Employers' Discriminatory Behavior and the Estimation of Wage Discrimination

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ABSTRACT

This paper considers the linkage of empirical estimates of wage discrimination between two groups, introduced by Oaxaca (1973), to a theoretical model of employers' discriminatory behavior. It is shown that, conditional on different assumptions about employers' discriminatory tastes, Oaxaca's estimators of wage discrimination can be derived. That the approach is more generally useful is demonstrated by deriving an alternative estimator of wage discrimination, based on the assumption that within each type of labor (e.g., unskilled, skilled) the utility function capturing employers' discriminatory tastes is homogeneous of degree zero with respect to labor inputs from each of the two groups. The estimators are compared empirically in an application to male-female wage differentials.

I. Introduction

This paper considers the linkage of empirical estimates of wage discrimination between two groups, as introduced by Oaxaca (1973), to a theoretical model of employers' discriminatory behavior.

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Oaxaca's widely used empirical technique estimates wage discrimination by determining how much of the wage differential between two groups is due to differences in coefficients of separately estimated wage regressions. (For examples, see Blinder 1973, Corcoran and Duncan 1979, Ferber and Green 1982, and Malkiel and Malkiel 1973.) The more general basis of this technique is a comparison of the wage structures for the two groups—captured in coefficients from separately estimated regressions—to an estimate of the wage structure that would prevail in the absence of discrimination (Reimers 1983). The component of the wage differential due to differences between the existing and the "no-discrimination" wage structure is attributed to discrimination. The remainder, which is due to differences in characteristics, is interpreted as nondiscriminatory.[1]

The paper argues that the no-discrimination wage structure used in this approach should be derived from a theoretical model of discriminatory behavior, and shows how this can be done using an extension of Becker's (1957) and Arrow's (1972b) model of employer discrimination. The particular focus is on the relationship between the form of employers' discriminatory tastes and the resulting estimate of wage discrimination. It is shown that under different assumptions about these tastes, embodied in particular characteristics of employers' utility functions, Oaxaca's estimators of wage discrimination can be derived. That the approach is more generally useful is demonstrated by deriving an alternative estimator of wage discrimination based on a different assumption about these tastes, specifically that within each of type of labor (e.g., unskilled, skilled) the utility function is homogeneous of degree zero with respect to labor inputs from each group. This alternative estimator is applied to data from the Young Men and Young Women samples of the National Longitudinal Survey (Center for Human Resources Research 1984), and the results are compared to those produced by Oaxaca's estimators.

Section II briefly reviews Oaxaca's approach to estimating wage discrimination, and shows that it uses two special cases of a more general decomposition of the wage differential. Section III develops the theoretical model of employer discrimination, and shows how different assumptions about employers' discriminatory tastes lead to Oaxaca's estimators, as well as to the alternative estimator of this paper. Section IV presents an empirical application of this alternative estimator, and compares its behavior to Oaxaca's estimators. Section V concludes the paper.

1. A related application of the comparison between current and "no-discrimination" wage structures is to the analysis of the effects of comparable worth (Ehrenberg and Smith 1987; Gregory and Ho 1985; Johnson and Solon 1986).

II. Oaxaca's Approach to Estimating Wage Discrimination

Let $\ln(\bar{w}_m)$ and $\ln(\bar{w}_f)$ be the means of the (natural) logs of male (m) and female (f) wages. (Any two groups could be used; the application to gender differentials in earnings is simply an example.) If the standard log wage model is estimated separately for males and females, then since regression lines pass through the means of the variables

$$(1) \quad \ln(\bar{w}_m) - \ln(\bar{w}_f) = \bar{X}_m' b_m - \bar{X}_f' b_f$$

where $\bar{X}_m$ and $\bar{X}_f$ are vectors containing the means of the variables for males and females, respectively, and b_m and b_f are the estimated coefficients.

