# Non-standard Employment and Wages in Australia

*Inga Laß and Mark Wooden*
*Melbourne Institute of Applied Economic and Social Research,*
*University of Melbourne*

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## Introduction

Recent decades have seen increased concern, both in Australia and elsewhere, about changes in the nature of work, and more specifically growth in non-standard forms of employment (such as part-time and casual work), and what this implies for the quality of jobs (e.g., ACTU, 2018; Buddelmeyer et al., 2015; Green et al., 2010; Kalleberg et al., 2000; McGovern et al., 2004; Richardson & Law, 2009; Watson, 2005). In particular, non-standard forms of employment are typically seen as 'precarious', with adverse consequences for workers flowing from greater economic insecurity (e.g., Campbell & Burgess, 2018; Markey & McIvor, 2018; Rogers, 1989; Tweedie, 2013; Vosko, 2006). Such concerns have received added impetus in recent years from fears about the impact of new digital technology and its potential to facilitate greater disintermediation of work tasks, and hence the rise in what is often referred to as 'gig work' (Healey et al., 2017).

Changes in the mix of 'standard' and 'non-standard' jobs may also have implications for wages. Andrew Haldane (2017), Chief Economist at the Bank of England, for example, identifies the changing nature of work, and especially growth in self-employment, temporary employment and zero-hours contracts, as a factor contributing to weak wages growth in the UK. Similarly, in Australia, Cassidy and Parsons (2017) point to the potential role that growth in the part-time employment share may have played in restraining real wages growth. More specifically, they point to both the concentration of part-time jobs in low paying occupations and industries and relatively low levels of bargaining power among part-time workers as factors that could drive down observed wages growth. In a similar vein, Pickering (2018) points to the rise not just in part-time jobs, but in jobs where workers have no paid leave entitlements (i.e., casual employment), as a likely contributing factor to relatively low real wages growth in recent years. That said, the wage floors created by minimum award wages may have shielded many of the lower paying non-standard jobs from any erosion in wages growth. Prima facie evidence for this is provided by growth in minimum award wages in recent years (15.6% in the five years to 2018) that has been considerably above the growth in consumer prices over that same period (just 9.1%).

Despite these arguments, we are unaware of any previous research that has examine the link between growth in non-standard employment and wages growth in Australia within a multivariate framework. Previous research has focused on associations with the wage level (e.g., Booth and Wood 2008; Green et al. 2010; Laß and Wooden, forthcoming; Watson 2005), but not with the rate of growth in wages. The aim of this paper is therefore to re-assess the strength and validity of arguments linking changes in the prevalence of employment types to wages growth in Australia. We begin by first defining what types of employment are covered by the term non-standard employment. We then use data from both the Australian Bureau of Statistics (ABS) Labour Force Survey and the Household, Income and Labour Dynamics in Australia (HILDA) Survey to describe the prevalence of, and trends in, the incidence of non-standard employment broken down into its major components over the period 2001 to 2017. Finally, we use HILDA Survey data to quantify associations between employment type and both the level of, and annual rate of growth in, real hourly wages.

We find that the share of non-standard employment in total employment has, perhaps surprisingly, not increased much since the turn of the millennium. However, a major factor working against growth in non-standard employment has been self-employment – the self-employment rate in Australia has been in long-term decline. If we restrict attention to employees, which seems appropriate for an analysis of wages, then we do observe an increase over time in the non-standard employment share. Further, all of this increase is concentrated in the years since the Global Financial Crisis (GFC) – over the period 2001 to 2008 the share of employees with non-standard employment contracts actually declined.

