Sex Discrimination in Faculty Salaries: Evidence from a Historically Women’s University

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Sex Discrimination in Faculty Salaries:
Evidence from a Historically Women’s University

By Barry T. Hirsch and Karen Leppel*

During the past several years, a number of studies have appeared in this Review and elsewhere examining the earnings and promotion of women faculty.1 A finding common to most of these studies is that the female-male salary differential is smallest at entry level, but widens over time.2 Indeed, George Johnson and Frank Stafford state, after presenting such evidence for Michigan State University: “We believe this qualitative result would be observed for any university in the United States for which the sample size is sufficiently large” (1974, p. 899).

While the qualitative findings of faculty salary studies have been similar, interpretations have differed. On the one hand, the human capital view attributes the widening sex differential, or flatter female earnings profile, to differences in acquired skill and productivity. This literature emphasizes differences in continuous labor market experience, hours of work, and the relative teaching/research division of labor. On the other hand, the discrimination view attributes the widening differential to increased labor market discrimination with respect to experience. The sex differential is smallest at entry level where universities must compete and pay prevailing salaries to attract incoming faculty members. However, discrimination is more easily exercised in the internal university labor market by male faculty and administrators as job mobility lessens with age (because of fixed costs, “tied” moves,3 university-specific job training, a shorter benefit span, etc). The strongest evidence supporting the discrimination view is the existence of significant unexplained salary differentials even where detailed data exist on research and teaching performance.

The purpose of this note is to examine the salary structure at a large university which was historically the state’s university for women, but is now fully coeducational. To our knowledge, no other such study is available. Comparison of results from such a study with those already in the literature can shed light on the interpretation of salary differential studies. We find, somewhat surprisingly, an exception to the finding by Johnson and Stafford and others of a widening female-male salary differential with experience. Apart from some qualifications discussed below, we find a small differential at entry, but little difference in the reward structure to men and women with respect to experience, ceteris paribus. Alternative interpretations of this evidence are provided. While we cannot clearly test between the human capital and discrimination explanations for salary differentials, our evidence strongly suggests that universities can and do exercise significant discretion in the awarding of salaries, discretion which is presumably made possible by the not-for-profit nature of these institutions. (See Armen Alchian and Reuben Kessel.)

Section I briefly describes the data source, while Section II examines specification and then tests for functional form. Section III

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*University of North Carolina-Greensboro. We thank Terry G. Seaks and Frank Stafford for helpful suggestions. We are equally responsible for the contents of the paper.

1A partial list includes George Johnson and Frank Stafford, with later comments by Steven Farber and by Myra Strober and Alice Quester; Nancy Gordon, Thomas Morton, and Ina Braden; Emily Hoffman; David Katz; Marianne Ferber and Jane Loeb; Ferber and Betty Kordick; Ferber, Loeb, and Helen Lowry; and James Koch and John Chizmar.

2There is debate over whether the differential begins to narrow late during the working life, as argued by Johnson and Stafford. See Johnson and Stafford (1977) and Ferber, Loeb, and Lowry.

3Viewing mobility within a household context (see Jacob Mincer) the potential for discrimination against married women faculty may be higher than with male or unmarried female faculty.
analyzes the empirical results, with conclusions following in Section IV.

I. Data

Our sample includes 487 full-time instructional faculty in tenure-track positions at the university during the 1980–81 academic year. Women comprise one-third of the sample. The school, previously the state university for women, has been coeducational since the early 1960’s. It currently enrolls 10,000 students, 69 percent of whom are female and 28 percent graduate students. The university, while currently being similar in structure and role to institutions studied previously, has historically been a “women’s college” and women faculty and administrators have played, and continue to play, an important role in decision making (the proportion of the faculty that is female did decline somewhat during the 1970’s). Thus, we believe our data provide a good test of whether university decision makers can and do exercise discretion with respect to sex in the awarding of faculty salaries.

