NBER WORKING PAPER SERIES

LABOR MARKET SEGMENTATION AND THE UNION WAGE PREMIUM

William T. Dickens

Kevin Lang

Working Paper No. 1883

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
April 1986

We would like to thank Richard Hansen and David Lilien for comments and Elizabeth Bishop, Philip Bokovoy and especially Daniel Dorrow for research assistance. We would also like to thank Paul Ruud for the use of his non-linear optimization program GNOME and the Institute of Industrial Relations at Berkeley and the National Science Foundation (grant no. SES-8409380) for generous research support. The research reported here is part of the NBER's research program in Labor Studies. Any opinions expressed are those of the authors and not those of the National Bureau of Economic Research.

NBER Working Paper #1883
April 1986

Labor Market Segmentation and the Union Wage Premium

ABSTRACT

Studies of the earnings of union workers have consistently shown that they earn considerably more than nonunion workers. This paper considers whether part of this observed union/nonunion differential is due to unions organizing high paying primary sector jobs. We extend our earlier work on the dual labor market in which we used an unknown regime switching regression to identify two labor market sectors -- a high wage primary sector and a low wage secondary sector. Here we estimate a model where worker's wages are determined by one of three wage equations: a union wage equation, a nonunion primary equation or a nonunion secondary equation. If individuals are in the union sector their sector is treated as known. If they are not then their sector is treated as unknown. Parameter estimates for this model suggest that union/nonunion differences are very large for average workers even when comparing union and nonunion primary workers.

We continue to find distinct primary and secondary sectors with wage equations similar to those that would be expected from the dual market perspective. Since it appears that union workers may be receiving large wage premiums it seems likely that there is non-price rationing of union jobs. If there is, our finding in previous papers of non-price rationing of primary sector jobs may have been due only to the rationing of union jobs. We test for the existence of non-price rationing of nonunion primary sector employment in this three sector model and continue to find evidence that at least black workers find it difficult to secure primary sector employment.

William T. Dickens
Sloan School of Management
M.I.T.
Cambridge, MA 02139
(617) 253-7782

Kevin Lang
School of Social Sciences
University of California
Irvine, CA 92717
(714) 856-6205

I. INTRODUCTION

There are many studies which demonstrate that workers in union jobs earn more than apparently equivalent workers in nonunion jobs (Ashenfelter, 1978; Mellow, 1981; Welch, 1980; Pencavel, 1970; Kahn, 1977; Schmidt, 1978; Farber, 1980; Leigh, 1981; Podgursky, 1980; Duncan and Stafford, 1980; Duncan and Leigh, 1980; Lee, 1978; Oaxaca, 1975). Many have argued that unions are not responsible for these higher wages, but instead unions form in jobs in which high quality workers are employed, that is, unions organize high wage jobs. One response to this argument is to develop models in which union status is endogenous and the unobserved factors determining union status are allowed to be correlated with unobserved worker attributes which affect wages (Duncan and Leigh, 1980; Lee, 1978; Pencavel, 1970; Kahn, 1977; Schmidt, 1978; Farber, 1979; Leigh, 1981; Duncan and Stafford, 1980). When union status is treated as endogenous, statistical analysis continues to indicate the existence of a union/nonunion wage differential. However, dual labor market theory (Doeringer and Piore, 1971) and a number of recent mainstream theories suggest that there is non-price rationing of high wage jobs (Calvo, 1979; Weiss, 1980; Stoft, 1982; Shapiro and Stiglitz, 1984; Bowles, 1985; Dickens, 1985). In this case, it is necessary to control not for the unobserved characteristics of the worker, but for the type of job a worker is in. This paper develops an approach which allows us to compare the wages received by workers in high wage primary sector nonunion jobs with the wages received by equivalent workers in union jobs. We continue to find that workers in union jobs receive a large wage premium.

Unless high union wages compensate union workers for undesirable job characteristics, there must be non-price rationing of union jobs; Abowd and Farber

(1982) and Farber (1983) provide evidence which supports the existence of rationing. It is possible that the rationing of high wage jobs which we observe in two previous papers (Dickens and Lang 1985a&b) is only the rationing of union jobs. Although the numeric importance of unions has declined in the U.S. economy, union workers continue to make up almost one-fifth of the civilian labor force and are more important among the full-time adult male workers who comprised our samples in our earlier studies. This paper extends our earlier work on testing for non-price rationing of high wage (primary) market jobs to a model in which there are three sectors of the labor force -- a union sector, and a nonunion sector consisting of two parts, a primary sector and a secondary sector. The definitions of the primary and secondary sectors are drawn from the literature on dual labor markets and are discussed in greater detail in our earlier work. In essence the primary sector consists of high wage jobs in which there are substantial returns to human capital variables (education and experience) while the secondary sector consists of low wage jobs in which there is little or no return to these variables.

We estimate wage equations for each sector and treat the sector of employment as endogenous. Thus we use an endogenous switching model. The innovation consists of developing a model in which it is known if individuals are in one of the regimes (the union sector), but if they are not in that sector, it is not known in which of the two remaining regimes -- the primary and secondary sectors -- they are employed. We estimate this system of equations using maximum likelihood. As in our previous work, the results provide support for the existence of two distinct types of nonunion employment and for the existence of non-price rationing of primary jobs.

