An Analysis of the Origins of Sex Differences in Australian Wages

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Different levels of measured skills, geographic location and demographic factors (such as marital status and country of birth) explain almost none of the hourly wage differences of Australian women and men in full-time employment. The major contribution to wage differences is apparently in the different returns paid by employers to men and women for observable characteristics. Usually this is considered as evidence for the existence of direct wage discrimination by employers but—at least for the data of this study—some questions remain as to the extent of this influence. Measurement issues related to both schooling and general labour market experience tend to exaggerate the role of direct employer discrimination, but even extreme assumptions as to the extent of mismeasurement of these variables do not eliminate such discrimination (although it is reduced from 13 to 3.5 per cent). Some part of the overall difference is likely to be a consequence of (unmeasured) occupation factors and, of much greater complexity, the possible influence of role-stereotyping and its interaction with family arrangements.

This paper investigates the determinants of the male–female wage differential. Until recently little has been known of the determinants of wage differences between women and men in Australia. Broadly based investigations have identified average wage differences by sex (Power 1973), and several analyses have used regression techniques to ascertain the influence of particular variables (Jones 1982; Chapman & Miller 1983; Chapman 1984; Haig 1982). This research, however, typically suffers from problems of specification and inadequate data. Consequently, strong conclusions concerning the nature and extent of so-called 'discrimination' are premature at this stage, as are calls for particular policy initiatives to remedy the situation.

This paper reports an empirical examination of an extremely rich and so far unutilized data set that provides new information on the origins of wage differences. A commonly used reduced-form econometric framework is employed to determine the separate contribution of a host of variables to average wage differences between men and women. The novelty and useful-

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ness of the exercise lies in the broadness of coverage of the data. The results imply that it is the differential employer treatment of women and men that explains sex differences in wages—measured variable differences between the sexes are seemingly unimportant.

After a brief literature review, the theoretical model typically employed in accounting for sex difference in wages is described and the way this framework may be used in an analysis of the data available is discussed. The characteristics of the data, reduced-form estimates of the model and the accounting for the origins of wage differences are then presented, and finally some qualifications are offered.

Literature review, the theoretical model and econometric methodology

Several analyses of sex differences in earnings in Australia have been undertaken, but they have important limitations. The Jones (1981) and Chapman and Miller (1983) investigations, for example, both used the 1976 census, which had information on income rather than earnings. Moreover, these data do not include information on regional distribution or the length of time individuals had been employed on the current job. Also, Jones used hours worked as an independent variable, a strategy implying specification error, as labour supply theory predicts that wage rates determine both hours and earnings. This endogeneity problem suggests interpretation difficulties. The Haig (1982) analysis also assumed hours to be exogenous and, further, used age as a proxy for experience. This latter specification necessarily biases the schooling coefficient towards zero, since individuals with high levels of schooling for given age will have been in the labour force for shorter periods and will thus have lower earnings. The issue is discussed formally in Rosenzweig and Morgan (1975).

Chapman's (1984) discussion of sex difference in earnings in the Australian public service is of little relevance for broadly based policy questions because it examines public sector employees only. And even within this category the analysis is narrowly defined because of the focus on persons employed in what is probably a single occupation, clerical and administrative officers of the third division.

Clearly, then, existing research suffers from problems of both methodology and coverage. The present investigation avoids these limitations, a point that becomes clear following the discussion of theoretical and econometric issues.

Investigations of sex (and race) wage differentials have typically used the Oaxaca (1973) methodological framework, which purports to break down overall differences in remuneration into productivity levels and returns. This method employs the earnings function, a reduced-form equation of the human capital genre, which may be developed as follows (Mincer & Polachek 1974):

$$E_t = E_{t-1} + rC_{t-1} \quad (1)$$

where E_t is gross earnings in period t, C_{t-1} is the dollar amount of net investment in period t-1 and r is the average rate of return to the individual's investment in human capital. If the ratio of investment expenditures to gross earnings, C_t/E_t, is given by k_t, then, by viewing investment in time-equivalent units:

$$E_t = E_{t-1}\ (1+rk_{t-1}) \tag{2}$$

Since $E_t = E_0\ (1+rk_0)\ (1+rk_1)\ \ldots\ (1+rk_{t-1})$, a logarithmic approximation of $\ln(1+rk) \doteqdot rk$ yields

$$\ln E_t = \ln E_0 + r \sum_{i=0}^{t-1} k_i \tag{3}$$

Analysing schooling and post-school experience, it is possible to separate the k terms, giving