The average salary is $24,771 for all faculty, $22,899 for women, and $25,766 for men. While the overall female–male (F-M) earnings ratio is .89, the earnings differential is generally smaller or nonexistent within rank or experience subclases. For instance, F-M is .96, .99, 1.01, and .94 for instructors, assistants, associates, and full professors, respectively. Within experience groupings, where EXP represents the number of years since completion of last degree, F-M is .88 for EXP through 6 years, .97 for EXP between 7 and 11 years, .92 for EXP between 12 and 18 years, and .87 for EXP greater than 18 years. As noted below, the widening salary differential among older faculty results not only from lower rewards to experience per se, but from differential rates of promotion (which can, of course, result from discrimination), a lower percentage of women with terminal degrees, and relatively fewer women who have held administrative positions. Among male (female) faculty with more than 18 years experience, 78 (67) percent hold terminal degrees, 71 (44) percent are full professors, and 29 (22) percent have held administrative positions.

II. Specification and Test for Functional Form

Previous studies have attempted to measure discrimination by estimation of functions relating earnings to productivity-related characteristics. Most of the studies cited above have used a semi-log specification, though many by economists and most by noneconomists have used a linear equation. The semi-log earnings function has substantial advantages; in addition to its theoretical underpinnings from the human capital framework, it facilitates examination of percentage differentials and allows comparisons across studies using data with different salary and price levels. Surprisingly, none of these studies have tested empirically for functional form, even though the Box-Cox test is suited for such a test.4

The most general form of the semi-log earnings function is

\[ \ln E_i = a + \sum_j \beta_{ij} X_{ij} + u_i, \]

while that for the linear is

\[ E_i = a + \sum_j b_{ij} X_{ij} + v_i, \]

where \( E_i \) represents the salary of faculty member \( i \), \( X_j \) is a vector of earnings determinants, and \( \beta_{ij} \) and \( b_{ij} \) are vectors of parameters which can vary by sex \( S \). The error terms \( u_i \) and \( v_i \) are assumed to be normal random variables with zero means and constant variances.

The vector \( X_j \) includes variables similar to those used in previous studies. Experience, \( EXP \), is measured by the number of years since completion of last degree. Neither time at the university (job tenure) nor age showed a highly significant effect on earnings, independent of experience.5 TERMDEG is a di-

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4See G. Box and D. Cox. James Heckman and Solomon Polachek have applied the Box-Cox test to earnings functions from a national sample.

5Job tenure had a negative, but insignificant, partial effect on earnings, reflecting higher salaries, ceteris paribus, for those incoming faculty with greater mobility and quality. Age had a very small, but significant positive effect. We excluded it, since its high correlation with \( EXP(\rho = .81) \) would make interpretation of the earnings-experience profile difficult.
chotomous variable equal to 1 for faculty members holding the highest attainable degree normally acquired in their field.° ADMIN equals 1 if a faculty member currently holds or previously held any administrative appointment.® The effects of rank are estimated using a set of dichotomous variables (ASST, ASSOC, FULL) with instructors as the reference group.® An additional set of dummy variables are included for eleven department groups or schools (results with respect to sex are similar using a finer thirty-two department breakdown).® There was not a sufficient number of nonwhite faculty to examine racial differences. No direct measures of research or teaching quantity/quality were available, apart from that measured by rank.

We first use the Box-Cox test for functional form to compare alternative specifications of (1) and (2). Initially, we restrict parameters β_j and b_j to be identical for men and women, except for an intercept dummy; that is, we compare

\[ (1a) \quad \ln E_i = \alpha + \beta_j X_{ij} + \gamma FEM_i + u_i; \]
\[ (2a) \quad E_i = a + b_j X_{ij} + \psi FEM_i + v_i, \]

where \( FEM = 1 \) if faculty member \( i \) is female. Search for the “optimal” transformation is pursued via application of the Box-Cox procedure, where the dependent variable is transformed by the operator \( \lambda \) to \( (E_i^\lambda - 1) / \lambda. \)

When \( \lambda = 0 \) the specification is semi-log; when \( \lambda = 1 \) the function is linear. The optimal value \( \hat{\lambda} \), determined by calculating the maximized log-likelihood function for alternative values of \( \lambda \), is found to be .06, extremely close to zero. The statistic \( 2[L_{\text{max}}(\hat{\lambda}) - L_{\text{max}}(\lambda)] \) is distributed \( \chi^2 \) with 1 df. The relevant \( \chi^2 \) value at the .05 significance level is 3.84. We obtain a value of \( \chi^2 = 83.32 \) for \( H_0: \lambda = 1 \), allowing us to easily reject the linear specification. A value of \( \chi^2 = .36 \) is obtained for \( H_0: \lambda = 0 \); thus, the hypothesis that the semi-log transformation is optimal cannot be rejected.

