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11 Academic Ability, Earnings, and the Decision to Become a Teacher: Evidence from the National Longitudinal Study of the High School Class of 1972

Charles F. Manski

11.1 Introduction

Perceived shortcomings in the quality of American education at the elementary and secondary school levels have drawn much public attention recently. In particular, concern with the composition of the teacher force has been prominent. This focus presumably arises out of the juxtaposition of three factors.

First, there is general acceptance of the proposition that educational achievement is influenced by the ability of the teachers who guide the learning process. (There is, of course, much less agreement about how educational achievement and teacher ability should be measured.) Second, there is an often-expressed dissatisfaction with the distribution of ability within the present teaching force. Third, there is a common perception that feasible changes in public policy can generate a shift in the ability distribution of the supply of teachers. In particular, it is asserted that merit pay, general increases in teacher salaries, and/or subsidization of the college education of prospective teachers would induce more college students of high ability to select teaching as a career.

Informed assessment of the various proposals for increasing the attractiveness of teaching is possible only if we can forecast the extent to which these proposals, if enacted, would influence the occupational choice decisions of high-ability young adults. Until now, there has been no basis for making such forecasts. In the absence of empirical analysis, we can only guess at the impact of changes in teacher salaries on the quality composition of the teaching force.

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The research reported here, through analysis of data from a national sample of college graduates, examines the relationship between academic ability, earnings, and the decision to become a teacher. The National Longitudinal Study of the High School Class of 1972 (NLS72) surveyed 22,652 high school seniors in the spring of 1972 and has subsequently followed this panel as its members have progressed through postsecondary education and into the labor force. The most recent survey took place in October 1979. At that time, contact was successfully made with 18,630 members of the panel. Of these, 3,502 reported they had completed a bachelor's degree in 1976 or 1977. Of this group, 2,952 reported they were working in October 1979. Of these, 510 reported they were employed as teachers.

The NLS72 data offer a valuable resource for description of the empirical pattern of ability, earnings, and occupations found in a recent cohort of American college graduates. Inspection of these data reveals the following:

—Among the working NLS72 respondents who have received a bachelor's degree, the frequency of choice of teaching as an occupation is inversely related to academic ability. This holds whether academic ability is measured by SAT score or by high school class rank. Conditioning on SAT score, however, the frequency of choice of teaching does not vary with class rank.

—Conditioning on sex and academic ability, the earnings of teachers are much lower, on average, than those of other working college graduates.

—Conditioning on sex, the earnings of teachers tend to rise only slightly, if at all, with academic ability. A relationship between earnings and ability is more noticeable in other occupations but remains weak. Academic ability explains only a small part of the observed variation in earnings within the cohort of NLS72 college graduates.

—Conditioning on academic ability and occupation, males consistently have higher earnings than do females. The sex differential in earnings is relatively small in teaching but quite pronounced in other occupations. Interestingly, the rate at which earnings rise with ability is very similar for males and females.

To evaluate policy proposals intended to influence the composition of the teaching force, it is necessary to go beyond descriptive analysis. The NLS72 data support estimation of an econometric model explaining occupation choice as a function of the earnings and nonmonetary characteristics associated with alternative occupations. Given this model, it is possible to forecast the consequences of policies that combine increases in teacher salaries with the institution of minimum academic ability standards for teacher certification. Forecasts presented in this paper suggest the following:

—In the absence of a minimum ability standard, increases in teacher earnings would yield substantial growth in the size of the teaching force but minimal improvement in the average academic ability of teachers. Under present conditions, the aggregate wage elasticity of the supply of teachers appears to be in the range of two to three. As wages increase, both high- and low-ability students are attracted into teaching, so the ability composition of the teaching force changes little.

—If teacher salaries are not increased, institution of a minimum ability standard improves the average ability of the teaching force but reduces its size. Establishment of a standard sufficient to raise the average academic ability of teachers to the average of all college graduates may reduce the size of the teaching force by 20 percent.

—The average ability of the teaching force can be improved and the size of the teaching force maintained if minimum ability standards are combined with sufficient salary increases. It appears that the average academic ability of teachers can be raised to the average of all college graduates if a minimum SAT score (verbal plus math) of 800 is required for teacher certification and if teacher salaries are raised by about 10 percent over their present levels. To achieve further improvements in average teacher ability without reducing the size of the teaching force would require a higher minimum ability standard combined with a larger salary increase.

Before proceeding, it is important to stress that the indicators of ability available for the NLS72 panel and used in this research are certain measures of academic success, namely SAT scores and high school class rank. It seems reasonable to assume that these variables are positively associated with performance as a teacher, but formal evidence for this proposition is lacking. (See, for example, the discussion in Weaver 1983.) The relevance of the analysis that follows to the debate over the quality of the teacher force depends on the extent to which academic ability and teaching ability coincide.

The plan of this chapter is as follows: Section 11.2 describes the NLS72 sample and the variables that measure occupation, academic ability, and earnings. Section 11.3 reports our descriptive analysis of the NLS72 data. The econometric model explaining occupation choices is developed and estimated in section 11.4. The model is applied to forecast the effects of policy proposals in section 11.5. Section 11.6 contains brief concluding comments.

11.2 Composition of the Sample and Definition of Variables

11.2.1 The Sample

The work in this chapter is based entirely on data for the 2,952 NLS72 respondents who, when interviewed in late 1979, reported that they had received a bachelor's degree in 1976 or 1977 and that they were working in October 1979. Some of the analysis is based on the subsample of respondents for whom complete academic ability and earnings data were available. A comprehensive description of the NLS72 data, including the sample design, questionnaires, and frequency counts of responses, is given in Riccobono et al. (1981).

11.2.2 The Occupation Variable

In all that follows, a respondent's occupation is taken to be his or her declared job type in October 1979 as coded by the NLS into the three-digit census classification system. In the cross-tabulations of tables 11.1, 11.2, and 11.4, these codes are aggregated into three occupation classes: (*a*) teachers, exclusive of college faculty (census codes 141–45); (*b*) professional, technical, and kindred workers, exclusive of teachers (census codes 001–140, 150–95); and (*c*) all other occupations (census codes 201–992). In the models of tables 11.3, 11.5, and 11.6, classes (*b*) and (*c*) are further aggregated into a single "nonteaching" occupation.

In principle, the census coding system distinguishes various categories of teachers. In practice, this detailed coding is ambiguous because 275 of the 510 teachers are not classified. Of the ones who are classified, 35 are reported to be nursery and kindergarten teachers, 104 to be elementary school teachers, 92 to be secondary school teachers, and 4 to be adult education teachers. These are small samples, particularly when disaggregated by sex.

