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Estimation of Wage Gains and Welfare Gains from Self-Selection Models

by

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THE ESTIMATION OF WAGE GAINS AND WELFARE GAINS FROM SELF-SELECTION MODELS

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Abstract

In this paper we consider the basic self-selection model for the effects of education, training, unions, and other activities on wages. We show that past models have ignored "heterogeneity of rewards" to the activity--i.e., differences across individuals in the rate of return to the activity--as a source of selection bias. We model such heterogeneity, show how its presence can be tested, and draw out its implications for the wage and welfare gains to the activity. An empirical application provides strong support for such heterogeneity.

THE ESTIMATION OF WAGE GAINS AND WELFARE GAINS FROM SELF-SELECTION MODELS

Economists are often interested in estimating the effect of various types of choices on wages. In labor economics, applications frequently have been made in four areas: (1) education, (2) unions, (3) manpower training, and (4) migration. Researchers on these subjects have become increasingly concerned with the potential self-selection that may arise, mainly because the decisions are made by the individuals themselves. In general, self-selection has been regarded as a disturbing problem for the issue under examination, for ordinary least squares (OLS) or otherwise unadjusted estimates of the parameters of interest are biased if self-selection is present. The usual remedies have been to control for self-selection either by applying the techniques developed by Maddala and Lee (1976), Heckman (1978, 1979), and Lee (1979) (see also Barnow, Cain and Goldberger, 1980), or by trying to avoid the problem by using panel data. Examples of the first approach are Willis and Rosen (1979) and Kenny et al. (1979) for education, Lee (1978) for unions, Nakasteen and Zummei (1980) for migration, and Mallar, Kerachsky, and Thornton (1980) for a jobs program. Examples of the second approach are Kiefer (1979), Bassi (forthcoming), and Nickell (1982) for manpower training.[1]

In this paper we demonstrate the importance of the selection mechanism per se in these types of problems. Our primary goal is to demonstrate the implications of interpreting the self-selection model as a basic model of <u>consumer demand</u>. In the context of the consumer-demand model we show that selection bias occurs because population heterogeneity constraints between the earnings equation and the education-choice equation which we impose in the estimation.

The presence of heterogeneity in the return also has strong implications for public policy, for it implies that those already participating are, in general, those with the highest return. Expanding the participant population--such as by providing educational subsidies or higher stipends to trainees--draws into the activity those who get less out of it. One of the strengths of our model is that it makes this point explicit. Indeed, with our model we can estimate both the mean rewards for those currently participating as well as the reward for those who would participate if the costs of participating were lowered.

In the next section we present our model. Then we provide an empirical illustration, using the case of manpower training. Our empirical results indicate rather dramatically that heterogeneity of rewards are present. We end with suggestions for future research.

I. HETEROGENEITY IN SELF-SELECTION MODELS

As a point of departure we let the individual maximize the utility function $U(Y_i - \phi_i T)$, where T is a dummy variable for participation in the activity, ϕ_i denotes the costs of participating in the activity for individual i, and Y denotes the wage. We assume that ϕ_i captures both monetary costs and a monetized utility component. Let α_i be the earnings gain from participation. Thus, the individual participates in the activity if

$$U(Y_i + \alpha_i - \phi_i) > U(Y_i) \text{ or } \alpha_i > \phi_i$$

where Y_i is now interpreted as earnings in the absence of participation.

terms of demand theory all individuals face the same price of non-participation. Hence some dispersion or heterogeneity in preferences or other costs must be present. Therefore we can only interpret $W_i\eta$ and v_i in the choice equation as observed and unobserved costs, respectively. This follows because the specification of the choice equation in terms of our framework should instead be:[4]

$$T_i^* = \alpha - W_i\eta - v_i,$$

where T_i^* is the net reward. The assumption of homogeneous rewards is in our view rather restrictive. In all the applications we have mentioned it can be argued that every individual is unique in terms of his skills and labor-market situation. Therefore it is reasonable to allow the wage gains to differ between individuals. A straightforward specification allowing for both observed and unobserved heterogeneity would be:

$$\alpha_i = Z_i\delta + u_i \tag{4}$$

where Z_i is a vector of observed variables, δ is its coefficient vector, and u_i is an error term. Reformulating the model with (4) gives us the following:

$$Y_i = X_i\beta + \alpha_i T_i + \varepsilon_i \tag{5}$$

$$\begin{aligned} T_i &= 1 \text{ if } T_i^* > 0 \\ T_i &= 0 \text{ if } T_i^* \leq 0 \end{aligned} \tag{6}$$

$$T_i^* = \alpha_i - \phi_i \tag{7}$$

$$\alpha_i = Z_i\delta + u_i \tag{8}$$

This formulation of the problem alters in a rather interesting fashion the interpretation of the estimated parameters from that usually given. First, note that there is no longer a single "effect" of the program since rewards are heterogeneous. Of course, we can speak of a mean reward, or the reward for an individual with a given characteristics vector Z_i. This could be calculated from the estimated parameter vector δ. The "average" reward is, we assume, that to which the usual constant-parameter estimate in the original equation (1) must correspond. But note that it could easily be negative, zero, or positive but small. Since an individual chooses T=1 only if the unobserved component u_i is sufficiently high, there is no reason for the mean reward to be positive. This obviously has major implications for the interpretation of the wage coefficients in previous studies. A more relevant measure of the rate of return in the mean size of the reward conditional upon choosing T=1:

$$\begin{aligned} E(\alpha_i \mid T_i^* > 0, Z_i\delta, W_i\eta) &= Z_i\delta + E(u_i \mid u_i - v_i > -Z_i\delta + W_i\eta) \\ &= Z_i\delta + (\sigma_{u,u-v}/\sigma_{u-v})[f(s)/(1-F(s))] \end{aligned} \tag{14}$$

where $s = (-Z_i\delta + W_i\eta)\sigma_{u-v}$, and f and F are the standard normal density and distribution functions, respectively (we have assumed normality for the errors). This expression is, we argue, the appropriate measure of the expected wage gain from the activity T.

Likewise, note that the T_i^* quantity in equation (6) is simply the dollar amount of the net reward (net of costs, that is). At mean values or any other values of Z_i and W_i in the population, it may be negative even if the reward α_i is positive. But we can use it to determine the

On the other hand, there is also selection bias in the β and δ coefficients if the unobserved costs, v_i, are correlated with the error term in the earnings equation, ε_i. This is the more usual case of selection bias. It will be eliminated by employing a first-difference technique if v_i is correlated only with some permanent component in the level error. However, note that even if there is no such correlation, the complete self-selection model (10)-(13) must still be estimated if we want to compute our measure of consumer surplus. Hence we conclude that the self-selection model is important per se irrespective of any bias of OLS estimates of the wage equation.

Identification and Estimation

The identification conditions in the full model (10)-(13) are virtually identical to those in the Lee (1979) model and therefore need little discussion. From our two earnings equations (10) and (11), it is clear that the coefficient vectors β and δ are identified, as are their error variances, $(\sigma_\varepsilon^2 + 2\sigma_{\varepsilon u} + \sigma_u^2)$ and σ_ε^2. In equation (13) the vector of parameters η is identified only if there is at least one variable in Z_i that is not in W_i (a similar condition appears in the Lee model). The variance in the same equation, $(\sigma_u^2 - 2\sigma_{uv} + \sigma_v^2)$ is also identified,[8] as are the covariances between the error in equation (13) and those in equations (10) and (11). From these composite variances it can be shown that some normalization is necessary for complete model identification.[9] We have chosen $\sigma_{uv} = 0$.[10] Subject to this normalization we can identify σ_ε, σ_u, σ_v, $\rho_{\varepsilon u}$, and $\rho_{\varepsilon v}$.

