Casual Employment and Long-term Wage Outcomes

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Abstract

Temporary and other forms of non-standard employment are an important feature of modern labour markets. Yet relatively little is known about how much and under what circumstances such employment arrangements impact on long-term wage outcomes. Using longitudinal data from the Household, Income and Labour Dynamics in Australia (HILDA) Survey spanning the period 2001 to 2013, we examine how employment status earlier in a working career is associated with subsequent wage dynamics. Particular attention is paid to how wage trajectories vary with gender and age. Estimates from a series of panel data models of real hourly wages reveal that among men there is an average long-run penalty from casual employment of about 10% that does not vary with experience, suggestive of scarring effects. This interaction with experience, however, varies with age, with the casual wage penalty diminishing over time among younger male workers. Among women the estimated average long-run wage penalty associated with casual employment is both much smaller and less robust. Further, with the exception of older female workers, this penalty disappears with time. We argue that expectations and norms about “ideal careers” may be an important explanatory factor underlying the casual employment wage penalty for men.

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Introduction

Recent decades have seen increased interest among researchers and policy-makers alike in the changing nature of work and, more specifically, the role played by temporary and casual employment. While these flexible forms of employment can play a useful role in helping firms adjust to external shocks and deal with uncertainty, they are also usually equated with poor-quality jobs (e.g., Dekker and van der Veen, forthcoming; Kalleberg et al., 2000; McGovern et al., 2004; Neinhüser and Matiaske, 2006; OECD, 2014). This reflects both the inherently insecure nature of temporary jobs, and the lower earnings and lesser access to other employment benefits that is often associated with them. The OECD (2015), for example, reports that, in almost all countries for which there are suitable data, there are marked wage gaps between workers on temporary or casual contracts and those on otherwise comparable standard, or so-called 'permanent', employment contracts. Despite the presence of such wage penalties, it does not follow that temporary jobs will necessarily be associated with unfavourable labour market outcomes in the future. This will depend on the extent to which temporary job spells affect subsequent earnings mobility.

While an extensive literature exists on the extent to which temporary employment is a persistent state, relatively little research has explicitly examined the link between non-standard forms of employment and long-run wage outcomes. Further, much of the research in this area has focused on youth and the consequences for future earnings of commencing a career in a temporary job (e.g., de Lange et al., 2014; Gebel, 2010; Pavlopoulos, 2013). Notable exceptions here are Booth et al. (2002) and Fuller and Stecy-Hildebrandt (2014). These two studies used panel survey data from Britain and Canada, respectively, reporting evidence that temporary and / or casual jobs are typically associated with wage penalties that in most situations are quite persistent (at least when measured over 5- to 7-year windows). But will such findings extend to other institutional settings and, more specifically, to

Australia? By international standards Australia is relatively unusual. First, it is casual employment, rather than fixed-term contracts, that is the most pervasive form of non-standard employment arrangement in Australia. Second, the extent of casual employment in Australia is relatively high, currently accounting for somewhere between 22% and 24% of dependent employment (depending on the data source and the definition used). Third, industrial laws in Australia require the payment of a substantial hourly wage premium to most casual workers (that is, those covered by industry awards), which helps to at least partly compensate for the loss of other benefits (notably paid leave entitlements).[1]

Another weakness of most previous studies is that the effects of temporary employment are typically captured by a dummy variable identifying a single job or employment spell, usually the first spell observed. As such, these studies will likely underestimate the effects of sustained patterns of temporary employment. After all, temporary jobs are, by definition, short lasting. We thus expect the effects of a single episode of temporary employment on earnings mobility over the long run to be weak. In contrast, we expect stronger associations with employment patterns that are characterised by frequent temporary employment spells over a sustained period.

We revisit the issue of whether or not non-standard employment is associated with detrimental impacts on future wage outcomes, focusing on the Australian labour market and the role of casual employment. We advance research by using longitudinal data from the Household, Income and Labour Dynamics in Australia (HILDA) Survey to examine how hourly wage outcomes over the period 2004 to 2013 vary with a summary measure describing employment status over the period 2001 to 2003. Our empirical approach improves on existing approaches in that it summarises employment status at multiple points in time (as opposed to a single episode). We also examine how wage differentials (between workers in standard and non-standard employment) both evolve over time and vary with age and gender.

Theoretical Insights

Non-standard Employment and Scarring Effects

A priori, it is unclear whether, and to what extent, temporary and other forms of non-standard employment hinder subsequent career development. As observed by Booth et al. (2002: F198), if temporary jobs are part of a search process during which workers, who are still unsure about their career and location preferences, gradually learn about which types of jobs best suit their skills and interests, then adverse long-run consequences should not exist. Indeed, econometric models of wages dynamics using US data have found positive returns to job shopping, though the magnitude of estimates are highly variable (cf. Altonji et al., 2013; Topel and Ward, 1992). Relatedly, if temporary employment is used by employers as a probationary device because of uncertainty about workers' abilities, then we would expect the size of any initial wage penalty to decline over time as experience is accumulated and the worker's true productivity revealed.

On the other hand, because the incentives for firms and workers to invest in training depend on the time period over which the benefits of training can be realised, persons in temporary jobs will be both less attractive training propositions to employers and more reluctant to participate in firm-specific training than persons employed on a more permanent basis. Survey evidence is mostly consistent with this hypothesis (e.g., Booth et al., 2002; Draca and Green, 2004; Neinhüser and Matiaske, 2006). Workers in temporary jobs will thus accumulate less human capital on the job than permanent workers, which in human capital models results in lower rates of future wages growth.

Increased wage penalties are also suggested by signalling theory. If workers are sorted into temporary and permanent jobs based on ability, with lower ability workers getting the temporary jobs, then securing a temporary job will send a negative signal to future employers about a worker's ability. This will adversely affect their future employment and wage prospects. Pedulla (forthcoming), for example, provides evidence from an experimental study in the US of employer responses to a sample of job applications in support of this hypothesis.

Very differently, it is also often argued that temporary and other forms of non-standard employment contracts are devices used by employers to further entrench differences between groups in the labour market. In its simplest form, it is argued that labour markets are segmented into a primary market with "good jobs" (providing high wages, fringe benefits, and good prospects for career development) and a secondary market with "bad jobs" (e.g., Doeringer and Piore, 1971), with temporary and other forms of nonstandard work playing a critical role in the allocation of workers to the secondary sector (e.g., Hudson, 2007). A key prediction of this theory, is that initial wage differences between temporary and permanent job holders should increase over time, as permanent job holders access opportunities for promotion and career development in the primary labour market that are not available to temporary job holders who are effectively trapped in the secondary labour market.