II. THE MODEL

Workers may be employed in either the union or nonunion sectors. If they are employed in the union sector, they receive a wage w_{ui} which is determined according to the following relation

$$(1)\ \ln w_{ui} = X_i B_u + e_{ui}.$$

where X_i is a vector of observed individual characteristics, B_u is a conformable vector of parameters and e_{ui} represents unobserved factors affecting the wage. Workers who are not employed in union jobs may be in either of two other sectors. We specify separate wage equations for these two sectors:

$$(2)\ \ln w_{pi} = X_i B_p + e_{pi}$$

$$(3)\ \ln w_{si} = X_i B_s + e_{si}$$

where the variables are defined analogously to those in equation (1).

In order to conform to the dual labor market view, wages in the primary sector should generally be higher than in the secondary sector and the return to education and experience should be higher in the primary sector.

Even if we were able to identify directly which nonunion workers were in primary jobs rather than in secondary jobs, equations (1) - (3) could not be estimated consistently by ordinary least squares since sector of employment may not be exogenous. If workers were free to choose in which sector they were employed, we would expect to find that workers with unusually high wages in union jobs would be more likely to be employed in those jobs and similarly for primary and secondary sector jobs. Consequently,

$$(4) \qquad E(e_{ji} | \text{ employed in sector } j) \neq 0$$

and the errors in each equation (the e_{ji}s) are likely to be correlated with the explanatory variables. This is a standard sample selection problem of the type discussed in Heckman (1976, 1979), Lee, Maddala and Trost (1980) and Griliches, Hall and Hausman (1978). Solutions to this problem are well known and can be based on either two-stage estimators as in the first three of these papers or on maximum likelihood techniques as in the last of these papers. In either case, one begins by specifying equations which determine the underlying tendency to be in each sector

$$(5) \qquad y^*_{ui} = X_i C_u + v_{ui}$$

$$(6) \qquad y^*_{pi} = X_i C_p + v_{pi}$$

$$(7) \qquad y^*_{si} = X_i C_s + v_{si}$$

where y*s represent unobserved variables measuring tendency to be in each sector, Cs are parameters and vs represent unobserved determinants of sectoral attachment. Individuals are in sector j if

$$(8) \qquad y^*_{ji} > y^*_{ki} \qquad \text{for all } k \text{ not equal to } j.$$

This mechanism may be given a choice theoretic interpretation. If individuals are free to choose their sector of employment, they will choose the sector which gives them the highest utility. Under a set of somewhat restrictive conditions, we can test the assumption of free choice by comparing the parameters of the wage equations with those of the switching equations (5-7). The proof is provided in Dickens and Lang (1985a) and is not repeated here; we limit ourselves to outlining the assumptions and conclusions.

We begin by assuming that individuals are wealth maximizers, i.e. that they are indifferent with respect to their sector of employment, and that they have perfect information regarding lifetime earnings in all sectors. It is then possible to place precise restrictions on the relation between parameters in the wage equations and the switching equations under either of two extreme assumptions: the return to experience in a sector is worth more in that sector than in any other sector and, as Lang and Ruud (1986) find, that individual discount rates are independent of the X_is or, alternatively, that in each sector the wage depends only on total experience and not on the sector in which that experience was acquired. In the first case, individuals will make once and for all decisions regarding their sector of employment and will remain in that sector throughout their work lives. In the second case, workers choose employment in whichever sector they receive the highest wage at that moment. It is possible to show in the first case, approximating the lengths of people's working lives as infinity, that the Cs all equal the Bs except for the experience term (for which C should equal zero) and for the constant term. In the second case, all of the Cs (including experience and constant terms) should equal the Bs.

It is therefore possible to test the free choice hypothesis by testing these cross-equation restrictions. Of course, the assumption that individuals do not care about their sector of employment is restrictive. In particular, we would expect that people care not only about their wages but also about their working conditions. For example, if all workers prefer primary sector employment to employment in the other sectors, the constant term in C_p will be greater than it otherwise would be. The C and B coefficients for schooling may diverge if more educated workers value primary sector employment more than less educated workers do. Thus, we would expect some divergence between the two sets of coefficients. However, we can check whether the divergence is compatible with other evidence on tastes for the nonpecuniary aspects of employment in the sectors.

Whether individuals are free to choose their sector of employment or whether there is some form of nonprice rationing mechanism, estimation proceeds in much the same way. If the individual's sector of employment were known, the system of equations (1) - (3), (5) - (7), could be estimated by a straight-forward extension of the techniques developed in the papers cited above. However, in the present case, estimation is further complicated by the absence of direct information on whether workers are employed in the primary or secondary sector if they are not employed in a union job. Estimation of switching models with unknown regimes is discussed in Goldfeld and Quandt (1976). Dickens and Lang (1985a&b) develop and estimate an endogenous switching model with unknown regimes.

The model developed here combines elements of switching models with known and unknown regimes. The likelihood for individuals who are in the union sector is the likelihood for workers in a particular sector for a model with three known regimes. The likelihood for the primary and secondary sectors consists of two elements of the likelihood for an endogenous switching model with three unknown regimes. The next section develops the likelihood function formally.

III. DERIVATION OF THE LIKELIHOOD FUNCTION

Before we can derive the likelihood function we must specify the distribution of the unobservable individual attributes. Defining $e_{1i}=v_{ui}-v_{si}$ and $e_{2i}=v_{pi}-v_{si}$ we will assume that

$$(9) \quad e_i = \begin{bmatrix} e_{ui} \\ e_{pi} \\ e_{si} \\ e_{1i} \\ e_{2i} \end{bmatrix} \text{ is distributed } N(0,S)$$

and that errors are independent across individuals or $E(e_i e_j')=0$ for all i not equal to j. We denote the elements of S as s_{ij}.

