# SEX DIFFERENCES IN WORKER QUITTING

W. Kip Viscusi*

## I. Introduction

THE stereotypical view of female employees is that they have relatively weak job attachment and that, in particular, they are especially prone to voluntary job separations. Although this notion is borne out by overall sex differences in aggregative quit rates, this evidence is at best only suggestive since it does not distinguish sex-specific differences in quit behavior from other factors, such as differences in job characteristics and wage rates.[1]

The principal study to date of sex differences in worker quitting is that of Barnes and Jones (1974), who analyzed differences in aggregative quit rates by sex. Although their findings were consistent with the view that females are more prone to quitting, the analysis was restricted to observations for only 19 two-digit industries for each sex so that there was not sufficient information in the sample to analyze many important patterns of interest.[2]

Quit rate studies that do not focus specifically on female quit behavior typically have included a variable reflecting the percentage of workers of a particular sex in the industry. While industries with larger percentages of female employees generally have been associated with higher levels of quitting,[3] these findings for samples of 47–52 two-digit industries are somewhat different from those found in other samples. Indeed, analysis of 95 3-digit industries by Viscusi (1979) reveals no significant sex effect on aggregative quit behavior. In this paper, I will utilize data for a large sample of individuals in an attempt to resolve the ambiguities in these earlier findings.

The most familiar economic motivation underlying potential male-female quit differences is that women often leave the labor market to bear and raise children. Moreover, since wives typically earn lower wage rates than do their spouses, they may serve as secondary earners, entering the labor force during periods of temporary economic needs and exiting thereafter. In addition, family migration decisions, such as those analyzed by Mincer (1978), may lead to quits by wives whose husbands have been transferred to new locales. There also may be important differences in the lifetime employment choice pattern related to the role of quitting as part of an adaptive choice process.[4] To the extent that women have less precise notions of their prospects for advancement and their working conditions, such as the presence of co-worker discrimination, they will be more likely to use the initial period of employment as a period of experimentation and then quit if their experiences are sufficiently unfavorable. An offsetting influence is the fact that males have a greater expected future period of work so that learning-induced quit behavior may offer greater potential gains even though the informational content of the on-the-job experiences may be less.[5] Finally, in situations in which workers are unable to "voice" their complaints effectively and have them settled through grievance procedures, they will adopt an alternative economic response of exiting from the undesirable job context.[6] Co-

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* Northwestern University and Council on Wage and Price Stability.

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[1] The importance of quit behavior to analyses of sex differences in employment and the inconclusive nature of existing studies is discussed by Reynolds (1978), especially on page 167, and by Pigors and Myers (1973).

[2] Their principal regressions included only two age variables and an industry wage variable. Inclusion of a worker education variable knocked out the wage effect for males. See footnote 16 on page 447 of Barnes and Jones (1974).

[3] See, for example, Burton (1969), Burton and Parker (1969), Parsons (1972), Pencavel (1970), and Stoikov and Raimon (1968) for aggregative results of this type. The signs for the worker sex variable are sometimes mixed or statistically insignificant.

[4] This adaptive framework is formalized in Viscusi (1979).

[5] For example, males may have sharper prior assessments of different job outcomes that are altered very little by their on-the-job experiences. Nevertheless, the benefits from quitting may be greater if the number of periods they expect to work are greater than for women. See Viscusi (1979) for formalization of the underlying analytic models.

[6] Freeman (1976, 1978) presents an extensive analysis of worker quitting along the lines of Hirschman's (1970) exit-voice model. His particular concern is with unions, which are

worker discrimination would diminish the effectiveness of grievance procedures for women and increase their quit rate.

It is likely that the importance of sex differences such as these has diminished as women have taken a more active role in the labor market. Indeed, whereas females in manufacturing industries quit 80% more often than did men in 1958, the discrepancy had dropped to 16% by 1968, the last year for which aggregative quit data by sex were collected.[7] Continued narrowing of this gap may have all but eliminated quit behavior as a principal area in which male and female employment patterns differ.

Although unemployment associated with greater labor force entry by women is the primary source of male-female unemployment differences, the greater rate of job leavers (quits) among females also contributes to the discrepancy in unemployment rates.[8] In addition, the presence of substantial differences in quit behavior may account in part for the lower wage rates received by women since firms will receive a lower expected return on their specific training investment than if women had greater job attachment.[9] This widely discussed linkage between wages and turnover hinges on sex differences in the level of quit rates.

However, quit behavior has another potentially important effect on worker wages. Consider a situation in which turnover is costly to the firm and worker quit rates are a continuous function of the wage level rather than a step function that jumps from 0 to 1 as the wage is reduced to some critical level. Responsiveness of this type would arise if, for example, workers differed in their learning about job characteristics or in their cost of changing jobs. Even if female employees had the same absolute level of quitting as did males at the wage rate paid to male workers, the optimizing firm would pay them a lower wage rate if their responsiveness to additional wage payments were less.[10] I will consequently be concerned not simply with determinants of the level of quit rates, but also with any sex differences in the responsiveness of quitting to financial incentives offered by the firm.

