Sex Differences in Labour Turnover in the Australian Public Service*

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This paper reports the use of a large data set of Australian government employees in an analysis of the influence of sex on voluntary labour turnover. Estimates derived from a conventional quit model are used to obtain ceteris paribus *sex specific quit probabilities. These suggest that once other relevant variables are held constant, the expected quit differences between men and single women are significantly reduced. The analysis indicates clearly that the use of aggregate statistics may substantially exaggerate turnover differences attributable to sex.*

I Introduction

The Federal government's recent policy initiatives on Affirmative Action provide further incentive for research into the status of women in the labour market. A better understanding of the forces which have in the past generated different outcomes for men and women enriches the policy debate.

This paper focuses on sex differences in voluntary labour turnover, in part because this factor could help explain the earnings differential between men and women. The justification for regarding turnover as a potential determinant of earnings derives from human capital theory with its explicit recognition of firm-specific investments associated with the hiring and training process (Becker, 1964; Oi, 1962). It is argued that voluntary labour turnover (quitting) imposes costs on employers in the form of foregone returns in their investments. Consequently, when employers judge a worker to have a higher than average probability of quitting they may seek compensation by paying less for any given level of qualification and experience (Chiplin and Sloane, 1976; Chapman, 1984).

The above point makes the basis for employer assessments of workers' turnover intentions a matter of some importance to the issue of wage determination. Yet individual quit probabilities are probably unknown and are likely to be extremely difficult to determine. Consequently, rational employers may resort to using broad generalizations and social stereotypes as the basis for their judgements. The implications of this behaviour for women are fairly clear. The traditional assumption of the association between sex and domestic duties invites the stereotyped view that women have only weak job attachment, a view which is reinforced by aggregate statistics which indicate a far higher quit rate for women than for men. Employees thus can experience statistical discrimination (Phelps, 1972; Spence, 1973).

However, aggregate quit rates are of little value in determining whether there is any inherent difference between men and women in quit behaviour. To isolate a sex differential both job and worker characteristics must be held constant. The study reported here addresses this task by examining male and female turnover behaviour within a single occupation, the clerical/administrative (CA) division of the Australian Public Service (APS).

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It should be noted at the outset that married women were excluded from the analysis and that all men were included as no data on their marital status were available. It was presumed that the greater importance of unobserved variables for married women would result in considerable specification problems. In particular, labour theory suggests that higher spouse's income increases the likelihood of quitting. Because married women with relatively high incomes also are more likely to have spouses with high incomes, the effect of salary on quitting is confused and biased. Consequently, data problems were considered to be too severe to allow meaningful analysis, a presumption confirmed by the poor results (available from the authors) found for this group.

The considerable difference which exists between the aggregate quit rates of the men and single women in the sample provides sufficient justification for investigation. This is even more the case given the possibility that both the behaviour of single women and employer attitudes to this group could be influenced by future expectation of marriage.

Section II reviews the literature and Section III considers theoretical issues relevant to the derivation of the quit function. In Section IV the data are discussed and logit estimations of the model are presented for men and single women. The fifth section compares quit determinants by sex to help establish the extent to which average quit differences between men and women can be explained by differences in labour market characteristics. A conclusion summarizes the major findings and notes their relevance for policy.

II The Literature

Labour turnover imposes obvious costs on business and it accordingly has attracted the interest of personnel and management researchers. Such studies often emphasize sex differences in turnover rates. In an analysis of separations from the APS Forster (1979, p.2) concluded that 'Separation rates are particularly sensitive to sex and age distributions' and went on to suggest that as the percentage of women in the APS increased over time, so too would the overall separation rate. Other Australian studies have continued to give prominence to sex in the presentation of highly aggregated turnover rates (Department of Employment and Industrial Relations, 1976).

Data bases which allow adjustment for sex differences in job and worker characteristics are rare in Australia and hence the bulk of the literature relates to US studies.[1] These reveal a wide divergence of opinion as to the influence of sex on quit behaviour. Mancke (1971), Burton and Parker (1969) and Barnes and Jones (1974) reported that industry quit rates were positively associated with the proportion of women employed. However, these were industry cross-section studies using relatively aggregated data and the results reflect to some extent the heavy concentration of women in industries and occupations with high quit rates for both sexes.

