

# On class differentials in educational attainment

Robert Erikson<sup>†</sup>, John H. Goldthorpe<sup>‡</sup>, Michelle Jackson<sup>‡§</sup>, Meir Yaish<sup>¶</sup>, and D. R. Cox<sup>‡</sup>

<sup>†</sup>Swedish Institute for Social Research, SE-106 91 Stockholm, Sweden; <sup>‡</sup>Nuffield College, Oxford OX1 1NF, United Kingdom; and <sup>¶</sup>University of Haifa, Haifa 31905, Israel

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**Social class differentials in educational attainment have been extensively studied in numerous countries. In this paper, we begin by examining class differentials in the progression to higher secondary education among 16-year-old children in England and Wales. As has been shown for other countries, the differentials result both from the primary effects of differing levels of academic performance of children of different class background and from the secondary effects of differences in the educational choices that these children make at given levels of performance. Through counterfactual analyses in which the performance distribution of one class is combined with the choice distribution of another, primary and secondary effects are decomposed and the former are shown to be roughly three times the size of the latter.**

academic performance | educational choice | logistic regression | counterfactual | social inequality

Social class differentials in educational attainment have been extensively studied. Children from more advantaged class backgrounds have higher levels of educational attainment than children from less advantaged class backgrounds. Sociological evidence suggests that there has been a relatively high degree of temporal stability in the association between class origin and educational attainment in modern industrial societies (although where change has occurred it has generally been in the direction of a weakening association) (1, 2). Explanations have focused on how class differences in economic, social, and cultural resources lead to differences in academic performance. However, Boudon (3) argued that, in addition to interclass differences in the distribution of academic performance, there are also interclass differences in the educational choices made at given levels of performance. He called these primary and secondary effects, respectively.

The distinction between primary and secondary effects has been largely neglected in empirical research, although a notable exception is a study of Swedish students by Erikson and Jonsson (4). In the present paper, we develop their method and apply it to educational choices made by 16-year-old students in England and Wales. At age 16, students complete compulsory education, and then choose whether they will continue in full-time education to take advanced level academic qualifications (A levels) or alternatively enter vocational education or the labor market. We first show that both primary and secondary effects are present in creating class differentials in entering A-level education.

To investigate the relative importance of primary and secondary effects, we use a counterfactual analysis implemented by numerical integration. We suppose an idealized process that goes from performance to choice to outcome. In our implied model, therefore, a student achieves a level of academic performance and then makes their choice about whether to continue to A-level education. It is assumed that the choice characteristics of students of one class can be combined with the performance distribution of students of another class to produce a counterfactual or potential outcome. That is, a hypothetical intervention is considered in which the choice characteristics change but the performance distribution is unchanged (and vice versa). We discuss the implications of this further in *Discussion*.

The method may have wider applications.

**Table 1. Specification of classes**

Our system	Seven category NS-SEC
Salariat	I, Higher managerial and professional occupations II, Lower managerial and professional occupations
Intermediate	III, Intermediate occupations IV, Small employers and own account workers V, Lower supervisory and technical occupations
Working	VI, Semi-routine occupations VII, Routine occupations

NS-SEC, National Statistics Socio-Economic Classification.

## Data

The data on which the following analyses are based are drawn from the Youth Cohort Study (YCS) of 2002. This is a government survey, commissioned by the U.K. Department for Education and Employment, and was first carried out in 1985, the 2002 YCS being the 11th in the series. A representative sample of English and Welsh students are first surveyed at age 16, and then again at ages 17 and 18, although in this paper we use data from only the first of the three sweeps. The response rate for this YCS is rather modest, at 55%. A detailed discussion of the effect of nonresponse on the findings has not been published; however, it is likely that class differences in educational attainment are underestimated.

For the analyses, we construct three variables for each student: class background, level of academic performance, and choice in the transition at age 16.

Class background is recorded by the reported class of father or other household head. In the YCS, respondents are asked about the occupations of their parents, and on the basis of this the parents are assigned a class position. Class positions are coded to the seven category version of the National Statistics Socio-Economic Classification (NS-SEC) (5). We collapse these seven categories into three: salariat (classes I and II), intermediate classes (classes III, IV, and V), and working class (classes VI and VII), as shown in Table 1.

Our measure of academic performance is based on results in the General Certificate of Secondary Education (GCSE), taken by all 16-year-old students as they complete compulsory education. Students take GCSE examinations in differing numbers and combinations of subjects. In the interests of comparability, we focus on results in English language and mathematics, the two subjects taken by almost all students. We assign points scores to the grades that students can achieve in these two subjects; these range from a high of A\* (scored 9) to a low of U (a fail, scored 1). The scores for English and mathematics are added, and then standardized to z scores (with a mean of 0 and standard deviation of 1).