The subsequent empirical analysis will consider these issues utilizing a large sample of individuals in the 1976 University of Michigan Panel Study of Income Dynamics (PSID). In addition to including thousands of individuals of each sex, this data set includes persons of all ages so that one can obtain a complete perspective on quit differences. The characteristics of the sample and the principal variables of interest are discussed in section II. In section III, I present the estimates of the determinants of the probability of worker quitting. This discussion will investigate issues such as the relative importance of different variables in influencing quit behavior and the existence of sex differences in the coefficients of the quit rate equation. Section IV utilizes the empirical results to provide a broader perspective on worker quitting. In particular, it assesses the extent to which quit differences are attributable to differences in jobs, differences in personal characteristics, or differences in behavior. The conclusions are summarized in section V.

## II. The Sample and the Variables

The empirical analysis will focus on the quit behavior from 1975–1976 of individuals in the University of Michigan Panel Study of Income Dynamics. Table 1 summarizes the variables' characteristics and the size of each of the subsamples, where all variables pertain to characteristics of the worker and his job in 1975. In addition to stratifying workers by sex, I also attempted to distinguish groups with differing labor market attachment. Whereas the first two columns of data in table 1 pertain to all employed workers, the latter two columns are associated with the subsample that includes only pre-elderly workers (i.e., less than age 65) who work full-time (i.e., 30 or more hours per week). The reason for this division is that retirees and those with relatively loose part-time job attachments may behave

found to reduce quitting. Since Freeman's analysis is quite extensive, I will not be concerned with the union effect here.

[7] These figures were calculated using data from the U.S. Bureau of Labor Statistics (1976).

[8] See, for example, the discussions in Hall (1972), Chiswick and O'Neill (1977), and in Vickrey (1977).

[9] For further discussion of turnover and specific human capital investment, see Oi (1962), Becker (1975), Pencavel (1972), Parsons (1972), and Viscusi (1979).

[10] This result can be generated using almost any of the turnover models cited in the previous footnote. The marginal wage cost of an increase in the wage of each sex will be identical. Other things equal, a marginal increase in the wage rate will offer a greater benefit if the quit rate of the group is more responsive to financial incentives, leading to the use of a higher wage rate for such a group.

differently than do full-time, prime age workers. For both the entire sample and for the full-time, pre-elderly subsample, there are over 3,000 males represented and over 2,000 females.

TABLE 1.—SUMMARY OF SAMPLE CHARACTERISTICS

| | Means and Standard Deviations | | | |
|---|---|---|---|---|
| | All Workers | | Full-Time Pre-Elderly Workers | |
| Variable | Males | Females | Males | Females |
| *QUIT* | 0.84 [a] | 0.167 [a] | 0.079 [a] | 0.172 [a] |
| *AGE* | 36.47 (12.83) | 35.49 (13.23) | 36.01 (12.26) | 34.32 (12.56) |
| *BLACK* | 0.285 [a] | 0.340 [a] | 0.287 [a] | 0.349 [a] |
| *EDUC* | 11.83 (3.79) | 11.93 (2.94) | 11.84 (3.75) | 12.00 (2.92) |
| *KIDS* | 1.46 (1.60) | 1.31 (1.50) | 1.49 (1.61) | 1.28 (1.47) |
| *MARRIED* | 0.872 [a] | 0.705 [a] | 0.875 [a] | 0.700 [a] |
| *HEALTH* | 0.081 [a] | 0.086 [a] | 0.077 [a] | 0.077 [a] |
| *TENURE* | 7.42 (8.19) | 3.93 (5.90) | 7.30 (7.99) | 3.71 (5.63) |
| *TENURE*1 | 0.276 [a] | 0.488 [a] | 0.277 [a] | 0.501 [a] |
| *WAGE* | 4.48 (3.03) | 2.86 (2.06) | 4.55 (3.00) | 2.92 (2.04) |
| *WAGEGAP* | 2.26 (2.86) | 1.44 (1.88) | — — | — — |
| *UNION* | 0.343 [a] | 0.141 [a] | 0.347 [a] | 0.147 [a] |
| *INJRATE* | 10.46 (5.39) | 7.04 (4.03) | 10.52 (5.40) | 7.20 (4.08) |
| *PFEM* | 29.79 (18.60) | 48.82 (19.02) | 29.61 (18.53) | 48.55 (19.40) |
| Sample Size | 3,178 | 2,609 | 3,075 | 2,233 |

[a] The standard deviations of the 0–1 dummy variables are omitted since they can be computed from their fraction $m$ in the sample, where the standard deviation is $(m - m^2)^{.5}$.

