THE MEASUREMENT OF RACE AND GENDER WAGE DIFFERENTIALS: EVIDENCE FROM THE FEDERAL SECTOR

GEORGE J. BORJAS*

This paper presents an empirical analysis of wage differentials based on race and gender in the federal bureaucracy. By focusing on the study of interagency variations in wage differentials, the author shows, first, that the use in earnings functions of a simple dummy variable to indicate race and gender leads to downward-biased estimates of the standardized wage differential. Second, across federal agencies there is a positive correlation between wage differentials based on race and those based on gender. Finally, the low relative wage of black females is more a result of their gender than of their race. This variety of empirical findings shows the promise of future studies that concentrate on the interfirm variance in employment policies that affect women and minorities.

THE analysis of wage differentials based on race and gender has evolved into a voluminous literature in the past decade.[1] Most of this work has focused on measuring standardized wage differences among the various racial groups and the two genders. The statistical procedure used is a decomposition of the observed wage differential into a portion "explained" by observable characteristics and an "unexplained" residual customarily labeled discrimination.[2] Labor economists usually find a sizable unexplained wage differential among the races and sexes in the data sets currently in general use.

This paper presents an empirical analysis of race and gender differences in earnings in the federal bureaucracy. This "firm" employed over 2.4 million workers in 1978, of whom 31.1 percent were women and 22 percent were classified as minorities.[3] These workers were employed by a myriad of federal agencies. To the extent that agen-

*The author is Professor of Economics at the University of California, Santa Barbara. His research was funded by a grant from the Employment and Training Administration, U.S. Department of Labor.

[1]The literature on wage differentials by race is surveyed by Richard B. Freeman, "Labor Market Discrimination: Analysis, Findings, and Problems," in Michael D. Intriligator and David A. Kendrick, eds., *Frontiers of Quantitative Economics* (Amsterdam: North-Holland, 1974), pp. 501 - 62; and Ray Marshall, "The Economics of Racial Discrimination: A Survey," *Journal of Economic Literature*, Vol. 12, No. 3 (September 1974), pp. 849 - 71. The literature on gender differences is surveyed by Cynthia B. Lloyd and Beth T. Niemi, *The Economics of Sex Differentials* (New York: Columbia University Press, 1979).

[2]The problems associated with this interpretation are well documented in the literature. A particularly clear discussion of these problems is given by Solomon W. Polachek, "Potential Biases in Measuring Male-Female Discrimination," *Journal of Human Resources*, Vol. 10, No. 2 (Spring 1975), pp. 205 - 29.

[3]U.S. Office of Personnel Management, *Equal Employment Opportunity Statistics* (Washington, D.C.: GPO, November 1978), p. 2.

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cies vary in their personnel policies, a new dimension can be added to the empirical study of wage differences, namely, the existence of interfirm variations in the hiring and placement of blacks and women.[4]

The objective of this paper is to show that the availability of interfirm data on race and gender wage differentials provides a unique opportunity to expand our understanding of the measurement and interpretation of these differences. In particular, the empirical work below will analyze the relationship among various measures of wage discrimination currently employed by economists. The analysis will also focus on the relationship among the "discrimination coefficients" for the different race and gender groups.

Measurement of Wage Differentials

This study analyzes data from the Central Personnel Data File (CPDF) compiled by the Office of Personnel Management, specifically data on civilian workers employed by the federal government on December 31, 1979. The sample is composed of a one percent random sample of personnel records in the eight largest federal agencies and of a 10 percent random sample for all other agencies.[5] Oversampling the smaller agencies ensures that even the relatively small federal agencies are represented in the analysis. Each record outlines such personal characteristics of the individual as education, race, and gender and also notes such employment characteristics as agency of employment, annual full-time earnings, and years of government service. The analysis here is restricted to permanent, full-time civilians working in the United States whose records report the key variables under study.

Table 1 presents statistics designed to measure the race and gender wage differential in both the pooled sample and in each of the 31 largest federal agencies. The 31 agencies employ over 95 percent of the total civilian work force of the federal government. The first column of the table indicates the average log of annual (full-time) earnings for white males ($\overline{\ln r^{wm}}$) in each sample. Since the average annual wage for white men is relatively high ($e^{9.9718}$ = \$21,414), it should be clear that even small percentage wage differentials actually reflect sizable dollar differentials.

Initially, a simple regression technique serves to measure the wage differential between white males and each of the other race and gender groups, holding observable skills constant. In particular, the following earnings function is estimated for each federal agency (and for the pooled sample):

$$(1) \qquad \ln r_h = Y_h \alpha + \beta R_h + \epsilon_h,$$

where r_h denotes the annual full-time earnings of individual h; Y_h is a vector of h's socioeconomic characteristics; and R_h is set equal to unity if the individual is a member of a particular race-gender group and zero if he is a white man. The initial estimates of unexplained wage differentials—as measured by the coefficient β—simply allow membership in a particular race-gender group to shift the intercept of the equation. Equation 1 will be estimated three times in each agency to make pairwise comparisons between white men and each of the other groups.