Given this result, the log wage differential can be decomposed in two ways. Letting $\Delta\bar{X}' = \bar{X}_m' - \bar{X}_f'$, and $\Delta b = b_m - b_f$, (1) can be written as:

$$(2) \quad \ln(\bar{w}_m) - \ln(\bar{w}_f) = \Delta\bar{X}' b_m + \bar{X}_f' \Delta b$$

or

$$(3) \quad \ln(\bar{w}_m) - \ln(\bar{w}_f) = \Delta\bar{X}' b_f + \bar{X}_m' \Delta b$$

The first term of either (2) or (3) is the part of the log wage differential due to different (average) characteristics of males and females, and the second term is the part of the differential due to different coefficients, or different wage structures. If in the absence of discrimination males and females would receive identical returns for the same characteristics, and differences in wages would therefore be due only to differences in pay-related characteristics, then this second term can be interpreted as the part of the log wage differential due to discrimination. This is the essence of Oaxaca's approach. A critical assumption, maintained in this paper as well, is that labor supply and individual characteristics are fixed, and would not respond to the changes in wages that would result from the elimination of discrimination (Butler 1982).

The question arises as to which of the two equations (2) and (3) to use in empirical work. In general they will yield different answers, and in some cases they are far apart. For example, in Oaxaca's original article, using data on whites, (2) yields an estimate of 52.9 percent of the male-female log wage differential as due to discrimination, while (3) yields an estimate of 63.9 percent. A much larger discrepancy arises in Ferber and Green (1982), in a study of pay discrimination for a sample of university professors, in which (2) gives an estimate of 2 percent of the wage differential as due to discrimination, while (3) gives an estimate of 70 percent.[2]

To understand the source of these differences, consider the assumptions made in using (2) or (3). In (2) it is assumed that in the absence of discrimination the male wage structure would prevail, since in estimating the wage differential that would then exist, the coefficients of the male wage structure are used to weight the differences in characteristics. Conversely, in (3) it is assumed that the female wage structure would prevail. While Oaxaca characterizes the choice of a no-discrimination wage structure as "the familiar index number problem" (1973, 697), a principal goal of this paper is to show that the choice actually hinges on the nature of discriminatory behavior. This, in turn, suggests that b_m and b_f need not be the only no-discrimination wage structures considered.

To see most clearly how the estimate of wage discrimination depends on the choice of a no-discrimination wage structure, consider a more general decomposition of the wage differential.

$$(4) \quad \ln(\bar{w}_m) - \ln(\bar{w}_f) = \Delta\bar{X}'b + [\bar{X}'_m(b_m - b) - \bar{X}'_f(b_f - b)]$$

where b is the no-discrimination wage structure.[3] In this decomposition, as before, the first term can be interpreted as the part of the log wage differential due to differences in characteristics. The second term, similarly, can be interpreted as the part due to discrimination.[4] If it is assumed that in the absence of discrimination the current male wage structure would prevail, then $b = b_m$, and the decomposition in (4) is identical to (2). If instead it is assumed that $b = b_f$, then (4) reduces to (3). Thus Oaxaca's estimators are two special cases of this more general decomposition. More generally, this points out that the critical issue is the choice of b, the no-discrimination wage structure.[5]

In the next section, a theoretical model of employer discrimination is developed, and the relationship between particular assumptions about employers' discriminatory tastes and the choice of the no-discrimination

2. They write, "The results show that under the male reward structure women would earn \$15,199, on average, instead of the \$15,101 they actually earn. Men, on the other hand, would earn only \$16,463 according to the female wage structure, although they actually earn \$19,699" (p. 557). I am grateful to David Bloom for pointing out this finding, which provided the initial motivation for this research.

3. The decomposition in (4) holds as long as (1) holds. This requires only that b_m and b_f are least squares estimates of the log-wage model for the separate samples.

4. The weights on the differences between the coefficients of the discriminatory and the no-discrimination wage structures are more natural in (4) than in Oaxaca's decompositions, in which another "index number problem" arises.

5. This issue was raised by Reimers (1983), who used $b = 0.5 \cdot b_m + 0.5 \cdot b_f$ as the no-discrimination wage structure.

wage structure is explored. The assumptions that justify using either b_m or b_f as the no-discrimination wage structure, as in Oaxaca's approach, are clarified. Based on a different assumption about employers' discriminatory tastes, an alternative, easily estimable no-discrimination wage structure is derived.

