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8

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The Persistence of Male-Female Earnings Differentials

With the passage of the 1963 Federal Equal Pay Act and the 1964 Civil Rights Act, it was hoped that the earnings differentials between the races and the sexes would be substantially reduced. In the case of the male-female earnings differential, it is clear that the high expectations of the federal legislation have yet to be realized. It has been ten years since the passage of the Civil Rights Act; yet the earnings gap between the sexes remains sizable.

This paper examines some of the factors responsible for the continued existence of sizable earnings differentials between the sexes. In and of itself, the existence of an earnings differential is not necessarily indicative of the extent of labor market discrimination. The crucial question is: What proportion of the observed differential is attributable to discrimination and what proportion is justifiable on some generally accepted grounds of equity. Looked at in this way, it is clear that the success or failure of legislation in reducing discrimination cannot be measured by changes in the gross earnings differential alone. Therefore, our analysis will focus on estimated changes in the proportion of the gross male-female earnings differential that result from discriminatory practices in the labor market.

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The outline of the paper is as follows: the underlying analytical model is developed in Section I; the empirical results are presented and discussed in Section II; Section III is a summary and conclusion; and the appendix contains supplementary regressions.

I. ANALYTICAL FRAMEWORK

The Economics of Discrimination

In the Becker model of discrimination, economic agents assert their propensities to discriminate against a given group of workers by acting as if the net cost of dealing with the group is greater than the direct money cost involved [Becker 1971]. For example, employers apply a psychic markup to the nominal wage that would be given to workers whom they have a distaste for hiring. Similarly, consumers apply a psychic markup to the price of a product or service which is produced or sold by workers whom they would prefer to avoid dealing with. In the case of workers, some may psychologically discount the wage they would receive from working with members of a group whom they would prefer to avoid. The extent of an economic agent's propensity to discriminate is measured by the percentage markup or discount applied to members of certain demographic groups. This percentage markup or discount is called the discrimination coefficient, and it measures the psychic costs incurred in dealing with workers from these groups.

Unless men and women are virtually identical with respect to the determinants of earnings, the magnitude of the gross earnings differential tells us little about the extent of discrimination in the labor market. This is because sex differences in the characteristics which determine earnings would generate sex differences in earnings even in the absence of discriminatory employment practices. Of course, such differences in the determinants of earnings reflect societal discrimination in terms of social conditioning from cradle to grave and unequal access to educational and vocational opportunities. In this sense, any difference in earnings would reflect discrimination in a larger context; however, we are interested here in discrimination that stems from the labor market. Thus, personal characteristics are taken as given. The overall effects of labor market discrimination can be measured by the market discrimination coefficient (D), which is defined to be the proportionate difference between the actual male/female earnings ratio and the ratio in the absence of discrimination. In terms of natural logarithms

$$\ln(D+1) = \ln(Y_m/Y_f) - \ln(Y_m^0/Y_f^0) \tag{1}$$

where Y_m and Y_f are the observed full-time earnings of men and women, respectively, and Y_m^0 and Y_f^0 are their respective full-time earnings in the absence of discrimination. According to Becker, the market discrimination coefficient will depend on such factors as the degree of substitutability between male and female workers, market structure, returns to scale, the relative supply of female workers, the average discrimination coefficient, and the dispersion in individual discrimination coefficients.

Suppose the production function can be characterized as $Z = F[L, K]$ and $L = \alpha_m L_m + \alpha_f L_f$ where Z represents output, L represents the labor input, K is vector of other inputs, and α_m and α_f are the efficiency parameters of male (L_m) and female (L_f) labor, respectively. The marginal products of men and women are denoted by MP_m and MP_f, respectively. Thus,

$$MP_m / MP_f = \alpha_m / \alpha_f$$

In the absence of discrimination, cost minimization brings about

$$Y_m^0 / Y_f^0 = \alpha_m / \alpha_f$$

but in a discriminating labor market, net cost minimization implies

$$Y_m / Y_f = (D+1)(\alpha_m / \alpha_f) \tag{2}$$

Let the observed gross male-female earnings differential (G) be defined by

$$G + 1 = Y_m / Y_f \tag{3}$$

Now substituting (3) in (2) and taking logs of both sides we have

$$\ln (G+1) = \ln (D+1) + \ln (\alpha_m / \alpha_f) \tag{4}$$

which is merely a rearrangement of equation 1. According to equation 4, the gross differential (in logs) can be separated into the effects of discrimination and the effects of male/female productivity differences. In the special case of perfect substitutes ($\alpha_m = \alpha_f$), any observed differences in earnings would be totally the result of labor market discrimination.

Cross-Section Model

The methodology employed in this section is derived from the author's previous study of male-female wage differentials [Oaxaca 1973]. The purpose of a cross-section analysis is to estimate the market discrimination coefficient at a point in time. Estimates of the effects of discrimination can be made for two cross sections, 1960 and 1970, and then compared.

We wish to specify and estimate a functional relationship between earnings and various socioeconomic variables: $Y = f(X_1, \ldots, X_k)$; where Y represents earnings, and the X's represent socioeconomic determinants of earnings. A commonly accepted functional form for the earnings relationship is that of the semilog specification which can be justified on the basis of human capital theory [Becker 1966]. Thus,

$$\ln(Y) = \Sigma_k \beta_k X_k$$

where β_k is the percentage effect on earnings from a change in X_k, other things held constant. Given a cross section of male and female workers, we can specify the corresponding statistical models as

$$\ln(Y_{jm}) = \sum_k \beta_{mk} X_{jmk} + \mu_{jm5} \qquad j = 1, \ldots, N_m \tag{5}$$

$$\ln(Y_{jf}) = \sum_k \beta_{fk} X_{jfk} + \mu_{jf6} \qquad j = 1, \ldots, N_f \tag{6}$$

where m and f index men and women, respectively; N_m and N_f are the number of men and women, respectively; j indexes the jth worker in each group; k indexes the kth variable; and μ_{jm5} and μ_{jf6} are the error terms.

We shall assume that in a nondiscriminatory labor market, the parameters in (5) and (6) would be identical for men and women, i.e., in the absence of discrimination, men and women would face a common earnings structure. Accordingly, differences between the sexes in personal characteristics provide the basis for a justifiable earnings differential; whereas male-female differences in the parameters provide the basis for measuring the degree of labor market discrimination. Since we have no way of knowing what the common earnings structure would be in the absence of discrimination, we are forced to make some assumption about this. We shall assume that the male earnings structure given by (5) would be the common earnings relationship that would apply to both men and women in the absence of discrimination. If sex discrimination against women exists, this assumption asserts that the currently observed average earnings of men is exactly what we should observe in the absence of discrimination, but the observed average earnings of women is below what they would receive in a nondiscriminating labor market. We could have assumed, instead, that the female earnings relationship given by (6) is the common earnings structure that would apply in the absence of discrimination. This assumption implies that the currently observed average earnings of women would be equal to their average earnings in the absence of discrimination; however, in this instance, discrimination against women would manifest itself as a situation in which men received, on the average, higher earnings than they would be entitled to in a nondiscriminating labor market.

Our assumption that the male earnings relationships would be the prevailing earnings structure in the absence of discrimination is not an entirely arbitrary assumption. The intents of the 1963 Federal Equal Pay Act and the 1964 Civil Rights Act were not to correct earnings disparities by lowering the earnings of white or male workers, but rather to bring about compliance with the goal of equal employment opportunity by increasing the earnings of affected minority and female workers. Consequently, the achievement of equal employment opportunity would most likely produce a common earnings structure more closely resembling the current male earnings relationship.

By a well-known property of ordinary-least-squares regression we have the following relationships:

$$\ln(\bar{Y}_m) = \sum_k \hat{\beta}_{mk}\bar{X}_{mk}$$

$$\ln(\bar{Y}_f) = \sum_k \hat{\beta}_{fk}\bar{X}_{fk}$$

where $\bar{Y}_m$ and $\bar{Y}_f$ are the geometric mean earnings of men and women, respectively; $\bar{X}_{mk}$ and $\bar{X}_{fk}$ are the average values of the kth variable for men and women, respectively; and $\hat{\beta}_{mk}$ and $\hat{\beta}_{fk}$ are the corresponding estimated coefficients. The gross earnings differential is calculated as

$$\ln(G+1) = \ln(\bar{Y}_m) - \ln(\bar{Y}_f) \tag{7}$$

Under our assumption about the common earnings relationship, we have

$$\ln(\hat{Y}^0_m) = \sum_k \hat{\beta}_{mk}\bar{X}_{mk} = \ln(\bar{Y}_m) \tag{8}$$

$$\ln(\hat{Y}^0_f) = \sum_k \hat{\beta}_{mk}\bar{X}_{fk} \tag{9}$$

where $\hat{Y}^0_m$ and $\hat{Y}^0_f$ are the respective estimated earnings of men and women in the absence of discrimination. If we substitute equations 7, 8, and 9 into the formula for the market discrimination coefficient given by (1), it can be shown that

$$\ln(\hat{D}+1) = \ln(\hat{Y}^0_f) - \ln(\bar{Y}_f) = \sum_k \Delta\hat{\beta}_k\bar{X}_{fk} \tag{10}$$

where $\Delta\hat{\beta}_k = \hat{\beta}_{mk} - \hat{\beta}_{fk}$, and $\hat{D}$ = the estimated market discrimination coefficient. The relative productivity of men is estimated from equations 8 and 9 as

$$\ln(\hat{\alpha}_m/\hat{\alpha}_f) = \ln(\hat{Y}^0_m/\hat{Y}^0_f) = \sum_k \hat{\beta}_{mk}\Delta\bar{X}_k \tag{11}$$

where

$$\Delta\bar{X}_k = \bar{X}_{mk} - \bar{X}_{fk}$$

The decomposition of the gross earnings differential into the effects of discrimination and the effects of productivity differences is accomplished through the substitution of equations 10 and 11 into equation 4:

$$\ln(G+1) = \ln(\hat{D}+1) + \ln(\hat{\alpha}_m/\hat{\alpha}_f) = \sum_k \Delta\hat{\beta}_k \bar{X}_{fk} + \sum_k \hat{\beta}_{mk} \Delta\bar{X}_k$$

The statistical significance of male-female differences in the estimated coefficients can be directly estimated by combining equations 5 and 6 into a single pooled regression in which the male observations are identified by a sex dummy variable:

$$\ln(Y_j) = \sum_k (\Delta\beta_k X_{jk} M_j + \beta_{fk} X_{jk}) + \mu_j \qquad j = 1, \ldots, N_f + N_m$$

where $M_j = 1$ if the worker is male and $=0$ otherwise.

The test for the overall structural difference between the male and female earnings relationships given by (5) and (6), respectively, is equivalent to a joint test of significance corresponding to the $\Delta\beta_k$'s.[1]

Annual rates of change in gross earnings differentials, discrimination, and relative productivity can be estimated from appropriate comparisons between 1960 and 1970 Census cross-section results. Let $\Delta \ln(G+1) = \ln(G+1)_{70} - \ln(G+1)_{60}$ and similarly for $\Delta \ln(D+1)$ and $\Delta \ln(Y_m^0/Y_f^0)$.[2] We then have

$$g = \partial \ln(G+1)/\partial t = \Delta \ln(G+1)/10$$

$$d = \partial \ln(D+1)/\partial t = \Delta \ln(D+1)/10$$

$$m - f = \partial \ln(Y_m^0/Y_f^0)/\partial t = \Delta \ln(Y_m^0/Y_f^0)/10$$

so that $g = d + m - f$: where g = the annual percentage rate of change in the male/female earnings ratio; d = the annual percentage rate of change in the male/female earnings ratio attributable to changes in discrimination; and $m - f$ = the annual percentage rate of change in the male/female earnings ratio attributable to change in relative productivity or labor quality (m and f = the rates of growth of male and female labor quality, respectively).

Time-Series Model

By comparing cross-section results from widely spaced periods in time, we hope to be able to make some judgment about trends in discrimination and relative productivity. The inferences drawn about trends derived from comparisons between two Census years implicitly assume either that cyclical conditions were the same in the two periods or that changes in discrimination are unaffected by the business cycle. This is an important consideration because what is believed to be a trend may in reality be a transitory phenomenon reflecting different cyclical conditions in the periods covered by the Census data. Consequently, a time-series approach will be pursued in a manner similar to that followed by Orley Ashenfelter in his study of changes in racial discrimination over time [Ashenfelter 1970].

The time-series analog of equation 2 is given by

$$Y_m(t)/Y_f(t) = [D(t)+1][\alpha_m(t)/\alpha_f(t)] \tag{12}$$

Suppose that the market discrimination coefficient has constant, trend, and cyclical components. One specification could be

$$D(t)+1 = \exp[C_0 + dt + a_1 V(t)] \tag{13}$$

where d is the percentage change in the male/female earnings ratio attributable to changes in discrimination; and $V(t)$ is a cyclical aggregate demand variable. In periods of tight labor markets, the cost of employment discrimination rises due to a general shortage of labor. Also, any worker resistance to equal employment opportunities for women may become less acute in periods of relative prosperity. A reasonable cyclical indicator is given by the unemployment rate of white males 35–44. Accordingly, we would expect $a_1 > 0$. Allowing for exponential growth in marginal productivities, we have

$$\alpha_m(t)/\alpha_f(t) = [\alpha_{0m}\exp(mt)]/[\alpha_{0f}\exp(ft)] \tag{14}$$

where m and f are the growth rates of male and female labor productivities, respectively. Ideally, it would be preferable to use wage rates rather than earnings in examining discrimination in rates of pay; however, time-series wage data are not disaggregated according to sex. The use of earnings data for year-round full-time workers is an effort to circumvent the lack of wage-rate data. Unfortunately, there still remains the problem that even among year-round full-time workers, men and women do not typically work the same number of hours. It has been estimated that men in this category may work as many as 10 percent more hours during a year than women [Council of Economic Advisers (CEA) 1973]. As an attempt to capture variations in relative hours worked, we introduce the ratio of the female unemployment rate, $U_f(t)$, to the male unemployment rate, $U_m(t)$, as a proxy for the relative use of female labor. If we now substitute equations 13 and 14 into 12, take logs of both sides, and add the labor utilization term and a disturbance term, we arrive at the operational representation of equation 4:

$$\ln[G(t)+1] = C + gt + a_1 V(t) + a_2[U_f(t)/U_m(t)] + \mu_{15}(t) \tag{15}$$

where $g = d + m - f$.