In other words, a regression will first be calculated for the sample of men, to estimate the unexplained wage differential between white men and black men.[6] The second regression is run on the sample of whites, thus providing an estimate of the unexplained wage differential between white men and white women. Finally, the sample for the third regression is restricted

[4]The fact that federal agencies differ in their personnel policies is discussed and documented in George J. Borjas, "Wage Determination in the Federal Bureaucracy: The Role of Constituents and Bureaucrats," *Journal of Political Economy*, Vol. 88, No. 6 (December 1980), pp. 1110-47.

[5]The eight largest agencies are the U.S. Departments of Defense; Agriculture; Justice; Health, Education and Welfare; Transportation; Treasury; and the Postal Service and the Veterans Administration. A coding error in the creation of the data led to an 11 percent random sample from the U.S. Department of the Interior.

[6]Throughout the paper, the sample of blacks includes individuals classified as minority in the CPDF.

Table 1. The Measurement of Interagency Wage Differentials.
(t-ratios in parentheses)

Agency	$\overline{\ln r_i^{wm}}$	$\hat{\beta}_{bm}$	$\hat{\beta}_{wf}$	$\hat{\beta}_{bf}$	Δ_{bm}	Δ_{wf}	Δ_{bf}	*Number of Observations*
Pooled[†]	9.9718	−.1188*	−.2179*	−.2702*	−.1743	−.2726	−.3123	19,939
		(−21.92)	(−41.17)	(−39.07)				
Cabinet Agencies								
Agriculture	9.9382	.0002	−.1722*	−.2190*	−.0333	−.2019	−.2434	898
		(0.00)	(−8.10)	(−5.36)				
Commerce	10.2021	−.1490*	−.2425*	−.3395*	−.1815	−.3103	−.3674	2,945
		(−9.62)	(−17.05)	(−19.41)				
Defense	9.9540	−.1161*	−.2897*	−.3494*	−.1366	−.3201	−.3721	7,702
		(−14.74)	(−38.33)	(−31.18)				
Energy	10.2619	−.2054*	−.3752*	−.4192*	−.2262	−.3977	−.4421	1,789
		(−7.82)	(−19.89)	(−14.99)				
Health, Education, and Welfare	10.0935	−.1403*	−.1282*	−.2504*	−.2092	−.2261	−.3050	1,378
		(−4.51)	(−6.25)	(−9.58)				
Housing and Urban Development	10.2594	−.1258*	−.3508*	−.4575*	−.1601	−.4156	−.5327	1,432
		(−5.80)	(−16.22)	(−18.74)				
Interior	10.0108	−.1059*	−.2206*	−.2822*	−.1434	−.2668	−.3174	6,188
		(−11.58)	(−23.42)	(−21.99)				
Justice	10.0725	−.0799*	−.2316*	−.2643*	−.0850	−.2725	−.2891	511
		(−2.33)	(−7.66)	(−6.83)				
Labor	10.1952	−.1220*	−.3861*	−.5353*	−.1414	−.4307	−.5789	2,161
		(−5.85)	(−21.46)	(−25.14)				
State	10.4508	−.1915*	−.1998*	−.3208*	−.2175	−.2014	−.2976	658
		(-5.44)	(−6.24)	(−8.26)				
Transportation	10.2181	−.0718*	−.4662*	−.5190*	−.1011	−.4939	−.5356	1,145
		(−1.99)	(−12.06)	(−8.23)				
Treasury	10.0638	−.1312*	−.2053*	−.3619*	−.1391	−.2631	−.3629	1,947
		(−4.38)	(−9.28)	(−12.42)				
Independent Agencies								
Environmental Protection Agency	10.1728	−.0076	−.1343*	−.2714*	−.0305	−.1823	−.3031	1,043
		(−.23)	(−6.38)	(−8.91)				
Equal Employment Opportunity Commission	10.1089	.0130	−.0984	−.1910*	—	—	—	257
		(.22)	(−.95)	(−2.30)				
Federal Communications Commission	10.2757	−.0131	−.0817	−.2390*	.0146	−.0817	−.1685	189
		(−.17)	(−1.43)	(−3.32)				

Table 1, Continued.