Evidence on whether associations between wage outcomes and different types of employment arrangements persist over long periods is limited. A brief summary of what we believe are the most relevant published studies is provided in Table 1. Our reading of this body of evidence suggests that for some of the labour market groups which are most exposed to temporary employment – new labour market entrants and low wage earners – effects on wages in the longer term are relatively modest, if not zero. Studies of broader population samples suggest more mixed findings. But, as previously discussed, the studies of both Booth et al. (2002) and Fuller and Stecy-Hildebrandt (2014) suggest the presence of temporary employment wage penalties that can be long lasting, and persisting long after the worker has secured more permanent employment. Further, the magnitude of such effects tends to vary with both gender and the nature of the employment arrangement.

Overall, we expect lower hourly wages for workers in temporary employment relative to their counterparts in regular employment. But whether this initial wage gap narrows or widens over time is ambiguous.

Heterogeneity in Wage Outcomes: A Gendered Perspective

The relatively high proportion of women in temporary and casual forms of employment fuels the question: why do more women than men select into non-standard forms of employment?[2] A prominent stream of literature on gender specialisation theorises about how (labour) markets encourage women to specialise in unpaid and household work and men in paid employment (Becker, 1981). This same specialisation logic can also be used to explain the apparent preference of women, and especially mothers, for non-standard forms of employment. In short, in the presence of children, the optimal strategy for maximising household earnings is for men, who usually have greatest (future) resources and economic advantage (higher starting wages and steeper wage growth), to concentrate on regular paid employment, leaving other (non-standard forms) employment options to their wives (for a review, see England, 2005). Indeed, women more than men, trade legal protection and employment stability for more flexible and insecure forms of employment to cope with the time consuming demands of work and family care. This preference of women to work in temporary employment is not only driven by unequal individual resources between partners, but may also reflect women's desires (perhaps driven by a sense of moral obligation) to retain primary responsibility for managing household activities and childcare. Indeed, many women may have strong preferences to retain control over a domain in which they feel expert in; preferences that are reinforced by internalised attitudes and norms about gender roles (Craig and Mullan, 2010; England, 2005; Hook, 2006; García-Manglano, 2015). In this case, both individual and household resources, together with norms and attitudes, govern how men and women select into specific forms of employment and non-employment. In countries with more traditional gender ideologies, for example, women will likely form male-breadwinner households and withdraw from employment during child-rearing periods (Craig and Mullan, 2011; Gornick and Meyers, 2009). In other countries with more egalitarian gender rules, it is more common that both men and women engage in full-time (regular) employment during child rearing periods (Gornick and Meyers, 2009). Despite progressive gender attitudes, Australia falls in between these two extremes, with men in most households working full-time and women often engaging in part-time and casual employment (Craig and Mullan, 2010). This suggests that men's and women's preferences about the timing and participation in temporary employment are mediated by cultural norms and attitudes about gender roles. These are then reinforced through social structures (such as work/family policies, gender equality schemes or work arrangements) that reflect the social organisation of work and care in a specific country (Craig and Mullan, 2011; Hook, 2006).

But if more women than men select into temporary employment, does temporary employment pay (even) less? The answer to this question is complex and subject to cultural influences and group-prejudices related to a specific sex. Focussing on women first, studies often show that occupations with a higher proportion of female workers tend to pay less, suggestive of a general devaluation of women's labour (for a review, see England et al., 2005). Additionally, in many countries, non-standard employment is clustered in low-paying sectors and occupations that offer limited promotion opportunities (OECD, 2014). On the one hand, these two factors combined suggest that temporary employment could become a "double burden" for women because of cultural bias against women's work and because non-standard employment is usually associated with poor quality jobs. On the other hand, the cultural bias related to temporary employment against women should carry less of a negative effect on their future wage outcomes if women adhere to the expected employment trajectories that are typical of women within a particular age category.

For men in temporary employment, stereotyping and group generalisations about their average performance will negatively influence employers' pay practices (England et al., 1998; Reskin and Roos, 1990). Men's engagement in temporary employment may raise red flags to employers and be perceived as a negative signal about their productivity, which translates into lower wages. In essence, a temporary or casual job is a violation of normative expectations with regard to a man's "ideal" employment career.

Overall, we expect casual employment to inflict negative wage outcomes among men if these clash with the broader career expectations. For women, the direction of the wage effects is more uncertain.

Heterogeneity in Wage Outcomes: A Life-Course Perspective

Expectations about the heterogeneity of wage outcomes over time are best understood within a life-course perspective (Giele and Elder, 1998). Developed to understand how disruptive events (such as divorce or job loss) modify people's positions as they transition from one event to the next, life-course theory emphasises the combination of the timing, sequencing and duration of such events in influencing future career and earnings dynamics. Within this perspective, the effect of temporary employment on subsequent wage outcomes is moderated by the age of the worker (timing), the time since the initial temporary employment (duration) and the order of employment patterns following the initial temporary contract (sequencing). As such, wage effects associated with temporary employment are expected to vary as people pass through different life cycle stages, transitioning from singles to couples, to families with young and adult children, and back to couples or "empty nesters" and retirement.

Workers transitioning from singles to couples, for example, are more likely to be in the beginning of their careers and more likely to use temporary jobs as part of an initial search process. At slightly older ages, and following the arrival of children, workers may use temporary employment to balance work and family responsibilities. Still other workers may consider temporary employment as a way to re-enter the labour market after a period of job interruption. And then as retirement age draws near, some workers may use temporary and casual employment as a vehicle for gradual withdrawal from the workforce. Employers, on the other hand, may use temporary jobs as a mechanism to identify the abilities of job applicants. As such, it follows that temporary employment among younger workers should be associated with relatively benign wage growth outcomes. For mothers, who use temporary employment to bridge the pre-school period of their children, wage penalties should be moderate and diminish as both children grow up and the mothers transition into regular employment. Temporary employment among older workers may result in faster re-employment on the one hand, but it also conveys a negative signal to the employers given such employment is often associated with both limited adaptability to new (technological) skills (Kuhn, 2002) and outdated skills that are no longer suited to newly created jobs (Blossfeld et al., 2006). This suggests that temporary employment will produce weak and moderate negative wage effects among respectively young and middle-aged workers, but much larger negative effects among older workers.

