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International Applications

Thus far, the empirical analysis of Part B has been concerned only with comparisons among regions within a country. The goal of this chapter is to discover whether the relationships found in the interregional data exist on an international level, too. To this end, the first section presents a comparative regression analysis for the United States, Canada, Puerto Rico, and Mexico, applying the procedures of the schooling model as in the preceding chapter. Next a different international data set is used to test the effect of schooling inequality and the level and growth rate of income per capita on the inequality of earnings of nonfarm males. In addition, the influences of some institutional arrangements and historical events on the parameters under review and on the mechanism through which income inequality is generated are examined, via the effects of mass immigration and distribution of capital (physical and human) in Israel and a comparison of Great Britain with the United States.

A FOUR-COUNTRY ANALYSIS

United States and Canada

Table 5-1 shows the results from regressing the log of earnings on schooling for males between the ages of twenty-five and sixty-

TABLE 5-1

Results from Regressing the Natural Log of Earnings on Schooling for Males, Twenty-five to Sixty-four, in the United States and Canada

	Summary Statistics				Regression: ln$E$ on S			
	SD(ln$Y$) (1)	SD(S) (2)	AV(ln$E$) (3)	AV(S) (4)	ln$E_{0,1}$ (5)	$\hat{r}_1$ (6)	Var$(U)_1$ (7)	R^2 (8)
United States	.70	3.66	1.54	10.31	.75 (.22)	.08 (.02)	.41	.15
White	.68	3.53	1.60	10.57	.83 (.25)	.07 (.02)	.40	.13
Non-South	.65	3.41	1.63	10.67	.94 (.23)	.06 (.02)	.38	.10
Non-South, white	.65	3.36	1.66	10.78	.96 (.26)	.06 (.02)	.38	.10
South	.76	4.03	1.32	10.42	.47 (.20)	.09 (.02)	.45	.22
South, white	.74	3.90	1.43	9.96	.60 (.23)	.08 (.02)	.44	.18
Canada, nonfarm	.68	3.38	1.36	8.70	.68 (.14)	.08 (.01)	.39	.15

Note: The U.S. data are based on a 5 per cent sample and contain 77 cells—7 schooling intervals and 11 earnings intervals. The Canadian data are for a 20 per cent sample of private nonfarm households and contain 78 cells—6 schooling intervals and 13 income intervals. Standard errors are in parentheses. See Chapter 4, pp. 50–51 for definitions of symbols.

four in the United States and nonfarm Canada.[1] Compared with the total South and the white South, Canada has lower inequalities of income, schooling, and residual income, as well as a lower rate of return and adjusted coefficient of determination. Compared with the total non-South and the white non-South, Canada has higher values for these parameters, except for the standard deviation of schooling. However, as indicated in Chapter 4, grouping causes a greater downward bias of this parameter in Canada than in the non-South. The total U.S. results are a weighted average of the North and South figures. The values of the parameters for the United States and Canada differ only very slightly (with one exception: the standard deviation of schooling), and no pattern emerges.

To summarize, these comparisons show that the inequalities of income, residual income, and schooling—as well as the rate of return—seem to be (1) higher in the South than in Canada, (2) approximately equal in Canada and the United States, and (3) higher in Canada than in the non-South.

1. The data are for individuals classified into schooling and income cells.

TABLE 5-1 (Concluded)

Regression: ln$E$ on £, \$, and H					
ln$E_{0,3}$ (9)	$\hat{r}_{£}$ (10)	$\hat{r}_{\$}$ (11)	$\hat{r}_H$ (12)	Var$(U)_3$ (13)	$\bar{R}^2_3$ (14)
.70	.09	.07	.08	.43	.13
(.48)	(.07)	(.06)	(.05)		
.84	.07	.07	.08	.42	.10
(.59)	(.09)	(.06)	(.06)		
1.06	.05	.06	.08	.39	.08
(.58)	(.08)	(.05)	(.05)		
1.09	.05	.06	.08	.39	.07
(.67)	(.09)	(.06)	(.06)		
.50	.08	.09	.09	.46	.20
(.40)	(.07)	(.07)	(.06)		
.66)	.07	.09	.09	.46	.16
(.50)	(.08)	(.07)	(.06)		
.63	.09	.07	.08	.40	.14
(.40)	(.06)	(.06)	(.05)		

Sources: (1) *U.S. Census of Population: 1960, Subject Reports. Occupation by Earnings and Education*, Washington, D.C., Tables 1, 2, and 3. (2) *Census of Canada: 1961, Population Sample. Incomes of Individuals*, Bulletin 4.1-2, Ottawa, Table B6.

Puerto Rico

Since Puerto Rico is a self-governing Commonwealth of the United States, the 1960 U.S. Census of Population includes an enumeration for the island. It contains a cross-classification of 1959 income data by schooling for all males, as well as urban males separately, twenty-five and over, with income.[2] Table 5-2 shows regression results based on these data for total and urban Puerto Rico, the United States, and nonfarm Canada, with the log of income as the dependent variable.

It is clear that Puerto Rico shows higher values than the United States for the three measures of inequality (income, residual income, and schooling), as well as for the rate of return, the explanatory power, and the education component. Presumably the same qualitative results would hold true if property income and aged males were excluded from all of the data. Note that, except for

2. *U.S. Census of Population: 1960, Vol. 1, Characteristics of the Population*, Washington, D.C., Part 53, p. LII and Table 117. The data are processed in the same manner as the U.S. data.