Coded as unclassified teachers are such groups as fine arts teachers and flying instructors as well as those school teachers whose response to the occupation question was insufficiently detailed to permit a more refined classification. Examination of the employer codes for the classified and unclassified teachers reveals that 59 percent of the former group and 60 percent of the latter group work for governmental units. The ability and earnings distributions of the two groups are also similar. These facts make it reasonable to assume that the unclassified group is composed primarily of elementary and high school teachers. Given this and given the small size of the classified group, the statistics presented here are computed using all respondents coded as teachers, not just those for whom a more detailed classification is available.

It should be noted that the NLS72 survey offers some alternatives to our identification of occupation with job type in October 1979. First, whenever a respondent reported that he had worked in October 1978 or October 1977, job type at these dates was reported. Second, when interviewed in 1979, each panel member was asked to anticipate his or her occupation at age thirty (that is, about five years into the future). Third, each respondent was asked to report the field in which he or she received a bachelor's degree. I have chosen to use the October 1979 job reports because they are the latest revealed preference data available for the NLS72 respondents. It would be interesting to redo the analysis using alternative definitions of occupation.

11.2.3 The Academic Ability Variables

As part of the base-year survey instrument administered in 1972, the NLS obtained from guidance personnel the percentile high school class rank of each respondent and, where available, each respondent's SAT or ACT score. A battery of IQ and aptitude tests was administered as well. In this paper, academic ability is measured by the class rank and SAT/ACT data. The NLS test battery data are not used here.

Among the 2,952 respondents, class rank information is available for 2,287. Either an SAT or ACT score is available for 2,468 respondents, with the former predominating. While the SAT and ACT examinations are distinct, I have, in the interest of using observations efficiently, converted each ACT score to an SAT equivalent by matching the tenth and ninetieth percentile scores and interpolating elsewhere. The rationale for using both the class rank and SAT score as measures of academic ability is that the two have previously been shown to have complementary explanatory power in predicting both college admissions decisions and college completion rates (Manski and Wise 1983).

11.2.4 The Earnings Variable

Each respondent working in October 1979 was asked to report gross pay per week at his or her primary job. Hours worked per week at the primary job were also reported. In the parts of this chapter concerned with earnings, I restrict attention to the 2,335 respondents whose reported hours worked per week are between thirty and sixty and whose pay per week is between $100 and $800. The restriction on hours worked is intended to limit attention to "normal" full-time jobs. The restriction on pay cuts off volunteer workers on the low end and, on the high end, a few respondents whose reported weekly pay seemed extraordinary for a twenty-five-year-old in 1979.

The reported pay per week is used as the measure of realized earnings. An obvious alternative measure is the hourly wage, computed by dividing gross pay by hours worked. The former measure seems preferable since most college graduates are paid on a salary rather than on an hourly basis. Empirically, the same patterns emerge whichever earnings measure is used. Note that all monetary figures in this paper are expressed in 1979 dollars.

11.3 Patterns of Academic Ability, Occupation, and Earnings

11.3.1 Academic Ability and Occupation

Considering males and females separately, table 11.1 partitions the sample respondents into four SAT score groups and, for each group, presents the observed distribution of occupations. In table 11.2, percentile class rank in high school is used as the measure of ability. These

Table 11.1 **Occupation as a Function of Sex and SAT Score**

	SAT Score (verbal + math)			
Occupation	400–800	801–1000	1001–1200	1201–1600
A. Males				
Teacher	.16	.11	.06	.05
Professional	.22	.23	.36	.55
Other	.62	.66	.58	.40
Number of respondents	148	400	501	221
B. Females				
Teacher	.34	.30	.21	.09
Professional	.14	.26	.36	.46
Other	.52	.43	.43	.45
Number of respondents	208	421	413	156

Table 11.2 **Occupation as a Function of Sex and High School Class Rank**

	Percentile Class Rank			
Occupation	1–50	51–75	76–90	91–100
A. Males				
Teacher	.11	.09	.08	.06
Professional	.20	.24	.34	.53
Other	.69	.66	.57	.40
Number of respondents	242	388	305	249
B. Females				
Teachers	.35	.31	.24	.20
Professional	.14	.23	.29	.39
Other	.51	.46	.47	.41
Number of respondents	116	244	336	407

data clearly corroborate the conventional wisdom that choice of teaching as an occupation is inversely related to academic ability. It does not matter whether we look at males or females, whether we take SAT score or class rank as the measure of academic ability. In each case, the frequency with which the NLS72 respondents enter teaching falls substantially as academic ability rises. In contrast, the frequency with which respondents work in professional or technical fields other than teaching consistently rises with ability, in fact dramatically so.

Other cross-tabulations of SAT scores and occupation based on NLS72 data have been presented in Vance and Shlechty (1982). Their criteria for inclusion in the sample and for classification of a respondent as a teacher were different than those used here. Their findings were similar.

Table 11.3 offers further perspective on the relationship between academic ability and occupation. Considering males and females separately, this table presents estimates for a simple probit model explaining the probability that, conditioned on SAT score and class rank, a working college graduate is a teacher. Inspection of the results indicates that when SAT score and class rank are conditioned on jointly, the partial effect of SAT score on the probability of entering teaching is almost identically negative and statistically significant for males and females. On the other hand, the partial effect of class rank is very weak and ambiguous in sign. In fact, it is reasonable to conclude that holding SAT score fixed, the probability of entering teaching does not vary with class rank.

11.3.2 Academic Ability, Occupation, and Earnings

Considering males and females separately, table 11.4 partitions the sample into twelve SAT score-occupation cells. Presented in each cell are (1) mean pay per week, (2) the number of respondents in the cell, and (3) the standard deviation of pay per week. I have computed alternative tables using hourly wage as the measure of earnings and class

Table 11.3 Probit Model of Teaching Occupation as a Function of Sex and Academic Ability

	Males		Females	
Variable	Coefficient	Asymptotic Std. Error	Coefficient	Asymptotic Std. Error
SAT score (200–1600)	−0.00115	(0.00036)	−0.00111	(0.00029)
Class rank (1–100)	0.00068	(0.00298)	−0.00228	(0.00275)
Intercept	−0.304	(0.317)	0.565	(0.240)
Sample size	1037		968	

rank as the measure of academic ability and have found patterns very similar to those in table 11.4. Among the many interesting features of table 11.4 are the following:

—Conditioning on sex and SAT score, mean pay per week is almost always highest for professional and technical workers and lowest for teachers, with workers in other occupations in between. For males, the differentials are more substantial than for females. For example, considering males with SAT score in the 801–1,000 range, the mean pay of professional workers is 1.48 times that of teachers. For females, the comparable number is 1.22.