We treat choice in the transition at 16 as simply binary, distinguishing students who stay in full time education after 16 to engage in A-level work from those who do not.

<sup>§</sup>To whom correspondence should be addressed. E-mail: michelle.jackson@nuf.ox.ac.uk.

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**Table 2. Summary information about transition rates**

Class	Percent making transition		Sample size
	Observed	Fitted	
Salariat	76.6	76.1	5,062
Intermediate	54.7	53.9	4,706
Working	39.7	39.5	1,870
All	62		11,638

Odds ratios from raw numbers are: Salariat (S)/Working (W), 4.96; S/Intermediate (I), 2.70; I/W, 1.84.

**Methods and Results**

Table 2 shows clear class differences in the probability of making the transition. Although 62% of all students continue to A-level work, only 40% of working class students do so as compared with 77% of students of salariat background. As indicated by the odds ratios in Table 2, the odds for the latter students to stay on are almost five times as high as those for the former.

To analyze the influence of primary and secondary effects in the transition to A-level work, we follow the approach of Erikson and Jonsson (4). We separate primary and secondary effects, and then show the results graphically as class-specific “transition probabilities” for all levels of previous performance.

We represent primary effects by estimating the mean and standard deviation of the performance scores separately for each class, as shown in Table 3. These are then used to fit a normal distribution for each class. The normal distributions are, of course, approximations, and there is a small inconsistency in using a normal distribution for each class component and for the overall scores, but this has a negligible effect on the conclusions. To represent secondary effects, we use separate binary logistic regressions for each class predicting whether a student with a given performance score, *x*, enters A-level education. With the probability of entering A-level written in the form

$$\Lambda(\alpha + \beta x) = e^{\alpha + \beta x} / (1 + e^{\alpha + \beta x}),$$

estimates of  $\alpha$  and  $\beta$  are given for each class also in Table 3, together with the estimate, namely  $-\alpha/\beta$ , of the value of *x* at which 50% entrance is achieved. Note that the log odds at argument *t* are  $\log[\Lambda(t)/\{1 - \Lambda(t)\}] = t$  so that  $\alpha$  is the log odds at  $x = 0$ , the overall mean.

The two sets of curves representing the primary and secondary effects are shown together in Fig. 1.

The systematic differences within the two sets of curves reveal the presence of both primary effects and secondary effects. The performance distributions show that students of salariat background have the highest level of performance, followed by students of intermediate class background, and then by students from the working class. Also, with a slight deviation in one tail, the logistic curves showing transition probabilities lie consistently above each other in the same class ordering. From these curves, it may be noted that the 50% point corresponds to an academic performance score of  $-0.47$  for students of salariat background, as against  $-0.20$  for working class students.

**Table 3. Estimated parameters ( $\bar{\mu}$ ,  $\bar{\sigma}$ ) for performance, ( $\bar{\alpha}$ ,  $\bar{\beta}$ ) for transition, and 50% point of transition,  $-\bar{\alpha}/\bar{\beta}$**

Class	$\bar{\mu}$	$\bar{\sigma}$	$\bar{\alpha}$	$\bar{\beta}$	50% point	<i>N</i>
Salariat	0.365	0.920	1.124	2.384	-0.471	5062
Intermediate	-0.185	0.952	0.681	2.271	-0.300	4706
Working	-0.530	0.959	0.448	2.226	-0.201	1870
All	0.00	1.00				11638

To develop the analysis further so as to quantify the relative contributions of primary and secondary effects and also to check on the adequacy of our model, we may argue as follows.

First, for a given class the proportion proceeding to A-level is, under the above assumptions,

$$\int \bar{\sigma}^{-1} \phi\left(\frac{x - \bar{\mu}}{\bar{\sigma}}\right) \Lambda(\bar{\alpha} + \bar{\beta}x) dx, \tag{1}$$

where  $\phi(x)$  is the standard normal density,  $(2\pi)^{-1/2} \exp(-x^2/2)$ .

This integral does not have a closed form and has to be evaluated numerically. We take  $10^3$  ordinates over the range  $(-4,4)$ . When the parameters for a particular class are used, the overall observed proportions are closely approximated (see Table 2).

The next step is to consider counterfactual questions such as: what if working class students retained their own level of performance but took on, say, the transition probabilities, conditional on performance, of students of salariat background? The resulting transition rate can be estimated by inserting in the integral the values of  $\bar{\mu}$ ,  $\bar{\sigma}$  for the working class and  $\bar{\alpha}$ ,  $\bar{\beta}$  for the salariat.