The estimation of the model is also no more or less difficult than the estimation of the usual self-selection model. The model can be

$$\lambda_1 = f(-V_i(\hat{\gamma}/\sigma_{u-v}))/[1-F(-V_i(\hat{\gamma}/\sigma_{u-v})]$$

$$\lambda_2 = -f(-V_i(\hat{\gamma}/\sigma_{u-v}))/F(-V_i(\hat{\gamma}/\sigma_{u-v}))$$

$$\psi_1 = \frac{\sigma_{\varepsilon+u,u-v}}{\sigma_{u-v}}$$

$$\psi_2 = \frac{\sigma_{\varepsilon,u-v}}{\sigma_{u-v}}$$

The results give consistent estimates of β and δ. In the third step a modified equation is estimated with probit:

$$T_i = 1 \text{ if } T_i^* > 0$$

$$T_i = 0 \text{ if } T_i^* < 0$$

$$T_i^* = c(Z_i\hat{\delta}) + cW_i\eta + u_i - v_i$$

where $c = 1/\sigma_{u-v}$. Since the coefficient on $(Z_i\hat{\delta})$ is one over the standard deviation, the parameter vector η can be obtained by dividing it into the probit coefficient vector $(\widehat{c\eta})$.

This provides estimates of all the coefficients. The composite variance parameters are also obtainable: σ_{u-v} from the aforementioned "c" coefficient, $\sigma_{\varepsilon+u,u-v}$ and $\sigma_{\varepsilon,u-v}$ thence from the estimates of ψ_1 and ψ_2, and σ_ε^2 and $\sigma_{\varepsilon+u}^2$ by a procedure explained in Lee (1979). (Note that the last two estimates are not needed for evaluation of the wage gain and welfare gain.) The underlying variance parameters are then obtainable from these composites (see n. 9).

The full-information technique is to be preferred for many reasons, most of all because it is more efficient than the limited-information provided by the Board. Information is now available on the persons in the sample who started manpower training from 1976 onwards; 470 persons in the total sample started manpower training from 1976 until May 1982.

Our basic model can be formulated both in terms of wage levels and first differences. In order to maintain comparability with recent American studies of manpower training (especially Kiefer, 1979; and Bassi, forthcoming) we have chosen the latter formulation, even though it has the disadvantage of reducing the sample substantially. The sample characteristics are presented in Appendix B. Our outcome variable is the difference of the log of wages between 1981 and 1974.[13]

For our X, Z, and W variables, we have only the standard choices. Among the X variables we have only included truly predetermined variables--experience, schooling, age, and sex. Experience and schooling are measured prior to training to avoid endogeneity problems with training.[14]

The exact same variables are included in Z, for we have no strong arguments for excluding any of them.[15] Our a priori assumption is that the skills provided by the courses are more useful for those with little general schooling and little experience. The ability to learn might also vary with age. Earlier studies have also shown that the gains are higher for women even though the reason is not clear (see Bassi, forthcoming). We will also experiment with health status and a dummy variable for immigrants among the Z variables.

The costs are more difficult to specify since items like preferences for schooling, foregone income, and size of the training stipend are not included in our data. However, it can be argued that age should be included because it determines the length of the horizon. Also, women

Table 1

Estimates of Full Model

	Full-Information Maximum Likelihood (FIML) (1)	(2)	Earnings Equation (OLS)	Three-Step Estimates[a]
Rewards ($Z\delta$):				
Constant	.023 (.441)	.004 (.441)	.724* (.255)	-3.324 (2.840)
Age	-.025* (.013)	-.025* (.013)	.022* (.008)	-.237 (.126)
Experience	-.016 (.012)	-.016 (.012)	.014 (.009)	-.171 (.109)
Female	-.056 (.127)	-.032 (.156)	.256* (.082)	-.879 (.667)
Schooling	-.081* (.030)	-.079* (.030)	-.013 (.017)	-.783 (.457)
Costs ($W\eta$):				
Constant	-.322 (.272)	-.328 (.285)	—	-5.577
Age	.003 (.010)	.002 (.010)	—	.018
Female	—	.033 (.128)	—	—
General Earnings Growth ($X\beta$):				
Constant	1.074* (0.119)	1.074* (0.124)	1.083	4.203
Age	-.004* (.002)	-.004* (.002)	-.004* (.001)	.0025 (.002)

-table continues-

Table 1, continued

	Full-Information Maximum Likelihood (FIML) (1)	(2)	Earnings Equation (OLS)	Three-Step Estimates[a]
$\rho_{\varepsilon(u-v)}$	-.081 (.865)	-.081 (.860)	—	-.46
Log likelihood	-908.28	-908.18	—	—

Asymptotic standard errors in parentheses.