Given our estimate of the adjusted growth rate of the male/female earnings ratio (g), we could estimate the percentage change attributable to discrimination if we had some independent estimate of the growth rate in the relative labor quality of males ($m-f$). Following Ashenfelter, we define the following index of labor quality:

$$Q_j(t)=\sum_s P_{sj}(t)Y_s, \qquad j=m,f$$

where $Q_j(t)$ is the labor quality index; P_{sj} is the proportion of workers in the jth sex group with s years of schooling; and Y_s is the earnings attributable to s years of schooling in some base period. Although the labor quality index for men should be constructed using male earnings, the choice of earnings figures for women is not so obvious. If we believe that male-female earnings differences within given schooling categories are solely the result of discrimination, it would be proper to use the male earnings figures to construct the female labor quality index. On the other hand, if we believe that sex differences in earnings within given schooling categories are solely the result of lower quality schooling for women, then it would be proper to use the earnings figures for women in the construction of their labor quality index. One could argue that women specialize in subject matters that raise their productivity in the home rather than in the market sector. Even here, a question arises as to what extent such specialization is a response to anticipated discrimination in the labor market. We thus arrive at one index for men and two alternative indexes for women:

$$Q_m(t)=\sum_s P_{sm}(t)Y_{sm}$$

$$Q'_f(t)=\sum_s P_{sf}(t)Y_{sm}$$

$$Q''_f(t)=\sum_s P_{sf}(t)Y_{sf}$$

Assuming that the labor quality indexes grow exponentially, we can posit the following time-series relationship:

$$\ln[Q_m(t)/Q_f(t)]=C_0+(m-f)t+\mu(t)$$

Given $\hat{g}$ and $(\hat{m}-\hat{f})$, it is then possible to obtain $\hat{d}$.

According to Becker, the market discrimination coefficient can be expected to increase with increases in the relative supply of female workers, the average propensity to discriminate, and the variance of individual discrimination coefficients. Thus, for example, an increase in the female proportion of the labor force would widen the male-female earnings differential, even though there were no changes in the propensity to discriminate. This could occur because the increased proportion of women must partly be absorbed by firms with discrimination coefficients above the former equilibrium market discrimination coefficient. Consequently, women would have to accept a lower relative wage which then raises the equilibrium market discrimination coefficient. Since the nonwhite proportion of the labor force has been fairly constant, Ashenfelter was able to ignore the relative supply factor in his analysis of race earnings differentials [Ashenfelter 1970]. Furthermore, the assumption of a nonshifting, perfectly inelastic nonwhite relative supply curve allowed him to interpret equation 15 as a relative demand curve. Thus, changes in the race earnings differential could be attributed solely to changes in relative demand. Unfortunately, this assumption is not tenable in the case of sex discrimination because the relative supply of females has been rising steadily. Therefore, equation 15 is interpreted as a reduced-form equation. In a structural relative demand equation, we would expect a positive relationship between the male-female earnings differential and the relative supply of women. This implies that our estimate of the change in the earnings differential, attributable to discrimination, $\hat{d}$, measures the combined effects of changes in relative female labor supply and changes in the expected value and variance of individual discrimination coefficients.

The most important source of the overall earnings differential between the sexes is probably their differing occupational distributions, as opposed to unequal pay for equal work. It is currently a matter of debate whether and to what extent the different occupational distributions are the results of labor market discrimination and to what extent different job preferences are responsible. Promising research is being conducted by psychologists studying whether there really are sex and race differences in motivation, aspirations, and expectations about career fulfillment [Gurin 1974] and [Laws 1974]. Such research provides the psychological and sociological backdrops for different occupational attachments and labor-supply elasticities. For the present, we confine ourselves to trends in occupational distributions and short-run cyclical determinants of occupational distributions. The time-series relationship is specified as

$$O_{ij}(t) = b_{0ij} + b_{1ij}t + b_{2ij}V(t) + \mu_{ij}(t)$$

where i denotes the ith occupational category, j denotes male or female, O_{ij} is the percent of the jth group's labor force who are in the ith occupation, t is the time-trend variable, $V(t)$ is the cyclical demand variable (the unemployment rate of white males aged 35–44), and $\mu_{ij}(t)$ is the disturbance term. The changes in male-female differences in occupational distribution can be directly estimated by

$$\Delta O_i(t) = \Delta b_{0i} + \Delta b_{1i}t + \Delta b_{2i}V(t) + \mu_i(t)$$

where

$$\Delta O_i(t) = O_{im}(t) - O_{if}(t),\ \Delta b_{1i} = b_{1im} - b_{1if}, \text{ and } \Delta b_{2i} = b_{2im} - b_{2if}.$$

It is clear that the effects of changes in occupational distributions on the male-female earnings differential depend on male-female earnings differentials within occupations. To see this explicitly, let us consider the following identity:

$$(16)\qquad \ln(Y_m/Y_f) \equiv \sum_i \ln(Y_{im})O_{im} - \sum_i \ln(Y_{if})O_{if}$$

where all earnings are geometric means, and O_{im} and O_{if} are the proportions of the male and all female labor forces in occupation i, respectively. Next, let $\ln(G_i + 1) = \ln(Y_{im}) - \ln(Y_{if})$, where G_i is the male-female earnings differential in the ith occupation. Now substitute for $\ln(Y_{if})$ in (16) the expression given by $\ln(Y_{if}) = \ln(Y_{im}) - \ln(G_i + 1)$. After collecting terms we have

$$(17)\qquad \ln(Y_m/Y_f) = \sum_i [\ln(Y_{im})\Delta O_i + \ln(G_i + 1)O_{if}]$$

Differentiating both sides of (17) with respect to time yields

$$(18)\qquad \frac{\partial \ln(Y_m/Y_f)}{\partial t} = \sum_i \left[\ln(Y_{im})\Delta b_{1i} + \frac{\dot{Y}_{im}}{Y_{im}}\Delta O_i + \ln(G_i + 1)b_{1if} + g_i O_{if}\right]$$

Thus, the contribution of each occupation to changes in the overall male/female earnings ratio can be estimated from equation 18.

Overall occupational dissimilarity between the sexes can be measured by means of an index used by the Council of Economic Advisers in their recent report on the economic role of women [CEA 1973]. For purposes of examining continuous movements in occupational dissimilarity, we compute the value of the index for each year as

$$I(t) = \tfrac{1}{2}\sum_i |\Delta O_i(t)|$$

where $I(t)$ is the index of occupational dissimilarity and $|\Delta O_i(t)|$ is the absolute value of the difference between the percentage of men in the ith occupation at time t and the percentage of women in occupation i at time t. If men and women were equally concentrated across all occupations, the index would equal zero. On the other hand, if each occupation were either all male or all female, the value of the index would equal one hundred. Intermediate values of the index indicate varying degrees of occupational dissimilarity. In order to detect trend and cyclical changes in occupational dissimilarity between the sexes, regressions of the following type are estimated: $I(t) = b_0 + b_1 t + b_2 V(t) + \mu(t)$.

II. EMPIRICAL RESULTS

Cross Section

The data for the cross-section analysis are urban worker subsamples drawn from the 1/1000 Public Use Samples of the 1960 and 1970 censuses. Of the variety of 1/1000 samples available from the 1970 Census, the 15 percent Neighborhood Characteristics sample was used in this study. Our analysis is confined to those who worked 50–52 weeks in the year preceding the Census year, who resided inside urban areas, and who were either government or private wage and salary workers. The gross differential in (geometric) mean earnings for whites rose from 79 percent in 1959 to 84 percent in 1969, while for blacks it declined from 95 percent to 60 percent over the same period.[2]

The earnings regressions corresponding to equations 5 and 6 are reported in Tables A-1 and A-2 in the appendix. We are more directly concerned with the resulting (log) decomposition of the sex earnings differentials into the effects of sex differences in the estimated coefficients and sex differences in the mean values of the independent variables. The former provides the basis for estimating the discrimination coefficient and the latter yields an estimate of labor quality or productivity differentials. These decompositions are reported in sum and also separately by independent variable in Tables 1 and 2. For each year there are two sets of decompositions: one set is based on earnings regressions which control for government employment, occupation, and industry, whereas the other set is based on earnings regressions which omit these variables. By not controlling for these variables, we hope to capture the effects of employment barriers on measured discrimination. Certainly, it is not implausible to argue that employment barriers have a larger impact on the overall earnings disparities between the sexes than do instances of unequal pay for equal work.

In Tables 1 and 2 the sum of the numbers in the columns headed by $\Delta\hat{\beta} \cdot \bar{X}_f$ is an estimate of the differential attributable to discrimination, $\ln(\hat{D}+1)$. The sum of the numbers in the columns headed by $\hat{\beta}_m \cdot \Delta\bar{X}$ is an estimate of the differential attributable to differences in personal characteristics. These sums and their components are expressed as percentages of the gross differential in logs.[3] As expected, the estimated effects of discrimination are always larger when calculated from the regressions that do not control for government, occupation, and industry. This result is more pronounced in 1960 and for blacks. As can be seen from Table 1, the estimated effects of sex discrimination among whites increased from 1960 to 1970 under both decompositions. Furthermore, the gross logarithmic earnings differential increased by less than the

TABLE 1 Decomposition of the Earnings Differential for Whites

	$\Delta\hat{\beta}\cdot\bar{X}_f$	Percent of $\ln(G+1)_{60}$	$\hat{\beta}_m\cdot\Delta\bar{X}$	Percent of $\ln(G+1)_{60}$	$\Delta\hat{\beta}\cdot\bar{X}_f$	Percent of $\ln(G+1)_{60}$	$\hat{\beta}_m\cdot\Delta\bar{X}$	Percent of $\ln(G+1)_{60}$
A. 1960: $\ln(G+1)_{60} = 0.5818$								
Experience	0.0248	4.3	0.0037	0.6	0.0428	7.4	0.0039	0.7
Education	0.1490	25.6	0.0044	0.8	0.0883	15.2	0.0061	1.0
Region	0.0010	0.2	0.0009	0.2	0.0005	0.1	0.0010	0.2
City size	0.0484	8.3	0.0065	1.1	0.0628	10.8	0.0079	1.4
Marital status	−0.0995	−17.1	0.0665	11.4	−0.1066	−18.3	0.0792	13.6
Children	0.0460	7.9	0.0000	0.0	0.0591	10.2	0.0000	0.0
Part-time	0.0137	2.4	0.0286	4.9	0.0270	4.6	0.0316	5.4
Recent move	−0.0091	−1.6	−0.0004	−0.1	−0.0064	−1.1	−0.0003	−0.1
Government	−0.0272	−4.7	−0.0002	−0.0				
Occupation	−0.2374	−40.8	0.0755	13.0				
Industry	−0.0245	−4.2	0.0560	9.6				
Constant	0.4528	77.8	0.0000	0.0	0.2796	48.1	0.0000	0.0
	0.3380	58.1	0.2415	41.5	0.4471	77.0	0.1294	22.2

TABLE 1 (concluded)

	$\Delta\hat{\beta} \cdot \bar{X}_f$	Percent of $\ln(G+1)_{70}$	$\hat{\beta}_m \cdot \Delta\bar{X}$	Percent of $\ln(G+1)_{70}$	$\Delta\hat{\beta} \cdot \bar{X}_f$	Percent of $\ln(G+1)_{70}$	$\hat{\beta}_m \cdot \Delta\bar{X}$	Percent of $\ln(G+1)_{70}$
			B. 1970: $\ln(G+1)_{70} = 0.6131$					
Experience	0.0356	5.8	0.0009	0.1	0.0686	11.2	0.0004	0.1
Education	0.1474	24.0	0.0157	2.6	0.0112	1.8	0.0203	3.3
Region	0.0068	1.1	0.0003	0.0	0.0190	3.1	0.0003	0.0
City size	0.0527	8.6	0.0034	0.6	0.0656	10.7	0.0045	0.7
Marital status	−0.1233	−20.1	0.0663	10.8	−0.1388	−22.6	0.0766	12.5
Children	0.0334	5.4	0.0000	0.0	0.0429	7.0	0.0000	0.0
Part-time	0.0205	3.3	0.0252	4.1	0.0312	5.1	0.0270	4.4
Recent move	−0.0039	−0.6	−0.0004	−0.1	−0.0080	−1.3	−0.0014	−0.2
Government	−0.0399	−6.5	−0.0039	−0.6				
Occupation	−0.2354	−38.4	0.0694	11.3				
Industry	0.0235	3.8	0.0446	7.3				
Constant	0.4741	77.3	0.0000	0.0	0.3943	64.3	0.0000	0.0
	0.3915	63.7	0.2215	36.1	0.4860	79.3	0.1277	20.8

SOURCE: The figures in this table are derived from earnings regressions corresponding to year-round urban workers from the U.S. Census 1/1000 Public Use Samples of 1960 and 1970. The Neighborhood Characteristics 15 percent sample was the particular 1/1000 sample used for the 1970 regressions.