—Conditioning on sex, mean pay per week in the nonteaching occupations tends to rise with SAT score but the pattern is weak. For teachers, there is little evidence of an earnings-ability pattern. A rela-

Table 11.4 Pay per Week as a Function of Sex, SAT Score, and Occupation

	SAT Score (verbal + math)				
Occupation	400–800	801–1000	1001–1200	1201–1600	Total
			A. Males		
Teacher mean	237	222	236	237	230
Count	14	40	21	11	86
Std. dev.	63	47	75	62	59
Professional	320	328	328	365	337
	26	78	155	89	348
	91	89	99	85	94
Other	271	283	288	286	283
	80	212	218	56	566
	101	104	97	92	100
Total	278	286	301	327	297
	120	330	394	156	1000
	98	100	100	97	100
			B. Females		
Teacher mean	199	223	227	216	219
Count	51	102	75	13	241
Std. dev.	54	53	53	65	55
Professional	272	272	271	309	279
	24	90	127	56	297
	153	70	69	84	83
Other	221	218	225	268	227
	86	149	142	54	431
	70	71	75	90	76
Total	221	234	243	281	241
	161	341	344	123	969
	86	69	72	89	78

Note: Mean pay is in dollars, reported in October 1979.

tionship becomes more apparent if we do not condition on occupation. Examination of the column marginals indicates clearly that mean pay does increase with SAT score. In particular, the mean pay of males with scores in the 1,200–1,600 range is 1.18 times that of those with scores in the 400–800 range. For females, the comparable number is 1.27.

—Conditioning on SAT score and occupation, males consistently have higher mean pay per week than females. This pattern persists in almost every SAT score-occupation cell but is least pronounced among teachers. To cite some examples, the mean pay of professional males with SAT scores in the 1,000–1,200 range is 1.21 times that of females with the same characteristics. Considering teachers with SAT scores in the same range, the mean income of the males is 1.04 that of the females. Recall that these data concern a sample of respondents all of whom graduated from high school in 1972, all of whom graduated from college in 1976 or 1977, and all of whom are working at least thirty hours per week and earning at least $100 per week in 1979. It is therefore difficult to attribute the observed differences in the pay of males and females to an unobserved determinant correlated with sex.

—Conditioning on sex and SAT score, the standard deviation of pay per week is consistently much lower for teachers than for the remaining two occupation groups. Conditioning on sex and occupation, the standard deviation is more or less invariant across ability groups. Conditioning on SAT score and occupation, the standard deviation is generally lower for females than for males.

Table 11.5 gives additional insight into the behavior of earnings. Conditioning on sex and occupation (teacher versus nonteacher), the table presents ordinary least squares estimates of a model explaining pay per week as a linear function of SAT score and high school class rank. Inspection of the table indicates that academic ability explains only a small part of the variation in observed earnings across this cohort of working college graduates. This fact, which was earlier noted in the analysis of table 11.4, is expressed succinctly in the R^2 statistics, which range from .03 to .06.

At the same time, the regressions uniformly show that conditioning on sex and occupation, earnings do increase with both SAT score and class rank. In fact, the estimated coefficients are reasonably similar across the four subsamples. To get a feel for magnitudes, consider a one hundred point increase in SAT score. The predicted effects on weekly earning across the four subsamples are $5.06, $7.26, $4.01, and $5.61 respectively. A ten percentile increase in class rank is associated with earnings increases of $2.85, $3.50, $4.19, and $3.69 respectively. The marginal statistical significance of the estimated coefficients should make one cautious in drawing sharp implications from these numbers. The general pattern, however, seems firmly based.

Table 11.5 Linear Model of Earnings as a Function of Sex, Occupation, and Academic Ability

	Male Teachers		Male Nonteachers	
Variable	Coefficient	Std. Error	Coefficient	Std. Error
SAT score (200–1600)	0.0506	(0.0475)	0.0726	(0.0234)
Class rank (1–100)	0.285	(0.421)	0.350	(0.189)
Intercept	158.	(41.)	204.	(21.)
R^2	.04		.03	
Sample size	64		748	

	Female Teachers		Female Nonteachers	
Variable	Coefficient	Std. Error	Coefficient	Std. Error
SAT score (400–1600)	0.0401	(0.0279)	0.0561	(0.0219)
Class rank (1–100)	0.419	(0.239)	0.369	(0.228)
Intercept	147.	(22.)	162.	(20.)
R^2	.06		.03	
Sample size	188		593	

Note: Earnings are in dollars per week, in 1979.

Comparison of the coefficients for males and females suggests that the earnings of males may be somewhat more sensitive to SAT score than are those of females but less sensitive to class rank. Again, these differences are relatively small. It seems more relevant to stress that the earnings of males and females tend to increase similarly with academic ability. The differences between male and female earnings that were seen in table 11.4 show up in these regressions as differences in the intercept coefficients. Those for males are higher than those for females, with the discrepancy much more pronounced in occupations other than teaching.

11.4 A Structural Interpretation of the Observed Patterns

The patterns of academic ability, earnings, and occupation reported in section 11.2 arise out of the interaction of the decisions of two sets of actors, college graduates and employers. In selecting occupations, college graduates presumably compare the expected earnings streams and nonmonetary characteristics associated with the available alternatives. In making job offers, employers may use measured academic ability as an indicator of potential job performance. To the extent that academic ability is perceived by employers to be positively associated with job performance, college graduates with high ability will be offered more attractive positions than will be offered those with low ability. To the extent that the return to ability differs across occupations, we should observe an empirical relationship between ability and occupation choice.

In this section we attempt to interpret the observed patterns in the NLS data in terms of a simple econometric model with two parts. One submodel explains occupation choice as a function of the earnings and nonmonetary characteristics associated with alternative occupations. The other explains occupation-specific earnings as a function of academic ability and other factors. With this done, it is possible in principle to predict the effect of changes in teacher salaries on the probability that a college graduate of given academic ability selects teaching as his or her occupation.

11.4.1 A Model of Occupation Choice and Earnings

Let $i = 1$ designate the occupation of teacher and let $i = 0$ represent all other occupations. Let T be the population of working college graduates and assume that each person t in T must select between the two classes of occupations. Assume that person t associates with teaching an expected present discounted earnings per week $y(t1)$ and an index of nonmonetary job characteristics $g + \gamma(t)$. Here g is a constant and γ varies with t. Person t aggregates the monetary and nonmonetary characteristics into a utility value

$$u(t1) = y(t1) + g + \gamma(t). \tag{1}$$

Nonmonetary job characteristics are unobservable to us, so we treat $\gamma(t)$ as a random variable distributed over T. Given the presence of the intercept g, we set $E[\gamma(t)] = 0$ without loss of generality.