Researchers using more disaggregated data or studying behaviour within a single occupation found little or no sex differences in quit rates (Parsons, 1972; Stoikov and Raimon, 1968; Parrish, 1965; Viscusi, 1980). In fact, Viscusi contended that when all relevant factors were held constant, women had a lower propensity to quit than men. The one indisputable fact to emerge from the American literature is that results are highly sensitive to the degree of aggregation of the data. To isolate a pure sex differential one needs, ideally, microdata which allow adjustment for all the personal and job characteristics influencing turnover.

The one published study in Australia in this area (Lewis, 1979) examined the inter-industry quit rates for men and women employed in the Illawarra region in August 1976. Lewis showed that much more of the differential could be explained by sex differences in occupational distributions but concluded that even after adjustment for this factor women still demonstrated a higher propensity to quit than men. The data available to Lewis, however, were highly aggregated. Of policy interest was his finding that employers consistently exaggerated the sex differential of their own firm's aggregate quit rates, ascribing a higher rate to women and a lower one to men than was indicated by the firm's employment records.

Lewis' finding assumes considerable importance if employers do in fact act against groups perceived to have relatively high quit rates. While the presumption of this behaviour is common in the literature (Lewis, 1979; Chiplin and Sloane, 1976; Blau and Kahn, 1981; Landes, 1977; Goldfarb and Hosek, 1976) very few studies have attempted to test the hypothesis directly. One exception is

[1] For a survey of turnover models see Parsons (1977). Recent additions to the literature may be found in Jovanovic (1979a, 1979b), Chapman (1982) and Viscusi (1980).

Chapman (1984) who developed a model of wage determination which incorporated employer expectations of employee quit probabilities. With some qualifications the results supported the hypothesis that employees experienced statistical discrimination. Both the Lewis and Chapman studies suggest the need for a better understanding of the determinants of quit behaviour and, in particular, the role played by sex.

III Theoretical Issues

Following Borjas (1984) and others, labour turnover determinants may be expressed as

$$I_i = \alpha_0 + \alpha_1 P_{ci} + \alpha_2 P_{ai} + \alpha_3 M_{mi} \qquad (1)$$

where, for individual i, I_i is a mobility index (or turnover probability), P_{ci} is expected discounted utility in the current job, P_{ai} is expected discount utility in alternative employment or activities, and M_{mi} is an index of the cost of mobility. This simple equation captures the essence of quite complex turnover models, such as job matching (Jovanovic, 1979a), and we expect that $\alpha_1 < 0$, $\alpha_2 > 0$ and $\alpha_3 < 0$. To understand the net effect of the unobserved variables in (1), it is assumed that they can be approximated by various observable socio-economic characteristics. If these effects vary by sex (for example, as a consequence of attitudinal or taste factors that may be influenced by perceptions of market and non-market opportunities), it follows that

$$P_{ci} = \delta^j x_i + \epsilon_{ci}$$

$$P_{ai} = \lambda^j Y_i + \epsilon_{ai}$$

and

$$M_{mi} = \beta^j Z_i + \epsilon_{mi} \qquad (2)$$

where X_i is a vector of variables affecting productivity on the current job, Y_i is a vector of variables affecting productivity in alternative employment or activities, Z_i is a vector of variables affecting mobility costs, the ϵ's are error terms, and the j's are male or female superscripts.

Substituting equation (2) into (1) yields the turnover reduced form

$$\begin{aligned} I^j_i &= \alpha_0 + \alpha_1(\delta^j X^j_i + \epsilon^j_{ci}) + \alpha_2(\lambda^j Y^j_i + \epsilon^j_{ai}) \\ &\quad + \alpha_3(\beta^j Z^j_i + \epsilon^j_{mi}) \\ &= \alpha_0 + \alpha_1\delta^j X^j_i + \alpha_2\lambda^j Y^j_i + \alpha_3 b^j Z^j_i \\ &\quad + (\alpha_1\epsilon^j_{ci} + \alpha_2\epsilon^j_{ai} + \alpha_3\epsilon^j_{mi}). \end{aligned} \qquad (3)$$

Equation (3) illustrates the complexity involved in establishing signs for the coefficients of the variables available to represent X, Y and Z. In part this is a consequence of some of these variables influencing more than one of the three parameters, a point that becomes apparent in the discussion of the quit functions to be estimated for each sex.