Agency	$\overline{\ln r_i^{wm}}$	$\hat{\beta}_{bm}$	$\hat{\beta}_{wf}$	$\hat{\beta}_{bf}$	Δ_{bm}	Δ_{wf}	Δ_{bf}	*Number of Observations*
Federal Deposit Insurance Corporation	10.1150	−.1620* (−2.95)	−.1828* (−4.57)	−.2517* (−3.93)	−.1542	−.1344	−.2519	316
Federal Home Loan Bank Board	10.2279	−.3548* (−3.63)	−.3648* (−4.16)	−.4798* (−5.44)	−.4521	−.4666	−.5667	133
Federal Trade Commission	10.4289	−.1154 (−1.66)	−.1185 (−1.88)	−.2606* (−2.91)	−.1857	−.1324	−.2904	162
General Accounting Office	10.3072	−.1108* (−2.90)	−.1079* (−3.76)	−.2514* (−5.77)	−.1053	−.1240	−.1393	486
General Services Administration	9.9153	−.1327* (−10.79)	−.2095* (−12.80)	−.3031* (−17.62)	−.1849	−.2531	−.3584	3,123
Government Printing Office	10.1562	−.3326* (−14.67)	−.4462* (−10.86)	−.5359* (−19.20)	−.3813	−.5270	−.5764	711
International Communications Agency	10.2751	−.2661* (−6.30)	−.1401* (−3.42)	−.3349* (−6.71)	−.3399	−.1800	−.3429	332
Interstate Commerce Commission	10.2934	−.2626* (−2.76)	−.1941* (−2.62)	−.5264* (−6.81)	−.3198	−.2748	−.4948	163
National Aeronautics and Space Administration	10.3578	−.0975* (−5.81)	−.2970* (−20.13)	−.3267* (−14.66)	−.1335	−.3568	−.3407	2,214
National Labor Relations Board	10.3834	−.1283* (−2.47)	−.0934* (−2.48)	−.2465* (−4.17)	−.0893	−.2191	−.2294	253
Nuclear Regulatory Commission	10.5284	−.1184* (−2.45)	−.2878* (−5.52)	−.4621* (−5.52)	−.1423	−.3934	−.4671	273
Office of Personnel Management	10.1284	−.1260* (−2.70)	−.1544* (−4.29)	−.2142* (−5.08)	−.1987	−.1974	−.2691	605
Postal Service	9.8103	−.0341* (−6.43)	−.0271* (−2.72)	−.0163 (−1.71)	−.0570	−.0341	−.0235	2,985
Securities and Exchange Commission	10.3641	−.0877 (−1.35)	−.1996* (−4.16)	−.3110* (−4.12)	−.0651	−.2353	−.2356	186
Smithsonian Institution	10.0345	−.2431* (−6.04)	−.3108* (−5.04)	−.4210* (−5.50)	−.2437	−.3170	−.4533	273
Veterans Administration	9.8170	−.1116* (−6.47)	−.0861* (−4.84)	−.1302* (−5.82)	−.1456	−.1291	−.1819	1,947

†The statistics for the pooled sample are estimated from a one percent random sample of all federal agencies.
*Significant at the .05 level in a two-tailed test.

to white men and black women, yielding the standardized wage differential between these two groups. This pairwise estimation procedure facilitates comparison between the empirical results in Table 1 and the literature on wage differentials based on race and gender.

The socioeconomic vector Y is composed of education (defined as a vector of 21 dummy variables indicating both the extent and nature of educational attainment[7]); years of experience in the federal sector; years of experience in the nonfederal sector (defined as age minus education minus experience in the federal sector minus 6); quadratics of and an interaction between the two experience variables; region of employment; status as a veteran; and status as physically handicapped.[8] Clearly, the variable for work experience previous to federal employment is likely to be measured with error in the case of women. The CPDF provides no information, however, on the work activities of the federal work force before entering the civil service. The potential biases that arise from measuring gender wage differentials with this experience variable will be discussed in detail below.

Columns 2 through 4 in Table 1 give estimates for $\hat{\beta}_{bm}$, $\hat{\beta}_{wf}$, and $\hat{\beta}_{bf}$, where $\hat{\beta}_{bm}$ is the estimated β for black males; $\hat{\beta}_{wf}$ is the statistic for white females; and $\hat{\beta}_{bf}$ is the statistic for black females. The coefficients in the pooled sample show that, holding skills constant, significant wage differentials by gender and race exist, such that white men earn approximately 12 percent more than black men and 22 to 27 percent more than all females.[9]

The regression technique described by Equation 1 assumes that the difference in earnings between white men and any of the other race-gender groups is caused by a constant shift in the level of the earnings profile among the groups. It is likely, however, that race and gender differences in the payoff to such variables as education and experience also determine the extent of race and gender wage differences. The methodology usually employed to allow for this possibility estimates an earnings function *within* each of the race-gender groups. These regressions are used to decompose the observed wage differential into explained and unexplained parts. Formally, suppose we believe that the white-male and black-male wage structures are different. Since the regression equation passes through the mean, we can write:

$$(2) \qquad \overline{\ln r_{wm}} = \overline{Y}_{wm}\hat{\alpha}_{wm}$$

and

$$\overline{\ln r_{bm}} = \overline{Y}_{bm}\hat{\alpha}_{bm},$$

where the bars denote sample averages, and the "^" denotes the estimated (vector of) coefficients. The observed percentage wage differential is given by:

$$(3) \quad \overline{\ln r_{bm}} - \overline{\ln r_{wm}} = \overline{Y}_{bm}\hat{\alpha}_{bm} - \overline{Y}_{wm}\hat{\alpha}_{wm}.$$

Adding and subtracting a term $\overline{Y}_{bm}\hat{\alpha}_{wm}$ to the right-hand side of Equation 3 yields:

$$(4) \quad \overline{\ln r_{bm}} - \overline{\ln r_{wm}} = (\overline{Y}_{bm} - \overline{Y}_{wm})\,\hat{\alpha}_{wm} + \overline{Y}_{bm}\,(\hat{\alpha}_{bm} - \hat{\alpha}_{wm}).$$

Equation 4 decomposes the observed average wage differential into two parts: one the result of differences in characteris-

[7]The definition of the dummy variables in the education vector is as follows: (1) some elementary school but did not complete; (2) elementary school completed but no high school; (3) some high school but did not graduate; (4) high school graduate or certificate of equivalency; (5) terminal occupational program but did not complete; (6) terminal occupational program completed; (7) some college but less than one year; (8) one year of college; (9) two years of college; (10) associate's degree; (11) three years of college; (12) four years of college; (13) bachelor's degree; (14) post-bachelor's work but no additional higher degrees; (15) first professional degree, such as dentistry, law, or medicine; (16) a degree beyond the first professional degree but no additional higher degree; (17) master's degree; (18) post-master's but no additional higher degree; (19) sixth-year degree, such as an advanced certificate in education, an advanced master's in education, or a certificate of advanced graduate study; (20) post-sixth year but no additional higher degree; (21) doctorate degree, including a doctor of education or doctor of juridical science, and Ph.D.; or (22) post doctoral work.

