The contented Australian female worker: Paradox lost, paradox found

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Abstract

In a seminal 1997 paper, Andrew Clark observed that British women report higher job satisfaction than their male counterparts, despite generally holding inferior jobs. To become known as the 'paradox of the contented female worker', Clark argued this was due to women having lower expectations, and that the phenomenon would disappear as women's positions in the labour market improved, a prediction supported by later evidence. This paper draws on data from the Household, Income and Labour Dynamics in Australia survey to investigate how the differential in women's job satisfaction relative to men's evolved in Australia between 2001 and 2022. Regression models suggest that a substantial job satisfaction premium for women gradually diminished over the first decade of this century. Unlike in Britain, however, it re-emerged and remained relatively constant from around 2013. Oaxaca-Blinder decompositions show the job satisfaction premium for women is primarily attributable to differences in the effects of variables, rather than differences in the mean characteristics of male and female workers or of their jobs. Changes in preferences relating to working hours and the effects of educational attainment on job satisfaction have particularly shaped the evolution of differences in job satisfaction by gender. Despite a convergence in the raw means of men's and women's job satisfaction assessments in recent years, the paradox of the contented female worker appears to be alive and well in the Australian labour market.

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Introduction

In a seminal 1997 paper, Andrew Clark observed that British women report higher levels of job satisfaction than their male counterparts, despite women generally holding jobs inferior to men's when assessed by objective standards, such as pay. To become known as the 'paradox of the contented female worker' (Bender, Donohue and Heywood 2005), this observation has since been the subject of scrutiny across numerous countries and over time, with some contention as to how widely the paradox holds. Job satisfaction studies based on data from the Household, Income and Labour Dynamics in Australia (HILDA) survey confirm that the contented female worker paradox applied in Australia in the 2000s and persisted well into the 2010's (Buchler and Dockery 2023; Kifle 2014; Kifle, Kler and Shankar 2014a; Long 2005).

Clark's (1997) preferred explanation for the paradox was that women had lower expectations for their jobs due to the disadvantage they had historically faced in the labour market, and that individuals assess job satisfaction, in part, relative to their expectations. Hence, Clark proposed the higher job satisfaction reported by women may be a transitory phenomenon, one which would disappear once women's relative labour market position improved to be more equal with men's. Clark's initial analysis was based on the first wave of the British Household Panel Survey (BHPS), taken in 1991. With the benefit of BHPS data through to 2014, Green *et al*. (2018) revisited Clark's prediction to find, indeed, the gender job satisfaction difference in Britain had vanished by 2012-14, as female workers became less satisfied as they aged and new cohorts of young female workers entered the labour market with relatively lower job satisfaction.

This paper draws on 22 waves of HILDA data to investigate how the differential in women's reported job satisfaction relative to men's evolved in Australia between 2001 and 2022. The motivation is, firstly, one of simple curiosity as to whether evidence of rising women's expectations in the labour market, as claimed for Britain, also applies to Australia. More broadly, however, understanding differences in job satisfaction by gender and the source of changes in the differential over time provides valuable insights into broader developments in gender equality in the labour market and changing social roles of men and women.

Australian women report higher average satisfaction with their jobs overall in all years up to and including 2019, with the gender gap closing in 2020, 2021 and 2022 (Figure 1). This followed a sharp upturn in average reported job satisfaction for both genders since around 2018, but which has been particularly pronounced for men. Using data from rolling 5-year intervals (2001-2005, 2002-2006, … 2018-2022) multivariate, random-effects panel regression models are fitted to estimate the gender differential in job satisfaction after controlling for a wide range of individual- and job-specific characteristics. This also allows an Oaxaca-Blinder style decomposition of the gender gap in job satisfaction and how it has changed over time into two different sources: 1. differences in the average characteristics of men's and women's jobs, and 2. differences in the way men and women value different job attributes.

The female job satisfaction 'premium' is found to have declined substantially through the 2000s, reaching a low in around 2011, to then recover somewhat and stabilise from around 2013. There is some evidence of rising expectations among female workers contributing to the closing of the gender gap in job satisfaction, in line with Clark's hypothesis. However, other factors have significantly shaped relative job satisfaction of men and women over the first two decades of the 21st century, particularly preferences around working hours.

The results show the paradox of the contented female worker persists, or at least has re-emerged, in the contemporary Australian labour market. Some implications of these findings for gender equality, discrimination and for policy are canvassed in the conclusion.

Background

There is a growing acceptance within economics of the validity of people's subjective assessments as an indicator of utility and one that offers important implications for decision-making by economic agents and for the purposes of policy formulation, as seen in the burgeoning 'happiness' literature (see Frey, 2008). Similarly, while economists traditionally focus on wages as the key indicator of job quality and the driver of workers' choices, subjective assessment of job satisfaction offers broader insight into workers' preferences and the value placed on different job attributes (Clark 1997, Hamermesh 2001). As a pertinent example, a very large body of empirical research focuses on the gender wage gap, when gender inequality may also manifest in a range of non-wage attributes, such as job security and part-time versus full-time status. Differences in job satisfaction by gender may also help to explain important labour market phenomena, such as occupational segregation by gender (Bender, Donohue and Heywood 2005, Buchler and Dockery 2023, Crompton and Harris 1998, Hakim 2000). Indeed, Goldin (2014) has suggested the gender wage gap can be largely attributed to gender differences in preferences for working long hours given productivity benefits associated with long hours in certain occupations.

Clark (1997) observed that women consistently report higher levels of job satisfaction despite the well-established gap in wages in favour of men and widespread evidence that women's jobs are inferior to men's on a range of other attributes such as security, promotion opportunity and sexual harassment. In line with earlier studies, Clark showed that the higher job satisfaction reported by women holds after controlling for a wide range of individual and job-related characteristics using data from Wave 1 of the British Household Panel Survey (BHPS). Explanations put forward for this finding include gender differences in work values, selection bias, and women having lower expectations of their jobs due to the relative disadvantage they faced within the labour market. Given the lower labour force participation rate of women, and particularly of married women, selection bias may create the impression women have higher job satisfaction because women are more likely than men to drop out of the labour force if they are dissatisfied with their jobs.

On the grounds of empirical tests, Clark dismissed the difference in work values and selection bias explanations, preferring the explanation that women have lower expectations, and people assess their job satisfaction, in part, relative to their expectations. Supporting this, he finds the gender satisfaction differential is not present for groups of women likely to have higher expectations: the young, more highly educated, whose mothers were professionals and who work in male dominated workplaces. Hence, Clark proposed the gender differential in job satisfaction may be a transitory phenomenon that would disappear as women's relative position in the labour market improved.

To become known as the 'paradox of the contented female worker' (Bender *et al*. 2005: 482), this observation and explanation has since been the subject of scrutiny across numerous countries. Along with the UK (Clark 1997, Gazioglu and Tansel 2006, Sloane and Willams 2000) evidence of higher job satisfaction for women has been observed in individual county studies for the US (Bender *et al*. 2005), Canada (Dilmaghani 2022), agricultural workers in Senegal (Fabry, Van den Broeck and Maertens, 2022), and Australia. For the latter, the availability of the HILDA dataset has clearly stimulated research in this area. Studies identifying higher job satisfaction for women in multivariate regressions analyses include Aletraris (2010) and Long (2005) based on Wave 1 of HILDA; Kifle (2014) using Waves 1-6; Kifle *et al.* (2014a) using Waves 1-10; and Buchler and Dockery (2023) using data from Waves 1–19. Numerous studies investigate other aspects of gender differences in job satisfaction based on the HILDA data, but without directly estimating a gender 'premium' (Booth and van Ours 2009; Buddelmeyer, McVicar and Wooden 2015; Fleming and Kler 2008, 2014; Kifle 2013; Kifle *et al.* 2014b; Mavramaros; Sloane and Wei 2012; Ong and Shah 2012).

A number of studies have found that the paradox does not apply for certain sub-groups of the workforce, notable younger cohorts and among the more educated, and this is seen to support the hypothesis that the gender differential is due to differences in expectations (Buchler and Dockery 2023; Clark 1997; Dilmaghani 2022; Donohue and Heywood 2004; Kifle *et al.* 2014a; Long 2005). There is also evidence that men's job satisfaction is more sensitive to relative or 'comparison' wages than is women's (Dilmaghani 2022, Donohue and Heywood 2004, Kifle 2013, Perugini and Vladisavljević 2019, Sloane and Williams 2000). Empirical tests undertaken in Hauret and Williams (2017) and Perugini and Vladisavljević (2019) confirm Clark's (1997) finding that the gender satisfaction differential is robust to controls for selection.

As noted, Green *et al.*'s (2018) *Paradox Lost* paper found support for Clark's suggestion that the female job satisfaction premium would disappear over time as women raised their labour market expectations to rival those of males. They find the gender job satisfaction differential had vanished by 2012-14. Cross-country studies also indicate the contented female worker paradox is not evident across all countries, and with mixed evidence on whether the paradox is disappearing over time. In a study of 28 countries based on the 1997 International Social Survey Programs (ISSP) data, Sousa-Poza and

Sousa-Poza (2000) found a gender difference in favour of women in only eight countries, with the largest difference observed for the US and the UK, leading them to suggest the contented female worker paradox is primarily an Anglo-Saxon phenomenon. Using 2015 ISSP data, however, Andrade, Westover and Peterson (2019) find a significant gender difference in job satisfaction for only one of 37 countries (Georgia), and no difference for the pooled sample (which included Australia).

