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INTERGENERATIONAL ECONOMIC MOBILITY AROUND THE WORLD

The Inheritance of Educational Inequality: International Comparisons and Fifty-Year Trends

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The Inheritance of Educational Inequality: International Comparisons and Fifty-Year Trends*

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Abstract

This paper estimates 50-year trends in the intergenerational persistence of educational attainment for a sample of 42 nations around the globe. Large regional differences in educational persistence are documented, with Latin America displaying the highest intergenerational correlations, and the Nordic countries the lowest. We also demonstrate that the global average correlation between parent and child's schooling has held steady at about 0.4 for the past fifty years.

KEYWORDS: intergenerational mobility, educational persistence

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The correlation between the socioeconomic status of parents and their adult offspring is positive and significant, in both the statistical and practical senses, in virtually every society for which estimates are available, and for many different measures of status. These correlations quantify the rate of transmission of inter-personal inequality from one generation to the next, and are often interpreted as measures of a society's failure to provide equality of opportunity to children from differing family backgrounds.[1] While many studies of status persistence, or its opposite, status mobility,[2] are available for the United States and Europe, our understanding of international differences and trends in this social statistic is far from complete, especially for the developing countries. This paper seeks to fill that gap by providing estimates of 50-year trends in two simple measures of status persistence – the coefficient from a regression of children's schooling against that of their parents, and the correlation between the two – for a sample of 29 developing and formerly communist economies. Comparable estimates are also presented for twelve Western European countries and the United States.

Early estimates of generational persistence in schooling in the U.S. include Spady (1967), Bowles (1972), Hauser and Featherman (1976), and Blake (1985). Recent comparisons among rich nations include Couch and Dunn (1997) on the U.S. and Germany, and de Broucker and Underwood (1998), who look at eleven nations. Studies of developing and transition economies include analyses for Panama (Heckman and Hotz, 1986), Malaysia (Lillard and Willis, 1994), South Africa (Thomas, 1996; Hertz, 2001), Brazil (Pastore and Zylberstajn, 1996), Brazil, Colombia, Mexico and Peru (Behrman, Gaviria, and Székely, 2001), Mexico (Binder and Woodruff, 2002), China (Sato and Shi, 2007), and six Eastern Bloc nations (Ganzeboom and Nieuwbeerta, 1999). Our contribution is to extend this list by creating comparable estimates for 42 countries, and improved estimates of long-run trends, so that we may begin to draw conclusions about global patterns in the inheritance of educational status.

[1] The validity of this interpretation has been challenged by Jencks and Tach (2005, page 2), who argue that "the size of the correlation between the economic status of parents and their children is not a good indicator of how close a society has come to equalizing opportunity," and describe situations in which increases in opportunity may coincide with increased persistence. We acknowledge that status persistence is an imperfect measure of the lack of opportunity, but hold that large differences in persistence between nations and over time are still informative.

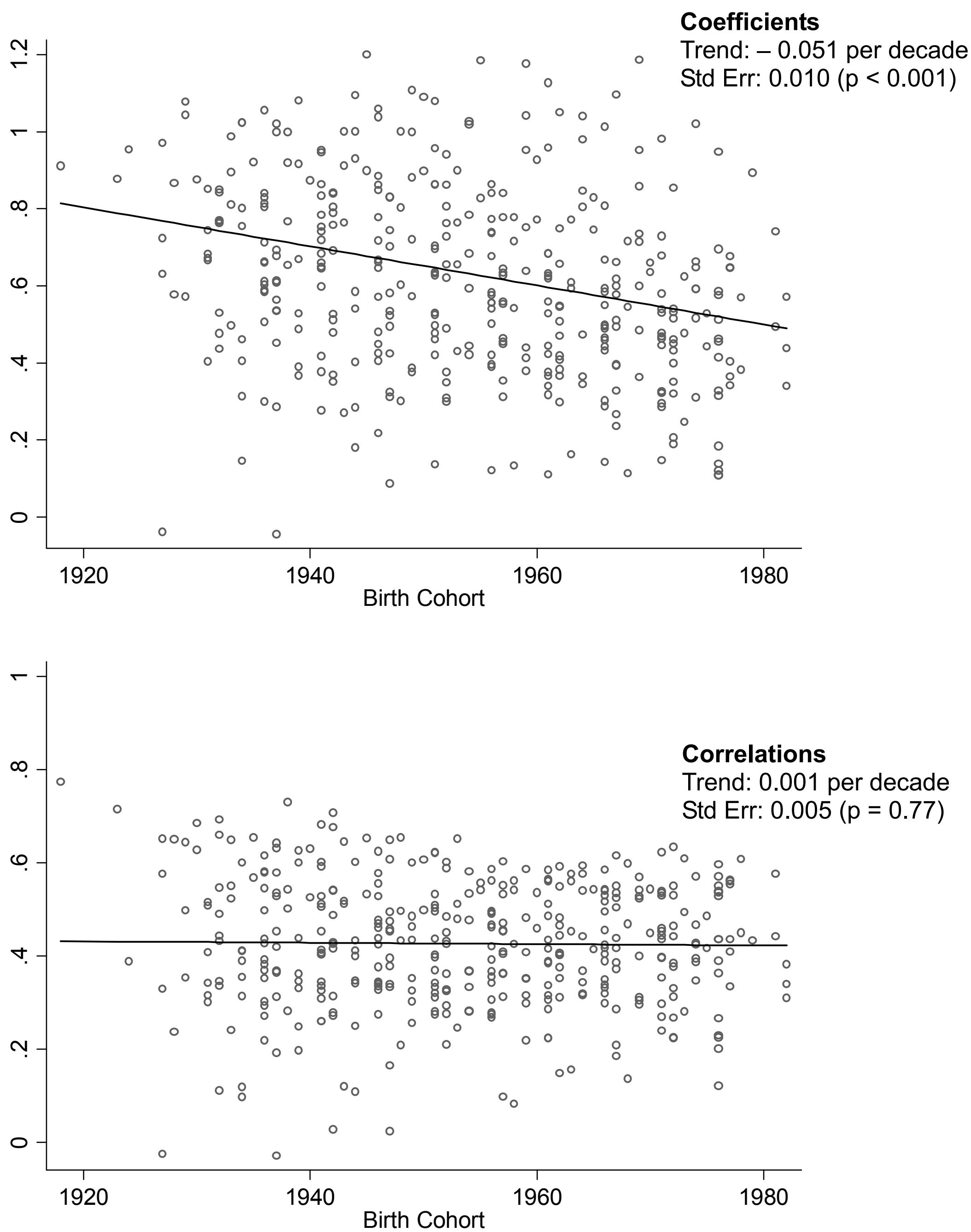
[2] It is common among economists to refer to both intergenerational regression coefficients and correlation coefficients as inverse measures of intergenerational mobility (Solon, 1999). We avoid this terminology because "mobility" may be defined in many ways, some of which are not closely related to these two measures of persistence. In particular, the popular notion of "upward mobility" is distinct from what we quantify here. We will use the terms "persistence," "transmission," and "inheritance" interchangeably, and always in a descriptive sense, with no attempt to analyze the biological and social processes that generate the socioeconomic resemblance between parents and children.

Our principal finding is both methodological and substantive in nature, and is illustrated in Figure 1. We demonstrate that for most countries, and for the sample as a whole, the *regression coefficient* of parents' education as a predictor of schooling in the next generation fell substantially over 50 years, indicating a long-run decrease in this basic measure of the intergenerational persistence of educational inequality. However, we do not observe any such trend in the *correlation* between parent and child schooling, which is a standardized measure of persistence, and which has risen in as many countries as it has fallen. It thus appears that a one-year difference in parents' education now corresponds to a smaller difference, on average, in the expected value of their children's schooling than it previously did, this being a statement about a weakened statistical association, not a diminished causal connection. Yet because the variance of parents' education has also increased relative to the variance of schooling in the second generation, the intergenerational correlation, or the R^2 from the parent-child schooling regression, has not fallen. On average for the countries in our sample, a one-standard-deviation difference in parental education corresponds to a schooling difference of about 0.4 standard deviations in the next generation, and this figure has held steady for half a century.[3] Otherwise put, around the world, parents' schooling, by itself, explains as much of the variance of children's schooling as ever.

Strikingly, the seven Latin American nations in our sample occupied the top seven positions among the 42 countries considered when ranked by their parent-child schooling correlations (see Table 2). These seven had an average parental schooling correlation, for adults between the ages of 20 and 69, of 0.60 (Table 3), compared to 0.41 for eight Eastern Bloc nations (Table 6), 0.39 for ten Asian nations (Table 4), 0.39 again for 13 Western nations (Table 7), and 0.36 for a small sample four African countries (Table 5). The Nordic nations stood out for displaying less persistence, on average, than the non-Nordic high-income nations, and less than the formerly communist countries.

[3] Standardized persistence is measured using Pearson's correlation coefficients for each cohort and country. These are generally close to Spearman's rank-order correlations, which sociologists often use to measure "exchange mobility," in part because education data are not plagued by large outliers. Yet the two are not identical: a given change in rank may amount to a larger or smaller change in units of the standard deviation, depending on the details of the cumulative distribution. Nor do we find empirical support for the claim that the normal curve, or indeed any other familiar density function, is shared by all educational distributions. Thus there is no invariant mapping from units of the standard deviation to educational percentiles. As an example, consider that in Ethiopia one can find 92 percent of parents within not two but rather one-half of a standard deviation from the mean (the mean number of years of education being 0.12).

FIGURE 1
INTERGENERATIONAL EDUCATIONAL REGRESSION COEFFICIENTS AND CORRELATIONS FOR 42 COUNTRIES, BY BIRTH COHORT



EDUCATION, INCOME, AND OCCUPATIONAL PRESTIGE

Educational attainment is, of course, not the only useful measure of social status. In this section we describe its relation to other commonly studied metrics, (namely, long-run average income or earnings, as well as occupation-based indices) and discuss their relative strengths and weaknesses for the study of intergenerational status transmission in developing countries.

Various composite measures of socioeconomic prestige, or class standing, may be found in the sociological literature, beginning with Otis Dudley Duncan's famous socioeconomic index (Duncan, 1961). The subsequent debate as to their relative merits and their ability to capture what we want to know about the process of social stratification has been a long and complicated one. One of the key disagreements has been over the weight that should be attached to each of the component parts of these indices, which usually include occupation-specific measures of earnings, average education, and the job's perceived desirability (Hauser and Warren, 1997), and the upshot would seem to be that there is no manifestly best measure of occupational status.

The relationship between such composite status measures and a simple count of the number of years of schooling is nicely summarized by Ganzeboom, Treiman, and Ultee (1991, page 284):

> The answer to the question of the extent to which educational attainment promotes social mobility thus turned out to be compound: Respondent's occupational status is more related to [own] education than to father's occupation, and most of the effect of education is independent of social origins, so the main role of education is to promote social mobility; but at the same time a majority of what social reproduction there is is transmitted through education, so education is also the main vehicle of social reproduction.

The validity of this conclusion as a statement about the relative magnitudes of the true causal effects of each aspect of family background is open to question: Bowles (1972), for example, demonstrates that it is easy to overestimate the independent causal impact of own schooling, and to underestimate the full effect of family background. Indeed, we could observe a strong correlation between schooling in two generations even if parental schooling *per se* had no independent causal impact on own schooling, as will be shown below. Yet this paragraph is correct in stressing that one's own education is a key correlate of occupational status, prestige, and income, and that this correlation is driven by two orthogonal components, one that is strongly related to family background (including parental education), and another that is not. The centrality of education as a marker of status in both generations is thus clear.

Parent-child schooling correlations are also related to parent-child correlations in long-run average (or "permanent") income, or earnings, with which economists have primarily been concerned (Solon, 1999). Moreover, this relationship could obtain even if parental schooling had no direct causal effect on children's education. Consider the standard intergenerational income equation:

$$y_1 = \alpha + \beta_y y_0 + \varepsilon \,, \tag{1}$$

where y_0 and y_1 are long-run average log incomes of the parents and children, and β_y is thus the intergenerational income elasticity, a widely-used measure of income persistence. Next, let β_s denote the parameter from a regression of own education (s_1) against parental education (s_0). The connection between the income and education regression coefficients can then be illustrated by adding a model of income determination in each generation. For parents we use a Mincer equation, in which long-run average log income is a function of education alone:

$$y_0 = \alpha_0 + p_0 s_0 + u_0 \,, \tag{2}$$

with p_0 being the percentage change in expected income associated with an extra year's schooling, and u_0 being a well-behaved error term, which represents the effects of any unobserved income-generating abilities that are orthogonal to schooling. In the second generation, we allow for a direct effect of own education (s_1), as well as direct effects of both parents' education and their incomes:[4]

$$y_1 = \alpha_1 + p_1 s_1 + \lambda s_0 + \delta y_0 + u_1 \,. \tag{3}$$

A final parameter is needed to capture the presumably positive covariance between the "income-generating ability" term in the parents' income equation (u_0) and the child's level of schooling (s_1). Such abilities could be transmitted either biologically or socially.

$$\phi = Cov(s_1, u_0) / Var(u_0) \tag{4}$$

The relation between the intergenerational income elasticity (β_y) and the coefficient from a regression of education on parents' education (β_s) is then:

$$\beta_y = \delta + \frac{(p_1 \beta_s + \lambda)}{p_0} R^2 + p_1 \phi (1 - R^2) \,, \tag{5}$$

where R^2 is the share of variance explained in the parents' income equation.

[4] Because we condition on parental income and education, p_1 is not the standard Mincerian estimator. Note also that we omit the effects of grandparents' income and education from the parents' equation. In structural terms this is a misspecification, but in structural terms the entire model is obviously misspecified: the model must be viewed descriptively, not structurally.

These parameters may be estimated for the U.S. using data from the Panel Study of Income Dynamics, a thirty-year study that allows for the measurement of both education and long-run average income in two generations. Using the data-set employed by Hertz (2005), we find that $\beta_y = 0.53$ and $\beta_s = 0.34$, and that they are related as follows (with β_y and β_s in boldface).

$$[6] \qquad \mathbf{0.53} = 0.34 + \frac{(0.13 * \mathbf{0.34} + 0.01)}{0.12} 0.31 + 0.13 * 0.59(1 - 0.31)$$

$$\text{or:} \qquad 0.53 = 0.34 + \quad 0.14 \quad + 0.05.$$

This reveals that the various channels that appear to operate via education, namely, the overall relation between parents' and children's schooling (β_s), the connection between parental education and children's income (λ), and the additional effect of "parental economic ability" (ϕ) on their children's schooling, together account for little more than a third of the intergenerational income elasticity (0.14 + 0.05 = 0.19). This leaves the "direct effect" of parents' income to account for the remainder, via $\delta = 0.34$ (coincidentally the same value as β_s).

Geographic and temporal differences in income persistence will thus be positively but far from perfectly correlated with the corresponding differences in educational persistence. Their trends and their national rankings might differ if the other parameters in equation [5] differ across countries or years.

Education-, income-, and occupational status-based measures all perform similar descriptive functions: they quantify the intergenerational association between conceptually and empirically distinct, but closely related, scalar measures of long-run socioeconomic status. There are a number of practical reasons why we chose to work with education alone. First among these is that education appears on virtually every household or labor force survey. However, only a minority of these also collect information on the education of one's parents when those parents are not household members, and it is this subset that we will work with.[5] To date we have identified ten suitable surveys from counties in Asia, seven from Latin America, four from Africa, and eight from the Eastern Bloc (see Table 1). We also present results for a selection of thirteen Western capitalist economies, including the U.K., the U.S., and four of the five Nordic nations.