[8]See Borjas, "Wage Determination," for a discussion of the effect of these variables on the wage structure in the federal bureaucracy.

[9]See Sharon P. Smith, *Equal Pay in the Public Sector: Fact or Fantasy* (Princeton, N.J.: Princeton University Press, 1977) for a comprehensive discussion of how race and gender wage differentials in the federal sector compare with those observed in the private sector.

tics, $(\overline{Y}_{bm} - \overline{Y}_{wm})\,\hat{\alpha}_{wm}$, and one the result of differences in structure, $\overline{Y}_{bm}\,(\hat{\alpha}_{bm} - \hat{\alpha}_{wm})$. The latter part measures the additional earnings the average black man would earn if he were "treated like" a white man. Define:[10]

$$(5)\ \Delta_j = \overline{Y}_j\,(\hat{\alpha}_j - \hat{\alpha}_{wm}),\ j = (bm,\ wf,\ bf).$$

Columns 5 through 7 in Table 1 give the calculated Δs for 30 of the agencies analyzed here.[11] Again, the pooled-sample estimates reveal that white men earn substantially more than the other race-gender groups.

Perhaps the most striking finding in Table 1 is the very large variance in the unexplained wage differentials (whether measured by $\hat{\beta}_j$ or Δ_j) across federal agencies. For example, the wage gap between white men and black men ranges from approximately zero percent (the Department of Agriculture, the Environmental Protection Agency, and the Equal Employment Opportunity Commission) to over 30 percent (the Federal Home Loan Bank Board, the Government Printing Office, the International Communications Agency, and the Interstate Commerce Commission). Similarly, the standardized log wage differential between white men and white women (as measured by Δ_{wf}) ranges from about 3 percent (the Postal Service) to 52 percent (the Government Printing Office). Finally, the standardized log wage differential between white men and black women ranges from 2 percent (the Postal Service) to about 58 percent (the Department of Labor and the Government Printing Office).

These results clearly indicate that inter-agency differences in personnel policies play an important role in the hiring and placement of blacks and women in the federal bureaucracy. An understanding of the causes of such differences requires specifying a behavioral model of the federal government. In a companion paper, these differences are explained by assuming that different federal agencies cater to different constituents and that these constituents' tastes for discrimination vary.[12] In the case of racial discrimination, for example, it is assumed that white (black) constituents prefer to "see" whites (blacks) employed by the federal bureaucracy. To the extent that the racial composition of the constituencies of federal agencies varies across agencies, it is clear that a government that wants to ensure political support will rationally employ blacks, or whites, in those agencies where it is politically advantageous to do so.

In fact, by identifying the agency's constituency as the set of employees of industries affected by the agency's redistribution policies, the empirical analysis in the companion paper revealed that the demand function for blacks and women in federal agencies depends on such factors as the percentage of blacks in the agency's constituency, the percentage of women in the agency's constituency, and the degree to which the agency enforces affirmative action programs in the private sector. These findings, therefore, show that the "unexplained" wage differentials documented in Table 1 can, in fact, be explained by the characteristics of the political sphere in which the agency operates.

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[10] Of course, it is also possible to measure the unexplained wage differential by calculating how much lower the earnings of the average white male would be if he faced the black wage structure. Generally, these two methodologies yield different measures of wage discrimination. The conceptually correct methodology calculates the earnings of a black male if he faced the white male wage structure that would exist in the absence of discrimination. It is likely, however, that if black males are a numerically small minority, they have a minor impact on the white-male wage structure. Thus, the second term in Equation 4 is a good approximation of the true unexplained wage differential.

[11] The decomposition procedure cannot be conducted for the Equal Employment Opportunity Commission, since, of the 257 observations in that sample, only 33 are of white males. This led to a very imprecise estimate of the white-male wage structure. It should also be noted that the explanatory power of the earnings functions summarized in Table 1 is much higher than those reported in the literature. For example, the white-male earnings function in the pooled sample yielded an R^2 of .588. Generally, the R^2 in all the regressions was greater than .5, with the exception of the Postal Service, which consistently showed an R^2 in the .2 to .3 range.

[12] George J. Borjas, "The Politics of Employment Discrimination in the Federal Bureaucracy," *Journal of Law and Economics,* Vol. 25, No. 2 (October 1982), pp. 271 - 99.

Table 2. Relationship Among Alternative Measures of Discrimination (N = 30).
(t-ratios in parentheses)

Dependent Variable	*Independent Variable*				R^2
	Constant	$\hat{\beta}_{bm}$	$\hat{\beta}_{wf}$	$\hat{\beta}_{bf}$	
Δ_{bm}	−.0056 (−.56)	1.1635* (18.72)	—	—	.926
Δ_{wf}	−.0313* (−2.08)	—	1.0592* (17.61)	—	.917
Δ_{bf}	.0052 (.23)	—	—	1.0599* (16.22)	.904

*Significant at the .05 level in a two-tailed test.