Data and Method

Method

We begin with a random-intercepts model of the form:

$$\ln w_{it} = \alpha + \beta' E_{i0} + \delta' x_{it} + \mu_i + \varepsilon_{it} \quad (1)$$

where w_{it} is the real hourly wage of worker i in year t (which covers the period 2004 to 2013), E_{i0} is a vector of variables identifying the initial employment status of individual i (which is measured using data from survey waves 1 to 3), and x_{it} is a vector of variables expected to be associated with hourly wages.

Consistent with the life-course perspective, the initial employment status variable summarises the timing and sequence of observed labour market states at multiple points in time. This is one point of difference between the approach used in this study and that employed in previous research, which, as noted earlier, has mostly relied on the first employment status observed.

A key statistical issue is non-random selection into the initial employment state that is correlated with unobservable traits. Purging the effects of such unobserved heterogeneity using conventional fixed-effects (FE) estimation, however, is not feasible given our primary variable of interest (E_{i0}) is time invariant. Instead, we estimate a model with correlated random effects, sometimes also known as the 'within-between estimator' (Bell and Jones, 2015) or the 'hybrid FE model' (Allison, 2009). The hybrid FE model decomposes each time-varying covariate into a within-person component (i.e., the deviation from that individual-specific mean) and a between-person component (i.e., the mean of each individual-specific variable). As Dieleman and Templin (2014) show, asymptotically this estimator is equivalent to the conventional FE estimator, and in small samples, while the FE estimator is the unambiguously preferred estimator for the purposes of prediction, the hybrid model generally produces marginal effects that are more precise. This inclusion of correlated random effects is another feature of this study that marks it as distinct from others in this literature.[3] This model takes the form:

$$\ln w_{it} = \alpha + \beta' E_{i0} + \delta'(x_{it} - \overline{x}_i) + \gamma' \overline{x}_i + \mu_i + \varepsilon_{it} \qquad (2)$$

In addition, the basic specification also includes time (i.e., survey year) dummies.

We estimate all models separately for men and women to test directly for gendered wage differences. Expectations about the shape of wage penalties (growing or shrinking over time) are tested through inclusion of interaction terms in equation (2) between the initial employment status variable E_{i0} and deviations from potential work experience. Similarly, wage outcomes across different age groups will be estimated through the inclusion of three-way interactions between the initial employment status variable, deviations from potential work experience and age.

Sample

The sample used for this analysis is restricted to persons observed in survey waves 4 through 13 who were aged 18 to 54, reported receiving wages or salary, and were not in full-time education. Furthermore, individuals also had to be observed at some point during waves 1 to 3. This left us with a working sample of 48,534 wage observations covering 8,168 persons.

Outcome variable

The dependent variable is the log of hourly wages in the main job deflated by the Consumer Price Index. Hourly wages were constructed by dividing estimated usual gross weekly wages and salaries by usual weekly working hours (where usual working hours includes both paid and unpaid overtime).[4]

Initial employment status

To measure initial labour force and employment status we created a series of categorical variables summarising observed labour force status at the time of interview across the first three survey waves. We distinguish between regular employment (defined here to include employees in ongoing or permanent jobs, as well as employees on fixed-term contracts), casual employment, self-employment, unemployment, and being outside the labour force (where the concepts of employment, unemployment and labour force are defined and measured in line with recommended International Labour Organization standards). We also distinguish between persons who are continuously in one state and those that switch states.[5]

After some experimentation, we opted for ten major groups, which between them described the employment patterns of over 97% of the selected sample.[6] As reported in Table 2, the numerically largest category is persons observed only in regular employment – 48% of men and 37% of women. Persons observed only in casual employment are a relatively small group, representing about 6.5% of men and 11% of women in our sample. There are, however, a considerable proportion who moved between casual and regular employment, but without being observed as jobless at any survey date (over 9% of men and over 12% of women), and between self-employment, casual employment and regular employment (11% of men and 8% of women). And then, of course, there are numerous persons who moved between casual employment and joblessness, though typically this also involves episodes of regular employment – persons only observed in both casual employment and joblessness are relatively rare (just 0.4% of men and 1% of women in our sample). We also separately identify those persons who move between regular employment and unemployment, and those persons observed as continuously jobless (most of whom were not looking for a job). The remaining cases (other persons who moved between joblessness and regular employment and / or self-employment) are lumped into a single "other patterns" category.

Somewhat contentious is our treatment of fixed-term contract workers as equivalent to permanent or ongoing employees. Nevertheless, simple regression equations indicate no significant differences in the wages of these two groups. This treatment is also consistent with other research that has found relatively little evidence that fixed-term contract workers fare worse in the Australian labour market than permanent employees (e.g., Buddelmeyer et al., 2015; Watson, 2005). This stands in marked contrast to labour markets in other developed countries, especially in Europe, where temporary jobs, defined by the presence of an employment contract with a fixed duration, have usually been found to be associated with large wage penalties (OECD, 2015). We suspect that a large part of the explanation for Australian exceptionalism here is simply the ease with which casual employment contracts can be used by employers in Australia. Casual employment is clearly the preferred mode of labour engagement for many low-wage jobs, presumably because of the greater flexibility it provides employers in terms of both the scheduling of working hours and dismissal and because of the absence of many entitlements, such as paid leave and paid public holidays. In contrast, use of fixed-term contracts in Australia is especially prominent in higher-paying professional occupations (Waite and Will, 2002), which is inconsistent with the notion that such contracts are used to reinforce labour market segmentation.

Control variables

Our models control for a range of differences in human capital, demographic and job characteristics. Human capital and demographic variables include: highest level of educational attainment (7 dummies representing 8 groups), years of job tenure (with the same employer), years of occupation experience, potential labour market experience (years since ceasing full-time education for the first time), country / region of birth (two dummies), the remoteness of the location in which the respondent resides (two dummies), the presence of a long-term health condition or disability that is work limiting, family and relationship status (four dummies identifying single persons, lone persons and couples with and without dependents), and the number of children (within two age bands). Controls for job characteristics include: the type of employment contract held at the time of the survey (i.e., whether employed in a casual job or not), major occupation group (8 dummies), whether working full-time or part-time hours (where full-time is defined as usual weekly work hours of 35 or more), union membership, industry division (18 dummies), and sector (with five dummies identifying whether the employer is a public or private enterprise and a commercial or non-commercial organisation). Summary statistics describing all variables used in the analysis are provided in Table 2.