*Significant at 10 percent level.

[a]Standard errors of η vector and covariances not obtainable in this method.

[b]Estimated $\sigma_v^2 = -26.72$.

[c]$\rho_{\varepsilon u} = \sigma_{\varepsilon u}/(\sigma_\varepsilon \sigma_u)$

[d]$\rho_{\varepsilon v} = \sigma_{\varepsilon v}/(\sigma_\varepsilon \sigma_v)$

values. Indeed, not only are the coefficient magnitudes often implausible, but the estimated variances are sometimes negative and the estimated correlation coefficients are sometimes greater than one in absolute value.[16] We have not conducted any systematic examination of the reasons for these results, but they may be related to the rather low trainee participation rates in the sample (about 5 percent). The three-step technique may be particularly unreliable when such a small tail of the distribution is being fitted.[17]

Table 2 shows the results of testing several of the restrictions regarding heterogeneity in which we are interested. The first column replicates the results from column (1) of Table 1. The second column tests the restriction that all cost parameters are zero ($\sigma_v = \rho_{\varepsilon v} = \eta = 0$). A likelihood ratio test indicates that the restriction is rejected at the 90 percent level but cannot be rejected at the 95 percent level. Thus the four cost parameters are, as a whole, barely significant. Note that in this case the wage gain equals the welfare gain, for participation is determined solely by the reward. Thus we can also conclude that the difference between the welfare gain and the wage gain in the full model is only barely significant.

In the next column we test the restriction that there is no unobserved heterogeneity of rewards ($\sigma_u = \rho_{\varepsilon u} = 0$). A likelihood ratio test overwhelmingly rejects this restriction ($\chi^2 = 56$). Note too that this restriction has a large effect on the δ parameters. This means that merely interacting T with other variables will not give correct estimates. Next we further restrict the model by having no observed heterogeneity of rewards—that is, we restrict the model to have only a constant wage effect of participation. The OLS estimates of this model,

Table 2 (cont.)

	Full Specification (FIML)	No Costs[a]	No Unobserved Heterogeneity of Rewards[b]	Complete Homogeneity of Rewards[c] FIML	OLS	Three-Step[d]
<u>Covariance Matrix</u>						
σ_ε	.315* (.005)	.321* (.003)	.323* (.003)	.321* (.003)	—	3.180
σ_u	.986* (.207)	.946* (.101)	—	—	—	—
σ_v	.228 (.692)	—	.338 (.299)	—[e]	—	—
$\rho_{\varepsilon u}$	-.268 (.524)	-.477* (.065)	—	—	—	—
$\rho_{\varepsilon v}$	-.673* (.239)	—	.037 (.872)	.083 (.404)	—	-.238
<u>Composite Variances</u>						
$\sigma(\varepsilon+u)$	.961* (.113)	.841* (.028)	.323* (.003)	.321* (.003)	—	3.180
$\sigma(u-v)$	1.012* (0.185)	.945* (.101)	.338 (.299)	—[e]	—	—
$\rho(\varepsilon+u)(u-v)$	.973* (.013)	.942* (.010)	-.037 (.872)	-.083 (.404)	—	.238
$\rho_\varepsilon(u-v)$	-.081 (.865)	-.477* (.065)	-.037 (.872)	-.083 (.404)	—	.238
Log Likelihood	-908.28	-912.91	-936.03	-947.43	—	—

Asymptotic standard errors in parentheses
*Significance at 10-percent level
[a] $\sigma_v = \rho_{\varepsilon v} = \eta = 0$
[b] $\sigma_u = \rho_{\varepsilon u} = 0$
[c] $\sigma_u = \rho_{\varepsilon u} = 0$, δ = constant. All η coefficients relative to $\sigma_v = .338$.
[d] Standard errors on covariances and on η vector not obtainable in this method.
[e] Not estimated; σ_v normalized at .338.

necessary in this result, and we expect that positive costs would occur in other applications and that therefore the welfare gain would be smaller than the wage gain.

