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The Existence of Gender-Specific Promotion Standards in the U.S.

Kathy A. Paulson Gjerde*

College of Business Administration, Butler University, Indianapolis, IN 46208-3485, USA

This paper is motivated by the claim that promotion probabilities are lower for women than men. Using data from the 1984 and 1989 National Longitudinal Youth Surveys, this paper tests this claim and two related hypotheses concerning training and ability. It is found that females are less likely to be promoted than males, and females receive less training than males. The relationship between promotion and gender varies across occupations, however, suggesting that the alleged glass ceiling faced by women and other minorities in the workplace is not uniform across all labor markets. Copyright © 2002 John Wiley & Sons, Ltd.

INTRODUCTION

Over the last few decades, labor economists have generated an extensive literature that explores the existence of wage differentials in the modern internal labor market. Not only do wage differentials exist between union and non-union workers, but there is evidence of public and private sector earnings differentials, as well as female and male differentials.[1] Although the existence of such differentials is well documented, the source of these differences remains a matter of debate.[2]

Focusing on the male–female earnings gap, one possible explanation is that men typically hold higher paying type jobs than women. For example, according to a 1991 survey, 'Fully 31% of full-time working women are in lower-paying administrative support jobs, compared with 6% of men. Furthermore, men are more likely to be in higher-paying skilled blue-collar trades (20% of men versus 3% of women)' (Crispel, 1991). Although this matching may take place before, during, or after the hiring stage, recent attention has centered on the existence of sorting after the hiring stage in the form of different promotion standards for men and women. Specifically, some argue that women face a 'glass ceiling' which blocks their upward mobility, at least in certain occupations.

As a result, the government has addressed this issue in a number of ways. For instance, the Glass Ceiling Act of 1991 proposed 'to create a 21-member commission to recommend ways to break through the 'glass ceiling' that has kept women and minorities for the most part out of top management jobs' (Skrzycki and Swoboda, 1991). The presence of a glass ceiling has become not only a congressional issue but also has received considerable attention from the Department of Labor. In August, 1991, Secretary Lynn Martin indicated that the Department of Labor 'would move ahead with a new, but long-simmering, plan to push employers to remove cultural and other barriers to promotion for female and minority executives' (Karr, 1991). The question remains, however, as to whether or not women and other minorities truly face a glass ceiling in the workplace in all settings or occupations.

Most empirical tests of the existence of a glass ceiling in the economics literature have been based on data from a single firm, organization, or profession. For example, evidence of an 'invisible ceiling' faced by females is found within a large Canadian service sector company (Cannings, 1988; Cannings and Montmarquette, 1991), within a

*Correspondence to: College of Business Administration, Butler University, 4600 Sunset Avenue, Indianapolis, IN 46208-3485, U.S.A.; E-mail: kpaulson@butler.edu

large high-tech Japanese firm (Ariga *et al.*, 1999), and within a medium-sized financial services firm (Paulin and Mellor, 1996). A similar result is found in the context of the legal (Spurr, 1990) and public school administration (Joy, 1998) professions. Although these studies shed light on the issue of gender-specific promotion standards, the majority are limited in scope and, as such, the robustness of their results questionable. A more comprehensive study across multiple labor markets is needed to understand the true scope and nature of the problem. This paper represents a first step toward determining how the relationship between gender and promotion varies across occupations.

Given the current emphasis on this issue, it is not surprising that a theoretical model of promotion capable of explaining this phenomenon has been developed in the economics literature. Lazear and Rosen (1990) present a model of gender-specific promotion standards which is consistent with the existence of a glass ceiling. According to Lazear and Rosen, a prime cause of the gender wage differential is different promotion rates for men and women. In support of their position, the authors state, 'For example, much of the data supplied by firms in job discrimination litigation appears to show that women have smaller probabilities of promotion into high-paying jobs then men of similar characteristics.' However, evidence presented in job discrimination litigation may also be a biased sample of the industry data. In this paper, in contrast, a more 'random' sample of the labor market is used in order to determine whether or not promotion probabilities differ by gender across occupations. Equally important, the hypotheses tested have a theoretical economic foundation.

In the following section, the Lazear and Rosen model, henceforth referred to as the LR model, is summarized and its results identified. The next two sections test the predictions of that model using data from the 1984 and 1989 National Longitudinal Youth Surveys. In the last section conclusions are presented.

THE LR MODEL OF JOBS

Lazear and Rosen consider a three-period model of employment.[3] In period 1, an individual works and is reviewed. On the basis of this review, the worker is reassigned to job A, a high productivity job, or to job B, a low productivity job. In period 2, the worker remains with the firm in his or her assigned position, but, in period 3, he or she has the option to quit. Interfirm mobility is ignored, such that in period 3 the worker's quit decision is whether to stay in or exit from the labor force.[4]

Each worker has an endowment ability level, δ, which influences his or her productivity. This ability level is revealed perfectly to everyone after one period of work. The output of a worker assigned to job A is δ in period 1, $\delta\tau_2$ in period 2, and $\delta\tau_3$ in period 3, where $\tau_2 < 1 < \tau_3$. In contrast, the output of a worker assigned to job B is δ in all periods. Thus, job A involves 'learning' and yields a higher future output, but there is an investment cost to assigning a worker to job A.