Such trends are suggestive of an association with low wages growth, with most indicators showing real wages growth declining to, and persisting at, quite low levels in recent years. Nevertheless, while it is true that some types of non-standard employment (notably casual work) are associated with relatively low wages, once we control for worker characteristics these differences disappear: Indeed, if anything, permanent part-time, casual and fixed-term contract workers earn hourly wage premiums. Such findings are in line with previous research using the HILDA Survey data (e.g., Booth & Wood, 2008). However, we also find clear evidence that employees in non-standard forms of employment have, throughout the period covered by this study (2001 to 2017), experienced relatively low rates of growth in hourly wages when compared with permanent full-time employees. Growth in the share of non-standard types of employment in total employee employment thus appears to be one factor contributing to slower rates of real wages growth in recent years. Nevertheless, a simple decomposition analysis suggests that the magnitude of this effect is small (and insignificantly different from zero).

## Defining and Measuring Non-standard Employment in Australia

Non-standard employment, or what has been referred to elsewhere as 'atypical work' (e.g., Addison & Surfield, 2009; Córdova, 1986; de Grip et al., 1997), 'alternative work arrangements' (e.g., Farber, 1999; Katz & Kreuger, 2016; Polivka, 1996) or 'flexible employment / contracts' (e.g., Green & Heywood, 2011; Guest, 2004; Houseman & Polivka, 2000), has most commonly been defined as any job that differs from full-time, permanent, dependent (i.e., wage and salary) employment (e.g., OECD, 2015: 138). This covers a broad and disparate array of employment arrangements, including self-employment, part-time work, and any job where there is no commitment on the part of the employer to a long-term relationship (e.g., fixed-term contracts, casual employment, and temporary agency work).

In Australia, as in most other advanced industrial countries, the identification and measurement of employment begins with the labour force framework developed by the International Labour Organization (ILO). In this framework, paid employment is based on the economic activity undertaken by individuals during a one-week reference period, with the basic prerequisite being just one hour of paid work. It is then conventional to distinguish between different types of job holders based on the relationship between the worker and the enterprise they work for; that is, between employees and the self-employed. One problematic subgroup within this framework is owner managers of incorporated businesses, who the Australian Bureau of Statistics (ABS) treats as employees of their own business. But while the legal status of a business has implications for who is held responsible in the event of insolvency, it has no bearing on the employment relationship – the owner of a firm is fundamentally different to other persons employed in that firm, not least because of the power the owner has over hiring and firing decisions and the allocation of tasks among workers. In this analysis, therefore, we treat all owner managers, regardless of the legal status of their businesses, as self-employed.[1]

There is a third category of employed persons – contributing family workers – who do not fit neatly into either the employee or self-employed groups. Workers in this category clearly fit the definition of non-standard employment but are relatively few.

Within the employee group we next categorise workers into different groups according to the nature of their employment contract. In Australia these take three main forms: (i) fixed-term contracts; (ii) casual employment; and (iii) permanent or ongoing contracts.

Fixed-term contracts cover all employment contracts that specify a specific date or event when employment will be terminated. In Australia, fixed-term contracts generally come with the same entitlements as permanent contracts (e.g., with respect to paid leave and paid holidays). Further, fixed-term contract workers have a general expectation of being employed at least for the duration of their current contract.

Less straightforward is the identification of casual employment. While a dictionary definition suggests that casual employees are hired for very short periods, with each engagement of work constituting a separate contract of employment (and indeed this is the definition most consistent with common law; Brooks, 1985), the reality is that many casual employees work regular hours for the same employer over long periods. Ultimately, the key defining feature of casual employment is the absence of any advance commitment on the part of the employer to both the continuity of employment and the number of days or hours to be worked (Creighton & Stewart, 2010: 198). It might therefore seem difficult to distinguish between casual employees and permanent employees given both essentially have open-ended employment contracts and only rarely are employment contracts truly permanent. Casual employment in Australia, however, is the subject of extensive regulation. It is specifically provided for in industry awards, with casual employees singled out as not having any legal entitlement to many forms of paid leave (notably annual leave and sick leave), paid public holidays, minimum periods of notice of termination, or severance pay. At the same time, a long-standing feature of award regulation is the requirement of the payment of a substantial hourly wage premium to casual workers, which helps to at least partly compensate for the loss of other benefits. Most casual employees in Australia should, therefore, be able to recognise that they are employed on a casual basis. In this analysis we thus mainly rely on data using a self-classification method to identify casual employees.