Table 4 summarizes the outcome of all such possible calculations for our three classes, the rows identifying the class whose performance distribution is used and the columns the class whose transition probabilities are used. The diagonal elements in Table 4 thus correspond to the real transition rates as observed and as estimated under our model; see Table 2.

Over the range in question, the log odds in Table 4 are an almost linear function of the probabilities, so that comparisons within the probability and log odds arrays give virtually identical conclusions. However, from some points of view, log odds are, as will be shown, more easily interpreted.

The estimated probabilities in Table 4, and more particularly the log odds, have a very simple form. First, they correspond to a close approximation to an additive structure in which each value is the sum of a row effect and a column effect. Equivalently, the interaction terms in the various  $2 \times 2$  contingency tables that could be formed from Table 4 are all very small. For comparison, the log odds calculated from the marginal means and assuming additivity are shown as the second entry. That is, each entry is the sum of the marginal means minus the overall mean. Discrepancies in log odds between the two values of log odds are typically  $\approx 0.01$ , corresponding to a 1% discrepancy in odds.

Secondly, the row differences in Table 4, representing primary effects, are almost proportional to the corresponding column differences, representing secondary effects.

Thus, however we choose to summarize the relative contribution of primary and secondary effects, the secondary effect is about one quarter of the total effect. More specifically, the primary effect as a percentage of total effect is 75.7 for salariat (S) vs. intermediate (I), 76.5 for S vs. working (W), and 77.7 for I vs. W.

An analytical explanation of some of these conclusions is given below.

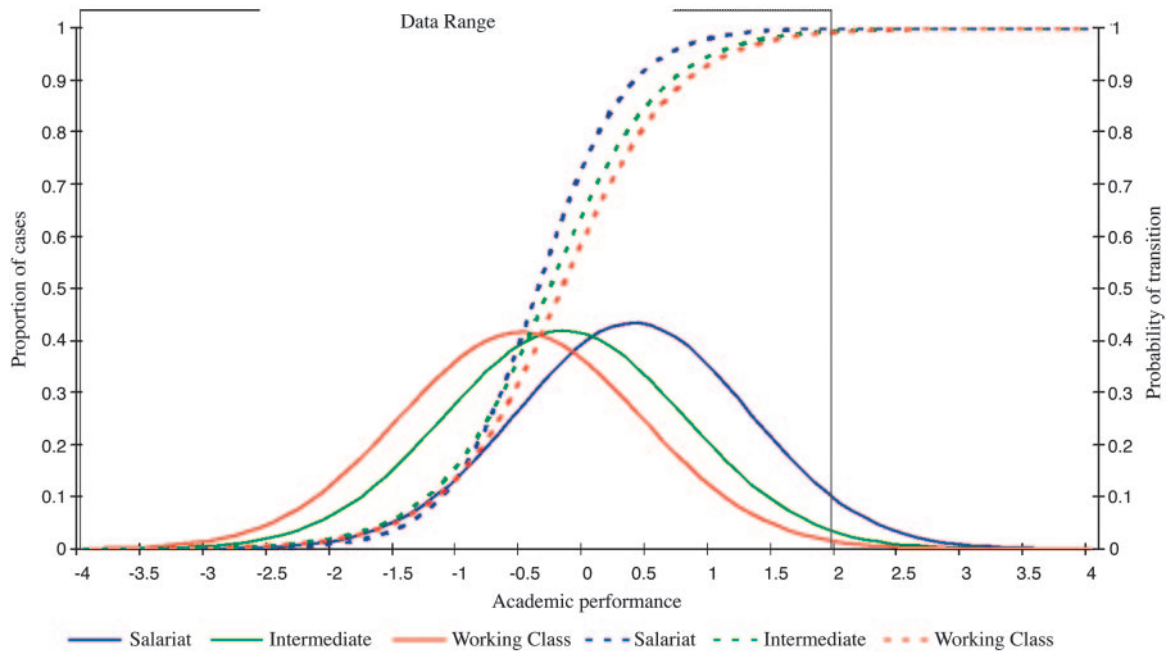


Fig. 1. Fitted normal distributions for performance within each of three classes. Fitted logistic curves for probability of transition as a function of performance.