Results

Table 3 presents the key results from estimation of both equations (1) and (2) for men and women (separately). The random effects (RE) model estimates suggest that there is a relatively large wage penalty associated with a history of extensive casual employment (i.e., more or less continuous over the three years preceding our estimation period), but these effects are larger and stronger for men (24.1%) than women (14.0%).[7]

Given selection into casual employment at time 0 is inversely correlated with unobserved ability (and other productivity-related traits), we expect the RE estimates of the coefficients on casual employment status to be biased downwards. Our results are consistent with this expectation, with the estimates on the casual employment history variables in the hybrid FE model being noticeably smaller in absolute terms. For men with a history of continuous casual employment, a coefficient of -.101 is obtained.[8] Compared with the reference category – men who were continuously in regular employment – this translates into a wage penalty of almost 11%, less than half the magnitude suggested by the simple random effects estimates. Where the earlier history of casual employment is interspersed with periods of regular employment there is also a significant but smaller wage penalty – just over 5%. Choice of estimator makes an even bigger difference for women, with the estimated casual employment wage penalty declining from 14% to less than 5%. However, and very different to men, there is no wage advantage for women who moved between casual and regular employment during the initial three-wave observation window.

But is continuous employment in regular employment the appropriate counterfactual? Following Gebel (2013), it might be argued that a better counterfactual is provided by those persons who opt to wait for a permanent job to come along rather than accept a casual job. We suggest that this reference group is better represented by those observed as both in unemployment and in regular employment (but no other state) during our initial 3-year window. When this group is used as the comparison then the penalty to male wages that is attached to casual employment remains virtually unchanged; the casual employment wage penalty remains close to 10%. In contrast, among women the penalty disappears entirely. Indeed, both the RE and hybrid FE results show that females observed in both regular employment and unemployment in waves 1 through 3 have lower future hourly wages than those observed only in casual employment at these same points in time. This difference is, however, not statistically significant.

Overall, these results lend support to our expectation that casual employment, at least if sustained over an extended period, is associated with long-term wage penalties. These penalties, however, are larger for men than women. Indeed, the magnitude of the estimated penalty for women is both relatively small and sensitive to the choice of comparator group.

Interactions with Experience and Age

We are also interested in knowing whether the wage penalty associated with casual employment grows or shrinks over time. To address this issue we re-estimated our hybrid models after including interaction terms between the initial employment status variables and the within-person deviations in potential work experience (as well as with mean experience). A pictorial summary of the results from this exercise is provided in Figures 1 and 2. These figures report, for men and women respectively, the adjusted predicted mean log of the real hourly wage by years of experience (or more precisely, the deviation from the mean of experience) for persons who only experienced casual employment in the first three years of the panel compared with those who only experienced regular employment in these same three years.

Focusing first on Figure 1, and consistent with what was reported in Table 1, men with a history of casual employment not only have significantly lower levels of hourly wages at almost all levels of observed experience, the rate of catch-up is relatively weak. Indeed, the relevant interaction term is not statistically significant ($\beta = .007$, s.e. = .005). For women, the situation is very different. As shown in Figure 2, the hourly wages of women with a history of casual employment converge quite rapidly on that of their counterparts with a history of regular employment. But it should be emphasised that among women the average wage penalty attached to casual employment is relatively small. That, combined with the imprecision of our estimates, means that at most levels of experience the wage differential between our two groups of women is not statistically significant.

While informative, the estimations underlying Figures 1 and 2 are still very restrictive in that they only allow patterns of wages growth to vary with gender. We also hypothesised that effects would vary across different age groups as people enter different life stages. We thus also tested for age-specific differences in wage progression by re-estimating our hybrid FE model after including interaction terms between the initial employment status variables, within-person deviations in potential work experience and age. We summarise these estimations in Figures 3 (men) and 4 (women).

Focusing first on men, we find that the trajectories in predicted mean wages vary markedly with age. For the youngest workers (those that enter our observation window aged between 25 and 34) wages rise, albeit at a declining rate, whereas for the oldest cohort (those aged 45 to 54 years) wages are falling. More importantly, we can see that the wage gap between those previously employed in a casual job and those previously employed in a regular job changes very little with experience for older men. In contrast, among the group of younger workers (aged 25 to 34) this wage gap clearly narrows with experience. Indeed, by the end of our 10-year estimation period those with a prior history of continuous casual employment earn, on average, virtually the same as those with a prior history of continuous regular employment.

For women, for all three age groups, no significant wage differences are found between those in casual and regular employment at different levels of work experience over the 10-year estimation period. That said, Figure 4 does suggest that for the younger cohorts of women (those aged 25 to 44) there are modest initial wage penalties associated with a history of casual employment that shrink with time. Indeed, by the end of the 10-year estimation period, among the youngest cohort, women with a history of casual employment actually earn slightly more than comparable women with a history of regular employment.

Robustness

We subjected our analyses to a number of robustness checks, results of which are reported in Table 4. First we checked whether our results were potentially affected by omitted variables bias by including the hourly wage in the first job observed during the panel. The idea here is that if our hybrid FE model is adequately capturing unobserved heterogeneity then including a variable that is likely correlated with important unobservables (such as ability) should not substantially affect our results. This expectation proved largely correct. While the first wage was strongly predictive of w_{it}, the estimated coefficients on the labour market history variables were little affected. Notably, and as reported in column (1a) of Table 4, for men the coefficient on the indicator of continuous casual employment barely changed at all. For women (Table 4, column 2a), however, there is more a noticeable change – the negative coefficient on this same variable changes from -.044 to -.029. While this change is still quite small, the coefficient is no longer significant. This suggests the possibility that the very modest casual wage penalty estimated for women in Table 3 may be overstated, and that the true effect may not be significantly different from zero.

Second, and very differently, it could be argued that some of the effects of casual employment may be reflected in other coefficients in our model. In particular, if casual employment is used as a device to reinforce labour market segmentation, then some of the negative effect on future earnings is likely captured by the occupation dummies. Nevertheless, exclusion of these occupation dummies had very little effect on the other coefficients (Table 4, columns 1b and 2b).

Third, it is possible that our estimates are biased due to endogenous response or selection. To deal with this, we repeated our estimation after including additional variables identifying whether the respondent was observed at the next survey wave ($t+1$), and if they were, whether they were either unemployed or not in the labour force, and hence excluded from our estimation sample. With one exception, the coefficients on these variables were insignificant in both the male or female equations, suggesting that non-random response and selection has no noticeable effect on our results. Among males, however, the control for non-response attracted a positive coefficient that was statistically significant, suggesting non-respondents are higher wage earners on average than those that remain. Nevertheless, inclusion of this variable (along with the other two controls for selection) had only a modest effect on the coefficients on most of the labour market history variables (Table 4, columns 1c and 2c). Most importantly, the estimated wage differential attached to continuous casual employment was little affected.