[5] A number of analyses (e.g. Hertz, Meurs, and Selcuk, 2007; Behrman, Birdsall, and Székely, 2000) have measured educational persistence by looking at the grade-for-age progression of school-aged children who still live at home. This approach can provide valuable insight into more recent developments, and it obviates the need for a separate (non-resident) parents' education question, thereby increasing the number of datasets that can be used. We opt to take the longer-run view, based on intergenerational comparisons among adults.

A second advantage is that the number of years of formal schooling is a reasonably unambiguous concept, although not all surveys measure it in great detail or with great precision, particularly for parents. Yet, recall-based information for parents' education is likely to be of higher quality than for income or wealth. Lacking information on the reliability of measured education in each country, we are forced to assume that reliability is roughly the same for all, and that it does not vary systematically with the age of the respondent. These assumptions are implicit in virtually all comparative work on this question.

Third, and of particular importance, formal schooling is usually fixed once one reaches adulthood, making it possible to reconstruct a long time-series of cohort-specific persistence estimates from data on a single representative cross-section of adults. This is much harder where long-run income is concerned, because of the volatile and age-dependent nature of income over the lifecycle (Jenkins, 1987; Haider and Solon, 2006). Indeed, Hertz (2007b) demonstrates that time trends in the persistence of permanent income cannot be estimated with much precision even with the thirty years of data contained in the Panel Study of Income Dynamics.[6]

One key limitation of education as an index of status, however, is its relative coarseness. In particular, in many countries a large number of people, or their parents, have no formal schooling, and so are all assigned the same status. This is a limitation we choose to live with: in particular, we do not treat these zero values as censored measures of unobserved human capital, but rather as the correct value of the outcome of interest. A second objection is that in some countries (for example, South Africa) those listed as having no education actually have higher average incomes than those with a year or two of schooling (Hertz, 2001). More generally, the relationship between education and income may be non-linear. While this weakens the correlation between education and income, or occupational status, it by no means eliminates it, nor does it render educational attainment uninteresting as a measure of status in its own right.

[6] Aaronson and Mazumder (2007) address this problem by means of a two-sample instrumental variables approach (TSIV), with state of residence as the instrument for parents' income, using the large cross-sectional datasets from the decennial U.S. censuses of 1940 to 2000. Grawe (2004) also uses TSIV, with parental education as the instrument, to estimate intergenerational income persistence in a number of developing countries. Dunn's (2007) TSIV analysis for Brazil is the only study of which we are aware that estimates a time trend in earnings persistence in a developing country.

SAMPLE AND VARIABLE DEFINITIONS

The countries appearing in this analysis are listed in Table 1. The data are drawn from World Bank Living Standards Measurement Surveys (LSMS), similar household surveys conducted by national statistical agencies, and country surveys affiliated with the European Social Survey (ESS), the International Social Survey Program (ISSP), and the International Adult Literacy Survey (IALS). Detailed sources appear in Appendix 2.

The survey years are the most recent available that include information on parental education; they fall between 1985 and 2004, with most having been conducted between 1994 and 2004. The age criterion of 20 to 69 then yields the dates of birth shown, which run from 1916 to 1984. Sample sizes range from 149,477 for Brazil to 1,047 for the Philippines; a total of more than 390,000 individual parent-child pairs are represented. We also report the size of the smallest five-year birth cohort for each country. For some countries, both ESS (or ISSP) and IALS surveys were available; if so, the IALS surveys were usually preferred, due to their larger sample sizes, and to some limitations in the coding of schooling in the ESS and ISSP, described below. In two cases (Chile and the U.S.) the IALS and ISSP surveys were comparable, and were pooled.

Information on the education of non-resident parents is usually available only for heads of household and their spouses; for some of the Latin American surveys it is available for all household members. In addition, if all household members are surveyed we are able directly to observe parental education for adults who still live with their parents, and these few extra observations are included to bolster the sample size of the youngest cohorts.[7] For the ESS and ISSP surveys, the respondent is usually a randomly selected family member over the age of 16 or 18. Each survey design should in principle yield an approximately representative sample of adults between the ages of 20 and 69.

Many of the surveys require weights to render them representative of the national non-institutionalized population, and these weights are used in all calculations. Note, however, that the surveys from Bangladesh, Belgium, China, Ethiopia, Indonesia, Norway, and South Africa are not fully national in scope (see notes to Table 1). Note also that many of the European surveys extend only to age 64 or 65, as opposed to 69; for these countries, there are nine cohorts rather than ten, and the oldest cohort is designated number two.

[7] In some surveys, the ID codes of the mother and father of each member are given, if they are resident, allowing parents' education to be calculated. In others, those listed as sons or daughters of the household head were assumed to be the children of both the head and the spouse, if the latter was present. Thus, some of our children are not the biological offspring of both of their "parents." Yet because we are not interested in questions of genetic mechanisms, we see no need to exclude step-parents and the like.

TABLE 1
COUNTRIES, DATES, SAMPLE SIZES, EDUCATION CODING, ENROLLMENT RATES, AND YEARS OF SCHOOLING

	Dates		Sample Sizes		Education Coding				Share Enrolled		Average Years of Education			
		Birth			Parents		Children		Ages	Ages	Parents		Children	
ASIA	Svy Year	Years	Total	Min	Min	Max	Min	Max	20-24	25-29	Cohort 1	Cohort 10	Cohort 1	Cohort 10
Bangladesh (Matlab)	1996	1927-76	5,439	287	0	17	0	17	0.30	0.08	0.8	2.9	2.1	5.9
China (Rural)	1995	1926-75	1,909	67	0	16	0	16	0.07	0.00	0.6	4.7	1.5	7.9
East Timor	2001	1932-81	4,027	92	0	19	0	19	0.10	0.01	0.1	1.1	0.3	6.2
Indonesia (Partial)	2000	1931-80	21,663	820	0	21	0	21	0.11	0.00	0.8	5.1	2.6	9.1
Malaysia*	1988	1939-68*	4,826	313	0	16	0	20	0.04	0.01	2.5	3.6	5.3	9.3
Nepal	2003	1934-83	12,708	506	0	18	0	18	0.13	0.02	0.1	1.4	0.7	5.5
Pakistan	1991	1922-71	14,696	480	0	21	0	21	0.09	0.01	0.2	1.6	0.9	4.3
Philippines	1999	1930-79	1,047	29	0	14	0	24	0.23	0.06	3.8	7.5	9.0‡	10.7
Sri Lanka	2000	1931-80	17,586	620	0	18	0	18	0.07	0.02	1.5	5.6	4.2	9.1
Vietnam	1998	1929-78	12,554	564	0	20	0	20	na	na	0.9	5.4	3.5	7.3
									Regional Average		***1.1***	***3.9***	***4.7***	***7.5***
AFRICA														
Egypt	1997	1928-77	6,815	310	0	16	0	16	0.19	0.02	0.6	3.4	2.1	8.2
Ethiopia (Rural)	1994	1925-74	3,382	120	0	13	0	16	0.09	0.03	0.0	0.3	0.1	3.1
Ghana	1998	1929-78	10,735	403	0	21	0	19	0.20	0.04	0.6	4.8	2.5	8.1
S.Africa (KZN**)	1998	1929-78	4,212	145	0	16	0	16	0.40	0.06	0.6	3.6	2.3	8.0
									Regional Average		***0.5***	***3.0***	***1.8***	***6.9***
LATIN AMERICA														
Brazil	1996	1927-76	149,452	5,449	0	17	0	17	0.22	0.07	1.5	3.7	3.1	7.0
Chile	1998-99	1930-79	3,975	138	0	18	0	21	na	na	4.3	7.7	7.1	11.6
Colombia†	1997	1928-77	17,479	757	0	16	0	19	0.00†	0.00†	3.0	4.4	4.2	7.3
Ecuador	1994	1925-74	9,745	289	0	17	0	21	0.22	0.08	3.4	5.3	3.8	9.0
Nicaragua	1998	1929-78	8,263	230	0	17	0	17	0.17	0.08	1.3	3.0	2.5	5.8
Panama	2003	1934-83	13,046	580	0	18	0	21	0.27	0.14	2.4	6.7	4.9	9.7
Peru	1985	1916-65	11,808	431	0	17	0	17	0.26	0.12	2.1	4.4	3.1	8.8
									Regional Average		***2.6***	***5.0***	***4.1***	***8.5***

* Missing parental education data for 1919-1938; remaining sample appears biased in favor of better-educated families.
** Survey covers those designated as African and Asian in KwaZulu-Natal only.
† Education was not recorded for those still enrolled. ‡ Small cell size explains anomalous result; the next younger cohort averaged 7.6 years.

TABLE 1, CONTINUED

	Dates		*Sample Sizes*		*Education Coding*				*Share Enrolled*		*Average Years of Education*			
		Birth			Parents		Children		Ages	Ages	Parents		Children	
EASTERN BLOC	Svy Year	Years	Total	Min.	Min	Max	Min	Max	20-24	25-29	Cohort 1|2	Cohort 10	Cohort 1|2	Cohort 10
Czech Republic*	1998	1933-78	2,821	247	***4***	17	0	21	na	na	9.3	12.0	11.8	12.4
Estonia	2004	1935-84	1,461	131	***2***	15	0	21	0.36	0.09	6.2	10.9	11.4	13.0
Hungary*	1998	1933-78	2,303	214	0	19	0	21	na	na	6.7	11.4	9.6	12.0
Kyrgyzstan	1998	1929-78	7,438	354	0	18	0	18	0.10	0.03	1.5	10.0	6.6	10.6
Poland*	1994	1930-74	2,596	219	0	17	0	21	na	na	5.5	9.6	8.6	11.9
Slovakia	2004	1935-84	1,131	54	***4***	17	0	21	0.31	0.08	8.9	12.0	11.2	12.4
Slovenia*	1998	1933-78	2,546	201	0	20	0	21	na	na	6.8	10.4	8.9	12.5
Ukraine	2004	1935-84	1,513	124	***3***	16	0	20	0.27	0.05	5.6	12.0	10.4	12.9
									Regional Average		***6.3***	***11.0***	***9.8***	***12.2***
WESTERN EUROPE, USA														
Belgium (Flanders)*	1996	1932-76	1,628	124	0	19	0	21	na	na	7.1	11.2	10.3	13.4
Denmark*	1998	1933-78	2,762	232	0	19	0	21	na	na	10.3	12.6	10.9	12.5
Finland*	1998	1933-78	2,593	218	***4***	17	0	21	na	na	6.5	10.9	8.9	13.1
Ireland*	1994	1930-74	1,988	154	0	16	0	21	na	na	7.2	8.6	8.9	11.8
Italy*	1998	1933-78	2,756	223	0	20	0	21	na	na	4.7	8.9	7.2	13.1
Netherlands	1994	1925-74	2,725	144	0	16	0	21	na	na	8.1	10.9	9.8	13.9
New Zealand*	1996	1931-76	2,675	196	0	17	0	21	na	na	9.9	11.4	11.2	13.2
Northern Ireland*	1996	1931-76	2,338	174	0	17	0	21	na	na	9.4	10.6	10.7	13.9
Norway (Bokmål)**	1998	1933-78	2,871	216	0	18	0	21	na	na	7.8	10.7	10.0	12.1
Sweden	1994	1925-74	2,464	144	0	17	0	21	na	na	6.9	10.1	8.7	12.1
Switzerland***	1994 & 98	1929-78	3,711	116	0	18	0	21	na	na	9.0	12.6	10.9	12.8
United Kingdom*	1996	1931-76	3,078	248	0	16	0	21	na	na	9.9	10.6	10.6	13.0
USA†	1994-2000	1929-80	3,351	137	0	20	0	21	na	na	10.5	12.6	13.1	13.0‡
									Regional Average		***8.3***	***10.9***	***10.1***	***12.9***

* Ages 20 to 64 or 65 only; nine cohorts.
** Survey covered speakers of Bokmål only, ages 20 to 65.
*** French, German and Italian language surveys were merged.
† The first and last cohorts include six rather than five years of birth. Ages are between 20 and 69.
‡ Many of these 20-24 years olds are clearly still in school. The average attainment of the 25-29 year olds is 13.6 years.

Education is coded as the number of years associated with the highest grade completed, assuming no grade repetition, as ascertained from the *World Education Encyclopedia* (Marlow-Ferguson, 2002). Parents' education is defined as the average for mothers and fathers, wherever possible.[8] For both generations, we often had to impute values for categories like "Some college," or "Incomplete primary." This was done by examining the modal number of years of education reported for people in that grade category, for those countries and cases where both questions were available. A further problem is that some parental education values were recorded only crudely (e.g. none, primary, secondary, college). Yet experimentation with grouping the more detailed codes into these four categories showed that the intergenerational correlations were not greatly affected; the regression coefficients, however, were more sensitive to these coding decisions. For four countries in the Eastern Bloc and for Finland, the lowest category permitted for parents was "Not completed primary," and the bold-faced entries in the table reflect our best estimates of the appropriate modal number of years.

The next columns list the shares of adults who are currently enrolled in school (who are included in the analysis). For those aged 20-24, these estimated shares sometimes run as high as 0.30 or 0.40, representing both delayed completion of secondary schooling and the pursuit of higher education. For this five-year cohort we might expect the intergenerational regression coefficient to be downwardly biased: if the children of better-educated parents have not yet had time to achieve their terminal levels of education, the covariance between (high) parental education and children's education might be lessened. On the other hand, it may be that some children who were born to poorly educated parents and who are taking longer than usual to achieve their final level of education are also represented in this age category, in which case the bias might cut the other way. In light of this uncertainty, we repeat the analysis of trends without this highly-enrolled age group. Note that enrollment rates for the next five-year cohort (25 to 29) are generally under ten percent. In either case, however, we expect that any biases due to the right-censored nature of these data should be fairly small, since the true value is likely to be just a year or two greater than the observed.

[8] All surveys asked about both mothers' and fathers' schooling, but response rates varied. Parents' education was sometimes entirely missing for a significant fraction of cases, often with higher rates of missing data for older cohorts, as might be expected due to problems of recall. Records with at least one parent reporting were usable. For these, father's education was available 87 percent of the time, and mother's education was present for 92 percent. We chose to use the average of both parents' education (or a single parent's reported value if need be) rather than only the father's education in order to maximize our sample size, Another option, however, is to use the more educated parent. The correlation coefficients are generally only slightly affected by this choice, whereas the regression coefficients usually rise when the two parents' education values are averaged.

Other problems arise if we extend the upper age limit too far, since education and longevity are correlated, as noted by Behrman, Gaviria, and Székely (2001). This should bias the reported mean *levels* of own schooling upwards for the earliest cohorts. Moreover, to the extent that their appearance in the sample is correlated with their own schooling, conditional on their parents' schooling (i.e. with their error term in the intergenerational regression) this sample selection process could lead to a downward bias in the intergenerational regression coefficient for the earliest cohorts.

The last columns in Table 1 report the average levels of schooling for children in the first and last five-year birth cohorts, and that of their parents. Subject to the proviso that these averages are often based on fairly crude underlying categories for parents, it is clear that the average level of education has increased rapidly over the past 50 years, by about three years for parents and three to five years for children in Asia, our four African nations, and Latin America. In the Eastern Bloc, average parental education started at a much higher base and then rose by about five years. Given that children born in the late 1930s already averaged 9.8 years of schooling in this region, whereas their parents averaged 6.3, it is clear that major gains in education took place between about 1930 and 1955. In the West, parental and child education also rose, by about three years. These global gains in mean schooling are important, and we must not lose sight of them as we go on to analyze how these trends have co-evolved with trends in intergenerational educational persistence. Nor, however, may we confuse the two: intergenerational persistence is about the similarity between parents' and children's positions in each generation's educational distribution, not about average attainment.

Although these are arguably representative samples of children, they are not representative samples of parents: those parents who have more children will be over-represented (provided their children survive long enough to appear in the survey). Given the common finding of a negative relationship between education and fertility, we would expect our sampled parents to have lower-than-average levels of schooling. This, however, is not a flaw in the sample's construction, but rather an implication of treating the child rather than the parent as the unit of analysis.

