# LABOUR TURNOVER AND WAGE DETERMINATION

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## I. Introduction

One possible contribution to the existence of wage differentials between individuals is that employers attach a premium to employment stability. However, since employees' labour turnover intentions are unknown, it is likely to be in the interest of the employer to form judgements of the quit probability on the basis of observable characteristics. Such judgements, once instituted, fall into a genre of behaviour known as statistical discrimination.[1] These judgements necessarily disadvantage those employees whose turnover intentions imply quit probabilities which are lower than the average of the group to which they have been subjectively allocated to by the employer.

Analyses of this phenomenon typically allude to the human capital model of investment in firm specific hiring and training (FST). This approach recognises that FST investments will be shared by firms and workers in order to minimise the probability of the other party instituting job separation proceedings. Employers are concerned with quits (and workers with layoffs) because such action implies foregone returns to FST outlays.

As yet theoretical examination of the question is limited,[2] and empirical tests of the main hypotheses focus on industry rather than individual data.[3] This paper addresses both aspects of the issue: a wage determination model based on the theory of human capital investment is expanded to incorporate employer expectations of employee quit probabilities, and is presented in section II. In section III the main hypotheses are subjected to tests

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[1]See E. Phelps (1972) and M. Spence (1973).

[2]See James F. Ragan and Sharon P. Smith (1981), Robert S. Goldfarb and James R. Hosek (1976) and Mary Corcoran and Greg J. Duncan (1979). The models analysed are either quite informal, or simply treat working time dichotomously, as training or post-training.

[3]James Ragan and Sharon Smith (1981) have analysed this issue with US data. Their model assumed that employers estimate quit probabilities on the basis of observation of sex-specific quit rates associated with industry. Support is found for the hypothesis through the discovery of a significant negative relationships between aggregate past industry quit probabilities and current earnings. The present investigation is a substantial improvement over the Ragan and Smith estimation because it ascribes to employers the more realistic expectation that many observable individual variables influence quit expectations. The Ragan and Smith test is narrowly defined in the sense that other observable individual characteristics, such as age, time on-the-job and education, play no part in the employers' presumption.

with a large sample of Australian women government clerical workers. While some support is found for the prediction of the model, several qualifications are noted. A concluding section summarises the main findings.

## II. The Theoretical Model[4]

Assume that there exists a class of jobs such that entry level wages are identical for all persons with equivalent stocks of measurable general human capital. These jobs are assumed to be characterised by identical levels of the initial hiring and training outlays, $T$, and workers pay the same share $\alpha$ (with $0 < \alpha < 1$) of investment costs. The assumptions imply that first period worker-financed specific investment outlays, in the form of foregone wages, are $\alpha T$ and that first period firm-financed specific investment outlays, in the form of foregone output, are $(1 - \alpha)T$.

Because rational workers should decrease training as the period to accrue returns shortens (*i.e.* as retirement or voluntary labour turnover approaches), an appealing investment function is one that declines over time. One such function is:

$$I(t) = Tr^t \tag{1}$$

where $I(t)$ is total investment outlays in the $t^{th}$ period of tenure, and $r$ is the rate of change of investments with tenure, with $0 < r < 1$.

This implies the following:

$$W_{1i} = bH_i - \alpha T \tag{2}$$

where, for individual $i, W_1, b$ and $H$ are the initial period wage, the rate of return to human capital (with $0 < b < 1$), and the pre-employment human capital stock respectively. The investment function implies that:

$$\begin{aligned} W_{ti} &= bH_i - \alpha Tr^t + b\alpha T(1 + r + r^2 + \dots + r^{t-1}) \\ &= bH_i - \alpha Tr^t + b\alpha T\left(\frac{1 - r^t}{1 - r}\right) \end{aligned} \tag{3}$$

Equation (3) implicitly assumes that there is no depreciation of human capital. Importantly for the task at hand, it assumes that all workers are expected to remain on the job until retirement. This assumption is relaxed in the following analysis.