$$\ln E_t = \ln E_0 + r \sum_{i=0}^{s-1} k_i + r \sum_{j=s}^{t-1} k_j \tag{4}$$

where k_i and k_j are respectively investment ratios during and after the schooling period. Assuming $k_i = 1$,[1]

$$\ln E_t = \ln E_0 + rs + r \sum_{j=1}^{t-1} k_j \tag{5}$$

Given that post-school investments are expected to decline over the lifetime (as retirement approaches, the expected return from investment falls), equation (5) may be approximated with the inclusion of a quadratic experience term. Thus the estimating equation becomes:

$$\ln W_i = a + b\ S_i + c\ GEXP_i + d\ GEXP_i^2 + e\ Z_i + \epsilon \tag{6}$$

where, for individual i, ln W is the logarithm of hourly wages, S is years of schooling, GEXP is length of time in the labour force and Z is a vector of other wage-determining variables. In this study, Z includes length of time in the current job (tenure), occupation, industry, region, union status, marital status and country of birth. This conventional (Oaxaca 1973; Corcoran & Duncan 1979; Jones 1981) specification implicitly makes several limiting assumptions noted below.

First, the coefficients on many of these variables are difficult to interpret in a single equation wage estimation, as the correct framework is likely to be a structural model treating them in a choice context. In particular, occupation, industry and, to a lesser extent, region, are likely to be influenced by years of schooling. As well, choices in some of these areas may be influenced by wage expectations. Ignoring these possibilities and estimating the reduced form is an accepted, albeit lamentable, practice in wage estimations, as theory is poorly developed *vis-ā-vis* a multi-equation system. The problems implied by this approach are unlikely to be severe, given the focus on average differences by sex, but recognition of the limitation is warranted in interpretation of results.

A second, and possibly more severe, limiting assumption of the model is the treatment of length of time in the current job as exogenous to wages. There are good reasons to believe that the firm's wage structure influences labour

1. This is equivalent to the assumption that tuition equals student earnings and scholarship.

turnover and thus tenure, an issue analysed in Beggs and Chapman (1985a, 1985b). They used an instrumental variables approach to correct this problem for data on Australian public servants. While their evidence is consistent with the view that tenure is endogenous in the wage equation, the correctly specified model had virtually no effect on decompositions of male-female wages (Beggs & Chapman 1985b). Since the decomposition is the purpose of the current exercise, we are relatively unconcerned with the simultaneity issue, although it is important to recognize its potential to undermine the validity of the approach.

Third, the inclusion of some, although not all, of the control variables is difficult to justify from economic theory. Labour economics has little to say concerning the influence of marital status on wages, for example. However, ethnicity can be taken to represent language or information problems;[2] region probably reflects local prices and other compensating differentials; and union membership is likely to proxy wage bargaining power.[3] Nevertheless, an integrated theoretical approach is lacking, which is more an indictment of the state of the art than it is a shortcoming specific to this analysis. Faced with this problem, we chose to estimate a relatively unrestrictive model rather than to impose zero coefficients on these variables.

A final comment concerning methodology is to note that the wage returns to occupation are implicitly constrained to be identical across industries, and vice versa. Indeed, the absence of interactions of other exogenous variables imposes similar restrictions. This implies that only simple levels effects are being identified. Because of the richness of the data, a large number of alternative specifications are possible. Simplicity of interpretation of average relationships has been accorded priority, an aggregation that may hide the true form. With these caveats in mind it is now appropriate to explain how equation (6) may be used in an understanding of male-female wage differentials.

By separately estimating the reduced form for each sex, some insight is gained into the following possibilities: first, that women earn less than men because they have lower measured skills; second, that employers reward variables differently depending on employees' sex; third, that women tend to be employed in low-paying jobs, as identified by union status, occupation and industry; and fourth, that women are disadvantaged in employment owing to 'demographic' factors such as country of origin, marital status and geographic location.

The average female wage can be predicted for four different scenarios. These are that women, all other things being equal, have the average (measured) productivity endowments of men; are paid the same reward for their observed variables as men are; are distributed across jobs (as determined by occupation, industry and union membership) in the same way that men are; and have the same regional, marital status and nativity distributions that men do. We refer

2. The empirical work of Stromback (1984), among others, supports this contention in the Australian context.
3. The union-non-union wage differential issue has been investigated substantially in the American literature. For an Australian analysis, see Mulvey (1986).

to these predictions respectively as wage differences attributable to endowments, coefficients, jobs and demography. Following Blinder (1973), coefficient differences are seen as indicators of employer sex discrimination. These decompositions may be illustrated in an alternative way.