Kaiser (2007) found a significant female job satisfaction premium in 10 of 14 European countries, and suggested the paradox is less likely to be observed in countries with more equal opportunities, such as the Scandinavian countries. Hauret and Williams' (2017) analysis of 2010 European Social Survey data comes to almost the opposite conclusion, finding the gender paradox to apply only in the Nordic group of the 14 countries. Finally, Perugini and Vladisavljević's (2019) analysis of gender job satisfaction differentials based on 2013 data for 32 European countries finds a significant job satisfaction difference in favour of women in the pooled data and in ten individual countries (the highest being for the UK). They also find the differential is lower for women exposed to more equal labour force participation rates by gender in their early life stages, consistent with Clark's (1997) hypothesis of women's lower expectations as the explanation for the contented female worker paradox.

In light of this on-going uncertainty over the existence, persistence and cause of the contented female worker paradox, and in the spirit of Green *et al.*'s *paradox lost* paper, this paper documents the evolution of the gender differential in job satisfaction in Australia over the past two decades. It builds on the existing literature by explicitly decomposing the gender differential in job satisfaction and the changes in that differential over time, into components attributable to observable differences in the characteristics of the jobs held by women and men, and to gender differences in preferences related to those characteristics.

Gender differences in job satisfaction in Australia

Descriptive overview

The data are from the first 22 waves of HILDA (2001–2022). HILDA is a panel survey of individuals from a representative sample of private households (Watson and Wooden, 2010). Within selected households all occupants aged 15 and over are surveyed annually. Around 13,000 individuals from over 7,000 households have responded in each year, with year-on-year attrition rates averaging below 10 per cent. In 2011 an additional top-up sample of 2,153 households encompassing 4,009 responding individuals was recruited to the survey sample (HILDA Survey Annual Report, 2012).[1] By definition, all observations included in the analysis are for persons aged 15 and over and who were employed at the time of the relevant HILDA survey. For all analyses the sample is also restricted to exclude multiple job holders and those who work as unpaid family helpers.

In each wave, the survey asks people who are employed how satisfied they are with a number of aspects of their job, before asking 'All things considered, how satisfied are you with your job?'. Responses are recorded on an 11-point scale ranging from 0 (totally dissatisfied) to 10 (totally satisfied). On the basis of the raw means of responses to that question, women have tended to report higher average satisfaction with their jobs. Across the full sample pooled over the 22 years the mean of people's assessment was 7.68 with a standard deviation of 1.63, indicating workers generally report quite high satisfaction levels.[2] The modal response is 8, with just over 80 per cent of workers reporting 7 or higher. For women, the mean was 7.74 compared to 7.63 for men. Looking at individual years, mean job satisfaction for women has been higher than for men in all years from 2001 up to and including 2019, with the gender gap closing in 2020 and 2021 (Figure 1). This followed an upturn in average reported job satisfaction for both genders commencing in 2018, but which was particularly sharp for men.

Figure 1. Mean overall job satisfaction rating by gender and year, HILDA waves 1-22

Notes: Job satisfaction assessed on an 11-point scale ranging from 0='totally dissatisfied' to 10='totally satisfied'. Number of observations per year range from 3,244 to 4,925 for women; 3,797 to 5,411 for men.

1 see http://melbourneinstitute.unimelb.edu.au/hilda for further details on the HILDA survey.

2 All reported means are calculated using HILDA's responding-person weights.

While the focus in this analysis is on the measure of overall job satisfaction, it is interesting to also look at gender differences in other job aspects. Leading up to the 'all things considered' assessment, respondents are asked to assess their satisfaction with their total pay, job security, the work itself (or 'what you do'), hours worked, and the flexibility to balance work and non-work commitments. Figure 2 shows trends in the mean assessments of satisfaction with these individual job aspects by gender. It appears that men and women have tended to provide very similar assessments of their satisfaction with total pay and the work itself. The former is somewhat surprising given the well-established gender wage gap in favour of men in Australia, of which only a portion can be explained by observable differences in productivity-related job and worker characteristics (see, for example, Duncan, Mavisakalyan, and Salazar 2022). For almost 20 years from 2001, women have consistently reported higher average satisfaction with job security, hours worked, and flexibility to balance work and non-work commitments. However, each of these appear to have contributed to the convergence in job satisfaction by gender. Mean satisfaction with flexibility to balance work and non-work commitments has been higher for men than for women since 2019, while mean satisfaction with job security has converged to be almost identical for Australian female and male workers in those years. Mean satisfaction with hours worked for men and women also converged between 2019 to 2021, but widened again in favour of women in the most recent (2022) wave of the HILDA survey.

Figure 2. Mean satisfaction with individual job aspects by gender and year, HILDA Waves 1-22

Multivariate estimates

Comparison of the raw means does not take into account differences in characteristics of jobs undertaken by women and men, or in the personal characteristics of workers, such as educational attainment. A standard approach to identifying a gender effect after controlling for other factors that may affect workers' satisfaction with their jobs, is to estimate a multivariate regression of the following form for panel data with random-effects:

$$JS_{it} = \alpha + \theta FEMALE_{it} + \beta X_{it} + u_i + \varepsilon_{it} \tag{1}$$

where: JS_{it} denotes the job satisfaction assessment of worker i in period t; ß is a vector of coefficients associated with the vector X of control variables, and with error terms including an individual component u_i and residual ε_{it} distributed with mean of zero across individuals and over time. The coefficient θ on a dummy variable indicating whether the individual is female represents the estimated gender effect on job satisfaction independent of the other control variables.[3]

To provide point-in-time estimates while also exploiting the panel dimension of the data to control for unobserved individual effects, models are fitted to data from rolling 5-year intervals (2001-2005, 2002-2006, … 2018-2022) to estimate the gender differential in job satisfaction. For convenience, these intervals are referenced by their midpoints (ie, 2003, 2004 …, 2020).

A wide range of individual-, job- and workplace-specific control variables are included in the models, as can be seen in the Appendix Table A1, which reports results from models estimated with the full 22-year panel. Guiding factors in the selection of independent variables to include are the need to control for factors previously found to have significant effects on job satisfaction, subject to the constraint that variables must be available on a consistent basis across all 22 waves.[4] For the key research question at hand, it is important to include variables likely to capture job expectations and, in particular, gender differences in job expectations. Among the individual-specific characteristics, marital status and the presence of dependent children are important

3 The time subscript has been included for the gender dummy to acknowledge that individuals may transition between genders. However, aside from a handful of exceptions, the variable is fixed for individuals in the HILDA panel, thereby precluding estimation by fixed-effects.

4 In Waves 1-21, employed persons were asked in the HILDA survey how many people worked at their workplace and whether their employer operates from more than one location in Australia. These questions were not included in Wave 22. To ensure consistent models are estimated across all periods, controls for these aspects of the workplace are not included in the modelling for this paper. Results relating to gender are insensitive to their inclusion or exclusion in models estimated on data from Waves 1-21.

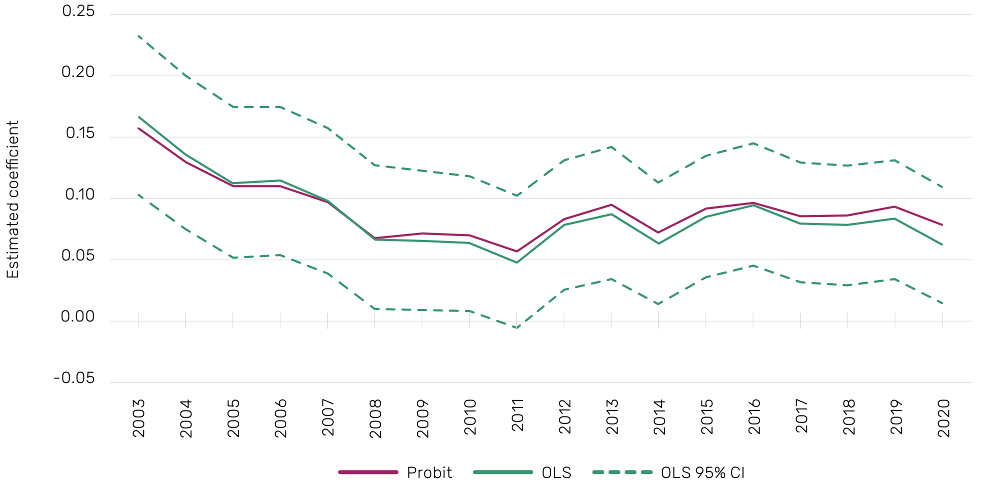
as Australian social norms around women taking on the 'secondary breadwinner' role are likely to shape expectations around work (Baxter and Hewitt 2013; Buchler and Dockery 2023). Educational attainment and neighbourhood socio-economic status are likely to affect who people compare themselves with, and therefore expectations of job quality. Region of residence has also been included as an individual effect, as Buchler and Dockery (2023) find job satisfaction is significantly higher for Australians living outside the major capital cities.

The estimated coefficients for the female dummy variable from models for the 5-yearly intervals are plotted in Figure 3. The ordered categorical nature of the dependent variable lends itself to estimation by the ordered probit model. However, the results are very similar whether the models are estimated by linear regression (ordinary least squares [OLS]) or the more technically correct ordered probit specification, as can be seen by the consistency in the significance, direction and general magnitude of the estimated coefficients reported in Appendix Table A1, and how closely the estimated gender effects from both specifications track one another in Figure 3. The subsequent analysis is therefore based on results from linear panel models, making interpretation of coefficient estimates more straightforward.

Results from OLS models with random effects return an estimated coefficient on the female dummy for the 2001-2005 sample of +0.17 ($p<0.01$), compared to +0.06 ($p<0.01$) for the final 2018-22 sample. It reached a low of +0.05 ($p=0.07$) for the 2009-2013 sample. Figure 3 also plots the 95 per cent confidence intervals for the OLS estimates, highlighting that at its low point, one could not dismiss the hypothesis of no gender differential at the 5 per cent level of significance. For all other years, however, the lower limit of the interval is above zero, and the estimate is highly significant in the early and later periods.[5] Hence, unlike Green *et al.*'s (2018) findings for Britain, the paradox of the contented female worker appears to have persisted, or at least re-emerged, in Australia.