In addition, we test functional form separately for male and female earnings functions (1) and (2), allowing all \( \beta_j \) and \( b_j \) to vary by sex. Again, the semi-log specification is found to be highly preferable to the linear. Searching by increments of .01, for the male earnings function the optimal \( \hat{\lambda} = 0.00 \), while for the female function, \( \hat{\lambda} = .27 \). In both cases, \( H_0: \lambda = 1 \) is decisively rejected, while \( H_0: \lambda = 0 \) cannot be.

### III. Empirical Results

How large is the female-male salary differential, after accounting for differences in experience, rank, department, terminal degree, and administrative appointment? Table 1 presents regression results from alternative specifications. The unexplained salary differential also can be measured in several ways. Estimation of (1a) produces a coefficient on \( FEM \) indicating an unexplained differential of 3.8 percent (\( t \)-statistic equals \( -3.25 \)). Exclusion of rank, as suggested by Hoffman, produces an estimate of 4.3 percent (\( -2.93 \)). Exclusion of department areas, but inclusion of rank, results in an estimated 3.0 percent (\( -2.67 \)) differential, indicating that the sex differential at this university cannot be explained by the presence of women in lower paying fields (these results not

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°A Ph.D. or Ed.D. is the terminal degree in all fields at the university other than nursing. Omission of nursing from our analysis did not substantially affect any of our results.

°°We included past as well as current administrative positions in constructing ADMIN since any salary compensation is spread over several years. Daniel Saks has examined the long-run effect of being a department head on salary at Michigan State.

°°°We compare these results with those where rank is excluded since differential rates of promotion are one source of discrimination. Lacking direct data on performance or quality, we believe the earnings functions including rank are generally preferable. Nine professors with endowed chairs were excluded from the sample since there were no women in name chairs and relevant comparisons would be impossible. Whether the lack of women in these positions is the result of discrimination could not be determined.

°Estimation of separate regressions by department/school groups indicated significant differences in intercepts; however, we could not reject the hypothesis of equal slopes.

10We implement the Box-Cox test using a method suggested by Terry Seaks.

11Following Peter Kennedy, the percentage sex differential is estimated by \( \{ \exp(\gamma - 1/2 \Var(\gamma)) - 1 \} \cdot 100. \)
Table 1—Determinants of Academic Salaries: Regression Results

<table>
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<tr>
<th></th>
<th>All Faculty</th>
<th>All Faculty</th>
<th>All Faculty</th>
<th>Males</th>
<th>Females</th>
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<td>.019</td>
<td>.039</td>
<td>.018</td>
<td>.021</td>
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<td>(.002)</td>
<td>(.002)</td>
<td>(.002)</td>
<td>(.004)</td>
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<tr>
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<td>(.00007)</td>
<td>(.00011)</td>
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<tr>
<td>FEM</td>
<td>-</td>
<td>-.039</td>
<td>-044</td>
<td>-</td>
<td>-</td>
</tr>
<tr>
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<td>(.012)</td>
<td>(.015)</td>
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<td>.039</td>
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<td>(.015)</td>
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<td>(.029)</td>
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<td>(.018)</td>
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<td>(.029)</td>
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<td>-</td>
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<td>(.028)</td>
<td>(.028)</td>
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<td>(.047)</td>
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<td>(.024)</td>
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<td>(.039)</td>
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<td>-</td>
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<td>(.020)</td>
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<td>(.028)</td>
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<td>included</td>
<td>included</td>
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<tr>
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<td>.875</td>
<td>.780</td>
<td>.877</td>
<td>.875</td>
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<tr>
<td>S. E. E.</td>
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<td>.100</td>
<td>.133</td>
<td>.099</td>
<td>.101</td>
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<tr>
<td>n</td>
<td>487</td>
<td>487</td>
<td>487</td>
<td>318</td>
<td>169</td>
</tr>
</tbody>
</table>

Note: Dependent variable is natural logarithm of salary (9-month basis). Standard errors are in parentheses. The department/school groups are Business and Economics; Education; Health, Physical Education, and Recreation; Home Economics; Music; Nursing; Natural Sciences; Humanities; Languages; Social Sciences; and Fine Arts (reference group).

shown). Alternatively, following Ronald Oaxaca, we can measure the unexplained differential from estimation of separate male and female earnings functions. Using the male coefficients characteristics we obtain a 2.7 percent unexplained differential, while using the female coefficients and male characteristics we obtain a 4.8 percent estimate of salary discrimination.