Given the above model and these assumptions about the distribution of e, the likelihood of observing someone with a union job and with characteristics X_i and e_{ui} is

$$\text{Prob}(\text{Union}|X_i,e_{ui})\ \phi(e_{ui},s_{uu})$$

where $\phi(\ )$ is the density function for a mean zero normal random variable with variance s_{uu}. The probability of union membership can be rewritten in terms of observables as

$$(10)\ \text{Prob}(\text{union}|X_i,z_i) =$$

$$B_u[(-X_i(C_u-C_s)-(z_i-X_iB_u)s_{u1}/s_{uu})/(s_{11}-(s^2_{1u}/s_{uu}))^{.5},$$
$$(-X_i(C_u-C_p)-(z_i-X_iB_u)(s_{1u}-s_{2u})/s_{uu})/(s_{11}+s_{22}-2s_{12}-(s_{1u}-s_{2u})^2/s_{uu})^{.5}]$$

where $z_i=\ln w_i$ and where B_u is the integral of the standard normal bivariate density function from the two arguments to infinity with correlation $[s_{11}-s_{12}-(s^2_{1u}-s_{2u}s_{1u})/s_{uu}]/[(s_{11}-s^2_{1u}/s_{uu})(s_{11}+s_{22}-2s_{12}-(s_{1u}-s_{2u})^2/s_{uu})]^{.5}$.

Workers who are not union members have their wages determined by one of two wage equations depending on which sector they are in. Since we do not know in which of the two sectors a nonunion worker is employed, the likelihood of observing someone with characteristics X_i, e_{pi} and e_{si} is

$$\text{(11)}\quad L_i = \text{Prob}(\text{Primary}|X_i,e_{pi})\phi(e_{pi},s_{pp}) + \text{Prob}(\text{Secondary}|X_i,e_{si})\phi(e_{si},s_{ss}).$$

Again, the conditional probabilities of sectoral attachment can be rewritten in terms of parameters and observables as

$$\text{(12)}\quad \text{Prob}(\text{Primary}|X_i,z_i) =$$
$$B_p[(-X_i(C_p-C_s)-(z_i-X_iB_p)s_{p2}/s_{pp})/(s_{22}-(s^2_{p2}/s_{pp}))^{.5},$$
$$(-X_i(C_p-C_u)-(s_{p2}-s_{p1})/s_{pp}(z_i-X_iB_p))/(s_{22}+s_{11}-2s_{12}-(s_{p2}-s_{p1})^2/s_{pp})^{.5}]$$

with correlation

$[s_{22}-s_{12}-(s^2_{2p}-s_{1p}s_{2p})/s_{pp}]/[(s_{22}-s^2_{2p}/s_{pp})(s_{22}+s_{11}-2s_{12}-(s_{2p}-s_{1p})^2/s_{pp})]^{.5}$

and

$$\text{(13)}\quad \text{Prob}(\text{Secondary}|X_i,z_i) =$$

$$B_s[(-X_i(C_s-C_u)+(s_{s1}/s_{ss})(z_i-X_iB_s))/(s_{11}-(s^2_{s1}\ /\ s_{ss})^{.5},$$
$$(-X_i(C_s-C_p)+s_{s2}/s_{ss}(z_i-X_iB_s))/(s_{22}-s_{s2}{}^2/\ s_{ss})^{.5}]$$

with correlation $(s_{12}-s_{1s}s_{2s}/s_{ss})/\ [(s_{11}-s^2{}_{1s}/s_{ss})(s_{22}-s^2{}_{2s}/s_{ss})]^{.5}$

IV. DATA, ESTIMATION AND HYPOTHESIS TESTING.

The model was estimated using data from the January 1983 wave of the Current Population Survey. The sample was restricted to employed male heads of households who had worked more than 1000 hours during the previous year in the private sector, who were over 20 years of age and under 65 and who earned more than the minimum wage. The sample was further restricted to those for whom information on all the following attributes was available: either the person's wage, or if that was not available his normal weekly earnings and his normal weekly hours of work, age, years of education, race, marital status, and whether or not he lived in an SMSA. This left us with a sample of 4392. For the purpose of estimation we constructed people's wages as being equal to their reported hourly wage if that was available. If not, it was set equal to their normal weekly earnings divided by their normal weekly hours of work. The standard measure of job experience was constructed for each person in the sample as his age minus years of schooling minus six. Dummy variables were constructed and set equal to one for those who were white, those who had never been married and those who lived in an SMSA.

Not all the parameters of the model are identifiable. Three of the elements of the variance covariance matrix cannot be identified in cross-section data --

s_{us}, s_{up}, and s_{ps} since we never observe individuals who are simultaneously in two sectors. In addition, only two of the C vectors can be independently identified so we normalize $C_s=0$. In the absence of any other restrictions on the Cs and Ss it is impossible to identify all the parameters of the switching equations. Consequently we normalize s_{11} and s_{22} to equal 1.

As explained above, the hypothesis of free choice between the sectors implies a constraint on the values of the Cs. With C_s constrained to zero, free choice between the primary and secondary sector implies that certain elements of $C_p=B_p-B_s$. However, this test would only be meaningful if tastes for primary versus secondary employment did not depend on people's observed characteristics. In fact, we would expect that they would. The difference can be thought of as the compensating differential necessary to make the average worker of a particular type indifferent between primary and secondary sector employment. In general it is difficult to know a priori how the Cs might differ from B_p-B_s, but an exception to this is the dummy variable for race. Previous studies provide evidence that blacks prefer more stable employment (Viscusi, 1979) and have greater demand for occupational safety (Kahn, 1983). These are characteristics which are commonly associated with primary employment so we would expect blacks to have a preference for primary employment. Thus we would expect that the element of C_p corresponding to the coefficient of white would be less than or equal to B_p-B_s. We would also expect that those with more education would be more likely to prefer primary employment and thus we would expect the relevant element of $C_p \geq B_p-B_s$.