The standard error on the wage gain is fairly large, equal to .402. This may seem at odds with our above results in the significance of heterogeneity of rewards, but the two findings are quite compatible, as illustrated in Figure 1. The mean reward in the population is -2.11, but the fraction participating (about 5 percent) occupies only the small upper tail of the reward distribution (the shaded region).[18] The conditional mean in that tail is not far from zero, partly because the conditional variance in the tail is naturally large, and partly because in our application we have estimated negative mean costs, as already mentioned--hence many individuals participate even though they have negative wage gains (although a few have very large wage gains).[19] Nevertheless, the likelihood ratio tests reported above indicate that the wage-gain distribution as a whole is a good explainer of participation and the wage gain from participation. A model which collapsed the distribution on the mean would be significantly worse.

The diagram in Figure 1 also shows how our model can be used to predict the effect on earnings of changing the participant population. For example, lowering costs--such as by paying stipends to a training program or providing scholarships for education--would shift $\bar{W}\eta$ to the left (as shown by the arrow) and enlarge the number of participants. Those brought into the program obviously would have smaller wage gains than those already in--hence the mean wage gain must fall. Mathematically, the effect on the mean wage gain of changing costs is:

$$\frac{\partial E(Z_i\delta + u_i \mid T_i = 1, Z_i\delta, W_i\eta)}{\partial(W_i\eta)} = \frac{\sigma_u^2 f(s_i)}{\sigma_{u-v}^2[1 - F(s_i)]}\left\{\frac{f(s_i)}{1 - F(s_i)} - s_i\right\} > 0$$

where the notation can be seen in the notes to Table 3. This expression must be positive because the term in curly brackets is positive (it is the expectation of (T_i^*/σ_{u-v}) conditional upon its being positive).

A related question with a somewhat different answer is what the effect on mean wages in the total population would be if costs were lowered and participation expanded, i.e., whether economy-wide productivity would increase. Expected wage growth in the total population is:

$$\begin{aligned} E(\Delta W_i \mid X_i\beta, Z_i\delta, W_i\eta) &= X_i\beta + \mathrm{Prob}(T_i{=}1)E[Z_i\delta + u_i \mid T_i = 1, Z_i\delta, W_i\eta] \\ &= X_i\beta + [1 - F(s_i)][Z_i\delta + (\sigma_u^2/\sigma_{u-v})f(s_i)/(1 - F(s_i))]. \end{aligned}$$

Hence

$$\partial E(\Delta W_i \mid X_i\beta, Z_i\delta, W_i\eta)/\partial(W_i\eta) = -f(s_i)[(Z_i\delta)\sigma_v^2 + (W_i\eta)\sigma_u^2]/\sigma_{u-v}^3.$$

This effect is a weighted average of rewards and costs, and hence is ambiguous in sign. In particular, the sign can differ between groups with different Z and W characteristics. In our sample, since the mean wage gain and mean costs are both negative, the expression is positive--hence lowering costs and increasing participation would lower mean wage growth. This is again because negative costs imply that many participants who are on the margin have negative wage gains. If costs were instead positive, subsidizing them would bring in participants with positive wage gains and hence could improve mean economy-wide wages. The strength of our model is that these effects can be calculated explicitly.