The utility of the nonteaching occupation is

$$u(t0) = y(t0). \tag{2}$$

Here, we have set the index of nonmonetary characteristics equal to zero in order to fix the origin of the utility function. Thus, the term $g + \gamma(t)$ appearing in equation (1) should be interpreted as indexing the nonmonetary characteristics of teaching relative to other occupations. Note that in equations (1) and (2), u is measured in the same units as y. This fixes the scale of the utility function as dollars.

We assume that person t selects teaching as an occupation if

$$u(t1) - u(t0) = [y(t1) - y(t0)] + g + \gamma(t) > 0. \tag{3}$$

Some obvious objections may be raised against equation (3). This specification of decision making ignores a host of dynamic considerations in the determination of career paths. Moreover, it aggregates broad arrays of heterogenous occupations into two fictitious, composite alternatives. Nevertheless, in the interest of enabling empirical analysis, we shall maintain equation (3) as a working hypothesis.

Empirically, we take the chosen occupation of an NLS respondent to be his or her reported occupation in October 1979. We do not directly observe expected earnings, but an indicator is sometimes available. That is, we observe reported weekly pay in October 1979 for the chosen occupation. Assume that the relationship between expected earnings y and reported pay, designated Y, is

$$Y(t1) = d_1 + y(t1) + \delta(t1), \tag{4}$$

$$Y(t0) = d_0 + y(t0) + \delta(t0), \tag{5}$$

where d_1 and d_0 are constants and $\delta(t1)$ and $\delta(t0)$ are random variables over T. Given the presence of the intercepts d_1 and d_0, we set $E[\delta(t1)] = E[\delta(t0)] = 0$.

Observe that d_1 and d_0 allow for the possibility that earnings vary systematically over the life cycle. In particular, if salaries tend to rise with seniority, then we should expect d_1 and d_0 to be negative since the NLS respondents are at the beginnings of their careers. The constants also allow for a population-wide difference between current and permanent income. In particular, we should expect d_0 and possibly d_1 to be lower in a recession year than in a boom year. With d_1 and d_0 picking up cohort-wide differences between reported and expected earnings, the random variables $\delta(t1)$ and $\delta(t0)$ represent person-specific deviations.

Let $S(t)$ and $R(t)$ be person t's observed SAT score and high school class rank. Assume that expected earnings in teaching is a linear function of these measures of academic ability and of other variables $\epsilon(t1)$, that is,

$$y(t1) = a_1 + b_1 * S(t) + c_1 * R(t) + \epsilon(t1), \tag{6}$$

where (a_1, b_1, c_1) are constants. Similarly, assume that expected earning in the nonteaching occupation is given by

$$y(t0) = a_0 + b_0 * S(t) + c_0 * R(t) + \epsilon(t0). \tag{7}$$

The coefficients (b_1, c_1) and (b_0, c_0) quantify the monetary returns to academic ability in the teaching and nonteaching occupations. The variables $\epsilon(t_1)$ and $\epsilon(t_0)$ represent worker-specific characteristics other than SAT score and class rank that are known to both employers and workers and are perceived as related to job performance. We do not observe these characteristics and so treat $\epsilon(t1)$ and $\epsilon(t0)$ as random variables over T. Given the presence of the intercepts a_1 and a_0, we set $E[\epsilon(t1)] = E[\epsilon(t0)] = 0$.

It follows from equations (4) through (7) that the reported pay of NLS respondent t is related to respondent's SAT score and high school class rank by

$$(8) \quad Y(t1) = (d_1 + a_1) + b_1{*}S(t) + c_1{*}R(t) + [\delta(t1) + \epsilon(t1)],$$

if the respondent is a teacher, and by

$$(9) \quad Y(t0) = (d_0 + a_0) + b_0{*}S(t) + c_0{*}R(t) + [\delta(t0) + \epsilon(t0)]$$

otherwise. It follows from equations (3), (6), and (7) that an NLS respondent chooses to be a teacher if and only if

$$(10) \quad (g + a_1 - a_0) + (b_1 - b_0){*}S(t) + (c_1 - c_0){*}R(t) + [\gamma(t) + \epsilon(t1) - \epsilon(t0)] > 0.$$

Conditional on S and R, the probability that a person is observed to choose teaching is

$$(11) \quad Pr(i = 1|S,R) = Pr(\eta < A + B{*}S + C{*}RIS,R),$$

where $A \equiv (g + a_1 - a_0)$, $B \equiv (b_1 - b_0)$, $C \equiv (c_1 - c_0)$, and $\eta(t) \equiv -[\gamma(t) + \epsilon(t1) - \epsilon(t0)]$.

Consider now a policy proposal whose sole effect is to change a person's expected earnings in teaching from $y(1)$ to $y(1) + X$, for some X. Under this proposal, the probability that the person will choose teaching as an occupation is

$$(12) \quad Pr(i = 1|S,R,X) = Pr(\eta < A + B{*}S + C{*}R + XIS,R,X).$$

If the parameters A,B, and C and the distribution of η are known, equation (12) provides an operational means of forecasting the impact of a proposed change in teacher salary on the occupation choice decision of a college graduate of given academic ability.

We shall estimate the probabilistic choice model (12) under the maintained hypothesis that conditional on (S,R),

$$(13) \quad [\gamma,\delta(1),\delta(0),\epsilon(1),\epsilon(0)] \sim N(0,V),$$

where V is a fixed but unrestricted variance-covariance matrix. The normality assumption aside, perhaps the most restrictive aspect of equation (13) is the condition $E(\gamma|S,R) = 0$. That is, on average, the nonmonetary returns to ability are the same in teaching and nonteaching.

Leaving V unrestricted provides important flexibility. For one thing, it allows for the possibility of compensating variations between the earnings and nonmonetary characteristics of a job. For example if, conditional on (S,R), teaching jobs that pay well tend to have poor working conditions and vice versa, then γ and $\epsilon(1)$ should be negatively correlated, all else equal.

The absence of restrictions on V also allows for any pattern of correlation between $\epsilon(1)$ and $\epsilon(0)$. Consider the possibility that employers in the teaching and nonteaching occupations value the same worker attributes. Then among workers with given values of S and R, a worker who expects relatively high earnings in teaching also should expect relatively high earnings in nonteaching. So $\epsilon(1)$ and $\epsilon(0)$ will be positively correlated, all else equal. On the other hand, it may be that the qualities valued in teaching are not valued in nonteaching. Then, $\epsilon(1)$ and $\epsilon(0)$ will be uncorrelated. Leaving V unrestricted allows for both possibilities.

Under equation (13), the random variable η is normally distributed with mean zero and unrestricted standard deviation σ, conditional on (S,R). Thus, the problem of estimating the probabilistic choice model (12) reduces to that of estimating the parameters A, B, C, and σ. For this to be possible, we must first establish that these parameters are identified.

