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ACADEMIC ABILITY, EARNINGS, AND THE DECISION TO BECOME A TEACHER: EVIDENCE FROM THE NATIONAL LONGITUDINAL STUDY OF THE HIGH SCHOOL CLASS OF 1972

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ABSTRACT

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. Informed assessment of the various proposals for increasing the quality of the teaching force 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.

The research reported here examines the relationships between academic ability, earnings, and the decision to become a teacher through analysis of data from a national sample of college graduates. Inspection of the data reveals that the frequency of choice of teaching as an occupation is inversely related to academic ability. 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.

An econometric analysis suggests that 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. If teacher salaries are not increased, institution of a minimum ability standard would improve the average ability of the teaching force but reduce its size. 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 + math) of 800 is required for teacher certification and teacher salaries are raised by about ten percent over their present levels.

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ACADEMIC ABILITY, EARNINGS, AND THE DECISION TO BECOME A TEACHER

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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.

The research reported here examines the relationship between academic ability, earnings, and the decision to become a teacher through analysis of data from a national sample of college graduates. 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 post-secondary 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, 3502 reported themselves as having completed a Bachelor's degree in 1976 or 1977. Of this group, 2952 reported that they were working in October, 1979. Of these, 510 reported that 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 non-monetary 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 twenty 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 + math) of 800 is required for teacher certification and teacher salaries are raised by about ten 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 the paper is as follows. Section 2 describes the NLS72 sample and the variables that measure occupation, academic ability, and earnings. Section 3 reports our descriptive analysis of the NLS72 data. The econometric model explaining occupation choices is developed and estimated in Section 4. The model is applied to forecast the effects of policy proposals in Section 5. Section 6 contains brief concluding comments.

2. COMPOSITION OF THE SAMPLE AND DEFINITION OF VARIABLES

A. The Sample

The work in this paper is based entirely on data for the 2952 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).

B. The Occupation Variable

In all that follows, a respondent's occupation is taken to be his declared job type in October, 1979 as coded by the NLS into the three-digit Census classification system. In the crosstabulations of Tables 1,2, and 4, these codes are aggregated into three occupation classes, being:

(a)teachers, exclusive of college faculty(Census codes 141-145)

(b)professional, technical, and kindred workers, exclusive of teachers(Census codes 001-140,150-195)

(c)all other occupations(Census codes 201-992).

In the models of Tables 3,5, and 6, classes (b) and (c) are further aggregated into a single 'non-teaching' 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 quite 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 occupation at age 30(that is about five years into the future). Third, each respondent was asked to report the field in which he received his 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 of interest to redo the analysis using alternative definitions of occupation.

C.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 2952 respondents, class rank information is available for 2287. Either an SAT or ACT score is available for 2468 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).

D.The Earnings Variable

Each respondent working in October, 1979 was asked to report his gross pay per week at his primary job. Hours worked per week at the primary job was also reported. In the parts of this paper concerned with earnings, I restrict attention to the 2335 respondents whose reported hours worked per week is between thirty and sixty and whose reported pay per week is between one hundred and eight hundred dollars. 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 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.

3.PATTERNS OF ACADEMIC ABILITY, OCCUPATION, AND EARNINGS

A.Academic Ability and Occupation

Considering males and females separately, Table 1 partitions the sample respondents into four SAT score groups and, for each group, presents the observed distribution of occupations. In Table 2, percentile class rank in high school is used as the measure of ability. These 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 crosstabulations of SAT scores and occupation basd on NLS72 data have been presented in Vance and Shlachty(1982). Their criteria for inclusion in the sample and for classification of a respondent as a teacher were diffferent than those used here. Their findings were similar.

Table 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 a working college graduate is a teacher conditional on his SAT score and class rank. 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.

B. Academic Ability, Occupation, and Earnings

Considering males and females separately, Table 4 partitions the sample into twelve SAT score-occupation cells. Presented in each cell are (i) mean pay per week, (ii) the number of respondents in the cell, and (iii) the standard deviation of pay per week. I have computed alternative tables using hourly wage as the measure of earnings and class rank as the measure of academic ability and have found patterns very similar to those in Table 4. Among the many interesting features of Table 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 801-1000 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 non-teaching occupations tends to rise with SAT score but the pattern is weak. For teachers, there is little evidence of an earnings-ability pattern. A relationship 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 score in the 1200-1600 range is 1.18 times that of those with score 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 score in the 1000-1200 range is 1.21 times that of females with the same characteristics. Considering teachers with SAT score 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 one hundred dollars 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 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 4, is expressed succinctly in the R2 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.