Workers are characterized by three variables: gender, market ability, and non-market alternative value of time, ω, where ω is a random variable, perfectly revealed to both the worker and the firm at the beginning of the third period. It is assumed that both male and female workers have the same distribution of labor market ability. However, the distribution function of ω for women stochastically dominates the distribution function of ω for men, i.e. $F_m(\omega) > F_f(\omega)$ for $\omega > 0$. In other words, from a social output perspective, women have a higher expected productivity in the home than men in the third period. Support for this assumption can be found in Gronau (1973a, b). Thus, men and women differ not in their market ability but in their non-market ability.[5]

Given these assumptions, the employer chooses a job assignment rule and pay scale that maximizes worker utility subject to a profit constraint. Lazear and Rosen show that the socially efficient promotion scheme involves promoting workers whose ability level is greater than some threshold ability level, δ^*, and paying workers their marginal product. Moreover, this efficient promotion policy dictates that the threshold ability standard for promoting a woman be higher than that for a man. The intuition behind this result is that women, being more productive in the home on average than men in the third period, are more likely to leave the firm. In other words, women have both a comparative and absolute advantage in the home from a social productivity perspective. Since a departure from the firm entails a greater social cost if the worker is in job A than if he or she is in job B, men are more likely than women to be promoted to job A in period 2. Thus, it is female workers' lack of attachment to the labor force that induces the employer to implement a promotion scheme resulting in occupational segregation. Empirical evidence of this lack of attachment is documented by Jones and Makepeace (1996).

Based on this somewhat limited model of jobs, the authors identify three empirically testable results. First, the male–female wage differential can be explained by the fact that women are less than proportionately represented in higher-paying jobs. That is, women are promoted less frequently than men. Second, since promotion and training are linked, women should receive less training than males. Although these predictions are not surprising, it is the third implication, namely that women who are promoted will be, on average, of higher ability than males who are promoted, that is of particular interest.[6]

EMPIRICAL ESTIMATION OF PROMOTION AND TRAINING

Description of Data

I test the three indicated hypotheses of the LR model using data from the National Longitudinal Study of Labor Force Behavior—Youth Cohort (1979–1989). The NLSY is an annual survey of 12, 686 young people throughout the U.S. Responsibility for the administration of the NLSY has been shared by two organizations: the Center for Human Resource Research at Ohio State University and NORC (formerly National Opinion Research Center) at the University of Chicago. NORC conducts a 1-h interview each year between January and May with the Youth Cohort respondents, completing, in the process, an extensive questionnaire which covers a wide range of topics, including marital history, schooling, military service, current labor force status, government training, fertility, childcare, and health. In addition to these topics, the NLSY includes variables relating to the Armed Services Vocational Aptitude Battery (ASVAB) tests which were administered to a total of 11,914 of the NLSY respondents in 1980. The ASVAB consists of a battery of ten tests that measure knowledge and skill in such areas as word knowledge, numerical operations, and paragraph comprehension. By combining scores from selected sections of the ASVAB, an Armed Forces Qualifications Test (AFQT) score can be calculated for each respondent.[7]

As discussed at length by Herrnstein and Murray (1994), the AFQT is one of the most highly 'g-loaded' mental tests in current use, i.e. measures 'general intelligence'. In addition, AFQT scores are highly correlated with a wide range of other mental test scores, lending further support to their use as a measure of IQ. For instance, Scullin *et al.* (2000), use the AFQT as a proxy for IQ in examining the relationship between general intelligence, educational attainment, and labor market outcomes. Similarly, the AFQT test scores serve as a proxy for general intelligence or ability in the following analysis. Thus, the NLSY data set provides us with the unique opportunity to explore patterns of promotion across multiple occupations and to link these patterns to a measure of general intelligence or ability typically lacking in other data sets.

Like all IQ tests, the AFQT is designed to be administered to a particular group of individuals. In particular, the AFQT is targeted at individuals in their late teens (18–19 years old) with a high school education. At the time the ASVAB was administered to the respondents in 1980, however, the age range of the cohort was 16–23. As expected, performance on the AFQT is positively correlated with both age (0.20) and education (0.60) in the sample. Regressing AFQT score on age and education, the marginal effect of education (6.01) is much greater than the marginal effect of age (0.83) on AFQT score. To handle this shortcoming, as suggested by Herrnstein and Murray (1994), both age and education are controlled for in the vector of independent variables used in the multivariate analysis to ensure that the AFQT score captures general intelligence or ability.