TABLE 2 Decomposition of the Earnings Differential for Blacks

	$\Delta\hat{\beta} \cdot \bar{X}_f$	Percent of $\ln(G+1)_{60}$	$\hat{\beta}_m \cdot \Delta\bar{X}$	Percent of $\ln(G+1)_{60}$	$\Delta\hat{\beta} \cdot \bar{X}_f$	Percent of $\ln(G+1)_{60}$	$\hat{\beta}_m \cdot \Delta\bar{X}$	Percent of $\ln(G+1)_{60}$
A. 1960: $\ln(G+1)_{60} = 0.6684$								
Experience	0.2738	41.0	0.0039	0.6	0.4499	67.3	0.0041	0.6
Education	0.1275	19.1	−0.0126	−1.9	−0.1001	−15.0	−0.0164	−2.5
Region	0.0253	3.8	0.0152	2.3	0.0036	0.5	0.0183	2.7
City size	0.0253	3.8	0.0078	1.2	0.0571	8.5	0.0077	1.2
Marital status	−0.0616	−9.2	0.0208	3.1	−0.0549	−8.2	0.0356	5.3
Children	0.0130	1.9	0.0000	0.0	0.0096	1.4	0.0000	0.0
Part-time	0.0440	6.6	0.0476	7.1	0.0718	10.7	0.0498	7.5
Recent move	−0.0030	−0.4	−0.0006	−0.1	0.0118	1.8	−0.0005	−0.1
Government	−0.0186	−2.8	0.0046	0.7				
Occupation	−0.0676	−10.1	0.0833	12.5				
Industry	0.0488	7.3	0.1074	16.1				
Constant	−0.0132	−2.0	0.0000	0.0	0.1174	17.6	0.0000	0.0
	0.3937	59.0	0.2774	41.6	0.5662	84.6	0.0986	14.7

TABLE 2 (concluded)

	$\Delta\hat{\beta}\cdot\bar{X}_f$	Percent of $\ln(G+1)_{70}$	$\hat{\beta}_m\cdot\Delta\bar{X}$	Percent of $\ln(G+1)_{70}$	$\Delta\hat{\beta}\cdot\bar{X}_f$	Percent of $\ln(G+1)_{70}$	$\hat{\beta}_m\cdot\Delta\bar{X}$	Percent of $\ln(G+1)_{70}$
			B. 1970: $\ln(G+1)_{70} = 0.4699$					
Experience	0.0916	19.5	0.0103	2.2	0.0911	19.4	0.0096	2.0
Education	−0.0917	−19.5	−0.0234	−5.0	−0.4830	−102.8	−0.0311	−6.6
Region	−0.0307	−6.5	0.0027	0.6	−0.0522	−11.1	0.0029	0.6
City size	0.0105	2.2	0.0001	0.0	0.0092	2.0	0.0002	0.0
Marital status	−0.1290	−27.5	0.0326	6.9	−0.1199	−25.5	0.0422	9.0
Children	0.0408	8.7	0.0000	0.0	0.0501	10.7	0.0000	0.0
Part-time	0.0006	0.1	0.0101	2.1	0.0143	3.0	0.0125	2.7
Recent move	−0.0012	−0.3	0.0011	0.2	−0.0009	−0.2	0.0013	0.3
Government	−0.0404	−8.6	−0.0001	0.0				
Occupation	0.0739	15.7	0.0173	3.7				
Industry	−0.0894	−19.0	0.0920	19.6				
Constant	0.4961	105.6	0.0000	0.0	0.9235	196.5	0.0000	0.0
	0.3311	70.4	0.1427	30.3	0.4322	92.0	0.0376	8.0

SOURCE: See Table 1.

increase in measured discrimination. This means that an increase in the relative productivity of white women, as evaluated in terms of the white male earnings regressions in each year, prevented the gross earnings gap from widening more than it actually did. Table 2 indicates that discrimination among blacks diminished over the decade, but by less than the reduction in the gross earnings differential (in logs). This implies that the decrease in the gross earnings differential between black men and black women would not have been as large if it were not for the rise in the relative productivity of black women. The information in Tables 1 and 2 can be used to estimate annual percentage rates of change in the male/female earnings ratio, discrimination, and relative productivity. These calculations have been made and are later presented and compared with time-series estimates.

An important aspect of labor market discrimination against women is the impact of government on sex earnings differentials. Our cross-section analysis sheds some light on the impact of government employment on male/female relative earnings. It is evident from the decomposition of the gross earnings differential that government employment reduces the average earnings of all men relative to the average earnings of all women in the labor market when compared with the male/female earnings ratio in the nongovernment sector. This occurs because the male-female earnings differential is less in government than nongovernment employment and because a larger proportion of female workers than male workers are employed by government. This narrowing of the overall earnings differential increased from 1960 to 1970 for both blacks and whites. The earnings regressions presented in the appendix to this paper reveal that between 1960 and 1970 the earnings advantage of government over nongovernment employment declined for all four race/sex groups. In the case of white men, earnings were actually lower in government as compared with nongovernment employment; furthermore, this discrepancy widened over the decade. This, of course, does not contradict the increased importance over the decade of government employment as a factor tending to narrow the male/female earnings ratio: the decrease in the earnings advantage of government over nongovernment employment was less for women than for men. These findings are consistent with the hypothesis that public sector employment is less discriminatory than employment in the private sector.

Before turning to the time-series results, a few comments are in order concerning the effects of childbearing on the earnings differential between men and women workers. Because the common earnings structure in the absence of discrimination is assumed to be the male earnings structure, the effects of childbearing must necessarily show up entirely as a contribution to the estimated effects of differences in coefficients. These effects have been included in our estimates of the discrimination coefficient.[4] One could argue that the wage effects of childbearing reflect discrimination in terms of access to positions offering on-the-job training and in terms of forced leave or quitting. An equally plausible argument against including the effects of childbearing in our estimates of discrimination is that childbearing reflects voluntary household decisions regarding the allocation of time between market and nonmarket work. If we were to accept this argument, then our estimates of the discrimination coefficient in each year would be lowered; however, the direction of change in the estimated discrimination coefficients between 1960 and 1970 would not be affected. In the case of white women, the increment in measured discrimination would rise because the effects of children on the sex earnings differential declined between 1960 and 1970. In the case of black women, the reduction in measured discrimination would be larger because the effects of children on the sex earnings differential increased over the Census decade.

Time Series

As has been previously stated, the advantage of the time-series analysis is that it may shed some light on the effects of cyclical movements on the male/female earnings ratio. Furthermore, yearly changes in magnitudes such as the sex earnings ratio and occupational distribution are directly estimated by time-series regressions. The time-series data are in the form of published averages and differ somewhat in concept from the Census microdata. Although some effort has been made to achieve comparability between the cross-section and time-series data, the remaining differences should be borne in mind when comparing the empirical findings.

The time-series earnings data are the median earnings of year-round full-time workers. The trends in gross differentials in median earnings were estimated from regressions of the type specified in equation 15. Variations on this specification were also estimated and the results are reported in Table 3. It is evident that the gross male-female earnings differential widened among whites and narrowed among nonwhites over the period 1955–71. The results also seem to suggest that sex earnings differentials move countercyclically as evidenced by the positive coefficients corresponding to the unemployment rate of white males 35–44; however, these coefficients are never statistically significant at the levels adopted in this study. Gross differentials widened with increases in the ratio of female/male unemployment, but this variable is statistically significant only for nonwhites.

TABLE 3 Estimated Reduced-Form Equations for Male-Female Earnings Differentials (1955–71)

Dependent Variable: $\ln(Y_m/Y_f)$[a]	Whites				Nonwhites			
Constant	0.3121†	0.3106†	0.3329†	0.4041†	0.1569	0.1701	0.1927	0.1947
	(0.0809)	(0.0735)	(0.0694)	(0.0516)	(0.1953)	(0.1823)	(0.1829)	(0.1857)
D(1966–71)		−0.0017				−0.0044		−0.0076
		(0.0009)				(0.0023)		(0.0088)
Time	0.0046†	0.0064†	0.0070†	0.0147†	−0.0207†	−0.0160†	−0.0153†	
	(0.0011)	(0.0013)	(0.0013)	(0.0021)	(0.0037)	(0.0044)	(0.0046)	
D · time			-0.132×10^{-3}*				-0.286×10^{-3}	
			(0.054×10^{-3})				(0.164×10^{-3})	
Time squared				-0.524×10^{-3}†				-0.696×10^{-3}
				(0.106×10^{-3})				(0.430×10^{-3})
Unemployment rate of white males 35–44[b]	0.0186	0.0154	0.0134	0.0035	0.0512	0.0410	0.0398	0.0388
	(0.0097)	(0.0090)	(0.0085)	(0.0065)	(0.0277)	(0.0265)	(0.0266)	(0.0273)
Ratio of female/male unemployment rates[b]	0.0828	0.0848	0.0683	0.0151	0.3146*	0.3084*	0.2845*	0.2694
	(0.0481)	(0.0437)	(0.0414)	(0.0317)	(0.1349)	(0.1258)	(0.1266)	(0.1302)
$\bar{R}^2$	0.80	0.84	0.86	0.93	0.77	0.80	0.80	0.79
Durbin-Watson	1.30	2.00	2.06	2.21	1.27	1.84	1.72	1.44

NOTE: Standard errors appear in parentheses.
*Significant at the 0.05 level in a two-tailed test.
†Significant at the 0.01 level in a two-tailed test.
[a]The median wage and salary earnings correspond to year-round full-time workers, i.e., those who work 50–52 weeks per year and 35 or more hours per week. These data were taken from various issues of the *Current Population Reports*, Series P-60, Consumer Income, U.S. Bureau of the Census.
[b]The unemployment statistics refer to the civilian noninstitutional population 16 years and older. These data were taken from the *Manpower Report of the President*, March 1973, pp. 148–149.

It is reasonable to suppose that sufficient time has elapsed since the passage of the 1963 Federal Equal Pay Act and the passage of the 1964 Civil Rights Act to detect any significant impact such legislation may have had on sex differentials in earnings. Are earnings differentials by sex any smaller than they would have been had there been no such legislation? Most likely, there would be no immediate impact of such legislation due to normal lags in the effective implementation of the laws.

To test for differences in the sex earnings ratios in the period following the federal legislation, a dummy variable representing the period 1966–71 was entered in the reduced-form equations for male-female earnings differentials. For both whites and nonwhites, the estimated coefficient was negative but not statistically significant. As an alternative specification, the dummy variable for the period 1966–71 was interacted with the time-trend variable. The estimated coefficient on the interaction term was negative for both whites and nonwhites but statistically significant only for whites; nevertheless, the negative effect was extremely small, indicating that after 1966 the white male/white female earnings ratio grew at a rate 0.01 percent per annum less than the pre-1966 rate. To pursue the possibility that the rate of change of the male/female earnings ratio was not constant over the perod 1955–71, a time-squared term was entered in the reduced-form equation.[5] Both the linear and quadratic terms of the time variable are statistically significant only for whites, and the corresponding coefficients imply that the log of the white male/white female earnings ratio attained a maximum in 1968 and began declining thereafter. In other words, the growth rate of the earnings ratio was positive but declining prior to 1968, and was negative after this period.

In order to estimate the rate of change in the sex earnings ratio attributable to changes in discrimination, we require independent estimates of changes in relative productivity. These latter estimates were obtained from equations in which the log of the ratio of male/female labor quality indexes was regressed on a time trend. In the case of whites, a quadratic specification of the time variable was also estimated. These regressions are presented in Table 4. The estimated annual percentage change in the male/female earnings ratio that can be attributed to changes in labor market discrimination ($\hat{d}$) is calculated as the difference between the gross annual percentage change in the earnings ratio ($\hat{g}$) and the annual percentage change in relative labor quality ($\hat{m} - \hat{f}$). The estimates of these parameters are first calculated from the regressions that include just the linear time-trend variable. These estimates are presented in Table 5 along with their counterparts derived from the cross-section results given in Tables 1 and 2. Despite the differences in data sources and methods, there is a broad consistency between the cross-section and

TABLE 4 Estimated Growth Rates of Relative Labor Quality (1952–71)

Group	Dependent Variable[a]	Constant	Time	Time Squared	$\bar{R}^2$
Whites	$\ln (Q_{wm}/Q'_{wf})$	−0.0314* (0.0029)	0.0021* (0.0002)		0.92
		−0.0382* (0.0019)	0.3949×10^{-2}* (0.0352×10^{-2})	-0.0085×10^{-2}* (0.0016×10^{-2})	0.98
	$\ln (Q_{wm}/Q''_{wf})$	0.5805* (0.0029)	0.0020* (0.0002)		0.91
		0.5738* (0.0020)	0.3779×10^{-2}* (0.0371×10^{-2})	-0.0083×10^{-2}* (0.0017×10^{-2})	0.98
Nonwhites	$\ln (Q_{nm}/Q_{nf})$	−0.0524* (0.0060)	0.0011* (0.0004)		0.40

NOTE: Standard errors appear in parentheses.
*Significant at the 0.01 level in a two-tailed test.
[a]The dependent variables are the natural logarithms of the ratio of the male labor quality index to the female labor quality index. The quality index for a given group in a given year is calculated by summing over the white male earnings in a given schooling category in 1959 weighted by the proportion of the group who fall in the particular schooling category in the given year. For white females the indexes Q'_{wt} and Q''_{wf} are calculated using white male and white female earnings, respectively. The earnings data correspond to individuals who worked 40 or more weeks in 1959 and were obtained from a special tabulation of the 1/1000 sample from the 1960 Census. These tabulations were kindly made available by Orley Ashenfelter. The schooling categories are as follows: less than 5 years, 5–8 years, 9–11 years, 12 years, 13–15 years, and 16 years or more. The data corresponding to the proportion of workers in these categories are reported for 1952, 1959, 1962, and 1964–71. They are available from the *Manpower Report of the President*, March 1973, Table B9, p. 177. Because of discontinuities in the data, Durbin-Watson statistics are not reported.