TABLE 5-2

Results from Regressing the Natural Log of Income on Schooling for Males, Twenty-five and Over, in Puerto Rico, the United States, and Canada

	Summary Statistics			Regression: $\ln Y$ on S		
	SD($\ln Y$)	SD(S)	$\frac{\text{AV}(\ln Y)}{\text{AV}(S)}$	$\ln Y_{0,1}$	$\hat{r}_1$	$\frac{\text{Var}(U_1)}{\bar{R}^2}$
	(1)	(2)	(3)	(4)	(5)	(6)
Puerto Rico	1.19	4.75	-0.05	-0.93	0.14	0.97
			6.17	(0.15)	(0.02)	0.32
Urban	1.14	5.01	+0.40	-0.62	0.13	0.89
			7.95	(0.16)	(0.02)	0.31
Canada	0.83	3.41	1.21	+0.40	0.10	0.58
			8.38	(0.09)	(0.01)	0.16
United States	0.91	3.95	1.34	+0.30	0.11	0.66
			9.80	(0.23)	(0.02)	0.20

Note: See notes to Tables 4-1 and 5-1. The Puerto Rico data are based on a 25 per cent sample of the population cross-classified in 117 cells, 9 schooling intervals, and 13 income intervals. Income is in thousand dollar units. A negative AV($\ln Y$) results from mean income being less than $1,000. Standard errors are in parentheses.

Sources: (1) *U.S. Census of Population: 1960*, Vol. 1, *Characteristics of the Population*, Part 53, Puerto Rico, Washington, D.C., Table 117. (2) *U.S. Census of Population: 1960, Subject Reports. Educational Attainment*, Washington, D.C., Table 6. (3) *Census of Canada: 1961*, Vol. 4, *Population Sample, Income of Individuals*, Bulletin 4.1-1, Ottawa, Dominion Bureau of Statistics, Table A.11.

schooling inequality, urban Puerto Rico seems to be less divergent from the United States than total Puerto Rico.

The data for Puerto Rico are defined in the same manner, and were obtained from the same sampling procedure, as those for the United States. Therefore, we can test the null hypothesis that the relationship between the education component and income inequality is the same in Puerto Rico as among the states.

This test is done in a three-stage procedure. First, income inequality is regressed on the education component for the states. Second, the observed value of the education component for Puerto Rico is inserted into the regression equation to obtain the predicted value of income inequality for Puerto Rico. If this predicted value is sufficiently different from the observed value, the hypothesis that the Puerto Rican data are derived from the same statistical universe as the data for the States is rejected. Whether the difference is sufficiently great is determined by a t-test. The third stage, then, is to compute the observed t-ratio, which is the difference between the observed and the predicted values for income inequality in Puerto Rico, divided by the standard error of prediction.

The computed (or observed) t-ratios are $t = 1.50$ for urban Puerto Rico and $t = 2.27$ for total Puerto Rico. For 49 degrees of freedom and a two-tailed test, the critical t-values are 2.0 for a 5 per cent level of significance and 1.7 for a 10 per cent level.

Thus, the relation between income inequality and the education component in Puerto Rico is consistent with the U.S. pattern for the urban male population, but not for the entire male population. This finding is quite reasonable, since the economic environment of urban Puerto Rico is more like that of the mainland than is that of the urban and rural sectors combined.

Mexico

Table 5-3 presents regression results for the log of average monthly earnings on schooling for a sample of 3,901 male workers taken in urban Mexico.[3] A comparison of Mexico (rows 1 and 2 of Table 5-3) with the United States and Canada (Table 5-1) indicates that the inequalities of income and schooling, the adjusted rate of return, and the coefficient of determination are all highest in Mexico.

Striking features of the Mexican data are the high adjusted rate

TABLE 5-3

Results from Regressing the Natural Log of Earnings on Schooling and Experience for Males in Urban Mexico

Age of Males	Summary Statistics			Regressions				
	SD(ln$E$) (1)	SD(S) (2)	$\frac{AV(\ln E)}{AV(S)}$ (3)	$\hat{r}$ (4)	b_1 (5)	b_2 (6)	Var(U) (7)	$\overline{R}^2$ (8)
All ages[a]	.89	3.98	6.94 6.57	.14 (.003)	—	—	.46	.42
Age 25+[a]	.79	4.26	7.25 6.79	.14 (.002)	—	—	.27	.56
All ages[b]	.89	3.98	6.94 6.57	.18 (.002)	.10 (.002)	-.0013 (.0000)	.22	.73

Note: E is in pesos per month, S is years of schooling, and T is age minus schooling minus five. Standard errors in parentheses.

Source: Sample provided by Martin Carnoy, including 3,901 individual, ungrouped observations, described in Appendix A-3.

[a] $\ln E_i = (\ln E_0) + (\hat{r}) S_i + U_i$.

[b] $\ln E_i = (\ln E_0) + (\hat{r}) S_i + b_1 T_i + b_2 T_i^2 + U_i$.

3. See Appendix A-3 for a discussion of the sample. Since the natural log of earnings is the dependent variable, the use of average monthly earnings (i.e., annual earnings divided by twelve) rather than annual earnings has no effect on the regression slope parameter or the measures of inequality.

of return and explanatory power of schooling. Even in Puerto Rico, where the adjusted rate of return is similar to that in Mexico, the explanatory power of schooling, while great, does not approach the magnitude of that for Mexico. The Mexican workers in the sample were all employed at the time and were asked to report their average monthly earnings. If, as seems plausible, they reported their full-time monthly earnings, then weeks worked were, in effect, held constant in the data. This may explain the high explanatory power of schooling.

The Mexican data contain the variable age (A), which permits an analysis of the effects of schooling (S) and years of labor market experience (T) ($T = A - S - 5$) on the natural log of earnings.[4] Two interesting conclusions emerge from this analysis (Table 5-3). First, when experience is held constant, the adjusted rate of return from schooling is increased from 14 to 18 per cent.[5] This implies that years of experience and schooling are negatively correlated. Second, the explanatory power of the model is increased from 42 to 73 per cent. That is, in this sample of urban Mexican workers nearly three-fourths of the individual variations in monthly earnings are explained by years of schooling and experience.