Given that the firm's share of investment outlays is $(1 - \alpha)$, the $(t + 1)^{th}$ period (undiscounted) investment return losses for the firm associated with worker $i$ leaving the job after $t$ periods ($IL_t^F$) is:

$$IL_{ti}^F = \frac{b(1 - \alpha)\, T\, (1 - r^t)}{1 - r} \tag{4}$$

Assuming a fixed working life of $z$ years, the total (undiscounted) loss for the firm of the worker quitting after $t$ periods is thus:

[4]This model is a modified and extended version of that developed in Bruce J. Chapman (1982). Malcolm Gray made some pertinent observations on its form.

$$TL_t^F = \frac{(z - a)\, b\, (1 - r^t)}{1 - r} \tag{5}$$

given $(z - a)$ periods remaining until retirement.

The total discounted loss to the firm of the worker quitting after acquiring $t$ periods of tenure ($TLD_{ti}$) is given by:

$$TLD_{ti} = \sum_{a=0}^{z} \frac{b(1 - \alpha)T(1 - r^t)}{(1 - r)\,(1 + d)^a} \tag{6}$$

where $d$ is the discount rate ($0 < d < 1$). If the quit probability in the next period is given by $q$, the firm's expectation of (discounted) investment return losses ($TLDQ_{ti}$) is given by:

$$TLDQ_{ti} = q\left\{\sum_{a=0}^{z} \frac{b(1 - \alpha)T(1 - r^t)}{(1 - r)\,(1 + d)^a}\right\} \tag{7}$$

In equilibrium a risk-neutral firm will adjust wages to reflect expected investment losses[5], so that in the presence of quitting:

$$W_t = bH - \alpha T r^t + b\alpha T\left(\frac{1 - r^t}{1 - r}\right)$$

$$- q\left\{\sum_{a=0}^{z} \frac{b(1 - \alpha)T(1 - r^t)}{(1 - r)\,(1 + d)^a}\right\} \tag{8}$$

Clearly,

$$\frac{\partial W_t}{\partial q} = -\left\{\sum_{a=0}^{z} \frac{b(1 - \alpha)T(1 - r^t)}{(1 - r)\,(1 + d)^a}\right\} < 0 \tag{9}$$

Result (9), the major empirical implication of the model, is that an individual's wage will be lower the greater has been the employer's expectation of a quit. This wage discount, presumably, will be manifested through lower promotion rates.

The analysis so far has implicitly assumed that $q$, the employer's expectation of the worker's quit probability, does not vary across individuals. However, it is credible that employers form expectations of quit probabilities on the basis of a host of worker variables such as sex, marital status, age, schooling and completed time-on-the-job. In order to clarify the conceptual and estimation issues, assume that the employer ascribes to each individual a quit probability $P_i$ in the next period described by the logistic form:

$$P_i = \frac{1}{1 + e^{-\beta X_i}} \tag{10}$$

[5]Note that this model implicitly assumes that the probability of a firm-initiated separation, or layoff, is zero. It would be straightforward to generate the comprehensive wage result by adding an analogous term or worker investment loss multiplied by the layoff probability.

where $\beta$ is a vector of coefficients and $X$ is a vector of explanatory variables,[6] for individual $i$. Estimation of equation (8) implies the necessity of incorporating equation (10) into the usual human capital wage equation.

## III. Estimating the Model

A simple extension of the human capital earning function[7] may be utilised to test the basic hypotheses concerning the relationship between wages and quit expectations by the employer. An estimating equation consistent with the model is:

$$lnW_i = a + b\ GEXP_i + c\ GEXP_i^2 + d\ OJEX_i + e\ OJEX_i^2 + f\ YOS_i$$
$$+ g(\frac{1}{1 + e^{-\beta X_i}}) + \varepsilon \quad (11)$$

where for individual $i$, $lnW$ is the logarithm of wages (or salary) $GEXP$ is length of time in the labour force (age-schooling-5), $OJEX$ is length of time in the current firm, $YOS$ is years of schooling, and $\varepsilon$ is an error term. The quadratic experience terms have been included to reflect the assumption in the model that training investments decline as retirement approaches. The predictions of this model are: $b,d,f > 0$ and $c,e,g < 0$.