The characteristics of the sample appear representative of the working population. Table 2 provides a detailed occupational breakdown for the full sample as well as quit rate information for these occupations. As one would expect, male workers are more likely to be farmers and farm managers, craftsmen and foremen, self-employed, managers, or operatives. The only categories in which female employees exhibit greater concentrations than do men are the low level white collar positions, such as those in the clerical and sales category and in the laborers, farm laborers, and service worker category, while male and female percentages exhibit relative parity for professional, technical, and kindred workers.

For all major occupational groups, male quit rates are considerably lower. The disparity is relatively low, however, for the two categories in which female workers are primarily concentrated—the clerical and sales category and laborers and service workers. These breakdowns do not imply that women in particular types of jobs are more likely to quit since there is substantial heterogeneity in job characteristics and rates of pay within these ten broad classifications. A principal purpose of the subsequent analysis will be to assess the extent to which differences in personal characteristics contribute to these observed differences.

The dependent variable of interest will be *QUIT*, which assumes a value of 1 if the worker quit his 1975 job and 0 otherwise. For both the entire sample and the full-time subsample, female workers quit roughly twice as frequently as did males.

The personal characteristic variables are quite extensive, including information regarding the worker's age in years (*AGE*), race (*BLACK*), years of schooling (*EDUC*), number of children (*KIDS*), marital status (*MARRIED*), health impairments (*HEALTH*), years of experience at the enterprise (*TENURE*), and union membership (*UNION*).[11]

Two wage variables were used. The first was the wage rate in dollars (*WAGE*). When included in the quit equation, this variable can be viewed as part of a larger recursive system in which workers' personal and job characteristics influence the wage rate, and these variables combine to influence quit decisions.[12] The second wage measure was the discrepancy between the actual and predicted wage for each worker (*WAGEGAP*). The predicted wage was determined by a regression for each sex of ln(*WAGE*) on an extensive group of personal and job characteristics.[13] Workers with large *WAGE*-

[11] The 0–1 dummy variables *BLACK*, *MARRIED*, *HEALTH*, and *UNION* were coded in the expected fashion.

[12] See Viscusi (1979) for a fuller articulation of the model.

[13] The explanatory variables included were *AGE*, *INJRATE*, *MARRIED*, *HEALTH*, *UNION*, *TENURE*, *EDUC*, *AGE* × *AGE*, 3 occupational dummy variables, and 3 regional economic conditions variables.

TABLE 2.—OCCUPATIONAL DISTRIBUTION OF QUIT RATES

| Occupation | Percentage in Occupation | | | Percentage Who Quit | | |
|---|---|---|---|---|---|---|
| | Males | Females | Males/ Females | Males | Females | Males/ Females |
| Professional, Technical, and Kindred | 14.3 | 13.4 | 1.07 | 4.8 | 14.9 | .32 |
| Managers, Officials, and Proprietors | 9.4 | 3.2 | 2.94 | 8.7 | 15.5 | .56 |
| Self-Employed | 4.2 | 0.8 | 5.25 | 7.4 | 15.0 | .49 |
| Clerical and Sales | 10.7 | 34.4 | 0.31 | 11.1 | 18.3 | .61 |
| Craftsmen, Foremen, and Kindred | 20.3 | 2.1 | 9.67 | 7.6 | 23.6 | .32 |
| Operatives and Kindred | 20.3 | 15.8 | 1.28 | 10.1 | 17.5 | .58 |
| Laborers, Farm Laborers, and Service Workers | 15.2 | 28.7 | 0.53 | 9.3 | 15.1 | .62 |
| Farmers and Farm Managers | 2.7 | 0.2 | 13.50 | 2.3 | 20.0 | .11 |
| Miscellaneous | 2.6 | 1.4 | 1.86 | 10.7 | 13.5 | .79 |

*GAP* values should be less likely to quit since they are paid more than predicted.

Other job characteristic variables include three occupational dummy variables and two variables created using information regarding the worker's two-digit industry—the percentage of female workers in the industry (*PFEM*) and the 1975 industry injury and illness rate (*INJRATE*).[14] The estimated ln(*WAGE*) equation revealed significant wage compensation for risk, suggesting an implicit value of an on-the-job injury for male workers of $14,000.[15] If, however, workers are not fully informed and compensated for the risk, there will be an additional *INJRATE* effect on worker quitting.[16]

Finally, all equations in the subsequent analysis included two regional dummy variables and an area unemployment rate variable. The expected signs and economic rationales underlying the selection of the principal explanatory variables of interest will be discussed in greater detail in section III.[17]

The final variable included is *TENURE*1, which assumes a value of 1 if the years of experience variable *TENURE* has a value not exceeding 1, and it takes on a value of 0 otherwise. The importance of distinguishing the first year of worker experience is indicated by the data in table 3. Almost half of all female employees have been at their jobs less than a year, as compared with just over one-fourth of the men. The experience distribution thereafter is remarkably similar, with the greatest relative disparity observed in the groups of workers with more than 20 years of experience. Differences of this type are to be expected in view of the increasing labor force activity on the part of women over the past few decades.