The Relationship Among Discrimination Coefficients

Although a strong behavioral assumption can offer important insights into federal personnel policies, the following analysis shows that even in the absence of behavioral assumptions, the results in Table 1 suggest a number of empirical regularities that can expand our understanding of the discrimination process. In particular, the large variance in the βs and Δs across federal agencies provides a unique opportunity to analyze several issues raised in the measurement of discrimination.

Table 1 presents two different methodologies to estimate race and gender wage differentials. One important question is whether there are important differences in the results obtained under the two methods. The empirical literature has not yet answered this question, since analysis is usually limited to estimating $\hat{\beta}_j$ and Δ_j for a single data set. Table 1 presents both of these statistics for 30 different agencies, allowing us to ascertain whether the variation in the conceptually "better" estimate, Δ, is systematically reflected in the simpler statistic, $\hat{\beta}$. This issue can be examined by estimating a regression of the form:

$$\Delta_j = \Theta_0 + \Theta_1\hat{\beta}_j + \epsilon_j, \quad j = (bm, wf, bf), \tag{6}$$

across the 30 federal agencies. The estimated regressions for each of the race-gender groups are presented in Table 2.

The regressions reveal two important findings. First, the simple dummy-variable coefficients $(\hat{\beta}_j)$ systematically underestimate the unexplained wage differential measured by Δ_j. This fact is indicated by the above unity estimates of Θ_1 and by the significantly negative constant term in the white-female regression.[13] Thus, the allowance of differential returns to labor market characteristics (such as education and experience) by race and gender tends to increase the standardized race and gender wage differentials.[14] Second, the regressions show an extremely high correlation between the two measures of unexplained wage gaps; in fact, the correlation coefficient between Δ_j and $\hat{\beta}_j$ exceeds 95 percent for all three race-gender differentials. A crucial result of the analysis, therefore, is that the variation in discrimination coefficients among agencies is captured equally well by Δ_j and $\hat{\beta}_j$.

The discrimination measures reported in Table 1 also allow us to investigate whether the gender wage differential is related to the race wage differential measured by the wage gap between white men and black men. Again, previous empirical studies of

[13]It should be noted, however, that of the three estimates of θ_1, only one (in the black-male regression) is significantly greater than unity.

[14]Joan G. Haworth, James Gwartney, and Charles Haworth obtained a similar finding in "Earnings, Productivity, and Changes in Employment Discrimination During the 1960's," *American Economic Review*, Vol. 65, No. 1 (March 1975), pp. 158-68.

Table 3. Relationship Among Gender and Race Discrimination Coefficients (N = 30).
(t-ratios in parentheses)

Dependent Variable	Independent Variable					R^2
	Constant	Δ_{bm}	$\hat{\beta}_{bm}$	Δ_{wf}	$\hat{\beta}_{wf}$	
Δ_{wf}	−.1766* (−4.48)	.5496* (2.72)	—	—	—	.209
$\hat{\beta}_{wf}$	−.1464* (−4.02)	—	.5581* (2.48)	—	—	.181
Δ_{bf}	−.2134* (−5.41)	.7887* (3.90)	—	—	—	.352
$\hat{\beta}_{bf}$	−.2177* (−5.97)	—	.8137* (3.62)	—	—	.319
Δ_{bf}	−.0546* (−2.35)	.2943* (2.89)	—	.8997* (10.60)	—	.875
$\hat{\beta}_{bf}$	−.0976* (−3.65)	—	.3562* (2.45)	—	.8198* (7.41)	.775

*Significant at the .05 level in a two-tailed test.

race and gender discrimination cannot answer such questions, because they estimate these statistics for only one data set. If we view the federal government as composed of agencies with different personnel policies, we can analyze whether agencies that have large race wage differentials also tend to have large gender wage differentials. Formally, this analysis is based on the regression:[15]

(7) $\Delta_j = \gamma_0 + \gamma_1 \Delta_{bm} + \nu, (j = wf, bf)$
and $\hat{\beta}_j = \gamma'_0 + \gamma'_1 \hat{\beta}_{bm} + \nu', (j = wf, bf)$.

The estimated coefficients are presented in rows 1 through 4 of Table 3. The results are striking. The negative and significant constant terms indicate that the unexplained gender wage differential would be between 15 and 20 percent (relative to white men) even in agencies where no race wage differential exists. Moreover, the positive and significant coefficients of Δ_{bm} and $\hat{\beta}_{bm}$ reveal that agencies that have larger standardized wage differentials based on race will have larger standardized wage differentials based on gender. In fact, the size of the coefficients γ_1 and γ'_1 indicates that a 10 percent fall in the black relative wage decreases the white-female relative wage by 5.5 percent and the black-female relative wage by 8 percent. Thus, Table 3 reveals the interesting finding that there is a strong positive correlation between race and gender wage differentials.