5 The estimates from the ordered probit models also reach a low in 2011, but with a higher level of significance ($ß=0.06$, $p=0.016$). This is the only interval for which the gender coefficient is not highly significant (ie. $p<0.01$).

Figure 3. Estimated coefficient on female dummy variable, rolling 5-year OLS and ordered probit panel regression models



Notes: Plotted against the midpoint of the 5-year intervals (eg. 2003 is the estimate for the 2001-2005 sample); models estimated using STATA's XTREG and XTOPROBIT commands.

As Figure 3 illustrates, Australia appeared to be following the pattern of the UK, with the paradox of the contented female worker gradually diminishing from the turn of the century to 2011, set to vanish in a supposed tide of rising female labour market expectations. However, where Green *et al*. (2018) found no significant gender difference in Britain by 2012-14, the job satisfaction premium in favour of women in Australia appears to have rebounded slightly after 2011 and since remained relatively constant.

Differences in expectations? Evidence from Oaxaca-Blinder decompositions

As is well known following the work of Oaxaca (1973) and Blinder (1973), and most commonly applied in the case of the gender wage differential (see, for example, Blau and Kahn 2017), consider a regression model of job satisfaction estimated separately for males (M) and females (F):

$$JS_{Fi} = \beta_F X_{Fi} + u_{Fi} \tag{2a}$$
$$JS_{Mi} = \beta_M X_{Mi} + u_{Mi} \tag{2b}$$

The difference in mean job satisfaction between women and men ($\overline{JS}_F - \overline{JS}_M$) can be decomposed into three components:

- Differences in the mean characteristics between women and men (differences in the X's: $\bar{X}_F - \bar{X}_M$).
- Differences between women and men in the estimated coefficients, or in the effects of characteristics on job satisfaction (differences in the ß's: $\hat{\beta}_F - \hat{\beta}_M$).
- Residual interaction effects capturing simultaneous differences in characteristics and coefficients.

Differences in the effects of variables on job satisfaction (i.e. differences in the coefficient estimates) can be seen as differences in women's and men's preferences for particular job attributes, or differences in the way individual characteristics, such as level of education or marital status, shape job satisfaction for men and women. In the literature on the gender wage gap, the earnings difference that can be attributed to differences in mean characteristics (differences in the X's) is often called the 'explained' component of the gender wage gap. The portion associated with differences in the coefficients is termed the 'unexplained' component, and often considered to proxy gender discrimination in wage setting. In the context of job satisfaction, the unexplained portion could be interpreted as differences in job preferences and expectations between male and female workers, rather than discrimination. However, experiences of discrimination may lower workers' expectations, so the two cannot be seen as unrelated.

The Oaxaca-Blinder decomposition was applied to each of the 5-year rolling samples of workers in the HILDA data. This allows the gender differences in mean job satisfaction in each 5-year interval to be decomposed into explained and unexplained portions. Specifically, the user-written OAXACA command for STATA is used (Jann 2008). Initially the 'pooled' option is used for the decomposition in which the benchmark $\hat{\text{ß}}$'s are estimated from a pooled regression over all persons, rather than choosing coefficients from either a regression for males or for females. This means there is no interaction component to the decomposition.

As shown in Figure 4, in the early 2000s the higher job satisfaction reported by women was due almost entirely to the differences in coefficients (the 'unexplained portion'), with differences in observed characteristics of individuals and their jobs contributing virtually nothing. That is, women's higher reported job satisfaction was due to different *effects* of characteristics rather than different characteristics of jobs and workers. The proportion of the gender gap in job satisfaction attributable to differences in preferences declined to reach a minimum in 2011, and in fact was statistically insignificant in models for samples centred from 2008 to 2011, before then increasing again.

In contrast, the proportion of the satisfaction differential attributable to differences in mean characteristics increased from close to zero to peak in 2008, before then declining such that, for the 2018-2022 sample, differences in mean characteristics worked to reduce the job satisfaction premium observed for women relative to men. That is to say, in the 2018, 2019 and 2020 samples (2016-2020 to 2018-22 intervals), women's job satisfaction would have been even higher if they were to have otherwise identical characteristics of men and their jobs. Possibly, the quality of women's jobs relative to men's jobs improved between 2001 and 2008, but that difference in job quality has since steadily eroded to now favour men.

Figure 4. Oaxaca-Blinder decomposition – explained and unexplained components of the female job satisfaction premium, HILDA Waves 1-22

It is possible to look in more detail at the specific characteristics and associated effects that drove these changes in relative job satisfaction. To do so the estimated gender job satisfaction differential can be decomposed down to the contribution of each individual variable in terms of the differences in means for men and women for that variable (contribution to the explained portion), and the difference in that variable's coefficients in models estimated separately for men and women (contribution to the unexplained portion). As gender differences in coefficients make the largest and most volatile contribution to the job satisfaction gap, we first consider how the contribution of coefficients for individual variables changed over time, and then turn the focus to the detailed effects of changing individual and job related characteristics (differences in means). Full results of OLS models estimated separately for male and female workers and associated variable means by gender are presented in appendix Tables A2-A5 for the pooled samples for 2003, 2008, 2011, and 2020, respectively.

'Unexplained' difference in job satisfaction – differences in the effects of variables

As noted above, the female job satisfaction premium was estimated to be largest for the initial 2001-2005 sample, with a difference of 0.174 points in favour of female workers on the 11-point scale used to assess job satisfaction. This may seem a small effect, but in fact represents a sizeable shift. As noted, the distribution of responses is highly clustered with almost one in three workers selecting a value of '8' near the highly satisfied end of the scale. The Oaxaca-Blinder decomposition indicates the female premium observed for the 2003 sample can be attributed almost entirely to gender differences in the coefficients: 0.172 of the 0.174 point difference. This declines to a minimum in 2011 and then rebounds over the timeframe of the study.

We can broadly group the explanatory variables into two categories: characteristics relating to the individual (e.g. age, level of education, family status etc.) and those relating to their job and working arrangements. Following this split, it is differences in coefficients on individual characteristics that drove the female job satisfaction premium in the early 2000s, with differences in those coefficients raising women's job satisfaction by an estimated 0.63 points on the scale. Differences in coefficients on job-related characteristics actually worked to increase men's job satisfaction relative to women's (by 0.10 pts). The difference in the intercept terms in the models estimated separately for men and women was also associated with higher satisfaction for men (0.36), but it is important to note this will also capture effects of omitted variables, not just differences in reporting tendencies between genders. The net gender differential attributable to differences in coefficients is comprised of these three components as shown in Table 1.

Table 1. Female job satisfaction 'premium': contribution of differences in coefficients (the ßs) [a positive effect is associated with higher job satisfaction for women compared to men]

	Estimated effects of coefficients on:			
Interval	Individual characteristics	Job-related characteristics	Intercept terms	Total 'unexplained' component
2003 (2001-05)	+0.634	-0.100	-0.361	+0.172
2011 (2009-13)	+1.093	-0.306	-0.744	+0.042
2020 (2018-22)	+0.709	+0.257	-0.892	+0.075

The major contributing differences in coefficient estimates from the 2003 sample that were associated with higher job satisfaction for women included:

- Age – age was included in the regression model along with its quadratic to capture second order effects. The resulting coefficients show both men and women's job satisfaction declining with age before recovering. The literal interpretation of the coefficients is consistent with men's job satisfaction reaching a minimum at 35.9 years of age, around two years later than for women (33.8). However, the first order (negative) effect for males is over twice the magnitude, indicating job satisfaction declines more rapidly for men after they enter the workforce than for women. This contributes a 1.90 gap in reported job satisfaction. The difference in the quadratic term offsets this (-1.04), but still leaves differences in the effects of age making a large net contribution (0.85), a multiple of around five times the total gender differential.
- Hours worked - usual weekly hours worked were included in the models as a series of mutually exclusive dummies: 0-15 hours, 16-30 hours, 31-38 hours, 45-54 hours and 55 hours or more. The omitted reference category is people who usually worked 39-44 hours per week. Effects associated with working up to 38 hours a week favoured women, with those working arrangements associated with higher job satisfaction than is the case for men. The effects of working long hours (45-54 hours, or 55 hours or more per week) partially offset this, as women appeared to have a lower preference (or greater distaste) for such working weeks, but the differences are very minor. The overall effect of coefficients on hours worked is to increase women's job satisfaction relative to men's.
- A smaller negative impact of casual status on job satisfaction for women.
- A slower decline in job satisfaction with time spent in the current occupation.

Differences in coefficients offsetting these effects in the 2003 model are:

- Wages – men's job satisfaction increased more rapidly with their hourly wage rate (modelled in logs)
- Union membership[6] – in the 2003 sample, being a member of a union membership was associated with significantly lower reported job satisfaction for women. The association for men was also negative, but much smaller and insignificant.
- Level of qualification – job satisfaction decreases with workers' level of education, so more highly qualified workers are less satisfied with their

6 Union membership status has been classified here as a job characteristic, although arguably it could be considered as a characteristic of the individual.

jobs, other things being equal. However, the gradient follows a much steeper decline for women.

- Neighbourhood socio-economic status – *ceteris paribus*, job satisfaction also declines with the SEIFA[7] decile of the neighbourhood in which the worker lives, presumably reflecting comparison effects. This decline was sharper for women in the 2003 sample, and in fact is only significant in the model for female workers ($ß=-0.02$; $p<0.05$).