The initial focus of the analysis is on the NLSY 1984 data, since this is the first year that specific questions relating to promotion are included in the questionnaire. The sample consists of 8143 respondents who were employed either at the time of the 1984 interview or had been employed in the previous year. Although the age range of the cohort in 1984 is 20–27, implying that the respondents have probably not had a great deal of labor force experience, it may still be possible to gain some initial insight into the promotion issue by looking at data from this particular year.

In addition to the NLSY 1984 data, a proxy for the amount of firm-specific training received within a given occupation is obtained from the 1989 NLSY survey. The sample used in the training analysis consists of 7461 respondents who were employed at the time of the 1989 interview or had been employed in the previous year. In 1989 respondents are asked how many months, excluding any regular schooling received, it would take the average new person to become fully trained and qualified to do a job similar to the respondent's job.[8] To simplify the analysis, respondents' jobs are classified into 19 occupational categories, and the mean response for each category is used as a proxy for the amount of firm-specific training within that classification of jobs. These categories, based on the 1980 census of population occupational classification system are: (1) executive, administrative, and managerial occupations; (2) professional specialty occupations; (3) teachers, except post secondary; (4) technicians and related support occupations; (5) sales occupations related to business goods and services; (6) sales occupations related to personal goods and services; (7) administrative support occupations; (8) clerical occupations; (9) private household occupations; (10) protective service occupations; (11) food preparation and service occupations; (12) health and personal service occupations; (13) farming, fishery, and forestry occupations; (14) mechanics and other repair occupations; (15) construction trades; (16) precision production occupations; (17) machine operators, assemblers, and inspectors; (18) transportation and material moving occupations; and (19) handlers, equipment cleaners, helpers, and laborers. Given that a similar training question was not asked in 1984 and, more importantly, many respondents changed jobs between 1984 and 1989, it is necessary to use the mean responses instead of the actual responses from the 1989 survey as a proxy for training in the 1984 probit models of promotion. Thus, by combining the NLSY 1984 promotion data with the NLSY 1989 training data, we are able to investigate a number of the LR model's predictions.[9]

Comparison of General Intelligence Scores

In order to test the LR model's prediction that for a woman to be promoted she must have higher ability than a man, we group mean AFQT scores for selected groups of workers. These mean scores are reported in Table 1. According to the proposed model of jobs, one would expect workers who are promoted to have higher general intelligence or ability on average than workers who are not promoted and, moreover, female workers who are promoted to have higher general intelligence or ability on average than male workers who are promoted. Not surprisingly, promoted females are found to have a higher average AFQT score than non-promoted females. Similarly, promoted males are found to have a higher average AFQT score than non-promoted males. But, as predicted by Lazear and Rosen, females who are promoted are

Table 1. Mean (Std. Dev.) of AFQT Score

	Females		Males	
	Promoted	Not promoted	Promoted	Not promoted
Pooled	72.93	67.72	69.55	63.61
	(18.386)	(20.154)	(21.169)	(23.602)
Sorted by race:				
White	77.08	72.72	74.32	69.52
	(16.463)	(18.667)	(18.964)	(22.500)
Black	57.48	54.53	51.76	48.72
	(17.191)	(17.748)	(19.814)	(19.402)
Non-white and non-black	64.95	56.80	61.08	54.06
	(18.603)	(19.504)	(21.053)	(22.212)
Sorted by union status:				
Union member	67.91	65.10	66.03	60.34
	(19.720)	(19.334)	(19.369)	(21.777)
Non-union member	73.85	68.11	70.53	64.30
	(17.990)	(20.250)	(21.552)	(23.918)

characterized by a higher average AFQT score than males who are promoted. Furthermore, using a simple equality of means test, we are able to reject at a 5% significance level the hypothesis that the average AFQT score of promoted females is equal to the average AFQT score of promoted males, lending support (albeit crude) to the LR model prediction. It is important to note, however, that the evidence is weak in that one gender's mean AFQT score falls within an interval of approximately ± 0.2 standard deviation of the other gender's mean AFQT score. We are also able to reject a similar hypothesis when comparing promoted white females to promoted white males, promoted Black females to promoted Black males, and promoted non-union member females to promoted non-union member males.

Model of Promotion

Given the observed overall differences in average intelligence scores for selected groups of workers, we turn to more formal tests of the proposed relationship between gender and promotion as well as the relationship between gender and training.[10] In order to gain further insight into the issue of gender-specific promotion standards, a probit model of promotion is estimated in which the dependent variable is equal to 1 if the respondent has been promoted at least once by his or her current employer.[11] The probability of being promoted depends on a range of both individual attributes and occupational factors (e.g., McDowell *et al.*, 1999; Ginther and Hayes, 1999). To maintain consistency with the LR model of jobs, the analysis initially focuses on the relationship between specific worker characteristics and the probability of promotion. Occupational differences are addressed in more detail in the next section. The vector of independent variables, defined in Table 2, controls for gender, race, union status, education, labor force experience,[12] and general intelligence or ability. Based on the LR model of jobs, one would expect the