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 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.

4.A STRUCTURAL INTERPRETATION OF THE OBSERVED PATTERNS

The patterns of academic ability, earnings, and occupation reported in Section 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 non-monetary 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 are 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 sub-model explains occupation choice as a function of the earnings and non-monetary 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 occupation.

A.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 non-monetary job characteristics $g+\gamma(t)$. Here g is a constant and γ varies with t. He aggregates the monetary and non-monetary characteristics into a utility value

(1a) $$u(t1) = y(t1) + g + \gamma(t).$$

Non-monetary 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 non-teaching occupation is

(1b) $$u(t0) = y(t0).$$

Here, we have set the index of non-monetary characteristics equal to zero in order to fix the origin of the utility function. Thus, the term

$g+\gamma(t)$ appearing in (1a) should be interpreted as indexing the non-monetary characteristics of teaching relative to other occupations. Note that in (1a)-(1b), 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 his occupation if

$$(2) \qquad u(t1) - u(t0) = [y(t1) - y(t0)] + g + \gamma(t) > 0.$$

Some obvious objections may be raised against (2). 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 (2) as a working hypothesis.

Empirically, we take the chosen occupation of an NLS respondent to be his 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

$$(3a) \qquad Y(t1) = d_1 + y(t1) + \delta(t1)$$

$$(3b) \qquad Y(t0) = d_0 + y(t0) + \delta(t0)$$

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

$$y(t1) = a_1 + b_1*S(t) + c_1*R(t) + \epsilon(t1) \tag{4a}$$

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

$$y(t0) = a_0 + b_0*S(t) + c_0*R(t) + \epsilon(t0). \tag{4b}$$

The coefficients (b_1,c_1) and (b_0,c_0) quantify the monetary returns to academic ability in the teaching and non-teaching occupations. The variables $\epsilon(t1)$ and $\epsilon(t0)$ 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 (3) and (4) that the reported pay of NLS respondent t is related to his SAT score and high school class rank by

$$\text{(5a)} \qquad Y(t1) = (d_1+a_1) + b_1*S(t) + c_1*R(t) + [\delta(t1)+\epsilon(t1)]$$

if the respondent is a teacher and

$$\text{(5b)} \qquad Y(t0) = (d_0+a_0) + b_0*S(t) + c_0*R(t) + [\delta(t0)+\epsilon(t0)]$$

otherwise. It follows from (2) and (4) that an NLS respondent chooses to be a teacher if and only if

$$\text{(6)} \quad (g+a_1-a_0) +(b_1-b_0)*S(t) + (c_1-c_0)*R(t) + [\gamma(t)+\epsilon(t1)-\epsilon(t0)] > 0.$$

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

$$\text{(7)} \qquad Pr(i=1|S,R) = Pr(\eta < A + B*S + C*R|S,R)$$

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 his occupation is

$$(8) \quad Pr(i=1|S,R,X) = Pr(\eta < A + B*S + C*R + X|S,R,X).$$

If the parameters A,B,C and the distribution of η are known, equation (8) 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 (8) under the maintained hypothesis that conditional on (S,R),

$$(9) \quad [\gamma,\delta(1),\delta(0),\epsilon(1),\epsilon(0)] \sim N(0,V)$$

where V is a fixed but unrestricted variance-covariance matrix. The normality assumption aside, perhaps the most restrictive aspect of (9) is the condition $E(\gamma|S,R) = 0$. That is, on average, the non-monetary returns to ability are the same in teaching and non-teaching.

Leaving V unrestricted provides important flexibility. For one thing, it allows for the possibility of compensating variations between the earnings and non-monetary 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 non-teaching 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 non-teaching. 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 non-teaching. Then, $\epsilon(1)$ and $\epsilon(0)$ will be uncorrelated. Leaving V unrestricted allows for both possibilities.