Summary

A comparison among the four countries indicates that the inequality of income is positively correlated with the rate of return, the inequality of schooling, and the explanatory power of schooling.

The proportion of the interregional variation in income inequality that can be explained by the education component, the residual variance, and their covariation can be computed using equations (3-21) and (3-22). For earnings data for the U.S. North, U.S. South, and nonfarm Canada, the education component ex-

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4. If it is assumed that the fraction of earnings invested in experience declines linearly, then the log of earnings is a parabolic function of years of experience. (See Mincer, *Schooling, Experience, and Earnings*, or Chapter 6 of this volume.) Experience is the number of years since leaving school, and is measured as age minus schooling minus five.

5. Recall from Chapter 3 that the slope coefficient of schooling is rk. Therefore, the rate of return is 0.14 when experience is not held constant, if $k = 1$. A $k = 1$ implies that the sum of direct plus opportunity costs of a year of schooling are equal to the earnings the student could have received in that year if he had not attended school. In other words, if the direct costs of the year of schooling equal the earnings the student received that year while attending school, then $k = 1$.

TABLE 5-4

International Explanatory Power of Schooling (per cent)

Regions	Education Component (1)	Residual (2)	Covariation (3)
U.S. North, U.S. South, Canada (earnings data)	30.2	20.9	48.8
United States, Canada, Puerto Rico (income data)	19.5	28.5	52.2
United States, Canada, Urban Puerto Rico (income data)	25.2	25.2	49.6

Note: See notes to Table 4-5.
Sources: Tables 5-1 and 5-2.

plains 30 per cent and the residual variance explains 21 per cent of the differences in earnings inequality among these three regions (see Table 5-4). For income data in the United States, Canada, and Puerto Rico, the education component explains 20 per cent of the differences in inequality. When the Puerto Rican data are restricted to urban males, the education component and the residual each explain 25 per cent. In a comparison of Table 5-4 with Table 4-5, we note that the education component's direct interregional explanatory power is smaller in the international data than among the fifty-one states (where it is 30.5 per cent).

EARNINGS INEQUALITY AND ECONOMIC GROWTH

Lydall performed the herculean task of collecting data and generating statistics on inequality of earnings and schooling of nonfarm males from several countries at various levels of development.[6] These data are analyzed here to provide additional evidence on the relationship between the distribution of schooling and the distribution of earnings.

Let us recall equations (3-13) and (3-14), which specified that earnings inequality is positively related to the level and inequality of schooling and rates of return from schooling.[7] To test this relationship, Lydall's data are used here for the inequality of earnings

6. Harold Lydall, *The Structure of Earnings*, Oxford, 1968, Chapter 5.
7. The equation is:

$$\text{Var}(\ln Y) = \bar{r}^2\ \text{Var}(S) + \bar{S}^2\ \text{Var}(r) + \text{Var}(r)\ \text{Var}(S) + \text{Var}(U).$$

and schooling. Data on GNP per capita are used as a proxy for the level of schooling. International data are scarce on the level of the rate of return from schooling, and nonexistent on the variance in rates of return.

Yet, there may be a relation between the level and inequality of rates of return from schooling and secular economic change. During periods of secular economic change, relative prices are in flux; if schooling promotes perceiving and adjusting to changed circumstances, it will raise the level of and inequality in the rates of return and thereby increase income inequality.[8] A discussion of the process through which secular economic change may influence the distribution of rates of return from schooling and the relevant empirical analysis are presented below.

Allocative Efficiency and Rates of Return

The skills created by school and postschool training can be thought of as falling into one of two not easily separable categories—worker efficiency and allocative efficiency.[9] Worker efficiency refers to the ability to perform a particular set of tasks; allocative efficiency refers to the ability to make correct decisions. Learning to use a drill press increases worker efficiency; learning how to decide on the appropriate grade of metal to drill increases allocative efficiency. Again, learning how to apply fertilizer to a field increases one's worker efficiency, while learning how to select a combination of chemicals to apply to a field increases one's al-

8. Note that this can be contrasted with the increase in income inequality during business cycle recessions. The decline in employment during recessions is experienced disproportionately by those with low levels of training. In terms of the schooling model of income distribution, the cyclical decline results in an increase in the level and inequality of rates of return from schooling computed from observed annual earnings. For theoretical and empirical analyses of the effect of the business cycle on the distribution of employment and income, see Becker, *Human Capital*, 1974, Part I; and Barry R. Chiswick and Jacob Mincer, "Time-Series Changes in Personal Income Inequality in the United States from 1939, with Projections to 1985," *Journal of Political Economy, Supplement*, May-June 1972.

9. The distinction was made originally by Theodore W. Schultz in *Transforming Traditional Agriculture*, New Haven, Yale University Press, 1964, Chapter 12. It was developed further by Richard Nelson and Edmund Phelps, "Investment in Humans, Technological Diffusion and Economic Growth," *American Economic Review*, May 1966, pp. 69-75; and Finis Welch, "Education in Production," *Journal of Political Economy*, January 1970, pp. 35-59. Welch uses U.S. agricultural data to show that schooling increases allocative efficiency.

locative efficiency. A particular job may, of course, involve varying combinations of allocative and worker efficiency. Allocative skills appear to be of relatively greater importance with higher levels of skill. Indeed, for most professionals, decision-making skills are of primary importance.