time-series results. Both show an annual percentage increase in the male/female earnings ratio among whites and a decrease among nonwhites. The estimated rate varies from 0.3 percent to 0.5 percent a year for whites and from −2.0 percent to −2.1 percent a year for nonwhites. Large differences in estimates occur with respect to $\hat{d}$ and $\hat{m} - \hat{f}$. This is to be expected in view of the crude nature of the labor quality index calculated from the time-series data available. Educational distribution alone is an extremely narrow basis on which to gauge changes in relative labor quality. In this regard, the Census microdata are clearly superior, because they provide detailed information on several different components of overall labor quality. Because the educational distribution of male workers has been improving over the years vis-à-vis female workers, it is not surprising that the time-series construction shows a rise in male relative to female labor quality. The tendency is to bias downward the estimated rate of change in the sex earnings ratio attributable to discrimination. In spite of this, discrimination is estimated to have increased among whites and to have fallen among nonwhites.[6] The cross-section estimates reveal a decrease in male/female relative labor

TABLE 5 Derived Rates of Change from Cross Sections and Time Series

	g	d		m − f	
Whites					
Cross section	0.0031[a]	0.0053[b]	0.0036[c]	−0.0022[b]	−0.0005[c]
Time series	0.0046[d]	0.0025[e]	0.0026[f]	0.0021[e]	0.0020[f]
Nonwhites					
Cross section	−0.0198[a]	−0.0063[b]	−0.0136[c]	−0.0135[b]	−0.0062[c]
Time series	−0.0207[d]	−0.0218[e]		0.0011[e]	

[a] $g = [\ln (G+1)_{70} - \ln (G+1)_{60}]/10.$

[b] $d = [\ln (\hat{D}+1)_{70} - \ln (\hat{D}+1)_{60}]/10$

and

$m - f = g - d$

where

$\ln (\hat{D}+1) = \sum_k \Delta\hat{\beta}_k \bar{X}_{fk} + [\ln (G+1) - \ln (\hat{G}+1)]/2$

These calculations are based on the cross-section Census earnings regressions that control for government, occupation, and industry.

[c] d and $m-f$ are derived as in footnote b except that the underlying regressions do not control for government, occupation, and industry.

[d] $g = \partial \ln [G+1]/\partial t$ estimated from the time-series male/female relative earnings regressions in which t enters in linear form.

[e] $d = g - (m-f)$; where $m - f = \partial \ln (Q_m/Q_f)/\partial t$ estimated from the time-series relative labor quality regressions which used white male earnings as weights and in which time enters in linear form.

[f] d and $m-f$ are derived as in footnote e except that the underlying relative labor quality regressions used white female earnings as weights.

quality for both whites and nonwhites. They also reveal an increase in discrimination among whites larger than the time-series estimates, and they show a decrease in discrimination among nonwhites smaller than the time-series estimates.[7]

Returning to the question of the impact of legislation on the male/female earnings ratio, we note that the quadratic specification of the time variable in the relative earnings and relative labor quality equations permit the treatment of $\hat{g}$ and $\hat{d}$ as functions of time. Let

$$g(t) = \partial \ln [Y_m/Y_f]/\partial t = a_1 + 2a_2 t$$
$$m(t) - f(t) = \partial \ln [Q_m/Q_f]/\partial t = b_1 + 2b_2 t'$$
$$d(t) = g(t) - [m(t) - f(t)] = a_1 - b_1 + 2(a_2 t - b_2 t')$$

where a_1, b_1, $a_1 - b_1 > 0$ and a_2, b_2, $a_2 - b_2 < 0$. Since the relative earnings regressions are estimated over the period 1955–71 and the relative labor quality regressions extend over the period 1952–71, we have the relationship $t' = t + 3$. Substituting for t' we have $d(t) = a_1 - b_1 - 6b_2 + 2(a_2 - b_2)t$. To calculate the year in which the sex earnings ratio attains a maximum as a result of discrimination, set $d(t)$ equal to zero and solve for t: $\hat{t} = (a_1 - b_1 - 6b_2)/2(b_2 - a_2)$. These computations were performed for whites. Using the translation $T = 1954 + t$, the results suggest that discrimination began to diminish in 1966.[8]

The evidence seems to suggest a small but discernible reversal in the trend toward increases in the white-male–white-female earnings differential at about the same time we would have expected government policy to start taking effect. The problems with interpreting these results as indicative of the effects of government policy are formally the same as those problems encountered in the Phillips curve literature on the effectiveness of wage-price guideposts. We may only be observing the net effect of many factors which are not otherwise accounted for by the regression in the particular period under study. Yet as slim as the evidence is that government policy was responsible for damping or reversing the growth in the sex earnings differential for whites, the results do point to a small reduction in discrimination for whatever reasons.

Turning now to male-female occupational changes over time, Table 6 presents the estimated effects of trend and cyclical factors on male-female differences in occupational distribution for eight occupational categories.[9] First examining the results for whites, we find that statistically significant trends in male-female differences in occupational concentration occurred in favor of men in the following occupations: professional & technical, and managerial, officials, and proprietors. On the other hand, statistically significant trends in white male-female differences in occupational concentration occurred in favor of greater relative female concentration in the following occupations: clerical & sales, and farming. Cyclical factors exerted no statistically significant independent influence on male-female differences in occupational concentration except for the laborers category. In this occupation, loose labor markets lead to increases in male concentration over female concentration. In the case of nonwhites, we find statistically significant trends in the concentration of male employment over the concentration of female employment among the following occupations: managerial, officials, & proprietors; craftsmen; and service & private household workers. Statistically significant trends in nonwhite male-female occupational concentration differences indicate a trend toward greater relative representation of nonwhite females among clerical & sales workers and laborers. In periods of loose labor markets, male concentration increases over female concentration among professional and technical workers and decreases among craftsmen. The estimated coefficients of the unemployment rate of white males 35–44 were not statistically significant in the remaining occupations.

TABLE 6 Estimated Equations for Changes in Sex Differences in Occupational Distribution (1958–71) (Dependent variable: Percent of total male employment in a given occupation minus percent of total female employment in the occupation[a])

	Whites					Nonwhites				
Occupation	Constant	Time	Unemployment Rate of White Males 35–44[b]	$\bar{R}^2$	Durbin-Watson	Constant	Time	Unemployment Rate of White Males 35–44[b]	$\bar{R}^2$	Durbin-Watson
Professional & technical	−0.9790 (0.4638)	0.0764* (0.0253)	−0.2085 (0.1133)	0.78	1.66	−3.1990† (0.4662)	−0.0252 (0.0254)	0.2954† (0.1139)	0.65	1.75
Managerial, officials, & proprietors	8.6062† (0.4542)	0.1125† (0.0248)	0.0593 (0.1109)	0.74	1.86	0.0102 (0.4126)	0.1821† (0.0225)	0.1821 (0.1007)	0.89	1.15
Clerical & sales	−28.8499† (1.0829)	−0.2229† (0.0590)	0.5459 (0.2645)	0.84	1.89	2.8764 (1.8642)	−1.1711† (0.1016)	−0.5570 (0.4553)	0.95	1.04
Craftsmen	19.4761† (0.5529)	0.0359 (0.0301)	−0.2116 (0.1350)	0.50	1.42	10.0941† (0.7328)	0.2974† (0.0399)	−0.5931† (0.1790)	0.95	2.04
Operative	3.8552† (0.5996)	0.0445 (0.0327)	0.0231 (0.1464)	0.10	1.77	10.7835† (1.2716)	0.0514 (0.0693)	−0.2101 (0.3106)	0.13	1.27
Laborers	5.2462† (0.4826)	−0.0127 (0.0263)	0.2753* (0.1179)	0.53	1.15	24.5461† (1.4435)	−0.6087† (0.0787)	−0.0062 (0.3526)	0.91	1.30
Service & private household	−12.9273† (1.2967)	0.1087 (0.0707)	−0.3981 (0.3167)	0.51	2.15	−50.6768† (3.3311)	1.3584† (0.1816)	1.1797 (0.8136)	0.88	0.68
Farming	6.0177† (1.2718)	−0.1635* (0.0693)	−0.1670 (0.3106)	0.35	0.95	5.5264 (2.5749)	−0.0772 (0.1404)	−0.3044 (0.6289)	−0.15	1.08

NOTE: Standard errors appear in parentheses.
*Significant at the 0.05 level in a two-tailed test.
†Significant at the 0.01 level in a two-tailed test.
[a]Data on occupational distribution were annual averages and were taken from *Current Population Reports* Series, P-50, Labor Force Characteristics, U.S. Bureau of the Census, for 1958–59, and from *Employment and Earnings and Monthly Report on the Labor Force*, Bureau of Labor Statistics, for 1960–71.
[b]Refer to footnote b in Table 3.

As has been discussed earlier, we can directly measure overall occupational similarity/dissimilarity through the use of a single index. The regressions relating this index to the time trend and cyclical variables are reported in Table 7. As the results show, there was no movement in

TABLE 7 Estimated Equations for Occupational Dissimilarity (1958–71) (dependent variable: index of occupational dissimilarity[a])

Group	Constant	Time	Unemployment Rate of White Males 35–44[b]	$\bar{R}^2$	Durbin-Watson
Whites	43.0401*	0.0279	0.132×10^{-3}	−0.04	1.88
	(0.6261)	(0.0341)	(0.1529)		
Nonwhites	51.5426*	−0.1751	−1.0637	0.12	1.23
	(2.2250)	(0.1213)	(0.5434)		

NOTE: Standard errors appear in parentheses.
*Significant at the 0.01 level in a two-tailed test.
[a]The index of occupational standing is calculated as one-half the sum of the absolute values of the difference between the percentages of males and females in each occupation. For the source of the occupational data used in the construction of the index refer to footnote a in Table 6.
[b]Refer to footnote b in Table 3.

the index for either whites or nonwhites over the period 1958–71. Since broad occupational categories were used, these findings do not rule out the possibility of changes in occupational dissimilarity occurring among more disaggregated classifications. The CEA index based on more disaggregated data was calculated as 62.9 in 1960 and 59.8 in 1970 for whites and nonwhites combined [CEA 1973]. This slight movement toward occupational similarity represents a reduction of 3.1 percentage points over a ten-year period, or an average reduction of only 0.3 percentage points a year. Thus, the use of more detailed data does not alter the conclusion that little progress has been made in the area of sex differences in occupational distribution.

It has been shown earlier that the effects of changes in occupational distributions on the overall male-female earnings differential depend on the earnings differentials within occupations. Unfortunately, yearly occupational earnings data are not available for each race/sex group separately. Such data are, however, available in Census years. Consequently, the data from our urban worker subsamples are used to calculate occupational effects on changes in relative earnings in accordance with equation 18. These estimated effects are shown in Table 8. Changes

TABLE 8 Effects of Occupational Developments on the Annual Rate of Change in the Overall Male/Female Earnings Ratio[a]

Group	Professional and Technical	Managers, Officials, and Proprietors	Clerical and Sales	Craftsmen	Operatives	Laborers	Services and Private Household	Farming[b]	$\sum_i \left[\frac{\partial \ln (Y_m/Y_f)}{\partial t}\right]_i$
Whites	0.0076	0.0138	−0.0276	0.0105	0.0053	0.0007	0.0028	−0.0099	0.0032
Nonwhites	−0.0031	0.0156	−0.0932	0.0292	0.0086	−0.0394	0.0655	−0.0045	−0.0213

[a]The effect of the ith occupation on the annual percentage change in the overall sex earning ratio is calculated from equation 18 as

$$\left[\frac{\partial \ln (Y_m/Y_f)}{\partial t}\right]_i = \ln (Y_{im}^{59})\Delta\hat{b}_{1i} + \frac{\dot{Y}_{im}}{Y_{im}}\Delta\hat{O}_i^{59} + \ln (G_i^{59}+1)\hat{b}_{1if} + g_i\hat{O}_{if}^{59}$$

$\Delta\hat{b}_{1i}$, $\hat{b}_{1if}$, $\Delta\hat{O}_i^{59}$, and $\hat{O}_{if}^{59}$ are derived from the occupational regressions in Tables 6 and A-3. $(\dot{Y}_{im}/Y_{im}) = (\ln Y_{im}^{69} - \ln Y_{im}^{59})/10$, and $g_i = [\ln (G_i^{69}+1) - \ln (G_i^{59}+1)]/10$. The occupational earnings data are from the urban worker subsamples of the 1960 and 1970 1/1000 Census Public Use Samples.

[b]Since the urban worker subsample in 1959 contained only one white woman and no black women in farming, the geometric mean earnings of female laborers in 1959 were used in lieu of farm earnings to calculate the partial effect of farmings.

within the clerical & sales occupation tended to substantially narrow the overall sex earnings differential for whites, while the managerial and craftsmen occupations contributed substantially to a widening of the differential. In the case of nonwhites, the clerical & sales and laborer occupations contributed significantly to a narrowing of the differential. The managerial, craftsmen, and services & private household occupations contributed to a substantial widening of the earnings gap among nonwhites. Because of the hybrid nature of these estimates, their sums do not exactly equal either the cross-section or time-series annual rate of growth in the male/female earnings ratio.

III. CONCLUDING REMARKS

From the middle of the 1950s to the beginning of the 1970s the male-female earnings differential for year-round workers increased among whites and decreased among blacks. Both the cross-section and time-series analyses point to an increase in sex discrimination among whites and a reduction in sex discrimination among blacks as the factors responsible for the trends observed in their respective gross earnings differentials. During this period there occurred little or no movement toward occupational similarity between men and women in either racial group.