The test of these two hypothesis is less than straight-forward because of the identification problems discussed above. In general, any one equality constraint on a parameter of a switching equation constitutes only a renormalization of the likelihood function since it is then possible to remove the normalization of the variance. Further, it is possible that one or both of the inequality constraints would not be binding. In our two previous papers the C associated with education was estimated to be less than the difference between B_p and B_s and the C associated with white was greater than the relevant difference when the variance of the switching equation was normalized to equal 1. Either of the free choice constraints could have been reconciled with the estimated values for the Cs and Bs by renormalizing, but not both at the same time. To test the joint hypothesis that more educated workers and blacks should have a non-negative compensating difference for taking secondary employment we imposed both the equality constraints, set the variance free, and constructed a likelihood ratio test with one degree of freedom. The natural interpretation is that one constraint constitutes a renormalization and the second a real constraint.

The situation may also arise in which renormalizing the switching equation to make an estimated C consistent with the values for the Bs would be impossible since it would require the standard deviation of the conditional switching equation error to be negative. Thus it is sometimes possible to test a single constraint if the values for the Bs together with the free choice equality constraint imply that the corresponding C has the wrong sign. For example, below we find that the C for the white dummy is positive while B_p-B_s is negative. There is no renormalization which can make the observed coefficients consistent with the joint hypothesis of free choice and that blacks prefer primary sector employment. If we can reject the joint hypothesis that either $B_p-B_s \geq 0$ or $C \leq 0$ for the dummy variable for white then either blacks do not have free choice or they prefer secondary employment relative to whites. To test this compound null hypothesis we find the values of B_p-B_s and C for the white dummy which minimize the value of the Wald statistic subject to the constraint that either $B_p-B_s=0$ or $C=0$. If we can reject for this value of the null we can reject for any value of the compound null.

The likelihood function was written in FORTRAN. The Berndt, Hall, Hall and Hausman (1974) algorithm was used to find the maximum. We used the implementation of this algorithm in Paul Ruud's GNOME program. We estimated the unconstrained model using both analytic and numeric derivatives. Due to problems with the accuracy of the approximation to the bivariate integral at extreme values we found that the maximization algorithm was better behaved when we used numeric derivatives. We also encountered several local maximums in estimating the model, though none were qualitatively different. The coefficient values reported below were those associated with the maximum with the highest likelihood value. For starting values we used coefficient estimates for the primary and secondary sector wage equations and C_p derived from our previous paper (Dickens and Lang 1985b) and with coefficients for C_u-C_p derived from a standard probit on union membership and coefficients for the union wage equation taken from OLS estimates using a union only sample. We obtained qualitatively similar results when we started with the same values for the primary and secondary sector wage equations and switching equation but used the primary sector coefficients for the union sector wage equation.

V. RESULTS

The first two columns in table 1 present OLS estimates of log wage equations for union and nonunion workers. Except for the slightly higher coefficient on white in the union equation these results are very similar to those obtained by others in past studies. In general personal characteristics have a smaller effect on union than nonunion wages.

Turning to the switching model we began by estimating the full system with experience and experience squared excluded from the equations measuring underlying tendency to be in the primary and secondary sectors. As discussed in section III, this specification will be appropriate if individuals are free to choose between primary and secondary employment and if there is "sector-specific" human capital. We then estimated the full system without these two constraints. Both the likelihood ratio test and the Wald test reject the constraints at the .01 level. We therefore limit ourselves to discussion of the unrestricted estimates.

As in our previous work, there is evidence of two distinct sectors in the nonunion labor market (see Table I). In the primary sector, there is a substantial return to education (6%) and to experience in a worker's early years. In the secondary sector, there are returns to both schooling and experience, but the estimated return to schooling is less than half that received by primary workers, and the returns to experience start out being less than half those received by primary workers and peak slightly earlier. The results are therefore supportive of the dual labor market typology although the two sectors are less distinct than those observed in our previous studies. It appears that removing union jobs from the analysis makes it more difficult to isolate the two sectors among nonunion workers.

We also confirm our earlier findings that there is non-price rationing of employment in the primary sector. If we were to interpret the coefficients of the primary/secondary sector switching equation as the difference between the wage parameters plus a compensating differential for secondary employment, our point estimates would indicate that the desire to be in the primary sector increases with education and is higher for whites than for blacks. This second finding contradicts empirical evidence presented above that blacks have stronger preferences for several characteristics of primary employment. Using the test described above, we can easily reject the hypothesis that either the C is negative or that $B_p - B_s$ is positive. The Wald test takes a minimum value of 11.87 in the space of the compound null hypothesis when $B_p - B_s$ is 0 and C is set to .598. The .01 critical value for the Wald statistic with two degrees of freedom is 9.21. Thus we can reject the hypothesis that blacks preferences for primary employment are greater than or equal to whites. Since we also wish to test the hypothesis that more educated workers prefer primary sector employment and since in this case the values of C and $B_p - B_s$ are in the space of the null hypothesis a conservative test of the two compound null hypotheses is to use the Wald value computed above and to compare it to the critical value for four degrees of freedom. In this case the test is no longer significant at the .01 level, but it is at the .05 level (critical value of 9.488).