METHODS

We first run regressions of s_1 on s_0, with no other covariates, and record the estimated regression coefficients and correlations. Survey sampling weights, if any, are used in these regressions.[9] We will refer to the regression coefficients as measures of "grade persistence," and the correlations as measures of "standardized persistence." We do this for each country, and each five-year birth cohort (c), denoting the results β_s^c (for grade persistence) and r_s^c (for standardized persistence). Recall that the relation between the two is given by $r_s^c = \beta_s^c(\sigma_0^c / \sigma_1^c)$, where σ_0^c and σ_1^c are the standard deviations of schooling in each generation. Next, we form the simple averages of these results across cohorts, for each country. The advantage of this approach as compared to running a single regression for all ages is that it does not give more weight to larger cohorts. It thus ignores questions of population growth and changes in fertility, and treats survivors-to-date as representative of their birth cohort. This corrects both for the fact that older cohorts are smaller, due to mortality, and that cohort shares may differ from population shares due to sample design, or sampling error.

To estimate time trends we run one set of simple linear regressions through all ten observations for each country, and another using nine observations, dropping the youngest cohort.[10] Finally, we estimate an overall global trend by pooling all cohorts and nations, and introducing country fixed effects.

Aggregation into five-year cohorts introduces a degree of arbitrariness, since one could always use larger or smaller intervals, or space them differently. However, aggregation should not bias our trend estimates unless the intervals are chosen with particular results in mind. We adopted the five-year age-bands for all surveys, and did not examine results under other aggregation schemes.

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[9] Deaton (1997) reminds us that if grade persistence is heterogeneous across groups in the population, and these group proportions are affected by the survey weights, then applying the weights does not yield the population-size-weighted average level of grade persistence. It does, however, produce a consistent estimator of the value that would be observed if the data were drawn as a simple random sample, or a full census. This is the statistic we seek to estimate.

[10] Another method, used for example by Ganzeboom and Nieuwbeerta (1999) or Mayer and Lopoo (2005), is to estimate the trend in a single pooled equation for all birth years, including both main effects of cohort (as a set of dummy variables) and an interaction of year of birth (as a continuous variable) with the parental attribute. This approach is often more efficient at detecting linear time trends, but only by assuming that grade persistence evolves linearly, which is often not true. Our approach allows it to vary freely by cohort, and then asks whether these results do in fact display a clear linear trend. (The pooled approach also has the disadvantage of placing greater weight on the larger cohorts. Note, however, that our approach does not give equal weight to each cohort either; the approach that *does* is simply to average the cohort-to-cohort changes, which, in this context, is no different than just reporting the change between the first and last birth cohorts. But this involves discarding eight-tenths of the data.)

RESULTS

Table 2 displays the long-run averages across cohorts of β_s^c (first column) and r_s^c (second column). As noted, the seven highest intergenerational schooling correlations are found in our seven Latin American countries: the regional average is 0.60, compared to values between 0.36 and 0.41 for the other four regions. Among Asian nations, Indonesia represents a high-outlier (at 0.55) while rural China's average value across cohorts was just 0.20. The formerly communist countries and the high-income Western nations vary between 0.28 (Kyrgyzstan) and Italy (0.54). The four Nordic nations (Denmark, Finland, Norway and Sweden) had an average parent-child schooling correlation of 0.34, which was significantly lower than the non-Nordic average (0.41, $p<0.01$), and also lower than the average for the Eastern Bloc (0.41, $p<0.01$).

The regression coefficients are much more volatile than are the correlations, but the two rankings are clearly related: the simple correlation between them is 0.51. The cross-regional differences just noted were generally not statistically significant when measured by the regression coefficients.

Figure 1 presented 403 country-cohort-specific estimates of the regression coefficients (top), and the correlations (bottom), plotted against the second generation's year of birth, along with their trend lines. Globally, we see a fall of 30 points in the estimated mean regression coefficient over 60 years, from 0.80 in 1920 to 0.50 in 1980. The correlation coefficients, however, display no significant time trend: they appear to have held steady, on average, at about 0.43 for the last half-century, meaning that parents' education continues to predict about 18 percent of the variance in children's education.

REGIONAL RESULTS: LATIN AMERICA

Table 3 reports our estimates of the trends in grade persistence and standardized persistence for Latin America; plots by cohort for each country may be found in Appendix 1. (The countries near the top of the table are those that saw the most rapid declines in standardized persistence among the 25-to-69 year-olds, reported in the final column). As already noted, long-run average persistence is high in Latin America; still, every country except Nicaragua showed a significant reduction in either grade or standardized persistence, or both, over time, and there were no significant positive trends for either measure for any country. These declines have lowered the average regression coefficient for the two most recent cohorts (the 20-29 year olds) to 0.60, and the correlation to 0.56, numbers that are still high by international standards (results not shown in table).

TABLE 2
COUNTRIES RANKED BY AVERAGE PARENT-CHILD SCHOOLING CORRELATION, AGES 20-69

Country	Coefficient	Rank	Correlation	Rank
Peru	0.88	6	0.66	1
Ecuador	0.72	12	0.61	2
Panama	0.73	11	0.61	3
Chile	0.64	18	0.60	4
Brazil	0.95	4	0.59	5
Colombia	0.80	8	0.59	6
Nicaragua	0.82	7	0.55	7
Indonesia	0.78	9	0.55	8
Italy†	0.67	17	0.54	9
Slovenia†	0.54	27	0.52	10
Egypt	1.03	2	0.50	11
Hungary†	0.61	20	0.49	12
Sri Lanka	0.61	19	0.48	13
Pakistan	1.00	3	0.46	14
USA	0.46	33	0.46	15
Switzerland†	0.49	30	0.46	16
Ireland†	0.70	15	0.46	17
South Africa (KwaZulu-Natal)	0.69	16	0.44	18
Poland†	0.48	31	0.43	19
Vietnam	0.58	23	0.40	20
Philippines	0.41	36	0.40	21
Belgium (Flanders)	0.41	35	0.40	22
Estonia	0.54	28	0.40	23
Sweden	0.58	26	0.40	24
Ghana	0.71	13	0.39	25
Ukraine	0.37	40	0.39	26
East Timor	1.27	1	0.39	27
Bangladesh (Matlab)	0.58	25	0.38	28
Slovakia	0.61	21	0.37	29
Czech Republic†	0.44	34	0.37	30
The Netherlands	0.58	24	0.36	31
Norway	0.40	38	0.35	32
Nepal	0.94	5	0.35	33
New Zealand†	0.40	37	0.33	34
Finland	0.48	32	0.33	35
Northern Ireland	0.59	22	0.32	36
Great Britain†	0.71	14	0.31	37
Malaysia	0.38	39	0.31	38
Denmark	0.49	29	0.30	39
Kyrgyzstan	0.20	42	0.28	40
China (Rural)	0.34	41	0.20	41
Ethiopia (Rural)	0.75	10	0.10	42

Surveyed between 1994 and 2004, except Peru (1985), Malaysia (1988) and Pakistan (1991).
† Ages 20 to 64 or 65 only.

TABLE 3
REGRESSION COEFFICIENTS AND CORRELATIONS, AVERAGES AND TRENDS: LATIN AMERICA

LATIN AMERICA	Average Across 10 5-Year Cohorts Coeff.	Correl.	Linear Trends (Per 5 Years) Coeff. Age 20+	Coeff. Age 25+	Correl. Age 20+	Correl. Age 25+
Colombia	0.80	0.59	-0.0377‡	-0.0294†	-0.0207‡	-0.0198‡
	(0.041)	(0.021)	(0.0087)	(0.0079)	(0.0017)	(0.0021)
Peru	0.88	0.66	-0.0204	-0.0057	-0.0177‡	-0.0172‡
	(0.040)	(0.021)	(0.0134)	(0.0140)	(0.0036)	(0.0050)
Ecuador	0.72	0.61	-0.0158	-0.0048	-0.0158‡	-0.0150‡
	(0.029)	(0.017)	(0.0100)	(0.0089)	(0.0029)	(0.0036)
Chile	0.64	0.60	-0.0416‡	-0.0377‡	-0.0099	-0.0106
	(0.050)	(0.024)	(0.0091)	(0.0109)	(0.0062)	(0.0094)
Brazil	0.95	0.59	-0.0489‡	-0.0372‡	-0.0064‡	-0.0069‡
	(0.055)	(0.007)	(0.0119)	(0.0110)	(0.0017)	(0.0019)
Panama	0.73	0.61	-0.0425‡	-0.0387‡	-0.0067†	-0.0066
	(0.044)	(0.011)	(0.0083)	(0.0092)	(0.0031)	(0.0040)
Nicaragua	0.82	0.55	-0.0030	0.0064	-0.0009	-0.0028
	(0.032)	(0.013)	(0.0132)	(0.0153)	(0.0041)	(0.0049)
Regional Average	**0.79**	**0.60**				
	(0.040)	(0.012)				

† Significant at 10%; ‡ Significant at 5%; robust standard errors in parentheses.

Our results for Panama are partially consistent with those derived from an earlier survey by Heckman and Hotz (1986). They report a father-son schooling correlation of 0.57 for sons who were born before 1965 and still lived at home in 1983; this is very close to the average of our results for the 1956, 1961 and 1966 cohorts, which works out to 0.58. However, they also report a higher estimate (0.68) for the schooling correlation between these fathers and their own fathers, which would imply that the correlation fell by eleven points over the course of a generation. This trend is far steeper than we detect: our estimates suggest the correlation fell by about four points over thirty years. This difference may be due to a difference in sample selection criteria: their father/grandfather pairs are selected on the basis of the fathers' having older children still living at home.

Behrman, Gaviria, and Székely (2001) present results for Brazil, Colombia, and Peru, using the same datasets we employ. Their estimates of the trends in grade persistence are similar to ours: they indicate a fall of 25 to 30 points in the regression coefficient over 35 years for Colombia and Brazil (about 3.6 to 4.3 points per five years) with smaller declines for Peru. Trends in standardized persistence, which has fallen much more slowly over time, are not discussed.

Although our estimated trends are similar, our reported levels of grade persistence are uniformly higher than those of Behrman, Gaviria, and Székely (2001). This difference is driven primarily by the fact that we average both parents' education levels, as opposed to selecting the better-educated parent, which alters the regression coefficient. This choice has a much smaller effect on the correlation coefficient, and this illustrates a more general point: in addition to being more stable over time, the correlations are less sensitive to small differences in the ways samples and variables are defined, which is a point in their favor.

Figure 2, below, provides an illustration of the long-run change in the joint and marginal distributions of schooling in Brazil, where grade persistence has declined steeply. It compares the 25-29 year olds (born 1967-71, right plot) to the same raw number of observations drawn at random from the first three cohorts (born 1927-41, left plot). Because the data are discrete, a random noise term is added to each point in the figure, so that they do not perfectly coincide; this allows us to get a sense of the density in that vicinity. We see that the slope of the regression line falls over time (from 1.1 to 0.74) and that the intercept rises significantly, from 1.5 years to 4.8, meaning that the children of uneducated parents are faring better than before. Over this period, however, the correlation coefficient only fell from 0.61 to 0.57, meaning that, in standardized terms, educational persistence is virtually unchanged. By either measure, Brazil stands out as a high-persistence nation, as noted by others (e.g. Pastore and Zylberstajn, 1996; Dunn, 2007).

REGIONAL RESULTS: ASIA

Average education in the ten Asian surveys grew from about one to four years for the parents, and from 4.7 to 7.5 years for the children, although there was considerable variation across countries in both starting and final values (see Table 1). Table 4, below, reports trends in persistence for each country; the detailed results may again be found in Appendix 1. Five of the ten trend estimates for grade persistence in the 25-69 age group are negative and significant at the ten percent level or better; one is significantly positive. However, only two nations (Rural China and Sri Lanka) show any significant evidence of a negative trend in standardized persistence, while for five countries this trend was positive. Indonesia and Vietnam provide the clearest examples of differences in the signs of these two trends. In Indonesia, for example, the regression coefficient fell by about 0.04 per five years, while the correlation coefficient rose by about 0.01 per five years.

Our correlation coefficient for Malaysia (0.31) is somewhat higher than was reported by Lillard and Willis (1994) for father-son and father-daughter correlations (0.19 and 0.23), most likely because of differences in our sample definitions. This points out the importance of using consistent methods across countries, but also confirms that, by either method, intergenerational educational persistence in Malaysia is fairly low by international standards.