If the regressions for men and women are given by:

$$ln\ W_i^m = \beta_o^m + \sum_{h=1}^{n} \beta_h^m X_{hi}^m + \sum_{k=1}^{l} \beta_k^m J_{ki}^m + \sum_{p=1}^{q} \beta_p^m D_{pi}^m + \epsilon_i^m \quad (7)$$

and

$$ln\ w_i^w = \beta_o^w + \sum_{h=1}^{n} \beta_h^w X_{hi}^w + \sum_{k=1}^{l} \beta_k^w J_{ki}^w + \sum_{p=1}^{q} \beta_p^w D_{pi}^w + \epsilon_i^w \quad (8)$$

where X_{hi} is a vector of productivity endowments, J_{ki} is the distribution of persons across particular jobs, and D_{pi} is the distribution of persons across particular demographic characteristics, then

$$W_e = \text{antilog}\ (\beta_o^w + \sum_{h=1}^{n} \beta_h^w X_h^m + \sum_{k=1}^{l} \beta_k^w J_k^w + \sum_{p=1}^{q} \beta_p^w D_p^w)$$

$$W_c = \text{antilog}\ (\beta_o^m + \sum_{h=1}^{n} \beta_h^m X_h^w + \sum_{k=1}^{l} \beta_k^m J_k^w + \sum_{p=1}^{q} \beta_p^m D_p^w)$$

$$W_j = \text{antilog}\ (\beta_o^w + \sum_{h=1}^{n} \beta_h^w X_h^w + \sum_{k=1}^{l} \beta_k^w J_k^m + \sum_{p=1}^{q} \beta_p^w D_p^w)$$

and

$$W_d = \text{antilog}\ (\beta_o^w + \sum_{h=1}^{n} \beta_h^w X_h^w + \sum_{k=1}^{l} \beta_k^w J_k^w + \sum_{p=1}^{q} \beta_p^w D_p^m)$$

where W_e, W_c, W_j and W_d are predicted average female wages attributable to endowments, coefficients, job distributions and demographic characteristics respectively, $\bar{X}_h^m$ and $\bar{X}_h^w$ are average male and female skill endowments; β_h^m and β_h^w are male and female endowment coefficients; β_k^m and β_k^w are average male and female job coefficients; $\bar{J}_k^m$ and $\bar{J}_k^w$ are average male and female job distributions; β_p^m and β_p^w are male and female demographic coefficients; and $\bar{D}_p^m$ and $\bar{D}_p^w$ are average male and female demographic characteristics. The estimated average wages can be compared with observed wages in an attempt to ascertain the probable influence of these four factors on the earnings differential. In other words, the greater are W_e, W_j and W_d compared with the actual female average wage, the greater is the contribution of male-female disparities in measured human capital, job distributions and demography respectively to an explanation of the wage gap between the sexes. Similarly, the greater is W_c compared with the actual female

average wage, the greater is the contribution of differential employer treatment of men and women to an explanation of the wage gap between the sexes.[5]

The data, estimations and statistical breakdowns

The data set used was gathered by the ABS in the course of Special Supplementary Survey No. 4 (SSS4) in 1982 and involved about 30 000 randomly selected households throughout Australia. Some adjustments were employed to allow implementation of the methodology outlined below. First, potential experience (PE) was approximated by the formula (age−schooling −5), which assumes that schooling is obtained before work begins, that individuals begin work immediately after finishing school, and that employment is continuous once commenced. These assumptions are clearly questionable for the sample. In particular, the distortions are likely to be greater for women, given this group's more intermittent labour force participation (an issue addressed in the statistical analysis).

Second, levels of education attained were converted into full-time years of schooling equivalents. The assumptions used for this conversion are noted in appendix 1.

Third, hourly wage was not directly available from the survey. This variable was constructed by dividing weekly earnings by weekly hours.

The analysis was restricted to full-time employees working in only one job, of which there were 8946 men and 4330 women. Appendix 2 presents the statistical characteristics of the data, where job tenure is number of years in the current job and persons not living in a city are allocated residency in their particular state.

Estimations of equations (7) and (8) were carried out, where UNION is a dummy variable equal to 1 if the individual belonged to a union, equal to 0 otherwise, and the following dummies were omitted to allow estimation: transport/service industries; 'separated' marital status; 'other' country of birth, living in South Australia outside Adelaide, and 'other' occupations. The results are presented in tables 1 and 2.