Table 2. Probit Analysis of Promotion for Pooled Sample

Variable	Summary statistics (1) Mean (Std. Dev.)	Probit analysis of promotion (2a) Coefficient (Std. Error)	(2b) Marginal effect	(3a) Coefficient (Std. Error)	(3b) Marginal effect
Constant	—	−1.672** (0.137)	—	−1.664** (0.138)	
FE (=1 if respondent is female; 0 otherwise)	0.476 (0.499)	−0.142** (0.034)	−0.053	−0.109** (0.34)	−0.041
ED (highest year of schooling completed as of May 1 survey)	12.544 (2.033)	0.029** (0.011)	0.011	0.019* (0.011)	0.007
EX (potential years of labor force experience = age–ED–6)	5.081 (2.709)	0.198** (0.019)	0.074	0.197** (0.019)	0.074
EXSQ (EX^2)	33.151 (32.268)	−0.014** (0.002)	−0.005	−0.014** (0.002)	−0.005
UNION (=1 if respondent's wages set by collective bargaining; 0 otherwise)	0.167 (0.373)	0.157** (0.038)	0.060	0.1500** (0.039)	0.057
BLCK (=1 if respondent is Black; 0 otherwise)	0.217 (0.413)	−0.119** (0.052)	−0.044	−0.105** (0.052)	−0.039
NONWH (=1 if respondent is non-white and non-Black; 0 otherwise)	0.051 (0.220)	0.080 (0.092)	0.030	0.087 (0.092)	0.032
FEBLCK (FE × BLCK)	0.101 (0.302)	0.024 (0.072)	0.009	0.016 (0.072)	0.006
FENONWH (FE × NONWH)	0.025 (0.156)	0.010 (0.131)	0.038	0.088 (0.131)	0.034
AFQT (respondent's Armed Forces Qualifications Test score)	67.631 (21.499)	0.007** (0.001)	0.003	0.007** (0.001)	0.003
TRAINING (number of months required for respondent to be fully trained in current job)	18.305 (12.134)	—	—	0.007** (0.001)	0.003

**One-tail test significant at the 2.5% level.
*One-tail test significant at the 5% level.

estimated coefficient for the gender term in this probit model of promotion (FE) to be negative in sign and statistically significant. Controlling for firm-specific training, however, gender should not have a significant effect on the probability of promotion.

Summary statistics of variables used in the analysis are found in the first column of Table 2. The estimated models of promotion, both controlling for and not controlling for training, are presented in columns (2a) and (3a) of Table 2. The associated marginal effects of the independent variables on the probability of promotion are also reported for each model in columns (2b) and (3b), respectively. The marginal effects measure the change in the probability of promotion for an infinitesimal change in each continuous independent variable. For each dummy variable, the marginal effect measures the change in the probability of promotion for a discrete change in the dummy variable from 0 to 1.

The reported results generally support the LR model's predictions concerning the relationship between gender and promotion. In particular, the estimated coefficient for FE in the probit model of promotion which does not control for training (column (2a) of Table 2) is statistically significant and negative in sign, suggesting that women have a lower probability of promotion than men. It is important to note, however, that gender is not the only significant factor affecting the probability of promotion. In particular, education (ED), labor force experience (EX;EXSQ), union status (UNION), race (BLCK), and test scores (AFQT) are also found to be significant. Somewhat surprisingly, the estimated coefficients for the race-gender interaction independent variables (FEBLCK;FE-NONWH) are not statistically significant.

Comparing the marginal effects of those independent variables which are statistically significant (column (2b) of Table 2), their magnitude falls within the range of 4–7%, with the exception of general intelligence (AFQT). For instance, additional labor force experience (EX) increases the likelihood of promotion by 7.4%, being female (FE) reduces the likelihood of promotion by 5.3%, and being Black (BLCK) reduces the likelihood of promotion by 4.4%. Thus, while gender may affect the probability of promotion, whether or not a worker is promoted is also influenced by changes in a number of other worker characteristics.

Given the statistical significance of FE in the probit model of promotion for the pooled sample, the data is sorted by gender and the probit model is re-estimated in order to compare the marginal effect of an increase in general intelligence or ability for females and males. The results are found in Table 3. Based on the LR model of jobs, one would expect the marginal effect of ability on the probability of promotion to be greater for