FIGURE 2
BRAZIL: RISING EDUCATIONAL ATTAINMENT, FALLING GRADE PERSISTENCE, BUT LITTLE CHANGE IN STANDARDIZED PERSISTENCE

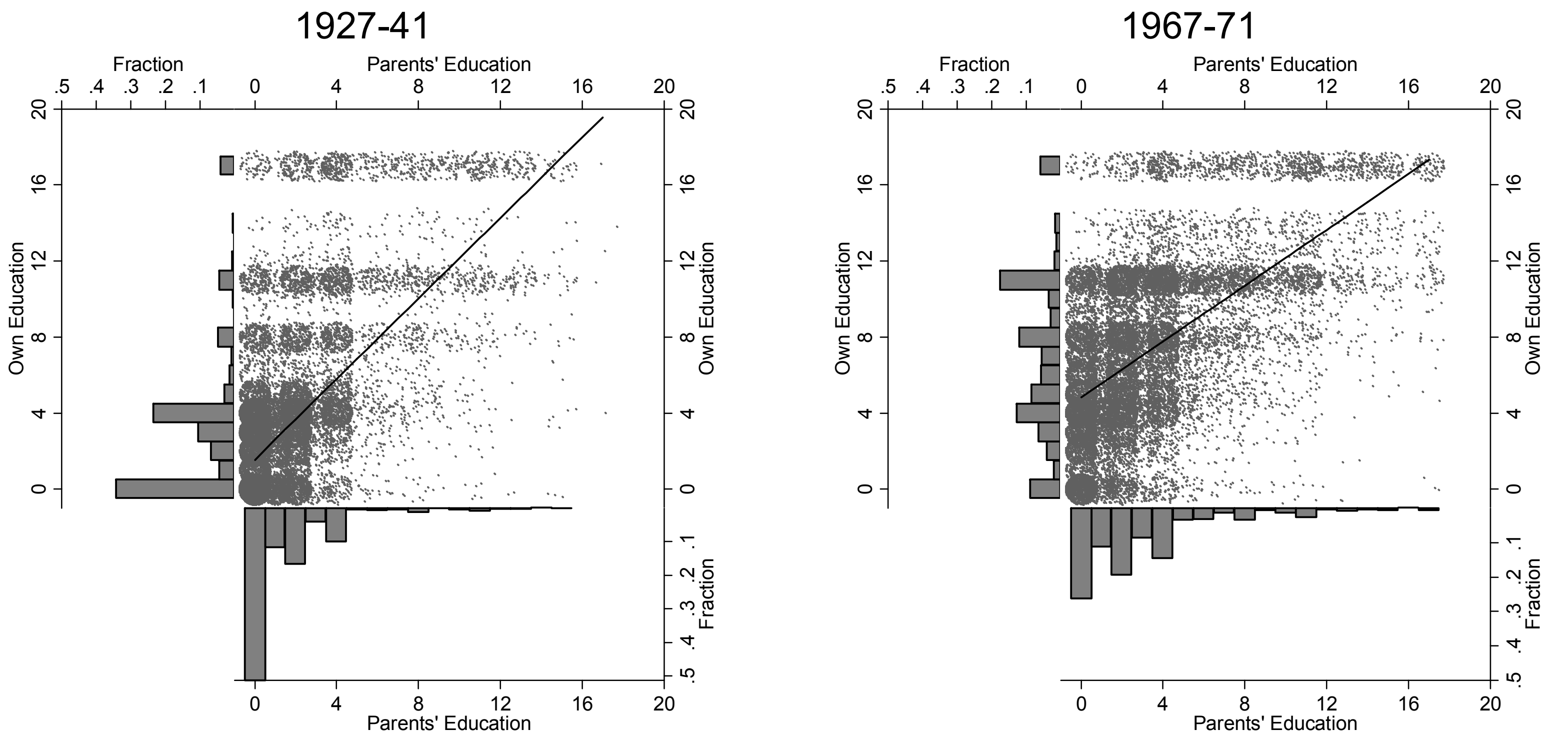


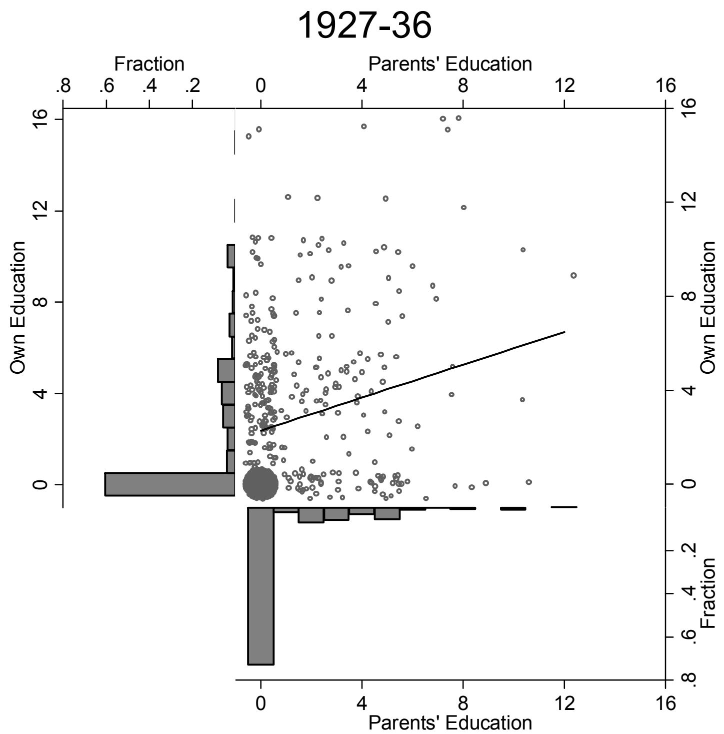
TABLE 4
REGRESSION COEFFICIENTS AND CORRELATIONS, AVERAGES AND TRENDS: ASIA

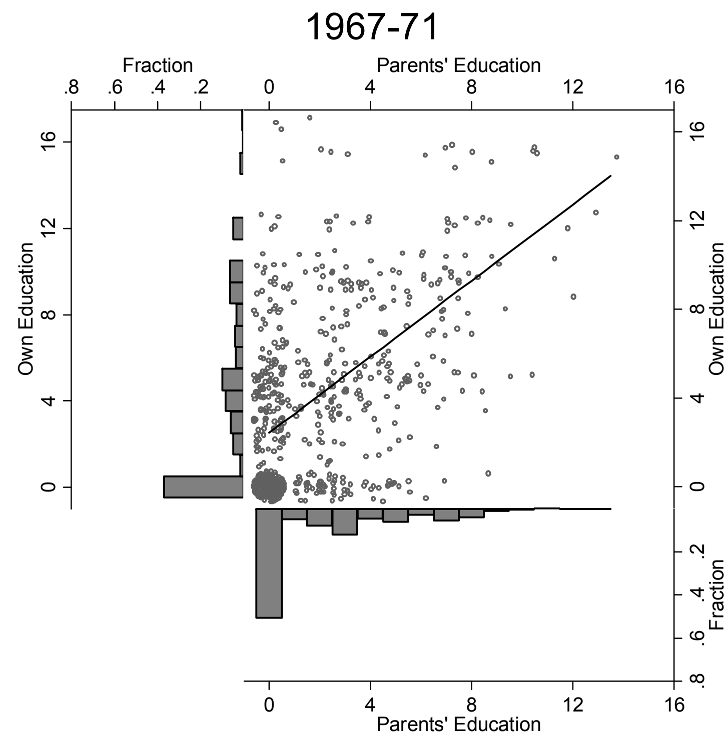
ASIA	Average Across 10 5-Year Cohorts Coeff.	Correl.	Linear Trends (Per 5 Years) Coeff. Age 20+	Coeff. Age 25+	Correl. Age 20+	Correl. Age 25+
China (Rural)	0.34	0.20	-0.0509‡	-0.0623‡	-0.0063	-0.0155‡
	(0.060)	(0.022)	(0.0117)	(0.0090)	(0.0080)	(0.0045)
Philippines	0.41	0.40	-0.0218‡	-0.0266‡	-0.0073	-0.0126
	(0.030)	(0.029)	(0.0078)	(0.0102)	(0.0108)	(0.0136)
Sri Lanka	0.61	0.48	-0.0431‡	-0.0400‡	-0.0075‡	-0.0079
	(0.043)	(0.013)	(0.0036)	(0.0032)	(0.0031)	(0.0044)
Malaysia	0.38	0.31	-0.0283‡	-0.0250	-0.0055	-0.0067
	(0.024)	(0.006)	(0.0077)	(0.0126)	(0.0030)	(0.0047)
East Timor	1.27	0.39	0.0088	0.0433	-0.0011	-0.0055
	(0.172)	(0.044)	(0.0669)	(0.0768)	(0.0206)	(0.0254)
Indonesia	0.78	0.55	-0.0418‡	-0.0395‡	0.0117‡	0.0112‡
	(0.041)	(0.014)	(0.0029)	(0.0036)	(0.0036)	(0.0046)
Pakistan	1.00	0.46	0.0023	0.0072	0.0124‡	0.0120†
	(0.014)	(0.016)	(0.0056)	(0.0044)	(0.0043)	(0.0055)
Vietnam	0.58	0.40	-0.0191‡	-0.0264‡	0.0186‡	0.0133‡
	(0.023)	(0.020)	(0.0061)	(0.0019)	(0.0044)	(0.0014)
Nepal	0.94	0.35	0.0059	0.0242	0.0235†	0.0242†
	(0.050)	(0.024)	(0.0233)	(0.0218)	(0.0023)	(0.0026)
	0.58	0.38	0.0459†	0.0559‡	0.0340‡	0.0394‡
Bangladesh (Matlab)	(0.064)	(0.045)	(0.0206)	(0.0236)	(0.0129)	(0.0155)
Regional Average	**0.69**	**0.39**				
	(0.052)	(0.023)				

† Significant at 10%; ‡ Significant at 5%; robust standard errors in parentheses.

Bangladesh's Matlab province stands out for having the most pronounced long-run increase in both grade and standardized persistence. As illustrated in Figure 3, about 70 percent of parents and children in the earliest cohorts had no schooling (left panel), and each additional grade of parental education contributed about 0.35 extra years to one's own expected value of education. In the later cohorts (right panel) the share with no schooling falls considerably, and roughly proportionally, in both generations. But now a one-year difference in parental education corresponds to a difference of 0.86 years in the expected value of own schooling. The data points also lie closer to the regression line in the later period, raising the correlation coefficient. The intercept remains unchanged at 2.4 years, indicating no long-term gains in schooling for the children of uneducated parents. This shows that rising mean attainment can coincide with rising intergenerational persistence and a lack of progress at the bottom of the educational distribution, a phenomenon that appears more likely when initial levels of schooling are very low, but can also occur among better-educated populations, as discussed below.

FIGURE 3
BANGLADESH: RISING EDUCATIONAL ATTAINMENT, AND RISING INTERGENERATIONAL PERSISTENCE





REGIONAL RESULTS: AFRICA

Three of the four African surveys paint fairly similar pictures: in Egypt, the KwaZulu-Natal province of South Africa, and in Ghana, average parental education grew from below one year to between 3.4 and 4.8 years, over similar time periods, while average education in the second generation grew from around two to around eight years (Table 1). As may be seen in the Appendix, in all three countries the regression coefficient linking parental to child schooling started quite high (above unity for Egypt and Ghana) but fell significantly for those born after about 1950 or 1955. In Egypt the correlation coefficient also fell, by about two points per five years, or from 0.62 to 0.40 over 50 years, indicating a reduction in standardized persistence (Table 5). In South Africa and Ghana, however, the correlation coefficient did not move, hovering at around 0.40.

TABLE 5
REGRESSION COEFFICIENTS AND CORRELATIONS, AVERAGES AND TRENDS: AFRICA

AFRICA	Average Across 10 5-Year Cohorts Coeff.	Average Across 10 5-Year Cohorts Correl.	Linear Trends (Per 5 Years) Coeff. Age 20+	Coeff. Age 25+	Correl. Age 20+	Correl. Age 25+
Egypt	1.03	0.50	-0.0967‡	-0.0839‡	-0.0210‡	-0.0216‡
	(0.097)	(0.022)	(0.0118)	(0.0099)	(0.0027)	(0.0037)
South Africa	0.69	0.44	-0.0591‡	-0.0497‡	-0.0054	-0.0013
(KwaZulu-Natal)	(0.063)	(0.014)	(0.0106)	(0.0105)	(0.0048)	(0.0050)
Ghana	0.71	0.39	-0.0543‡	-0.0536‡	0.0081	0.0070
	(0.066)	(0.015)	(0.0162)	(0.0197)	(0.0047)	(0.0058)
Rural Ethiopia	0.75	0.10	0.0511	0.0616	0.0256‡	0.0248‡
	(0.185)	(0.031)	(0.0690)	(0.0814)	(0.0056)	(0.0070)
Regional Average	**0.80**	**0.36**				
	(0.079)	(0.089)				

† Significant at 10%; ‡ Significant at 5%; robust standard errors in parentheses.

The fourth country, Ethiopia, produces erratic initial results: in the earliest cohorts nearly 98 percent of parents had no schooling, rendering the estimated slope of the regression line essentially meaningless. From 1952 on, there is enough dispersion in parental education to generate more stable estimates. For this period, the average regression coefficient stands at 0.92, but the correlation is only on the order of 0.18. In other words, small differences in parental education make a large difference (nearly one-for-one) in the expected value of children's education, but parental education does not vary a great deal across households and can explain only a modest share of the overall variation in child schooling.

Our results from South Africa agree nicely with those of Thomas (1996), who uses a different dataset covering a larger portion of the country. We both observe unchanging grade persistence for those born before 1950, but falling coefficients thereafter (see Appendix 1). The fact that standardized persistence has barely changed in 50 years, however, is not remarked upon by Thomas, or by anyone else, to our knowledge. This stems in part from a difference in method, reflecting a difference in research objectives. Thomas and others, such as Alain-Désiré and Vencatachellum (2007) who have also produced estimates for South Africa, seek to quantify the structural effect of parental education and are thus concerned with controlling for factors such as social capital and unobservable household characteristics. By contrast, our aim, and that of much of the mobility literature, is purely descriptive.[11]

REGIONAL RESULTS: THE EASTERN BLOC

With the apparent exception of Kyrgyzstan, the eight Eastern Bloc countries surveyed all had higher average initial levels of education than were found in Latin America, Asia, or Africa. Table 2 reveals that five of our eight former communist countries fall below the full sample's median in terms of their long-run average intergenerational correlation coefficients, while three nations, Slovenia, Hungary, and Poland, have above-median correlations (of 0.52, 0.49, and 0.43). Results using the ESS surveys instead of the IALS were similar.

Given that many communist countries made explicit efforts to raise the educational levels of the children of working class parents, we might expect to see significant declines in educational persistence over the post-War period. Some evidence of this is found in Table 6, in terms of grade persistence, for Slovenia, Kyrgyzstan, Poland and Hungary. However, only Slovenia saw a reduction in standardized persistence, and this statistic rose significantly in Hungary, the Ukraine and the Czech Republic.

Four of the nations studied here, namely, Hungary, Poland, the Czech Republic, and Slovakia also figure in an analysis by Ganzeboom and Nieuwbeerta (1999). Using larger samples (numbering between 3400 and 5500 per country, as compared to our 1100 to 2800) they find significant negative trends in grade persistence during this period for Hungary and weak evidence of a negative trend for Poland (as do we). However, they also observe clear negative trends for the Czech Republic and for Slovakia (where we find no trend). Experimentation with different methods of estimating the trend suggests that the discrepancy between our findings and theirs is primarily due to a difference in specification. Ganzeboom and Nieuwbeerta report that it makes little difference whether the parental education parameter is assumed to follow a linear trend, or allowed to vary freely across cohorts, but in our data this choice makes a clear difference,

[11] Similarly, both of the authors just mentioned disaggregate by race, and compare within-group results. As argued by Hertz (2005; 2007a), this is inappropriate if one seeks to describe group members' mobility in relation to the full education (or income) distribution.

with the linear assumption producing steeper negative trends than does our preferred specification.[12]

TABLE 6
REGRESSION COEFFICIENTS AND CORRELATIONS, AVERAGES AND TRENDS: EASTERN BLOC

EASTERN BLOC	Average of 9 or 10 5-Year Cohorts Coeff.	Correl.	Linear Trends (Per 5 Years) Coeff. Age 20+	Coeff. Age 25+	Correl. Age 20+	Correl. Age 25+
Slovenia*	0.54	0.52	-0.0310‡	-0.0353‡	-0.0102	-0.0152‡
	(0.031)	(0.016)	(0.0054)	(0.0062)	(0.0056)	(0.0061)
Kyrgyzstan	0.20	0.28	-0.0288‡	-0.0323‡	-0.0067	-0.0023
	(0.032)	(0.014)	(0.0061)	(0.0078)	(0.0048)	(0.0039)
Estonia	0.54	0.40	-0.0121	-0.0085	-0.0051	0.0001
	(0.042)	(0.030)	(0.0141)	(0.0176)	(0.0108)	(0.0127)
Poland*	0.48	0.43	-0.0184†	-0.0121	0.0018	0.0074
	(0.032)	(0.032)	(0.0097)	(0.0121)	(0.0111)	(0.0150)
Slovakia	0.61	0.37	0.0092	0.0156	0.0108	0.0141
	(0.062)	(0.039)	(0.0237)	(0.0286)	(0.0139)	(0.0179)
Hungary*	0.61	0.49	-0.0230‡	-0.0226‡	0.0179‡	0.0208‡
	(0.027)	(0.022)	(0.0055)	(0.0071)	(0.0055)	(0.0067)
Ukraine	0.37	0.39	0.0084	0.0140	0.0125	0.0211†
	(0.037)	(0.023)	(0.0160)	(0.0223)	(0.0095)	(0.0092)
Czech Republic*	0.44	0.37	-0.0068	0.0028	0.0181‡	0.0240‡
	(0.026)	(0.023)	(0.0097)	(0.0091)	(0.0062)	(0.0057)
Regional Average	**0.47**	**0.41**				
	(0.049)	(0.026)				

† Significant at 10%; ‡ Significant at 5%; robust standard errors in parentheses.
* Ages 20 to 64 or 65 only; nine cohorts.

REGIONAL RESULTS: WESTERN EUROPE AND THE USA

Table 7 reports our results for thirteen high-income Western nations. Our estimate of the intergenerational schooling correlation for the U.S. (0.46) is a few points higher than Couch and Dunn's (1997) results, which are derived from the Panel Study of Income Dynamics and range between 0.40 and 0.43; this discrepancy is likely explained by differences in sample selection criteria. By our estimate, the rate of standardized educational persistence in the U.S. is comparable to that found in Ireland and Switzerland, and among high-income nations is exceeded only by the results from Italy (Table 2).

[12] Their methods are otherwise very similar to ours: they take the average of parents education; they group children into five-year cohorts; and they include adults aged 21 to 69. However, they disaggregate by gender; they impute a final number of years of education for those who are still in school; and they include controls for the size of the town the respondent lived in. Again, the issue of standardized persistence is not discussed.