The male and female quit percentages in the final columns of table 3 are particularly striking. For workers with a year or less experience, women quit twice as often as do men. However, the relative quit rates fluctuate considerably for all subsequent experience levels, as women exhibit lower quit rates for 5 of the 11 categories. Indeed, for workers with more than a year of experience, the quit percentage for women is 5.9, as compared with 6.4 for men. Once past the initial work period, women are more stable employees than are male workers. The implications of this pattern will be investigated more fully in the following section.

## III. Empirical Results

Worker quit probabilities $p$ were assumed to be characterized by the logistic form,

$$p = \frac{1}{1 + e^{-\beta X}},$$

where $\beta$ is the coefficient vector and $X$ is a vector of explanatory variables. Since the maximum likelihood estimation procedure for this large sample exceeded the computer limits, I em-

[14] The data were from the U.S. Bureau of the Census (1973) and U.S. Bureau of Labor Statistics (1977), respectively.

[15] This value is comparable to that found in earlier studies. For the pre-OSHA BLS injury rates, which were lower in frequency and perhaps more severe, the implicit value of an injury was about $13,000–$13,500. See Viscusi (1979).

[16] The rather aggregative matchups between the workers in the PSID sample and the *INJRATE* values (based on industry listings at the 2-digit level) create substantial measurement error problems, biasing the estimates downward. See Viscusi (1979) for stronger empirical results and presentation of the underlying theory.

[17] I will not, however, dwell on the *UNION* variable since doing so would duplicate Freeman's (1978) analysis. In order to have a comparable union variable for wives and family heads, I used union membership rather than coverage by a collective bargaining agreement as the variable.

TABLE 3.—QUIT RATES AND WORK EXPERIENCE

| | Percentage in Category | | | Percentage Who Quit | | |
|---|---|---|---|---|---|---|
| TENURE ($T$) | Males | Females | Males/ Females | Males | Females | Males/ Females |
| $0 \leq T \leq 1$ | 27.6 | 48.8 | 0.57 | 13.6 | 28.0 | 0.49 |
| $1 < T \leq 2$ | 9.0 | 9.3 | 0.97 | 10.5 | 5.4 | 1.94 |
| $2 < T \leq 3$ | 10.1 | 6.8 | 1.49 | 7.5 | 8.5 | 0.88 |
| $3 < T \leq 4$ | 4.4 | 5.4 | 0.81 | 6.5 | 7.1 | 0.92 |
| $4 < T \leq 5$ | 6.5 | 4.9 | 1.33 | 6.8 | 5.4 | 1.26 |
| $5 < T \leq 6$ | 4.6 | 5.0 | 0.92 | 4.1 | 5.3 | 0.77 |
| $6 < T \leq 7$ | 4.7 | 3.4 | 1.38 | 5.3 | 5.7 | 0.93 |
| $7 < T \leq 8$ | 3.9 | 2.3 | 1.70 | 4.8 | 3.4 | 1.41 |
| $8 < T \leq 9$ | 3.0 | 2.5 | 1.20 | 4.2 | 0 | [a] |
| $9 < T \leq 10$ | 3.0 | 1.7 | 1.76 | 6.3 | 4.5 | 1.40 |
| $10 < T \leq 20$ | 13.9 | 7.1 | 1.96 | 4.3 | 6.5 | 0.68 |
| $20 < T$ | 9.4 | 2.9 | 3.24 | 7.0 | 8.0 | 0.88 |

[a] Ratio is not calculated since female percentage is zero.

ployed the following mixed estimation procedure. After dividing the sample into $k$ random subsamples, I estimated the logit equation for each of the $k$ sets of data.[18] Let $\hat{\beta}_i$ be the estimated parameter vector for the $i^{\text{th}}$ subsample and $\hat{V}_i$ be the associated estimated covariance matrix. The full sample estimates indicated by $\beta^*$ and $V^*$ were obtained using the covariance matrix as weights, or

$$\beta^* = \left( \sum_{i=1}^{k} \hat{V}_i^{-1} \right)^{-1} \left( \sum_{i=1}^{k} \hat{V}_i^{-1} \beta_i \right),$$

and

$$V^* = \left( \sum_{i=1}^{k} \hat{V}_i^{-1} \right)^{-1},$$

where $\beta^*$ is approximately multivariate normal. This procedure can be viewed as an application of the Theil-Goldberger mixed estimation technique where subsample estimates are weighted by the covariance matrices to obtain estimates for the entire sample.[19]

All of the analyses below will be undertaken for three variants of the model. First, the quit probability equation will be estimated for the entire sample using the *WAGE* as the financial rewards variable. Second, the full sample estimates will be obtained using the *WAGEGAP* variable instead of *WAGE*. Third, the full-time pre-elderly estimates will be estimated using the *WAGE* variable. The general spirit of the empirical results was not particularly sensitive to either the nature of the sample or the financial rewards variable employed.