Finally, it might be argued that three data points (spread over a two-year window) is too few to derive a reliable measure of sustained employment patterns. We thus re-estimated our models over 8 years (2006 to 2013), rather than 10, with the indicators of employment status recalculated using five data points (2001 to 2005) rather than three. Again we found little effects for men (Table 4, column 1d). For women, however, the estimated wage penalty attached to continuous casual employment declines to close to zero (Table 4, column 2d).

In summary, the estimated wage differentials among males are very robust to variations in specification. All of the different specifications tested produced estimates of an average casual employment wage penalty lying in the range of 10 to 12%. In contrast, there is much less certainty attached to the equivalent estimate for women. Our initial estimate of this penalty is much smaller for women (just 4.5%) and in at least two alternative specifications we found estimates that were not significantly different from zero.

Conclusions

At any point in time a relatively large proportion of the Australian workforce is engaged in casual employment. Such jobs are inherently more insecure than other forms of employment and provide fewer non-wage entitlements. While in Australia such differences are at least partly offset by industrial regulations that require the payment of a substantial wage premium, the concern remains that for many workers casual employment is a labour market state that is difficult to escape. While the extent to which casual employment in Australia is or is not a permanent state has been the subject of research (cf. Buddelmeyer and Wooden, 2011; Watson 2013), no previous Australian study has attempted to quantify the possible consequences of casual employment for future earnings. This was the objective of this study. More specifically, we examined how employment status, measured over multiple points in time, influences hourly wage outcomes much later in a worker’s career. We also tested whether patterns of wage differentials (between workers with a history of casual and noncasual employment) evolve over time and the extent to which these patterns over time vary with age cohort and gender.

The results reported here suggest two major conclusions. First, casual employment has a much larger detrimental impact on the long-run earnings prospects of men than of women. Indeed, among women the average wage penalty associated with a history of continuous employment is relatively small and not robust. That casual employment is generally found to be associated with lower future wages is in line with theories predicting scarring effects from non-standard employment. The marked gender difference, on the other hand, is consistent with arguments that emphasise the importance of gender-based normative expectations about what constitute "ideal careers" for men and women.

Second, the magnitudes of these wage differentials have a tendency to shrink with increased work experience, but patterns of wage differentials are not universal and vary with age and gender. In particular, wage differentials are relatively persistent for workers that experience casual employment later in their working life. In contrast, for younger workers (under 35 years of age) these wage gaps shrink with experience, and by the end of our 10-year observation period have largely disappeared. Within a life-course perspective, these results confirm that the size and persistence of wage penalties stemming from temporary employment are strongly moderated by the timing in which this type of employment occurs.

It should, however, be noted that in presenting our results we have placed emphasis on what is arguably an atypical group of casual employees; those observed in a casual job at each of our first three annual interview dates. As previously noted, most persons observed in casual employment do not fall into this group. Almost 34% of the men in our sample, and almost 49% of the women, were observed in a casual job at the time of one the first three survey dates, but only around one in five of this group (19% of men and 23% of women) were in a casual job at all three. This is of large importance given the wage effects we estimate are mostly smaller for other employment patterns involving casual employment.

Finally, we emphasise the importance of heterogeneity and, in particular, non-random selection into employment states. We found very large differences in the magnitude of the estimated wage penalties in simple random effects models compared with those from our preferred hybrid FE model; differences which we believe reflect the superiority of the hybrid model in dealing with unobserved heterogeneity. Further, a simple added-variable test suggests our findings are robust. Nevertheless, it might still be argued that some of the estimated differences reflect systematic differences between persons that have not been entirely eliminated, and as such our estimated wage differentials might be overstated.

Notes

1 From July 2014 this premium became a standard 25% in all awards. Prior to that date there was considerable variation across awards, but with 20% widely recognised as the norm (Watson, 2005).
2 In the sample used here, for example, 27.9% of women were employed on a casual basis at the time of interview, which compares with 17.5% of men (see Table 2).
3 Booth et al. (2002) also included both the deviations from within-person means in time-varying control variables and the within-person means of these variables in their specification, but in contrast to our approach, treated them as instruments. Booth et al. claim that these are valid instruments given they are, by construction, uncorrelated with the error term, but say nothing about the power of the instruments. We suspect they are very weak instruments for employment contract status.
4 The reported value of usual gross weekly and wages and salary is an estimate in the sense that: (i) respondents are permitted to report gross wages over any period (e.g., week, fortnight, month, or year) and hence weekly estimates have to be derived for some respondents; (ii) some respondents are only able to report take-home pay and hence gross wages for these individuals have to be estimated using income tax scales; and (iii) not all employed respondents provide an answer each year (about 3% do not), and in these cases their wages have been imputed.
5 This classification is complicated by the fact that not all persons respond at every survey wave. We thus could also distinguish between persons with complete information (balanced panel respondents) and incomplete information (the unbalanced panel). This distinction, however, proved not to be very important, and hence persons with incomplete information are assigned into categories based solely on the years for which information is available.
6 Respondents only needed to be observed in at least one of the first three survey waves to be classified.
7 The estimated casual employment wage penalty is derived from $\exp(-\beta) - 1$, where β is the coefficient on the casual employment only variable.
8 While we use the term ‘continuous’, the data does not actually enable the duration of spells of employment by contract type to be identified. Use of the term ‘continuous’ is thus just a convenient short-hand to describe workers who are only observed in the first three survey waves in one employment or labour market state.

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Table 1. Previous studies of long-term wage effects of non-standard employment

Study	Data / Sample	'Treatment' variable	Observation period	Method	Key findings on 'treatment' group
Ferber and Waldfogel (1998)	NLSY, 1979-1993 (USA); persons in employment at end of panel (N = 2870). Sample aged 28 to 36 by 1993.	Ever worked in a temporary job.	Up to 14 years	Ordinary least squares.	Hourly wages at end of period are approx. 7% lower among males who had ever held a temporary job, and 4% lower for females. Models of 5-yearly wage growth, however, reveal no differences.
Booth et al. (2002)	BHPS, 1991-1997 (Britain); employees aged ≤60 who had left school (N = 3721).	Number of seasonal-casual jobs, and number of fixed-term contract jobs.	6 years	Random effects with instrumental variables.	Wages decline with number of casual / seasonal jobs and with number of fixed-term contract jobs. Negative effects are larger for casual jobs and for men.
Gash and McGinnity (2007)	ECHP, 1994-2001 (France and Germany); employees with less than 2 years job tenure (N = 421 / 692).	Employment status (fixed-term contract vs permanent employment).	4 years	Propensity score matching.	Germany: Higher wages for women, but only significant at *t+2*. France: Some evidence of a wage penalty for men at *t+4*.
Andersson et al. (2009)	Sub-sample of Longitudinal Employer Household Dynamics program, 1993-2001 (USA); prime-private non-farm wage and salary workers in five US states aged 25 to 55 in 1993 who were persistently low wage earners (N = 1384).	Temporary agency employment during three-year period, differentiated by whether primary employment or not.	6 years	Fixed effects.	Temporary agency employment is associated with lower subsequent earnings (3-6% less six years later). For persons who escape temporary agency work a substantial premium (11 to 13%) exists.
Gebel (2010)	BHPS, 1991-2007 (UK) / GSOEP, 1991-2007 (Germany); persons aged <33 years who had left education and transitioned into a permanent or temporary contract (N = 1669 / 1338).	Initial employment status (fixed-term contract vs permanent employment).	5 years	Propensity score matching.	Germany: Large initial wage penalty that diminishes with time (and become insignificant for women). UK: Insignificant differences for men. For women, the penalty diminishes over time.