III. The Relationship between Employers' Discriminatory Tastes and the No-Discrimination Wage Structure

The model used to explore this question is an extension of the employer discrimination model of Arrow (1972b) and Becker (1957). For expositional purposes most of what follows extends this previous work by allowing for two types of labor; the empirical implementation, however, is based on a generalization to any number of different types of labor.

An economy of identical firms produces output with the technology $f(A,B)$, where A and B are two types of labor (e.g., unskilled and skilled, or blue-collar and white-collar); $f(\cdot,\cdot)$ is strictly concave and increasing. A and B labor may have different productivities, but within each type males and females are homogeneous inputs. Due to the possibility of discrimination, four labor prices exist, w_{MA}, w_{MB}, w_{FA}, and w_{FB}, where w_{MA} is the price of type A male labor, etc. The price of output is normalized to one. Letting M_A, M_B, F_A, and F_B be the inputs of each type of labor, differentiated by gender, profits of each firm are

$$(5) \quad \pi = f(M_A + F_A, M_B + F_B) - w_{MA}M_A - w_{FA}F_A - w_{MB}M_B - w_{FB}F_B$$

Discrimination arises because employers derive utility not only from profits, but also from the gender-composition of their labor force. Employers have identical, strictly concave utility functions

$$(6) \quad U(\pi, M_A, F_A, M_B, F_B)$$

with $U_{MA} \geq 0$, $U_{MB} \geq 0$, $U_{FA} \leq 0$, and $U_{FB} \leq 0$. At least one of these must hold with strict inequality to capture discrimination against women and/or nepotism toward men.

Assume that the supplies of each type of labor, for each gender, are fixed (as in the empirical estimation of wage discrimination) and that firms are at interior solutions. Then the equilibrium quantities of labor in each firm are the total supplies divided by the number of firms. Indexing types of labor by $j = A,B$, the first order conditions for each firm's optimum are

$$(7) \quad U_{\pi}(f_j - w_{Mj}) + U_{Mj} = 0$$
$$U_{\pi}(f_j - w_{Fj}) + U_{Fj} = 0$$

Letting

$$(8) \quad d_{Mj} = \frac{-U_{Mj}}{U_{\pi}} \qquad d_{Fj} = \frac{-U_{Fj}}{U_{\pi}}$$

the first order conditions can be rewritten as

$$(9) \quad w_{Mj} = f_j - d_{Mj}$$
$$w_{Fj} = f_j - d_{Fj}$$

d_{Mj} and d_{Fj} are similar to Becker's (1957) "discrimination coefficients," differing only in that they are not constant, but instead are functions of the partial derivatives of the utility function. Given the signs of the first derivatives of the utility function, (9) implies that

$$(10) \quad w_{Mj} \geq f_j \geq w_{Fj}$$

with strict inequality holding at least once.

Having developed the general model, different assumptions about the structure of employers' discriminatory tastes, and their different implications for the no-discrimination wage structure, can be explored. First, suppose $d_{MA} = d_{MB} = 0$, so that there is no nepotism toward males, but only discrimination against females. Then

$$(11) \quad w_{Mj} = f_j$$
$$w_{Fj} = f_j - d_{Fj}$$

Males are therefore receiving their marginal products. Thus if discrimination were eliminated their wages would not change at all, while wages of females (within each type of labor) would rise to those of males.

To apply this result to data on earnings and characteristics, in order to obtain a no-discrimination wage structure in terms of a set of wage regression coefficients, identify types of labor by different sets of values for the independent variables. Then in the pure discrimination case captured in (11), letting a type j worker be identified by a given set of characteristics X_j, and generalizing to any number of types of labor, implies that the current male wage structure, b_m, is the appropriate no-discrimination wage structure to use in estimating wage discrimination.

The polar opposite case is when there is no discrimination against females, but only nepotism toward males. Then

$$(12) \quad w_{Mj} = f_j - d_{Mj}$$
$$w_{Fj} = f_j$$

In this case male employees take the whole gain from employers' discriminatory behavior, and in its absence their wages would fall to female wages. This characterizes the situation in which it is appropriate to use the current female wage structure, b_f, as the no-discrimination wage structure.