Table 3. Probit Analysis of Promotion for Data Sorted by Gender

Variable	Females (1a) Coefficient (Std. Error)	Females (1b) Marginal effect	Males (2a) Coefficient (Std. Error)	Males (2b) Marginal effect
Constant	−1.981**	—	−1.594**	—
	(0.205)		(0.186)	
ED (highest year of	0.045**	0.017	0.016	0.006
schooling completed as of May 1 survey)	(0.016)		(0.015)	
EX (potential years of labor force	0.209**	0.077	0.209**	0.080
experience = age–ED–6)	(0.031)		(0.025)	
EXSQ (EX^2)	−0.016**	−0.006	−0.014**	−0.005
	(0.003)		(0.002)	
UNION (= 1 if respondent's	0.125**	0.047	0.168**	0.065
wages set by collective bargaining; 0 otherwise)	(0.060)		(0.050)	
BLCK (= 1 if respondent is	−0.099*	−0.036	−0.120**	−0.045
Black; 0 otherwise)	(0.060)		(0.055)	
NONWH (= 1 if respondent is non-white and non-Black;	0.193**	0.073	0.067	0.026
0 otherwise)	(0.096)		(0.092)	
AFQT (respondent's Armed	0.007**	0.002	0.007**	0.003
Forces Qualifications Test score)	(0.001)		(0.001)	

**One-tail test significant at the 2.5% level.
*One-tail test significant at the 5% level.

males than for females. Since women are assumed to be less attached to the firm, they must meet a higher ability standard than their male counterparts in order to be promoted. So, a given increase in ability will be more likely to allow a male to meet his threshold ability requirement and be promoted than a female. The results reported in Table 3 are consistent with this observation, in that the marginal effect of an increase in general intelligence (AFQT) on the probability of promotion is 0.2% for females and 0.3% for males. The magnitude of this marginal effect, however, is extremely small for both females and males, as it is in the pooled sample. Thus, although general intelligence is a statistically significant predictor of the probability of promotion, the impact of this independent variable is minimal.

It is interesting to note that the observed results are consistent with other theories of promotion presented in the literature. According to the LR model of jobs, different promotion patterns are observed for men and women because of differing non-market abilities. The significance of the gender variable in the probit model of promotion, however, may be the result of other differences between men and women. Browne (1998), for example, suggests that women are less likely to be promoted than their male counterparts due to differences in traits such as aggressiveness, competitiveness, desire for status, and risk tolerance, traits which are highly correlated with success in the workplace. In other words, men and women may differ in their market ability instead of their non-market ability. In such a situation, men may be more likely to base their decision to leave a job on promotion prospects compared to women. When this occurs, the men who remain on the job have a higher promotion rate than their female counterparts due to self-selection bias (Kahn and Griesinger, 1989). The probit model of promotion as presented in Table 2 does not allow us to distinguish between these competing explanations. Instead, the results simply support the idea that there are differences between men and women, either inherent or produced through societal conditioning, which result in different propensities to be promoted.

Model of Training

In order to explore the hypothesized negative relationship between gender and training, a training profile is estimated for the NLSY 1989 data. For the purpose of this regression, only data from the NLSY 1989 survey is used. Instead of using the mean response for each of the 19 occupation categories as a proxy for the amount of training received within that classification of jobs, the dependent variable, TRAINING, consists of the actual responses given by the 7461 respondents in the 1989 sample. In particular, TRAINING is regressed on ability as well as gender, race, union status, education, and labor force experience. These results are found in Table 4.

The estimated coefficient for FE in the OLS training profile for the pooled sample is statistically significant and negative in sign, suggesting that women typically report that less training is required for their jobs than do men. Assuming that reported training requirements are highly correlated with training actually received, the implication is that women receive less training than their male counterparts. This result is consistent with other findings in the literature (Olsen and Sexton, 1996).[13] Given this observation, as well as the fact that promotion is found to be positively correlated with training in the NLSY 1984 data, the implication is that controlling for training should reduce the effect of gender on the probability of promotion. Weak support for this prediction is found in the estimated probit model of promotion which controls for training (column (3a) of Table 2). There is a fall in the absolute value of the coefficient of FE when training is included in the promotion equation. Note, however, that the marginal effect of being female on the probability of promotion only falls from its previous absolute value of 5.3 to 4.1% as a consequence of controlling for training.

Firm-specific training, as it is formulated and used in the probit model of promotion, can be thought of as a proxy for occupation. While training is an important aspect of an occupation, it is not the only job-related factor that may affect the probability of promotion. For instance, the amount of noise inherent in the worker assessment process affects the optimal salary structure within that occupation. Production uncertainty or measurement error reduces the incentive of workers to put forth effort. To counter this effect, employers can increase the wage spread between high-level and low-level jobs in order to encourage employees to compete more vigorously against each other. Additional levels can also be introduced into a firm's hierarchy when sorting workers is