This analysis can be extended by asking a completely unexplored question: to what extent is the low relative wage of black women due to their race, to their sex, or to both? This question can be studied by estimating regressions of the form:

(8) $\Delta_{bf} = \lambda_0 + \lambda_1 \Delta_{bm} + \lambda_2 \Delta_{wf} + \mu$
$\hat{\beta}_{bf} = \lambda'_0 + \lambda'_1 \hat{\beta}_{bm} + \lambda'_2 \hat{\beta}_{wf} + \mu'$.

The estimated regressions are presented in rows 5 through 6 of Table 3, and they also yield interesting results. The relative magnitudes of λ_2 and λ_1 (or λ'_2 and λ'_1) *indicate that the fact that black women are* women is 2.3 to 3 times more important than the fact they are black in causing their low relative wage. A major finding of this

[15]Equation 7 is not derived from any structural model of discriminatory behavior. Rather, its purpose is to measure the correlation among the various race and gender wage differentials. Moreover, since the regressors and the dependent variables in that equation are *themselves estimated with error, it is likely that the* estimator of the coefficient vector γ is biased. The direction of the bias cannot be ascertained, however, unless assumptions are made about the magnitudes of γ_1 and γ'_1.

analysis, therefore, is the conclusion that gender is much more important than race in determining the wage gap between white men and black women.[16]

It is worth pausing at this point to discuss the substantive implications of the findings in Table 3. The positive correlation between race and gender wage differentials suggests that some set of factors leads agencies systematically to pay lower wages to *both* blacks and women. In the context of the political model discussed earlier, this finding implies that common variables exist that enhance both blacks' and women's political value in some federal agencies and not in others; it also suggests that an understanding of the evolution of race and gender wage differentials in the federal bureaucracy could be obtained by a careful analysis of the political forces influencing the agency.

More generally, the results in Table 3 imply that studies of *interfirm* differences in discrimination in the private sector can improve our understanding of discriminatory behavior. For instance, it is well known that the standardized wage differentials $\hat{\beta}_j$ and Δ_j can be interpreted in two extreme ways. They either measure true discrimination or they capture unobserved productivity differences among races and sexes employed by the firm. The positive correlation between race and gender wage differentials documented in Table 3 indicates that if $\hat{\beta}_j$ and Δ_j measure true discrimination, firms that discriminate more than other firms against blacks also discriminate more than others against women. Alternatively, if $\hat{\beta}_j$ and Δ_j measure unobserved productivity differences, the results show that to some extent the same unobserved characteristics are responsible for both race and gender wage differentials.

Note that under either interpretation, the analysis suggests a strategy for further research. In particular, Table 3 proves the existence of agency-specific fixed effects resulting either from employer behavior (such as discrimination) or from technological factors that differentiate blacks and women from other workers. The results in Table 3 therefore show that our understanding of wage discrimination can be greatly expanded by the analysis of differences in the characteristics of firms and of the markets in which firms operate.

Measurement of Labor-Force Experience

One important criticism can be made of the findings in Tables 1 through 3. In particular, how robust are the results with respect to the selection and definition of the independent variables in the earnings function?

For example, the regressions control for years of federal service and not for years of employment in the specific job held by the individual. Similarly, none of the regressions summarized in Table 1 controls for the type of work conducted by the worker. Thus, it could be argued that many of the interagency differences documented in Table 1 (and analyzed in Tables 2 and 3) are the result of differences in occupational characteristics by race and gender across federal agencies.

It should be noted, however, that discrimination, if it exists, can work either through wage differentials within the same job or by placing selected groups of people in particularly low-paying or high-paying occupations. Controlling for occupation in the regressions, therefore, would ignore an important channel in the discrimination process, and the resulting estimates would not measure the full effect of discriminatory behavior in the federal sector. More generally, controlling for variables that can be used by the employer to accelerate or hamper the progress of some employees — some examples being job placement, years on the particular job, and promotion decisions — biases measurement of the full effect of wage discrimination.

A more serious measurement problem is that the labor-force experience of women may be measured with substantial error. In particular, the CPDF provides no information on the work activities of women before

[16]Note that this finding shows that the weights attached to sex discrimination are greater than the weights attached to race discrimination in the determination of the black-female wage. A probable explanation of this result lies in the fact that gender wage differentials are, on the average, greater than race wage differentials in the federal bureaucracy.

their entry into the federal civil service. Recent work on the determinants of female earnings provides strong evidence that the intermittent labor-force participation of married women can explain a substantial portion of the male/female wage differential.[17] Although the CPDF does not give any information on the marital status or the experience of women outside the labor force, it does give data on the number of years of their employment in the federal civil service. This variable can be extremely useful in calculating an upper bound to the bias created by ignoring the intermittent labor-force participation of married women.