The first matter of interest is to ascertain whether one can pool the males and females, i.e., whether one cannot reject the hypothesis that $\beta_F = \beta_M$, where $\beta_F$ and $\beta_M$ indicate the entire $j$-dimensional coefficient vectors for males and females, respectively. The chi-squared statistic for this test is 48.7 for the entire sample, 63.8 for the *WAGEGAP* estimates for the entire sample, and 59.11 for the full-time sample, where these values are distributed approximately $\chi^2$ with 19 degrees of freedom. The hypothesis that male and female coefficients are identical can be rejected at all usual significance levels.[20]

Although one can reject the hypothesis that $\beta_F = \beta_M$, males and females may nevertheless respond in identical fashion to the explanatory variables but differ solely in the value of the intercept term. Let $\beta_F{}^0$ and $\beta_M{}^0$ indicate the $j - 1$ dimensional coefficient vectors that are identical to $\beta_F$ and $\beta_M$ except that the intercept is excluded. The test statistic for the hypothesis that $\beta_F{}^0 = \beta_M{}^0$ is 47.3 for the entire sample, 48.5 for the *WAGEGAP* quit equations and 53.4 for the full-time sample. The critical $\chi^2$ value is considerably below these levels at conventional significance levels.[21]

This sequence of tests indicates that male and female quit behavior is of a different nature, and

[18] The value of $k$ was set at 3 or 4 for the subsequent analyses.

[19] See Theil (1971). The logit program used was QUAIL, which was developed by Berkman, Brownstone, Duncan, and McFadden (1978). The consistency and efficiency of the mixed estimation technique in this context is formalized by Duncan (1978).

[20] With 19 degrees of freedom, the critical value is 30.1 for $\chi^2_{.05}$ and 38.6 for $\chi^2_{.005}$.

[21] With 18 degrees of freedom, the critical value is 28.9 for $\chi^2_{.05}$ and 37.2 for $\chi^2_{.005}$.

that it cannot be captured by simply adding a sex-specific constant term to the analysis. The nature of these differences can be seen by examining the maximum likelihood estimates of the quit equations for each sex, which are presented in table 4.

Consider first the role of workers' personal characteristics. Since the gains to worker mobility diminish with worker age, one would expect *AGE* to have a negative impact on worker quitting. For both sexes, the elasticity of the quit probability with respect to worker age is substantial, where elasticity estimates in this context reflect the change in the conditional proportion of workers choosing to quit as the explanatory variable is increased. Female quit behavior is about as responsive as males' to age in the samples of all workers, exhibiting elasticities of $-0.77$ and $-0.83$ in the *WAGE* and *WAGEGAP* equations, as compared with $-0.77$ and $-0.97$ for men. Once retirees are excluded from the analysis in the full-time sample, males exhibit a much greater change in their stability with age (a male age elasticity of $-1.33$ as compared with $-0.68$ for women). Due to the more intermittent nature of female employment, the greater stabilizing influence of age for male workers is to be expected.

The effect of worker race on quit behavior is ambiguous from a conceptual standpoint since on-the-job discrimination may increase the incentive to quit while the greater difficulty of locating a new job would tend to diminish quits of blacks. The consistently negative *BLACK* coefficients for each sex suggest that the latter effect is dominant and that racial differences in turnover rates are not responsible for the lower earnings received by black workers.

Worker education has a variety of influences on quit behavior, as it impinges on present and future job opportunities both inside and outside the firm. Schooling has no significant effect on male quit behavior except in the *WAGEGAP* quit equation in which there is an elasticity of $-0.49$ of the quit probability with respect to years of schooling. For females, the education effect is