In order to compare the shapes of the earnings profiles for men and women faculty, we follow Johnson and Stafford and include interaction terms FEM·EXP and FEM·EXP^2, continuing to control for department/school, terminal degree, and administrative position. Whereas Johnson and Stafford found a significantly flatter and less-concave profile for women, we find an opposite pattern, though differences by sex are not generally significant. Below are partial regression results; in addition to the other control variables, equation (3) includes rank while (4) does not (standard errors in parentheses):

$$
(3) \quad \ln E = 0.019 \cdot EXP - 0.00027 \cdot EXP^2
\quad (0.0002) \quad (0.0006)
- 0.043 \cdot FEM + 0.002 \cdot FEM \cdot EXP
\quad (0.022) \quad (0.003)
- 0.00009 \cdot FEM \cdot EXP^2;
\quad (0.00011)
$$

$$
(4) \quad \ln E = 0.038 \cdot EXP - 0.00052 \cdot EXP^2
\quad (0.003) \quad (0.0008)
- 0.040 \cdot FEM + 0.004 \cdot FEM \cdot EXP
\quad (0.029) \quad (0.004)
- 0.00025 \cdot FEM \cdot EXP^2.
\quad (0.00014)
$$

12Oaxaca’s measure of discrimination is \( D = (\text{Wm}/\text{Wf} - (\text{Wm}/\text{Wf})^{10}/(\text{Wm}/\text{Wf})^0 \), where \( \text{Wm}/\text{Wf} \) is the actual male-female earnings ratio, and \( (\text{Wm}/\text{Wf})^0 \) is the ratio that would exist in the absence of discrimination. The latter is calculated in two different ways. The first method is based on the assumption that the non-discriminatory salary structure would be that which is observed for men, while the second employs the salary structure that is observed for women.
Unlike Johnson and Stafford, who argue sex differentials first widen with experience, but may narrow in later years, we find that women faculty at this university receive rewards from experience during early years at least as large as men, but that the differential widens somewhat during later years. For instance, coefficient estimates from (3) indicate that male earnings peak at 35.2 years of experience, while women's peak at 29.2 years (the corresponding estimates from (4) are 36.5 and 27.3 years, respectively).

For precise comparison to Johnson and Stafford we reestimated their equation, (p. 899) including faculty from only six fields (sociology is the reference group) and excluding instructors (and one additional non-Ph.D.):

\[
\ln E = 9.735 + .040 \text{ EXP} - .0004 \text{ EXP}^2 \\
(0.008) \\
- .072 \text{ FEM} + .030 \text{ FEM} \cdot \text{EXP} \\
(0.083) \\
- .0012 \text{ FEM} \cdot \text{EXP}^2 + .200 \text{ ECON} \\
(0.0004) \\
+ .016 \text{ ANTHRO} + .080 \text{ PHYS} \\
(0.057) \\
+ .035 \text{ MATH} - .034 \text{ BIO} \\
(0.049)
\]

\[R^2 = .797; \, S.E.E. = .112; \, n = 67.\]

Again we obtain results suggesting that women faculty at this university have a lower but more concave earnings profile than men.

An alternative way of comparing the earnings determination process for male and female faculty is to test for equality of coefficients across earnings function (1) estimated separately for men and women. We test jointly and separately for both differences in intercepts and in slopes, using both standard F tests for linear restrictions and a weaker MSE test developed by T. D. Wallace.\(^{13}\) The Wallace test uses an average squared distance criterion in order to capture tradeoffs between bias and variance. The calculated \(F\)-statistic comparing the residual sum of squares between the restricted (pooled) and unrestricted (separate) earnings functions is \(F_{(18,451)} = 1.98\), which allows us to reject the hypothesis of equality at the .05 significance level using the standard Chow test (critical \(F_{(0.05,18,451)} = 1.63\)), but not using the weaker Wallace test (critical \(F_{(0.05,18,451)} = 3.07\)).