Further evidence for the existence of non-price rationing of primary jobs is provided by comparison of the wages individuals receive in the primary and secondary sectors. An individual with the average characteristics of a nonunion worker would receive \$11.47 per hour in the primary sector but only \$6.98 in the secondary sector. Moreover, the expected wage in the primary sector is greater for virtually all categories of workers. For example, a white with no labor market experience who had never married and lived outside an SMSA would receive a higher wage in the primary sector if he had at least seven years of education. Workers with labor market experience or those who have been married, live in an SMSA or who are black receive even higher wages in the primary sector relative to the secondary sector. We would expect that most workers would prefer primary employment. Unless the marginal worker's preferences are very different from the average worker's or those who are in secondary employment all would receive very low primary wages, our results indicate that people are not free to choose between the sectors.

Although primary workers generally earn considerably more than secondary workers, we continue to find evidence of a union/primary sector wage differential. The average person in our sample would earn $14.26 in a union job but only $11.32 in a primary job, a difference of 26%. Since secondary wages are even lower than those in the primary sector our estimate of the union/nonunion difference is among the higher estimates found in the literature (Freeman and Medoff, 1981).

Our union wage equation also differs from those obtained by most previous researchers. The return to schooling in the union sector is not only higher than in the secondary sector, it is even higher than in the primary sector. In contrast to the work cited in Freeman and Medoff (1981), we find no evidence to support the view that unions reduce the return to schooling and thus reduce the variance of earnings among workers with the same amount of labor market experience. It is interesting that although our results differ notably from the "standard" finding in the literature, they are consistent with evidence on the union status of workers. Farber (1983) finds no evidence that the desire for a union job or the probability of being selected from the queue for union jobs is related to education. Using dummy variables for several categories of educational attainment Farber's point estimates suggest that the most educated workers are more likely than others to desire a union job and less likely to be selected from the queue although neither coefficient is statistically significant and there is not a monotonic relation between education and these probabilities. If the union/nonunion wage differential decreased with education, we would expect more educated workers to be less likely to choose to enter the queue for union jobs.

The similarity between Farber's and our findings and their divergence from those of other researchers may reflect the more complicated sector selection models we use. Farber explicitly models both the choice to enter the union queue and the selection of workers from that queue. If the highest productivity workers tend not to enter the queue and the lowest productivity workers are not chosen from the queue, a single selection equation may not provide a good approximation to the "true" model. Although our model does not explicitly represent this selection process, the presence of two selection equations allows the highest productivity workers to be assigned to the primary sector and the lowest productivity workers to be assigned to the secondary sector and thus our selection process may be similar to the one in Farber's work. Analysis of our data indicates that the average worker in the primary sector has more "human capital" than the average worker in the union sector since the average nonunion worker has more education than the average union worker and the model indicates that more educated workers are most likely to be in the primary sector.

In contrast to our results for education, our estimates do confirm previous work which suggests that the return to experience is higher in the nonunion sector than in the union sector. The estimated return to experience in the union sector is about half the return in the primary sector and similar to the return obtained in the secondary sector. As can be seen in figures 1 and 2, wages peak at approximately the same time in all three sectors. Maximum earnings are estimated to be at 32 years of experience in the primary sector, 31 years in the union sector and 29 in the secondary sector.

Our results for experience might be taken as evidence in support of the "wage standardization" hypothesis that unions reduce within union skill differentials. However, our results for schooling contradict that hypothesis. Consequently they suggest that the less steep lifetimes earnings profile in union jobs should be attributed to some other factor. One possibility is that unionized firms negotiate for less steeply sloped wage profiles. If upward sloping wage-earnings profiles are paid, at least in part, to solve problems of effort elicitation, quitting, or worker quality (Lazear 1981) then the payment of a union wage premium provides similar incentives and thus lessens the need for the age-earnings profile. (Lang, 1984).

The final results of interest are those concerning race. In the OLS results the coefficient on the dummy variable for race is roughly equal for the union and nonunion samples. With our three sector model the black-white wage differential varies considerably across sectors. Whites earn over 25% more than blacks in the union sector and about 16% more in the secondary sector. In the primary sector the difference is small and statistically insignificant. In contrast to these results, the coefficient on the race dummy is small and insignificant in the switching equation for the union/secondary sector choice and blacks are more likely to be in the union sector relative to the primary sector. As noted above, it seems to be more difficult for blacks to obtain primary sector employment than for whites. Interpreting these results literally it would seem that there is wage discrimination against blacks in the union and secondary sectors, and hiring discrimination in the primary sector. However, these results are not entirely consistent with our past work.

In our first study (Dickens and Lang 1985a) we found no evidence of wage discrimination in the primary market and weak evidence that blacks received higher wages than whites in the secondary market. In our second paper (Dickens and Lang 1985b) we found blacks earning statistically significantly lower wages in the primary sector and higher wages in the secondary sector. Since the coefficient estimates in the first paper were imprecise the results, though different, are easily reconciled. The results for the current study with respect to race in the union and nonunion primary sector may also be consistent with our previous findings. Though we find no evidence of significant wage discrimination in the nonunion primary sector in this study, the two previous studies confounded union and nonunion primary so their finding of significant wage discrimination could have been due to the discrimination in the union sector. However, the finding here of significant wage discrimination against blacks in the secondary sector is not consistent with past studies.

VI. CONCLUSIONS

Our results indicate that the union/nonunion differential cannot be ascribed to the tendency of unions to organize high wage primary jobs. Most workers would earn more in the union sector than they do in the primary sector of the nonunion labor market. Thus, our results are consistent with observations that there is non-price rationing of union jobs.

The results presented here also provide support for our earlier findings that there are two distinct sectors of the labor market, a high wage primary sector with substantial returns to education and experience and a low wage secondary sector with little or no return to education and experience. Evidently our previous findings were not due to our failure to distinguish between primary and union jobs. Even when we take account of the existence of a union sector, there are still two sectors with these distinct characteristics in the nonunion part of the labor market.