To see that the parameters are identified, inspect the reduced form equations (8), (9), and (10). The identifiable parameters in equations (8) and (9) include $[(d_1 + a_1), b_1, c_1]$, $[(d_0 + a_0), b_0, c_0]$, and certain functions of the matrix V. The identifiable parameters in equation (10) are $[\{(g_1 + a_1 - a_0)/\sigma\}, \{(b_1 - b_0)/\sigma\}$, and $\{(c_1 - c_0)/\sigma\}]$. It follows that of the parameters A, B, C, and σ appearing in the forecasting model (12), A/σ, B, and C are always identified. σ is identified if either $b_1 \neq b_0$ or $c_1 \neq c_0$.

The condition for identification of σ can be explained. If $b_1 = b_0$ and $c_1 = c_0$, the monetary returns to academic ability are identical in the teaching and nonteaching occupations. Then the probability of choosing teaching is invariant with respect to academic ability. In this case, we cannot infer from the empirical pattern of ability and occupation choice the impact of salary on occupation choice.

11.4.2 Estimation of the Parameters

In principle, equations (8), (9), and (10) can be estimated by the full information maximum likelihood method. (See Maddala 1983, 283 for details.) To obtain the maximum likelihood estimate, a more or less standard iterative optimization algorithm was written. The routine uses the outer product of the score function to generate a search direction. It performs a linear search along this direction using an iterative quadratic inter(extra)polation method. The score function is calculated by applying two-sided numerical derivatives to the log-likelihood function.

Unfortunately, the estimation of switching regressions with endogenous selection is often more difficult in practice than in principle.

Applying the optimization program from a number of alternative starting values, I have not been able to achieve convergent estimates. It turns out that the likelihood is very flat in some regions of the parameter space and has sharp ridges in others. As a consequence, the algorithm produces sequences of estimates that "hang up" in the flat regions and swing wildly across the parameter space in the regions with sharp ridges. Apparently this behavior is not atypical. Several colleagues have reported that they have sometimes experienced similar difficulties in applying maximum likelihood to endogenous switching models.

A simple alternative to maximum likelihood is the two-step approach of Heckman (1976); also see Maddala 1983, 223). The first step ignores the presence of observations of reported earnings and estimates the identifiable parameters of equation (10), namely A/σ, B/σ, and C/σ, by maximum likelihood. We have already reported these estimates in table 11.3.

The second step estimates the identifiable parameters of equation (8) from the subsample of teachers, by least squares regression of $Y(1)$ on an intercept, S,R, and an estimate of the "Mills ratio." The identifiable parameters of equation (9) are estimated in the same manner. The validity of the second step derives from the fact that conditional on S,R, and on being selected into the sample, the expected values of the disturbances $\delta(t1) + \epsilon(t1)$ and $\delta(t0) + \epsilon(t0)$ are

$$(14) \quad E[\delta(1) + \epsilon(1)|S,R,\ \eta < A + B*S + C*R] = -\lambda_1*M(1);$$

$$(15) \quad E[\delta(0) + \epsilon(0)|S,R,\ \eta > A + B*S + C*R] = \lambda_0*M(0).$$

Here $\lambda_1 = E[\{\delta(1) + \epsilon(1)\}*\eta]$, $\lambda_0 = E[\{\delta(0) + \epsilon(0)\}*\eta]$, and $M(1)$ and $M(0)$ are the Mills ratios

$$(16) \quad M(1) = \phi[(A + B*S + C*R)/\sigma]/\Phi[(A + B*S + C*R)/\sigma];$$

$$(17) \quad M(0) = \phi[(A + B*S + C*R)/\sigma]/\{1 - \Phi[(A + B*S + C*R)/\sigma]\}.$$

ϕ is the standard normal density and Φ is the standard normal distribution function. To estimate $M(1)$ and $M(0)$, one uses the first step results.

Note that the least squares estimates reported earlier in table 11.5 differ from the second-step estimates in that they omit the Mills ratio variables. The table 11.5 estimates are inconsistent for the parameters of equations (8) and 9 unless $\lambda_1 = \lambda_0 = 0$. Given that $\eta \equiv -[\gamma + \epsilon(1) - \epsilon(0)]$, the λ coefficients are generally nonzero unless $\epsilon(1)$ and $\epsilon(0)$ are identically zero. But this result holds only if expected earnings in teaching and nonteaching are determined solely by SAT score and high school class rank. Such a sharp restriction is implausible.

Execution of the second step of the two-step method always yields a numerical estimate. As with maximum likelihood, however, application can be less gratifying than the theory suggests. In particular, the fact that S and R are highly collinear with the Mills ratio variables suggests that if the values of λ are far from zero, large samples may be required to obtain useable second-step estimates.

In fact, the second-step estimates obtained on our sex-disaggregated samples were not very credible and had large reported standard errors. Given this, it was natural to consider pooling the samples for males and females in an attempt to obtain more precise estimates. Pooling seemed justified because the slope coefficients of the occupation choice and earnings functions reported in tables 11.3 and 11.5 are very similar for males and females. This suggests that we can safely constrain the slope parameters of equations (8) and (9) to be equal for males and females.

Estimates based on the pooled samples are given in table 11.6. The numbers listed in the "Reported Standard Error" columns do not correct for heteroskedasticity nor for the fact that the Mills ratios have themselves been estimated. Nevertheless, they should at least indicate the orders of magnitude of the true standard errors.

The results in table 11.6 are amazingly sensible, especially given the estimation difficulties described above. Our primary interest is in the estimates of the returns to academic ability. First observe that the partial return to high school class rank is almost identically positive in the teaching and nonteaching occupations, that is, $c_1 \approx c_0 > 0$. This

Table 11.6 Revised Linear Model of Earnings as a Function of Sex, Occupation, and Academic Ability

	All Teachers		All Nonteachers	
Variable	Coefficient	Reported Std. Error	Coefficient	Reported Std. Error
SAT score (400–1600)	0.004	(0.060)	0.127	(0.031)
Class rank (1–100)	0.389	(0.207)	0.326	(0.145)
Intercept	118.	(60.)	128.	(39.)
Sex dummy (1 for females)	12.8	(35.5)	−92.3	(16.6)
Mills ratio $(0 - \infty)$	42.2	(59.7)	140.0	(61.3)
R^2	.07		.11	
Sample size	253		1344	

Note: Earnings are in dollars per week, in 1979.

accords well with our estimates of equation (10), reported in table 11.3. There we found that all else equal, the frequency of choice of teaching as an occupation does not vary with class rank, that is $(c_1 - c_0)/\sigma \approx 0$.

Second, observe that table 11.6 and table 11.3 are in agreement in their estimates of the returns to SAT score. In table 11.3 we saw that all else equal, the frequency of choice of teaching as an occupation falls as SAT score rises, that is $(b_1 - b_0)/\sigma < 0$. In table 11.6 we find that there is no partial return to SAT score in teaching and a positive return in nonteaching, that is $0 \approx b_1 < b_0$.