Whether one should accept b_m, b_f, or some other set of coefficients as the no-discrimination wage structure, and if so which one, therefore depends on the nature of employers' discriminatory behavior. The true nature of discrimination may be empirically indiscernible. Previous work by Arrow (1972a) and Goldberg (1982) considers whether discrimination or nepotism is the more likely cause of wage differentials, the central focus being their long-run sustainability.[6]

While the issue remains unresolved, the alternative no-discrimination wage structure derived in this paper is generated by an assumption about employers' discriminatory tastes that imposes neither pure discrimination nor pure nepotism. Instead, it is assumed that employers can be both nepotistic toward males and discriminatory toward females. This avoids a strong asymmetry in employers' tastes such that, for example, they require higher profits to compensate for hiring females, but are not willing to accept lower profits to hire males. Thus d_{MA} and d_{MB} can be negative, and d_{FA} and d_{FB} can be positive, implying that in the absence of discrimination wages of males would fall, while those of females would rise.

The cost of relaxing the pure nepotism or pure discrimination assumption is that some other restriction must be imposed on employers' tastes in order to derive an estimable no-discrimination wage structure. The restriction imposed on the utility function is that, within each type of labor, it is homogeneous of degree zero with respect to male and female labor inputs. Equivalently, if the numbers of male and female workers of a given type are increased or decreased proportionately, the employer's utility is unchanged. Intuitively, this means that employers care only about the relative proportions of males and females, and not absolute numbers. The assumption is less restrictive than it may seem, though,

6. Goldberg (1982) concludes that only nepotism is sustainable, since there is a utility gain from staying in business that compensates the employer for the pecuniary costs of his discriminatory behavior.

since homogeneity of degree zero must hold only within each type of labor, so that the employer's utility is affected by the distribution of workers, by gender, across types of labor.

Given this assumption, the alternative no-discrimination wage structure and its estimator are derived in three steps. First, a theoretical implication of the assumed form of the utility function of employers is derived. Second, as above, a method of applying this result to data on the earnings and characteristics of workers, in order to estimate the no-discrimination wage structure as a set of coefficients in a wage regression, is presented. Third, it is shown that implementation of this estimator is often simple, requiring only the estimation of the log wage regression for the full sample.

Under the assumed form of the utility function, an expression for the no-discrimination wage structure, in terms of wages in the presence of discrimination, emerges. Homogeneity of degree zero for each type of labor implies, by Euler's Theorem, that for each j

$$(13) \quad U_{Mj} \cdot M_j + U_{Fj} \cdot F_j = 0$$

Dividing through by $-U_\pi$ yields

$$(14) \quad d_{Mj} \cdot M_j + d_{Fj} \cdot F_j = 0$$

Using the first order conditions in (9), this implies

$$(15) \quad (f_j - w_{Mj})M_j + (f_j - w_{Fj})F_j = 0$$

or

$$(16) \quad f_j = \frac{w_{Mj}M_j + w_{Fj}F_j}{M_j + F_j}$$

But f_j, the marginal product of a type j worker (of either gender) is, of course, equal to the no-discrimination wage. This shows that, for each type of labor j, the no-discrimination wage can be expressed as the weighted average of the wages for males and females of that type in the presence of discrimination.

An appealing feature of this result is that changes in wages in going from the discriminatory to the nondiscriminatory wage structure are sensitive to the gender-composition of each type of labor. This is clear from (16). Thus the no-discrimination wage for a type of labor that is predominantly female will be relatively closer to the current wage for women, and vice versa. With Oaxaca's method, in contrast, the no-discrimination wage structure is insensitive to these numbers.[7]

7. This problem with Oaxaca's method was raised by Chiplin (1979), who suggested that since there were more males in the labor force, b_m should be used as the no-discrimination wage structure.