Table 1 (cont'd)

Study	Data / Sample	'Treatment' variable	Observation period	Method	Key findings on 'treatment' group
Gebel (2013)	BHPS, 1991-2009 (UK) / GSOEP, 1991-2009 (West Germany) / SHP 1999-2009 (Switzerland); unemployed persons aged 15-54.	Temporary work vs remaining unemployed for at least one more month.	5 years	Propensity score matching.	Wage premiums that vary from 3-5% in Switzerland to 8-11% in Germany. The premium does not decline with time.
Pavlopoulos (2013)	BHPS, 1991-2007 (UK) / GSOEP, 1984-2008 (Germany); persons aged 16-30 years who entered the labour market for the first time during the panel in either a permanent or fixed-term contract job (N = 5674 / 5023).	Initial employment status (fixed-term contract vs permanent employment).	Unclear	Random effects, with sample selection correction.	Germany: Small but significant wage penalties that decline with experience. UK: Insignificant differences for men. For women, a large initial penalty declines with experience.
De Lange et al. (2014)	OSA Labour Supply Panel, 1988-2008 (Netherlands); persons employed at first interview after leaving full-time education (N = 973).	Initial employment status (standard employment vs flexible employment).	8 years	Random effects.	7-8% lower wages, but statistically insignificant.
Fuller and Stecy-Hildebrandt (2014)	SLID, 1999-2004 (Canada); persons aged 18 to 60 who were in temporary or permanent employment at wave 2 (of each 6 wave panel) (N = 3694).	Initial employment status (temporary vs permanent employment).	5 years	Propensity score matching / Growth curve model (random effects).	Large initial wage penalties (~14-17% for men and 23-25% for women) that decline over time. Extent of catch-up dependent on specification.
Fernández-Kranz et al. (2015)	Spanish social security records, 1996-2006; female employees born between 1961 and 1973 with at least 3 years in full-time employment.	Years of fixed-term contract employment.	Up to 10 years	Random effects; simultaneous equations model.	Small wage premium for full-time fixed-term contract employment that changes very little with time in that state.

Notes: BHPS = British Household Panel Survey; ECHP = European Community Household Panel; GSOEP = German Socio-Economic Panel; NLSY = National Longitudinal Study of Youth; SHP = Swiss Household Panel; SLID = Survey of Labour and Income Dynamics; TANF = Temporary Assistance for Needy Families. Ns refer to the number of unique individuals (and not person-wave observations).

Table 2. Summary descriptive statistics

	Men		Women	
	Mean	Std. dev.	Mean	Std. dev.
Log of real hourly wage	3.094	0.491	2.993	0.433
Initial labour force / employment status				
Regular employment only	0.481	0.500	0.366	0.482
Casual employment only	0.065	0.246	0.110	0.313
Self-employed (SE) only	0.107	0.309	0.043	0.203
Casual + Regular	0.093	0.290	0.122	0.328
Casual + SE or Casual + Regular + SE	0.112	0.316	0.084	0.278
Casual + Regular + (Unemployed or NILF)	0.061	0.240	0.157	0.364
Casual + (Unemployed or NILF)	0.004	0.066	0.010	0.101
Other patterns involving Casual employment	0.003	0.050	0.003	0.057
Regular + Unemployed	0.025	0.155	0.016	0.127
NILF or unemployed (but never employed)	0.025	0.155	0.060	0.237
Other patterns	0.024	0.154	0.027	0.163
Casual employment at time of interview	0.175	0.380	0.279	0.449
Occupation (ANZSCO)				
Managers	0.176	0.381	0.100	0.300
Professionals	0.210	0.407	0.299	0.458
Technicians and trades workers	0.227	0.419	0.040	0.197
Community and personal service workers	0.061	0.240	0.146	0.354
Clerical and administrative workers	0.073	0.260	0.240	0.427
Sales workers	0.049	0.215	0.097	0.296
Machinery operators and drivers	0.104	0.305	0.012	0.109
Labourers	0.100	0.300	0.064	0.245
Industry				
Agriculture, forestry and fishing	0.041	0.198	0.019	0.135
Mining	0.030	0.170	0.006	0.077
Manufacturing	0.134	0.341	0.049	0.216
Electricity, gas, water and waste	0.015	0.122	0.004	0.065
Construction	0.134	0.340	0.020	0.139
Wholesale trade	0.042	0.201	0.023	0.149
Retail trade	0.072	0.258	0.104	0.306
Accommodation and food services	0.038	0.191	0.058	0.235
Transport, postal and warehousing	0.066	0.248	0.020	0.140
Information media and telecommunication	0.027	0.162	0.023	0.150
Financial and insurance services	0.036	0.187	0.046	0.210
Rental, hiring and real estate	0.012	0.110	0.015	0.121
Professional, scientific and technological services	0.079	0.270	0.078	0.268
Administrative and support services	0.025	0.156	0.036	0.186
Public administration and safety	0.086	0.280	0.063	0.243
Education and training	0.051	0.221	0.155	0.362
Health care and social assistance	0.047	0.211	0.225	0.418
Arts and recreation services	0.019	0.136	0.016	0.124
Other services	0.040	0.196	0.034	0.181
Employer type				
Private sector for profit organization	0.776	0.417	0.603	0.489
Government business enterprise	0.048	0.213	0.045	0.207
Other commercial	0.002	0.046	0.003	0.052
Private sector not for profit	0.036	0.186	0.096	0.295
Other governmental organization	0.135	0.342	0.246	0.431
Other non-commercial organization	0.002	0.049	0.005	0.073
Full-time work	0.898	0.303	0.542	0.498
Union membership	0.256	0.437	0.264	0.441