The data consist of information on female clerical/administrative (CA) officers employed in the Third Division of the Australian Public Service (APS) in 1969.[8] 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 specialised variety were excluded from the analysis.[9]

Information on this group of workers is available for the years 1969 and 1974, thus allowing an examination of quit and wage determinants over a five year period. The analysis was limited to individuals aged less than 45 in years in 1969 in order to minimise the possibility that separations (identified only by the individual's absence in 1974[10]) 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 realistic. The above limitations restricted the sample size to 5,236. Table I 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-

[6]The absence of an error term in this equation reflects the assumption that the employer forms judgements of the quit probability on the basis only of observable characteristics. This is an appropriate approach in a world of statistical discrimination.

[7]For the derivation of its usual form, see J. Mincer and S. Polachek (1974).

[8]Data were also available for males, but were less rich, so the estimation was confined to females.

[9]For example, Auditors, Computer System Officers, Naval Stores Officers and Interpreter/Translators.

[10]Workers were identified across time periods by numbers, so name changes associated with marriage or separation would not result in an incorrect presumption of a job separation.

schooling equivalents. Appendix I presents information on the distribution of qualification 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 (fairly narrow) Class level of the individual. Neither the schooling or salary adjustments are believed to introduce important distortions.

TABLE I

*Statistical Characteristics of the Data (as of 1969)*

| | Single† Women | | | Married† Women | | |
|---|---|---|---|---|---|---|
| | Mean | | Standard Deviation | Mean | | Standard Deviation |
| Annual Salary ($) (*SAL*) | 3,008 | | 1,065 | 3,127 | | 892 |
| Age (years) (*AGE*) | 22.93 | | 5.65 | 22.37 | | 3.64 |
| Internal Experience* (Years (*OJEX*) | 2.88 | | 4.47 | 2.40 | | 2.70 |
| Education** (Years) (*YOS*) | 12.39 | | 1.04 | 12.28 | | 0.90 |
| *CAN**** | 27.69 | | | 26.40 | | |
| Percent Quitting by 1974 | | 52.45 | | | 50.68 | |
| Observations | | 2,221 | | | 3,015 | |

* Time since joining the CA division.
** See explanation of derivation above.
*** Percentage of workers employed in Canberra.
† As of 1968.

The basic hypotheses of the model were tested in the following way. First, logit quit estimations were run on the entire sample, using the model:

$$P_i = \alpha 0 + \alpha_1 AGE_i + \alpha_2 AGE_i^2 + \alpha_3 YOS_i + \alpha_4 OFF_i + \alpha_5 MAR_i$$
$$+ \alpha_6 OJEX_i + \alpha_7 OJEX_i^2 + \varepsilon \tag{12}$$

where, for individual $i$, $P_i$ is dichotomous (equal to 1 if the individual has left the job by 1974, equal to 0 otherwise), *AGE* is age in years in 1969, *YOS* is years of schooling, *MAR* is a marital status dummy variable (equal to 1 if married in 1968, equal to 0 otherwise) and *CAN* is a location dummy variable (equal to 1 if the individual was employed in Canberra in 1969, equal to 0 otherwise).

The coefficients obtained from the estimation of equation (12) were then used to generate the employers' quit expectation probability for each individual, $\hat{P}$. This approach thus assumes that employers' quit expectations are *ex ante* correct, on average, at least as far as the influence of observed variables is concerned.

In a second stage the generated regressor $\hat{P}$ was used in the 1974 salary equation estimation (11). Because this estimation is restricted to persons who did not quit, the *ex post* quit probability expectation is necessarily incorrect. Interpretation of the coefficient on $\hat{P}$ in the wage equation is thus as follows. If $\frac{\partial lnW}{\partial \hat{P}} < 0$ this implies that non quitting individuals are penalised for possessing observable characteristics that on average are associated with quit tendencies.[11] The model thus allows a test of statistical discrimination, as well as the modified human capital relationships detailed in section I.

Table II presents the results of the logit estimation of equation (12).