Moreover our evidence continues to suggest the existence of non-price rationing of primary jobs in the nonunion sector. The rationing of primary jobs which we found in our earlier work cannot be ascribed to the confounding of the primary and union sectors and to the rationing of union jobs. The higher wages paid in the union sector indicate that there is probably rationing of union jobs and that the typical worker would prefer a union job to a primary job. However, if unable to obtain a union job, he would prefer a primary job to a secondary job.

REFERENCES.

Akerlof, George A. "Labor Contracts as Partial Gift Exchange," Quarterly Journal of Economics 87, November 1982, p543-69.

Ashenfelter, Orley, "Union Relative Wage Effects: New Evidence and a Survey of Their Implications for Wage Inflation," in Economic Contributions to Public Policy R. Stone and W. Peterson eds. (N.Y.: Macmillan, 1978)

Abowd, J. M. and H. S. Farber, "Job Queues and the Union Status of Workers," Industrial and Labor Relations Review, 35, April 1982, 354-67.

Berndt, E.K., R.H. Hall, B. Hall, J. Hausman, "Estimation and Inference in Non-Linear Structural Models," Annals of Economic and Social Measurement, 3, October 1974, p653-65.

Bowles, S."The Production Process in a Competitive Economy: Walrasian, Neo-Hobbesian and Marxian Models," American Economic Review 75, March 1985, p16-36.

Bulow, J. I. and Summers, L. H. "A Theory of Dual Labor Markets with Application to Industrial Policy, Discrimination and Keynesian Unemployment," NBER working paper #1666 (July,1985).

Calvo, G. "Quasi-Walrasian Theories of Unemployment," American Economic Review, May 1979, p102-7.

Dickens, W.T. "Wages, Employment and the Threat of Collective Action by Workers," Berkeley mimeo 1985.

Dickens, W.T. and K. Lang, "A Test of Dual Labor Market Theory," American Economic Review, 75 (1985a) p792-805.

------------ "Testing Dual Labor Market Theory: A Reconsideration of the Evidence," NBER working paper #1670 (1985b).

Doeringer P. and M. J. Piore, Internal Labor Markets and Manpower Analysis, Lexington, Mass.: D.C. Heath (1971)

Duncan, G.H. and D.E. Leigh, "Wage Determination in the Union and Nonunion Sectors: A Sample Selectivity Approach," Industrial and Labor Relations Review 34 (1980), 24-34.

Duncan, G.M. and F.P. Stafford, "Do Union Members Receive Compensating Wage Differentials?" American Economic Review, 70 (June 1980), 355-71.

Farber H. S., "Unionism, Labor Turnover, and Wages of Young Men," Research in Labor Economics, 3, 1980, 33-53.

--------, "The Determination of the Union Status of Workers," Econometrica, 51, September 1983, 1417-38.

Freeman, R.B. and J.L. Medoff, "The Impact of Collective Bargaining: Illusion or Reality?" in U.S. Industrial Relations 1950-1980: A Critical Assessment, J. Stieber, R.B. McKersie and D.Q. Mills eds. Madison Wisconsin: IRRA 1981.

Goldfeld S.M. and R.E. Quandt, "Techniques for Estimating Switching Regressions," in their Studies in Non-Linear Estimation Cambridge: Ballinger, 1976.

Griliches, Z. , R.H. Hall and J. A. Hausman, "Missing Data and Self- selection in Large Panels," Annales de l"insee, 30-31, April- September 1978, 137-176.

Heckman, J. D., "The Common Structure of Statistical Models of Truncation, Sample Selection, and Limited Dependent Variables and A Simple Estimator for Such Models," Annals of Economic and Social Measurement, 5, 4:1979, 475-92.

-----------, "Sample Selection as Specification Bias," Econometrica, January 1979, 153-62.

Kahn , L. M., "Union Impact: A Reduced Form Approach," Review of Economics and Statistics, 59, November 1977, 503-7.

Kahn, S. "Occupational Safety and Worker Preferences," unpublished doctoral dissertation, M.I.T. 1983.

Lang, K. "Unions, Firms, and the Returns to Seniority," Journal of Labor Research, 5, Winter 1984, 81-92.

Lang, K. and Ruud P. A. "Returns to Schooling, Implicit Discount Rates and Black-White Wage Differentials," Review of Economics and Statistics, 68, 1986, forthcoming.

Lazear, E. "Agency, Earnings Profiles, Productivity, and Hours Restrictions," American Economic Review, (September 1981), 606-620.

Lee, L-F. "Unionism and Wage Rates: A Simultaneous Equations Model with Qualitative and Limited Dependent Variables," International Economic Review, 19, (June 1978), p415-33.

Lee, L-F., G. S. Maddala and R. P. Trost, "Asymptotic Covariance Matrices of Two-Stage Probit and Two-Stage Tobit Methods for Simultaneous Equations Models with Selectivity," Econometrica, 48, March 1980, 491-503.

Leigh, D.E. "Do Union Members Receive Compensating Wage Differentials? Note," American Economic Review, 71 (December 1981), 1049-55.

Mellow, W. "Unionism and Wages: A Longitudinal Analysis," Review of Economics and Statistics, 63, (February 1981), 43-52.

------------ "Employer Size and Wages," The Review of Economics and Statistics, 64, (August 1982), 495-501.

Oaxaca, R. L. "Estimation of Union/Nonunion Wage Differentials Within Occupational/Regional Subgroups," Journal of Human Resources, 10, Fall 1975, 529-36.