Government has been shown to have some narrowing effect on earnings differentials. Government employment has had the effect of reducing the economy-wide male-female earnings differential below the private sector differential. That is to say, the male-female earnings differential would be larger in the absence of public-sector employment. There is also some suggestion that legislation may also have had an effect on the sex earnings differential. This is especially true for whites. Over the entire period of 1955 through 1971 there was, on the average, an increase in the earnings differential due to discrimination; however, the rate of increase was not constant. Starting around 1966, the adjusted growth rate of the male-female earnings differential among whites diminished or perhaps even became negative due to reductions in discrimination. Government policy stemming from the 1963 Federal Equal Pay Act and the 1964 Civil Rights Act may have been responsible for the downturn in the steady growth in the earnings differential due to discrimination. Although the adjusted white-male/white-female earnings ratios were smaller after 1966 than what would have been predicted on the basis of pre-1966 relationships, the effects were fairly small.

We have seen that sex earnings differentials rise during periods of loose labor markets and fall during periods of tight labor markets. This supports the expectation that efforts toward greater equal employment opportunity will enjoy more success during periods of relative prosperity. At such times, there is a maximum of flexibility for making needed changes in employment practices where women and minorities are concerned. In this regard, it is interesting to note that the unemployment rate of white males aged 35–44 averaged 1.9 percent over the period 1966–71 as compared with 3.0 percent over the period 1955–65. Any efforts toward promoting equal employment opportunities during 1966–71 surely benefited from coincidence with full employment.

There are certain nagging problems with the residual approach to measuring discrimination. Space limitations permit only brief mention of these problems, but they do merit careful attention by students of the economics of discrimination. We have assumed that a common earnings structure would exist in the absence of discrimination. It can be argued, however, that specialization within households could lead to different earnings structures for men and women even in the absence of discrimination. This has implications for sex differences in the acquisition of on-the-job training and occupational choice. It is extremely difficult to give any precise information about the allocation of time within households that would exist if there were no sex discrimination in the market sector.

Clearly, the legislative intent of the 1964 Civil Rights Act was to go beyond the narrow definition of discrimination inherent in the 1963 Equal Pay Act, which prohibited unequal pay for equal work. It is debatable whether the wider view of discrimination is adequately reflected by earnings regressions that control for broad occupational categories. One could argue that adjusted earnings differentials within broad occupational groupings mainly reflect discriminatory employment practices, whereas, sex differences in broad occupational affiliation mainly reflect voluntary labor supply decisions explainable in terms of a household life-cycle maximization model. It is the author's contention that the traditional occupational choices of women have been in large part conditioned by the rational expectation of labor market discrimination. Therefore, the estimates of discrimination obtained from the inclusion and exclusion of occupation, industry, and government employment are best viewed for policy purposes as lower and upper bounds to measured discrimination.

Although the residual approach to measuring discrimination avoids the difficult questions regarding the interactions of social and political institutions in determining sex roles, the approach does possess a certain operational practicality. The objective is to arrive at a set of regression control variables that reasonably reflect legislative intent with regard to implied definitions of unlawful discrimination. The residual approach allows us to measure discrimination in accordance with legislative intent and to gauge progress in terms of this standard. It is obvious that there has been a growing impatience over the race and sex earnings differentials that have long persisted in our society. While there is a genuine scholarly interest in researching the formation of these roles, public policy toward discrimination has not awaited, and should not await, the definitive treatment of the subect.

APPENDIX

Background regressions which are used in the construction of Tables 1, 2, and 8 in the main text are reported in this appendix. The cross-section earnings regressions are reported in Tables A-1 and A-2, and the estimated equations for changes in occupational distribution are given in Table A-3.

A brief explanation will be given of the procedures used in connection with the cross-section earnings regressions. As is stated in the text, the cross-section data are drawn from the 1960 and 1970 Census 1/1000 Public Use Samples. In 1960, there was only one type of Public Use Sample available, but in 1970 there were six subcategories to choose from. For this study, the 1970 Neighborhood Characteristics 15 percent sample was used. Our data refer to those who worked 50–52 weeks in the year prior to the year in which the Census was conducted, who also resided inside urban areas, and who also were either government or private wage and salary workers.

Those familiar with Census data know that the Bureau of the Census establishes cutoff levels beyond which household and individual incomes are not reported. The cutoffs and their associated open intervals were $25,000 in 1960 and $50,000 in 1970. The problem confronting the researcher is what income should be assigned to individuals whose incomes are reported to lie somewhere in these open-ended intervals. A commonly accepted practice, and one which is followed in this study, is to estimate the upper tail of the income distribution as a Pareto distribution. From the Pareto distribution we can estimate the mean income of the above $25,000 or $50,000 class.

Assume that the Pareto density function describes the distribution of income above some level of income Y_0:

$$n(y) = \alpha Y_0^{\alpha} y^{-(1+\alpha)} \qquad \text{for } Y_0 < y < \infty$$

$$n(y) = 0 \qquad \text{for } y \leqq y_0$$

where $\alpha > 1$.

TABLE A-1 Census Earnings Regressions for Whites

Dependent Variable: ln (earnings)	1960				1970			
	Male		Female		Male		Female	
	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2
EXP	0.0224†	0.0241†	0.0224†	0.0253†	0.0268†	0.0310†	0.0256†	0.0284†
	(0.0010)	(0.0011)	(0.0015)	(0.0016)	(0.0010)	(0.0010)	(0.0016)	(0.0017)
$EXP^2/10$	−0.0031†	−0.0033†	−0.0034†	−0.0042†	−0.0038†	−0.0043†	−0.0039†	−0.0044†
	(0.0002)	(0.0002)	(0.0003)	(0.0003)	(0.0002)	(0.0002)	(0.0003)	(0.0003)
EDUC	0.0390†	0.0537†	0.0258†	0.0459†	0.0564†	0.0732†	0.0438†	0.0722†
	(0.0015)	(0.0014)	(0.0028)	(0.0027)	(0.0018)	(0.0015)	(0.0033)	(0.0029)
NEAST	0.0201	0.0109	0.0383*	0.0358	0.0318†	0.0435†	0.0542†	0.0583†
	(0.0114)	(0.0119)	(0.0180)	(0.0194)	(0.0121)	(0.0124)	(0.0193)	(0.0201)
NCENT	0.0862†	0.0848†	0.0546†	0.0409*	0.0676†	0.0841†	0.0186	0.0028
	(0.0116)	(0.0121)	(0.0186)	(0.0201)	(0.0122)	(0.0124)	(0.0195)	(0.0204)
WEST	0.0773†	0.0605†	0.0844†	0.0753†	0.0227	0.0096	0.0190	−0.0011
	(0.0129)	(0.0136)	(0.0204)	(0.0222)	(0.0129)	(0.0134)	(0.0210)	(0.0220)
U499	−0.0599†	−0.0614†	−0.1292†	−0.1561†	−0.0936†	−0.0964†	−0.1447†	−0.1572†
	(0.0109)	(0.0115)	(0.0164)	(0.0179)	(0.0127)	(0.0132)	(0.0195)	(0.0205)
UR499	−0.0232	−0.0162	−0.1056†	−0.1337†	−0.0378*	−0.0393*	−0.1370†	−0.1624†
	(0.0145)	(0.0153)	(0.0247)	(0.0268)	(0.0151)	(0.0156)	(0.0253)	(0.0265)
UR500	0.0888†	0.1083†	−0.0088	−0.0118	0.0570†	0.0786†	−0.0336*	−0.0360*
	(0.0093)	(0.0098)	(0.0151)	(0.0164)	(0.0102)	(0.0106)	(0.0161)	(0.0169)
SINGLE	−0.2900†	−0.3385†	−0.0189	−0.0385*	−0.4050†	−0.4722†	−0.0215	−0.0329
	(0.0132)	(0.0138)	(0.0169)	(0.0184)	(0.0139)	(0.0143)	(0.0191)	(0.0201)
DS	−0.0935†	−0.1377†	0.0146	−0.0271	−0.1360†	−0.1646†	0.0238	0.0177
	(0.0226)	(0.0238)	(0.0207)	(0.0226)	(0.0208)	(0.0216)	(0.0211)	(0.0222)
WID	−0.1768†	−0.1999†	−0.0145	−0.0440	−0.1493†	−0.1596†	−0.0017	−0.0150
	(0.0331)	(0.0350)	(0.0226)	(0.0247)	(0.0406)	(0.0422)	(0.0254)	(0.0267)

TABLE A-1 (continued)

Dependent Variable: ln (earnings)	1960				1970			
	Male		Female		Male		Female	
	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2
CHILD			−0.0415†	−0.0533†			−0.0235†	−0.0302†
			(0.0049)	(0.0053)			(0.0049)	(0.0051)
PARTIME	−0.2741†	−0.3024†	−0.3494†	−0.4509†	−0.2575†	−0.2762†	−0.3781†	−0.4600†
	(0.0148)	(0.0156)	(0.0164)	(0.0176)	(0.0166)	(0.0171)	(0.0184)	(0.0190)
MOVE	0.0124	−0.0103	0.0519†	0.0349	−0.0051	−0.0209	0.0218	0.0339
	(0.0108)	(0.0113)	(0.0183)	(0.0200)	(0.0110)	(0.0112)	(0.0198)	(0.0208)
GOVT	−0.0360		0.1793†		−0.1084†		0.1470†	
	(0.0190)		(0.0280)		(0.0157)		(0.0255)	
PROF	0.1158†		0.3669†		0.1355†		0.3960†	
	(0.0188)		(0.0369)		(0.0192)		(0.0372)	
MANAGE	0.2994†		0.4047†		0.2480†		0.4286†	
	(0.0179)		(0.0396)		(0.0186)		(0.0428)	
CLER	−0.1199†		0.2021†		−0.1102†		0.1937†	
	(0.0185)		(0.0295)		(0.0202)		(0.0308)	
CRAFT	−0.0146		0.1821†		0.0090		0.2910†	
	(0.0173)		(0.0550)		(0.0181)		(0.0546)	
OPER	−0.1327†		0.0689*		−0.0778†		0.1217†	
	(0.0175)		(0.0346)		(0.0190)		(0.0386)	
PRIV	−0.9112†		−0.6654†		−0.9597†		−0.5995†	
	(0.2291)		(0.0624)		(0.2305)		(0.0862)	
SERV	−0.2031†		−0.0622		−0.1372†		−0.0540	
	(0.0226)		(0.0348)		(0.0231)		(0.0356)	
LABOR	−0.1963†		−0.2058		−0.2356†		−0.0124	
	(0.0253)		(0.1095)		(0.0258)		(0.0746)	

FARM	−0.4458†	−0.7368	−0.4672†	−0.3817
	(0.1197)	(0.4867)	(0.1341)	(0.2713)
NOCC	−0.0179	0.1126		
	(0.0346)	(0.0604)		
AG	−0.0081	0.2878	0.1569	−0.0255
	(0.0853)	(0.1631)	(0.0937)	(0.1778)
MINE	0.1676†	0.4882†	0.2518†	0.4337†
	(0.0596)	(0.1287)	(0.0538)	(0.1215)
CON	0.2240†	0.0692	0.2997†	0.3302†
	(0.0218)	(0.0576)	(0.0218)	(0.0627)
DURMAN	0.2396†	0.2912†	0.2722†	0.2659†
	(0.0153)	(0.0268)	(0.0160)	(0.0290)
NONDUR	0.2175†	0.2060†	0.2131†	0.1713†
	(0.0163)	(0.0266)	(0.0180)	(0.0301)
TRANS	0.2094†	0.2824†	0.2699†	0.2894†
	(0.0173)	(0.0304)	(0.0184)	(0.0336)
WTRADE	0.1837†	0.2116†	0.2222†	0.2146†
	(0.0202)	(0.0365)	(0.0203)	(0.0380)
FINANCE	0.1113†	0.1419†	0.2114†	0.2021†
	(0.0211)	(0.0269)	(0.0221)	(0.0275)
BUSREP	0.0941†	0.1317†	0.1199†	0.1607†
	(0.0263)	(0.0445)	(0.0261)	(0.0430)
PERSER	−0.0696*	0.0080	−0.0699	−0.0853*
	(0.0334)	(0.0401)	(0.0385)	(0.0414)
REC	0.0799	0.0630	0.1010*	0.0298
	(0.0502)	(0.0709)	(0.0502)	(0.0822)
PROFSER	−0.0423	0.0010	0.1168†	0.0709†
	(0.0225)	(0.0270)	(0.0208)	(0.0258)
PUBADM	0.1860†	0.1118†	0.3448†	0.2049†
	(0.0266)	(0.0422)	(0.0238)	(0.0421)

TABLE A-1 (concluded)

Dependent Variable:	1960				1970			
	Male		Female		Male		Female	
ln (earnings)	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2
NOIND	0.1127† (0.0376)		0.1877† (0.0610)					
CONSTANT	7.7381	7.6938	7.2853	7.4142	7.8670	7.8042	7.3929	7.4099
S.E.E.	0.4522	0.4786	0.4576	0.5016	0.5556	0.5775	0.6098	0.6422
R^2	0.32	0.24	0.37	0.24	0.33	0.27	0.26	0.18
NOBS	13,996	13,996	5,650	5,650	17,746	17,746	8,293	8,293

SOURCE: U.S. Census 1/1000 Public Use Samples of 1960 and 1970. The Neighborhood Characteristics, 15 percent sample was the particular sample 1/1000 sample used for the 1970 regressions.