Table 3

Expected Wage and Welfare Gains[a]

	FIML	Three-Step Technique
Total Population $\bar{Z}$, $\bar{W}$:		
Exp. Wage Gain$_{T=1}$	.065 (.402)	5.306 —[b]
Exp. Wage Gain$_{T=0}$	-2.177 (.919)	-24.791 (10.143)
Exp. Welfare Gain$_{T=1}$	.343 (.353)	3.643 (3.297)
Exp. Welfare Gain$_{T=0}$	-1.969 (.934)	-14.809 (9.126)
Participant Population $\bar{Z}$,$\bar{W}$:		
Exp. Wage Gain$_{T=1}$	.103 (.427)	4.821 —[b]
Exp. Welfare Gain$_{T=1}$	.434 (.384)	4.023 (3.583)
Non-Participant Population $\bar{Z}$,$\bar{W}$:		
Exp. Wage Gain$_{T=0}$	-2.189 (0.920)	-24.902 (10.162)
Exp. Welfare Gain$_{T=0}$	-1.981 (0.941)	-19.929 (9.139)
High-Reward Population $\bar{Z}$,$\bar{W}$:[c]		
Exp. Wage Gain$_{T=1}$	.187 (.484)	4.110 —[b]
Exp. Welfare Gain$_{T=1}$	.534 (.453)	4.922 (4.215)

-table continues-

structure on participation probabilities and mean rewards of participants. Our empirical application to a Swedish manpower training program provides strong evidence of the existence of heterogeneity of rewards.

There are several areas of additional research on this topic. First, it would be interesting to incorporate uncertainty into the model, for participation decisions are presumably based upon some guess about the future returns—the actual return is not known. Another extension would be the incorporation of involuntary nonparticipation into the model, such as would occur if an individual desires to be a member of a union and cannot get a union job, or if an individual desires to enroll in an educational or training program but cannot.[20] Finally, it would of course be interesting to see this model applied to wage equations for education, unions, migration, and other training programs.

been estimated in the applied literature, our formulation only provides an alternative interpretation of the various correlations (albeit an economically important one). The importance of the interpretation of the error terms can be seen by comparing our interpretation to the union model of Lee (1978). Such a comparison has been made recently by Björklund (1983), who shows that in the context of our model Lee's parameter estimates have very different implications for the magnitude of union-wage effects than he supposed.

[6]Both hypotheses are nested in the full model. See notes 9 and 10.

[7]To see this, note that the earnings obtainable by participating and by not participating in the activity respectively can be denoted

$$Y_{ti} = Y_{si} + X_i\beta + Z_i\delta + u_i + \varepsilon_i \text{ if } T_i = 1$$

$$Y_{ti} = Y_{si} + X_i\beta + \varepsilon_i \qquad \text{if } T_i = 0$$

where t and s represent time periods after and before the activity.

[8]Note that the variance of the T* equation is identified, unlike that of a probit equation. The reason is important. The variance is identified because the wage gain $Z_i\delta$ appears in the T* equation with a coefficient of one. This is our restriction from theory--that the participation decision must be a direct function of the dollar wage gain. T* is thus measurable in dollar terms and its scale can be fixed. This also relates to the identification condition on the W and Z vectors. In the Lee model, the same condition appears as a requirement that the coefficients on Y in the selection equation be identified--the variance cannot be identified and is normalized to one. In our model, the theoretical restriction we impose on those coefficients allows us to identify

[13]The concept of fixed effect in this model is thus in relative terms.

[14]In wage change equations it is common to include the change in experience and schooling on the right-hand side. In our case we find that inappropriate because both variables are endogenous; the choice beween participation and nonparticipation implies a choice between different changes in experience and schooling.

[15]Note that there is no identification requirement on X and Z.

[16]The lambda variables in these equations and those in Table 2 are all significant at the 10 percent level.

[17]Another source of imprecision in the model may lie in our not having any variables in the selection equation that can be reasonably excluded from the earnings equation.

[18]We assume $v = 0$ for illustration.

[19]Such would occur in any case, of course, since v ranges to minus infinity. But clearly a negative mean cost results in more participants with negative wage gains than would be the case if costs were positive.

[20]Such a specification would lead to a disequilibrium model of participation (Moffitt, 1981). The bivariate probit model with partial observability would be applicable (Poirier, 1980).

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Appendix A (cont.)

$$r_1 = (z_2 - \rho_{ef} z_1)/(1 - \rho_{ef}^{2})^{1/2}$$

$$r_2 = (z_2 - \rho_{\varepsilon f} z_3)/(1 - \rho_{\varepsilon f}^{2})^{1/2}$$

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