TABLE 4.—MAXIMUM LIKELIHOOD ESTIMATES OF QUIT RATE EQUATIONS

| Independent Variables | Coefficients and Standard Errors | | | | | |
|---|---|---|---|---|---|---|
| | All Workers | | | | Full-Time, Pre-Elderly Workers | |
| | Males | Males | Females | Females | Males | Females |
| *AGE* | −.023 | −.029 | −.026 | −.028 | −.040 | −.024 |
| | (.007) | (.007) | (.005) | (.006) | (.008) | (.006) |
| *BLACK* | −.704 | −.777 | −.443 | −.492 | −.730 | −.448 |
| | (.210) | (.189) | (.157) | (.155) | (.194) | (.166) |
| *EDUC* | −.011 | −.045 | +.119 | +.069 | −.030 | +.132 |
| | (.026) | (.024) | (.029) | (.027) | (.026) | (.031) |
| *KIDS* | −.095 | −.121 | −.051 | −.074 | −.069 | −.059 |
| | (.055) | (.054) | (.045) | (.045) | (.054) | (.049) |
| *MARRIED* | −.382 | −.419 | −.682 | −.461 | −.304 | −.618 |
| | (.204) | (.189) | (.157) | (.164) | (.201) | (.169) |
| *HEALTH* | +.547 | +.801 | +.383 | +.656 | +.574 | +.445 |
| | (.242) | (.222) | (.207) | (.203) | (.242) | (.223) |
| *TENURE* | −.0035 | +.012 | −.014 | −.048 | +.0046 | −.027 |
| | (.012) | (.012) | (.021) | (.022) | (.013) | (.025) |
| *TENURE*1 | +.473 | +.590 | +1.250 | 1.030 | +.545 | +1.316 |
| | (.180) | (.168) | (.199) | (.203) | (.175) | (.221) |
| *WAGE* | −.226 | — | −.389 | — | −.214 | −.412 |
| | (.028) | — | (.037) | — | (.028) | (.039) |
| *WAGEGAP* | — | −.204 | — | −.360 | — | — |
| | — | (.026) | — | (.036) | — | — |
| *INJRATE* | +.024 | +.013 | +.050 | +.044 | +.025 | +.052 |
| | (.020) | (.019) | (.018) | (.018) | (.019) | (.019) |
| *PFEM* | +.0088 | +.006 | +.0056 | +.0058 | +.0063 | +.0050 |
| | (.0052) | (.005) | (.0039) | (.0038) | (.0051) | (.0041) |

Note: Each equation also includes three occupational dummy variables, two regional dummy variables, an area unemployment rate variable, a unionization variable, and a constant term.

consistently positive and substantial.[22] Although possible explanations for this effect may include greater initial uncertainty and learning-induced quits by better educated women in traditionally male-dominated professions, the precise cause of the discrepancy is unclear.

Marriage and children appear to be stabilizing influences for both groups, although the effect is somewhat stronger for women. Being married reduces the female quit probability by 0.10 in all cases and the male quit probability by 0.03 in the equations with the *WAGE* variable and by 0.04 for the *WAGEGAP* equation. Unmarried women may be especially prone to turnover since a change in their marital status may lead to migration or withdrawal from the labor force to raise a family.

The presence of a health impairment (*HEALTH*) will diminish the worker's ability to switch to a job alternative but will also increase the possibility that a particular job is not well-suited to his particular needs and capabilities. The job experimentation effect appears dominant and of substantial magnitude since health limitations approximately double worker quit rates for workers of both sexes.[23]

The most important personal characteristic variable is the worker's experience at the firm. Although total years of experience (*TENURE*) is never significant, the *TENURE*1 dummy variable for those with a year or less experience exerts a pivotal influence on worker quit probabilities.

This variable reflects three types of economic impacts that one would expect to be most pronounced during the early period of on-the-job experience. First, low tenure workers have acquired less enterprise-specific human capital so that the foregone opportunities after changing jobs will be less.[24] In contrast, those with substantial experience and seniority will be more reluctant to leave their jobs. Second, one would expect that the greatest period of worker learning about the properties of a particular job and one's future prospects will be during the initial period of work.[25] After substantial experience at the firm, additional information is less likely to alter his probabilistic judgments sufficiently to lead him to quit. Finally, those with substantial work experience have revealed themselves to be non-quitters so that *TENURE*1 may reflect this self-selection phenomenon. This variable would, for example, capture women who planned to work only for short periods due perhaps to periodic economic needs.[26]

The strong and diverse conceptual underpinnings of the initial experience variable are reflected in the magnitude of its impact, particularly for females. Male workers with not more than a year of experience have a quit probability ranging from 0.04–0.05 greater than more experienced male workers in the samples, while females in both samples had a quit probability increase of 0.15–0.18 with low levels of experience. Women with not more than one year of experience are about three times more likely to quit than their more experienced counterparts.

Sex differences in the mean level of *TENURE*1 and its magnitude account for a mean sex difference in quitting of 0.08 for each sample using *WAGE* and 0.09 for the *WAGEGAP* quit equation, magnitudes that reflect almost the entire observed difference. Roughly half of this 0.08 value (0.04 for the entire sample with *WAGE*, 0.03 with *WAGEGAP*, and 0.05 for the full-time sample) is attributable to sex differences in the mean value of *TENURE*1. Thus the dramatic influence of *TENURE*1 appears to depend almost equally on differences in the magnitude of the explanatory variable and the size of its coefficient.

As in the analysis of table 3, a principal difference in male and female quit propensities appears to be the greater concentration of women in the low experience group and their greater quit propensities in that experience category. It should be emphasized that this pattern does not simply reflect the loose job attachments of women who work part-time since the results for

[22] The elasticity estimates are 1.18 for the entire sample, 0.69 for the entire sample with *WAGEGAP*, and 1.31 for the full-time sample.