The demand for allocative skills is higher the more decisions there are to be made and the higher the payoff is for "better" decisions. In a long-term stagnant economy where relative prices are unchanged, the most efficient procedures would have long been discovered and the knowledge spread to all relevant members of the system. In this situation allocative skills would not be very valuable.[10]

Suppose we now introduce a development that changes relative prices. Those with the greatest allocative skills are the first to learn of this change and the first to implement an appropriate response; as a consequence they receive higher incomes—now allocative skills have achieved more economic value. If this is the only change (and a once-and-for-all change), the premium for this skill will decline as the knowledge spreads to those with lesser allocative skills.[11]

In an environment of continuous economic change, there is the possibility of economic gain by combining factors of production in a different way or by producing a different output. In such a situation premiums for superior allocative efficiency can (and presumably do) persist even in the long run. Allocative efficiency is, therefore, assumed to have a larger payoff the more change there is in the economic environment. Periods of secular economic growth are periods in which the economic environment is in flux and in which gains may be had from recombining factors of production.

Let us return to the simplified internal rate of return formula, $r = d/C$, where d is the annual increment in earnings due to an investment in schooling and C is the dollar cost of the investment. The differential d can, in principle, be decomposed into an allocative efficiency (d_a) and a worker efficiency (d_w) differential, $d = d_a + d_w$. The costs of the investment are the direct costs (C_d) and the opportunity costs, where the latter is the sum of an allocative

10. Schultz argues in *Transforming Traditional Agriculture*, p. 8, that the "critical feature of traditional agriculture is the low rate of return to investment in agricultural factors of the type that farmers have been using for generations." For an elaboration of this point, see ibid., Chapter 6.

11. For an analysis of the spread of an innovation, see Zvi Griliches, "Hybrid Corn and the Economics of Innovation," *Science*, July 29, 1960, pp. 275-280.

(C_a) and a worker efficiency (C_w) component. Then, $C = C_d + C_a + C_w$. The rate of return is written as:

$$r = \frac{(d_a + d_w)}{C_d + C_a + C_w}. \tag{5-1}$$

Suppose a change in the economic environment increases the return to allocative efficiency by 100t per cent. The new rate of return is

$$r' = \frac{(1 + t)\, d_a + d_w}{C_d + (1 + t)\, C_a + C_w}, \tag{5-2}$$

if direct costs are not affected by the change. This implies an increase in the level of the rate of return.[12] If we allow for individual differences in worker and allocative efficiency, the variance in rates of return is likely to increase when the premium for allocative skills increases.[13]

Thus, it is hypothesized that more rapid economic change increases opportunities for a larger premium to allocative skills. Since allocative skills differ among individuals, and since these skills are likely to increase in relative importance for those with higher levels of skill, the expanded opportunities increase both the level and variance in rates of return from schooling—and consequently income inequality.[14] Hence, since economic growth is associated

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12. The rate of return increases as long as $r' > r$, $\frac{(1 + t)\, d_a + d_w}{C_d + (1 + t)\, C_a + C_w} > \frac{d_a + d_w}{C_d + C_a + C_w}$, or $\frac{d_a}{d_w} > \frac{C_a}{C_d + C_w}$. If C_d is nonnegative, $\frac{d_a}{d_w} > \frac{C_a}{C_w}$, if allocative skills rise in importance as skill level increases. Thus, r' is greater than r as long as allocative skills do not fall in relative importance as skill level increases.

13. This is easy to show if we make several assumptions. Let us assume those without the investment have no allocative efficiency, and the increased premium is a constant proportion 100t per cent. The new rate of return is $r' = \frac{(1 + t)\, d_a + d_w}{C}$. The variance in rates of return is $S^2(r) = (1/C)^2 [S^2(d_a) + S^2(d_w) + 2\,\mathrm{Cov}(d_a, d_w)]$, and $S^2(r') = (1/C)^2 [(1 + t)^2 S^2(d_a) + S^2(d_w) + 2(1 + t)\,\mathrm{Cov}(d_a, d_w)] = (1/C)^2 [S^2(d_a) + S^2(d_w) + 2\,\mathrm{Cov}(d_a, d_w)] + (1/C)^2 [(t^2 + 2t)\, S^2(d_a) + 2t\,\mathrm{Cov}(d_a, d_w)]$. For a positive t, $S^2(r')$ is greater than $S^2(r)$ as long as $(t + 2)\, S^2(d_a) + 2\,\mathrm{Cov}(d_a, d_w) > 0$. This necessarily holds if individual differences in allocative and worker skills are not negatively correlated.

14. The scanty international data that are available on rates of return support this hypothesis. Although there appears to be no relation between the level of output per capita and the rate of return, in Latin America at least,

with a changing economic environment, growth may be a *cause* of greater inequality in labor market income.

Empirical Analysis[15]

We can now test the hypotheses that earnings inequality is greater the larger the inequality of schooling, the higher the level of schooling, and the more rapid the secular economic change.

Lydall's measure of earnings is money wages and salaries before taxes for nonfarm males.[16] He uses three percentile measures of relative income inequality. The percentile measure of income inequality $P(X)$ is the ratio of the income of the individual in the X^{th} percentile from the upper end of the distribution ($p(X)$), relative to the median income (i.e., $p(50)$), multiplied by 100.[17]