The particularly high intergenerational schooling correlation estimates for Italy, which have been quite steady over time (see Appendix 1), are consistent with Piraino's (2007) finding that Italy's level of *income* persistence across generations is also high by international standards. Piraino then demonstrates that only about one-third of the father-son income correlation can be attributed, in a descriptive sense, to the link between the father's income and the son's education, and concludes that most of the transmission of income occurs through other channels. This, however, does not contradict the finding that educational persistence is strong; indeed, Piraino's "other channels" would include any direct causal effect of parents' schooling on that of their children.

Standardized educational persistence in Great Britain, by contrast, is not especially high. The fact that it is much lower than in the U.S. is noteworthy, given that the U.S. and the U.K. have similar values of income persistence, which are the highest observed among the rich nations (Corak, 2006; Jäntti, *et al.*, 2006). However, although they differ in terms of persistence levels, the U.S. and the U.K. are alike in being the only two countries in Table 7 to display a statistically significant increase over time in standardized persistence, a phenomenon that we have previously seen only in countries at very low levels of schooling. Moreover, for Britain this trend accords with work by Blanden, *et al.* (2004), who observe that rising educational attainment between the late 1950s and early 1970s was driven primarily by increases in tertiary education among the children of higher-income, well-educated families, and thus represented an increase in intergenerational educational persistence. They go on to argue that this helps explain an increase in income persistence over the same period.[13]

[13] The estimates of grade persistence (regression coefficients) for Great Britain are extremely volatile over cohorts, due primarily to the concentration of the majority of parents in a single educational category. This coding anomaly casts some doubt on the quality of the data, but, as already noted, has much less impact on estimates of standardized persistence, and its trend, than on grade persistence.

TABLE 7
REGRESSION COEFFICIENTS & CORRELATIONS, AVERAGES & TRENDS: WESTERN EUROPE & USA

WESTERN EUROPE, USA	Average of 9 or 10 5-Year Cohorts		Linear Trends (Per 5 Years)			
	Coeff.	Correl.	Coeff. Age 20+	Coeff. Age 25+	Correl. Age 20+	Correl. Age 25+
Netherlands	0.58	0.36	-0.0649‡	-0.0569‡	-0.0290‡	-0.0285‡
	(0.073)	(0.037)	(0.0121)	(0.0157)	(0.0078)	(0.0107)
Ireland*	0.70	0.46	-0.0620‡	-0.0700‡	-0.0208	-0.0281
	(0.066)	(0.031)	(0.0132)	(0.0169)	(0.0128)	(0.0156)
Finland*	0.48	0.33	-0.0557‡	-0.0484‡	-0.0223‡	-0.0211‡
	(0.057)	(0.027)	(0.0116)	(0.0118)	(0.0048)	(0.0057)
Denmark*	0.49	0.30	-0.0627‡	-0.0482‡	-0.0213†	-0.0119
	(0.065)	(0.029)	(0.0165)	(0.0147)	(0.0104)	(0.0098)
Switzerland	0.49	0.46	-0.0475‡	-0.0348‡	-0.0221‡	-0.0110
	(0.054)	(0.030)	(0.0107)	(0.0086)	(0.0095)	(0.0075)
Northern Ireland*	0.59	0.32	-0.0126	-0.0188	0.0013	-0.0092
	(0.039)	(0.023)	(0.0141)	(0.0181)	(0.0113)	(0.0104)
Sweden	0.58	0.40	-0.0648‡	-0.0582‡	-0.0114	-0.0050
	(0.076)	(0.024)	(0.0183)	(0.0205)	(0.0101)	(0.0098)
New Zealand*	0.40	0.33	0.0072	-0.0020	0.0038	-0.0046
	(0.023)	(0.017)	(0.0109)	(0.0085)	(0.0091)	(0.0062)
Italy*	0.67	0.54	-0.0432‡	-0.0339‡	-0.0022	-0.0026
	(0.043)	(0.008)	(0.0078)	(0.0048)	(0.0032)	(0.0041)
Norway*	0.40	0.35	-0.0552‡	-0.0506‡	-0.0057	0.0006
	(0.053)	(0.016)	(0.0081)	(0.0087)	(0.0070)	(0.0064)
Belgium (Flanders)*	0.41	0.40	-0.0068	0.0013	0.0069	0.0115
	(0.026)	(0.022)	(0.0093)	(0.0088)	(0.0079)	(0.0102)
USA	0.46	0.46	-0.0081	-0.0011	0.0219‡	0.0222‡
	(0.018)	(0.023)	(0.0073)	(0.0056)	(0.0028)	(0.0034)
Great Britain*	0.71	0.31	0.0166	0.0306	0.0260‡	0.0283†
	(0.075)	(0.035)	(0.0263)	(0.0321)	(0.0107)	(0.0136)
Regional Average	**0.54**	**0.39**				
	(0.031)	(0.021)				

† Significant at 10%; ‡ Significant at 5%; robust standard errors in parentheses.
* Ages 20 to 64 or 65 only; nine cohorts.

Finally, we observe that one important general finding in the income-mobility literature is reproduced in the education data, namely, that the Nordic countries have low rates of intergenerational status persistence. The average of the cohort-specific intergenerational schooling correlations for Denmark, Finland, Norway, and Sweden is 0.34, compared to 0.41 for the non-Nordic high-income Western nations; this difference is statistically significant at the one-percent level. The Nordic nations also display significantly lower levels of standardized educational persistence, on average, than do the Eastern Bloc countries.

DISCUSSION

Taking a global view, we have one measure, the regression coefficient, by which the transmission of educational attainment from parent to child has decreased markedly over the past 50 years, and another, the correlation coefficient, by which it has not changed. What accounts for this difference? First, note the factors that do *not* distinguish between the two. Most importantly, both are simple linear measures of statistical association, not of the true causal effect of parental schooling on children's schooling, *ceteris paribus.* In particular, the regression coefficient necessarily overstates that structural parameter, due to the omission of many other socioeconomic factors that are positively correlated with schooling in both generations. Thus its secular decline does not prove that the causal influence of parental education is lessening. Second, both statistics assume that education by itself is a meaningful measure of status, and both are content to assign equal status to all those with equal grade levels (in a given generation), even if some such categories contain the majority of the population. Last, both statistics are unconcerned with differences in mean schooling across generations, countries or cohorts, or with the distinction between upward and downward mobility.[14]

Where the two differ is that one measures interpersonal differences in status by the difference in grade, while the other divides the grade difference by the standard deviation of education in that generation. Whether this produces a more appropriate measure of status differences is a matter of opinion.[15] We may ask whether the welfare effect of a one-grade difference in the expected value of schooling strikes us as more important in a society in which education does not vary a great deal among people. This is not the question of whether we measure our status in relation to that of others, or to the mean, but rather whether we evaluate our status differences with others in relation to the overall dispersion of outcomes. If interpersonal differences are deemed more important when they represent a larger proportion of the typical observed difference among people, then this would provide a basis for preferring the correlation coefficient, which takes the dispersion of status into account for each generation. As a practical matter, the correlation coefficient has the advantage of being more robust to alternative coding assumptions, and less volatile over time.

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[14] Fields (2000) provides a typology of mobility/persistence measures that characterizes the correlation coefficient as a measure of time dependence, or of symmetrical status movement.

[15] Analogously, we might debate whether it is desirable to take logs of income, so that income differences are measured in percentages; or whether household income should be divided by the number of household members; or whether parental education should be summed or averaged, or assigned the value of the better educated parent. In each case, the issue is *not* which choice is structurally correct, but which best accords with our conception of status, or status difference.

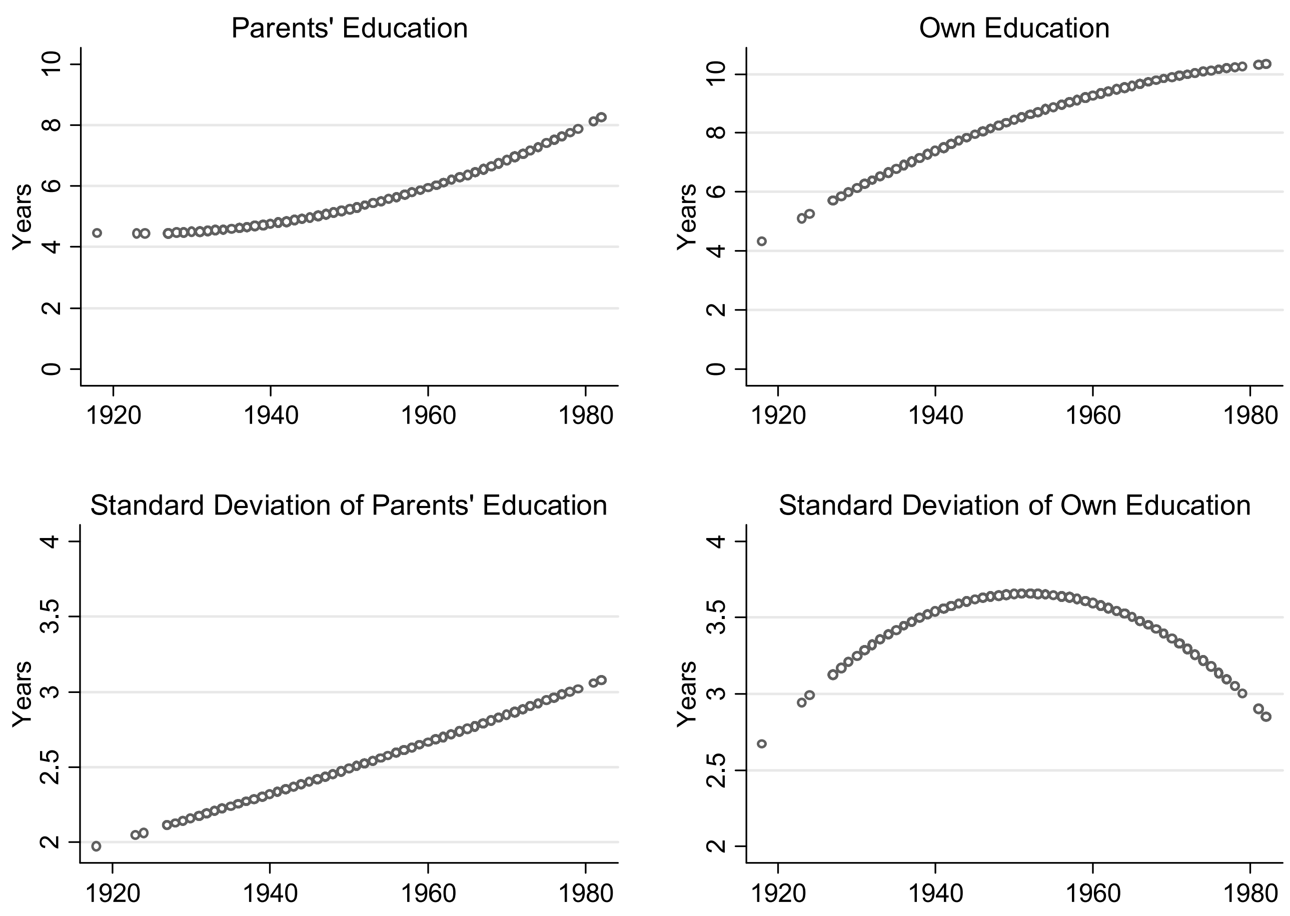
These considerations relate to the interpretation of the difference in the trends in grade persistence and standardized persistence. A second task is to describe the origins of that difference, and Figure 4 provides the explanation.[16] Average education has grown steadily over time in our sample, first at an increasing rate, and then at a decreasing rate. (Note that we could roughly concatenate the parental time series to that of the children if we were to shift the parents' graphs 25 or 30 years to the left, to reflect their own dates of birth rather than their children's.) The standard deviation of schooling, however, at first increased, then decreased, confirming Ram's (1990) finding of a Kuznets-type relationship between the level and dispersion of education.[17] This finding is intuitive: if nearly all initially have no education, and then a minority gain access to schooling, the variance of education will increase. That this rate of increase should slow, and eventually turn negative, is also intuitive, unless education grows without bound. If the share of people with no, or very little, schooling falls sufficiently as the mean and mode rise, this will reduce the mass in the left tail of the distribution enough to reduce its variance. And finally, if the process converges to a steady state, the trend in the variance of education becomes flat, by definition.

The increase in the variance of schooling means that the ratio of the standard deviation of parents' education to that of their children, born some 15 to 55 years later, will be less than one, and this in turn implies that r_s^c lies below β_s^c. As the standard deviation of schooling continues to grow more or less linearly over time, this ratio will converge to unity from below, driving β_s^c and r_s^c towards each other, as observed in many countries in Appendix 1. Finally, as growth in the standard deviation of education begins to slow and eventually turns negative, the correlation may come to exceed the regression coefficient, as it has in many nations at higher levels of schooling.

[16] The figure plots the predicted values from regressions of the variable indicated against year of birth and its square, from our panel of 42 countries over a maximum of 10 cohorts. Country fixed effects are included, meaning that the trends represent within-country changes over time. In the plots the country fixed effects are suppressed but the grand mean is included.

[17] Ram finds that the standard deviation of education rises with the mean until the mean reaches 6.3 years, in a cross-section of 74 developing countries. A longitudinal analysis of the developing countries in our sample, with country fixed effects, puts the turning point at a mean of 6.2 years. For the full sample of 42 countries, the turning point lies at 7.4 years.

FIGURE 4
TRENDS IN MEANS AND STANDARD DEVIATIONS OF SCHOOLING, BY CHILDREN'S YEAR OF BIRTH



This process explains why the slope of the time trend in grade persistence is almost invariably less than that of standardized persistence (i.e. more negative, or negative when the latter is positive, or a smaller positive number), but it does not constrain the sign of either of those trends, or explain what causes them. To understand this process we should start by noting the basic fact that grade persistence will fall if increases in average education are driven primarily by increases among the sons and daughters of poorly educated parents, as typically happens when primary schooling is expanded. On the other hand, if the children of better-educated parents are the first to take advantage of new educational opportunities, grade persistence will rise, as has happened in Bangladesh. Matters are more complicated for standardized persistence: as we have seen, increases in the correlation coefficient can coincide with decreases in the slope of the regression line if the data become more tightly clustered around that line.

There is a large body of empirical research that documents the link between changes in the economic and policy environment and changes in educational attainment. An increasing number of such papers take the further step of evaluating these schooling changes in relation to the level of education of the parents, thereby qualifying as studies of educational persistence. First, as noted above, Ganzeboom and Nieuwbeerta argue that communist educational policies led to reductions in grade persistence, although we are only able to replicate this finding for half of our Eastern Bloc nations. Our research on this topic continues, using more sources of data, and looking at broader definitions of social mobility.

Second, Hertz and Jayasundera (2007) study the natural experiment created by Indonesia's massive school construction program in the 1970s and find evidence that it raised attainment the most for the sons of poorly educated parents, which is to say, it reduced grade persistence. (Interestingly, for girls it appears that the reduction in persistence did not occur until primary school fees were eliminated.) Conversely, Hertz and Meurs (2007) argue that the collapse of state spending on education and the general economic crisis in Bulgaria in the mid-1990s were responsible for the sharp increase in educational persistence for young adults that occurred between 1995 and 2001. This increased persistence was accompanied by absolute reductions in attainment for the children of parents with an eighth grade education or less. Finally, our finding that the intergenerational transmission of educational status is lower in the Nordic countries, which echoes the results from the income-based research, is also suggestive of a role for government policy, given these countries' levels of political commitment to social welfare provision.