The coefficients reported in tables 1 and 2 may be used to calculate the contribution to changes in hourly wages given variations in the independent variables. The percentage change in the dependent variable is shown in table 3 for one-year increases in schooling, time in the labour force, and time in the current job (calculated at the mean), and for unit changes in the dummy variables.

5. An analogous exercise could have been undertaken using female endowments evaluated by the male equation and comparing results with the average male wage. The choice is essentially arbitrary and while results will differ to a small extent because of the index number problem, we followed the convention established by Blinder (1973) and practised by, among others, Jones (1981).

Table 1 OLS wage estimation: women

Variable	*Coefficient*	*Variable*	*Coefficient*
Intercept	0.894**	*Country of birth*	
Union	0.0604**	UK or Ireland	0.0249
Education		Australia	0.0141
Schooling	0.0402**	Italy	0.0287
Experience		Greece	−0.0389
Job tenure	0.0114**	Yugoslavia	−0.0284
Job tenure²	−0.000148**	Holland	−0.0102
Potential experience	0.0273**	West Germany	0.0582
Potential experience²	−0.000554**	*Place of residence*	
Industry		Sydney	0.0650*
Agriculture	−0.452**	Melbourne	0.0652*
Mining	0.194**	Brisbane	0.0202
Food	0.0487	Hobart	0.0585
Textiles	−0.0748**	Darwin	0.175**
Wood	−0.0878	Perth	0.0166
Chemicals	0.0689	Canberra	0.0942**
Metals	−0.00101	Newcastle	0.103
Electricity	0.130*	Adelaide	0.0406
Construction	−0.0194	New South Wales	−0.0104
Retail	−0.0690**	Victoria	0.00888
Transport	0.0219	Tasmania	−0.0260
Communications	0.0893**	Western Australia	0.0128
Finance	0.00397	Northern Territory	0.108
Public administration	0.0951**	Queensland	−0.0143
Community services	−0.0257	*Occupation*	
Entertainment	−0.0666**	Professional	0.231**
Marital status		Administrative	0.133**
Married	0.00666	Clerical	0.0968**
De facto	0.0506**	Sales	−0.00805
Widowed	0.00407	Farmer	−0.067
Single	−0.0801**	Tradesperson/labourer	−0.00226

Notes: $R^2 = 0.40$

** Significant at 0.05 per cent level.

* Significant at 0.10 per cent level.

Table 2 OLS wage estimation: men

Variable	*Coefficient*	*Variable*	*Coefficient*
Intercept	0.929**	*Country of birth*	
Union	0.0816**	UK or Ireland	0.0735**
Education		Australia	0.0528**
Schooling	0.0452**	Italy	−0.0537**
Experience		Greece	−0.0163
Job tenure	0.00956**	Yugoslavia	−0.0630**
Job tenure²	−0.000122**	Holland	0.0168
Potential experience	0.0252**	West Germany	0.0242
Potential experience²	−0.000452**	*Place of residence*	
Industry		Sydney	0.0737**
Agriculture	−0.227**	Melbourne	0.0516**
Mining	0.350**	Brisbane	0.0488**
Food	−0.0255	Hobart	−0.00105
Textiles	−0.0803**	Darwin	0.0223**
Wood	−0.121**	Perth	0.0180
Chemicals	0.0344	Canberra	0.114**
Metals	−0.0177	Newcastle	0.119**
Electricity	0.0748**	Adelaide	0.00219
Construction	0.0852**	New South Wales	0.0185
Retail	−0.112**	Victoria	−0.00957
Transport	−0.0564**	Tasmania	−0.0350
Communications	−0.00407	Western Australia	0.0438*
Finance	−0.0392*	Northern Territory	0.111*
Public administration	−0.00155	Queensland	0.0245
Community services	−0.115**	*Occupation*	
Entertainment	−0.112**	Professional	0.192**
Marital status		Administrative	0.193**
Married	0.0700**	Clerical	0.108**
De facto	0.0650**	Sales	0.0251
Widowed	0.0917*	Farmer	−0.166**
Single	−0.0639**	Tradesperson/labourer	−0.00862

Notes: $R^2 = 0.43$

** Significant at 0.05 per cent level.

* Significant at 0.10 per cent level.