An alternative, and longer-standing, approach is to infer casual employment status from survey data on the receipt of paid annual and sick leave entitlements, with employees reporting receiving neither paid annual leave nor paid sick leave entitlements being classified as casual workers. We thus also report on data that show how sensitive estimates of the level of casual employment are to these definitional differences.

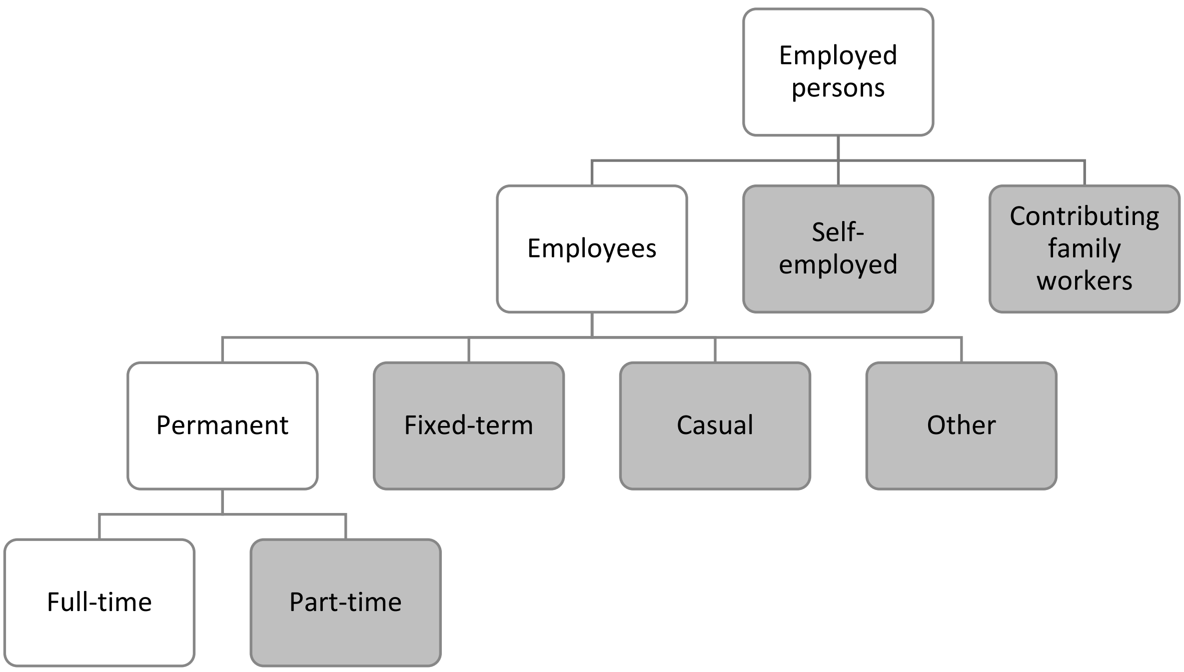
The final dimension we use to identify non-standard employees is usual hours of work. That is, among employees working on a permanent or ongoing basis we distinguish between full-time and part-time workers. There is, however, no internationally accepted definition of part-time work. Definitions used by official national statistics agencies, for example, vary in

[1] This is consistent with ILO guidelines, which, as revised in 2013, specifically identify owner managers of incorporated enterprises as a group that may be classified either as employees or as self-employed, but then suggested that classification as self-employed will generally be best for labour market analyses (ILO, 2013: 18).

terms of: (i) whether a weekly hours threshold is used or whether part-time is self-defined; (ii) if a weekly hours threshold is used, which threshold is chosen (usually either 30 or 35 hours); (iii) whether usual hours or actual hours worked in a specific week are used (or a combination of both); and (iv) whether the hours cover all jobs or just the main job (van Bastelaer et al., 1997). In Australia, the ABS in its monthly Labour Force Survey (LFS) defines part-time workers as employed persons "who usually worked less than 35 hours a week (in all jobs) and either did so during the reference week, or were not at work in the reference week". The reference to hours worked in all jobs, however, is problematic for identifying workers in standard and non-standard employment. Most obviously, there are persons holding multiple part-time jobs who will be classified as full-time workers, and hence as being in standard employment, when none of the jobs they hold meet the full-time criterion. And indeed, in supplements to the LFS (notably the Characteristics of Employment [CoE] Survey now conducted every August), the ABS distinguishes between part-time workers in all jobs and part-time workers in the main job. We thus define part-time jobs based on the number of hours worked in the main job and, in line with ABS practice, use a 35 weekly hours threshold to identify full-time work.