In particular, the earnings functions estimated in Table 1 hold constant a measure of "previous" experience *(PREV)* defined as age minus education minus years of government service minus six. This variable has, as expected, a positive effect on the earnings of federal workers. The average value of *PREV* for women is about ten years. Undoubtedly, a part of this time was spent working in other jobs and another part of it was spent outside the labor market. The number of years in the labor market will have a positive effect on current earnings, whereas the nonmarket years may have a negative effect on earnings, because of the depreciation of market skills. The simplest way to take these facts into account is to assume that the two effects exactly outweigh each other, so that women enter federal employment with zero years of labor-force experience. The wage differentials obtained for the various samples after setting *PREV* equal to zero for women are reported in Table 4. The reestimated $\hat{\beta}_k^0$ and Δ_k^0 are somewhat smaller (in absolute value) than the more conventional estimates, $\hat{\beta}_k$ and Δ_k, presented in Table 1.[18]

In fact, we can analyze precisely how the two sets of estimates are related by estimating regressions of the form:

$$(9) \qquad \Delta_k^0 = \delta_0 + \delta_1 \Delta_k + \omega \qquad (k = wf,\ bf),$$

$$\hat{\beta}_k^0 = \delta_0' + \delta_1' \hat{\beta}_k + \omega'.$$

The regressions are presented in Table 5. Note that the constant term is not significantly different from zero in any of the regressions. Moreover, the coefficient of Δ_k or $\hat{\beta}_k$ is positive and significant and lies between .75 and .85. The results thus indicate that controlling for the intermittent labor-force participation of married women is responsible for, at most, 15 to 25 percent of the unexplained wage differentials documented in Table 1.

This estimate is likely to be an upper estimate of the true effect for two reasons. First, the assumption that the negative depreciation effect will cancel out the positive effect of labor-force experience is probably too strong, since there is some debate about whether nonmarket time indeed has a negative effect on current earnings.[19] Second, the estimates suffer from selectivity bias, since only women currently working are included in the analysis. Suppose there is a positive correlation over time in a woman's probability of participating in the labor force, such that women currently working are more likely to have worked in the past. This assumption implies that women who are currently working will have an above average number of years of labor market experience before starting federal employment.[20]

The results in Table 5 provide strong evidence, therefore, that the bias introduced by the intermittent labor-force participation of married women is not very

[17]Jacob Mincer and Solomon Polachek, "Family Investments in Human Capital: Earnings of Women," *Journal of Political Economy,* Vol. 82, No. 2, Part 2 (March/April 1974), pp.S76 - S108.

[18]This is not surprising, since assuming that women have no experience is likely to overstate the actual differences in labor market experience and hence reduce the unexplained wage differentials.

[19]See, for example, Steven H. Sandell and David Shapiro, "The Theory of Human Capital and the Earnings of Women: A Reexamination of the Evidence," *Journal of Human Resources,* Vol. 13, No. 1 (Winter 1978), pp.103 - 17.

[20]James J. Heckman and Robert J. Willis provide empirical evidence of the serial correlation in the probability of labor-force participation in "A Beta-Logistic Model for the Analysis of Sequential Labor Force Participation by Married Women," *Journal of Political Economy,* Vol. 85, No. 1 (February 1977), pp.27 - 58.

Table 4. Measurement of Gender Wage Differentials Assuming No Previous Experience. (t-ratios in parentheses)

Agency	$\hat{\beta}^0_{wf}$	$\hat{\beta}^0_{bf}$	Δ^0_{wf}	Δ^0_{bf}
Pooled†	−.1946* (−31.82)	−.2487* (−27.43)	−.2035	−.2552
Cabinet Agencies				
Agriculture	−.1583* (−6.06)	−.1960* (−4.46)	−.1406	−.2032
Commerce	−.1636* (−10.56)	−.2559* (−13.66)	−.1834	−.2674
Defense	−.2776* (−31.87)	−.3349* (−27.69)	−.2714	−.3328
Energy	−.2670* (−11.69)	−.3165* (−10.28)	−.2680	−.3364
Health, Education, and Welfare	−.1012* (−4.14)	−.2006 (−6.62)	−.1363	−.2122
Housing and Urban Development	−.2591* (−9.75)	−.3790* (−13.56)	−.2905	−.4420
Interior	−.1543* (−14.58)	−.2041* (−14.73)	−.1614	−.2194
Justice	−.1723* (−4.85)	−.2144* (−4.93)	−.2029	−.2223
Labor	−.2177* (−10.05)	−.4228* (−17.25)	−.2680	−.4702
State	−.1197* (−3.58)	−.2307* (−5.80)	−.0892	−.1852
Transportation	−.4618* (−10.96)	−.4989* (−7.63)	−.4599	−.5035
Treasury	−.2338* (−8.92)	−.3792* (−11.31)	−.4124	−.3288
Independent Agencies				
Environmental Protection Agency	−.0738* (−3.04)	−.2123* (−6.53)	−.0910	−.2305
Equal Employment Opportunity Commission	−.0330 (−0.23)	−.1076 (−0.99)	—	—
Federal Communications Commission	−.1170 (−1.95)	−.2759* (−3.67)	−.1081	−.2040
Federal Deposit Insurance Corporation	−.1798* (−4.28)	−.2562* (−3.83)	−.1257	−.2610
Federal Home Loan Bank Board	−.2922* (−3.03)	−.4431* (−4.30)	−.4057	−.5430
Federal Trade Commission	.0017 (0.02)	−.1081 (−1.02)	−.0349	−.1550
General Accounting Office	−.0736* (−2.47)	−.2099* (−4.68)	−.0673	−.0867
General Services Administration	−.1785* (−9.22)	−.2782* (−13.53)	−.1807	−.2934
Government Printing Office	−.3688* (−7.82)	−.4683* (−13.15)	−.3911	−.4549
International Communications Agency	−.1008 (−1.84)	−.3029* (−5.26)	−.1331	−.3200
Interstate Commerce Commission	−.0702 (−0.86)	−.4355* (−4.91)	−.1915	−.3819

Table 4 Continued.