Table 4. OLS Training Profile

Variable	Pooled sample (1) Coefficient (Std. Error)	Females (2) Coefficient (Std. Error)	Males (3) Coefficient (Std. Error)
Constant	−15.013*	−19.287	−21.391*
	(8.846)	(12.132)	(12.852)
FE (=1 if respondent is	−10.530**	—	—
female; 0 otherwise)	(1.735)		
ED (highest year of schooling	2.692**	3.107**	2.341**
completed as of May 1 survey)	(0.525)	(0.727)	(0.755)
EX (potential years of labor	−1.540	−2.656**	−0.272
force experience = age–ED–6)	(0.953)	(1.293)	(1.399)
EXSQ (EX^2)	0.123**	0.166**	0.072
	(0.048)	(0.067)	(0.070)
UNION (=1 if respondent's	−2.249	0.590	−4.894*
wages set by collective bargaining; 0 otherwise)	(1.886)	(2.700)	(2.639)
BLCK (=1 if respondent is	−6.587**	−3.998	−5.324*
Black; 0 otherwise)	(2.508)	(2.572)	(2.760)
NONWH (=1 if respondent is	0.652	−3.777	1.325
non-white and non-Black; 0 otherwise	(4.312)	(5.124)	(4.553)
FEBLCK (FE × BLCK)	4.296	—	—
	(3.326)		
FENONWH (FE × NONWH)	−3.551	—	—
	(6.834)		
AFQT (respondent's Armed Forces	0.159**	0.076	0.225**
Qualifications Test score)	(0.047)	(0.069)	(0.066)

**One-tail test significant at the 2.5% level.
*One-tail test significant at the 5% level.

particularly important, either because the cost of a mismatch between worker and job is high or the work force is heterogeneous (Lazear, 1998). Thus, one would expect the probability of promotion to vary across firms and occupations, an observation which is supported by comparing the frequency of promotion across the 19 occupational categories defined in the analysis. The percentage of respondents promoted in each of these categories ranges from 69% in sales occupations related to business goods and services to 8% in private household occupations. Since men and women are not equally represented in all occupations, it is critical to explicitly model occupational choice when examining patterns of promotion.

JOINT MODEL OF OCCUPATIONAL CHOICE AND PROMOTION

The previous results support the LR model's claims concerning promotion, training, and ability. However, at this point it is important to highlight a key feature of the proposed model of jobs, namely firms characterized by a single entry-level position. Although the job assignment in period 1 is not explicitly specified, the output of a worker in this period is δ regardless of job assignment. Job assignment, therefore, is irrelevant, and it is as if only one entry-level position exists. By restricting their model in this manner, the possibility of worker segregation either before or at the hiring stage is eliminated. In other words, the LR model does not address the potential biased flow of workers into certain occupations.

Returning to the original probit model of promotion, one might argue that the observed statistical significance of gender is at least partially the result of failing to control for occupational choice.[14] In order to account for this possible selection problem, the data are sorted into 19 occupational categories and a bivariate probit model of promotion which controls for sample selection is estimated for each category. In the context of this paper, the bivariate probit model is formulated as

$$y_{i1}^* = \beta_1{}'\mathbf{x}_{i1} + \varepsilon_{i1}, \tag{1}$$

$$y_{i2}^* = \beta_2{}'\mathbf{x}_{i2} + \varepsilon_{i2}, \tag{2}$$

where ε_{ij}, $j = 1,2$ has a bivariate standard normal distribution with correlation ρ, y_{i1}^* is the propensity to be promoted within a particular occupation, and y_{i2}* is the propensity to enter that particular occupation. We do not know these propensities, but we do know whether or not the respondent was promoted and the respondent's occupational choice. Thus, although y_{i1}* and y_{i2}* are not observed, we do observe a signal for each dependent variable, which is coded $y_{ij}=1$ if y_{ij}* is sufficiently great, $y_{ij}=0$ otherwise, $j=1,2$. While Equations (1) and (2) can be estimated separately by univariate probit methods, this is inefficient due to the fact that it ignores the correlation between the disturbances. Within the bivariate probit setting, it may be the case that data on y_{i1} are observed only when y_{i2} equals 1. In other words, a selectivity problem may exist.

In the context of this paper, I consider a model of joint determination of whether an individual enters a particular occupation and whether that individual receives a promotion within that occupation.[15] Obviously, the latter outcome is observed only if the individual actually enters the particular occupation. As a result, in a bivariate probit model for the two outcomes, the observed data are non-randomly selected. This selectivity problem is corrected for by estimating a bivariate probit model with sample selection. Using the notation defined in the previous paragraph, if we let y_{i1} be equal to 1 if the respondent has been promoted at least once by his or her current employer and equal to 0 otherwise, Equation (1) becomes a model of promotion. Similarly, if we define y_{i2} to be equal to 1 if the respondent is employed in a particular occupational category and equal to 0 otherwise, Equation (2) becomes a model of occupational choice for that specific job classification. I assume that the vectors $\mathbf{x}_{i1}$ and $\mathbf{x}_{i2}$ both control for gender, race, union status, education, labor force experience, and general intelligence or ability and that the correlation between the disturbance terms of the promotion and occupational choice equations is ρ. Sorting the NLSY 1984 data into nineteen occupational categories, this bivariate probit model of promotion and occupational choice which controls for sample selection is estimated for each category.