The classification system described above is summarised in Figure 1. Workers in standard employment are those in the box in the bottom left corner of this diagram – permanent full-time employees. All other employees and workers (those in the shaded boxes) have non-standard jobs.

***Figure 1: Classification of employment types***




Employees could also be classified according to whether or not they are employed through an intermediary or agency. Such workers are generally thought of as having a non-standard employment relationship since the employing organisation is not the organisation for which the labour services are performed. However, given such positions are most often the result of a short-term need to meet some temporary peak in demand or to cover absences of regular employees, most agency workers will be employed on a fixed-term or casual basis and so already covered by the typology in Figure 1. Nevertheless, there will be a minority of agency workers (e.g., those with highly valuable but specialised skills) who are employed on an ongoing contract with the agency who arguably should be classified as being in non-standard jobs due to the peculiar, tripartite nature of the employer-employee relationship. In this analysis, however, we ignore this distinction and so treat these agency workers on permanent contracts as being in standard employment.[2]

## The Prevalence of Non-standard Employment in Australia: ABS and HILDA Survey Estimates Compared

Table 1 presents cross-sectional population estimates of the prevalence of non-standard employment in Australia in 2016 where employment status is as reported by workers.[3] It reports data from two sources: (i) the ABS CoE Survey, which, as already noted, is conducted as a supplement to the August round of the monthly LFS; and (ii) the Household, Income and Labour Dynamics in Australia (HILDA) Survey, a household panel survey that commenced in 2001, and which has been extensively used by researchers to study the characteristics, correlates and consequences of different forms of non-standard employment (e.g., Buddelmeyer et al., 2015; Buddelmeyer & Wooden, 2011; Booth & Wood, 2008; Green et al., 2010; Laß & Wooden, forthcoming; McVicar et al., 2019; Mooi-Reci & Wooden, 2017; Productivity Commission, 2006; Richardson et al., 2012; Watson, 2005, 2013).

According to both sources, less than half of the employed Australian workforce was in a standard employment relationship (i.e., had a permanent full-time job) in 2016 – 48.6% when using the CoE Survey and 45.1% when using the HILDA Survey. And both surveys indicate that the most prevalent form of non-standard employment is casual employment, accounting for almost one in every five workers (19.5% or 19.6%) when using a self-classified measure. Further, using the alternative proxy measure of casual employment based on the absence of leave entitlements also produces two similar estimates – about 25% of all employees are estimated to be without paid leave entitlements when using the CoE Survey and 26% when using the HILDA Survey. These estimates are equivalent to about 21% and 22% of all employed persons respectively (so around 1.5 to 2.5 percentage points higher than the rates indicated by the self-classification methods). The two data sources also produce very similar estimates of the prevalence of permanent part-time work – just over 12% of all employed.

The lower permanent full-time employment share in the HILDA Survey is driven mainly by a much larger estimate of fixed-term contract employment – 8.7% compared with just 2.8% in the CoE Survey (though the latter percentage rises to 3.4% if all persons reporting to be employed on both a casual basis and on a fixed-term contract are counted as fixed-term contract workers). This marked difference in the shares of fixed-term contract workers between ABS Surveys and the HILDA Survey has also been found in previous years and

[2] According to HILDA Survey data this group represented less than 0.4% of all employed persons in 2017.

[3] While more recent data from 2017 are available, the questionnaire used in the 2017 round of the CoE Survey did not include the self-identification question about casual employment status.