Table 3 Percentage change in wage given changes in the independent variables

Variable	*Women*	*Men*	*Variable*	*Women*	*Men*
Union	6.04	8.16	*Country of birth*		
Education			UK or Ireland	2.49	7.35
Schooling	4.02	4.52	Italy	2.87	−5.37
Experience			Australia	1.41	5.28
Job tenure	1.00	0.77	Greece	−3.89	−1.63
Potential experience	1.08	0.80	Yugoslavia	−2.84	−6.30
Industry			Holland	−1.02	1.68
Agriculture	−45.20	−22.70	West Germany	5.82	2.42
Mining	19.40	35.00	*Place of residence*		
Food	4.87	−2.55	Sydney	6.50	7.37
Textiles	−7.48	−8.03	Melbourne	6.52	5.16
Wood	−8.78	−12.10	Brisbane	2.02	4.88
Chemicals	6.89	3.44	Hobart	5.85	−0.11
Metals	−0.10	−1.77	Darwin	17.50	22.30
Electricity	13.00	7.48	Perth	1.66	1.80
Construction	−1.94	8.52	Canberra	9.42	11.40
Retail	−6.90	−11.20	Newcastle	10.30	11.90
Transport	2.19	−5.64	Adelaide	4.06	0.22
Communications	8.93	−0.41	New South Wales	−1.04	1.85
Finance	0.40	−3.92	Victoria	0.89	−0.96
Public administration	9.51	−1.55	Tasmania	−2.60	−3.50
Community services	−2.47	−11.50	Western Australia	1.28	4.38
Entertainment	−6.66	−11.20	Northern Territory	10.80	11.10
			Queensland	−1.43	2.45
Marital status			*Occupation*		
Married	0.67	7.00	Professional	23.10	19.20
De facto	5.06	6.50	Administrative	13.30	19.35
Widowed	0.41	9.17	Clerical	9.68	10.81
Single	−8.01	−6.39	Sales	−0.81	2.51
			Farmer	−6.72	−16.64
			Tradesman/labourer	−0.23	−0.86

The results of table 3 suggest the following. Women union members earn about 6 per cent more than women non-union members. For women, an additional year of schooling adds about 4 per cent to the hourly wage, and additional years of experience in the current job and in the labour force are associated with about 1 and 1.1 per cent increases in hourly wage respectively. Women in de facto relationships earn about 5 per cent more than separated, married and widowed women, and single women earn about 8 per cent less than women in these last three categories. For women there are virtually no differences in earnings by country of birth, but residency in cities is associated with relatively high earnings (in particular, in Darwin, Canberra, Sydney and Melbourne). Women in professional, administrative and clerical occupations earn relatively more than women employed in sales, trades or farm occupations. Industries paying relatively high hourly wages for women are mining, electrical goods, commerce and the public sector, while retail trade, wood, agriculture, textiles and entertainment pay relatively low hourly wages.

With a few exceptions, these results are similar to those obtained for men. Union members receive 8 per cent higher hourly wages, an additional year

of schooling adds 4.5 per cent to earnings and additional years of time in the current job and in the labour force both add about 0.8 per cent to hourly earnings. Married, widowed and de facto men do well, and single men poorly, relative to divorced and separated men. Men born in the United Kingdom or Ireland, and Australia, receive 5–7 per cent higher earnings than 'others'; men born in Italy and Yugoslavia receive about 5 and 6 per cent less respectively. Men living in cities also earn relatively more, and this is particularly true for residents of Darwin, Canberra, Newcastle, Sydney and Melbourne. Men employed as professionals, administrators or clerical workers earn relatively more, and trades and farm workers earn relatively less.

The coefficients reported in tables 1 and 2 were used to calculate W_e, W_c, W_j and W_d. These data are presented in table 4.

Table 4 Predicted female hourly wage under various scenarios ($)

Observed average male wage	*Observed average female wage*	*We*	*Wc*	*Wj*	*Wd*
6.68	5.65	5.78	6.44	5.44	5.64

The results shown in table 4 imply the following. If women had the same schooling and experience levels as men, they would have earned W_e, which is 2.27 per cent higher wages than they did earn, or 14.45 per cent less than men. If women had been rewarded for their measured human capital levels and other variables in the same way as men (in other words, received the male coefficients), then, other things being equal, they would have earned W_c, which is 13.07 per cent more than they did earn, or 5.14 per cent less than men. If women had been distributed across jobs in the same way as men, then, other things being equal, they would have earned W_j, which is 3.79 per cent less than they did earn, or 21.93 per cent less than men. Finally, if women had the same demographic characteristics as men, then, other things being equal, they would have earned the same as they did earn.