Changes in the unexplained component: 2003 to 2011

The 'unexplained' component of the female job satisfaction premium fell from 0.172 in 2003 to a low of just 0.042 in 2011. The largest contribution to this change in the coefficient estimates related to the constant (or intercept) terms for men and women, and hence the shift is largely unexplained by effects of the included variables (see Table 1). Potentially, it may capture growing expectations of women relative to men with respect to the general quality of their jobs, and hence lower satisfaction reported by women in 2011, relative to 2003, for given job characteristics. However, note again that those differences will include any effects of omitted variables. The change in the intercept terms equates to a 0.383 point reduction in job satisfaction for women compared to men from 2003 to 2011. This more than accounts for all of the fall in the unexplained component between 2003 and 2011 (down by 0.129), so we cannot ascribe that development to any particular change in the effects of individual characteristics or relative preferences relating to job characteristics.

Aside from the effects of the intercept terms, coefficients on individual characteristics served to increase women's job satisfaction relative to men's (+0.460), while changes in the effects of job-related characteristics reduced the gap (-0.206). Changes in estimated coefficients for age and qualifications between 2003 and 2011 in fact accentuated the female satisfaction premium, as did the effect of being married with children aged 5-14 years, which was associated with higher job satisfaction among women in the later period.

Some of that effect was offset by changes in the effect of time spent in current occupation, primarily a negative and significant effect observed for men in 2003 became insignificant in 2011. Changes in the estimated preferences for working hours also played a small role (-0.055), primarily through the estimated effect of working 16-30 hours per week (as opposed to 39-44 hours) converging for men and women. In 2003 women were significantly more satisfied working 16-30 hours per week than with full-time work, but this was not the case in 2011.

7 Socio-economic Indices for Areas (SEIFA) are measures generated from census data by the Australian Bureau of Statistics (see Adhikari 2006). SEIFA deciles of respondents' postcode are available as derived variables in the HILDA data. The decile for the 2001 SEIFA index of relative socio-economic advantage and disadvantage is used in the modelling in this paper.

The negative effect of living in a neighbourhood of higher socio-economic status that was observed in 2003 had vanished in the estimates for 2011, such that the estimated effect was small and insignificant for both men and women. So while this effect acted to reduce women's job satisfaction relative to men's in 2003, no such effect was present in 2011.

Changes in the unexplained component: 2011 to 2020

So, if there was a general fall in women's reported satisfaction with otherwise 'like' jobs between 2003 and 2011, what would account for the subsequent rebound in women's relative satisfaction? The unexplained portion of that total gender gap rebounded from 0.042 to 0.092 in 2016 and stood at 0.075 at the end of the analysis time frame (2020, estimated for the 2018-22 sample).

Changes in the intercept terms of the models for men and women were associated with a further deterioration in relative job satisfaction for women between 2011 and 2020. Differences in the effects of individual characteristics continued to make a positive contribution to the female job satisfaction premium in 2020, but to a lesser degree than in 2011, such that changes in the effects of individual characteristics between 2011 and 2020 reduced women's job satisfaction relative to men's. However, changes in the effects of job characteristics offset both those effects, such that in total the 'unexplained' female premium increased back to 0.075 (see Table 1). Unlike in 2003 and 2011, estimated coefficients on variables capturing aspects of jobs and workplaces contributed positively to women's relative jobs satisfaction in 2020. Recall, this relates to differing preferences for job characteristics, not differences in the actual characteristics of those jobs.

In terms of the effects of individuals' characteristics, a more rapid decline in job satisfaction with age for women in the 2020 sample was the main source of the fall in the female satisfaction premium associated with individual characteristics. Working against this, the effects of qualifications further moved to increase women's job satisfaction. In the early 2000s, having a degree or post-graduate qualification was associated with lower job satisfaction for both genders, but the effect was stronger for women. Those coefficients for men remained quite stable across the period, but by the end of the analysis timeframe the coefficients for women had converged toward the male coefficients. Highly qualified women no longer faced a steeper decline in job satisfaction. Optimistically, this may reflect a reduced 'glass ceiling' effect.

The differential effect of living in a neighbourhood of higher socio-economic status again contributed to the female premium. Between 2011 and 2020, the change was in the estimated coefficient for men, such that living in a higher status neighbourhood had a significant and negative effect on men, with the estimate for females insignificant. In summary, neighbourhood comparison effects shifted from women to men over the period of analysis. What appear to be the negative effects on satisfaction from neighbourhood 'rivalry' were observed only for women in 2003; for neither men nor women in 2011, and only for men in 2020.

In terms of the effects of job characteristics, the estimated elasticity of job satisfaction associated with hourly wages was higher for women than for men in 2020, whereas in 2003 and 2011 the estimate was larger for men. The results for the earlier periods are consistent with previous literature suggesting women place higher relative value on intrinsic job rewards, such as the nature of the work, while men place higher relative value on extrinsic rewards, such as pay and promotion (see Andrade, Westover and Peterson 2019; Dilmaghani 2022). This shift in responsiveness to wages contributed substantially to the growing female premium between 2011 and 2020. Changing effects of time spent in the current occupation and of having supervisory responsibilities also contributed. In the latter case, in the initial (2003) period having supervisory responsibilities had a significant and positive effect on job satisfaction for men, but was smaller and insignificant for women. In 2011 the effects were almost identical for both genders, and by 2020 were positive and significant for women only. Again, optimistically, this shift in the effect of managerial responsibilities on job satisfaction from one which favoured men to one which favoured women may also reflect a lessening of 'glass ceiling' effects in the labour market.

For the latter interval, men also appear to have developed a stronger preference for working standard full-time hours. In 2020, job satisfaction was significantly lower for men who worked part-time compared to those who worked roughly 40 hours per week, an effect that was not evident in either 2011 or 2003. By 2020 men were also more dissatisfied when working long hours, with significant negative effects observed for men working 45 to 54 hours per week and particularly those working 55 or more hours per week. This effect was also more pronounced than in the earlier years, though women still display a stronger distaste for longer working hours.

Explained difference in job satisfaction – differences in means

As outlined above, the gender gap in job satisfaction that could be 'explained' in an Oaxaca-Blinder decomposition by gender differences in the mean characteristics of individuals and their jobs was close to zero in 2003 (+0.003), increased to a peak in 2008 (+0.060), and then fell to -0.026 in the final period (2018-2022) such that differences in observable characteristics favoured men. The summary provided in Table 2 shows that the increase in the explained component of the female job satisfaction premium to 2008 and subsequent decline can primarily be attributed to changes in the mean characteristics of men's and women's jobs. Differences in mean individual characteristics shifted to reduce the female premium over the full period.

Table 2. Female job satisfaction 'premium': contribution of differences in means (the Xs) [a positive effect is associated with higher job satisfaction for women compared to men]

Interval	Contribution of differences in means of: Individual characteristics	Job-related characteristics	Total 'unexplained' component
2003 (2001-05)	-0.014	+0.017	-0.003
2008 (2006-10)	-0.018	+0.078	+0.060
2020 (2018-22)	-0.039	+0.013	-0.026

While the net effect of the differences in means across all variables was close to zero in 2003, some individual variables did make substantial but offsetting contributions. The higher proportion of women with a degree qualification (25 per cent versus 18 per cent for men) contributed to lower average job satisfaction for women, but this was offset by the much lower proportion of women holding a certificate level III/IV qualification. The higher proportion of women working 0-15 hours per week and in the public sector contributed to greater women's job satisfaction, since in the pooled sample both those characteristics are associated with higher job satisfaction. The higher proportion of women working as casuals detracted from women's satisfaction.

Changes in the explained component: 2003 to 2008

From 2003, the contribution of differences in coefficients declined while the contribution of differences in means to the female satisfaction premium increased, suggesting either the characteristics of female workers were changing or the quality of women's jobs were improving. Comparing the change in the 'explained' contributions from 2003 to 2008, when the effects of differences in means peaked, there were two main developments. Both appear to arise from changes in preferences between the two periods, rather than changes in characteristics of workers or jobs:

- Hours worked – between 2003 and 2008 there was very little change in the gender distribution across hours worked, with women clearly over-represented in part-time work and under-represented among long hours workers in both periods (see Table 3). However, in 2008, part-time work attracted a marginally higher job satisfaction premium than was the case in 2003, and long hours of work a marginally greater penalty. Hence, in the 2008 model, the differences in working times between men and women, notably the higher proportion of women working 1–15 hours, accounted for substantially more of the overall female job satisfaction premium. Essentially a shift in preferences towards a shorter working week and away from long hours of work meant that the pre-existing hours distribution was now more favourable to female workers.

- The smaller proportion of women working in the occupation of machinery operators and drivers – for both the 2003 and 2008 samples, 11 per cent of men and just one per cent of women worked in jobs in this major occupational grouping. However, for the latter period, working in this occupation was significantly associated with lower job satisfaction, and hence women's under-representation in the occupation contributed to higher job satisfaction relative to men.

Changes in the explained component: 2008 to 2020

As shown in Figure 3, from 2008 the component of the female job satisfaction premium that could be explained by gender differences in means steadily declined. By 2020 the differences in mean characteristics are estimated to favour men. Again, that change can be traced primarily to gender differences in means of job characteristics, with only a minor contribution from changes in differences in person characteristics by gender (see Table 2). Based on the coefficients for the pooled sample for 2008, differences in the means of the job characteristics of men and women contributed +0.78 points to women's job satisfaction relative to men. This fell to a contribution of +0.013 for women in 2020, a net change of -0.065 corresponding to a lower job satisfaction premium for women in 2020 compared to 2008. Over the same period, there was a net change of -0.021 associated with differences in the means of individual characteristics, further adding to the shift in job satisfaction in favour of men.