Pencavel J., An Analysis of the Quit Rate in American Manufacturing Industry, Princeton, N.J., IR Section Princeton: 1970.

Podgursky, Trade Unions and Income Inequality Ph.D. thesis University of Wisconsin, Madison, 1980.

Schmidt, P. "Estimation of a Simultaneous Equations Model with Jointly Dependent Continuous and Qualitative Variables: The Union Earnings Question Revisited," International Economic Review, 19, (June 1978), 453-65.

Shapiro, C. and J.E. Stiglitz, "Equilibrium Unemployment as a Worker Discipline Device," American Economic Review, 74, (June 1984), 433-44.

Stoft, S. "Cheat Threat Theory: An Explanation of Involuntary Unemployment," mimeo, Boston University, May 1982.

Viscusi, W.K. Employment Hazards: An Investigation of Market Performance, Cambridge: Harvard University Press, 1979.

Welch, S.W. "Union-Nonunion Construction Wage Differentials," Industrial Relations, 19, (Spring 1980), 152-62

Weiss, Andrew, "Job Queues and Layoffs in Labor Markets with Flexible Wage Expectations," Journal of Political Economy, 88, (June 1980), 526-38.

Table 1

OLS and Three Section Switching Model of Union and Non-Union Wages

	OLS NON-UNION	OLS UNION	SWITCHING MODEL: WAGE EQUATIONS			SWITCHING MODEL: SWITCHING EQUATIONS	
			UNION	PRIMARY	SECONDARY	UNION-SECONDARY	PRIMARY-SECONDARY
Constant	.477	1.446	1.014	1.215	1.271	-1.629	-4.074
	(.053)	(.086)	(.122)	(.122)	(.084)	(.057)	(.173)
SMSA	.133	.046	.048	.087	.058	.371	.273
	(.015)	(.021)	(.024)	(.023)	(.032)	(.056)	(.060)
Never Married	-.086	-.004	-.030	.010	-.080	-.275	-.235
	(.025)	(.041)	(.042)	(.044)	(.037)	(.102)	(.124)
School	.081	.036	.085	.060	.026	.058	.219
	(.003)	(.005)	(.007)	(.006)	(.007)	(.004)	(.008)
White	.160	.178	.255	.010	.162	.051	.336
	(.027)	(.031)	(.033)	(.034)	(.034)	(.060)	(.074)
Experience	.030	.018	.017	.027	.013	.067	.050
	(.002)	(.004)	(.004)	(.004)	(.005)	(.002)	(.008)
Experience Squared	-.476E-3	-.299E-3	-.263E-3	-.425E-3	-.221E-3	-.956E-3	-.694E-3
	(.053E-3)	(.074E-3)	(.077E-3)	(.078E-3)	(.087E-3)	(.034E-3)	(.153E-3)
Covariance with Union-Secondary Switching Error	--	--	.023	.000	-.060	--	.784
			(.039)	(.002)	(.109)		(.106)
Covariance with Primary-Secondary Switching Error	--	--	-.336	-.002	-.238	.784	--
			(.054)	(.009)	(.135)	(.106)	
Standard Error	.426	.337	.419	.324	.336	1*	1*
Log Likelihood	-1870.18	-355.48			-4340.67		
N	3309	1083			4392		