***Table 1 Estimates of the prevalence of non-standard employment, Australia 2016***

| | *Characteristics of Employment (CoE) Survey* | *HILDA Survey* |
|---|---|---|
| Employment status (% of all employed persons) | | |
| Permanent full-time | 48.6 | 45.1 |
| Permanent part-time | 12.1 | 12.1 |
| Fixed-term contract | 2.8 | 8.7 |
| Casual | 18.9 | 19.6 |
| Both casual and fixed-term | 0.6 | n.a. |
| Self-employed | 17.0 | 13.9 |
| Contributing family worker | n.a. | 0.3 |
| Total | 100.0 | 100.0[(a)] |
| % of employees without leave entitlements[(b)] | 25.1 | 26.0 |
| Respondent | Any responsible adult in household | Worker |
| Reference unit | Main job | Main job |
| Time of observation | August | Mainly August to December |
| Coverage[(c)] | All employed except contributing family workers and members of the permanent defence forces (est. pop. = 11,831,700) | All employed (est. pop. = 11,885,000) |
| Source | Data from ABS, *Characteristics of Employment, Australia, August 2016* (ABS cat. no. 6330.0), and extracted using Table Builder. | HILDA Survey Release 17, confidentialised unit record data file (Department of Social Services / Melbourne Institute of Applied Economic and Social Research 2018). |

Notes:

(a) Includes a small number of cases (about 0.3% of all employed persons) who selected the option 'other' when asked about their employment contract type and who could not be back-coded to one of the three main contract types. While the employment contract type of these employees cannot be determined, they are very likely to be in a non-standard relationship.

(b) Persons reporting that they 'don't know' whether they have entitlements to paid annual leave and / or paid sick leave are treated as not having any entitlements.

(c) The CoE Survey also excludes all persons living in non-private dwellings. The HILDA Survey also excluded such persons at the first wave (but not students living in boarding schools and university colleges) but not at subsequent waves. Given the focus on employment, however, these exclusions and differences should have a negligible impact on the estimates.

discussed by the Productivity Commission (2006: 131-132).[4] They pointed to the administration of the LFS survey (and its supplements) on an 'any responsible adult' basis and the use of a two-part question that first requires respondents to indicate whether their employment has a set end date or event as potential sources of underestimation within the ABS surveys. However, the absence of any clear definition of a fixed-term contract in the HILDA Survey may lead to inconsistencies in reporting across individuals in that data set.

There is also a marked divergence in the two surveys in their estimates of self-employment – the CoE suggests that the self-employed accounted for 17% of total employment whereas the HILDA Survey estimate is just 13.9%. While differences here could again reflect classification errors in the CoE Survey stemming from interviewing one responsible adult in the household rather than every adult, such arguments are less convincing in the case of self-employment. Indeed, if anything we would expect other household members to be more likely to mistakenly assume a self-employed household member (especially if a contractor) is an employee. We thus are drawn to the conclusion that self-employment may be under-enumerated in the HLDA Survey. It is certainly the case that this is a group that, when the HILDA Survey commenced in wave 1, had relatively low response rates. Employment status is not, however, a characteristic used in the construction of population weights and hence it is possible (if not likely) that the weighting procedure used on the HILDA Survey does not adequately correct for this under-enumeration.