Again, the major shift related to gender differences in hours of work with preferences playing a key role, and this continued the shift observed from 2003 to 2008. Compared to the sample for the 2008 estimates, women in the 2020 sample were less likely to be working 0-15 hours per week: 18 per cent in 2008 compared to 13 per cent in 2020 (see Table 3). The proportion of women working 31-38 hours increased by 5 percentage points (25 per cent to 30 per cent). More importantly, while part-time workers reported substantially higher job satisfaction in 2008, this no longer applied in 2020, as reported in Table 3. This was most apparent for short hours workers: the estimated coefficient on working between 0 and 15 hours per week in the pooled 2008 sample was +0.337 ($p<0.01$), but was small and insignificant for the pooled 2020 sample ($ß=+0.043$, $p=0.31$). Hence, the major factor driving the rise and then fall of the 'explained' component of the female job satisfaction premium was an initial growing preference for part-time work, which favoured women's satisfaction but vanished between 2008 and 2020, aided by a decline in the proportion of women working part-time. Growing dissatisfaction with long working weeks partially offset this effect. By 2020, the difference in distribution of males and females across all hours of work categories only marginally favoured women in terms of the association between those working hours and job satisfaction (+0.023 points), whereas in the earlier periods it had substantially favoured women (+0.052 in 2003 points and +0.082 points in 2008).

Table 3. Distribution of hours worked by gender, and associated estimated coefficients

Usual hours worked/week	Interval	Mean		Coefficient (signif.)
		Women	Men	
0 to 15 hours	2003 (2001-05)	0.19	0.07	0.32*** (0.00)
	2008 (2006-10)	0.18	0.07	0.34*** (0.00)
	2020 (2018-22)	0.13	0.07	0.04 (0.31)
16 to 30 hours	2003 (2001-05)	0.26	0.07	0.15*** (0.00)
	2008 (2006-10)	0.25	0.07	0.18*** (0.00)
	2020 (2018-22)	0.27	0.09	-0.05* (0.08)
31 to 38 hours	2003 (2001-05)	0.23	0.18	0.05 (0.20)
	2008 (2006-10)	0.25	0.19	0.00 (0.89)
	2020 (2018-22)	0.30	0.24	0.02 (0.48)
39 to 44 hours	2003 (2001-05)	0.18	0.26	n.a.[a]
	2008 (2006-10)	0.18	0.27	
	2020 (2018-22)	0.17	0.27	
45 to 54 hours	2003 (2001-05)	0.11	0.26	0.03 (0.35)
	2008 (2006-10)	0.11	0.26	-0.03 (0.34)
	2020 (2018-22)	0.10	0.23	-0.13*** (0.00)
55 hours or more	2003 (2001-05)	0.04	0.15	-0.12** (0.02)
	2008 (2006-10)	0.04	0.14	-0.15*** (0.00)
	2020 (2018-22)	0.03	0.10	-0.29***(0.00)

Notes:
*** p<0.01, ** p<0.05, * p<0.10;
a. No coefficient estimate as 39-44 hours per week is the omitted or comparison category in the regressions.

In terms of individual characteristics, the contribution of differences in means was to decrease women's job satisfaction relative to men's, and this became more pronounced over time. This can largely be attributed to nuanced changes relating to educational attainment. The satisfaction 'penalty' associated with higher level qualifications fell over time, and this change was driven by a changing association between educational attainment and job satisfaction among women. However, the proportion of female workers with post-school qualifications at all levels increased, including certificate III/IV, diplomas, bachelor degrees and post-graduate qualifications, and each of these is associated with lower job satisfaction relative to workers who had not completed Year 12.

A sudden convergence?

As shown in Figure 1, the raw means of job satisfaction assessments of men and women converged to be virtually identical in 2020 and 2021, before very marginally drifting in favour of women again in 2022. Trends in satisfaction with individual aspects of workers' jobs (Figure 2) suggest growing satisfaction of male workers with their hours of work, job security, and the flexibility to balance work and non-work commitments all contributed to this convergence. With mean assessments of men's and women's job satisfaction finally equalising in 2020 and 2021, an appealing first explanation is that COVID-19 had an effect on men's expectations around work, and possibly led to changes in working arrangements that enhanced satisfaction. However, there seems to have been a quite general up-tick in male workers' satisfaction across all domains starting from around 2017, well before the impacts of COVID-19 hit the Australian labour market in 2020.

Having selected rolling 5-year intervals as the basis for analysis, the change within this interval will be only partially captured in the reported estimates. Cross-sectional regressions based only on the 2022 HILDA sample do still suggest a positive and moderately significant female worker premium in ordered probit models (ß=0.06; p=0.02), but the estimate is not significant in the linear model (ß=0.06; p=0.13).

To investigate this more recent development in detail, the Oaxaca-Blinder decomposition was conducted using the sample of male workers from 2017 and comparing them to the sample of male workers in 2022. On the 11-point scale, mean job satisfaction for male workers increased from 7.58 in 2017 to 7.86 in 2021 (+0.28), by far the largest jump over a four-year period in the 22 years of HILDA data. Again, changes in mean characteristics play a trivially small role (+0.03), and most of the rise in job satisfaction is associated with the differences in the estimated effects of person or job attributes.

Among the differences in mean characteristics contributing to the small 'explained' component of the change between 2017 and 2022 are an increase in the proportion who do some of their usual hours of work from home (from 20 per cent in 2017 to 30 per cent in 2022), a small increase in real hourly wages, and a drop in the proportion of males working 55 hours or more per week (from 12 per cent to 9 per cent). The higher incidence of working from home, which is associated with higher job satisfaction, is likely to be a legacy of the pandemic. However, it did not substantially contribute to gender differentials analysed above, indicating this change in working arrangements had a similar effect on job satisfaction for men and for women.

The differences in estimated coefficients are the source of most of the increase in job satisfaction for men between 2017 and 2022 (+0.51 excluding the intercept term). There is a decline in the elasticity of job satisfaction with respect to real hourly wages that reduces job satisfaction in 2022 (-0.23), and that more than offsets the effect of the actual increase in wages. The largest effect comes from a lower decline in job satisfaction with age for the 2022 sample of men (+1.28 net effect when combined with the change in the estimate for the quadratic terms). The intercept term (+0.65) for 2021 compared to 2017) is also a major contributor, reflecting either changes in men's assessments for like jobs (expectations), or changes in the means and effects of omitted variables.

On the evidence here, the convergence in the raw means in job satisfaction ratings of male and female workers in 2020 and 2021 may constitute another fluctuation in the magnitude of the female job satisfaction premium, but probably not its demise. Future waves of data will be needed to see if the increase in male job satisfaction observed since around 2017 continues, and continues beyond what can be accounted for by observable characteristics and changing preferences. Indeed, data from the very latest wave of HILDA (2022) show male job satisfaction levelling off to be marginally eclipsed by female's assessments of their jobs.

Conclusion

The paradox of the contented female worker was alive and well in the 2018-2022 period, the most recent of the 5-year intervals analysed in this paper. This is in the sense that, after controlling for a wide range of individual, job, and workplace characteristics, the estimated effect of a worker being female was to increase job satisfaction, and that estimate is highly significant using both linear and ordered probit models. For that interval, higher female job satisfaction can be primarily attributed to gender differences in the effects of workers' own characteristics on job satisfaction, notably the differential effect of age for men and women.

In line with reported developments in the UK and other countries, and with Clark's (1997) prediction, the female satisfaction premium in Australia appeared to be steadily dissipating over the first decade of this century. Moreover, given the variables included, the modelling approach used here could not attribute that fall in the female worker premium to either changes in observable characteristics or their estimated effects, with much of the change associated with differences in the gender intercept terms. This is consistent with the hypothesis that the narrowing gender gap was a result of women's rising expectations around their position in the labour market, but it may also reflect changing effects of variables that have not been included in the modelling.

However, the paradox did not continue to fade away in the second decade of the 2000s. Rather, it re-emerged and remained relatively constant from around 2013, albeit with the gender gap still smaller than at the turn of the century. Results of Oaxaca-Blinder decompositions are potentially consistent with female expectations continuing to rise, however, other factors have been at work to maintain a positive female job satisfaction premium. In terms of labour market developments, the most important of these appear to be changes relating to working hours and educational attainment. In both cases, it is changes in the effects of these variables on job satisfaction that most impact upon the overall gender difference in job satisfaction, rather than changes in the distributions of working hours or educational attainment. Presumably, in the case of hours of work, this reflects the changing preferences of workers and, in the case of educational attainment, changing expectations of workers.

The actual distribution of working hours by gender did not drastically change over the two decades of the HILDA survey. Both women and men saw an increase of around 6 percentage points in the proportion working 31-38 hours per week; but for women this was associated with a decline in the proportion working 0-15 hours per week, and for men a decline in the proportion working 55 hours per week or more. However, initial preferences for part-time work enhanced women's job satisfaction in 2003 relative to men's, and more so in 2008. Preferences then changed in favour of typical full-time working weeks, including an apparent growing distaste for long working weeks. With the proportion of women working 0-15 hours per week declining significantly between 2008 and 2020, these changing preferences over the time period contributed to the proportion of the female job satisfaction premium that could be explained by gender differences in working hours increasing up until 2008, but then falling again.

There were offsetting effects with regard to educational attainment. In all periods, and for both male and female workers, higher levels of educational attainment are associated with lower job satisfaction. This is to be expected given we are controlling for job characteristics: due to expectations, a worker with higher educational attainment is likely to be less satisfied with a given job than a less educated worker. However, this effect was much larger for women in the earlier part of the timeframe with, in particular, a relatively larger negative effect of having university level qualifications. By the final interval the effects (coefficients) of different levels of education were similar for men and women. However, average educational attainment of women continued to rise, adding to the reduction in the 'explained' component of the female satisfaction premium.