The clear conclusion of the above results is that holding constant industry, occupation, marital status, union status, country of birth and geographic location does not change the estimate of the average wage difference between the sexes. What appears to be of substantial importance is the returns paid to particular levels of the independent variables (plus the intercept term). This coefficient interpretation is usually interpreted as evidence for employer discrimination. The extent to which the result can be justified on this basis alone is complex, however, and is considered in some depth below.

Problems in interpretation of the results

There are several problems inherent in using the results of this exercise as indicative of direct wage discrimination by employers. The first category of issues concerns measurement inadequacies, a possibility that is particularly important if the data understate the true productivity of males relative to women. This may be the case for schooling and is highly likely to be the situation for the measurement of general labour market experience. They will be considered in turn.

Using estimated years of schooling as a reflection of formal general training introduces a bias if men attend higher quality education institutions than women, or if male tertiary education of equivalent length to female tertiary education is of a form that results in greater work productivity. The former case is unlikely in a country such as Australia where the variance of school quality is low relative to the United States, for example, where both secondary and tertiary education qualities differ substantially. Even in the United States correction for institution quality has only a minor influence in explaining earnings disparities (see Griliches 1976). Thus it is difficult to argue that significant problems are inherent in Australian studies incorporating no quality adjustments for this variable.

Another possibility *vis-a-vis* schooling is that males undertook formal training that was of higher specific vocational value than was the case for females (for example, a law degree compared with a more broadly based university education). While this may be true for particular individuals, for the issue at hand the question is the empirical importance of this possibility on average. Given that 38 per cent of females undertook post-secondary education, an extreme assumption is that this factor results in an overstatement of true female schooling productivity of 10 per cent. This implies that, on average, female tertiary education is worth two-thirds of male tertiary education of equivalent length. It is reasonable to interpret this assumption as understating the true value of female education, a point illustrated in appendix 3. If the adjustment is extreme, it allows insight into a lower bound of W_C for this component.

The above assumption resulted in a revised W_C of \$6.10, which is 7.67 per cent higher than the actual average female wage and 9.08 per cent lower than the actual average male wage. Thus the (probably extreme) adjustment for the possible lower labour market productivity of female education decreases the estimate of W_C by about 5.5 percentage points (from 13 to 7.5), or by about half.

A further complication related to the schooling variable is the possibility that female students are accorded less attention than male students, a factor that may result in their receiving relatively low education quality. No adjustment is incorporated for this factor. It should be noted that, if it exists, it represents a societal barrier to females reaching in full their employment potential. In this sense it is discrimination, but it should not be classified within the general rubric of direct employer prejudice.

Perhaps the most complicated problem relates to the measurement of general labour market experience. The use of the formula (age − schooling − 5) must overstate the actual labour market experience of those groups with relatively intermittent labour force participation (in this case, women). If that is the case, the estimate of W_C will be biased upwards since there is an exaggeration of relative female labour market productivity. Consequently the extent of the earnings difference interpreted in the context of employer discrimination (i.e. different pay by sex holding constant productivity) will be over-estimated.

A possible solution to this bias problem is to adjust (age − schooling − 5) with an estimate of α, the proportion of time since leaving school that

persons actually spend in the full-time labour force. For the estimations of this paper the following adjustment has been used. It is assumed that:

$$ij = ER_{ij}$$

where ER_{ij} is the employment rate for group i in year j. Given that employment rates have varied over time (particularly for women), it is important to adjust measured experience to reflect the ER of persons of the relevant sex and age group for the year in which they were that age, and not to impose 1982 cross-sectional ERs. This strategy results in the following formula for adjusted experience:

$$AE_{kni} = ER_{ji} + ER_{(j+1)i} + ER_{(j+2)i} + \ldots + ER_{ni}$$

where AE_{kni} is the accumulated experience of individual k of group i in year n, and ER_{ji} is the employment rate of group i in the year in which individual k was aged j, the first year since the completion of formal schooling.

For the purposes of recalculation of W_C, it is necessary only to adjust the average level of female PE, rather than to re-estimate the model.[6] Assuming an average level of schooling of 12 years, and an average level of PE of 15 years, it is possible to recalculate the 'corrected' average general labour market experience of women from the 1976 census[7] (see Chapman & Miller 1983). The adjusted female experience level is 7.20 years, which results in a revised W_C of \$6.18, which is 8.96 per cent higher than the actual female average wage and 7.78 per cent lower than the actual average male wage. Thus the adjustment for female general labour force experience decreases the estimate of W_C by about 4 percentage points (from 13 to 9), or by about one-third. This estimate is very close to that obtained by Borjas (1984) using a similar adjustment on United States data.