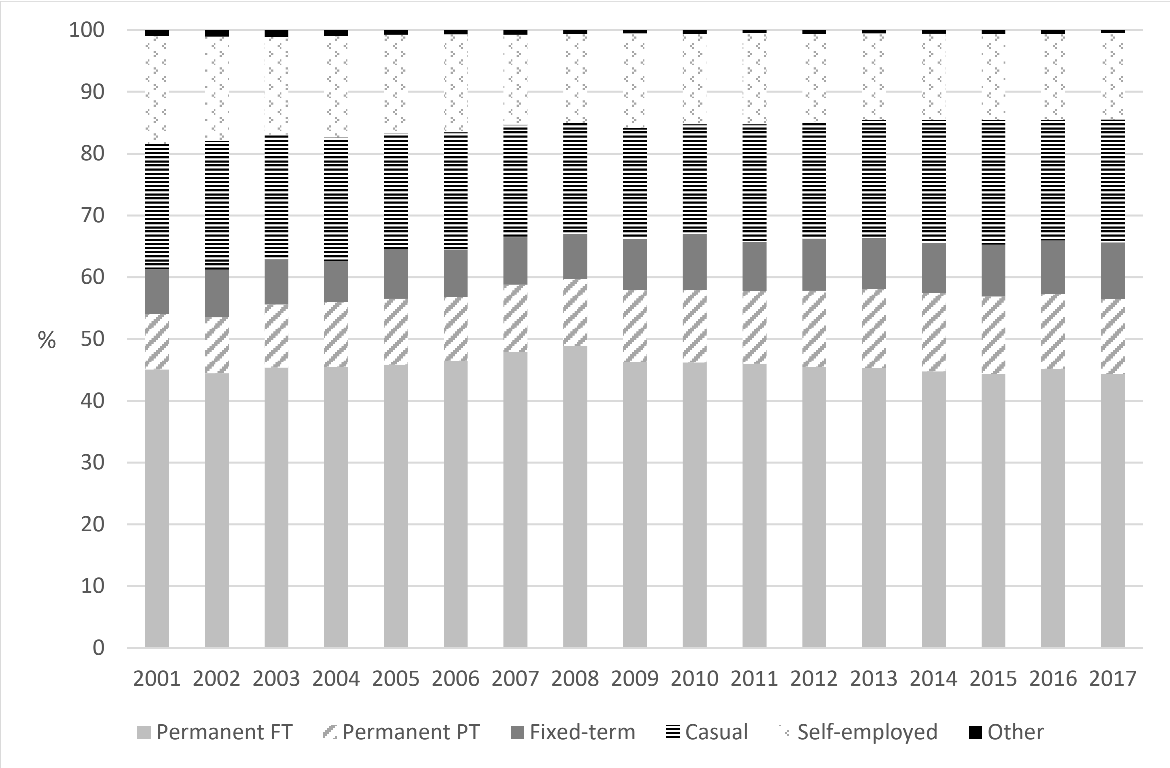
## Recent Trends

Figure 2 uses data from the HILDA Survey to show how the share of standard and non-standard forms of employment, and its composition by employment type, has evolved since the beginning of the millennium. The overall share of non-standard employment in the most recent figures – 55.6% – is as high as it has ever been over the period covered. Nevertheless, this proportion is only slightly higher than the level at the start of the century. Significant changes, however, occurred in the intervening period, with the share falling markedly during the first half of the period – from 54.9% in 2001 to 51.2% in 2008. Since then, and perhaps influenced by the Global Financial Crisis, the trend has reversed.

This pattern of first decline and then increase is not, however, mirrored in the trends for each of the different sub-types of non-standard employment. The permanent part-time employment share has exhibited close to constant growth over much of the period, rising from 9% in 2001 to 12.7% by 2013, before subsiding slightly in the most recent years (12.1% in 2017). There has also been a clear upward trend in the fixed-term contract employment share which has risen from 7.2% in 2001 to 9.1% in 2017 (though growth has been very uneven over the period). In contrast, the self-employment share has moved in the opposite direction, falling from 17.5% in 2001 to 14.0% by 2013, before stabilising in more recent years. Only the casual employment share has exhibited a pattern of first declining (from 20.3% in 2001 to 17.9% by 2010) and then rising (to 20.2% by 2015) so that today's share (19.9%) is similar to (in fact, slightly below) the level of 2001.

[4] The Productivity Commission (2006) was making comparisons between data from the HILDA Survey and the ABS Forms of Employment (FOE) Surveys conducted in 1998, 2001 and 2004. However, the questions used in the FOE Survey to identify fixed-term contract workers were the same as those used in the CoE Survey.