In addition to potential changes in the expectations of female workers, the results suggest a possible amelioration of 'glass ceiling' effects for women. The falling penalty associated with higher educational qualifications for women suggests a closer fit of actual job quality to expectations for men and women with higher qualifications in more recent times. A job satisfaction premium associated with having supervisory duties was also observed only for men in the initial period, but had emerged for women in the 2018-22 interval. Both may reflect women becoming more comfortable – or less likely to feel discomfort – when taking on more senior or authoritative roles.

Much of the focus on gender equity in the Australian labour has been on the gender wage gap (Duncan *et al*. 2022) and rates of promotion to more senior or executive positions (Cassells and Duncan 2021). The results relating to job satisfaction provide added insights into the nature of labour market inequality. If indeed the disappearance of the 'paradox of the contented female worker' in Britain was because women improved their relative position in the labour market, and thus revised their expectations to be equal to those of men, the same has not occurred in Australia. Does this mean Australian women's relative positions in the labour market have not improved as it may have in other countries?

Lower expectations of job quality for women could arise due to systemic discrimination in the workplace, or from differences in social norms around gender division in work and home production. There is evidence that lower expectations due to discrimination within the labour market ameliorated over the first two decades of this century. Between the 2001-05 to the 2018-22 samples, the gender gaps in mean wages, the proportion with supervisory responsibilities, and the proportion working part-time reduced marginally. As noted, there is some evidence of reduced 'glass ceiling' effects. Recall also, the female satisfaction premium at the end of the period can be attributable purely to differences in coefficients rather than mean characteristics of individuals and their jobs. These findings suggest policy to promote gender equity by targeting discrimination within the labour market must be accompanied by policies that address social expectations of women as secondary breadwinners, particularly after having children (see Buchler and Dockery 2023, Kifle *et al*. 2014b). Such policies may include provisions to provide greater flexibility to balance caring roles with paid work, and to encourage men to take up such provisions in greater proportion.

Limitations and avenues for further research

Changing estimates of the effect of age on job satisfaction for men and women had substantial effects on the estimates. Whether this represents genuine gender differences or genuine changing social trends, as opposed to sensitivity of the estimates to model specification, warrants further research. Estimates may be less sensitive if a series of dummy variables capturing age categories were used instead of the inclusion of years of age as a linear variable and its quadratic. More detailed analyses of changes in job satisfaction for university educated women also seems a potentially fruitful avenue for further research. There have clearly been important changes in relative job satisfaction of women and men with university and higher qualifications. What is not clear is the extent to which this change is due to changes in the quality of jobs secured by university qualified men and women, or revised expectations associated with gaining a degree.

A possible explanation for the higher job satisfaction observed for women is selection. Given wider social acceptance of roles outside of paid employment for women, women may be more likely than men to leave employment when they are dissatisfied with their jobs. As noted above, a number of other studies have formally tested for selection effects and found these not to be important (Clark 1997; Hauret and Williams 2017; Perugini and Vladisavljević 2019), but this study does not directly test for selection effects. This is a limitation of the current paper and a potential extension for future work.

The sudden convergence of mean job satisfaction ratings for male and female workers in Australia from 2020 to 2022 also warrants further investigation as future waves of HILDA data become available. The timing of the increase in relative job satisfaction for men suggests it is not associated with the COVID-19 pandemic. There is no clear evidence, as yet, that this convergence will mark the demise of the paradox of the contented Australian female worker.

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Appendix Table A1. Job satisfaction: OLS and ordered probit regression results, HILDA Waves 1–22

Independent variable	OLS	Ordered Probit[a]
Female	0.073*** (0.000)	0.073*** (0.000)
Age	-0.037*** (0.000)	-0.032*** (0.000)
Age-squared	0.001*** (0.000)	0.001*** (0.000)
Has disability	-0.093*** (0.000)	-0.065*** (0.000)
Born in:		
Australia	–	–
English speaking country	-0.122*** (0.000)	-0.068*** (0.005)
Non-English speaking country	-0.113*** (0.000)	-0.080*** (0.000)
Highest qualification:		
Post-graduate	-0.500*** (0.000)	-0.475*** (0.000)
Degree	-0.465*** (0.000)	-0.441*** (0.000)
Diploma	-0.292*** (0.000)	-0.285*** (0.000)
Certificate III/IV	-0.217*** (0.000)	-0.209*** (0.000)
Completed Year 12	-0.214*** (0.000)	-0.215*** (0.000)
Did not complete Year 12	–	–
Lives in:		
Major capital city	–	–
Inner regional	0.124*** (0.000)	0.109*** (0.000)
Outer regional/remote	0.180*** (0.000)	0.169*** (0.000)
SES of neighborhood (decile)	-0.006** (0.019)	-0.006*** (0.001)
Family status:		
Married, no children	–	–
Married, child aged 0-4	0.078*** (0.000)	0.065*** (0.000)
Married, child aged 5-14	0.100*** (0.000)	0.082*** (0.000)
Married, child age 15-24	0.093*** (0.000)	0.071*** (0.000)
Single, no children	0.024 (0.119)	0.009 (0.442)
Single, child aged 0-4	0.138** (0.018)	0.127*** (0.008)
Single, child aged 5-14	0.107*** (0.003)	0.094*** (0.002)
Single, child age 15-24	-0.010 (0.794)	0.004 (0.898)

Appendix Table A1. continued

Independent variable	OLS	Ordered Probit[a]
Occupation:		
Manager	-0.057*** (0.001)	-0.044*** (0.002)
Professional	–	–
Technician/trade worker	-0.012 (0.595)	0.002 (0.918)
Community/personal service worker	-0.028 (0.207)	-0.014 (0.464)
Clerical/administrative worker	-0.078*** (0.000)	-0.052*** (0.001)
Sales worker	-0.209*** (0.000)	-0.180*** (0.000)
Machinery operator/driver	-0.165*** (0.000)	-0.125*** (0.000)
Labourer	-0.278*** (0.000)	-0.206*** (0.000)
Firm sector:		
Private for-profit	–	–
Private not-for-profit	0.200*** (0.000)	0.160*** (0.000)
Government business	0.191*** (0.000)	0.161*** (0.000)
Public sector	0.252*** (0.000)	0.205*** (0.000)
Other	0.182*** (0.001)	0.184*** (0.000)
Employment contract:		
Self-employed/employer	0.257*** (0.000)	0.223*** (0.000)
Fixed term contract	-0.029** (0.046)	-0.025** (0.035)
Casual contract	-0.084*** (0.000)	-0.065*** (0.000)
Permanent/ongoing	–	–
Other	-0.173* (0.067)	-0.062 (0.393)
Usual no. hours worked per week:		
0 to 15 hours	0.053** (0.015)	0.083*** (0.000)
16 to 30 hours	-0.002 (0.903)	0.017 (0.214)
31 to 38 hours	-0.016 (0.201)	-0.011 (0.284)
39 to 44 hours	–	–
45 to 54 hours	-0.054*** (0.000)	-0.058*** (0.000)
55 hours or more	-0.167*** (0.000)	-0.151*** (0.000)

Appendix Table A1. continued

Independent variable	OLS	Ordered Probit[a]
Real hourly wage (log of)	0.251*** (0.000)	0.198*** (0.000)
Union member	-0.095*** (0.000)	-0.083*** (0.000)
Years in current occupation	-0.014*** (0.000)	-0.013*** (0.000)
Years in occupation squared/100	0.028*** (0.000)	0.028*** (0.000)
Years with current employer	-0.023*** (0.000)	-0.022*** (0.000)
Years current employer squared/100	0.041*** (0.000)	0.040*** (0.000)
Works non-standard hours	-0.145*** (0.000)	-0.123*** (0.000)
Works some hours from home	0.123*** (0.000)	0.105*** (0.000)
Employed by labour hire firm	-0.115*** (0.000)	-0.079*** (0.001)
Has supervisory responsibilities	-0.017* (0.092)	-0.020** (0.012)
Constant[a]	7.739*** (0.000)	
Observations	169,728	169,728
Number of individuals	24,609	24,609
Obs/individual: min	1	1
Average	6.9	6.9
Max	22	22
Wald chi-square	2,852 (0.000)	3,044 (0.000)

Notes: Robust p-values in parentheses; *** p<0.01, ** p<0.05, * p<0.10; a. The ten intercept terms for the cut-points in the ordered probit model not reported.

