74551
pol1987	994	900	1894
pol1987i	1841	2102	3943
pol1988	2702	3152	5854
pol1991	870	965	1835
pol1991i	505	558	1063
pol1991j	696	846	1542
pol1992g	739	908	1647
pol1992i	734	902	1636
pol1993g	727	922	1649
pol1994	1680	1840	3520
pol1994g	731	878	1609
pol1994z	959	1168	2127
pol1995g	717	886	1603
pol1995l	1431	1569	3000
pol1997g	1051	1351	2402
pol1999g	984	1298	2282
pol1999i	479	656	1135
pol2002e	1032	1078	2110
pol2004e	833	883	1716
pol2006e	815	906	1721
pol2008e	764	855	1619
39. Portugal:			
por1999i	533	611	1144
por2002e	1260	1762	3022
por2004e	821	1231	2052
por2006e	863	1359	2222
por2008e	926	1441	2367
40. Quebec (Canada):			
que1960	470	530	1000
que1962	469	529	998
que1965m	174	0	174
que1965w	0	170	170
que1973	4367	4645	9012
que1977	2782	1049	3831
que1986	3178	3845	7023
que1994	1175	1459	2634
que1994l	734	975	1709
que1999i	120	53	173
41. Romania:			
rom2006e	1020	1119	2139
rom2008e	966	1180	2146
42. Russia:			
rus1991j	784	948	1732
rus1992i	886	1097	1983

Country Study	Male Total	Female Total	Total
rus1992w	898	1251	2149
rus1993	2005	2997	5002
rus1998	1998	2820	4818
rus1999i	778	927	1705
rus2006e	983	1454	2437
rus2008e	989	1523	2512
43. Scotland:			
sco1963e	81	113	194
sco1964e	88	90	178
sco1966e	87	101	188
sco1969e	17	15	32
sco1970e	93	90	183
sco1970ex	55	56	111
sco1974e	556	622	1178
sco1975	4887	0	4887
sco1987i	56	55	111
sco1999i	40	40	80
sco2002e	82	102	184
sco2004e	73	94	167
sco2006e	111	124	235
sco2008e	67	97	164
44. Slovakia:			
slo1984	940	1126	2066
slo1991j	168	200	368
slo1992i	196	227	423
slo1993	2412	2508	4920
slo1999i	513	569	1082
slo2004e	743	734	1477
slo2006e	841	896	1737
slo2008e	685	1116	1801
45. Slovenia:			
sln1967t	280	320	600
sln1968	1009	1073	2082
sln1973	1089	1004	2093
sln1980	911	844	1755
sln1989	960	902	1862
sln1991i	978	1085	2063
sln1991j	602	769	1371
sln1992i	530	519	1049
sln1998	491	516	1007
sln1999i	490	516	1006
sln2001	486	610	1096
sln2002e	723	795	1518
sln2004e	648	762	1410
sln2006e	667	809	1476
sln2008e	596	690	1286
46. South Africa:			
saf1991	4610	4104	8714
47. Spain:			
spa1990	1143	1257	2400
spa1991	5745	6255	12000
spa1992	5758	6242	12000
spa1993	5789	6211	12000
spa1994	1739	1861	3600

Country Study	Male Total	Female Total	Total
spa1995	2311	2489	4800
spa1999i	595	616	1211
spa2002e	818	911	1729
spa2004e	849	813	1662
spa2006e	902	974	1876
spa2008e	1222	1354	2576
48. Suriname:			
sur1993	173	169	342
sur1993l	97	108	205
sur1993s	166	141	307
49. Sweden:			
swe1960	770	833	1603
swe1968	2992	2929	5921
swe1972	497	508	1005
swe1972l	383	355	738
swe1974	2799	2815	5614
swe1980w	669	476	1145
swe1981	2819	2794	5613
swe1991	2683	2623	5306
swe1992i	381	368	749
swe1999i	542	608	1150
swe2002e	1014	981	1995
swe2004e	981	967	1948
swe2006e	951	975	1926
swe2008e	918	912	1830
50. Switzerland:			
swi1976p	641	649	1290
swi1987i	596	378	974
swi1999i	574	684	1258
swi2002e	981	1059	2040
swi2004e	948	1193	2141
swi2006e	815	988	1803
swi2008e	822	997	1819
swi1972	895	1022	1917
51. Taiwan:			
tai1970	1173	1049	2222
tai1970l	809	0	809
tai1990a	1242	1289	2531
tai1990b	1191	1340	2531
tai1991a	1225	1263	2488
tai1991b	597	542	1139
tai1992a	1099	1278	2377
tai1992b	754	654	1408
tai1994a	870	983	1853
tai1994b	824	1038	1862
tai1995a	1073	1020	2093
tai1995b	1068	1013	2081
tai1996a	1008	916	1924
tai1996b	1452	1379	2831
tai1997a	1290	1306	2596
tai1997b	1444	1391	2835
tai1997s	813	904	1717
tai1999a	1029	919	1948
tai1999b	962	963	1925