To apply this result to data on earnings and characteristics, consider first a simplified case, in which there are only two types of workers, with type A workers described by the vector X_A, and type B workers X_B. While for each gender the wage of a worker with given X may vary, due to the usual causes of residual variation in a wage equation, the most obvious way to estimate, say, the wage for a type A, male worker, is to use the fitted value from the wage regression for the male sample, at $X = X_A$. This leads to four fitted wages, $\hat{w}_{MA}$, $\hat{w}_{FA}$, $\hat{w}_{MB}$, and $\hat{w}_{FB}$. In this simple example the no-discrimination wage structure could be constructed by finding the two points w_A and w_B (or equivalently f_A and f_B) that satisfy (16). (Actually, to preserve linearity, the wage model should be in terms of log wages; this issue is taken up below.) If there are J types of workers instead, each described by X_j, and receiving a (fitted) wage $\hat{w}_{Mj}$ or $\hat{w}_{Fj}$, a similar procedure could be followed. The problem is that there is no reason for the J wages similar to (16) to lie on a line. A reasonable approach is to adopt a least squares criterion in fitting a line to these J points, weighting by the number of workers of each type.

The rationale for using this procedure to estimate the no-discrimination wage structure is as follows. Equation (16) gives the wage for each type of labor that would prevail in the absence of discrimination. As in the Oaxaca decompositions, the implicit simplifying assumption is that individual characteristics are fixed, and would not change if discrimination were eliminated. As long as the form of the standard wage equation would still hold in the absence of discrimination, then it makes sense to estimate the coefficients of this equation in the usual way, by minimizing the sum of squared residuals of the predicted wages around the surface generated by the independent variables.[8]

A minor point that requires discussion is that the desired specification of the wage regression is log-linear, while the theoretical result in (16) is in terms of wages. Taking the logarithm of the no-discrimination wage derived in (16) gives

$$(17)\quad \ln(w_j) = \ln\left[\frac{M_j \cdot w_{Mj} + F_j \cdot w_{Fj}}{M_j + F_j}\right]$$

What is needed instead is

$$(18)\quad \ln(w_j) = \frac{M_j \cdot \ln(w_{Mj}) + F_j \cdot \ln(w_{Fj})}{M_j + F_j}$$

8. A slightly different approach is to minimize this sum over types of labor, rather than individuals, in which case there would be no need to weight by the number of workers of each type. This could lead to somewhat different results if some individuals have identical X vectors.

By Jensen's Inequality, the first expression is greater than the second. For reasonable ranges of the data, though, these numbers are very close. Further, the constant component of this approximation error gets loaded onto the constant term of the wage structure. But Equation (4) shows that the constant term does not affect the decomposition, so that only biases in the slope coefficients are relevant.

It turns out that this estimator of the no-discrimination wage structure can be implemented simply, as the coefficients estimated from the log wage regression for the whole sample, using fitted wages from the separate wage regressions as the dependent variable. To see this, let there be J types of workers, indexed by $j = 1, \ldots, J$, with each type of labor defined by a unique K-vector X_j. Then expressions similar to (16) can be derived for each type of labor j. To determine the wage structure that the fitting procedure described above produces, let M_j and F_j be the number of workers of each gender of type j, $\ln(w_{Mj})$ and $\ln(w_{Fj})$ their respective log wages, and $\ln(\hat{w}_{Mj})$ and $\ln(\hat{w}_{Fj})$ their fitted wages. Denote by Λ the J-vector with jth element

$$(19) \quad \Lambda_j = \left[\frac{M_j}{M_j + F_j}\right] \cdot \ln(\hat{w}_{Mj}) + \left[\frac{F_j}{M_j + F_j}\right] \cdot \ln(\hat{w}_{Fj})$$

and by X the $(J \times K)$ matrix describing each type of worker. Lastly, let $\Omega = diag(M_1 + F_1, \ldots, M_J + F_J)$. Under the least squares criterion discussed above the function to be minimized in estimating the no-discrimination wage structure β is

$$(20) \quad (\Lambda - X\beta)'\Omega(\Lambda - X\beta)$$

Minimizing this with respect to β yields the solution

$$(21) \quad b = (X'\Omega X)^{-1}(X'\Omega\Lambda)$$

Consider instead the simpler approach in which the fitted wages from the separate regressions are regressed on X for the individual workers in this model. Type j workers are still characterized by the same "bundle" of characteristics X_j, but indexing by $i = 1, \ldots, N$ for individual workers, the minimand becomes

$$(22) \quad \sum_{i=1}^{N} (\ln(\hat{w}_i) - X_i\beta)^2$$

Given that X_i is equal across all workers of the same type, and $\ln(\hat{w}_i)$ is equal across same-gender workers of the same type, this is equivalent to

$$(23) \quad \sum_{\substack{men \\ j=1}}^{J} M_j(\ln(\hat{w}_{Mj}) - X_j\beta)^2 + \sum_{\substack{women \\ j=1}}^{J} F_j(\ln(\hat{w}_{Fj}) - X_j\beta)^2$$