Appendix Table A2. OLS regression coefficients and means by gender, pooled 2001-2005 sample

	Males		Females		Mean	
	Coef.	P>t	Coef.	P>t	Males	Females
Job satisfaction					7.55	7.73
Age	-0.094	0.00	-0.043	0.00	37.47	37.25
Age-squared	0.001	0.00	0.001	0.00	1,569	1,547
Has disability	-0.162	0.00	-0.272	0.00	0.13	0.12
Born in:						
Australia	–		–		0.79	0.80
English speaking country	-0.129	0.06	0.059	0.41	0.10	0.09
Non-English speaking country	-0.072	0.30	-0.063	0.34	0.11	0.11
Highest qualification:						
Post-graduate	-0.439	0.00	-0.723	0.00	0.04	0.03
Degree	-0.455	0.00	-0.618	0.00	0.18	0.25
Diploma	-0.263	0.00	-0.496	0.00	0.09	0.10
Certificate III/IV	-0.189	0.00	-0.296	0.00	0.27	0.13
Completed Year 12	-0.119	0.05	-0.292	0.00	0.16	0.18
Did not complete Year 12	–		–		0.26	0.32
Lives in:						
Major capital city	–		–		0.69	0.69
Inner regional	0.140	0.01	0.198	0.00	0.20	0.20
Outer regional/remote	0.225	0.00	0.217	0.00	0.11	0.11
SES of neighborhood (decile)	-0.006	0.45	-0.019	0.01	5.77	5.92
Family status:						
Married, no children	–		–		0.27	0.27
Married, child aged 0-4	0.129	0.03	0.144	0.04	0.15	0.10
Married, child aged 5-14	0.046	0.47	-0.038	0.57	0.16	0.17
Married, child age 15-24	0.116	0.13	0.137	0.07	0.07	0.08
Single, no children	0.023	0.66	-0.052	0.35	0.33	0.30
Single, child aged 0-4	0.634	0.09	-0.162	0.36	0.00	0.01
Single, child aged 5-14	0.157	0.45	-0.058	0.56	0.01	0.04
Single, child age 15-24	-0.200	0.47	-0.076	0.58	0.01	0.03
Occupation:						
Manager	0.056	0.34	-0.044	0.57	0.14	0.08
Professional	–		–		0.19	0.26
Technician/trade worker	0.048	0.45	0.118	0.22	0.21	0.04
Community/personal service worker	0.088	0.35	-0.032	0.66	0.06	0.14
Clerical/administrative worker	-0.245	0.00	-0.053	0.38	0.08	0.25
Sales worker	-0.105	0.22	-0.256	0.00	0.07	0.13
Machinery operator/driver	-0.025	0.76	0.118	0.49	0.11	0.01
Labourer	-0.187	0.02	-0.244	0.01	0.13	0.08

Appendix Table A2. continued

	Males		Females		Mean	
	Coef.	P>t	Coef.	P>t	Males	Females
Firm sector:						
Private for-profit	–		–		0.76	0.62
Private not-for-profit	0.342	0.00	0.241	0.00	0.03	0.08
Government business	0.079	0.33	0.124	0.15	0.06	0.05
Public sector	0.262	0.00	0.252	0.00	0.14	0.23
Other	0.192	0.20	0.339	0.00	0.01	0.02
Employment contract:						
Self-employed/employer	0.307	0.00	0.309	0.00	0.08	0.04
Fixed term contract	0.065	0.26	-0.035	0.54	0.08	0.09
Casual contract	-0.224	0.00	-0.070	0.19	0.16	0.26
Permanent/ongoing	–		–		0.67	0.60
Other	0.067	0.80	-0.793	0.03	0.00	0.00
Usual no. hours worked per week:						
0 to 15 hours	0.162	0.08	0.280	0.00	0.07	0.19
16 to 30 hours	-0.083	0.30	0.145	0.01	0.07	0.26
31 to 38 hours	-0.012	0.81	0.070	0.18	0.18	0.23
39 to 44 hours	–		–		0.26	0.18
45 to 54 hours	0.054	0.22	0.036	0.56	0.26	0.11
55 hours or more	-0.088	0.15	-0.106	0.29	0.15	0.04
Real hourly wage (log of)	0.232	0.00	0.198	0.00	3.42	3.33
Union member	-0.048	0.29	-0.239	0.00	0.29	0.28
Years in current occupation	-0.013	0.04	-0.007	0.31	9.40	7.87
Years in occupation squared	0.000	0.05	0.000	0.06	186	137
Years with current employer	0.009	0.18	0.005	0.53	6.91	5.79
Years current employer squared	0.000	0.12	0.000	0.46	114	79
Works non-standard hours	-0.190	0.00	-0.252	0.00	0.25	0.25
Works some hours from home	0.021	0.65	0.100	0.04	0.21	0.20
Employed by labour hire firm	-0.294	0.01	-0.114	0.30	0.03	0.03
Has supervisory responsibilities	0.084	0.03	0.016	0.67	0.53	0.42
Constant	8.424	0.00	8.062	0.00		
Observations	16,277		14,913			
Individuals	5,537		5,311			
R-squared	0.05		0.05			

Notes: p-values based on robust standard errors.

Appendix Table A3. OLS regression coefficients and means by gender, pooled 2006-2010 sample

	Males		Females		Mean	
	Coef.	P>t	Coef.	P>t	Males	Females
Job satisfaction					7.61	7.71
Age	-0.091	0.00	-0.025	0.02	37.83	37.87
Age-squared	0.001	0.00	0.000	0.00	1,614	1,615
Has disability	-0.073	0.13	-0.222	0.00	0.12	0.13
Born in:						
Australia	–		–		0.82	0.83
English speaking country	-0.189	0.01	0.020	0.78	0.09	0.08
Non-English speaking country	-0.005	0.94	-0.059	0.32	0.09	0.09
Highest qualification:						
Post-graduate	-0.411	0.00	-0.466	0.00	0.04	0.04
Degree	-0.359	0.00	-0.415	0.00	0.19	0.26
Diploma	-0.305	0.00	-0.278	0.00	0.08	0.10
Certificate III/IV	-0.073	0.19	-0.180	0.00	0.27	0.16
Completed Year 12	-0.140	0.02	-0.148	0.01	0.17	0.18
Did not complete Year 12	–		–		0.24	0.26
Lives in:						
Major capital city	–		–		0.68	0.68
Inner regional	0.131	0.01	0.138	0.00	0.20	0.20
Outer regional/remote	0.104	0.07	0.181	0.00	0.12	0.11
SES of neighborhood (decile)	0.002	0.78	-0.020	0.01	5.84	5.86
Family status:						
Married, no children	–		–		0.30	0.29
Married, child aged 0-4	-0.044	0.42	0.069	0.25	0.15	0.10
Married, child aged 5-14	-0.016	0.78	0.059	0.34	0.14	0.15
Married, child age 15-24	-0.082	0.24	-0.008	0.91	0.07	0.09
Single, no children	-0.063	0.18	-0.082	0.11	0.33	0.29
Single, child aged 0-4	-0.248	0.62	0.060	0.67	0.00	0.01
Single, child aged 5-14	-0.171	0.38	0.020	0.85	0.01	0.04
Single, child age 15-24	-0.367	0.17	-0.095	0.39	0.01	0.03
Occupation:						
Manager	0.036	0.50	-0.126	0.05	0.14	0.08
Professional	–		–		0.19	0.27
Technician/trade worker	0.029	0.62	-0.098	0.34	0.23	0.04
Community/personal service worker	0.062	0.46	-0.060	0.34	0.06	0.15
Clerical/administrative worker	-0.223	0.00	-0.040	0.46	0.08	0.24
Sales worker	-0.176	0.03	-0.207	0.00	0.07	0.14
Machinery operator/driver	-0.268	0.00	-0.187	0.23	0.11	0.01
Labourer	-0.124	0.09	-0.316	0.00	0.12	0.07

Appendix Table A3. continued

	Males		Females		Mean	
	Coef.	P>t	Coef.	P>t	Males	Females
Firm sector:						
Private for-profit	–		–		0.77	0.61
Private not-for-profit	0.392	0.00	0.173	0.00	0.04	0.09
Government business	0.160	0.04	0.111	0.13	0.05	0.05
Public sector	0.253	0.00	0.172	0.00	0.14	0.24
Other	0.412	0.05	0.259	0.06	0.01	0.01
Employment contract:						
Self-employed/employer	0.177	0.01	0.351	0.00	0.09	0.04
Fixed term contract	-0.028	0.58	-0.133	0.02	0.09	0.09
Casual contract	-0.251	0.00	-0.065	0.23	0.16	0.24
Permanent/ongoing	–		–		0.67	0.62
Other	-0.243	0.40	-1.250	0.00	0.00	0.00
Usual no. hours worked per week:						
0 to 15 hours	0.206	0.01	0.322	0.00	0.07	0.18
16 to 30 hours	0.080	0.27	0.169	0.00	0.07	0.25
31 to 38 hours	-0.067	0.14	0.035	0.43	0.19	0.25
39 to 44 hours	–		–		0.27	0.18
45 to 54 hours	-0.018	0.65	-0.038	0.49	0.26	0.11
55 hours or more	-0.090	0.09	-0.288	0.00	0.14	0.04
Real hourly wage (log of)	0.225	0.00	0.206	0.00	3.52	3.41
Union member	-0.052	0.21	-0.205	0.00	0.25	0.26
Years in current occupation	-0.006	0.31	-0.021	0.00	9.50	8.14
Years in occupation squared	0.000	0.63	0.001	0.00	196	152
Years with current employer	0.008	0.22	0.005	0.52	6.90	6.07
Years current employer squared	0.000	0.47	0.000	0.88	114	88
Works non-standard hours	-0.207	0.00	-0.179	0.00	0.23	0.24
Works some hours from home	0.066	0.09	0.092	0.04	0.20	0.19
Employed by labour hire firm	-0.219	0.02	-0.036	0.76	0.03	0.02
Has supervisory responsibilities	0.064	0.05	0.081	0.02	0.53	0.44
Constant	8.538	0.00	7.638	0.00		
Observations	17,308		16,269			
Individuals	5,520		5,351			
R-squared	0.05		0.05			

Notes: p-values based on robust standard errors.