[23] A health impairment increases the male quit probability by 0.08 for the equation including *WAGEGAP* and by 0.05 for those including *WAGE*. Female quit probabilities are increased by 0.06 and 0.07 for the entire and full-time *WAGE* specifications, respectively, and by 0.11 for the *WAGEGAP* equation.

[24] See, particularly, Parsons (1972), Pencavel (1972), Becker (1975), and Oi (1962) for discussions of specific training.

[25] See Viscusi (1979).

[26] At the time this study was initiated, longitudinal quit data were not available for wives so that a variance-components analysis of person-specific quit differences could not be undertaken. This issue is presently being analyzed.

the full-time subsample were almost identical to the findings for all workers.

In all optimizing models of individual job choice, the worker's wage is a central determinant of the quit decision. As was indicated in section I, the responsiveness of worker quitting to the *WAGE* variable also has important implications for market discrimination since groups whose quit propensities are less sensitive to financial incentives will be paid less, other things equal. The consistently significant *WAGE* coefficients reflect remarkably similar behavior. For the entire sample, the elasticity of worker quit probabilities with respect to wage increases is $-0.93$ for both sexes, while for the full sample this elasticity is $-0.90$ for males and $-1.00$ for females, a discrepancy well within the bounds of error.

Similar patterns for each sex were also reflected in the quit equations using *WAGEGAP*, which is the discrepancy between the actual and predicted wage levels. The male and female quit probability elasticities of $-0.42$ and $-0.48$ did not suggest that women were less responsive to financial rewards. In short, male and female workers respond almost identically to financial incentives, suggesting that the cause of the male-female wage gap lies elsewhere.

The principal nonpecuniary job characteristic variable is the injury rate for the worker's industry (*INJRATE*). The analysis in Viscusi (1979) indicates that individuals will display a systematic preference for hazardous jobs whose implications are dimly understood (i.e., loose priors for any given mean value of the prior) and that sufficiently unfavorable on-the-job experience will lead them to quit if not compensated sufficiently. Due to the considerable measurement error involved in matching up the two-digit industry-wide injury rate average to particular individuals, the resulting coefficients undoubtedly understate the true effects. Nevertheless, *INJRATE* is both significant and of substantial importance for females, whose quit probability displays an elasticity with respect to *INJRATE* of 0.3 for all workers and 0.5 for full-time employees. This variable may be of greater relative importance for women since high injury industries tend to include the types of jobs for which women may have less precise notions of the appropriateness of the work tasks to their preferences and capabilities.

Finally, a variable reflecting the percentage of women in the industry (*PFEM*) was included as a test for the possibility of co-worker discrimination against women that might lead to quit behavior after they learned about their unfavorable job conditions. This variable does not display the expected negative sign in the equations for females, but instead performs consistently for all workers and is generally statistically insignificant.

## IV. Comparisons of Quit Behavior

Although analysis of particular variables, most notably *TENURE*1, provides important insights, it is also instructive to obtain a broader view of quit behavior. The three quit difference measures presented in table 5 for the full sample equations with *WAGE* correspond to the rows of that table. The first measure is the predicted difference in the quit probability. The second measure is the wage compensation the women in the sample would require to have the same quit rate as did the men. This statistic translates the behavioral difference into a compensating wage differential, providing a different metric for assessing the extent of the gap. The third statistic is the entropy measure, a widely used information measure that in this context will reflect the degree of surprise from the fact that women do not quit in the same manner as do men. Unlike the first two measures that rely on mean values, this statistic is calculated on an individual basis and averaged for the entire female sample.

Each of these statistics is calculated for four assumptions concerning female quitting as represented by the columns in table 5. First, females are assumed to have the estimated coefficients for their sex and their sex's values of the explanatory variables. These estimates serve as the frame of reference for analyzing the determinants of the predicted quit differences. In the second column, women continue to have their personal characteristics but now have the male coefficient vector. If women behaved in the same manner as men, how would their quit propensities be affected? The third column assumes that women have their sex's coefficient vector but that they have the males' average characteristics and jobs. If women had the male set of characteristics, would their quit behavior be diminished? The last column of the table isolates