Country Study	Male Total	Female Total	Total
tai2000a	993	967	1960
tai2000b	952	943	1895
tai2001a	1000	979	1979
tai2002a	971	1021	1992
tai2002b	977	1006	1983
tai2004a	915	866	1781
tai2005a	1073	1073	2146
tai2005b	1106	1065	2171
52. Turkey:			
tur1978	4431	0	4431
tur2004e	826	1029	1855
tur2008e	1127	1289	2416
53. Ukraine:			
ukr2004e	747	1283	2030
ukr2006e	776	1226	2002
ukr2008e	690	1155	1845
54. United States:			
usa1947	1416	1505	2921
usa1956e	787	975	1762
usa1958e	667	783	1450
usa1960e	535	646	1181
usa1962o	20329	0	20329
usa1964e	806	1028	1834
usa1966e	572	719	1291
usa1967	1232	1317	2549
usa1968e	724	949	1673
usa1970e	728	966	1694
usa1972g	807	806	1613
usa1973g	701	803	1504
usa1973o	33613	0	33613
usa1974g	691	793	1484
usa1974p	736	983	1719
usa1975g	670	820	1490
usa1976g	669	830	1499
usa1977g	693	837	1530
usa1978g	643	889	1532
usa1980g	641	827	1468
usa1980w	875	885	1760
usa1982g	779	1081	1860
usa1983g	690	909	1599
usa1984g	598	875	1473
usa1985g	688	846	1534
usa1986g	621	849	1470
usa1987g	778	1041	1819
usa1987i	664	900	1564
usa1988g	638	843	1481
usa1989g	660	877	1537
usa1990g	604	768	1372
usa1990w	1129	1359	2488
usa1991g	636	881	1517
usa1991j	627	787	1414
usa1993g	685	921	1606
usa1994g	1290	1702	2992

Country Study	Male Total	Female Total	Total
usa1994l	1437	1608	3045
usa1996g	1285	1619	2904
usa1998g	1232	1600	2832
usa1999i	551	721	1272
usa2000g	1229	1588	2817
usa2002g	1228	1537	2765
usa2004g	1280	1532	2812

APPENDIX B - COEFFICIENTS FOR MODEL 5 WITH ALTERNATIVE MEASURES OF DEVELOPMENT AND EDUCATIONAL EXPANSION

	Alternative Macro Indicator			
	G[a]	A_1	A_2	E_2
Micro intercept:				
Macro intercept	9.95 (.000)	10.16 (.000)	10.05 (.000)	10.49 (.000)
(D) Development	.085 (.000)[b]	.050 (.000)	.051 (.000)	-.940 (.000)
(E) Educational expansion	5.89 (.000)	5.72 (.000)	5.09 (.011)	.974 (.000)
(I) Educational inequality .	-.157 (.076)	.006 (.930)	-.029 (.793)	.217 (.000)
(T) Cohort	.059 (.000)	.035 (.000)	.039 (.000)	.002 (.351)
(C) Communist regime . . .	.334 (.020)	.333 (.006)	.901 (.000)	-.052 (.473)
(C*T) Comm.*Cohort	-.017 (.013)	-.029 (.000)	-.028 (.000)	-.012 (.001)
Joint sig. of C & C*T	(.012)	(.000)	(.000)	(.000)
Father's education:				
Macro intercept	.338 (.000)	.357 (.000)	.352 (.000)	.339 (.000)
(D) Development	-.0075 (.000)	.0000 (.903)	-.0006 (.153)	.0800 (.006)
(E) Educational expansion	-.601 (.000)	-.580 (.000)	-.600 (.000)	-.042 (.000)
(I) Educational inequality .	.057 (.000)	.076 (.000)	.073 (.000)	.033 (.000)
(T) Cohort	-.0000 (.958)	-.0022 (.000)	-.0017 (.000)	-.0001 (.747)
(C) Communist regime . . .	-.092 (.000)	-.062 (.000)	-.070 (.000)	-.051 (.000)
(C*T) Comm.*Cohort	.0003 (.700)	.0005 (.423)	.0005 (.490)	-.0004 (.412)
Joint sig. of C & C*T	(.000)	(.000)	(.000)	(.000)
Father's ISEI:				
Macro intercept	.522 (.000)	.516 (.000)	.526 (.000)	.518 (.000)
(D) Development	-.0058 (.000)	-.0023 (.000)	-.0023 (.000)	-.0819 (.042)
(E) Educational expansion	-.548 (.000)	-.573 (.000)	-.392 (.010)	-.024 (.000)
(I) Educational inequality .	.098 (.000)	.084 (.000)	.092 (.000)	.077 (.000)
(T) Cohort	-.0045 (.000)	-.0044 (.000)	-.0045 (.000)	-.0036 (.000)
(C) Communist regime . . .	-.049 (.030)	-.024 (.225)	-.066 (.002)	-.031 (.128)

continued

	Alternative macro indicator			
	G	A_1	A_2	E_2
(C*T) Comm.*Cohort	.0004 (.733)	-.0002 (.812)	.0005 (.577)	.0002 (.802)
Joint sig. of C & C*T	(.017)	(.140)	(.001)	(.159)
Female:				
Macro intercept	-.509 (.000)	- .546 (.000)	-.587 (.000)	-.468 (.000)
(D) Development	.0512 (.000)	.0114 (.000)	.0081 (.000)	1.296 (.000)
(E) Educational expansion	1.406 (.002)	1.335 (.008)	1.246 (.021)	.0192 (.334)
(I) Educational inequality .	-.160 (.000)	-.231 (.000)	-.239 (.000)	-.109 (.003)
(T) Cohort	.0114 (.000)	.0146 (.000)	.0175 (.000)	.0076 (.000)
(C) Communist regime . . .	.549 (.000)	.398 (.000)	.504 (.000)	.509 (.000)
(C*T) Comm.*Cohort	-.0064 (.145)	-.0115 (.000)	-.0101 (.001)	-.0075 (.076)
Joint sig. of C & C*T	(.000)	(.000)	(.000)	(.000)

Notes:

[a] G = gross national product per capita (in 000's) (from Haber and Menaldo 2011); A_1 = proportion of labor force not engaged in agriculture or retail sales (calculated from our micro data); A_2 = proportion of labor force not engaged in agriculture (from CNTS Archive 2002); and E_2 = the mean years of schooling of the previous cohort within the same nation (calculated from our micro data). See text for additional detail.

[b] Shaded bars indicate the coefficients associated with alternative measurements.