Letting Λ_M and Λ_F be J-vectors of fitted male and female log wages for each type of labor, $\Omega_M = diag(M_1, \ldots, M_J)$, and similarly for Ω_F, this can be written as

$$(24)\quad (\Lambda_M - X\beta)'\Omega_M(\Lambda_M - X\beta) + (\Lambda_F - X\beta)'\Omega_F(\Lambda_F - X\beta)$$

Minimizing this with respect to β yields

$$(25)\quad b_{LS} = [(X'\Omega_M X) + (X'\Omega_F X)]^{-1}[(X'\Omega_M\Lambda_M) + (X'\Omega_F\Lambda_F)]$$

It turns out, however, that b and b_{LS} are identical. To see this, note that the first matrices of (21) and (25) are equivalent by inspection. The kth element of the second matrix (which is a vector) in (25) is

$$(26)\quad \sum_{j=1}^{J} (X_{jk}M_j \cdot \ln(\hat{w}_{Mj}) + X_{jk}F_j \cdot \ln(\hat{w}_{Fj}))$$

The kth element of the second matrix in (21) is

$$(27)\quad \sum_{j=1}^{J} X_{jk}(M_j + F_j)\left[\frac{M_j \cdot \ln(\hat{w}_{Mj})}{M_j + F_j} + \frac{F_j \cdot \ln(\hat{w}_{Fj})}{M_j + F_j}\right]$$
$$= \sum_{j=1}^{J} (X_{jk}M_j \cdot \ln(\hat{w}_{Mj}) + X_{jk}F_j \cdot \ln(\hat{w}_{Fj}))$$

which is identical to (26).[9]

Lastly, it is straightforward to show that unless sample weights are being used, identical coefficient estimates result from using actual log wages for individuals instead of fitted values from the separate regressions, so that the no-discrimination wage structure is simply the set of coefficients from the pooled regression. Thus, given the assumed form of the utility function, the model leads to an alternative, easily estimable no-discrimination wage structure to use in estimating wage discrimination.

IV. An Empirical Application

In this section the alternative estimator of wage discrimination is applied to data from the National Longitudinal Survey of Young Men and Young Women (NLS). It turns out that this alternative estimator

9. This proof could only be carried out using a characterization of different types of workers that is an oversimplification, in associating with each type of worker the "bundle" of characteristics X_j. But the result implies that the technique can be applied to actual data in which the regression cannot feasibly be estimated for workers assigned to each of J categories.

yields a lower estimate of discrimination than either of the Oaxaca decompositions. The reason for this lies in the implications of the different assumptions underlying the various estimators of wage discrimination.

Most of the data are taken from the 1980 questionnaires of the NLS, with work history data from earlier surveys also used to construct a measure of actual work experience. Four factors result in final sample sizes of 1,819 men and 1,505 women: (1) attrition from the sample, (2) missing data on some subset of the independent variables, (3) deletion of self-employed and others not working for a wage, and (4) missing wage due to not working or working without pay. The detailed results that follow are given for the data uncorrected for sampling weights. Final estimates of wage discrimination, however, are also given (in Table 3) for data using sampling weights in the NLS, as well as an adjustment for gender-specific attrition rates.

Table 1 presents descriptive statistics and variable definitions. The log wage differential of 0.502 is typical, translating, at the means, into female wages being 61 percent of male wages. The mean values for schooling and experience are higher for males, which is, however, partly due to males being more than 1.5 years older because of the earlier starting date for men in the NLS. This age difference is also probably reflected in the higher proportion married in the male sample. These figures also reflect oversampling of blacks, leading to a relatively high proportion non-white and from the South.

Table 2 presents regression results for one specification of the log wage equation.[10] At the bottom of the table these results are combined with those of Table 1, to estimate wage discrimination based on the two Oaxaca decompositions (denoted b_f and b_m) and the alternative decomposition of this paper (denoted b_{LS}).