NOTE: Standard errors appear in parentheses. The difference between Equation 1 and Equation 2 is that Equation 2 does not control for government employment, occupation, or industry.

*Significant at the 0.05 level in a two-tailed test.

†Significant at the 0.01 level in a two-tailed test.

TABLE A-2 Census Earnings Regressions for Blacks

Dependent Variable:	1960				1970			
	Male		Female		Male		Female	
ln (earnings)	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2
EXP	0.0260†	0.0297†	0.0106	0.0021	0.0325†	0.0325†	0.0271†	0.0288†
	(0.0036)	(0.0037)	(0.0059)	(0.0062)	(0.0040)	(0.0040)	(0.0054)	(0.0055)
$EXP^2/10$	−0.0038†	−0.0044†	−0.0025*	−0.0016	−0.0047†	−0.0048†	−0.0043†	−0.0049†
	(0.0006)	(0.0006)	(0.0010)	(0.0011)	(0.0007)	(0.0007)	(0.0010)	(0.0010)
EDUC	0.0209†	0.0271†	0.0074	0.0376†	0.0320†	0.0425†	0.0405†	0.0871†
	(0.0043)	(0.0041)	(0.0083)	(0.0084)	(0.0064)	(0.0059)	(0.0096)	(0.0089)
NEAST	0.1670†	0.1959†	0.1835†	0.2673†	0.0553	0.0610	0.2107†	0.2618†
	(0.0357)	(0.0367)	(0.0604)	(0.0631)	(0.0435)	(0.0440)	(0.0541)	(0.0557)
NCENT	0.2418†	0.2923†	0.1403*	0.2095†	0.2120†	0.2321†	0.1506†	0.1923†
	(0.0352)	(0.0357)	(0.0597)	(0.0636)	(0.0431)	(0.0426)	(0.0548)	(0.0566)
WEST	0.2731†	0.3162†	0.1754	0.2580†	0.1613†	0.1592†	0.1898*	0.2241†
	(0.0533)	(0.0550)	(0.0927)	(0.0989)	(0.0595)	(0.0603)	(0.0751)	(0.0789)
U499	−0.1152†	−0.1300†	−0.2470†	−0.3794†	−0.1292†	−0.1384†	−0.0956	−0.1168*
	(0.0349)	(0.0361)	(0.0585)	(0.0624)	(0.0452)	(0.0458)	(0.0570)	(0.0596)
UR499	−0.3979†	−0.3906†	−0.2810*	−0.3089*	0.0582	0.0068	−0.1639	−0.1295
	(0.1017)	(0.1050)	(0.1267)	(0.1362)	(0.0910)	(0.0904)	(0.1132)	(0.1191)
UR500	0.0191	−0.0030	0.0129	−0.0621	−0.0036	0.0168	−0.0697	−0.0456
	(0.0440)	(0.0456)	(0.0795)	(0.0856)	(0.0477)	(0.0484)	(0.0613)	(0.0645)
SINGLE	−0.2183†	−0.2500†	−0.0482	−0.1092	−0.3325†	−0.3824†	0.0315	0.0031
	(0.0430)	(0.0447)	(0.0719)	(0.0776)	(0.0491)	(0.0494)	(0.0612)	(0.0645)
DS	−0.0117	−0.0478	0.0332	0.0024	−0.0670	−0.1059	0.0917	0.0332
	(0.0465)	(0.0482)	(0.0572)	(0.0615)	(0.0546)	(0.0555)	(0.0516)	(0.0541)
WID	−0.1030	−0.1761*	0.0729	−0.0245	−0.1778	−0.2153	0.0586	−0.0692
	(0.0856)	(0.0886)	(0.0729)	(0.0782)	(0.1107)	(0.1129)	(0.0773)	(0.0806)

TABLE A-2 (continued)

Dependent Variable: ln (earnings)	1960				1970			
	Male		Female		Male		Female	
	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2
CHILD			−0.0083	−0.0062			−0.0196*	−0.0241*
			(0.0122)	(0.0129)			(0.0097)	(0.0102)
PARTIME	−0.2860†	−0.2993†	−0.4362†	−0.5445†	−0.1258*	−0.1565*	−0.1296*	−0.2469†
	0.0400)	(0.0414)	(0.0508)	(0.0539)	(0.0613)	(0.0620)	(0.0569)	(0.0588)
MOVE	−0.0535	−0.0475	−0.0146	−0.1985*	0.0370	0.0419	0.0492	0.0512
	(0.0474)	(0.0491)	(0.0878)	(0.0928)	(0.0507)	(0.0507)	(0.0680)	(0.0718)
GOVT	0.0776		0.1988*		−0.0084		0.1499*	
	(0.0510)		(0.1010)		(0.0538)		(0.0638)	
PROF	0.0758		0.3076		0.2602*		0.2872	
	(0.1364)		(0.2279)		(0.1139)		(0.1535)	
MANAGE	0.0477		0.1143		0.2824*		0.2164	
	(0.1656)		(0.3587)		(0.1249)		(0.1961)	
CLER	0.0086		0.3035		0.0144		−0.0073	
	(0.1193)		(0.2085)		(0.0989)		(0.1389)	
CRAFT	−0.0036		−0.1594		0.0524		−0.1950	
	(0.1176)		(0.3003)		(0.0973)		(0.2094)	
OPER	−0.0106		0.0357		−0.0239		−0.0995	
	(0.1139)		(0.2166)		(0.0943)		(0.1507)	
PRIV	−0.3197		−0.2672		−0.2065		−0.6454†	
	(0.2349)		(0.2202)		(0.4331)		(0.1675)	
SERV	−0.1367		−0.0925		−0.1556		−0.1740	
	(0.1144)		(0.1973)		(0.0943)		(0.1390)	

LABOR	−0.0961	0.3793	−0.1437	0.0463
	(0.1162)	(0.3465)	(0.0978)	(0.2237)
FARM	−0.5651		−0.9071*	0.5971
	(0.4311)		(0.4595)	(0.8580)
NOCC	0.0376	−0.1266		
	(0.1373)	(0.2843)		
AG	0.1090		−0.0164	−0.4323
	(0.3431)		(0.3637)	(0.7392)
MINE	0.2018		−0.0326	−0.0814
	(0.4783)		(0.3259)	(0.7314)
CON	0.2281†	0.0137	0.2049*	0.5193
	(0.0755)	(0.6271)	(0.0860)	(0.2960)
DURMAN	0.3661†	0.4443†	0.2310†	0.2979†
	(0.0506)	(0.1646)	(0.0645)	(0.1133)
NONDUR	0.3007†	0.1841	0.1707*	0.1913
	(0.0537)	(0.1365)	(0.0726)	(0.1103)
TRANS	0.2340†	0.2070	0.1861†	0.2611*
	(0.0569)	(0.2342)	(0.0719)	(0.1275)
WTRADE	0.2526†	0.1447	0.0465	0.2684
	(0.0824)	(0.2945)	(0.0907)	(0.1860)
FINANCE	−0.0262	−0.0276	0.0542	−0.0506
	(0.0843)	(0.1582)	(0.0970)	(0.1107)
BUSREP	−0.0591	−0.2281	0.0280	−0.0303
	(0.0942)	(0.2298)	(0.1017)	(0.1505)
PERSER	−0.0481	−0.1158	−0.1633	0.0210
	(0.0716)	(0.1146)	(0.1094)	(0.1057)
REC	−0.1631	0.2221	−0.7850†	0.5806*
	(0.1345)	(0.3620)	(0.2222)	(0.2833)
PROPSER	0.2074†	0.1001	0.0215	0.1476
	(0.0731)	(0.0986)	(0.0817)	(0.0832)

TABLE A-2 (concluded)

Dependent Variable: ln (earnings)	1960				1970			
	Male		Female		Male		Female	
	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2	Equation 1	Equation 2
PUBADM	0.2565†		0.2030		0.2527†		0.2799*	
	(0.0754)		(0.1573)		(0.0857)		(0.1152)	
NOIND	0.1158		0.2105					
	(0.0960)		(0.2266)					
CONSTANT	7.3907	7.4536	7.4039	7.3362	7.7694	7.7791	7.2733	6.8556
S.E.E.	0.4755	0.5003	0.6043	0.6598	0.7116	0.7296	0.7231	0.7682
R^2	0.33	0.24	0.44	0.31	0.19	0.14	0.28	0.17
NOBS	1,374	1,374	768	768	1,981	1,981	1,315	1,315

SOURCE: See Table A-1.
NOTE: Standard errors appear in parentheses. The difference between Equation 1 and Equation 2 is that Equation 2 does not control for government employment, occupation, or industry.
*Significant at the 0.05 level in a two-tailed test.
†Significant at the 0.01 level in a two-tailed test.

TABLE A-3 Estimated Equations for Changes in Occupational Distribution (1958–71) (Dependent variable: percent employment in a given occupation[a])

Occupation	Constant	Time	Unemployment Rate of White Males 35–44[b]	$\bar{R}^2$	Durbin-Watson	Constant	Time	Unemployment Rate of White Males 35–44[b]	$\bar{R}^2$	Durbin-Watson
	A. White Males					B. White Females				
Professional and technical	11.3643†	0.2736†	−0.2278	0.95	0.94	12.3432†	0.1972†	−0.0193	0.92	1.72
	(0.5480)	(0.0299)	(0.1338)			(0.4514)	(0.0246)	(0.1102)		
Managerial, officials, and proprietors	13.2259†	0.1093*	0.2450	0.32	1.74	4.6197†	−0.0033	0.1856	0.27	1.74
	(0.7473)	(0.0407)	(0.1825)			(0.4684)	(0.0255)	(0.1144)		
Clerical and sales	13.6593†	−0.0416	−0.0124	0.23	1.83	42.5092†	0.1812†	−0.5583*	0.84	1.72
	(0.4492)	(0.0245)	(0.1097)			(0.9368)	(0.0511)	(0.2288)		
Craftsmen	20.3804†	0.0574	−0.1986	0.64	1.33	0.9042†	0.0215*	0.0130	0.54	1.76
	(0.5230)	(0.0285)	(0.1277)			(0.1285)	(0.0070)	(0.0314)		
Operatives	22.9887†	−0.1853†	−0.9266†	0.77	1.42	19.1335†	−0.2298†	−0.9497†	0.91	1.77
	(0.5801)	(0.0316)	(0.1417)			(0.3786)	(0.0206)	(0.0925)		
Laborers	5.2030†	0.0180	0.3614†	0.44	1.10	−0.0432	0.0306†	0.0861	0.40	2.20
	(0.5441)	(0.0297)	(0.1329)			(0.1748)	(0.0095)	(0.0427)		
Service and private household	4.3243†	0.1155*	0.3495	0.35	2.27	17.2516†	0.0068	0.7475*	0.53	1.46
	(0.7091)	(0.0387)	(0.1732)			(1.0984)	(0.0599)	(0.2683)		
Farming	9.1708†	−0.3609†	0.3567†	0.99	1.69	3.1531*	−0.1974*	0.5238	0.71	0.97
	(0.2923)	(0.0159)	(0.0714)			(1.4126)	(0.0770)	(0.3450)		

TABLE A-3 (concluded)

Occupation	Constant	Time	Unemployment Rate of White Males 35–44[b]	$\bar{R}^2$	Durbin-Watson	Constant	Time	Unemployment Rate of White Males 35–44[b]	$\bar{R}^2$	Durbin-Watson
	C. Nonwhite Males					D. Nonwhite Females				
Professional and technical	2.5417†	0.3769†	0.0172	0.98	2.12	5.7407†	0.4021†	−0.2782	0.96	1.64
	(0.3678)	(0.0201)	(0.0898)			(0.6636)	(0.0362)	(0.1621)		
Managerial, officials, and proprietors	0.6109	0.2406†	0.4184†	0.86	1.78	0.6007	0.0586*	0.2362*	0.26	1.84
	(0.5462)	(0.0298)	(0.1334)			(0.4320)	(0.0236)	(0.1055)		
Clerical and sales	5.8846†	0.2473†	0.0474	0.77	1.27	3.0082	1.4183†	0.6044	0.95	0.61
	(0.9903)	(0.0540)	(0.2419)			(2.2699)	(0.1237)	(0.5544)		
Craftsmen	10.3353†	0.3378†	−0.5422*	0.95	1.96	0.2413	0.0404†	0.0509	0.71	2.32
	(0.7423)	(0.0405)	(0.1813)			(0.1560)	(0.0085)	(0.0381)		
Operatives	26.9486†	0.2436†	−1.1152†	0.92	2.02	16.1651†	0.1921	−0.9050	0.70	0.87
	(1.0229)	(0.0558)	(0.2498)			(1.7777)	(0.0969)	(0.4342)		
Laborers	24.7543†	−0.5894†	0.1466	0.90	1.26	0.2082	0.0194	0.1528*	0.21	2.76
	(1.5901)	(0.0867)	(0.3883)			(0.2760)	(0.0150)	(0.0674)		
Service and private household	15.7325†	−0.0622	−0.1116	−0.14	0.75	66.4093†	−1.4205†	−1.2912	0.76	0.54
	(2.0518)	(0.1119)	(0.5011)			(5.1175)	(0.2790)	(1.2499)		
Farming	13.3805†	−0.8018†	1.0908†	0.98	1.98	7.8541*	−0.7245†	1.3951	0.84	1.21
	(1.1225)	(0.0612)	(0.2742)			(3.1927)	(0.1740)	(0.7797)		

*Significant at the 0.05 level in a two-tailed test.
†Significant at the 0.01 level in a two-tailed test.
[a]Refer to footnote a in Table 6 of the text.
[b]Refer to footnote b in Table 3 of the text.