Recall that σ is identified if either $b_1 \neq b_0$ or $c_1 \neq c_0$. Based on the estimates in tables 11.3 and 11.6, it seems well founded to conclude that the former condition holds and the latter does not. A consistent estimate for σ can be formed by evaluating the identity

$$\sigma = (b_1 - b_0)/[(b_1 - b_0)/\sigma] \tag{18}$$

at the estimates of b_1 and b_0 given in table 11.6 and the estimate of $(b_1 - b_0)/\sigma$ given in table 11.3. We obtain the estimates 0.004 and 0.127 from table 11.6 and -0.0011 from table 11.3. Therefore, our estimate for σ is 111.8.

Now let us consider some other aspects of table 11.6. We find that in teaching, males and females have essentially the same intercepts in their earnings functions. In nonteaching, the intercept for females is \$90 per week lower than for males. This corroborates the pattern of sex differentials observed in table 11.4.

The estimates of the Mills ratio coefficients satisfy $0 < -\lambda_1 < \lambda_0$. This pattern is easily explainable. Observe that

$$-\lambda_1 = E[\{\delta(1) + \epsilon(1)\}*\{\gamma + \epsilon(1) - \epsilon(0)\}] \tag{19}$$

and that

$$\lambda_0 = E[\{\delta(0) + \epsilon(0)\}*\{\epsilon(0) - \epsilon(1) - \gamma\}], \tag{20}$$

and consider the case in which the random variables are mutually independent. Then equations (19) and (20) reduce to $-\lambda_1 = \mathrm{Var}(\epsilon(1)] > 0$ and $\lambda_0 = \mathrm{Var}[\epsilon(0)] > 0$. Moreover, we know from table 11.4 that conditioning on academic ability, the variance of reported earnings in nonteaching is larger than in teaching. This suggests that $\mathrm{Var}[\epsilon(1)] < \mathrm{Var}[\epsilon(0)]$. Thus, there is an inherent predisposition toward the pattern $0 < -\lambda_1 < \lambda_0$. To alter this pattern, the random variables must be mutually dependent in a sufficiently strong and perverse manner.

We earlier pointed out that if $-\lambda_1$ and λ_0 are nonzero, the least squares estimates of table 11.5 are biased. We can with some confidence predict the nature of the bias. Given that $C \equiv c_1 - c_0 \approx 0$, the Mills ratios $M(1)$ and $M(0)$ defined in equations (16) and (17) do not vary with the class rank variable R. We should therefore expect only a small bias, if any, in the estimates of c_1 and c_0 given in table 11.5. Given that $B \equiv b_1 - b_0 < 0$, $M(1)$ is an increasing function of S and $M(0)$ is a decreasing one. Since $-\lambda_1$ and λ_0 are positive, we should expect that the estimate of b_1 in table 11.5 is biased upward and that of b_0 is biased downward.

Comparison of tables 11.5 and 11.6 supports all of these predictions. The estimated returns to class rank are in the neighborhood of 0.35 in both tables. On the other hand, the estimated returns to SAT score differ substantially between the two tables. The estimates of b_1 drop from ~0.045 in table 11.5 to 0.004 in table 11.6. The estimates of b_0 rise from ~0.064 in table 11.5 to 0.127 in table 11.6.

11.5 The Impact of Earnings and Ability Standards on the Teaching Force

In this section we apply the estimated model of occupation choice and earnings to forecast the consequences of some plausible policy proposals. Many parties have suggested that the size and quality of the teaching force can be influenced by combining increases in teacher salaries with the institution of minimum academic ability standards. We shall evaluate policies that combine an across-the-board salary increase of X dollars per week with a minimum SAT score M for certification as a teacher. In practice, the SAT itself would probably not be used as criterion for teacher certification. Our forecasts are of interest if a certification test similar to the SAT is invoked.

Let $D(S,M) = 1$ if $S > M$; $D(S,M) = 0$ otherwise. As earlier, let Φ be the standard normal distribution function. Under equations (12) and (13), the probability that a member of the NLS72 cohort with SAT score S and class rank R is eligible to teach and chooses teaching as his or her occupation is

$$\psi(S,R,X,M) \equiv \Phi[(A + B*S + C*R + X)/\sigma]*D(S,M). \tag{21}$$

To obtain an operational version of equation (21), we use the estimates reported in tables 11.3 and 11.6 and accept the evidence that $C = 0$. Let $F = 1$ if the respondent is female and $F = 0$ if male. Then

$$\psi(F,S,X,M) \equiv \Phi(-0.304 + 0.869*F - 0.0011*S + 0.0089*X)*D(S,M) \tag{22}$$

predicts the probability that a working NLS72 college graduate with SAT score S and sex F would have become a teacher under the policy characterized by (X,M).

Given equation (22), we can easily predict the aggregate behavior of the NLS72 cohort. Let $n = 1, \ldots, N$ designate the NLS72 respondents. Then

(23) $$\hat{\psi}(X,M) \equiv \frac{1}{N}\sum_{n=1}^{N}\hat{\psi}(F_n,S_n,X,M)$$

estimates the fraction of the cohort that would have become teachers under policy (X,M). The average SAT score of those who would have become teachers can be estimated by

(24) $$\Omega(X,M) \equiv \frac{1}{N}\sum_{n=1}^{N} S_n * \hat{\psi}(F_n,S_n,X,M)/\hat{\psi}(X,M).$$

To the extent that the cohort of working NLS72 college graduates are representative of the population from which teachers are drawn, computations of $\hat{\psi}(X,M)$ and $\Omega(X,M)$ provide forecasts of the nationwide effect of policies combining salary increases with academic ability standards.

Table 11.7 reports forecasts for thirty values of the (X,M) pair. The following major results emerge:

—In the absence of a minimum ability standard, increases in teacher earnings yield substantial growth in the size of the teaching force. This result is seen by inspection of the top row of table 11.7. Setting X = \$25

Table 11.7 Forecast Supply and Ability of Teachers as a Function of Earnings and Standards

Minimum SAT Score and Fraction of Cohort above Minimum	Change in Earnings per Week (1979 dollars) +0	+25	+50	+75	+100
400 1.00					
Supply of teachers[a]	.19	0.24	0.30	0.37	0.44
Average SAT score	950	956	961	966	972
600 0.98					
	0.18	0.23	0.29	0.36	0.43
	965	970	974	979	984
700 0.94					
	0.17	0.22	0.27	0.34	0.41
	989	992	996	1000	1004
800 0.88					
	0.15	0.19	0.25	0.31	0.37
	1017	1020	1023	1026	1029
900 0.73					
	0.12	0.15	0.19	0.24	0.30
	1064	1067	1069	1072	1074
1000 0.54					
	0.08	0.10	0.13	0.17	0.20
	1126	1127	1129	1130	1132

[a]Fraction of the cohort of working NLS72 college graduates who have SAT scores above the minimum and are forecast to choose teaching.

is predicted to raise the supply of teachers from 19 percent of the cohort to 24 percent. Setting $X = \$100$ is predicted to raise the supply of teachers to 44 percent of the cohort.