Based on the Wald test statistic, the hypothesis that ρ is equal to zero can be rejected at a 5% significance level in the following six occupational categories: (1) executive, administrative, and managerial occupations; (2) professional specialty occupations; (3) teachers, except post secondary; (4) sales occupations related to personal goods and services; (5) precision production occupations; and (6) machine operators, assemblers, and inspectors. The implication of this observation is that occupational selection either before or at the hiring stage is a significant factor in these job categories, and, as a result, the LR model of jobs does not directly apply. Furthermore, in three additional categories, namely (1) sales occupations related to business goods and services, (2) private household occupations, and (3) protective service occupations, it is not possible to estimate a bivariate probit model of promotion. Our failure to estimate is most likely due to the fact that these categories consist largely of female (private household occupations) or of males (sales occupations related to business goods and services; protective service occupations), suggesting a biased flow of workers into these jobs as well.

Thus, omitting the 9 occupational categories listed above, only 10 occupational categories remain which seem to fit the proposed model of gender-specific promotion standards. Re-estimating the original univariate probit model of promotion defined in column (2a) of Table 2 for each of the ten job classifications, we are able to determine whether or not gender is a significant determinant of the probability of promotion within each category. Due to space limitations, only the estimated coefficient and marginal effect associated with the gender dummy variable FE for each of these 10 categories are reported in Table 5. Looking at Table 5, we see that gender has a significant effect on the probability of promotion in three job classifications: (1) technicians and related support occupations; (2) food preparation and service occupations; and (3) farming, fishing, and forestry occupations. In particular, being female reduces the probability of promotion by 12.7, 20.3, and 11.8% in these three categories respectively, an observation which is consistent with Lazear and Rosen's predictions. As discussed previously, these differences in the probability of promotion may be due to the lack of attachment of female workers to the firm, but they may also be the consequence of some other difference, either inherent or learned, between men and women.

Given that the sample consists of such a young cohort with a relatively short work history, one would not expect to find much labor force information in the data pertaining to promotion. Thus, the fact that a significant negative relationship

Table 5. Probit Analysis of Promotion for Data Sorted by Occupation

Occupation		Estimated coefficent for FE (Std. Error)	Marginal effect for FE
(1)	Technicians and related support occupations	−0.339** (0.165)	−0.127
(2)	Administrative support occupations	0.035 (0.099)	0.014
(3)	Clerical occupations	0.398 (0.266)	0.144
(4)	Food preparation and service occupations	−0.581** (0.116)	−0.203
(5)	Health and personal service occupations	−0.188 (0.136)	−0.058
(6)	Farming, fishery, and forestry occupations	−0.693* (0.392)	−0.118
(7)	Mechanics and other repair occupations	−0.485 (0.518)	−0.176
(8)	Construction trades	−0.306 (0.242)	−0.113
(9)	Transportation and material moving occupations	−0.655 (0.425)	−0.189
(10)	Handler, equipment cleaners, helpers, and laborers	−0.238 (0.183)	−0.074

**One-tail test significant at the 2.5% level.
*One-tail test significant at the 5% level.

between gender and promotion is observed in three out of 10 categories provides some, albeit limited, support for the existence of gender-specific promotion standards within certain occupations. What is important to note is that the relationship between promotion and gender varies across occupations.[16] This finding suggests that more research in this area is necessary to determine the extent and form of the glass ceiling faced by women and other minorities in the U.S.

CONCLUSION

This paper is motivated by the claim that promotion probabilities are lower for women than for men. Focusing on the theory proposed by Lazear and Rosen, this paper tests this claim and two related hypotheses concerning training and ability. As predicted, it is found that females are less likely to be promoted than males, and females receive less training than males. While these results are consistent with other findings in the literature, what is of particular interest is the observation that the relationship between gender and promotion varies across occupations. In contrast to the view typically expressed in the mass media, it is found that females are held to higher promotion standards than their male counterparts as indicated by ability level in only certain occupations. Thus, this study provides some evidence as to the existence of the alleged glass ceiling faced by females as well as to the form that this barrier may take, but suggests that such a ceiling does not apply uniformly to all labor markets.

Although the results presented in this paper are significant in a statistical sense, they do not necessarily provide strong evidence that women are treated significantly different from their male counterparts in the case of promotion in all settings. Clearly, the relationship between promotion and gender is complex and warrants further study. The analysis presented in this paper is merely suggestive of the types of questions that should be examined. A more detailed look at this issue of gender-specific promotion standards is needed. At the very least, such an analysis should include a measure of worker attachment to a single firm as well as a measure of worker attachment to the labor force. Other measures of workers' market and non-market abilities should also be incorporated in order to determine the nature of the differences between men and women that drive different promotion patterns. Ideally, such an analysis would be based on a more recent data set in order to reflect the current state of the work place.