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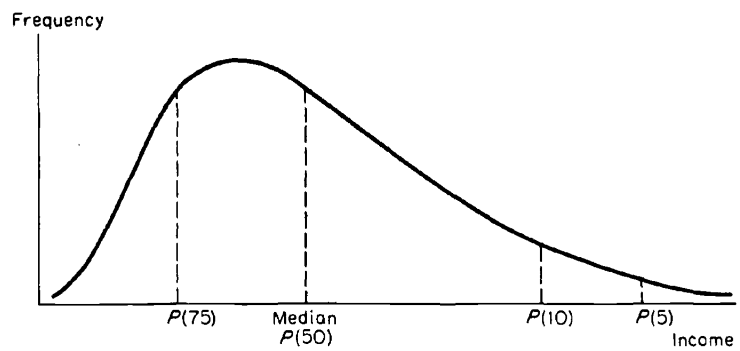
rates of return are higher in the more rapidly growing countries. (See Martin Carnoy, "Rates of Return to Schooling in Latin America," *Journal of Human Resources*, Summer 1967, pp. 354-374, and T. Paul Schultz, "Returns of Education in Bogota, Colombia," *RAND* Memorandum, Santa Monica, 1968, Table 9.) Using time series data, Schultz found that the partial correlation between income inequality and the *secular* growth rate of output was positive in the United States and the Netherlands, although it was significant only for the latter. ("Secular Trends and Cyclical Behavior of Income Distribution in the United States: 1944-1964," in Lee Soltow, ed., *Six Papers on the Size Distribution of Income and Wealth*, New York, 1969, p. 87.) In his cross-sectional analysis of farm family income in the United States, Gardner used "research and extension expenditures and the growth of output per farm" to serve as an "indicator of the dynamism of a state's agriculture." When schooling inequality, migration, and several other variables were held constant, Gardner's variables for "dynamism" had a positive effect on income inequality. (Bruce R. Gardner, "An Analysis of U.S. Farm Family Income Inequality, 1950-1960," Ph.D. dissertation, University of Chicago, 1968, pp. 62-63 and Tables 6 and 7.)

15. The empirical analysis is based on my "Earnings Inequality and Economic Development," *Quarterly Journal of Economics*, February 1971, pp. 21-39.

16. *The Structure of Earnings*, pp. 60 and 153.

17. A Schematic Representation of the Distribution of Income



($P(X) = [p(X)(100)]/[p(50)]$.) $P(5)$ and $P(10)$ are measures of inequality in the upper half of the distribution, their values exceed 100, and larger values imply greater inequality. $P(75)$ is a measure of inequality in the lower half of the distribution, has values less than 100, and larger values imply less inequality. $P(5)$ (or $P(10)$) and $P(75)$ are highly negatively correlated, which implies that regions with greater inequality in the upper half of the earnings distribution tend to have greater inequality in the lower half, and therefore tend to have greater overall earnings inequality.[18]

Lydall's measure of schooling inequality is the "Lorenz coefficient," ($LC(S)$), a measure of relative rather than absolute dispersion. He computed it for nonfarm males in ten countries for which he calculated personal earnings inequality.[19] Data on the level of schooling are not available, so GNP per capita is used as a proxy. The average per capita GNP is in constant 1967 prices (dollar equivalents) for the 1950–1960 period, except for Japan, where the period covered is 1952-1960.[20] The secular rate of change is measured by the per cent change in per capita GNP in constant 1967 prices (dollar equivalents) for the same periods.

The cross-sectional data on level and growth rate of output are available for only nine of the ten countries for which Lydall calculated Lorenz coefficients of schooling, and for two of these there are no estimates of $P(75)$. Thus, the empirical analysis of the effects of income level (Y'), growth rate of output ($\%\Delta Y$), and schooling inequality ($LC(S)$) on $P(5)$ and $P(10)$ is performed for nine countries, and on $P(75)$, for seven countries.[21] Due to extremely small samples the results can only be suggestive.

Table 5-5 contains the analysis of earnings inequality for the upper tail of the distribution, $P(5)$ and $P(10)$. The simple correlations between average income and $P(5)$ and $P(10)$ are negative and

18. The correlation coefficients for twenty-two countries are $R[P(5), P(10)] = 0.98$, $R[P(5), P(75)] = -0.80$, $R[P(10), P(75)] = -0.83$.

19. Lydall, *The Structure of Earnings*, pp. 209-211.

20. U.S. Agency for International Development, *Gross National Product, Growth Rates and Trend Data by Regions and Countries*, Documents No. RC-W-138, April 25, 1969 (Statistics and Reports Division, Office of Program and Policy Coordination). An index of per capita consumption constructed by W. Beckerman and R. Bacon ("International Comparisons of Income Levels: A Suggested New Measure," *Economic Journal*, Vol. 76, September 1966, pp. 519-536) was tried as a substitute for GNP per capita, but the results were not significantly different and are not reported.

21. $LC(S)$, the Lorenz coefficient of schooling, is a measure of relative inequality; higher coefficients imply a greater inequality of schooling.

The seven are Argentina, Canada, France, Japan, Mexico, United Kingdom, and United States. The two additional countries are Brazil and Chile.

TABLE 5-5

Analysis of $P(5)$ and $P(10)$

	$P(5)$	$P(10)$	Y'	$\%\Delta Y$	$LC(S)$	$\bar{R}^2$
			Correlation Matrix			
$P(10)$	0.97					
Y'	-0.67	-0.68				
$\%\Delta Y$	0.03	0.02	-0.26			
$LC(S)$	0.80	0.73	-0.61	-0.27		
			Regressions			
Dependent variable $P(5)$			-0.016 (-0.43)	87.3 (0.66)	0.339 (2.12)	0.56
Dependent variable $P(10)$			-0.017 (-0.75)	21.1 (0.26)	0.137 (1.36)	0.40

Note: The larger $P(5)$ and $P(10)$, the larger the earnings inequality. There are 9 observations; t-ratios are in parentheses. Critical values for correlations $R_{0.95} = 0.58$; for regression coefficients $t_{0.95} = 2.02$. Symbols are defined as follows:

$P(i) = \dfrac{100\,pi}{p_{50}}$, where pi is earnings at the ith percentile from the upper end of the distribution, and p_{50} is the median, for nonfarm males.

Y'—average per capita GNP in constant 1967 prices (dollar equivalents) for 1950-1960 (Japan, for 1952-1960).

$\%\Delta Y$—percentage change in GNP per capita in constant 1957 prices (dollar equivalents) from 1950 to 1960 (Japan, 1952 to 1960).