If we separately test for equality of the intercepts and slopes we find that the significant differences between male and female earnings structure are due to differences in intercepts and not in rewards to earnings determinants (slopes). The \(F\)-statistic for the test of unequal slopes is \(F_{(17,451)} = 1.41\), which implies that we cannot reject the hypothesis of equal slopes using either the standard or Wallace criterion. If we then test for unequal intercepts, maintaining that slopes are equal, \(F_{(4,468)} = 11.424\) (the square of the \(t\)-statistic on the FEM coefficient), allowing us to reject easily the hypothesis of equal intercepts using either criterion. Thus, we are able to conclude that earnings differences between male and female faculty at this university stem primarily from differences in entry salaries (intercepts), but that the reward structure within the university to experience and other earnings determinants is similar.\(^{14}\)

Other results from our model are similar to those obtained in previous studies. Possession of the terminal degree increases salary by about 15 percent, holding rank constant. When rank is excluded the terminal degree increases salary by approximately 36 percent since possession of the degree accelerates (makes possible) promotion. Women who have held administrative positions earn about 4 percent more than those who have not, while the comparable figure for men is about 13 percent. The reason for this difference is unclear. Rewards by rank are similar to those in other studies. Relative to instructors, the salaries of assistant, associate, and full professors are about 16, 35, and 60 percent.

\(^{13}\)See Gregory Chow, Wallace, and James Goodnight and Wallace.

\(^{14}\)Stafford suggested that this entire school may have a "female salary structure" since the experience coefficients estimated here are somewhat lower than those estimated by Johnson and himself.
higher, respectively. The rewards to female faculty from promotion are slightly higher than for men, though rates of promotion are somewhat slower. Among department/school groupings the lowest salaries, ceteris paribus, are in the languages and humanities. Relative to these groups, faculty in natural sciences earn about 2 percent more; in the social sciences, nursing, and fine arts, about 8 percent more; in music and health, physical education, and recreation, about 11 percent more; in education and home economics, 13 percent more; and in business and economics, approximately 23 percent more.

IV. Conclusion

Previous studies have found significant unexplained salary differences between male and female faculty, differences which are smallest at entry but widen with experience. In this study, we have examined the salary structure at a large state university which was historically the state university for women but is now fully coeducational. If the salary structure at such a university was found to be similar to that elsewhere, it would indicate either than such differentials reflect true productivity differences (the human capital view), or that this university has a similar “taste” for discrimination and that the not-for-profit nature of the institution allows it to be exercised (the discrimination view), or that the university is an efficient cost minimizer which pays no more than market wages (this would imply that it should hire and retain a predominately “low-priced” female faculty).

Rather than finding a salary structure identical to other universities, however, we found significant differences. Whereas Johnson and Stafford, as well as others, have found that women faculty have earnings profiles which are lower, flatter, and less concave than men, we find that women at this university have profiles which are lower, initially steeper, and more concave than men. Ceteris paribus, women are paid less than men; however, this reflects primarily differences in market salary levels and not significant differences in treatment once at the university. This evidence, like that obtained from other studies, is consistent with alternative explanations. If the human capital explanation for the commonly observed sex differential is correct, then it implies that this university has chosen to treat male and female faculty similarly (at least until later years) despite real differences in productivity. On the other hand, if the discrimination view is correct, it suggests that sex discrimination is lower at an institution which historically has been a university for women.15 Regardless of the explanation (and, of course, the alternatives are not mutually exclusive), our evidence implies that this university need not, and does not, pay salaries which fully reflect existing market differentials (i.e., it does not fully price discriminate in order to minimize the wage bill for a given quality and sex composition of the faculty). Rather, we find that the university, by virtue of its not-for-profit character, is able and willing to exercise discretion in the awarding of salaries by sex.

15We have not addressed directly the possibility that salary differentials have been significantly affected by affirmative action. The university drafted an affirmative action plan in 1974 which presented nonbinding employment goals. If affirmative action were to explain the observed salary structure, we would expect to see the narrowest (or no) differentials for young faculty, but remaining differentials for other faculty. That we observe the smallest differentials for middle-age faculty (7 ≤ EXP ≤ 11) appears inconsistent with such an explanation.

REFERENCES