The alternative decomposition produces considerably lower estimates of the percentage of the wage differential due to discrimination. Using b_m as the no-discrimination wage structure, it is estimated that 70 percent of the log wage differential between men and women is due to discrimination, while using b_f leads to an estimate of 69 percent. But when the decomposition based on b_{LS} is used, the estimate drops to 57 percent.

Table 3 gives summary results for other specifications, and for the weighted analysis. Quantitatively and qualitatively, the results are very similar. In both the weighted and unweighted analyses, when industry and occupation dummy variables are added, the estimate of discrimination

10. For all of the specifications considered, the hypothesis that a single regression could be fitted to the data for males and females was rejected at the 1 percent significance level. When quadratic terms in *EXPER* and *AGE* were included, they did not have a statistically significant effect, probably due to the relative youth of the sample.

Table 1
Descriptive Statistics

Variable:	Mean (Standard Deviation) Full Sample (N = 3,324)	Female Sample (N = 1,505)	Male Sample (N = 1,819)	Variable Definition
LNWAGE	6.467 (0.515)	6.192 (0.434)	6.694 (0.464)	Natural logarithm of hourly wage
SCHOOL	13.146 (2.565)	12.793 (2.353)	13.437 (2.694)	Highest grade completed
EXPER	8.469 (4.767)	6.763 (3.458)	9.881 (5.218)	Post-schooling actual work experience
AGE	31.670 (3.264)	30.753 (3.073)	32.429 (3.224)	
URBAN	0.696 (0.460)	0.694 (0.461)	0.698 (0.459)	Dummy variable = 1 if lived in SMSA
SOUTH	0.415 (0.493)	0.429 (0.495)	0.403 (0.491)	Dummy variable = 1 if lived in South
UNION	0.321 (0.467)	0.277 (0.448)	0.358 (0.480)	Dummy variable = 1 if wages set by collective bargaining
MST	0.701 (0.458)	0.630 (0.483)	0.760 (0.427)	Dummy variable = 1 if married, spouse present
WHITE	0.749 (0.433)	0.710 (0.454)	0.782 (0.413)	Dummy variable = 1 if white

falls. This is not surprising, given that there is some tendency for women to be concentrated in lower-paying industries, and a strong tendency for the same to be true across occupations. The question of whether industry or occupation dummy variables should be included in regressions to estimate wage discrimination hinges on the extent to which the distribution of men and women across industries and occupations is itself a result of discrimination. The range of results indicates that, together with the choice of the no-discrimination wage structure, the answer to this question has a strong effect on the ultimate estimate of wage discrimination.

The explanation of the lower estimates of discrimination produced by the alternative decomposition of this paper lies in the effects of the distribution of characteristics on the no-discrimination wage structure. The implication of the assumption that leads to the alternative decomposition, as (16) shows, is that the effect of discrimination is to redistribute wages only within each type of labor. Thus the resulting estimate of wage discrimination is sensitive to differences in the distribution of characteristics across men and women. The no-discrimination wage structures used in

Table 2
Regression Estimates and Wage Differential Decompositions[a]

Variable:	Full Sample	Female Sample	Male Sample
CONSTANT	4.418	5.173	4.817
	(0.083)	(0.113)	(0.109)
	[−1.247]	[−1.034]	[−1.006]
SCHOOL	0.088	0.072	0.062
	(0.003)	(0.004)	(0.004)
	[−0.517]	[−0.398]	[−0.369]
EXPER	0.033	0.034	0.011
	(0.002)	(0.003)	(0.002)
	[−0.791]	[−0.661]	[−0.440]
AGE	0.008	−0.009	0.015
	(0.003)	(0.003)	(0.003)
	[2.697]	[2.323]	[2.144]
URBAN	0.175	0.129	0.205
	(0.016)	(0.021)	(0.020)
	[0.185]	[0.154]	[0.149]
SOUTH	−0.077	−0.092	−0.070
	(0.016)	(0.021)	(0.020)
	[0.033]	[0.027]	[0.028]
UNION	0.172	0.138	0.134
	(0.016)	(0.021)	(0.020)
	[−0.015]	[−0.005]	[−0.004]
MST	0.100	−0.001	0.153
	(0.016)	(0.020)	(0.022)
	[0.364]	[0.340]	[0.274]
WHITE	0.169	0.106	0.218
	(0.018)	(0.023)	(0.024)
	[0.291]	[0.255]	[0.225]
R^2	.360	.292	.317
Coefficients used in decomposition	b_{LS}	b_f	b_m
Percent due to characteristics	.43	.31	.30
Percent due to discrimination	.57	.69	.70