The cumulative distribution is given by

$$N(y) = \int_{Y_0}^{y} n(v)dv = 1 - Y_0^{\alpha} y^{-\alpha}$$

where $N(y)$ is the proportion of workers earning incomes less than or equal to y. Let $G(y)$ be the proportion of workers earning incomes greater than y, then

$$G(y) = 1 - N(y) = Y_0^{\alpha} y^{-\alpha}$$

or in terms of logs,

$$\ln G(y) = \alpha \ln (Y_0) - \alpha \ln (y)$$

Let $G(Y_1)$ and $G(Y_2)$ be the proportion of workers earning incomes greater than Y_1 and Y_2, respectively. By substituting these pairs of values into the equation for $\ln G(y)$ and solving for α, we obtain

$$\hat{\alpha} = -\ln [G(Y_1)/G(Y_2)]/\ln (Y_1/Y_2)$$

We have for 1960 $Y_1 = \$15{,}000$ and $Y_2 = \$25{,}000$, but for 1970 $Y_1 = \$35{,}000$ and $Y_2 = \$50{,}000$.

The average income of those earning above $Y_2(\bar{Y}_h)$ is calculated as the conditional mean of y for $y \geqq Y_2$:

$$\bar{Y}_h = E(Y|y \geqq Y_2) = \int_{Y_2}^{\infty} yn(y)\,dy \Big/ \int_{Y_2}^{\infty} n(y)\,dy = \frac{\hat{\alpha}}{\hat{\alpha}-1} Y_2$$

where $Y_2 = \$25{,}000$ and $\$50{,}000$ for 1960 and 1970, respectively. Values of $\bar{Y}_h$ were estimated in each year for the combined sample of all four race/sex groups. Each worker whose income was reported in the open-ended interval was assigned an income of $\bar{Y}_h$. For the independent variables, these workers were assigned the mean values corresponding to the members of their particular race/sex group whose incomes were in the open-ended interval.

The independent variables used in the earnings regressions are briefly defined here. EXP = potential experience calculated as age − education − 6 years. EDUC = years of schooling completed. NEAST, NCENT, and WEST are regional dummy variables representing the Northeast, Northcentral, and West, respectively. The South is the regional reference group. $U499$, $UR499$, and $UR500$ are urban city size dummy variables representing residence in a central city located in an urban area of 50,000–499,999, residence in the remainder of an urban area 50,000–499,999, and residence in the remainder of an urban area of 500,000 or greater. Residence in a central city located in an urban area of 500,000 or greater is the city size reference group. SINGLE, DS, and

WID are marital status dummy variables representing never married, divorced or separated, and widowed. Married workers form the marital status reference group. CHILD = the number of children born to the female worker. PARTIME is a dummy variable which identifies individuals who worked less than thirty-five hours during the week prior to the Census week. MOVE is a dummy variable which identifies individuals who have changed county, state, or national residence since the age of eighteen during the previous five years. GOVT is a dummy variable which identifies government workers. The occupational dummy variables are defined as follows: PROF = professional, technical, and kindred workers; MANAGE = managers and administrators, except farm; CLER = clerical and kindred workers; CRAFT = craftsmen and kindred workers; OPER = operatives (including transport equipment operatives); PRIV = private household workers; SERV = service workers; LABOR = laborers, except farm; FARM = farmers, farm managers, farm laborers, and farm foremen; and NOCC = no occupation reported. Sales workers form the occupational reference group. The industry dummy variables are as follows: AG = agriculture, forestry, forestry, and fisheries; MINE = mining; CON = construction; DURMAN = manufacturing, durable goods; NONDUR = manufacturing, nondurable goods; TRANS = transportation, communications, and utilities and sanitary services; WTRADE = wholesale trade; FINANCE = finance, insurance, and real estate; BUSREP = business and repair services; PERSER = personal services; REC = entertainment and recreation services; PROFSER = professional and related services; PUBADM = public administration; and NOIND = industry not reported. Retail trade workers form the industry reference group.

NOTES

1. Computer delays and difficulties precluded carrying out these tests. For additional information on these testing procedures see [Oaxaca 1974].
2. The earnings data from the 1960 and 1970 censuses correspond to 1959 and 1969, respectively.
3. Because of rounding-off errors, these mutually exhaustive decompositions do not, in general, add up to the exact value of the gross differential in logs.
4. Had we instead assumed the female earnings structure to be the common earnings structure, the effects of childbearing would have shown up entirely as the effects of differences in the mean values of the independent variables, e.g., as a source of earnings differentials attributable to worker productivity differences.
5. This specification was suggested to the author by Andrea Beller.

6. The time-series estimates of d along with their associated standard errors are given below.

Whites		Nonwhites
0.0025*	0.0026*	−0.0218†
(0.0011)	(0.0011)	(0.0037)

*Significant at the 0.05 level in a two-tailed test.
†Significant at the 0.01 level in a two-tailed test.

7. To handle the discrepancies from rounding-off errors in the cross-section decompositions, the difference between the actual and calculated gross differentials in logs was apportioned equally between the measures of discrimination and productivity differences reported in Tables 1 and 2.
8. The results also predict that the male/female labor quality ratio for whites should peak in 1974.
9. The separate occupational regressions are reported by race and sex in Table A-3 of the appendix.

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8 COMMENTS

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Ronald Oaxaca has applied a previously developed methodology to analyze trends in male-female earnings differentials in the period 1955–71.[1] The approach is an attempt to sort out the effect of discrimination from other factors that might account for the earnings gap and to test hypotheses about changes in sex discrimination since 1955.

Oaxaca concludes from his analysis that for white year-round, full-time workers, the male-female differential in median earnings (adjusted for hours worked) increased by 15 percent from 1955 to 1971 with about 55 percent of this attributable to increased discrimination and the rest due to the relative increase in an index of male to female labor quality. He found a small rate of decline in the rate of increase in discrimination after 1965 which he suggests may be related to more favorable government policies toward women reflected in the 1963 Federal Equal Pay Act and the 1964 Civil Rights Act.

For nonwhite year-round, full-time workers, Oaxaca found a significant decline in the male-female earnings differential, from 73 percent in 1955 to 33 percent in 1971 and concludes that practically all of this gain for black women represented a decline in sex discrimination. He makes the interesting, and I believe accurate, suggestion that this decline in the black sex differential not attributable to differences in labor quality was more likely due to black women gaining access to jobs formerly held by white women than their penetration of male occupations. If this is true, it is, of course, inappropriate to refer to the effect as a decline in sex discrimination per se, but instead it should be attributed to differential rates of change of racial discrimination by sex.

Oaxaca found very little change in male-female occupational structure over the period. Employment of women relative to men increased most rapidly in the sales and clerical fields that were already traditionally dominated by women, supporting the view that the relative deterioration of female wages during a period of rapid influx into the labor force is due to crowding into occupations where their wages are low because of their exclusion from other kinds of work.[2] White females lost representation relative to white males in the Professional and Technical, Managerial, and Officials and Proprietors categories, suggesting that efforts to penetrate higher status occupations have not been successful. Oaxaca also found that the sex earnings gap falls in tight labor markets, suggesting that prosperity promotes equality and recession fosters discrimination.

The tone of Oaxaca's conclusions is hedgingly optimistic. We still have sex discrimination, particularly among whites, but its rate of increase is on the decline. And black women are faring much better relative to black men, albeit at the expense of white women. And if we can regain prosperity and full employment we can expect even more success.

Oaxaca's paper addresses an extremely important issue, particularly in view of an explicitly stated governmental policy to reduce discrimination in the United States. Unfortunately, I cannot concur with the author's optimistic interpretation of the trends.

My objections are concerned with both his methodology and interpretation. I have always been uneasy about studies that treat discrimination as an unexplained residual, since we have available a number of hypotheses about the way discrimination is transmitted that are certainly susceptible to empirical testing. But if we can ignore that issue—and let me say parenthetically that I think it is extremely dangerous to let this research be disseminated outside of econometric conferences with the label "discrimination" applied to the unexplained residual—we must, at least, be sure that we have accounted for all the factors that might have been responsible for changes in the sex earning gap before we draw conclusions about trends in discrimination. This isn't to say everything must be included—but the important things must be.

To use an index of years of schooling as a measure of labor quality and then to use this index as the sole source of nondiscriminatory wage change (other than hours worked) may be permissible for a study of black-white earnings differentials, a la Ashenfelter,[3] but its extension to an analysis of sex discrimination is not warranted. Oaxaca observes that a major difference between race and sex factors is that female labor force participation has increased sharply relative to that of men, whereas this has not been the case for blacks relative to whites. What Oaxaca does not observe is that the age composition of the female labor force has also undergone a radical change over the period, particularly for white women, with an important shift occuring around 1965 that increased the relative participation of women under 25. We are all aware that the age profiles of male-female earnings, particularly for whites, are not very far apart until around age 25 and then widen rapidly until age 55 when they begin to close again. There are a number of reasons for this, involving barriers to the acquisition of human capital and to managerial jobs for women, but I do not have time to explore these issues here. The point is that these discrepancies in the age-profiles of the male-female earnings rates, together with a shift in the age composition of new labor force entrants around 1965, can more than account for the so-called decline in the rate of increase in discrimination Oaxaca observes when his data are not adjusted for age.

Table 1 shows the percent change in labor force participation rates for white women over the period 1955–71, and for the subperiods 1955–65 and 1966–71.

TABLE 1 Percentage Change in Labor Force Participation Rates of White Females by Age, 1955–71

	Age								
Period	16+	16–17	18–19	20–24	25–34	35–44	45–54	55–64	65+
1955–71	23.5	21.7	5.8	26.4	32.9	25.8	25.8	33.6	−11.4
1955–65	10.4	−4.0	−2.7	7.4	10.7	11.0	16.9	26.7	−7.6
1966–71	11.8	26.8	8.7	17.7	20.1	13.3	7.6	5.5	−4.1
Female/male earnings ratio (1966)	.57	.87	.87	.68	.55	.48	.50	.55	.62

SOURCE: Labor Force Participation Rates: U.S. Department of Labor, *Manpower Report of the President*, 1973. Earnings Ratio: Isabel V. Sawhill, "The Economics of Discrimination Against Women: Some New Findings," *Journal of Human Resources* 8 (Summer 1973), pp. 383–395. Estimated from the *Current Population Survey*, 1966.

In the earlier period, the growth in participation for young women was considerably below the average for all women, with the highest rates of entry coming in the age group with the largest male-female earnings gap. Since 1966–71 the rate of entry is much higher for young women, who experience much less discrimination with respect to earnings. However, and this is extremely important, just because these young women have higher earnings relative to men than older women, there is no evidence that they will not face exactly the same sort of wage discrimination for the same reasons when they get older. Hall and others have suggested that older black workers as a group have relatively low earnings because of a "vintage effect," that is, they have less education than younger workers.[4] This implies that the improved earnings position of younger black workers represents a real gain due to a decline in discrimination in educational opportunity. But educational opportunities have not changed significantly for women over the period and lower wages for older women undoubtedly reflect their inability or unwillingness to acquire human capital or to attain managerial status that would increase their earnings profile as they age. This means that an increase in overall female earnings as a result of a reduction in the average age of the female work force does not imply that discrimination has declined or that any individual woman has a higher expected lifetime earnings stream than before, even though a different interpretation could be made for blacks. It may well be that consciousness-raising has produced an incentive for younger women to penetrate male occupations, but there is not evidence of this and Oaxaca's occupational results do not suggest such a trend.

In Table 2, I show some crude estimates of the impact of changes in the age structure of the female labor force that would have occurred if there had been no change in the male-female earnings ratio by age. *ER* is the female-male earnings ratio by age for full-time workers computed from the Current Population Survey in 1966 and adjusted for hours worked. Over the entire 1955–71 period, for white females, changes in their age distribution alone should have *increased* the overall *ER* by .61 percent per year. In the earlier period, 1955–65, the potential annual rate of increase was .41 percent, while in 1966–71 it was 1.2 percent! Now these estimates are admittedly crude and the use of 1966 earnings data was arbitrary. However, Fuchs has shown that there is not much difference in the age profiles of the sex earnings ratios between 1960 and 1970.[5] On the other hand, there is likely to be some interaction in the labor quality index that would reduce the absolute marginal impact of the age effect.

Apart from the implications for the absolute changes in sex discrimination (that now will increase by the amount of the potential increase in *ER* associated with changing age composition) what is most interesting is the huge difference in this effect before and after 1965. Not only has the rate of increase in the wage gap attributable to discrimination increased by about .6 percent more per year overall than Oaxaca estimates (less any interaction with the labor quality index), but the rate of increase in discrimination would increase substantially (by .8 percent per year) after 1965 rather than decrease. Not only is all the slowdown Oaxaca observes explained by the changing age distribution of the female labor force, but it is more than accounted for and by a large margin, indicating an increase rather than decline in the *rate of increase* of discrimination. The influx of young men into the labor force since the deescalation in Vietnam has undoutedly worked in the same direction, but I have not tried to estimate that effect.

Table 3 shows similar computations for black females. Changing age distribution had very little effect on their earnings relative to black males, except after 1965 when a changing age composition had the potential effect of increasing the female-male earnings ratio by .9 percent per year. This implies that after 1965, Oaxaca overstated the amount of reduction in sex discrimination (if one can properly call it that) in the black labor force.