TABLE 5.—MEASURES OF QUIT BEHAVIOR DIFFERENCES

| | Assumptions for Female Behavior | | | |
|---|---|---|---|---|
| Index of Male-Female Differences | $\beta_i^F, X_i^F$ | $\beta_i^M, X_i^F$ | $\beta_i^F, \bar{X}_i^M$ | $\beta_i^F, X_i^F$ for Personal Characteristics, $\bar{X}_i^M$ for Job and Region[a] |
| Difference in Male-Female Quit Probabilities, i.e., $\bar{p}^F - \bar{p}^M$ | 0.043 | 0.060 | −0.029 | −0.005 |
| Wage Compensation for Females to Have Male Quit Probabilities, i.e., $\dfrac{\ln\left(\frac{\bar{p}^F}{1-\bar{p}^F}\right) - \ln\left(\frac{\bar{p}^M}{1-\bar{p}^M}\right)}{-\beta^F_{\text{wage}}}$ | \$1.31 | \$1.72 | −\$1.37 | −\$0.20 |
| Entropy Measure, i.e., $\frac{1}{N}\sum_{i=1}^{N} p_i^F \ln\left(\frac{p_i^F}{\bar{p}^M}\right) + (1-p_i^F)\ln\left(\frac{1-p_i^F}{1-\bar{p}^M}\right)$ | 0.147 | 0.078 | 0.007 | 0.036 |

[a] For these calculations, the following variables assumed the values for the females in the sample: *AGE, BLACK, EDUC, TENURE, TENURE*1, *KIDS, HEALTH,* and *MARRIED*. The following variables took on the mean male values: *WAGE, UNION, INJRATE, PFEM*, three occupational dummy variables, and three regional variables.

job-specific differences in the explanatory variables. If women had the same types of jobs and lived in the same regions as did men but otherwise had their own personal characteristics and quit equation coefficients, would the quit difference be narrowed? The experience variables are included with the personal characteristics since these measures compound job differences, such as different specific human capital, and personal differences, such as persistent quit propensities. Consequently, the measure of the relative role of job differences understates the actual contribution of the difference attributable to the types of jobs held by men and women.

The predicted overall quit difference of 0.043 is roughly half of the actual mean difference so the estimated equations understate the actual observed difference in behavior. A relatively modest additional wage premium of \$1.31 per hour would equalize the predicted quit rates. If women behaved as did men, the difference in their quit rates would increase as women would display greater turnover. For the female values of the explanatory variables, female quit behavior actually results in less turnover than would occur if they had males' quit behavior. Similarly, if female employees behaved in the manner predicted by the female quit equation but had the mean value of the males' personal characteristics, they would quit less than would men and would have to incur a \$1.37 wage decrease to equalize their quit behavior. Finally, consider the last column of table 5 in which female quit behavior is altered only by assuming that they have the same types of jobs and live in the same regions as do men. These job differences alone eliminate differences in quit rates.

Moreover, the reduction in the entropy measure from equating simply the jobs and region of the two sexes is 76%, which is only slightly less than the 95% reduction achieved by assuming that all explanatory variables had the male value and considerably greater than the 47% entropy reduction from assuming that women had the male coefficient vector.

These results suggest that the primary difference in male and female quit behavior is a difference in their jobs' characteristics rather than a difference in quit behavior or personal characteristics. Indeed, women exhibit less turnover than they would if they followed the male quit equation. These findings would be reinforced if the role of *TENURE*1 were treated as a job characteristic, such as the enterprise's specific training investment, rather than as a personal characteristic.

## V. Conclusion

Although women quit more both overall and within major occupational groups than do men, this observation is not particularly informative due to the substantial heterogeneity of worker characteristics and job characteristics. Analysis of a sample of almost 6,000 male and female workers suggests that sex differences in quitting have been overdrawn in many previous discussions.

Female quit behavior differs from that of males by more than the addition of a sex-specific intercept term. For example, women are more likely to quit work in hazardous industries due to the likely greater uncertainty regarding their appropriateness to such jobs. Unlike their male counterparts, better educated women are more likely to quit their jobs, perhaps because of the greater uncertainties associated with jobs traditionally held by men. Conventional notions regarding female quitting are reflected by the lower stabilizing effect of age on their quit rates.

Certainly the most important single difference is that female employees are more likely to have no more than a year of experience and within this low experience category they display greater quit rates. The source of the *TENURE*1 difference is not clear since it reflects specific human capital investments, learning about job characteristics that alters the position's attractiveness, as well as periodic labor force attachments other than those reflected through work on a part-time basis (since inclusion of this influence did not substantially affect the results). After the initial year of work, male and female quit rates are roughly identical.

Almost the entire predicted male-female quit difference and half of the actual difference can be explained by differences in their jobs and regional economic conditions. If women had the same job characteristics and the same percentage with more than one year of experience at the firm, their predicted quit rate would be below that for men and their mean quit rate for the sample would be equal to that of men after adjusting for these influences.

Indeed, women display greater stability than they would if characterized by the coefficients in the male quit equation. Coupled with the almost identical response of each group's quit rates to additional wage payments, these findings suggest that the overall quit rates resulting from somewhat different behavior leads to turnover rates more similar than earlier studies have suggested. As a consequence, sex differences in wages and unemployment should not be so readily attributed to turnover-related differences in the behavior of male and female employees.

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