a. Standard errors are in parentheses. For the full sample estimates the standard errors and R^2 are for the regression using actual log wages, not fitted values. Proportional contributions to the discriminatory component of the log wage differential, as estimated using coefficients in the column as the no-discrimination wage structure, are given in square brackets.

Table 3
Estimates of Wage Discrimination for Alternative Specifications, and after Adjustment for Sample Weights[a]

Coefficients Used in Decomposition	b_{LS}	b_f	b_m
Unweighted			
Industry Dummy Variables Included			
Percent due to characteristics	.59	.41	.44
Percent due to discrimination	.41	.59	.56
Industry and Occupation Dummy Variables Included			
Percent due to characteristics	.69	.47	.52
Percent due to discrimination	.31	.53	.48
Weighted			
Specification as in Table 2			
Percent due to characteristics	.47	.32	.32
Percent due to discrimination	.53	.68	.68
Industry Dummy Variables Included			
Percent due to characteristics	.63	.40	.49
Percent due to discrimination	.37	.60	.51
Industry and Occupation Dummy Variables Included			
Percent due to characteristics	.72	.46	.54
Percent due to discrimination	.28	.54	.46

a. Industry and occupation dummy variables are for one-digit 1960 *SIC* and *SOC* codes.

Oaxaca's decompositions, however, are insensitive to these differences. The relevance of this to the resulting estimates of wage discrimination can be demonstrated explicitly for the simple case of a model with one explanatory variable, X, and a constant which is the same for men and women. In this case it can be shown that whenever the more highly paid group has the greater endowment of X, the alternative estimator of the no-discrimination wage structure produces a higher coefficient of X than the coefficient in either of the regressions estimated separately for males and females, and hence a lower estimate of wage discrimination than either of Oaxaca's estimators.[11]

11. This analytical result does not hold for the multiple regression model, since each coefficient of the estimated alternative no-discrimination wage structure depends on all of the coefficients from the separate regressions. But experimentation with more parsimonious specifications than the model in Table 2 shows that it tends to hold empirically.

The theoretical material developed above suggests that one should not adopt the estimate provided by the alternative estimator simply because it yields an unambiguous answer, as opposed to Oaxaca's approach. At the same time, the empirical results show that one should not accept the two estimates produced by the Oaxaca decompositions as "a range of possible values" (Oaxaca 1973, 697) for the "true" answer. Jointly, the results demonstrate that decomposing the wage differential between two groups to estimate wage discrimination should not be considered an algebraic exercise independent of the nature of the underlying discriminatory behavior.

V. Conclusion

This paper has considered the problem of linking empirical estimates of wage discrimination to a theoretical model of discriminatory behavior by employers, by deriving implications of the nature of this behavior for the structure of wages that would prevail in the absence of discrimination. It utilized a simple model of employer discrimination to show how particular assumptions about employers' discriminatory tastes can, in this framework, justify Oaxaca's widely-used estimators. But it also showed that different assumptions about these tastes can lead to different estimators of wage discrimination. Furthermore, the application demonstrates that the implications for estimates of wage discrimination are empirically substantial. The application of the general approach to other models of discrimination, and to other assumptions about employers' discriminatory tastes, will be taken up in future research.

At this stage, the only possible conclusion is that there is some arbitrariness in using decompositions of wage differentials to estimate wage discrimination. While it was demonstrated that different estimates can be linked to theoretical models of employer discrimination, the choice of the theoretical model to use remains an open question. But there is no reason for this state of affairs to last. Empirical as well as theoretical explorations of the various models should be pursued. The dependence of these empirical measures of wage discrimination on the underlying model provides an important motivation for doing so.

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