Agency	$\hat{\beta}^0_{wf}$	$\hat{\beta}^0_{bf}$	Δ^0_{wf}	Δ^0_{bf}
National Aeronautics and Space Administration	−.2409* (−15.98)	−.2615* (−11.35)	−.2482	−.2636
National Labor Relations Board	−.1030* (−2.52)	−.2436* (−4.00)	−.1561	−.2141
Nuclear Regulatory Commission	−.0836 (−1.23)	−.3015* (−3.17)	−.1407	−.3227
Office of Personnel Management	−.1456* (−3.42)	−.1878* (−3.78)	−.1239	−.2077
Postal Service	−.0444* (−4.18)	−.0254* (−2.48)	−.0464	−.0300
Securities and Exchange Commission	−.0047 (−0.09)	−.1379 (−1.74)	−.0555	−.0703
Smithsonian Institute	−.1715* (−2.38)	−.2677* (−2.84)	−.1590	−.2770
Veterans Administration	−.0394 (−1.77)	−.0803* (−3.01)	−.0535	−.1128

†The statistics for the pooled sample are estimated from a one percent random sample of all federal agencies.
*Significant at the .05 level in a two-tailed test.

Table 5. Relationship Between $\Delta_k(\hat{\beta}_k)$ and $\Delta^0_k(\hat{\beta}^0_k)$ (N = 30).
(t-ratios in parentheses)

Dependent Variable	*Independent Variable*					R^2
	Constant	Δ_{wf}	Δ_{bf}	$\hat{\beta}_{wf}$	$\hat{\beta}_{bf}$	
Δ^0_{wf}	.0257 (.85)	.7835* (7.61)	—	—	—	.674
Δ^0_{bf}	.0208 (.87)	—	.8491* (13.17)	—	—	.861
$\hat{\beta}^0_{wf}$	.0129 (.53)	—	—	.7843* (8.14)	—	.703
$\hat{\beta}^0_{bf}$	.0025 (.09)	—	—	—	.8285* (10.94)	.811

*Significant at the .05 level in a two-tailed test.

important in the CPDF data. Moreover, the strong positive correlation between race and gender wage differentials is not affected by the correction used in Table 5. In particular, Table 6 replicates the analysis in Table 3 by relating the gender wage differentials as measured by Δ^0_j and $\hat{\beta}^0_j$ to the wage gap between black men and white men. As in Table 3, the most important result is the strong positive correlation among the various gender and race wage differentials. In other words, a common factor is partly responsible for the low relative wages of both women and blacks.

Summary

The empirical analysis of interagency race and gender wage differentials in the federal bureaucracy yields several important findings. First, the use of a simple dummy variable to indicate race and gender in the regressions leads to downward-biased measures of the standardized wage differential. Second, federal agencies that have larger race differentials are likely to have larger gender differentials. Third, the low relative wage of black women is more the result of the fact

Table 6. Relationship Among Gender and Race Discrimination Coefficients, Assuming No Previous Experience (N = 30).
(t-ratios in parentheses)

Dependent Variable	*Independent Variable*					
	Constant	Δ_{bm}	$\hat{\beta}_{bm}$	Δ^0_{wf}	$\hat{\beta}^0_{wf}$	R^2
Δ^0_{wf}	−.1147* (−2.91)	.4180* (2.07)	—	—	—	.133
$\hat{\beta}^0_{wf}$	−.1135* (−3.15)	—	.3542 (1.59)	—	—	.083
Δ^0_{bf}	−.1616* (−4.29)	.6626* (3.44)	—	—	—	.297
$\hat{\beta}^0_{bf}$	−.1774* (−5.08)	—	.6776* (3.14)	—	—	.261
Δ^0_{bf}	−.0701* (−2.91)	.3293* (2.83)	—	.7972* (7.86)	—	.786
$\hat{\beta}^0_{bf}$	−.0943* (−3.47)	—	.4181* (2.78)	—	.7325* (5.99)	.683

*Significant at the .05 level in a two-tailed test.

they are women than the fact they are black. Finally, the intermittent labor-force participation of women is responsible for no more than 25 percent of the unexplained gender wage differential in the federal sector.

It is worthwhile to emphasize an important implication of these findings for understanding federal personnel policies and regulations. In particular, as with any empirical study of race and gender wage differentials, there is no unambiguous, ironclad proof that discrimination exists, since, of course, wage differentials can always be dismissed as resulting from unobserved skill differentials. Nonetheless, the empirical finding of a strong positive correlation between race and gender wage differentials across federal agencies implies the existence of a common set of factors responsible for the wage "discrimination coefficients." That finding opens up a new line of research in discrimination analysis: the identification, *at the firm level,* of variables that account for interfirm differences in employment policies regarding both minorities and women.

More generally, the type of research conducted in this paper suggests that similar studies of interfirm differences in the relative wages of the races and the sexes in the private sector may lead to extremely useful insights. In particular, if interfirm variation in these variables could be explained by such factors as the racial (gender) composition of the firm's customers, the degree of physical proximity between the workers and the customers, and the race or sex of the firm's ownership or management, the empirical analysis of wage discrimination will have advanced substantially from its basically descriptive state to a study of the determinants of discriminatory behavior.