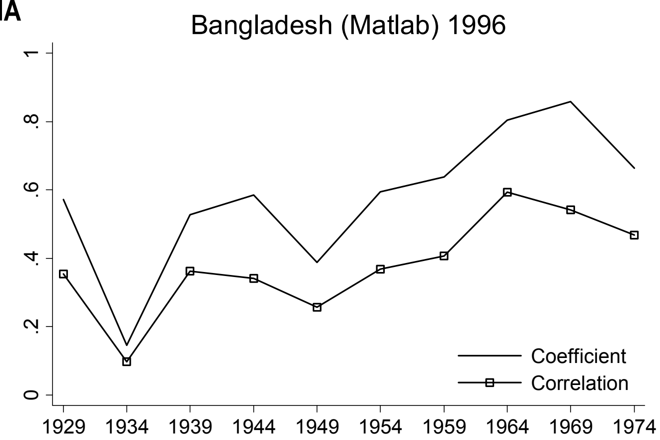
The most comprehensive empirical study of the determinants of educational persistence of which we are aware is that of Behrman, Birdsall, and Székely (2000), which covers sixteen Latin American nations, using an index based on grade-for-age schooling gaps among school-aged children. They demonstrate that greater government expenditures on primary schooling (per student of primary age) and higher average levels of education of the teachers both work to reduce their measure of educational persistence. Macroeconomic factors are also shown to play a role: in particular, greater financial depth (M2 over GDP), which they take as an indication of the degree to which the economy is "marketized," is associated with less intergenerational persistence of schooling.

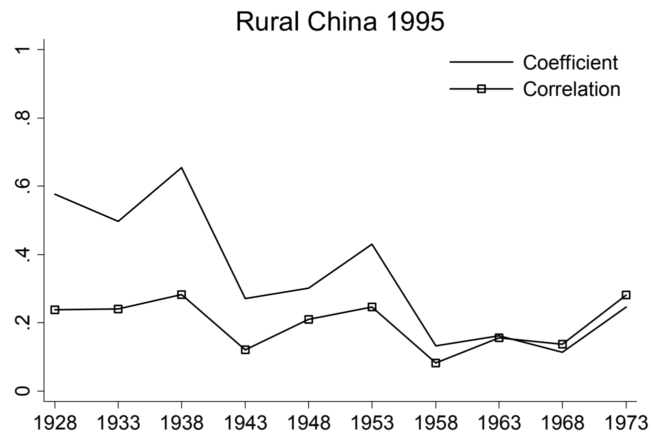
Less is known about the origins of long-run differences in educational persistence between nations, in particular, the striking, albeit slowly shrinking, difference that we have documented between Latin America and the rest of the world. One hypothesis that is currently being pursued by one of the authors is that the origins of this difference lie in the colonial past, and relate to the way public and private schooling systems operate in the presence (or absence) of deep ethnic divisions. More generally, we are looking for historical and institutional explanations for why the full effect of family background on one's access to education might vary across continents.

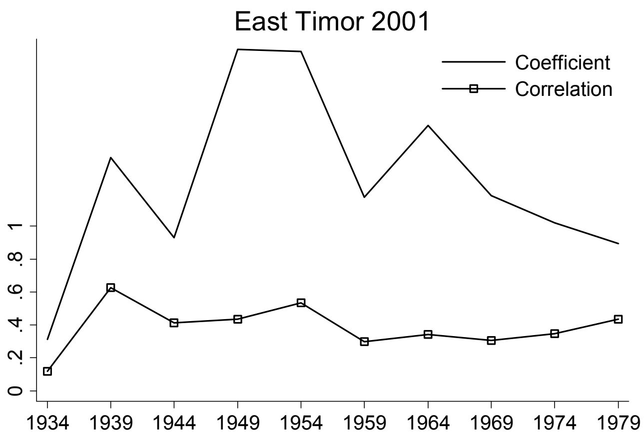
Similarly, there is little empirical evidence as to the *consequences* of educational persistence, which might plausibly be a determinant of economic growth, not merely an omnibus measure of intergenerational equity. One hypothesis we are seeking to test is that an excessive level of intergenerational transmission of inequality could act as a drag on economic growth, if it means that talented children from the lower social strata are denied the opportunity to reach their full economic potential.

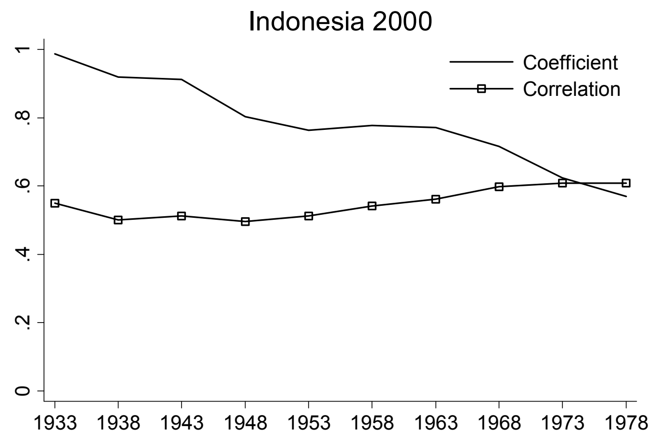
APPENDIX 1: INTERGENERATIONAL CORRELATIONS AND COEFFICIENTS, BY COUNTRY AND COHORT

ASIA









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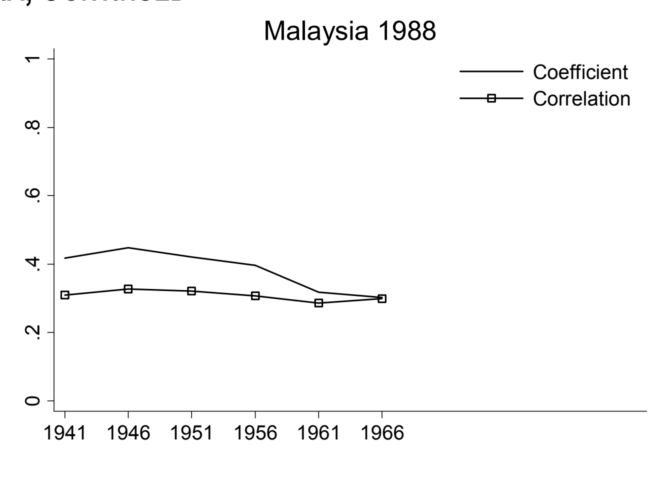

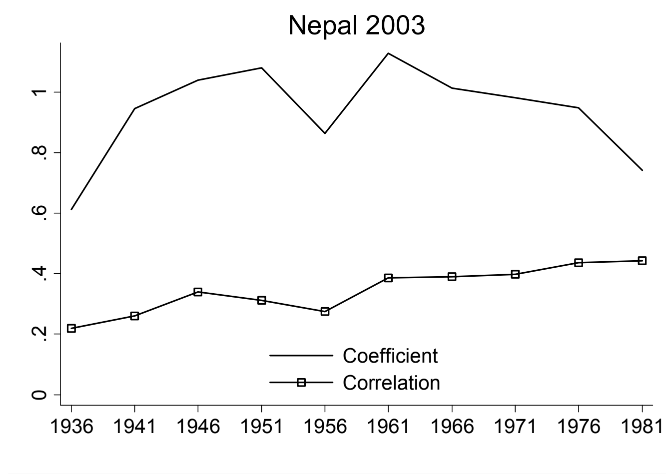

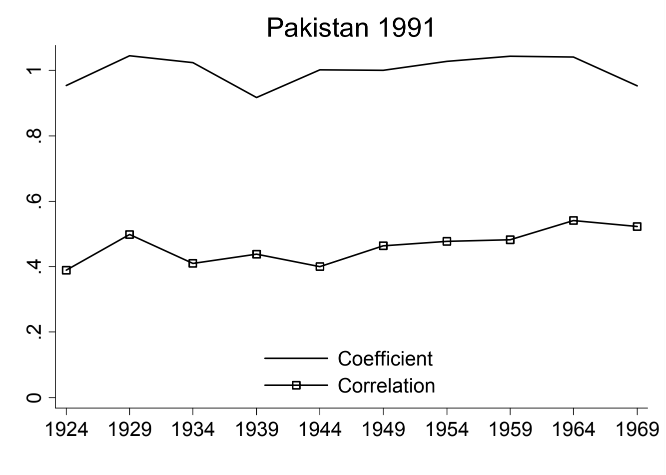

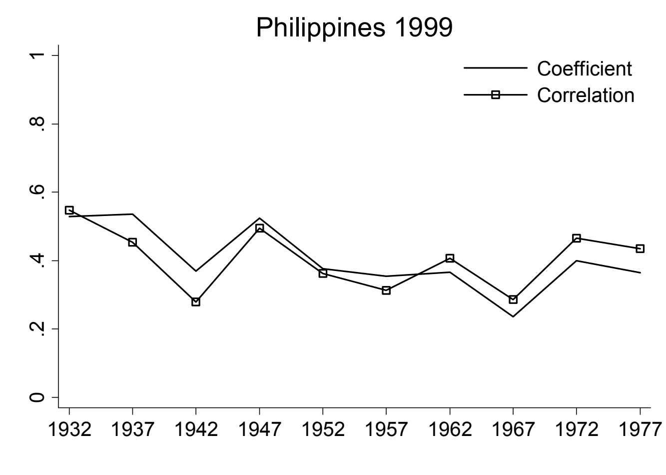

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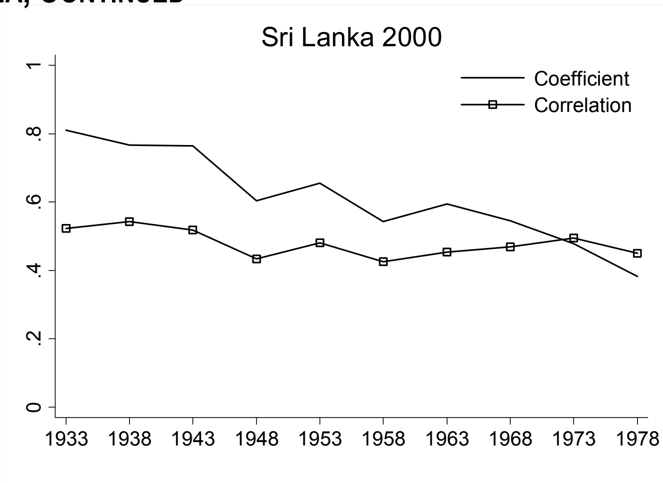
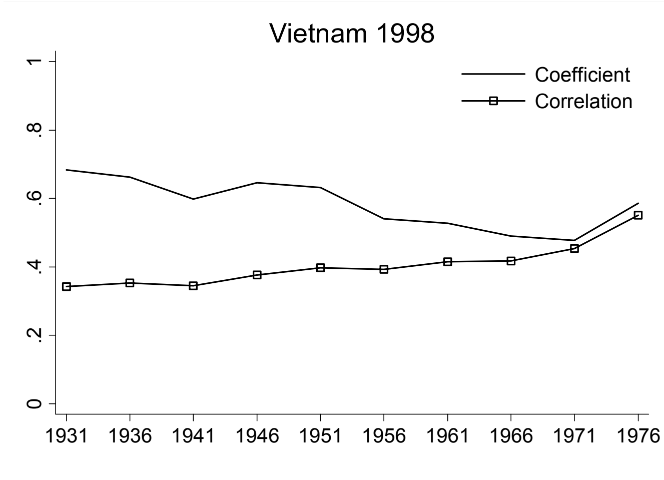
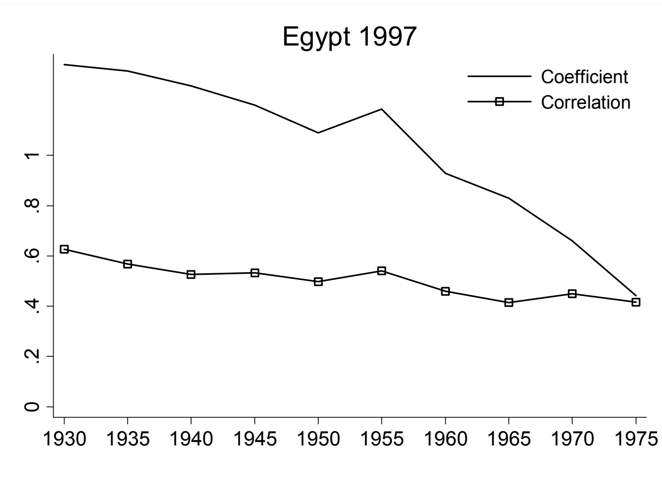