Recall from table 11.4 that the mean reported earnings in 1979 of the NLS72 teachers was about $225 per week. Allowing for the fact that reported earnings may be somewhat lower than expected earnings, $25 is about a 10 percent increase in expected earnings and $100 is about a 40 percent increase. This implies that the aggregate wage elasticity of the supply of teachers ranges from about 2.4 for small increases in salary to about 3.2 for large changes.

—In the absence of a minimum ability standard, increases in teacher earnings yield only a minimal improvement in the average ability of the teaching force. The top row of table 11.7 predicts that as expected earnings increase, the average SAT score of those who choose to teach rises only very slightly, from 950 to 972. This result is easily explained. Increases in expected earnings attract more high-ability students into teaching, but the increases also attract more low-ability students. Overall, the relative growth in low- and high-ability recruits turns out to be comparable to the initial composition of the teaching force.

—If teacher salaries are not increased, institution of a minimum ability standard improves the average ability of the teaching force but reduces its size. The first column of table 11.7 predicts the magnitude of these effects. In particular, requirement of a minimum SAT score of 800 for teacher certification is predicted to raise the average SAT score of the teaching force from 950 to 1,017 but to reduce the supply of teachers from 19 percent to 15 percent of the NLS72 cohort. The average SAT score of all college graduates is not far from 1,017. Thus, setting 800 as the minimum score for certification succeeds in raising average teacher ability to the national average, at the cost of a 20 percent decline in the size of the teaching force.

—The average ability of the teaching force can be improved and the size of the teaching force maintained if minimum ability standards are combined with sufficient salary increases. The entries in table 11.7 reveal that if 800 is established as the minimum SAT score for certification, salaries must be increased by $25 per week in order to maintain the size of the teaching force at 19 percent of the NLS72 cohort. Then the average SAT score of the teaching force is predicted to be 1,020.

If the minimum SAT score is set at 1,000, prevention of a reduction in the size of the teaching force is predicted to require a salary increase of around $90 per week. In this case, the average SAT score of the teaching force is predicted to be about 1,130. Observe that setting the minimum SAT score at 1,000 leaves only 54 percent of the NLS72 cohort eligible to be teachers. Thus, for 19 percent of the cohort to become teachers, about 35 percent of all the eligible, high-ability college graduates must choose to enter teaching. It should not be surprising that a substantial increase in salaries is needed to induce such a large shift from present patterns of behavior.

11.6 Conclusion

Evaluation of proposals to improve the quality of the teaching force requires credible forecasts of the consequences of these proposals. Credible forecasting requires an empirical understanding of the determinants of occupation choices. In this chapter, we have attempted to provide the needed empirical analysis and have offered forecasts derived from it.

Our interpretation of the NLS72 data rests on a number of maintained assumptions. We have taken care to call attention to these assumptions. We have also noted difficulties experienced in executing certain approaches to parameter estimation. Clearly, our analysis should be accepted with caution. At the same time, the analysis should prove useful. In the past, discussion of policies intended to induce more high-ability students to enter teaching has been conducted in a vacuum. Now, some quantitative forecasts have been laid on the table.

Note

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Comment Herman B. Leonard

It seems fitting that we should consider a chapter about what we do—that is, about teaching. We can collectively bemoan the well-established low correlation between academic ability and earnings among teachers, which Manski has ably demonstrated once again. Or we can try to think of counterexamples to his finding—again a familiar theme—that people with greater academic ability have lower probabilities of choosing teaching as a profession. It is hard not to be reminded of the old adage that if you can't do, teach, and if you can't teach, consult. We can at least take some solace in the fact that this chapter goes relatively easy on consultants. It is with our teaching hats on that we should examine this work, and it is good to see a roomful of professional teachers take this problem seriously.

And that is exactly what this chapter does. It takes a real policy problem—it is not an understatement to call it one of the pressing questions on the current national agenda—and takes seriously the task of saying something concrete and intelligent and empirical about it. The chapter is commendable on a variety of grounds. It is technically sound and creative and instructive. It is engaging. It is organized neatly into empirical stages. But what I find most commendable is that one has the sense that the chapter considers its problems to be "for real."

This chapter is also a nice illustration of how difficult serious policy work can be. Nothing ever quite fits together when we look at a real issue. The data are not quite what we want. The model we can fit cannot represent an effect we think is important. We cannot separately identify two forces we would like to distinguish. This work faces all of those problems. The test of a policy paper is whether, when push comes to shove, it bends the issue to fit the empirical technology or the technology to fit the issue. What distinguishes this chapter—and, more broadly, this conference volume—is that the outcome was never in doubt. From start to finish, Manski has concentrated on finding out what he could say about the issue, rather than finding an issue he had something to say about.

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If we stick to the task, this kind of work is difficult. Manski's issue involves the simplest and most direct policy question, and its resolution should turn on the simplest of empirical findings. How much more do we have to pay teachers in order to attract a teaching cadre with better teaching ability? This is what the world would like to know, and it sounds like it should be an emminently approachable empirical question. But of course it is not. We have no reliable measures of teaching quality; we must settle instead for studying the relation between the academic prowess of prospective teachers (as measured by test scores such as the SAT, taken years before a teaching career would start) and the salaries we would have to pay to attract them. Even then, the relationship is difficult to discern in the best data we could reasonably expect to have available. If a question this seemingly simple is this hard, it is no wonder, perhaps, that as researchers we so often seek questions less related to what the world wants to know and more closely linked to what we can find out.

Manski's overall results have the surface plausibility that comes from consonance with what microeconomic choice theory would have predicted. If we raise salaries for teachers without changing the hiring standards, we will wind up with more teachers. If we raise hiring standards without raising salaries, we will wind up with fewer. If we raise both together, we can get the same number but better teachers. The question is, how much do we have to pay to get how much better teachers?