Under (9), 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 (8) 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 this is so, inspect the reduced form equations (5a)-5(b) and (6). The identifiable parameters in (5a)-(5b) include $[(d_1+a_1),b_1,c_1]$, $[(d_o+a_o),b_o,c_o]$, and certain functions of the matrix V. The identifiable parameters in (6) are $[\{(g_1+a_1-a_o)/\sigma\}$, $\{(b_1-b_o)/\sigma\}$, and $\{(c_1-c_o)/\sigma\}]$. It follows that of the parameters A,B,C,σ appearing in the forecasting model (8), A/σ, B, and C are always identified. σ is identified if either $b_1 \neq b_o$ or $c_1 \neq c_o$.

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

B.Estimation of the Parameters

In principle, (5a)-(5b) and (6) can be estimated by the full information maximum likelihood method. See Maddala(1983), P.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 kind of behavior is not atypical. Several colleagues have reported that they have sometimes experienced similar difficulties in the application of maximum likelihood to endogenous switching models.

A simple alternative to maximum likelihood is the two-step approach of Heckman(1976). Also see Maddala(1983), P.223. The first step ignores the presence of observations of reported earnings and estimates the identifiable parameters of equation (6), namely A/σ, B/σ, and C/σ, by maximum likelihood. We have already reported these estimates in Table 3.

The second step estimates the identifiable parameters of equation (5a) 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 (5b) are estimated in the same manner. As is well known, 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

$$(10a)\quad E[\delta(1)+\epsilon(1) | S,R,\ \eta < A+B*S+C*R] = -\lambda_1 * M(1)$$

$$(10b)\quad E[\delta(0)+\epsilon(0) | S,R,\ \eta > A+B*S+C*R] = \lambda_0 * M(0).$$

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

$$(11a)\quad M(1) = \phi[(A+B*S+C*R)/\sigma]/\Phi[(A+B*S+C*R)/\sigma]$$

$$(11b)\quad M(0) = \phi[(A+B*S+C*R)/\sigma]/\{1-\Phi[(A+B*S+C*R)/\sigma]\}.$$

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 5 differ from the second step estimates in that they omit the Mills ratio variables. The Table 5 estimates are inconsistent for the parameters of

(5a)-(5b) 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 holds only if expected earnings in teaching and non-teaching 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 3 and 5 are very similar for males and females. This suggests that we can safely constrain the slope parameters of (5a)-(5b) to be equal for males and females.

Estimates based on the pooled samples are given in Table 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 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 non-teaching occupations, that is $c_1 \approx c_0 > 0$. This accords well with our estimates of equation (6), reported in Table 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 6 and Table 3 are in agreement in their estimates of the returns to SAT score. In Table 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 6, we find that there is no partial return to SAT score in teaching and a positive return in non-teaching, 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 3 and 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

$$(12)\quad \sigma = (b_1 - b_0)/[(b_1 - b_0)/\sigma].$$

at the estimates of b_1 and b_0 given in Table 6 and the estimate of $(b_1 - b_0)/\sigma$ given in Table 3. We obtain the estimates 0.004 and 0.127 from Table 6 and -0.0011 from Table 3. Therefore, our estimate for σ is 111.8.

Now let us consider some other aspects of Table 6. We find that in teaching, males and females have essentially the same intercepts in their earnings functions. In non-teaching, the intercept for females is ninety dollars per week lower than for males. This corroborates the pattern of sex differentials observed in Table 4.

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

$$(13a) \qquad -\lambda_1 = E[\{\delta(1)+\epsilon(1)\}*\{\gamma+\epsilon(1)-\epsilon(0)\}]$$

$$(13b) \qquad \lambda_0 = E[\{\delta(0)+\epsilon(0)\}*\{\epsilon(0)-\epsilon(1)-\gamma\}].$$

and consider the case in which the random variables are mutually independent. Then (13a)-(13b) reduce to $-\lambda_1 = Var[\epsilon(1)] > 0$ and $\lambda_0 = Var[\epsilon(0)] > 0$. Moreover, we know from Table 4 that conditioning on academic ability, the variance of reported earnings in non-teaching is larger than in teaching. This suggests that $Var[\epsilon(1)] < Var[\epsilon(0)]$. Thus, there is an inherent predisposition towards 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 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 (11a)-(11b) 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 Tables 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 5 is biased upward and that of b_0 is biased downward.