TABLE II
*Quit Estimation*†

| Variable | |
|---|---|
| Constant | -4.691 (6.37) |
| *AGE* | 0.451 (8.19) |
| $AGE^2$ | -0.00870 (8.49) |
| *YOS* | -0.00870 (1.48) |
| *CAN* | 0.132 (2.01) |
| *MAR* | -0.200 (3.44) |
| *OJEX* | -0.0687 (4.25) |
| $OJEX^2$ | 0.000659 (1.03) |
| Bookmaker specification test | .0059* |

† Absolute *t*-statistics in parentheses.

* This test statistic is unbounded. If its absolute size exceeds 1.96, this implies the null hypothesis that the model is a good fit of the data is rejected. For this test the model apparently performs well [Beggs (1983)].

[11]For these data, the penalty will be manifested through variations in promotion rates.

Table III presents interpretive estimates of the coefficients, showing the percentage point change in the quit probability associated with one year increases in schooling, age and time-on-the-job, and a one unit increase in office and marital status, calculated at the mean.

TABLE III
*Interpretive Estimates of Quit Coefficients*

| | |
|---|---|
| *AGE* | 0.0143 |
| *YOS* | -0.0118 |
| *OJEX* | -0.0163 |
| *CAN* | 0.0330 |
| *MAR* | -0.0499 |

The coefficients imply the following. The quit probability, *ceteris paribus*, is maximised at 25.9 years of age, and individuals employed in Canberra are about 3 percent more likely to quit than those not in Canberra. Married women are about 5 percent less likely to quit than single women, a finding probably reflecting the former group's greater attachment to the labour force. This result could mean that single women are more likely to quit their jobs after marriage. On-the-job experience and schooling coefficients are similar to those found in alternative tests of turnover.[12] An extra year of experience decreases the quit probability by 1.6 percent at the mean and an extra year of schooling by about 1.2 percent, although the latter is not statistically significantly different from zero.

The coefficients reported in Table 2 were used to generate the regressor $\hat{P}$, interpreted to be the employers' expectation of a quit contingent on the observable variables. That is, the conventional human capital earnings function including $\hat{P}$ was estimated for those remaining in the job in 1974, around 49 percent of the original sample. The statistical characteristics of $\hat{P}$ are presented in Table IV and the earnings function results in Table V.

TABLE IV
*Statistical Characteristics of* $\hat{P}$

| Mean | Standard Deviation | Minimum | Maximum |
|---|---|---|---|
| 0.4993 | 0.1022 | 0.02583 | 0.6727 |

[12]For a review of these studies, see D. Parsons (1977). More recent work confirming these results is in Richard Freeman (1980).

TABLE V
*Earnings Function Estimations*†

| Variables | (i) | (ii) | (iii) |
|---|---|---|---|
| Constant | 7.997<br>(207.04) | 7.989<br>(302.38) | 8.0463<br>(165.83) |
| *YOS* | 0.0678<br>(23.66) | 0.0681<br>(23.67) | 0.0675<br>(23.48) |
| *OJEX* | 0.0132<br>(8.19) | 0.0134<br>(8.27) | 0.00761<br>(6.14) |
| $OJEX^2$ | -0.000154<br>(4.06) | -0.000157<br>(4.13) | -0.000139<br>(3.58) |
| *GEXP* | 0.0174<br>(7.64) | 0.0173<br>(7.57) | 0.0223<br>(6.08) |
| $GEXP^2$ | -0.000500<br>(7.45) | -0.000495<br>(7.35) | -0.000681<br>(5.35) |
| *OFF* | 0.129<br>(21.71) | 0.129<br>(21.73) | 0.131<br>(21.47) |
| *MAR* | | 0.00642<br>(1.15) | 0.00108<br>(0.51) |
| $\hat{P}$ | | | -0.121<br>(1.71) |
| $\bar{R}^2$ | .41 | .41 | .41 |

† Absolute t-statistics in parentheses.

Number of observations = 2543.