*Normalized to one

Figure 1

Returns to Experience, OLS

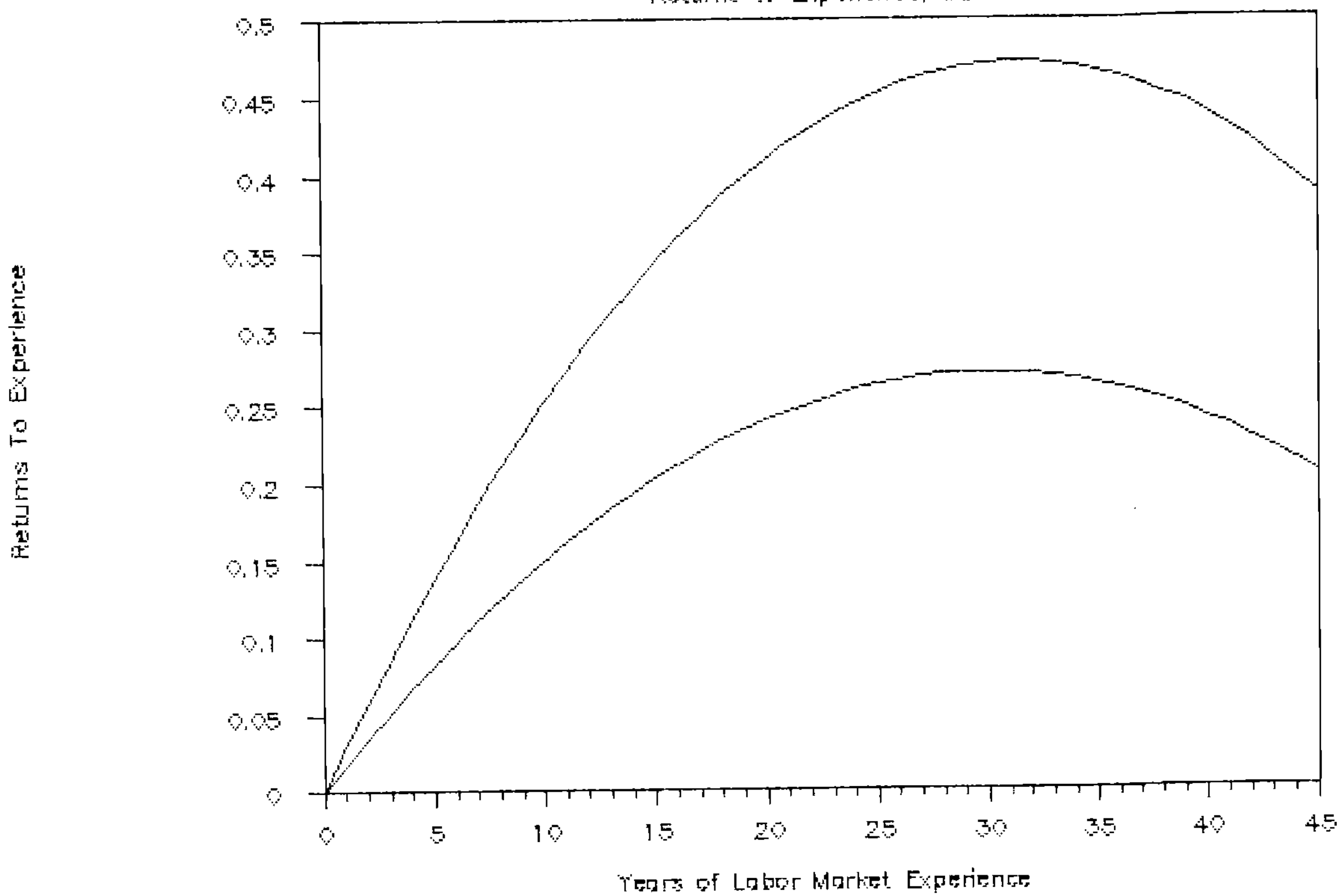
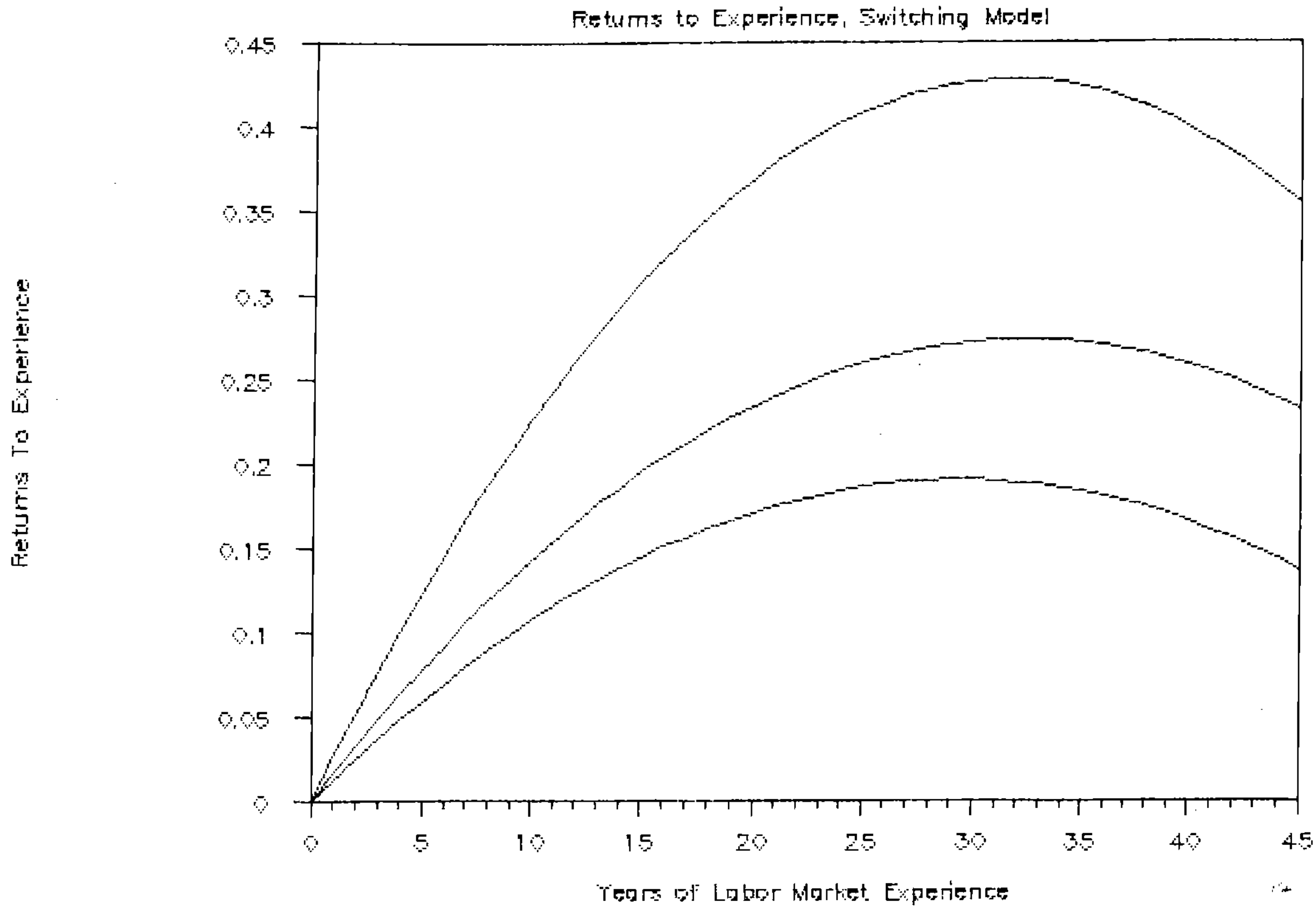


Figure 2


Returns to Experience, Switching Model
Returns To Experience
0
0.05
0.1
0.15
0.2
0.25
0.3
0.35
0.4
0.45
0
5
10
15
20
25
30
35
40
45
Years of Labor Market Experience