To find out, Manski specifies the simplest model that adequately represents a plausible parsimonious set of relevant influences. His model represents choice between teaching and nonteaching as determined by the (possibly differing) relationships between ability and earnings in the two professions and by an individual characteristic, the relative desirability of teaching compared to the alternative.[1] This model is fit using the sample of nearly 3,000 college graduates in the NLS 1972 high school cohort who were working in 1979. This group included about 500 teachers. These data provide a well-suited test bed for the model and Manski's hypotheses about the effects of earnings on professional choice and the effects of ability on earnings. In particular, the NLS provides about the best we can hope for in measures of academic ability, which, for reasons of observability, we are forced to use in place of teaching ability. We are required to make the imperfect but

1. The one restrictive feature of the model is its requirement that this characteristic be uncorrelated with earning ability in teaching and nonteaching, conditional on academic ability. This is required to separately identify the impact of earnings on choice of profession. It is not a strong restriction—once measured ability is taken into account, it is not obvious why relative earning ability as a teacher should be related to the individual's perceived nonsalary benefits from teaching.

not unreasonable assumption that teachers who were more capable students are more capable teachers as well. At a minimum, it is hard to imagine that higher academic ability results in worse teaching, other things equal.

In technical terms, the results are mixed, though in policy terms they are reasonably clear. As is not uncommon in endogenous switching models of this kind, the likelihood function is not particularly well behaved; it is flat over large regions of the parameter space and has sharp ridges in others—a common econometrician's nightmare. The maximum likelihood estimation approach thus fails, providing one area where the work might be extended later. Additional refinements in programming techniques are not likely to yield better results—the likelihood function simply appears to be ill behaved. This tells us that, with these data and in the context of this model, we know much more about some combinations of the parameters than we do about any particular parameter. It might be worthwhile to examine *what* combinations the likelihood function is tighter on and whether they provide any interesting limits for the parameter values of interest. In a nonlinear model of this type, where the reduced form is being estimated and the parameters of interest are nonlinear combinations of the estimated values, this approach may well not yield any additional insight, but it may be worth a try.

The maximum likelihood results are, not to put too fine a point on it, disappointing—there simply are none. Manski had hoped to estimate the model in the approved way, but the cutting edge of this technology is often dull on problems of this sort. If this chapter were on econometric theory, we might regard this as a technical failure and move on to another problem. But this chapter takes the policy problem seriously; reading it as a failure of econometric technique would be a dramatic underestimate of its contribution both as a policy comment and as an example of how policy-relevant research can be conducted.

As a substitute for maximum likelihood estimates, Manski presents results based on a two-step estimation process. In spite of the difficulties encountered in the maximum likelihood estimation process, the two-stage procedure yields what Manski refers to as "sensible" results. The pattern of coefficients is as expected, and the magnitudes are plausible. The returns to ability are found to be smaller for teachers than for others, as are the returns to class rank. This result is consistent with the overall findings that those with higher academic ability differentially choose not to be teachers.

Manski's description of the results as "sensible" illustrates an important feature of how we learn from the outputs of sophisticated model specifications. We would reject out of hand any results that did not accord fairly closely with our expectations—we would simply decide that the estimation method had not worked. This means that there is a decidedly limited extent to which we are willing to have our beliefs modified by observing results from sophisticated models such as these. This is as it should be because it is hard to tell how robust these results are. But the two-stage estimates do look rather sensible, and we can learn something from them as a result.

The results suggest that there is a high supply elasticity of teachers, on the order of 2 or 3. Manski finds through his simulations based on the two-stage estimates that a 10 percent increase in pay, from $225 to $250, raises the fraction of the cohort predicted to choose teaching as a profession from about one-fifth to about one-fourth. This suggests that for not very much additional money, the degree of selectiveness that can be applied in choosing teachers can be increased fairly substantially; Manski illustrates this by examining how much the minimum standard for SAT scores for teachers could be raised. These results are based on a model that was difficult to estimate, and they must be regarded as tentative. They do, however, suggest a higher sensitivity of job selection to earnings than we might have anticipated.

But to focus on these results only is to miss the force of the essential contribution of Manski's paper as a discussion of a real policy issue and as an example of how policy research can add to the debate. His analysis covers the spectrum from the most basic to the most sophisticated. No one piece of evidence, by itself, would be convincing. But the combined force is substantial. And much of it comes from the power of the most basic observations. Based on the simplest descriptive tabulations, Manski observes: 1. teachers are the less able members of the cohort. 2. Twenty percent of teachers have combined SAT scores below 800. The first is lamentable on general principle; the second drives home its significance. Those who are deeply immersed in the education literature are at least vaguely aware of facts such as these. But stated to a broader audience—and related, as in Manski's paper, to the other findings—they form a part of a startling pattern. Manski goes on to observe, now using more sophisticated modeling strategies (that confirm results obtainable through less sophisticated but not quite correct statistical approaches): 3. In teaching, there is little or no payoff for having higher ability, but there is in other professions. Not only, then, are salaries lower overall for teachers—which could be explained away, for example, by the assertion that teachers have better working conditions—but salaries are lower by a greater margin for the highly able than for the less able. This is a salient plausible justification for the first two findings. But the sophisticated results can give us one more observation that rounds out the story: 4. The selection of teaching as a profession is sensitive to wages offered to teachers. This last result is, of course, the least certain of the conclusions these data lead us to;

it is based on the more complex structure of underlying assumptions and is subject to potentially substantial errors in estimation. It is highly suggestive, however, and if it is right its implications are of dramatic importance.

Some might see this chapter as being merely an elaborate econometric exercise designed to provide an answer about the fourth observation—and, of course, much of the effort in the paper is dedicated to studying that issue. But to be a serious policy paper, it has also to show why the issue is material—and that is exactly what the first three observations establish. Observations (1) and (2) tell us there is a problem. Observation (3) says that we have not explored an obvious potential solution. Observation (4) then provides the hopeful message that there may really be a solution. It is the combination of these findings that could alter the quality and form of the policy debate on this question.

Of course, education policy is likely to be adjusted incrementally. We may decide to raise teacher salaries and to raise the selection standards applied in choosing from among those who apply. We do not need to know exactly what the wage elasticity of supply of teachers is, because we can—and will—adjust salaries gradually to get roughly the number we need. Does this mean that there is no value in trying to find out what the elasticity is? It does not, for two reasons. First, a crucial part of the policy debate is deciding how much more academic quality among teachers we want to buy. We have to decide what adjustments to make in the selection standards for those who aspire to teach. Having some sense for how the trade-off between desired academic standards and teacher salaries might work can help us avoid excessively expensive programs, on the one hand, and ineffective programs, on the other. Second, a finding that the supply elasticity of teachers is high is fundamentally optimistic. It suggests that quality can be improved without enormous additional expenditures. That finding in itself may change the character of the discussion.

Manski's chapter makes contributions on two levels. It provides an illuminating tour of the data on the relation between earnings and ability in teaching, and capably presents an accessible set of important results about an issue of demonstrable practical importance. It is also an example of a serious policy paper. It presents results drawn from techniques ranging from the most simple and direct interpretations of the data to the most sophisticated methods available. It shows the kind of contribution that providing results integrated across a wide spectrum of sophistication can make.