These results imply the following. Those individuals with relatively high *ex ante* quit probabilities derived from the turnover model receive relatively low wages even though *ex post* they are revealed not to quit, a result which is confirmed at about the 10 percent level of significance. If the methodology employed in the estimations is correct, this finding confirms (weakly) a basic tenet of statistical discrimination: observable variables are apparently used by the employer as a reflection of individual behaviour presumably because of the ascribed association between individual and group behaviour. This implies that if a person has the characteristics of a group that tends on average to have a high quit propensity, that individual will experience relatively low promotion rates and consequently lower wages.

The results imply that an increase in the expected turnover probability from 0 to 1 over a 5 year period results in a 12.1 percent decrease in wages (about $369 at the mean), and a one standard deviation increase in the expected turnover probability decreases wages by 1.2 percent. In a sense, workers remaining are compensating the employer for those employees of similar observable characteristics who choose to separate from the firm.

Econometrically the statistical discrimination results is are not unambiguously powerful. The overall explanatory power of the equation is not affected by the inclusion of $\hat{P}$, although it is widely recognised that $\bar{R}^2$ is not a very useful test statistic. As well, the standard error on the $\hat{P}$ coefficient is not small, a result that could reflect some collinearity with included regressors, particularly *OJEX*. On the other hand, the stability of other coefficients increases confidence that the finding is of interest.

The human capital coefficients of the earnings functions are similar to those reported in other studies. One year increases in schooling, one-the-job experience and general labour market experience are associated with about 6.8 percent, 1.3 - 0.06 percent and 2.2 - 1.74 percent increases respectively. Being employed in Canberra adds about 13 percent to earnings, an issue addressed more comprehensively in Chapman (1985).[13] It is clear, however, that the inclusion of $\hat{P}$ decreases estimates of rates of return to FST, at least as approximated by the coefficient on *OJEX*.

The results should be treated cautiously because of a conceptual issue concerning workers' motivations. Earnings function estimations implicitly assume that utility maximisation involves earnings maximisation which in essence is an assumption that workers are identical in their motivation to pursue promotions. However, if this is not the case, the error term includes workers' earnings ambitions, a variable which is likely to be associated with actual earnings. This results in biased coefficients if this variable is related to $\hat{P}$.

For the following reasons the association alluded to above is not fanciful. Those workers with observable characteristics typically associated with relatively high quit probabilities may have believed in 1969 that they are unlikely to remain in employment, and consequently pursued promotions less vigorously than persons with greater employment commitment. If they were still employed in 1974, their relatively low earnings may be a consequence in part of their own and not just their employers' action. Thus some part of the association between earnings and expected quit probabilities may be independent of employer statistical discrimination. This implies that $\frac{\partial lnW}{\partial \hat{P}}$ is an upper bound of the extent of the earnings implications of employer misinformation.

## IV. Conclusions

This paper has presented a theoretical model of the FST variety in an attempt to understand the wage consequences of employer expectations of voluntary labour turnover. The model predicts that workers with personal characteristics that tend to be associated with relatively high quit probabilities will receive relatively low wages. The result implies misspecification of the models typically used to estimate earnings determinants.

[13]Note that if the null hypothesis that the coefficient on $\hat{P}$ is zero is rejected, the $t$-statistics on other regressors are incorrect because $\hat{P}$ is a generated regressor. In general, this problem can be solved through bootstrapping (Pagan, 1984), a procedure that is not feasible with such a large sample.

Econometric investigation of the relationships support the model, albeit stastically not strongly, for a large group of female Australian government employees in the 1969-74 period.

The results offer support for models of statistical discrimination in that they suggest the relevance to wages of employer's ascribing to individuals quit probability expectations derived from *ex ante* expectations of groups. The important qualification to the results is that some part of the relationship is, in theory, attributable to workers making errors concerning their expected job tenure.

## Appendix I

*Education Qualification of the Sample*

| Qualifications | Per Cent of Sample | Assumed Year-of-Schooling Equivalent |
|---|---|---|
| Matriculation | 78.18 | 12 |
| Diploma | 5.06 | 14 |
| Bachelor – Ordinary | 13.80 | 15 |
| Bachelor – Honours | 2.37 | 16 |
| Masters | .53 | 18 |
| Doctorate | .06 | 20 |

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