The experience adjustment introduces a bias of its own, for two reasons. First, women in the labour force at one point in time are likely to experience employment rates greater than those of the female population in general. If this is the case, AE systematically understates the true experience of female workers. Second, a similar adjustment has not been incorporated for male potential experience. Thus it is implicitly assumed that, for males, (age − schooling − 5) accurately reflects male labour force experience. The small bias inherent in this approach serves also to understate relative female experience. These possibilities are an advantage of the adjustment, as it provides a lower bound for relative female experience, and thus are likely to result in W_C estimates that are biased downwards in absolute size. This helps identify the lower bound for W_C.

6. This is not the case for W_e, W_j and W_d, since mismeasurement of PE leads to upwardly biased estimates of the coefficient on PE in the earnings function (see Chapman & Miller 1983). Adjustment of PE would result in some decrease of these estimates. Given that the role of endowments is apparently insubstantial and that further correction would further diminish this role, the exercise is left to future analysis.
7. This approach thus assumes that women observed in 1982 experienced the same employment ratios as women of equivalent age in 1976.

Incorporating the schooling and general market experience adjustments to form a revised estimate of W_C results in a new W_C of $5.85, which is 3.48 per cent higher than the actual average female wage, and 13.25 per cent lower than the actual average male wage.

A second set of issues concerning the relative size of W_C is that the occupational groupings are very broad. It is likely that the estimate of the role of job type would increase given more disaggregated information. To illustrate this point, note that, for example, nurses, doctors and university vice-chancellors are all included in the 'professionals' category, a procedure that results in an under-estimate of the role of job segmentation in explaining average wage differences, and an over-estimate of W_C, since unmeasured occupational wage differences will be manifested in the intercept term. It follows that it is naive to interpret the results as strong evidence against the role of broadly based forces influencing occupational choice or occupational allocation. More disaggregated investigations have found that occupational concentrations have a substantial influence on earnings differences between the sexes (Jones 1981). Although the present study does not support this view, the aggregated nature of the analysis implies that estimates of W_j are likely to be under-stated.[8]

Third, using the earnings function technique implicitly equates the utility maximization assumption with that of earnings maximization. But this is a strong assumption in a world in which particular groups may have different job aspiration or family-related constraints. If, for example, full-time working women still bear disproportionate family responsibilities they may be less interested in or less able to pursue promotions. This possibility will be manifested in lower average wages for women, other things being equal, for factors that are beyond the direct control of individual employers.

Appendix 1
The Variable 'Schooling'

Schooling is measured by a continuous variable indicating the number of years a person has spent in formal schooling. Years spent in primary and secondary schooling were directly measured in the survey. Years spent in post-secondary schooling were not directly measured in the survey but, instead, the higher educational qualification (including an apprenticeship) was recorded for each individual. It is possible to attribute an approximate number of years of study to each such qualification. The basic conventions used were taken from Chapman and Miller (1983), hence diplomas and bachelor's degrees were attributed three years each while higher degrees were attributed five years. The schooling component of an apprenticeship was estimated to be three months. This was intended to measure the 'schooling' element, as opposed to the on-the-job training element, which is proxied for all workers by the experience variables. Those individuals with none of the qualifications listed were attributed a value of 0 for post-secondary schooling.

8. An indirectly related issue to that of occupational groupings is the difference between job estimated male-female earnings differences. These are presented in appendix 3 below, where the numbers are the coefficients on sex dummy variables (equal to 1 for males, 0 for females) estimated for the full model separately by occupation.

Appendix 2
Statistical Characteristics of the Data

	Men		*Women*	
Variable	*Mean*	*Standard deviation*	*Mean*	*Standard deviation*
Hourly wage ($)	6.68	1.53	5.65	1.46
Schooling (years)	11.44	2.20	11.93	2.34
Job tenure (years)	7.60	8.67	4.60	5.63
Potential experience (years)	19.02	13.52	14.61	12.67
	Industry distribution		*Industry distribution*	
	Number	*%*	*Number*	*%*
Agriculture	241	2.69	38	0.87
Mining	294	3.29	26	0.60
Food	388	4.34	100	2.31
Textiles	95	1.06	169	4.36
Wood	199	2.22	32	0.74
Chemicals	263	2.94	63	1.45
Metals	1197	13.38	203	4.69
Electricity	249	2.78	20	0.46
Construction	682	7.62	42	0.97
Retail	1470	16.43	773	17.85
Transport	646	7.22	102	2.36
Communications	268	3.00	98	2.26
Finance	584	6.53	593	13.69
Public administration	569	6.36	259	5.98
Community services	1012	11.31	1386	32.00
Entertainment	297	3.32	238	5.50
Other	224	2.30	150	1.68
	Marital status distribution		*Marital status distribution*	
	Number	*%*	*Number*	*%*
Married	5797	64.79	2153	49.71
De facto	322	3.60	237	5.47
Widowed	46	0.51	88	2.03
Single	2575	28.78	1655	38.21
Separated	206	2.30	197	4.55