To avoid, among others, the problems of heteroscedasticity, logit is an appropriate functional form for estimations of equation (3). Thus for each sex, the quit function is given by

$$P^j_i = 1/(1 + e^{-\beta^j Q^j_i}) \qquad (4)$$

where, for individual i of sex j, P is the probability of quitting during the period (in this case, 1969–74), β is a vector of quit coefficients and Q is a vector of independent variables. For the data available, Q is made up of 1969 annual salary (*SAL*), length of time employed in the firm (*OJEX*), years of schooling (*YOS*), age in years (*AGE*) (entered as five-year interval dummies) and a dummy variable (*CAN*) equal to one if the individual was employed in Canberra, and to zero otherwise. These variables are discussed further in Section IV.

In attempting to determine coefficient signs it is important to specify the assumptions typically imposed in the derivation of the mobility index I (equation (1)). Search theories of turnover characterize the world as one in which workers are less than fully informed about either the optimality of the match of their current job to their (unobserved) skills (Jovanovic, 1979a) or the range and desirability of opportunities outside current employment (Parsons, 1977).

Conventionally, labour market analysis presumes that a major constraint to mobility is the existence of worker-financed firm specific training investments.[2] In this view, workers quitting forego future returns to the present specific human capital stock since such skills are valueless in alternative employment. With the above framework in mind, it is useful to venture some expectations of the signs of the coefficients on Q in equation (4).

Current salary operates through its effects on discounted utility on the current job, and thus its sign in the quit equation depends on the interaction $\alpha_1\delta_0$ from equation (3). A higher salary increases utility, $\delta_0 d_0 > 0$. Since $\alpha_1 < 0$, $\alpha_1\delta_0 < 0$, and the quit probability is predicted to decrease as *SAL* increases. This specification is not ideal, however, because the appropriate variable to be employed is not wage, but wage relative to alternatives. It has

[2] For the original theory, see Becker (1964) and for empirical confirmation of the hypothesis, see Chapman and Tan (1980).

been demonstrated that the use of the former tends to bias towards zero the absolute size of the coefficient in the quit function (Chapman, 1982). Data inadequacies precluded the use of the more correct variable.

Time in the current firm also increases productivity (utility) in current employment because of its positive correlation with specific training investments (that is, from equation (3) $\delta > 0$ and thus $\alpha_1\delta_0 < 0$). However, *OJEX* may also increase productivity in alternative employment because of its association with general skills accumulation, implying that $\alpha_2\lambda > 0$ if outside employers reward these skills relatively highly (Beggs and Chapman, 1985). Further, some individuals with high levels of CA experience may be more aware of and attracted to outside job opportunities, further implying that $\alpha_2\lambda > 0$. The lack of access to relative wage data may mean that *OJEX* will reflect this influence to some extent.

For the above reasons signing the coefficient on *OJEX* in the quit function is difficult. We believed, however, that specific training factors, coupled with self-selection (individuals with relatively high tenure have demonstrated an unwillingness to quit in the past, and are thus less likely to quit in the future), would swamp other complications and result in a negative relationship between *OJEX* and the quit probability. In other words, $\alpha_1\delta + \sigma_2\lambda < 0$, because $\alpha_1\delta < 0$ and $|\alpha_1\delta| > |\alpha_2\lambda|$.

Education is presumed to affect productivity on both the current and in alternative jobs, and the costs of mobility for reasons explained below. Thus the sign of the coefficient on *YOS* in the quit equation depends on, from equation (3), $\alpha_1\delta + \alpha_2\lambda + \alpha_3\beta$. Education will increase current job productivity more than alternative job productivity if individuals with more schooling acquire job-specific skills relatively rapidly. That is $|\alpha_1\delta| > |\alpha_2\lambda|$ and, with no mobility cost implications, the coefficient would be expected to be negative. However, if schooling decreases search costs, $\beta < 0$, and thus $\alpha_3\beta > 0$. But because the minimum level of schooling is high it is doubtful that there are significant differences in the sample in overall search efficiency. Thus, while ambiguous from theory, we expected higher levels of schooling to diminish quit probabilities.