Not only does Oaxaca fail to correct for age, but the assumptions he makes about his labor quality index may also bias his results. For instance, in order to obtain the result that the rate of discrimination declined after 1965 he had to assume that the male and female labor quality indexes grew at a constant rate. Since his entire paper was concerned with trend analysis, I am puzzled that he did not test for a nonlinear trend in the labor quality indexes also. Since there is, for all practical purposes, a finite limit to the amount of formal education one receives (statistically it is the 16+ category), then labor quality cannot grow indefinitely. More than likely, the relative educational gains of white males is a once and for all phenomenon. Furthermore, as Oaxaca observes, if labor quality

TABLE 2 Estimated Effect of Changing Age Composition of the White Female Labor Force on the Female/Male Earnings Ratio, 1955–71

Age (Group i)	Change in Labor Force of Age Group (Thousands) (1)	Change in (LF)$_i$ Relative to Overall Change in Labor Force (2)	$\frac{ER_i}{ER}$ (3)	(2) × (3) (4)
	1955–71 (Average Annual Rate of Change 0.61 Percent)			
16+	10,103	1.00	1.00	
16–17	634	0.063	1.534	.097
18–19	783	0.078	1.534	.120
20–24	2,285	0.226	1.199	.271
25–34	1,422	0.141	0.970	.137
35–44	952	0.094	0.846	.080
45–54	2,160	0.214	0.881	.189
55–64	1,631	0.161	0.970	.156
65+	236	0.023	1.093	.025
				Σ = 1.073
	1955–65 (Average Annual Rate of Change 0.41 Percent)			
16+	4,870	1.00	1.00	
16–17	286	0.059	1.534	.091
18–19	439	0.090	1.534	.138
20–24	773	0.159	1.199	.191
25–34	22	0.005	0.970	.005
35–44	745	0.153	0.846	.129
45–54	1,378	0.283	0.881	.249
55–64	1,046	0.215	0.970	.209
65+	159	0.033	1.093	.036
				Σ = 1.047
	1966–71 (Average Annual Rate of Change 1.2 Percent)			
16+	4,287	1.00	1.00	
16–17	266	0.062	1.534	.095
18–19	119	0.028	1.534	.043
20–24	1,299	0.303	1.199	.363
25–34	1,236	0.288	0.970	.279
35–44	189	0.044	0.846	.037
45–54	633	0.148	0.881	.130
55–64	454	0.106	0.970	.103
65+	91	0.021	1.093	.023
				Σ = 1.074

SOURCE: See Table 1.

TABLE 3 Estimated Effect of Changing Age Composition of the Black Female Labor Force on the Female/Male Earnings Ratio, 1955–71

Age (Group *i*)	Change in Labor Force of Age Group (Thousands) (1)	Change in (LF)$_i$ Relative to Overall Change in Labor Force (2)	$\frac{ER_i}{ER}$ (3)	(2) × (3) (4)
1955–71 (Average Annual Rate of Change 0.15 Percent)				
16+	1,439	1.00	1.00	
16–17	57	0.040	1.29	.052
18–19	95	0.066	1.25	.083
20–24	342	0.238	1.12	.267
25–34	259	0.180	1.02	.184
35–44	198	0.138	0.917	.127
45–54	256	0.178	0.899	.160
55–64	194	0.135	0.951	.128
65+	41	0.028	1.07	.030
				Σ = 1.029
1955–65 (Average Annual Rate of Change 0.0 Percent)				
16+	801	1.00	1.00	
16–17	27	0.034	1.29	.044
18–19	37	0.046	1.25	.057
20–24	147	0.184	1.12	.206
25–34	55	0.069	1.02	.070
35–44	171	0.213	0.917	.195
45–54	181	0.226	0.899	.203
55–64	148	0.185	0.951	.176
65+	36	0.045	1.07	.048
				Σ = 1.000
1966–71 (Average Annual Rate of Change 0.9 Percent)				
16+	505	1.00	1.00	
16–17	12	0.024	1.29	.031
18–19	24	0.048	1.25	.060
20–24	183	0.362	1.12	.405
25–34	188	0.372	1.02	.379
35–44	8	0.016	0.917	.015
45–54	53	0.105	0.899	.094
55–64	35	0.069	0.951	.066
65+	2	0.004	1.07	.004
				Σ = 1.055

SOURCE: See Table 1.

is related to experience as well as to education, then the growing attachment of females to the labor force ought to raise relative female labor quality. If these factors were taken into account, male labor quality relative to female labor quality would not have increased at a constant rate over the period and the earnings gap attributable to discrimination would have been greater. Furthermore, if the rate of relative gains of men slowed later in the period, then the decline in the rate of growth of discrimination cited by Oaxaca would not have been observed.

In general, I find it very difficult to be optimistic about trends in sex discrimination when so little progress seems to be made in reducing occupational segregation by sex. We all know that instances of unequal pay for equal work are rare and that the Equal Pay Act of 1963 merely ratified existing practice. The Civil Rights Act of 1964 was more specifically addressed to the issue of occupational discrimination, but the evidence is that women are losing rather than gaining representation in high status occupations.

There are other problems. The paper fails to address the consequences of inability of many women to get full-time work. The interpretation of the cyclical results also has to be viewed with caution. My own work, for instance, has shown female unemployment to rise relative to men's in tight labor markets,[6] suggesting that prosperity does not unequivocally improve the labor force status of women.

Most of my remarks have been addressed to sex differentials among whites, in part because discrimination on the basis of sex seems to be more of a problem and a growing problem within that group. However, while I have less quarrel with the data for sex differences between blacks, I have difficulty accepting the interpretation that sex discrimination per se is any less present. Black women are not penetrating male occupations; racial discrimination is simply less severe among women. Undoubtedly this is because no women—black or white—gain much from seniority. Hierarchical relationships in female employment are less prevalent and black women and white women are systematically excluded (voluntarily or involuntarily) from acquiring human capital and attaining managerial positions. To the extent that black women drop out of the labor force less frequently than white women, they may even have an advantage. It is disturbing indeed if this equalization among women is cited as a triumph for racial equality and as a sign of lessening of racial discrimination. There are simply fewer channels through which racial discrimination can be transmitted among females than among males. And my own personal opinion is that women workers have less of a Beckerian taste for discrimination[7] in general, in part because they feel less threatened by the encroachment of racial minorities on their jobs or pay standards.

Despite these reservations about the interpretation of the experience of the 1960s, for a number of reasons I, personally, see some grounds for optimism about the future relative income position of women. First and foremost are the social changes and the growing awareness of young women that sex roles based on outdated needs and institutions need not limit their labor force experience to being a nurse, secretary, domestic worker, saleslady, or elementary school teacher. I think young women are growing increasingly to expect a more or less continuous labor force commitment and that the men they live with will, in turn, see that these trends are to their advantage as well. Most of the evidence that I have seen suggests that if progress is being made along these lines it is at a snail's pace. The ideology of the women's movement has developed much faster than its practical implications. But to evaluate these prospects in a rigorous and scientific way, as opposed to an impressionistic one, one needs an underlying theoretical rationale for discrimination on which to base a structural model. For instance, one approach suggests that women as a group will only attain equality with men if they penetrate male occupations at an early age, stay in the labor force, and acquire human capital and assume managerial responsibilities conducive to substantial growth of their wages over the life cycle. In other words, occupational choices of *young* women may be the key to forecasting the future age-earnings profile of all female workers. The empirical analogue to this model would be to examine changes in the sex composition of the various occupational categories by age group to determine whether or not the observed tendency of the female-dominated occupations to absorb most of the increases in the female labor force is due to the choices of young women or instead to those of older women whose career horizons may be more limited. The implications for the future course of male-female wage differentials may be much more interesting than any simple extrapolation in the trend of some unexplained residual—labeled "discrimination" only for want of a better word, and perhaps to elicit some emotional reaction. Such pure empiricism can only run up against some of the problems I have described. Again, I applaud the effort to measure discrimination and to ascertain whether or not it is waning. But to design effective social policy to eliminate discrimination, one must develop a theoretical explanation, not just a chart of its course.

NOTES

1. Oaxaca's paper extends the methodology developed by Gary S. Becker in *The Economics of Discrimination* (Chicago: University of Chicago Press, 1957) and applied by Orley Ashenfelter in "Changes in Labor Market Discrimination over Time," *The Journal of Human Resources* 5 (Fall 1970), pp. 403–430, to analyze racial discrimination.
2. This hypothesis is suggested by Barbara Bergmann in "The Effect on White Incomes of Discrimination in Employment," *Journal of Political Economy* 74 (March-April 1971), pp. 294–313.
3. Ashenfelter, "Changes in Labor Market Discrimination."
4. See Robert E. Hall's discussion of a paper by Richard B. Freeman, "Changes in the Labor Market for Black Americans," *Brookings Papers on Economic Activity* (1973: 1), p. 126.
5. Victor R. Fuchs, "Short-Run and Long-Run Prospects for Female Earnings," presented at the Annual Meeting of the American Economic Association, December, 1973.
6. Nancy S. Barrett and Richard D. Morgenstern, "Why Do Blacks and Women Have High Unemployment Rates?" *The Journal of Human Resources* 9 (Fall 1974), pp. 452–462.
7. Becker, *The Economics of Discrimination.*

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Oaxaca's paper represents an enormously innovative and up-to-date measurement of the discrimination component in the earnings differential between women and men. He also provides a useful framework for tracking its changing behavior over time.

Oaxaca adapts a discrimination model measuring the adjusted growth rate of male-female earnings differences, with a cyclical measure of labor market tightness, and time as the main arguments accounting for changes in the behavior differential. The gross earnings differential provides a residual which, when adjusted by an index employed to account for male-female labor quality differences, purports to "explain" the amount of the widening pay differential between men and women over the 1957–71 period attributable to discrimination.

The same determinants, i.e., a cyclical component and time trend, are also arguments in the differential changes in male-female occupational distribution over time and in measuring the effects of changes in occupational distribution on the gross male-female earnings differential.

Oaxaca finds that the gross male-female earnings differential rose 15 percentage points, from 55 percent in 1955 to 70 percent by 1971, and his results estimate that more than one-half of the differential was due to sex discrimination, with the discrimination component on the increase up to 1966. He finds that a reduction in discrimination from 1966 onward probably operated to check further widening in the earnings gap after 1968.

He estimates that part of the slowdown in the widening pay gap may have been caused by the lagged impact of federal equal pay legislation (1963), and the 1964 Civil Rights Act, but notes that the legislation's enforcement machinery operated concurrently with other factors (e.g., labor market tightness and the military draft) which were probably far more significant.

At any rate, Oaxaca's own results suggest that Equal Employment Opportunity Commision and equal pay enforcement effects were minimal in this respect. On the whole, Oaxaca's findings are impressive. His estimate (of over 50 percent) certainly exceeds a different one recently made for the Council of Economic Advisers, which claimed that only about 20 percent of the pay difference between men and women was due to sex bias.

This is the more remarkable since Oaxaca's index of relative labor productivity is constructed in a manner that fails to account adequately for recent shifts in the labor market behavior of women in terms of their labor force participation and attachment, and in the rising proportion of the total U.S. labor force comprised of women (from 30 percent in 1950 to 41 percent in 1973).

Now, during this period males gained on females in terms of school years completed, but male labor force participation and the proportion of men in the labor force have declined since 1960, whereas, for females, dramatic rises occurred in both these statistics. Oaxaca's index therefore overstates the importance of male labor productivity changes relative to females and correspondingly probably underestimates the amount of the male-female pay differential properly ascribable to discrimination.

This point concerning the rising importance of female labor in the U.S. economy deserves further elaboration, since it has crucial implications for the future of female job segregation and unemployment. Neoclassical theories regarding women's "weak" attachment to the labor market are founded on nothing much stronger than outmoded sexist notions of women's "proper" social role. They satisfy what Galbraith terms a "convenient social virtue" of facilitating growing household consumption. Whether or not one can accept this view, it is significant that almost every projection made in recent years regarding labor force growth has been consistently understated, usually by the amount of the substantial underestimate in the growth of the female labor force.

Now, given the growing segregation of women in jobs predominantly "female," such as elementary and secondary school teacher and clerical worker, combined with the inflationary surge certain to "push" more women out of the home and into the labor market, unless male labor force participation declines more drastically, relative unemployment for women can be expected to move sharply upward in the coming years. This may lead to further widening in the amount of the male-female differential due to discrimination, because the additional women entrants must be absorbed in jobs more restrictive in terms of sex bias and must try to get hired by employers with higher than average "discrimination coefficients." One possible counter to forestall this outcome would be a more determined federal enforcement against, and monitoring of, sexist labor market practices.

A final point, Oaxaca offers two alternative interpretations of the reduction in the nonwhite male-female earnings differential—a marked contrast to the white relationship. One is greater relative reduction in sex discrimination against nonwhite females relative to a reduction in racial discrimination against nonwhite males. The other interpretation, which he feels to be the more likely, and which I would like to underscore, is that while nonwhite females have not really invaded traditionally nonwhite male jobs, they have probably been relatively more successful in "nailing down" a larger share of traditional white female jobs (e.g., clerical and health workers).

Moreover, given the unaccountable decline in the black male labor force participation, the stronger labor force attachment of black women relative to white women, and finally the persistence of black male segregation in jobs such as nonfarm laborers and operatives—employment highly sensitive to the level of aggregate activity—it is hardly surprising that the diminution of the male-female pay differential has been more marked among nonwhites.

Oaxaca mentions the sort of "chicken or egg" dispute among economists as to the extent to which differing sex occupational distributions are due to job market bias, and the extent to which they may be ascribed to differing job preferences between men and women. He points to some "promising" research conducted by psychologists testing whether there are "real" differences by sex as to motivations, aspirations, and expectations regarding career fulfillment. I, for one, am not too hopeful that any illuminating result will emerge from such findings. Labor market sex bias operates, on the one hand, directly to constrict female job opportunities, while generating, on the other hand, intricate feedback which conditions female job and career aspirations, expectations, and motivations from adolescence to old age. I know of no generally acceptable econometric technique which could adequately model these joint tendencies.