$LC(S)$—Lorenz coefficient of quantity of schooling.

See footnote 21 for a list of countries involved.

Sources: $P(i)$—Harold Lydall, *The Structure of Earnings* (Oxford, 1968), p. 153, Table 5.5.

Y' and $\%\Delta Y$—*Gross National Product, Growth Rate and Trend Data by Regions and Countries*, Document No. RC-W-138 (April 25, 1969), Statistics and Reports Division, Office of Program and Policy Coordination, U.S. Agency for International Development.

$LC(S)$—Lydall, *The Structure of Earnings*, p. 211, Table 7.1.

significant. When the rate of output growth and the Lorenz coefficient of schooling are held constant, however, the partial relation remains negative but becomes insignificant. Thus, although countries with lower levels of development are associated with greater earnings inequality, low income does not seem to have a direct effect, but operates through other variables (like the inequality of schooling, for example).

The growth rate of output has positive but insignificant simple and partial correlations with $P(5)$ and $P(10)$. The positive simple correlations of schooling inequality with $P(5)$ and $P(10)$ are significant, and although the significance declines, the partial relation remains positive, and is significant at the 5 per cent level for $P(5)$.

The dependent variable $P(75)$ declines with greater earnings inequality. Thus, the positive simple correlation shown in Table

TABLE 5-6

Analysis of $P(75)$

	$P(75)$	Y'	$\%\Delta Y$	$LC(S)$	$\bar{R}^2$
			Correlation Matrix		
Y'	0.63				
$\%\Delta Y$	-0.68	-0.43			
$LC(S)$	-0.47	-0.48	-0.24		
			Regressions		
Dependent variable $P(75)$		-0.0008 (-0.46)	-24.55 (-3.81)	-0.027 (-3.05)	0.81

Note: The smaller $P(75)$, the larger the earnings inequality. There are 7 observations; t-ratios are in parentheses. Critical values for correlations $R_{0.95} = 0.67$; for regression coefficients $t_{0.95} = 2.35$. For definition of symbols, see note to Table 5-5.

Sources: See footnote to Table 5-5.

5-6 between $P(75)$ and average income means that a smaller earnings inequality is associated with a higher level of income. This coefficient is barely significant at a 5 per cent level. The partial relation between Y' and $P(75)$ is not significant. The growth rate of output has a significant, and schooling inequality, an insignificant negative simple correlation with $P(75)$. Schooling inequality becomes highly significant and the growth rate of output increases in significance in the multiple regression.

The multiple regression analysis for $P(75)$ gives the expected signs for the three independent variables, and the schooling inequality and growth rate variables are significant. The results for $P(5)$ and $P(10)$ are weaker, but the signs for schooling inequality and the growth rate of output are consistent with our expectations. The differences between the $P(75)$, the $P(5)$, and $P(10)$ regressions may be due to the different samples or to the different aspects of inequality reflected in the three measures. To separate these effects, the regressions for $P(5)$ and $P(10)$ were run for the seven countries used in the $P(75)$ analysis. Although the results for $P(5)$ and $P(10)$ become more consistent with our expectations, they are not as consistent as the $P(75)$ results.[22] Thus, the differences between the $P(75)$ seven-country and the $P(5)$ and $P(10)$ nine-country analyses are partly due to the different sections of

22. For $P(5)$ the seven-country partial slope coefficient for schooling inequality is positive and significant ($t = 2.59$), and insignificantly positive for the growth rate of output ($t = 1.49$) and the level of output ($t = 0.29$). For $P(10)$ the three coefficients are positive, significant for schooling inequality ($t = 2.83$), but insignificant at the 5 per cent level (one-tailed test) for the growth rate of output ($t = 1.96$) and level of output ($t = 0.23$).

the earnings distribution encompassed by these variables. $P(75)$ measures dispersion closer to and below the median, while $P(5)$ and $P(10)$ reflect the extreme upper end of the distribution. The model, therefore, seems a better predictor for the center of the distribution of earnings than for the upper tail.

Conclusions

Although it is based on a very small sample, the international cross-sectional analysis with data on earnings inequality of nonfarm males tends to confirm the hypotheses derived from the schooling model of income distribution. The observed negative simple correlation between the earnings inequality of nonfarm males and the level of development does not appear to be directly due to development, since the partial relation is insignificant. If the level of development has an effect on earnings inequality it operates through other variables. A greater inequality of schooling and a higher growth rate of output appear to increase earnings inequality. The model is more successful in explaining earnings inequality around the median than between the median and the upper tail of the distribution.

MASS MIGRATION AND INCOME INEQUALITY IN ISRAEL

The variables examined in this volume are affected by historical events. It has been well established, for example, that historical events shaped the pattern of income distribution among the Jewish population of Israel. The exogenous changes in labor supply due to mass migration resulted in changes in Israeli income inequality, the rate of return from schooling, and the correlation of income with schooling which are consistent with our model.

Trends over Time

At the time of Israel's establishment in 1948, its Jewish population was composed primarily of immigrants from Europe.[23] Their migration was largely motivated by noneconomic considerations.[24] They came in large numbers to an underdeveloped econ-

23. See A. Hovne, *The Labor Force in Israel*, Jerusalem, Falk Project for Economic Research in Israel, July 1961, p. 19.

24. R. Bachi, "Immigration into Israel," in B. Thomas, ed., *Economics of International Migration*, London, Macmillan, 1958, pp. 315-318.

omy and often found that their previous training was of little economic value. The result: a low skill differential and, consequently, a low rate of return from schooling and a low correlation between earnings and schooling.[25]

Few immigrants brought capital with them, and previous Jewish settlement had been too recent for the development of a substantial group of property owners. National, public, and cooperative ownership of land and development enterprises tended to dominate the economic scene. The result was a small inequality among Jews in the distribution of nonlabor wealth and property income. The low rate of return from schooling and small inequality of property income led to a small inequality of total income.[26]

The residual variance, in turn, was affected by the small inequality of wealth and low rates of return from human capital. Skill differentials were sufficiently low to permit residual factors (such as ability and luck) to exert a relatively important effect on income distribution and to result in a small explanatory power of schooling.