Table 2 (cont'd)

	Men		Women	
	Mean	Std. dev.	Mean	Std. dev.
Tenure in the same occupation	9.675	9.237	7.972	8.248
Tenure in the same occupation squared	178.9	282.7	131.6	233.1
Tenure in the same firm	7.167	7.819	5.923	6.584
Tenure in the same firm squared	112.5	217.4	78.42	164.7
Potential experience	20.40	10.27	21.03	10.33
Potential experience squared	521.7	420.3	548.9	425.4
Education				
Year 11 or less	0.190	0.393	0.210	0.407
Year 12	0.163	0.369	0.174	0.379
Certificate III or IV	0.299	0.458	0.168	0.374
Advanced diploma	0.092	0.289	0.108	0.311
Bachelor or honors	0.153	0.360	0.203	0.403
Graduate diploma, graduate certificate	0.051	0.221	0.087	0.282
Postgraduate	0.051	0.220	0.049	0.215
Undetermined or missing	0.000	0.012	0.000	0.015
Country of birth				
Australia	0.816	0.388	0.816	0.387
Overseas: main English-speaking country	0.087	0.282	0.080	0.271
Overseas: other country	0.097	0.296	0.104	0.305
Remoteness				
Major city	0.624	0.484	0.632	0.482
Inner regional	0.222	0.416	0.219	0.413
Outer regional and remote areas	0.154	0.361	0.150	0.357
Work-limiting long-term health condition or disability	0.077	0.267	0.086	0.280
Family / relationship status				
Single	0.200	0.400	0.132	0.338
Lone parent w/o dependents	0.004	0.063	0.016	0.124
Lone parent	0.015	0.122	0.100	0.300
Couple without dependents	0.249	0.432	0.282	0.450
Couple with dependents	0.532	0.499	0.470	0.499
Children				
# of children < 12 years	0.642	0.971	0.574	0.894
# of children aged 12-17 years	0.342	0.686	0.392	0.709
Survey year				
2004	0.118	0.323	0.114	0.317
2005	0.118	0.322	0.117	0.322
2006	0.114	0.318	0.114	0.318
2007	0.107	0.309	0.108	0.310
2008	0.103	0.304	0.104	0.305
2009	0.098	0.298	0.099	0.298
2010	0.093	0.290	0.094	0.292
2011	0.089	0.284	0.089	0.285
2012	0.083	0.275	0.083	0.276
2013	0.077	0.267	0.078	0.268

N = 48,534 (Men = 24,580; Women = 23,954)
SE = Self-employed; NILF = Not in the labour force.

Table 3. The effect of initial employment status on log of hourly wages, by gender, 2004-2013: Estimates from random effects (RE) and hybrid fixed effects (FE) models

	Men		*Women*	
	RE	*Hybrid FE*	*RE*	*Hybrid FE*
Regular employment only	Reference group			
Casual employment only	-0.216***	-0.101***	-0.131***	-0.044*
	(0.022)	(0.024)	(0.017)	(0.019)
Self-employment only	-0.336***	-0.365***	-0.156**	-0.166***
	(0.033)	(0.034)	(0.049)	(0.049)
Casual + Regular	-0.098***	-0.053**	-0.095***	-0.045**
	(0.020)	(0.019)	(0.015)	(0.015)
Casual + SE or Casual + Regular + SE	-0.117***	-0.087***	-0.071***	-0.033
	(0.021)	(0.021)	(0.021)	(0.020)
Casual + Regular + (Unemployed or NILF)	-0.213***	-0.062*	-0.118***	-0.015
	(0.026)	(0.027)	(0.017)	(0.019)
Casual + (Unemployed or NILF)	-0.151*	0.045	-0.155***	-0.031
	(0.076)	(0.071)	(0.037)	(0.038)
Other patterns involving casual employment	-0.065	0.007	-0.122	-0.057
	(0.121)	(0.104)	(0.096)	(0.091)
Regular + Unemployed	-0.097**	-0.009	-0.137***	-0.061
	(0.032)	(0.031)	(0.040)	(0.038)
NILF or unemployed (but never employed)	-0.236***	-0.035	-0.242***	-0.078**
	(0.043)	(0.048)	(0.024)	(0.029)
Other patterns	-0.235***	-0.114**	-0.145***	-0.052
	(0.040)	(0.038)	(0.032)	(0.029)
R-squared (within-person)	0.109	0.119	0.088	0.094
R-squared (between-person)	0.339	0.428	0.333	0.400
R-squared (overall)	0.315	0.399	0.302	0.353
# of wage observations	18895	18895	18778	18778
# of workers	3488	3488	3655	3655
Sigma μ	0.322	0.315	0.276	0.268
Sigma ε	0.249	0.249	0.263	0.263
Rho	0.625	0.615	0.524	0.508

Notes:

1. The dependent variable in all models is the log of real hourly wages.
2. All models include controls for: the type of employment contract at time of interview; occupation; employer type (e.g., private vs public); industry; full-time work hours; job tenure and job tenure squared; occupation experience and occupation experienced squared; potential experience and experience squared; union membership; education level; country / region of birth; remoteness; family / relationship status; number of children aged <12 years, and between 12 and 17 years; presence of a work-limiting long-term health condition; and survey year. In the hybrid models, all time-varying variables are specified as deviations from the within-person mean, which are included alongside the person-specific mean.
3. *** p<0.001; ** p<0.01; * p<0.05.

Table 4. Sensitivity checks – Coefficients estimates from alternative specifications (hybrid FE models)