### An Analytical Approximation

We have assumed that, within a given class, measured performance,  $x$ , is normally distributed with mean  $\mu$  and standard deviation  $\sigma$  and that the probability of an individual progressing to A-levels given  $x$  is conveniently specified in the form,  $\Lambda\{\gamma + \beta(x - \mu)\}$ , where in the previous notation  $\alpha = \gamma - \beta\mu$ . Then the proportion progressing to A-level is given by the integral (Eq. 1). The integral, as noted above, cannot be evaluated in closed form, but it is known that to a close numerical approximation,  $\Lambda(t) = \Phi(kt)$ , where for any value of  $y$ , the function  $\Phi(y)$  is the standard normal integral,

$$\Phi(y) = \int_{-\infty}^y \phi(x) dx,$$

and  $k$  is a tuning constant; with  $k = 0.61$  the approximation has small relative error except in the extreme tails (6). We then replace Eq. 1 by

$$\int \sigma^{-1} \phi\{(x - \mu)/\sigma\} \Phi\{k\gamma + k\beta(x - \mu)\} dx = \Phi\{k\gamma/\sqrt{(1 + k^2\sigma^2\beta^2)}\}. \quad [2]$$

Table 4. Estimated probabilities and odds and log odds, direct and under additive assumption, of transitions for real and counterfactual combinations of estimated distributions of performance (rows) and transition probabilities (columns)

	Salarial	Intermediate	Working
Estimated probabilities and odds			
Salarial	0.761; 3.182	0.710; 2.448	0.681; 2.135
Intermediate	0.596; 1.475	0.539; 1.169	0.507; 1.028
Working	0.481; 0.927	0.426; 0.742	0.395; 0.653
Log odds			
Salarial	1.158; 1.140	0.895; 0.901	0.758; 0.770
Intermediate	0.389; 0.394	0.156; 0.155	0.028; 0.024
Working	-0.076; -0.063	-0.298; -0.303	-0.426; -0.434

Here, the first integral has been evaluated in closed form. If now the normal integral is approximated by a logistic function by applying the above approximation in reverse order, we have that the probability of transition is

$$\Lambda\{\gamma/\sqrt{(1 + k^2\sigma^2\beta^2)}\}.$$

That is, the log odds are

$$\frac{\gamma}{\sqrt{(1 + k^2\sigma^2\beta^2)}} = \frac{\alpha + \beta\mu}{\sqrt{(1 + k^2\sigma^2\beta^2)}}. \quad [3]$$

Now, when this is applied to different social classes, different values of  $\sigma$  and  $\beta$  have been used for each. Had the same  $\sigma$ ,  $\beta$  been used throughout, and this is a relatively small approximation as Table 2 shows, the log odds would have been

$$c(\alpha/\beta + \mu),$$

where  $c$  is a constant. That is, if the attainment mean  $\mu_i$  for class  $i$  were combined with the logistic curve for class  $j$ , the counterfactual log odds would have a simple additive structure, being a multiple of the difference between the attainment mean and the 50% point,  $-\alpha_j/\beta$ . The small departures from additive structure in Table 4 are a consequence of using different  $\alpha, \beta$  for the different classes.

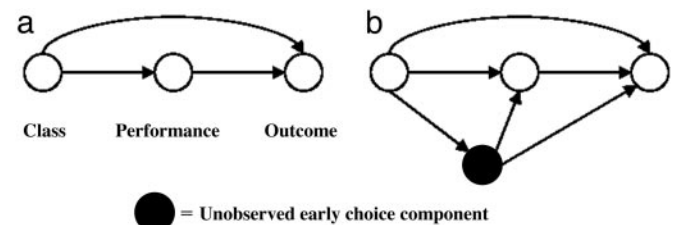


Fig. 2. Path diagrams for the analysis in this paper (a) and with unobserved early choice variable (b).

It can be shown that in this instance the numerical error involved in interchanging logistic and integrated normal functions is <1% of the tabulated values; such good results would not be achieved if small (or large) overall probabilities critically dependent on the tails of the distribution were involved.

## Discussion

There has been extensive use of counterfactual arguments or potential responses in the context of empirical studies of potential causality in the observational social sciences and in epidemiology. Much of the work stems from Rubin (7) (for a wide-ranging discussion see ref. 8; for recent reviews, see refs. 9 and 10; and for causality in sociology, see ref. 11). In the present context, the role of the counterfactual argument is essentially to place the contributions of performance and choice on a comparable scale so that their relative importance can be assessed.

In its simplest form, the implicit potentially causal model is that of Fig. 2*a*, where edges in the graphical representation go from class to performance and from class and performance to outcome. There are two paths from class to outcome and our objective has been to assess their relative strengths. A more realistic model has an unobserved component, early choice, shown in Fig. 2*b*. Marginalization over the unobserved component inserts extra dependencies in the graph, in particular between class and performance and the implication here is that the overall importance of choice may well have been underestimated. For example, it is possible that an anticipatory decision not to enter A-level courses would lead to less work being done in preparing for the examinations used to measure performance and hence to lower performance. Indeed, we have shown that such decisions are likely to lead to underestimation of the secondary effects (unpublished data). Future work is needed to address this issue using more detailed data as well as to examine the development over time in the features studied here.

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