Thus, it appears that the exogenous increase in the supply of skilled workers depressed the rate of return from schooling and generated a small inequality of income and a low correlation between income and schooling—a pattern which is consistent with our model.

The mass immigration after independence in 1948 once again was not motivated by conventional economic forces.[27] The new immigrants had less schooling than the immigrants who had come in the 1920s and 1930s.[28] They were also at a disadvantage as to other forms of human capital: many came from refugee camps in Europe and poor countries of the Middle East and North Africa, and we may assume that their level of health was lower than that of prestatehood immigrants. Moreover, knowledge of the language and "way of life" of a country and information concerning labor markets are also part of human capital. Hovne finds that "prestate immigrants of the same age, level of education, and continent of

25. Giora Hanoch, "Income Differentials in Israel," *Fifth Report, Falk Project for Economic Research in Israel*, 1961, p. 43.

26. Hanoch (ibid., Chapter 1) shows that in 1957-58, income inequality was relatively small, yet larger than nearly a decade earlier.

27. Bachi, "Immigration into Israel," pp. 319-320.

28. In 1957 the mean and variance of years of schooling for all male Jews of fifteen years and over were 6.6 and 17.9, respectively, and for those designated as "veterans" (in Palestine in 1947), 7.8 and 17.4, respectively.

The five schooling intervals were assigned the following numerical values:

origin were at an advantage,"[29] presumably because they had already acquired these forms of capital.

The increased supply of unskilled workers decreased the average level of schooling and other human capital, and changed relative scarcities.[30] As would be expected, skill differentials increased. Bahral reports that "the relative wage differentials of workers performing different jobs (when comparing high and low wage groups), on the average, widened during the first ten years of the state. This is true for the first years of mass immigration and for the period of the second wave of immigration, 1955-58."[31] He adds that "this relative price of highly paid labor services in Israel should be stressed in view of the downward trend of occupational differentials found in most modern economies and in mandatory Palestine up to the end of the second World War."[32] Hanoch, too, shows that income inequality increased in the first decade after independence.[33]

Since the average rate of return increased and the inequality of schooling remained nearly unchanged (see footnote 29 below), the education component (the product r^2 Var (S)) increased. It seems plausible that the residual variance increased too, due to higher returns on other forms of capital and a possible growth in the inequality of human capital other than schooling. The effects of some variables in the residual (ability and luck, for example) probably prevented it from rising at the same rate as the education component, and thereby resulted in an increase in the correlation between income and schooling. This is consistent with Hanoch's

Schooling Interval	Years
1. Did Not Attend School	0.0
2. Partial Primary Education	3.0
3. Completed Primary Education	7.0
4. Completed Secondary Education	12.0
5. Completed Higher Education	16.0

Source: Computed from *Statistical Abstract of Israel: 1957-58*, Jerusalem, Central Bureau of Statistics, 1958, Section S, Education, Table 28, p. 365.

29. Hovne, *The Labor Force in Israel*, p. 45.

30. Hanoch reports that "from a relative abundance of persons with secondary and higher education, and of experts, there developed a quite serious shortage." See his "Income Differentials in Israel," p. 44.

31. Uri Bahral, *The Effects of Mass Migration on Wages in Israel*, Jerusalem, Falk Project for Economic Research in Israel, 1965, pp. 5-6.

32. Ibid., p. 6.

33. "Income Differentials in Israel," pp. 44-52.

finding a negligible correlation between income and schooling in 1948 and a significant one in 1957.[34]

Thus, the postindependence migration lowered the level of schooling, raised the rate of return, left the variance of schooling unchanged, and consequently increased the education component (r^2 Var (S)). The results, as predicted by our analysis, were an increase in the inequality of income, in the explanatory power of schooling, and possibly also in the residual variance. The Israeli experience demonstrates that the parameters under study were influenced by historical events, and that the direction of the changes in the parameters was consistent with our previous findings.

Comparison with the United States

It is interesting to compare Israel in the late 1950s with the United States of approximately the same period. For urban heads of households in Israel in 1957, Klinov-Malul calculates internal rates of return (based on total money costs to society) of 17.0 per cent for eight years of elementary education, 6.5 per cent for secondary education, and 7.5 per cent for higher education.[35] These are considerably lower than rates of return in the United States.[36]

The variance of schooling in 1957 for male Jews of fifteen years of age and over was 17.9 years-squared and the average was 6.6 years. (See footnote 29 on p. 99). With the same grouping for schooling, the variance for white U.S. males between the ages of twenty-five and sixty-four in 1960 was 12.0 years-squared,[37] or two-thirds the Israeli inequality. Comparisons of U.S. and Israeli rates of return and schooling inequality indicate that the education component, r^2 Var (S), is smaller in Israel than in the United States,[38] and that the same is true of the level of schooling.

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34. For 1948, see Hanoch, p. 43. For 1957, see footnote 40 below.

35. Ruth Klinov-Malul, "Profitability of Investment in Education," *Fifth Report, Falk Project for Economic Research in Israel*, Jerusalem, 1961, pp. 138–146.

36. For U.S. internal rates of return, see Table 4-2.

37. *United States Census of Population, 1960, Subject Reports: Educational Attainment*, Washington, D.C. Bureau of the Census, Table 6.