	Men				*Women*			
	Plus Initial wage	*Minus occupation*	*Plus controls for NR*	*5-year observation window*	*Plus Initial wage*	*Minus occupation*	*Plus controls for NR*	*5-year observation window*
	(1a)	*(1b)*	*(1c)*	*(1d)*	*(2a)*	*(2b)*	*(2c)*	*(2d)*
Casual employment only	-0.099***	-0.111***	-0.100***	-0.098**	-0.029	-0.056**	-0.046*	-0.009
	(0.025)	(0.025)	(0.024)	(0.031)	(0.018)	(0.020)	(0.019)	(0.021)
Self-employment only	-0.250***	-0.343***	-0.381***	-0.340***	-0.122*	-0.157**	-0.171**	-0.176**
	(0.035)	(0.034)	(0.036)	(0.037)	(0.052)	(0.050)	(0.052)	(0.054)
Casual + Regular	-0.045*	-0.065***	-0.054**	-0.060**	-0.041**	-0.056***	-0.048**	-0.030
	(0.018)	(0.020)	(0.020)	(0.020)	(0.014)	(0.015)	(0.015)	(0.017)
Casual + SE or Casual + Regular + SE	-0.061**	-0.083***	-0.093***	-0.077**	-0.027	-0.032	-0.045*	-0.008
	(0.020)	(0.021)	(0.022)	(0.023)	(0.019)	(0.021)	(0.021)	(0.022)
Casual + Regular + (Unemployed or NILF)	-0.036	-0.053	-0.049	-0.066**	-0.027	-0.007	-0.015	0.010
	(0.028)	(0.028)	(0.027)	(0.026)	(0.020)	(0.020)	(0.020)	(0.017)
Casual + (Unemployed or NILF)	0.085	0.072	0.028	-0.052	-0.032	-0.033	-0.041	-0.020
	(0.071)	(0.071)	(0.071)	(0.055)	(0.040)	(0.038)	(0.039)	(0.030)
Other patterns involving casual employment	-0.001	-0.033	-0.039	-0.155**	-0.054	-0.059	-0.073	-0.033
	(0.099)	(0.112)	(0.113)	(0.049)	(0.072)	(0.093)	(0.092)	(0.040)
Regular + Unemployed	-0.020	0.012	-0.045	-0.027	-0.032	-0.047	-0.066	-0.098
	(0.028)	(0.034)	(0.031)	(0.031)	(0.041)	(0.039)	(0.037)	(0.051)
NILF or unemployed (but never employed)		0.009	-0.047	-0.014		-0.056	-0.074*	-0.048
		(0.050)	(0.051)	(0.052)		(0.030)	(0.031)	(0.033)
Other patterns	-0.080*	-0.102**	-0.129***	-0.036	-0.059*	-0.055	-0.043	-0.031
	(0.037)	(0.037)	(0.039)	(0.022)	(0.030)	(0.031)	(0.030)	(0.018)
Initial wage	0.338***				0.221***			
	(0.018)				(0.019)			
Non-respondent at *t+1*			0.032**				0.006	
			(0.010)				(0.010)	
Unemployed at *t+1*			0.012				0.010	
			(0.024)				(0.023)	
Not in labour force at *t+1*			0.002				-0.001	
			(0.025)				(0.013)	

Notes:

1. The dependent variable in all models is the log of real hourly wages.
2. Most models include controls for: the type of employment contract at time of interview; occupation; employer type (e.g., private vs public); industry; full-time work hours; job tenure and job tenure squared; occupation experience and occupation experienced squared; potential experience and experience squared; union membership; education level; country / region of birth; remoteness; family / relationship status; number of children aged <12 years, and between 12 and 17 years; presence of a work-limiting long-term health condition; and survey year. In the hybrid models, all time-varying variables are specified as deviations from the within-person mean, which are included alongside the person-specific mean.
3. *** p<0.001; ** p<0.01; * p<0.05.

Figure 1. Adjusted predicted log wages by experience and initial employment status (with 95% CIs): Men

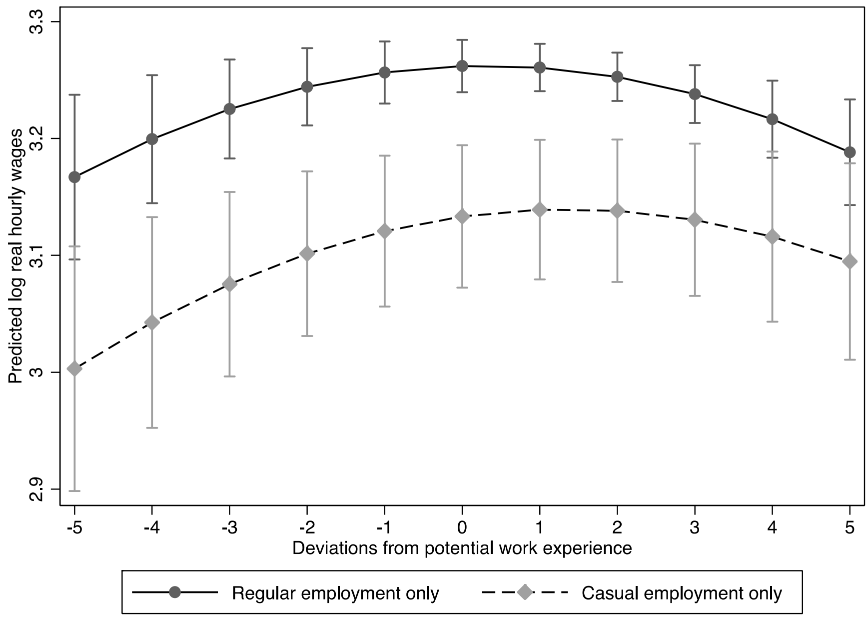

Figure 2. Adjusted predicted log wages by experience and initial employment status (with 95% CIs): Women

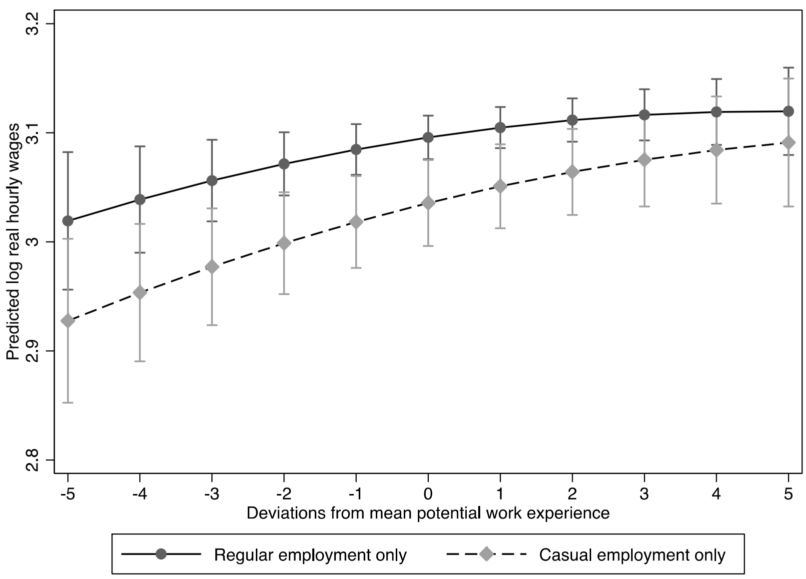

Figure 3. Adjusted predicted log wages by experience, initial employment status and age: Men

Figure 4. Adjusted predicted log wages by experience, initial employment status and age: Women