Appendix Table A4. OLS regression coefficients and means by gender, pooled 2009-2013 sample

	Males		Females		Mean	
	Coef.	P>t	Coef.	P>t	Males	Females
Job satisfaction					7.61	7.69
Age	-0.110	0.00	-0.054	0.00	38.426	38.403
Age-squared	0.001	0.00	0.001	0.00	1,664	1,665
Has disability	-0.082	0.08	-0.211	0.00	0.113	0.125
Born in:						
Australia	–		–		0.80	0.82
English speaking country	-0.159	0.01	0.043	0.49	0.094	0.080
Non-English speaking country	-0.052	0.39	-0.123	0.03	0.102	0.103
Highest qualification:						
Post-graduate	-0.364	0.00	-0.511	0.00	0.050	0.054
Degree	-0.336	0.00	-0.375	0.00	0.197	0.268
Diploma	-0.228	0.00	-0.310	0.00	0.085	0.098
Certificate III/IV	-0.086	0.09	-0.155	0.01	0.285	0.179
Completed Year 12	-0.157	0.01	-0.137	0.01	0.175	0.183
Did not complete Year 12	–		–		0.21	0.22
Lives in:						
Major capital city	–		–		0.69	0.69
Inner regional	0.172	0.00	0.195	0.00	0.194	0.200
Outer regional/remote	0.188	0.00	0.245	0.00	0.112	0.107
SES of neighborhood (decile)	-0.004	0.54	-0.007	0.33	5.793	5.832
Family status:						
Married, no children	–		–		0.31	0.30
Married, child aged 0-4	0.043	0.38	0.130	0.02	0.160	0.103
Married, child aged 5-14	0.047	0.37	0.183	0.00	0.132	0.133
Married, child age 15-24	0.019	0.76	0.084	0.17	0.076	0.095
Single, no children	-0.028	0.53	-0.098	0.03	0.309	0.289
Single, child aged 0-4	-0.171	0.56	0.200	0.09	0.001	0.012
Single, child aged 5-14	0.074	0.69	-0.069	0.48	0.008	0.041
Single, child age 15-24	-0.061	0.75	-0.063	0.56	0.008	0.025
Occupation:						
Manager	0.066	0.17	-0.116	0.05	0.154	0.090
Professional	–		–		0.20	0.27
Technician/trade worker	0.030	0.57	-0.083	0.33	0.219	0.042
Community/personal service worker	0.057	0.47	0.005	0.93	0.062	0.153
Clerical/administrative worker	-0.158	0.01	-0.027	0.59	0.077	0.237
Sales worker	-0.074	0.30	-0.155	0.02	0.061	0.132
Machinery operator/driver	-0.127	0.07	-0.070	0.63	0.108	0.012
Labourer	-0.101	0.13	-0.205	0.02	0.119	0.063

Appendix Table A4. continued

	Males		Females		Mean	
	Coef.	P>t	Coef.	P>t	Males	Females
Firm sector:						
Private for-profit	–		–		0.77	0.60
Private not-for-profit	0.394	0.00	0.172	0.00	0.041	0.109
Government business	0.125	0.07	0.083	0.23	0.047	0.046
Public sector	0.252	0.00	0.158	0.00	0.140	0.244
Other	0.598	0.06	-0.081	0.76	0.001	0.002
Employment contract:						
Self-employed/employer	0.175	0.00	0.310	0.00	0.080	0.037
Fixed term contract	-0.071	0.12	-0.065	0.16	0.087	0.095
Casual contract	-0.323	0.00	-0.116	0.02	0.157	0.226
Permanent/ongoing	–		–		0.67	0.64
Other	0.003	0.99	-0.324	0.33	0.002	0.001
Usual no. hours worked per week:						
0 to 15 hours	0.125	0.09	0.192	0.00	0.070	0.168
16 to 30 hours	0.072	0.27	0.056	0.23	0.078	0.262
31 to 38 hours	-0.085	0.04	-0.017	0.68	0.208	0.261
39 to 44 hours	–		–		0.27	0.17
45 to 54 hours	-0.067	0.06	-0.096	0.07	0.245	0.100
55 hours or more	-0.164	0.00	-0.354	0.00	0.129	0.037
Real hourly wage (log of)	0.280	0.00	0.239	0.00	3.577	3.458
Union member	-0.092	0.02	-0.197	0.00	0.245	0.264
Years in current occupation	0.002	0.75	-0.016	0.01	9.725	8.417
Years in occupation squared	0.000	0.74	0.000	0.01	202	158
Years with current employer	0.009	0.11	0.000	0.99	7.030	6.434
Years current employer squared	0.000	0.16	0.000	0.85	117	96
Works non-standard hours	-0.179	0.00	-0.121	0.00	0.226	0.222
Works some hours from home	0.120	0.00	0.118	0.00	0.194	0.188
Employed by labour hire firm	-0.193	0.03	-0.212	0.09	0.025	0.018
Has supervisory responsibilities	0.058	0.06	0.056	0.07	0.525	0.439
Constant	8.700	0.00	7.955	0.00		
Observations	20,619		19,326			
Individuals	6,847		6,674			
R-squared	0.05		0.05			

Notes: p-values based on robust standard errors.

Appendix Table A5. OLS regression coefficients and means by gender, pooled 2018-2022 sample

	Males		Females		Mean	
	Coef.	P>t	Coef.	P>t	Males	Females
Job satisfaction					7.80	7.85
Age	-0.068	0.00	-0.046	0.00	39.43	39.50
Age-squared	0.001	0.00	0.001	0.00	1,748	1,756
Has disability	-0.178	0.00	-0.203	0.00	0.11	0.12
Born in:						
Australia	–		–		0.83	0.82
English speaking country	-0.065	0.25	-0.091	0.10	0.08	0.07
Non-English speaking country	-0.033	0.54	-0.087	0.09	0.09	0.11
Highest qualification:						
Post-graduate	-0.353	0.00	-0.353	0.00	0.07	0.09
Degree	-0.325	0.00	-0.234	0.00	0.22	0.31
Diploma	-0.203	0.00	-0.171	0.01	0.09	0.11
Certificate III/IV	-0.115	0.03	-0.057	0.33	0.29	0.19
Completed Year 12	-0.165	0.00	-0.067	0.24	0.18	0.16
Did not complete Year 12	–		–		0.16	0.13
Lives in:						
Major capital city	–		–		0.69	0.69
Inner regional	0.069	0.06	0.171	0.00	0.21	0.21
Outer regional/remote	0.147	0.01	0.269	0.00	0.10	0.10
SES of neighborhood (decile)	-0.015	0.01	-0.001	0.79	5.70	5.75
Family status:						
Married, no children	–		–		0.31	0.30
Married, child aged 0-4	0.061	0.13	0.066	0.17	0.17	0.13
Married, child aged 5-14	0.044	0.33	0.085	0.08	0.14	0.15
Married, child age 15-24	-0.066	0.22	0.114	0.04	0.07	0.08
Single, no children	-0.065	0.10	-0.044	0.26	0.29	0.27
Single, child aged 0-4	-0.133	0.72	-0.057	0.67	0.00	0.01
Single, child aged 5-14	-0.245	0.17	0.030	0.70	0.01	0.04
Single, child age 15-24	-0.211	0.25	-0.205	0.03	0.01	0.03
Occupation:						
Manager	0.041	0.34	-0.067	0.16	0.17	0.11
Professional	–		–		0.21	0.31
Technician/trade worker	0.022	0.65	0.115	0.13	0.21	0.04
Community/personal service worker	0.065	0.36	0.059	0.23	0.07	0.16
Clerical/administrative worker	-0.145	0.02	0.028	0.54	0.06	0.20
Sales worker	-0.160	0.02	-0.036	0.54	0.06	0.10
Machinery operator/driver	-0.057	0.36	0.104	0.50	0.11	0.01
Labourer	-0.030	0.63	-0.053	0.51	0.11	0.05

Appendix Table A5. continued

	Males		Females		Mean	
	Coef.	P>t	Coef.	P>t	Males	Females
Firm sector:						
Private for-profit	–		–		0.77	0.58
Private not-for-profit	0.276	0.00	0.096	0.03	0.05	0.12
Government business	0.155	0.03	0.118	0.08	0.04	0.04
Public sector	0.210	0.00	0.129	0.00	0.14	0.26
Other	0.237	0.41	0.370	0.15	0.00	0.00
Employment contract:						
Self-employed/employer	0.309	0.00	0.538	0.00	0.08	0.04
Fixed term contract	-0.114	0.01	0.041	0.30	0.07	0.10
Casual contract	-0.073	0.13	0.064	0.14	0.16	0.19
Permanent/ongoing	–		–		0.69	0.67
Other	0.372	0.15	0.199	0.40	0.00	0.00
Usual no. hours worked per week:						
0 to 15 hours	-0.134	0.06	0.082	0.14	0.07	0.13
16 to 30 hours	-0.195	0.00	-0.023	0.57	0.09	0.27
31 to 38 hours	-0.024	0.47	0.046	0.20	0.24	0.30
39 to 44 hours	–		–		0.27	0.17
45 to 54 hours	-0.107	0.00	-0.150	0.00	0.23	0.10
55 hours or more	-0.262	0.00	-0.338	0.00	0.10	0.03
Real hourly wage (log of)	0.220	0.00	0.271	0.00	3.65	3.57
Union member	-0.096	0.01	-0.183	0.00	0.19	0.24
Years in current occupation	-0.007	0.08	-0.009	0.06	9.96	8.95
Years in occupation squared	0.000	0.17	0.000	0.04	208	175
Years with current employer	0.002	0.76	-0.005	0.37	7.30	6.83
Years current employer squared	0.000	0.77	0.000	0.45	124	106
Works non-standard hours	-0.157	0.00	-0.208	0.00	0.20	0.20
Works some hours from home	0.121	0.00	0.082	0.01	0.27	0.30
Employed by labour hire firm	-0.109	0.17	-0.143	0.17	0.02	0.01
Has supervisory responsibilities	-0.030	0.27	0.057	0.04	0.50	0.42
Constant	8.634	0.00	7.742	0.00		
Observations	21,584		21,527			
Individuals	6,630		6,798			
R-squared	0.03		0.04			

Notes: p-values based on robust standard errors.