The coefficients on *AGE* are more difficult to interpret. The tendency for 'job-shopping'[3] to decrease with age suggests that workers acquire greater knowledge of their individual abilities with time and are thus less likely to change jobs as they grow older. Reduced expectations from job-shopping imply, from equation (3), that $\alpha_1\delta < 0$, since $\delta > 0$. As well, we would expect the influence of these factors to decrease with age.

Predicting the sign of the *CAN* coefficient is also complex. Being employed in Canberra has offsetting effects on expected productivity in the current job: promotion prospects are enhanced and diminished respectively by the greater number of upper echelon positions and the likely greater ability of competitors for those positions. That is, $\lambda > 0$. On the other hand, fewer alternative job prospects in Canberra may discourage quitting (i.e., $\lambda < 0$, and thus $\alpha_2\delta < 0$). These ambiguities and no strong priors concerning relative magnitudes discourage a prediction of the sign of this variable.

Difficulties in interpreting coefficients are not of overriding concern. This is because the major motivation for the estimation of equation (4) is to allow insights into the following important questions. First, given control for other variables associated with quitting, how large is the sex difference in the probability of labour turnover? Second, are the determinants of turnover for Australian government employees similar to those estimated for other samples? Both issues are important in a statistical sense, and the first has fundamental policy implications.

IV The Data and the Logit Estimations

The data consist of information on all full-time male and single female CA officers employed in the Third Division of the APS in 1969. The individuals involved performed public sector white collar duties such as the preparation of reports on various aspects of government, and the administration of government projects and personnel. CA officers of a more specialized variety were excluded from the analysis.[4] Thus it is not a gross over-simplification to treat the group as relatively homogeneous. Nevertheless, conclusions reached concerning the existence of quit differences in what is assumed to be a single occupation should be interpreted with some caution.

Information on this group of workers is available for the years 1969 and 1974, thus allowing an examination of quit determinants over a five-year

[3] Job-shopping might be understood as the acquisition of information as to both individual-specific abilities and the requirements of alternative jobs.

[4] For example, auditors, computer systems officers, naval stores officers and interpreter/translators. We are grateful to Dr George Rotham for this suggestion.

TABLE 1
Statistical Characteristics of the Data (as of 1969)

Variable	Men		Women	
	Mean	Standard Deviation	Mean	Standard Deviation
Annual salary ($) (*SAL*)	4410.2	1442.2	2983.7	1037.2
Age (years) (*AGE*)	27.37	7.62	22.87	5.59
Internal experience* (years) (*OJEX*)	5.70	5.94	2.70	4.25
Education** (years) (*YOS*)	12.46	1.10	12.32	0.95
Canberra (*CAN*)***	31.85		27 47	
Percent quitting by 1974	20.67		52.52	
Observations	16534		2119	

* Time since joining the CA division; ** see explanation of derivation in text; *** percentage of workers employed in Canberra.

period. The analysis was limited to individuals aged less than 45 years in 1969 in order to minimize the possibility that separations (identified only by the individual's absence in 1974) resulted from age or invalidity retirements. Given that discharges are virtually non-existent in the APS, if individuals were not in the sample in 1974 the assumption that they had voluntarily separated is not considered to be unrealistic.[5] The above limitations, and the exclusions of observations with inconsistent or incomplete data, restricted the sample sizes to 16534 men and 2119 single women. Table 1 presents the statistical characteristics of the data.

The calculation of several variables given the available data required the imposition of restrictive assumptions. First, the individual's education was recorded as the highest qualification as of 1969. These educational qualifications were converted to year-of-schooling equivalents. Table 2 presents information on the distribution of qualifications and the assumptions imposed in order to derive years of education. Second, as actual earnings are not available in the data they were computed by taking the midpoint of the salary range for the class level of the individual. Neither the schooling nor salary adjustments are believed to introduce important distortions, although some measurement error is introduced by these simplifications.

[5] The category 'voluntary separation' does not necessarily mean that the employees initiated turnover moves. An employer, or spouse, can encourage quits.