AFRICA


Egypt 1997
Coefficient
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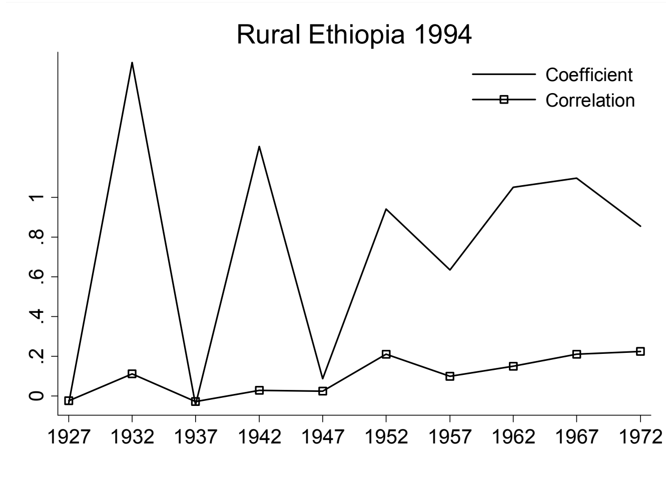
Rural Ethiopia 1994
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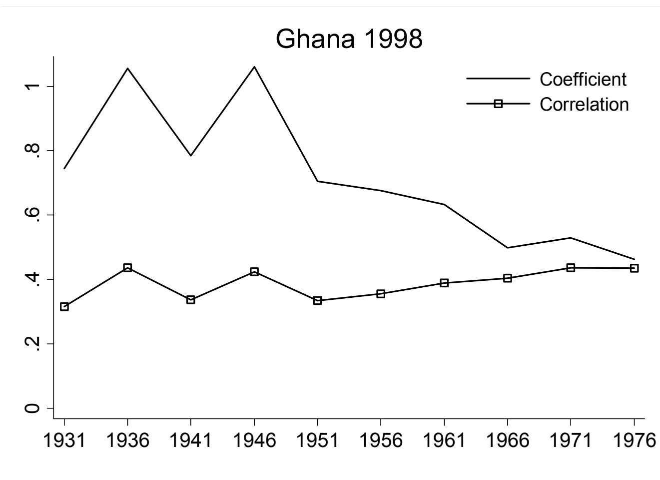
Ghana 1998
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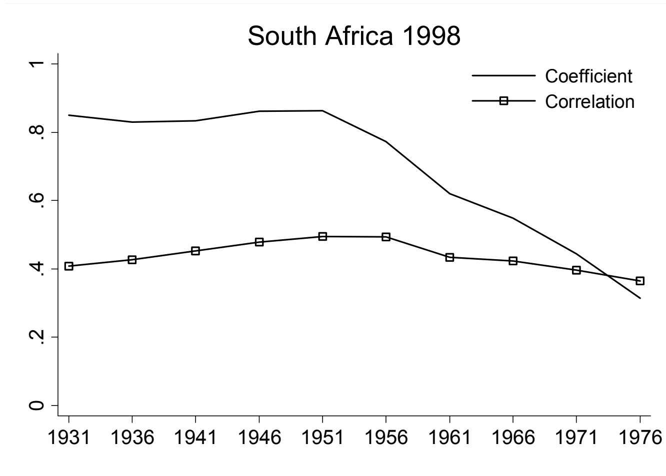
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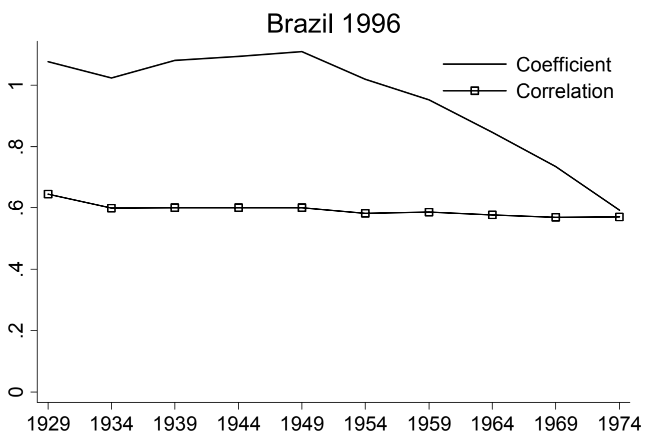
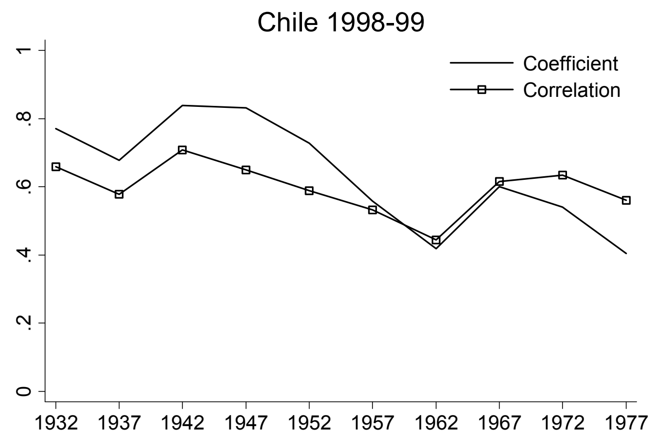
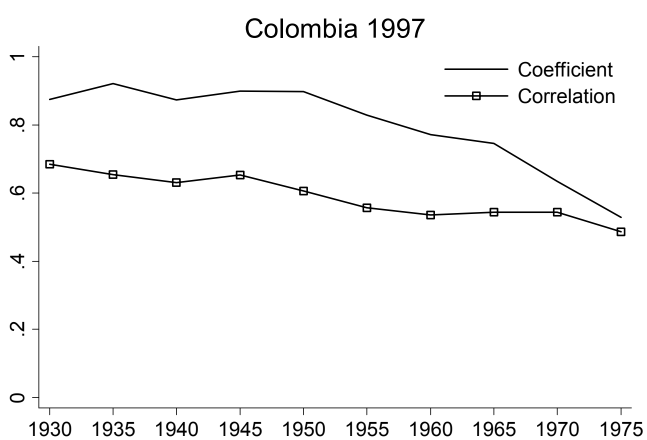
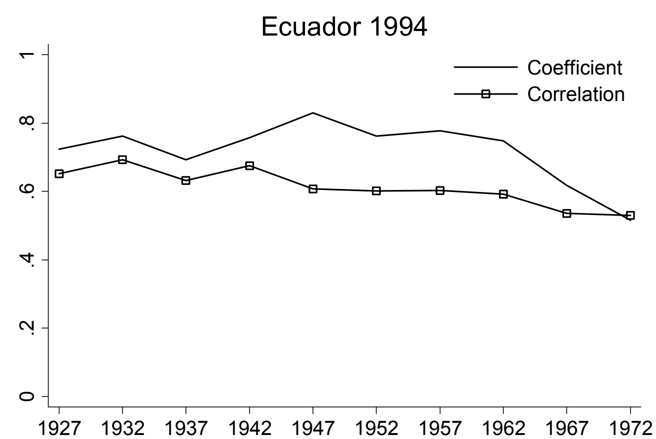
LATIN AMERICA









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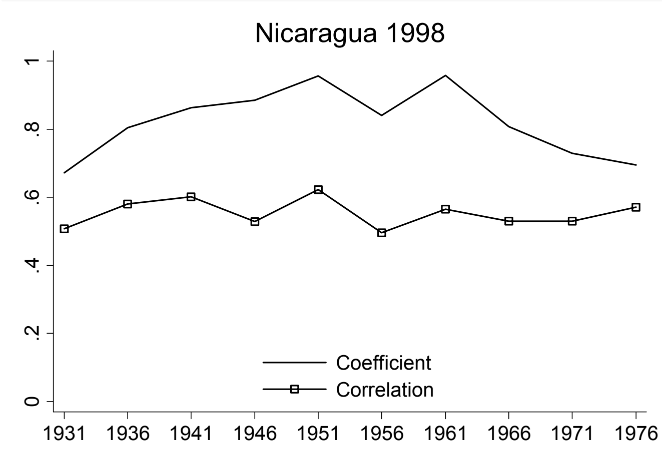
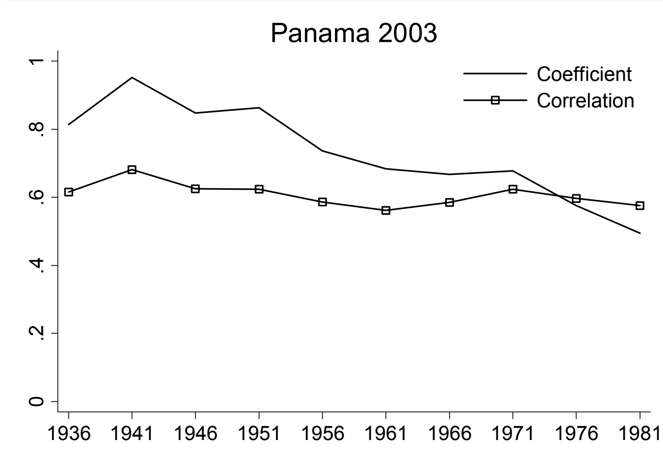
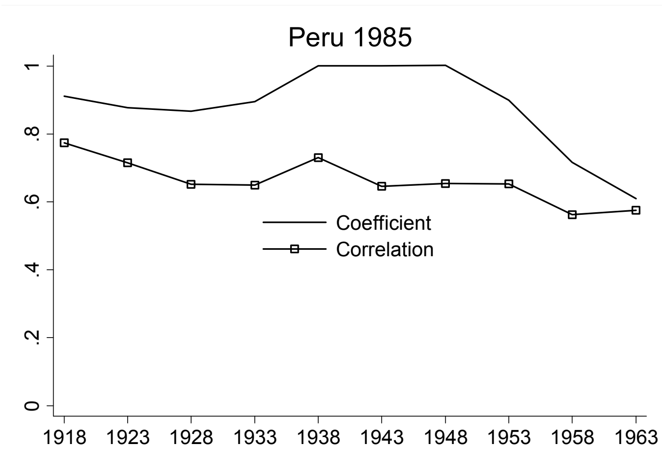
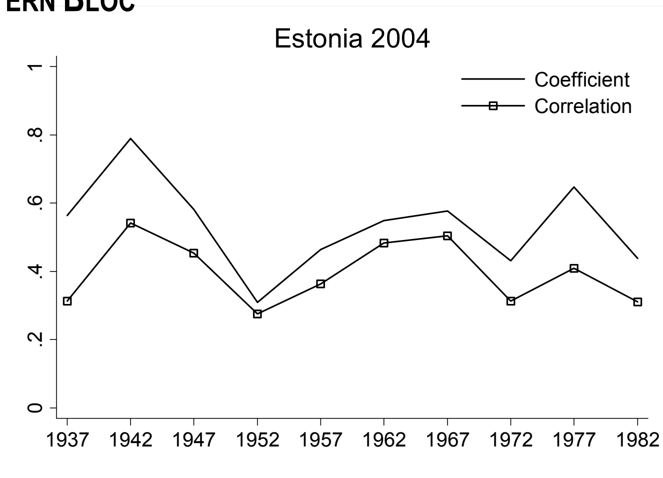






EASTERN BLOC



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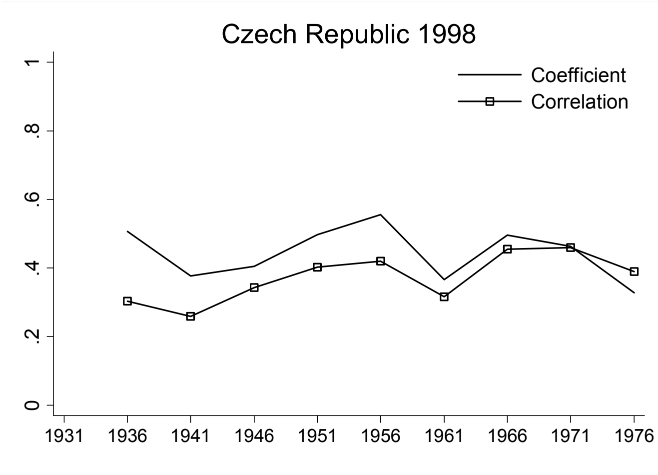

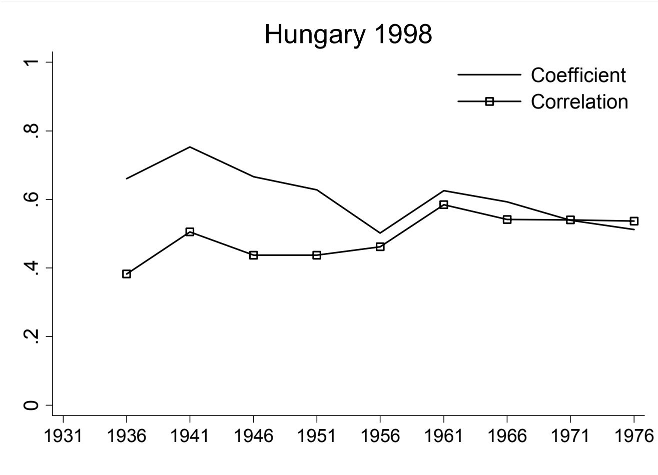

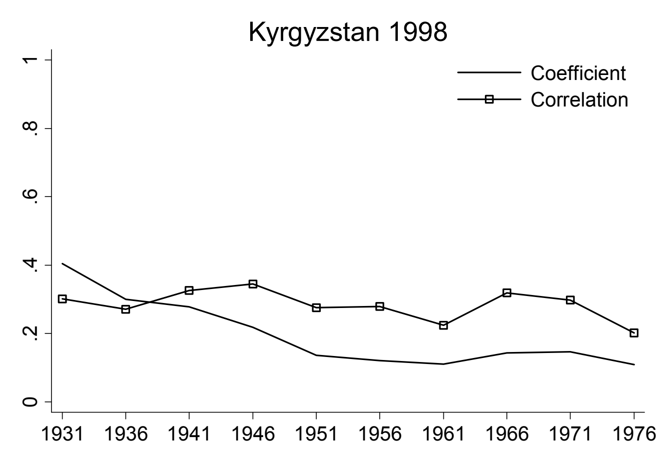

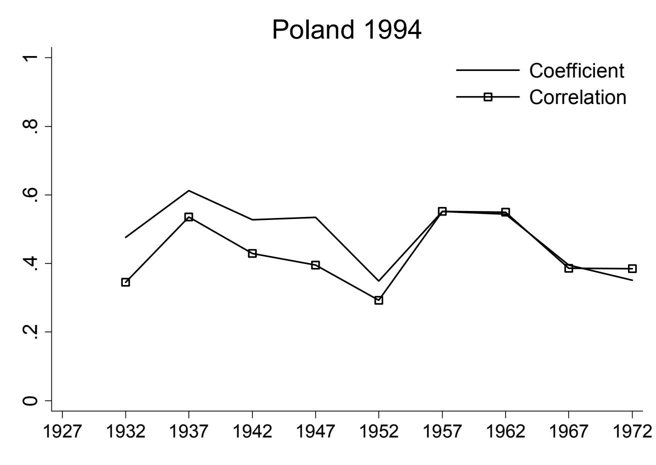

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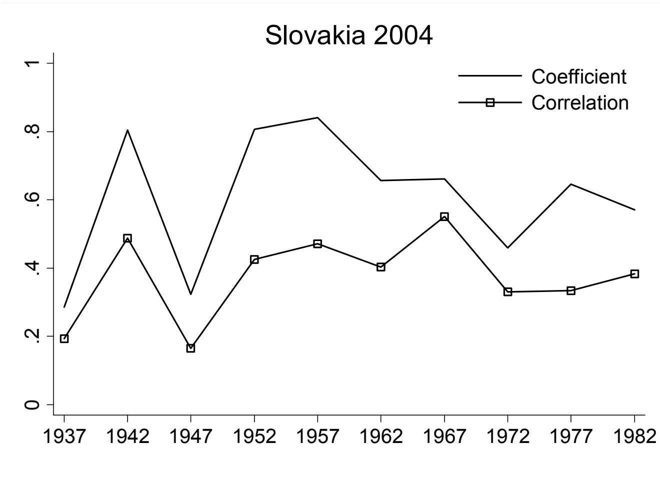
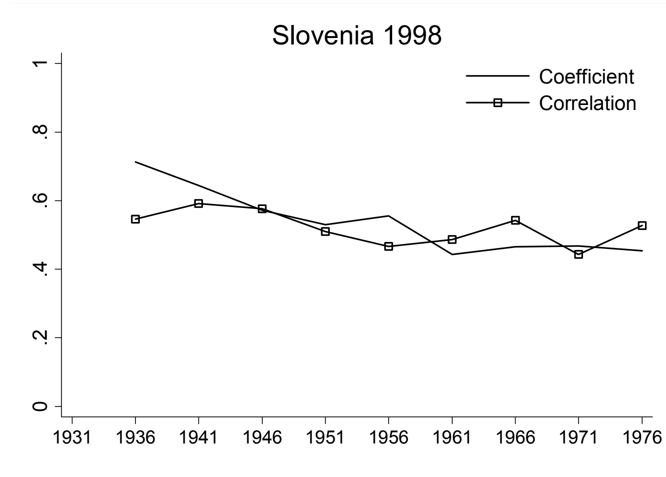
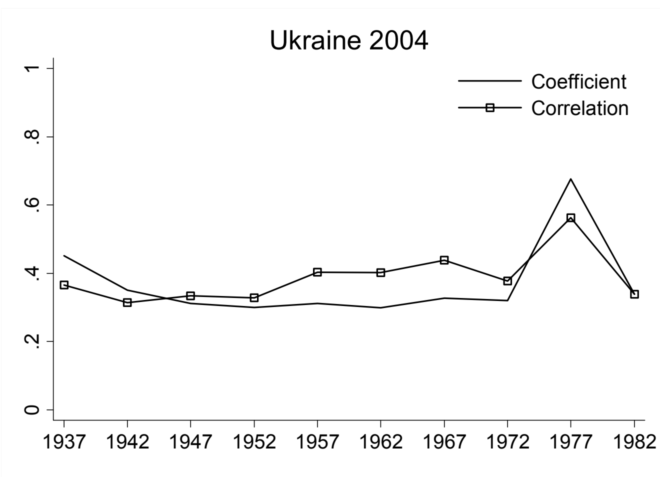
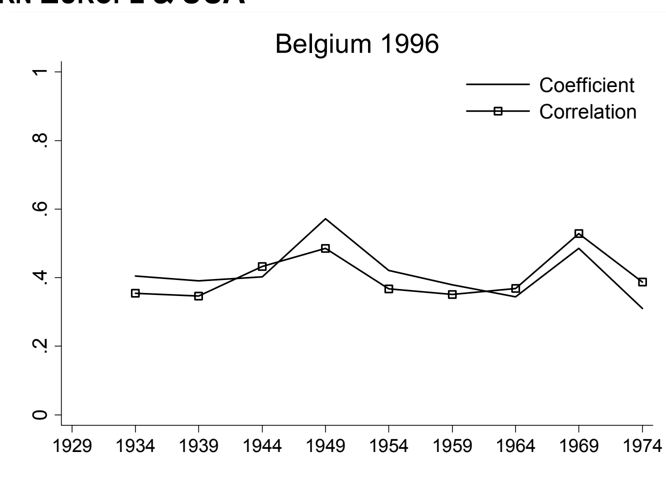