Comparison of Tables 5 and 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 5 to 0.004 in Table 6. The estimates of b_0 rise from ~0.064 in Table 5 to 0.127 in Table 6.

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 (8) and (9), 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 occupation is

$$(14) \qquad \Psi(S,R,X,M) \equiv \Phi[(A + B*S + C*R + X)/\sigma]*D(S,M).$$

To obtain an operational version of (14), we use the estimates reported in Tables 3 and 6 and accept the evidence that C=0. Let F=1 if the respondent is female and F=0 if male. Then

$$(15)\quad \hat{\psi}(F,S,X,M) \equiv \Phi(-0.304 + 0.869*F - 0.0011*S + 0.0089*X)*D(S,M)$$

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 (15), we can easily predict the aggregate behavior of the NLS72 cohort. Let n=1,...,N designate the NLS72 respondents. Then

$$(16)\qquad \hat{\Psi}(X,M) \equiv \frac{1}{N}\sum_{n=1}^{N} \hat{\psi}(F_n,S_n,X,M)$$

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

$$(17)\qquad \hat{\Omega}(X,M) \equiv \frac{1}{N}\sum_{n=1}^{N} S_n * \hat{\psi}(F_n,S_n,X,M)/\hat{\Psi}(X,M).$$

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 $\hat{\Omega}(X,M)$ provide forecasts of the nationwide effect of policies combining salary increases with academic ability standards.

Table 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 is seen by inspection of the top row of Table 7. Setting X = $25

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 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 7 predicts that as expected earnings is increased, the average SAT score of those who choose to teach rises only very slightly, from 950 to 972. This is easily explained. Increases in expected earnings do attract more high ability students into teaching but they 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 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 1017 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 1017. 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 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 1020.

If the minimum SAT score is set at 1000, 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 1130. Observe that setting the minimum SAT score at 1000 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.

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 paper, 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.

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Table 1: Occupation as a Function of Sex and SAT Score

A. Males

Occupation	SAT Score (Verbal + Math) 400 - 800	801 - 1000	1001 - 1200	1201 - 1600
Teacher	.16	.11	.06	.05
Professional	.22	.23	.36	.55
Other	.62	.66	.58	.40
Number of Respondents	148	400	501	221

B. Females

Occupation	400 - 800	801 - 1000	1001 - 1200	1201 - 1600
Teacher	.34	.30	.21	.09
Professional	.14	.26	.36	.46
Other	.52	.43	.43	.45
Number of Respondents	208	421	413	156

Table 2:Occupation as a Function of Sex and High School Class Rank

A.Males

Occupation	Percentile Class Rank 1 - 50	51 - 75	76 - 90	91 - 100
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

Table 3: Probit Model Of Teaching Occupation As A Function of Sex and Academic Ability

	A. Males		B. Females	
Variable	Coefficient	Asymptotic Stnd. Error	Coefficient	Asymptotic Stnd. 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	

Table 4:Pay Per Week as a Function of Sex, SAT Score and Occupation

A.Males

Occupation		SAT Score(Verbal + Math) 400-800	801-1000	1001-1200	1201-1600	Total
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

Occupation		400-800	801-1000	1001-1200	1201-1600	Total
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

*dollars, reported in October, 1979

Table 5: Linear Model of Earnings as a Function of Sex, Occupation, and Academic Ability*

Variable	A. Male Teachers Coefficient	A. Male Teachers Stnd. Error	B. Male Non-Teachers Coefficient	B. Male Non-Teachers Stnd. 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 squared	.04		.03	
Sample Size	64		748	

Variable	C. Female Teachers Coefficient	C. Female Teachers Stnd. Error	D. Female Non-Teachers Coefficient	D. Female Non-Teachers Stnd. 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 squared	.06		.03	
Sample Size	188		593	

*Earnings is in dollars per week, in 1979.

Table 6:Revised Linear Model of Earnings as a Function of Sex, Occupation, and Academic Ability*

	All Teachers		All Non-Teachers	
Variable	Coefficient	Reported Stnd. Error	Coefficient	Reported Stnd. 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 - *	42.2	(59.7)	140.0	(61.3)
R squared	.07		.11	
Sample Size	253		1344	

*Earnings is in dollars per week, in 1979.

Table 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*	.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

*fraction of the cohort of working NLS72 college graduates who have SAT score above the minimum and are forecast to choose teaching