TABLE 2
Highest Educational Qualifications of the Sample

Qualifications	Per Cent of Sample	Assumed Year-of-Schooling Equivalent
Matriculation	76.2	12
Diploma	5.6	14
Bachelor – ordinary	14.4	15
Bachelor – honours	3.1	16
Masters	0.6	18
Doctorate	0.1	20

The quit model used followed the conventional form of equation (4), and was estimated separately by sex. Quit status in 1974 was modelled to be a function of the 1969 value of the independent variables, results of these estimates being reported in Table 3.

The coefficients of Table 3 may be used to investigate how quit probabilities vary in response to changes in the independent variables. These calculations are presented for men and single women in Table 4, which shows the percentage point change in the quit probability for various explanatory variable changes, calculated for individuals who have sample mean characteristics of their sex. Because of the nonlinear estimation method, these marginal impacts will change for different explanatory variable levels.

TABLE 3
*Logit Estimations of the Quit Model**

Variable	Men	Women
Intercept	−0.064 6	1.087
	(0.26)	(1.67)
SAL	−0.000 0642	−0.000 178
	(2.57)	(2.39)
OJEX	−0.099 8	−0.115
	(13.82)	(5.00)
AGE 20-24	−0.195	0.110
	(2.83)	(0.86)
AGE 25-29	−0.358	−0.373
	(4.10)	(1.99)
AGE 30-34	−0.435	−0.338
	(3.80)	(1.14)
AGE 35-39	−0.342	−0.540
	(2.57)	(1.29)
AGE 40-44	−0.155	−1.00
	(1.24)	(2.06)
YOS	−0.026 6	−0.016 5
	(1.25)	(0.29)
CAN	0.044	0.254
	(0.89)	(2.26)
Standard errors of regression equations	0.006 45	0.023 3
Goodness of Fit Criteria:		
Bookmaker Specification test	0.004 1	0.004 6
McFadden's R^2	0.056	0.079
Number of wrong predictions	0	0

*Absolute *t* statistics in parentheses.

The results are similar, in sign, to those estimated on overseas data. Wage, or salary, and time with the current firm are typically negatively associated with quitting (Parsons, 1977; Viscusi, 1980; Chapman, 1982). Schooling is negatively, but not significantly, related to quitting for both groups. The male quit probability is minimized at age 30-34, and the single female quit probability is minimized at age 40-44. Being employed in Canberra had no effect on male quit probabilities and was positively associated with quitting for females.

The goodness of fit criteria statistics are a mixed bag.[6] The bookmaker specification tests suggest that the models perform reasonably, and both estimations had a perfect record in terms of within-sample prediction. As well, the standard errors of the regression equation were fairly small. On the other hand, relatively low McFadden R^2s were obtained.

Coefficient sizes vary from those obtained in other studies, in part because of dissimilarities in functional form and data availability. Nevertheless, the coefficients reported above are not markedly different to those estimated elsewhere. It is worth noting that the female marginal impacts for both *SAL* and *OJEX* are substantially higher than those estimated for males, and that *AGE* is a more important quit determinant for men.

[6] For information on these tests, see Amemiya (1981), Beggs (1983) and McFadden (1974).

TABLE 4
*Percentage Point Change in Quit Probability for a Given Change in Independent Variables**, ++

Variable	Men		Women	
	(i)	(ii)	(i)	(ii)
SAL	−0.42	−1.82	−1.33	−8.62
OJEX	−0.85	−1.46	−0.78	−2.88
AGE 20-24	−3.22		2.74+	
AGE 25-29	−5.59		−9.25	
AGE 30-34	−6.63		−8.39+	
AGE 35-39	−5.36		−13.25+	
AGE 40-44	2.57+		−23.32	
YOS	−0.50**	−0.40**	−0.51**	−0.41**
CAN	−0.75**		−6.30	

* Equation (i) is for 10 per cent increases in the continuous variables.
Equation (ii) is for a $1 per hour increase in salary and one year increases in *OJEX* and *YOS*.

** Not significantly different from zero at the 5 per cent level.

\+ Not significantly different from $AGE<20$ at the 5 per cent level.

++ Figures based on individuals with the sample mean characteristics of their sex.