	Country of birth distribution		*Country of birth distribution*	
	Number	*%*	*Number*	*%*
Australia	6408	71.62	3190	73.66
UK or Ireland	974	10.89	452	10.44
Italy	274	3.06	84	1.94
Greece	83	0.93	61	1.41
Yugoslavia	149	1.67	73	1.69
Holland	89	0.99	18	0.42
West Germany	98	1.10	46	1.06
Other	871	9.74	406	9.38
	Occupational distribution		*Occupational distribution*	
	Number	*%*	*Number*	*%*
Professional	1230	13.78	923	21.31
Administrative	691	7.72	101	2.33
Clerical	971	9.74	1737	40.11
Sales	458	5.12	358	8.28
Farmer	296	3.31	35	0.87
Tradesman/labourer	4245	47.45	536	12.38
Transport/service	1155	12.91	640	14.79
	Residency distribution		*Residency distribution*	
	Number	*%*	*Number*	*%*
Sydney	1566	17.50	893	20.62
Melbourne	1359	15.19	817	18.86
Brisbane	762	8.52	394	9.10
Hobart	249	2.78	110	2.54
Perth	900	10.06	410	9.47
Darwin	70	0.78	38	0.87
Canberra	216	2.41	126	2.91
Newcastle	47	0.53	17	0.39
Adelaide	926	10.35	473	10.92
New South Wales	759	8.48	257	5.93
Victoria	429	4.79	179	4.13
Tasmania	364	4.07	132	3.05
Western Australia	329	3.68	120	2.77
N. Territory	33	0.37	16	0.37
Queensland	686	7.67	261	6.03
South Australia	251	2.81	87	2.01

	Union membership		*Union membership*	
	Number	*%*	*Number*	*%*
Union	4910	54.88	2124	49.05
Non-union	4036	45.12	2206	50.95
Total	8946		4330	

Appendix 3
Illustrating the extremity of the schooling adjustment assumption

To gain some insight into the possible extremity of this assumption, note the following. In 1983,[9] 182 000 employed males possessed degrees in law, humanities, education or science, respectively numbering 25 800, 45 000, 38 400 and 73 600. For employed females the numbers were 125 800 (total), 4000, 46 500, 54 600 and 20 700. Making law and science degrees numeraires of four and three years tertiary education respectively, and assuming that a humanities degree is as valuable as a degree in education, it is possible to calculate the time-value of the female-dominated degrees of humanities and education implied by the two-thirds assumption.

That is, the unadjusted value of women's relative education expressed in tertiary years assuming a four-year law degree and three-year humanities, education and science degrees is [4000.4 + (46 500 + 54 600 + 20 700).3] = 381 400. The equivalent figure for men is [25 800.4 + (45 000 + 38 400 + 73 600).3] = 574 200. This ratio is 0.6642 which, under the two-thirds assumption is adjusted downwards to two-thirds or 0.4428. It is now possible to calculate the time equivalence of humanities and education degrees (x), since

$$\frac{(46\,500 + 54\,600)x + 3(20\,700) + 4(4000)}{(45\,000 + 38\,400)x + 3(73\,600) + 4(25\,800)} = 0.4428$$

Therefore x = 1.02 years

In other words, under the assumptions noted, a humanities or education degree is accorded the equivalent of 34 per cent of the value of a science degree. We consider this to be fairly low and, because women's tertiary qualifications are concentrated in the former categories, this reinforces the opinion that the value of the schooling adjustment is severe.

9. Australian Bureau of Statistics, *Labour Force Status and Educational Attainment*, Catalogue No. 6235.0, February 1983.

Appendix 4
Sex Dummy Coefficients by Occupation*

Occupation	*Coefficient* %
Professional	5.29
Administrative	20.65
Clerical	13.37
Sales	10.25
Farmer	28.82
Tradesman/labourer	15.63
Other	16.17

Note: * The numbers represent the percentage wage difference between males and females by occupation, as estimated by the inclusion of a sex dummy variable in the general model run separately by occupation.

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