By developing a more explicit model of earnings and promotion, the goal would be to quantify how much of the earnings differential is due to occupational differences at the time of hiring versus after hiring, due to unequal training and promotion opportunities. The source of these unequal training and promotion opportunities after the hiring stage should also be carefully examined. For example, differences in training and promotion opportunities may be attributable to differences in worker attachment to the firm and, thus, be efficient. In contrast, differences in training and promotion opportunities may be attributable to informational or institutional factors such as discrimination and may not be efficient.

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NOTES

1. For example, see Blau and Ferber (1991), Groshen (1991), Light and Ureta (1990), Paglin and Rufolo (1990), Orazem, Mattila, and Yu (1990), Sorenson (1989), Gronau (1988), Gyourko and Tracy (1988), Lindsay and Maloney (1988), Blau and Belier (1988), Robinson and Tomes (1984), Baugh and Stone (1982), Duncan and Leigh (1980), Corcoran and Duncan (1979), Smith (1976, 1977), Polachek (1975), Blineder (1973), and Oaxaca (1973).
2. For a discussion of various theories of discrimination, see Milgrom and Oster (1987).
3. Krowas (1993) extends this three-period model to a five-period model in order to explore how changes over time in outside opportunities for women can lead to an elimination of gender-specific promotion standards. The intuition behind the model is the same, however.
4. Bernhardt (1995) develops an alternative model of promotion that focuses on the information content of promotion. In this model, interfirm mobility plays a critical role in determining patterns of promotion.
5. An alternative explanation of different promotion patterns for men and women may be based on the idea that men and women differ in terms of other traits that affect their relative market ability. This point will be addressed in the next section in more detail.
6. In a related paper, Kaestner (1994) presents empirical evidence which supports two additional predictions of the LR model, namely, the prediction that the incidence of employee separations from the firm influences the promotion decision and the prediction that the difference in promotion rates between men and women diminished with ability. It is important to note, however, that Kaestner's results are based on data drawn from only a single Fortune 500 company. In addition, Kaestner uses a different proxy for ability than that found in this paper. Support for the latter prediction is also found in a large financial company (Jones and Makepeace, 1996).
7. The AFQT score for each respondent is calculated based on the methodology developed and used by the Department of Defense prior to 1989. In particular, the raw scores from the following four sections of the ASVAB are summed for each respondent: Section 2 (arithmetic reasoning), Section 3 (word knowledge), Section 4 (paragraph comprehension), and one half of the score from Section 5 (numerical operations).
8. Of the 7461 respondents who were employed at the time of the 1989 interview or had been employed in the previous year, 120 indicated that they are 'never fully trained' in their current job. Although 480 months is used as a proxy for this response, the results reported in this paper are not dependent upon this assumption. In particular, repeating the analysis with the exclusion of these 120 observations as well as repeating the analysis over a range of values for 'never fully trained' does not change the nature of the findings.
9. It is important to note that the NLSY 1984 data seem to be generally consistent with previous research pertaining to the male-female wage differential. In particular, females do earn less on average than their male counterparts. As suggested by Lazear and Rosen, this average difference in earnings can be (partially) accounted for by the fact that males typically hold higher-paying type jobs than females (e.g. controlling for occupation reduces the magnitude of the earnings difference).
10. Note that promotion and earnings are positively correlated.
11. This specification of the dependent variable ignores differences in the number of promotions among promoted respondents. The average number of promotions for men in the sample is 0.792 and the average number of promotions for women is 0.623. Although this difference in means is statistically significant, it is important to note that approximately 85 and 81% of the male and female respondents, respectively, had either not been promoted or had received only one promotion. Limiting the sample to those respondents who had been promoted at most one time does not change the nature of the reported results.
12. Note that labor force experience is defined as potential years of labor force experience. Technically, it is the difference between the respondent's current age and the respondent's age at the

completion of his/her highest year of schooling. Defining experience in this way allows us to implicitly control for age in the analysis.

13. An alternative explanation is that women and men in the same job may have different perceptions and, thus, respond differently when asked about how much training is necessary to become fully trained and qualified to do a particular job. Interpreted in this manner, the results reported in Table 4 do not address differences in training actually received but simply suggest that women report that less training is required than men. Using such a gender-biased response as the basis for creating the TRAINING proxy variable used in the probit analysis of promotion would result in an overstatement of the effect of controlling for training on the marginal impact of gender.
14. According to an interesting study by Fields and Wolff (1991), the degree of occupational segregation decreased in the U.S. during the 1970s. However, in order to determine if sample selection is a problem in the current sample, it is necessary to estimate a bivariate probit model of promotion.
15. As the data set does not contain information about whether or not a promotion moves a respondent from one occupational category to another, the analysis assumes that the respondent is in the same occupational category after promotion. Such an assumption is reasonable given the broad definition of the occupational categories and the relatively young age of the cohort.
16. Groot and Maassen van den Brink (1996) report a similar result using data from the British Household Panel Survey.

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