38. If Var $(S)_{US} = (2/3)$ Var $(S)_I$, the education component would be smaller in Israel if r_{US}^2 Var $(S)_{US} > r_I^2$ Var $(S)_I$, $(2/3)\, r_{US}^2 > r_I^2$, or $(0.775)\, r_{US} > r_I$. Comparisons of Klinov-Malul's internal rates of return with estimates for the United States by Becker, Hansen, and Hanoch (see Table 4-2) indicate that the internal rates are sufficiently low in Israel to result in a smaller education component.

There is considerable evidence that income inequality is smaller in Israel than in the United States.[39] This is consistent with the smaller education component and lower level of schooling. As for the residual variance, the small inequality of personal wealth discussed above would preclude a relatively large variance, and indications are that it is, in fact, smaller in Israel.

Parameters given in Hanoch's paper on Israel permit a calculation of the coefficient of determination between schooling and earnings for urban heads of households in 1957. The unadjusted explanatory power for earnings (not log of earnings) is 16.9 per cent.[40] When earnings in 1959 are regressed on schooling of U.S. adult males (all males as well as white males), the unadjusted explanatory power is 12.2 per cent and 11.0 per cent, respectively.[41] It appears that for earnings the explanatory power of schooling is larger in Israel than in the United States. A greater explanatory power but a smaller income variance than in the United States implies that the residual variance is smaller in Israel.

To sum up, in Israel (1957-58) the inequality in ownership of nonhuman wealth, the rate of return, the level of schooling, and the education component are smaller than in the United States (1959), while the inequality of schooling is larger. The smaller education component and dispersion of nonlabor income result in a smaller dispersion of income in Israel. The explanatory power of schooling is greater in Israel. Altogether, the results of the time series analysis and the comparison with the United States are consistent with the interregional relationships found in Chapter 4 and the model developed in Chapter 3.

GREAT BRITAIN VERSUS THE UNITED STATES

In view of the fact that the inequality of wealth is greater in Great Britain than in the United States, many observers are under

39. See Hanoch, "Income Differentials in Israel," 1961, pp. 39-42; and Irving B. Kravis, *The Structure of Income: Some Quantitative Essays*, Philadelphia, The Wharton School, University of Pennsylvania, 1962, pp. 244-249.

40. Designating earnings by 1, continent of origin and place of immigration by 2, and level of education by 3, Hanoch states that $R^2_{12.3} = 0.064$ and $R^2_{12.3} = 0.220$. (See "Income Differentials in Israel," p. 105, Table 24.)

It can be shown that $R^2_{12.3} = (R^2_{12.3} - R^2_{13})/(1 - R^2_{13})$. (See T. Johnston, *Econometric Methods*, New York, McGraw-Hill, 1963, p. 60.) Then, $R^2_{13} = 0.169$.

41. *U.S. Census of Population, 1960, Subject Reports, Occupation by Earnings and Education*, Washington, D.C., Table 1.

the impression that Britain also has a greater inequality of income, and are surprised to find that income inequality is approximately the same in the two countries.[42] The analysis developed in this study can explain this pattern.

Note, first, that the inequality of earnings is smaller in Britain than in the United States.[43] Secondly, inequality is related to the rate of return from schooling and the level and inequality of schooling. Private rates of return from schooling in Britain have been estimated at 13 per cent for those who stay at school three years beyond the age of fifteen and at 14 per cent for those with three additional years of schooling.[44] These are similar to the calculated rates of return for high school and college in the United States (see Table 4-2). As can be seen from Table 5-7, in Great Britain the level of schooling is lower and schooling is more equally

TABLE 5-7

Distribution of the Labor Force by Years of Schooling in the United States and the United Kingdom

Years of Schooling	Per Cent Distribution	
	U.S., 1957	U.K., 1951
0 to 7	17.3	2.1
8	16.1	7.6
9	6.1	63.1
10	7.3	12.2
11	5.9	7.8
12	29.5	2.8
13	3.4	1.6
14	3.5	0.4
15	1.7	0.6
16	5.5	0.6
17 or more	3.7	1.3
	100.0	100.0

Source: Edward F. Denison, "Measuring the Contribution of Education (and the Residual) to Economic Growth," in *The Residual Factor and Economic Growth*, Paris, OECD, 1964, p. 43, Table 10.

42. See Kravis, *The Structure of Income*, pp. 249–250; and Lydall and Lansing, "A Comparison of the Distribution of Income and Wealth in the United States and Great Britain," *American Economic Review*, March 1959.

43. Lydall and Lansing, ibid., or Lydall, *The Structure of Earnings*, Table 5.5, p. 153.

44. D. Henderson-Stewart, "Appendix: Estimate of the Rate of Return to Education in Great Britain," *Manchester School*, September 1965, pp. 252–261.

distributed.[45] The similar rate of return and the lower level and inequality of schooling are consistent with the smaller inequality of earnings in Britain.

Thus, income inequality is approximately the same in Great Britain as in the United States because the smaller inequality of earnings in Britain (due to a lower level and inequality of schooling) is combined with a greater inequality of property income.[46] These findings are consistent with my schooling model of income inequality.

45. The table is for the United Kingdom (i.e., Great Britain plus Northern Ireland), but a similar education distribution would emerge if Northern Ireland were removed from the data.

46. If Y were total income, E earnings, and W property income, $Y = E + W$, and $\mathrm{Var}\,(Y) = \mathrm{Var}\,(E) + \mathrm{Var}\,(W) + 2\,\mathrm{Cov}\,(E,W)$. If, for simplicity, the covariance term were assumed to be the same in both countries, a larger inequality of property income combined with a smaller inequality of earnings could result in the same inequality of total income.

For an analysis of the effect of minimum schooling laws in Great Britain on the parameters under study, see my "Minimum Schooling Legislation and the Cross-Sectional Distribution of Income," *Economic Journal*, September 1969, pp. 495–507.