WESTERN EUROPE & USA



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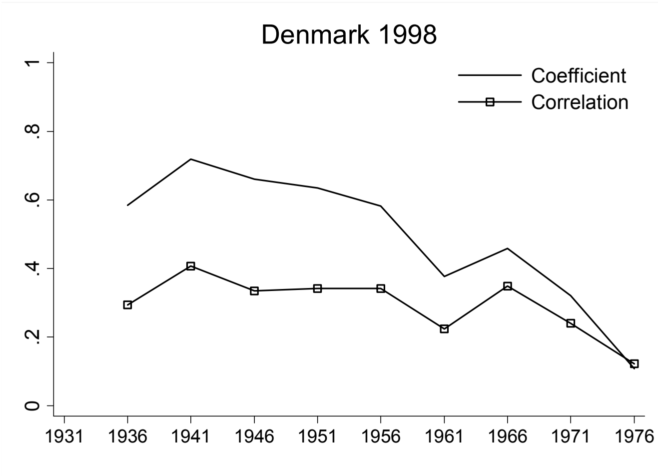

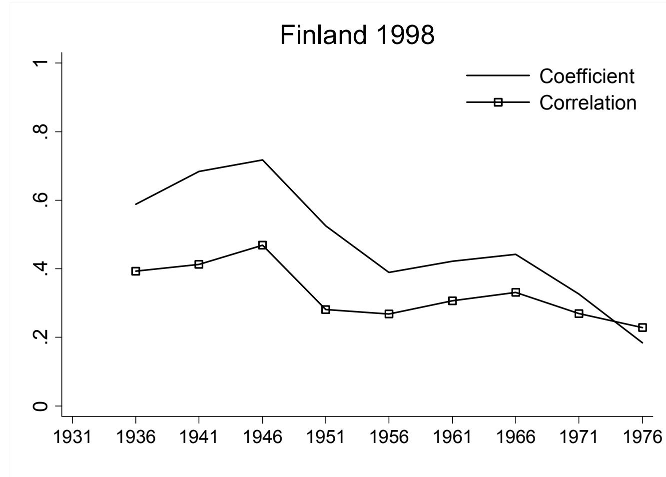

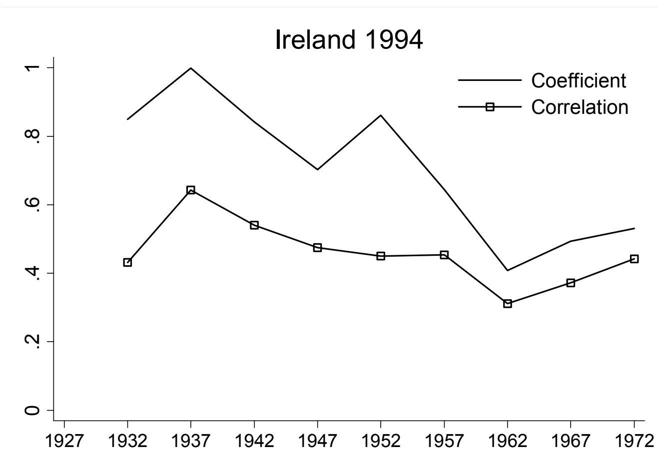

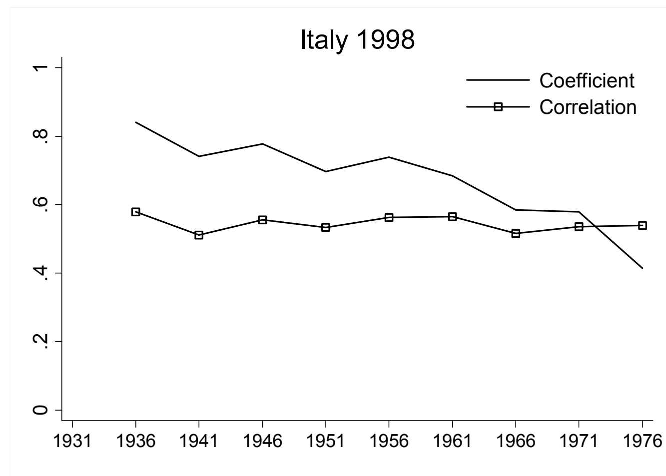

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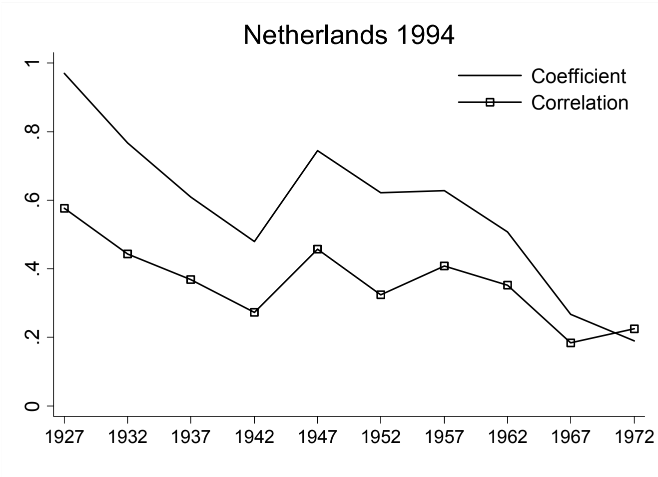

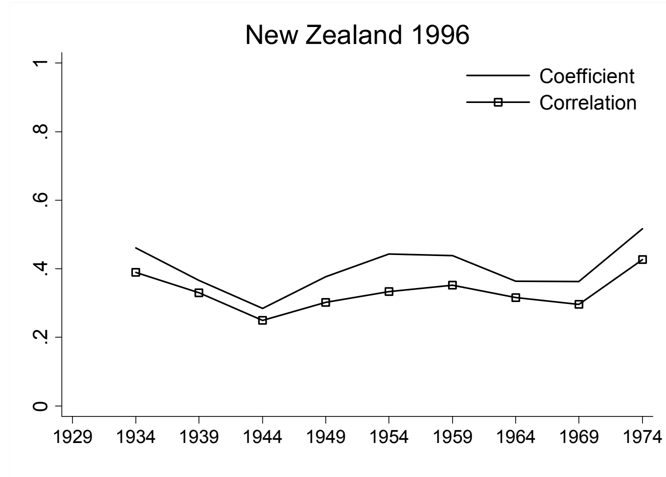

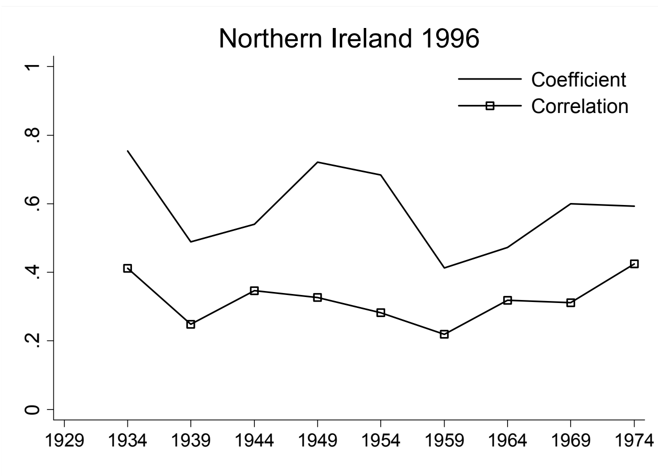

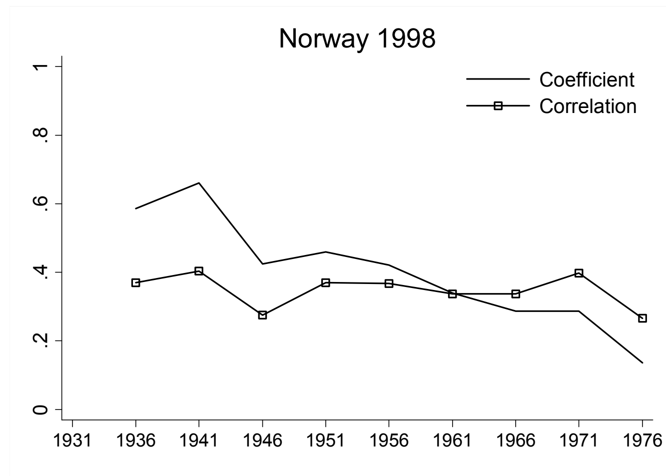

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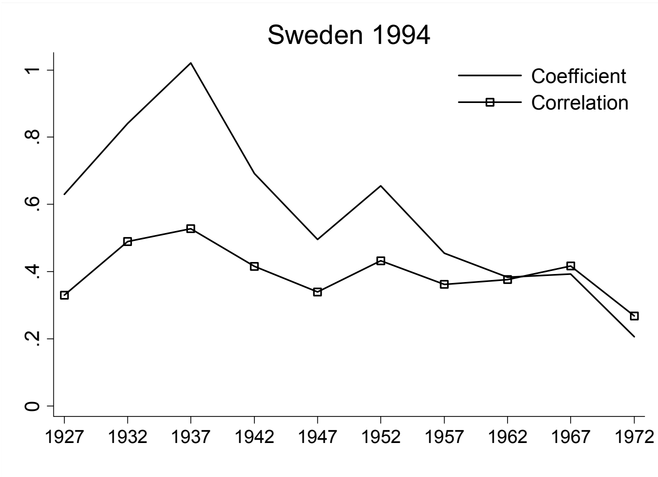

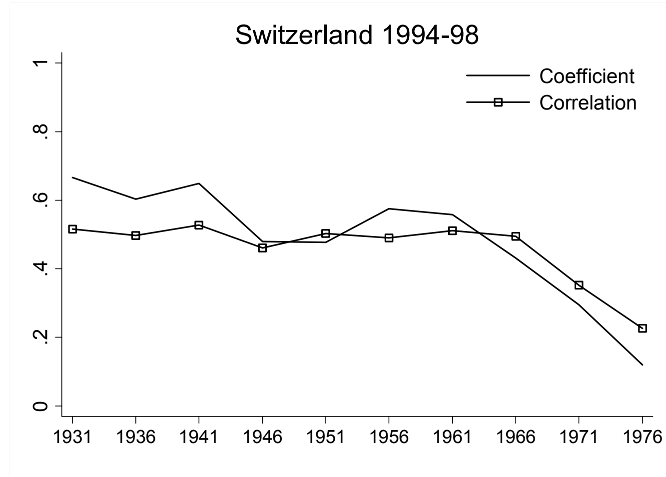

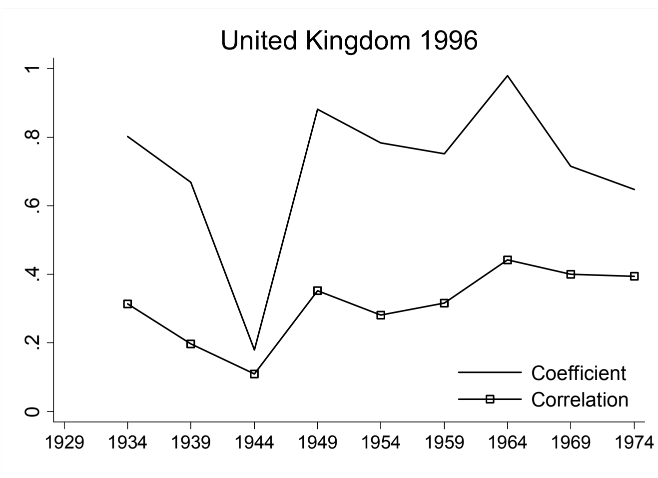

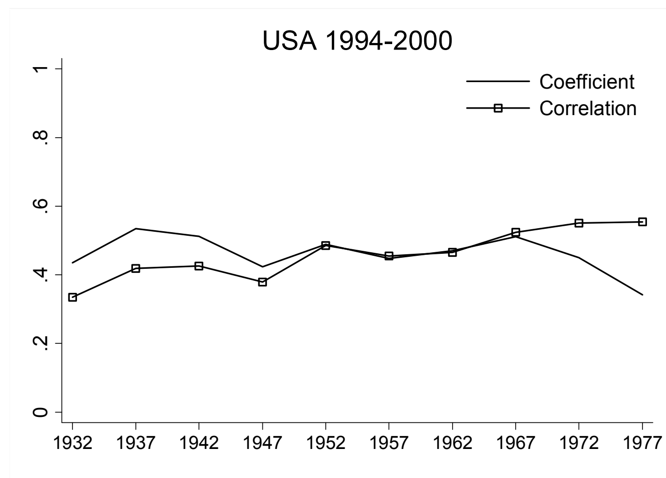

APPENDIX 2: SOURCES OF SURVEY DATA

Country	Year	Survey Name	Data Source
Bangladesh	1996	Matlab Health and Socio-Economic Survey	RAND Corp.
Belgium	1996	International Adult Literacy Survey	Statistics Canada
Brazil	1996	Pesquisa Nacional por Amostra de Domicílios	IBGE
Chile	1998 & 99	International Adult Literacy Survey & ISSP	Statcan & ISSP
China	1995	Living Standards Measurement Surveys	World Bank
Colombia	1997	Encuesta Nacional de Hogares	DANE
Czech Republic	1998	International Adult Literacy Survey	Statistics Canada
Denmark	1998	International Adult Literacy Survey	Statistics Canada
East Timor	2001	Living Standards Measurement Surveys	World Bank
Ecuador	1994	Living Standards Measurement Surveys	World Bank
Egypt	1997	Egypt Integrated Household Survey	IFPRI
Estonia	2004	European Social Survey	ESS
Ethiopia	1994	Egypt Integrated Household Survey	IFPRI
Finland	1998	International Adult Literacy Survey	Statistics Canada
Ghana	1998	Living Standards Measurement Surveys	World Bank
Hungary	1998	International Adult Literacy Survey	Statistics Canada
Indonesia	2000	Indonesia Family Life Survey	RAND Corp.
Ireland	1994	International Adult Literacy Survey	Statistics Canada
Italy	1998	International Adult Literacy Survey	Statistics Canada
Kyrgyzstan	1998	Living Standards Measurement Surveys	World Bank
Malaysia	1988	Malaysian Family Life Surveys	RAND Corp.
Nepal	2003	Living Standards Measurement Surveys	World Bank
Netherlands	1994	International Adult Literacy Survey	Statistics Canada
New Zealand	1996	International Adult Literacy Survey	Statistics Canada
Nicaragua	1998	Living Standards Measurement Surveys	World Bank
Northern Ireland	1996	International Adult Literacy Survey	Statistics Canada
Norway	1998	International Adult Literacy Survey	Statistics Canada
Pakistan	1991	Living Standards Measurement Surveys	World Bank
Panama	2003	Living Standards Measurement Surveys	World Bank
Peru	1985	Living Standards Measurement Surveys	World Bank
Philippines	1999	International Social Survey Programme	ISSP
Poland	1994	International Adult Literacy Survey	Statistics Canada
South Africa	1998	KwaZulu-Natal Income Dynamics Study	IFPRI
Slovakia	2004	European Social Survey	ESS
Slovenia	1998	International Adult Literacy Survey	Statistics Canada
Sri Lanka	2000	Sri Lankan Integrated Survey	Ministry of Finance
Sweden	1994	International Adult Literacy Survey	Statistics Canada
Switzerland	1994 & 98	International Adult Literacy Survey	Statistics Canada
Ukraine	2004	European Social Survey	ESS
United Kingdom	1996	International Adult Literacy Survey	Statistics Canada
USA	2000	International Social Survey Programme	Statcan & ISSP
Vietnam	1998	Living Standards Measurement Surveys	World Bank

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