TABLE 5
Adjusted Quit Probabilities

Men			Women		
Actual	Adjusted	95% Confidence Interval	Actual	Adjusted	95% Confidence Interval
.207	0.266	0.256–0.277	0.525	0.329	0.287–0.374

V Using the Results to Estimate Sex Differences in Quitting

The above results imply that economic variables are significantly associated with the probability of quitting for men and single women. This suggests that using average quit differences by sex to ascertain the expectation of quitting is an inappropriate approach if the measures of these economic variables vary between the groups. Only through an adjustment of the data is it possible to ascertain the underlying sex differential. That is, some part of the average turnover probability by sex is attributable to differences in the average level of determinants rather than through gender-specific behaviour. This may be demonstrated as follows.

From equation (4) it is possible to determine what the male and single female average quit probabilities would have been if the determinants of quitting were at the same level as the other group by calculating

$$AQP_F = 1/(1 + e^{-\beta^F Q^M})$$

and

$$AQP_M = 1/(1 + e^{-\beta^M Q^F}) \qquad (5)$$

where, respectively for females and males, AQP_F and AQP_M are adjusted quit probabilities, β^F and β^M are estimated quit coefficients, and $\bar{Q}^F$ and $\bar{Q}^M$ are average levels of independent variables. These data and their 95 per cent confidence levels are presented in Table 5.

The data of Table 5 should be interpreted as follows. If men had had the same average levels of the independent variables as single women, their average quit probability would have been 26.6 per cent, instead of the observed 20.7 per cent. If single women had had the same average levels of independent variables as men, their average quit probability would have been 32.9 per cent, instead of the observed 52.5 per cent.

From these estimations the underlying sex differential in labour turnover probabilities over the five-year period averages 19.1 ([(52.5 − 26.6) + (32.9 − 20.7)]/2) per cent, rather than the observed (uncorrected) difference of 31.8 (52.5 − 20.7) per cent. Women would still have had greater quit rates than men if they had had the same economic characteristics as men, but the differences are reduced sustantially after controlling for these characteristics. The 95 per cent confidence intervals reveal that the adjusted quit probabilities are statistically significantly different from the unadjusted quit probabilities for both men and women.

These results are important to an understanding of gender-specific behaviour with regard to labour turnover. The adjusted quit probabilities provide an answer to the question of the extent of male-female differences holding constant other determinants. They suggest that a significant part of the single women–men average quit difference is a consequence of lower female salaries and on-the-job experience. Unadjusted quit probabilities exaggerate the *ceteris paribus* expectations of single women quitting compared to men by about 50 ([(31.8 − 19.1) / (31.8 + 19.1) / 2].100) per cent. These estimates demonstrate that using averages to infer gender-specific turnover probabilities exaggerates sex differences for this sample in the 1969-74 period.

VI Conclusions

This paper has used a large data set of CA employees to test some important hypotheses on labour turnover in the 1969-1974 period. It was found that the conventional quit model performs well for men and single women. The quit estimations are used to ascertain *ceteris paribus* sex-specific turnover probabilities. The data reveal that once relevant economic variables are held constant, the expected quit differences between men and single women are reduced significantly. Thus, for this sample, the use of average quit data appears to exaggerate substantially turnover differences attributable to sex.

Partly for historical reasons, such as the 'marriage bar' [7] and age restrictions on recruitment, the women in this sample were on average younger, less experienced and lower paid than the men. These factors contributed significantly to their higher average quit rate, but adjustment for these effects indicated that women still had a higher propensity to quit than men. The results may be misleading if translated to other situations since the findings apply only to workers employed in the APS in 1969 when attitudes to work and perceptions of female career possibilities might have been influenced by prior conditions such as the marriage bar and contemporary socialization processes. The dramatic change in female employment patterns in Australia since 1969 suggests the potential for a significant change in these relationships.[8] Exploration of this hypothesis must await further research.

What is clear is that the use of aggregate statistics may significantly overstate sex differences in quit rates, with potentially adverse implications for the employment experience of women. On both efficiency and equity grounds, it could thus be of substantial benefit to ensure that employers have access to the correct information.

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[7] The marriage bar referred to Section 49 of the *Public Service Act*, under which 'every female officer shall be deemed to have retired from the Commonwealth Service upon her marriage'. This restriction was lifted on 18 November 1966. See Public Service Board, *Annual Report*, 1970.

[8] It should be noted that Viscusi (1980) found a steady decrease over time in the US in the discrepancy between male and female aggregate quit rates.

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