***Figure 2 Trends in employment arrangements, 2001-2017 (% of all employed persons)***



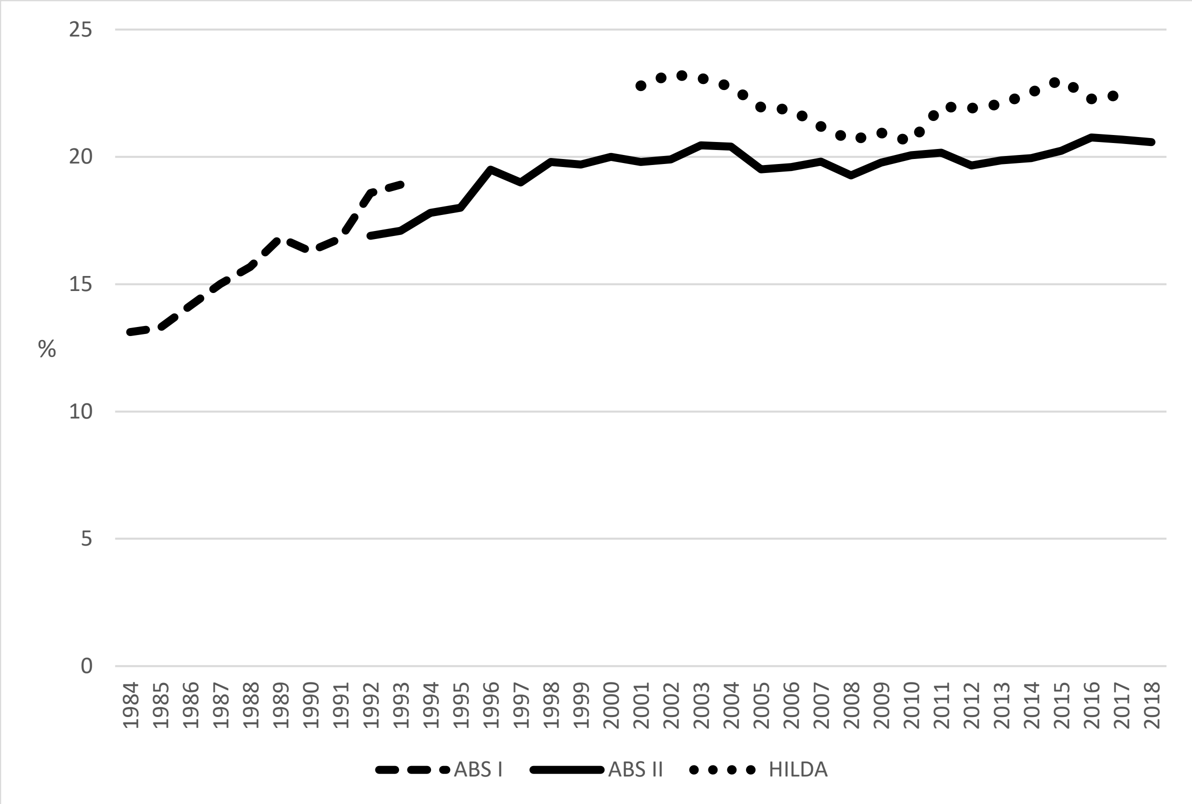

Notes: n=160,174. Data weighted using responding person weights.

Source: HILDA Survey Release 17 (Department of Social Services / Melbourne Institute of Applied Economic and Social Research 2018).

*Rising Workforce Casualisation?*

At first glance the long-term trend in casual employment depicted in Figure 2 appears to contrast with the often-heard claim that work in Australia is becoming increasingly casualised (e.g., ACTU, 2011; Carney and Stanford, 2018). Indeed, arguably the most striking feature of the trend in the casual employment share is how little has changed since 2001. Further, it is possible that the HILDA Survey data overstates how much volatility there has been in the casual employment share. As show in Figure 3, ABS data from the LFS, where casual employment is proxied by the absence of paid leave entitlements, exhibits far more stability over the past two decades than the comparable measure from the HILDA Survey (possibly reflecting lesser sampling variability in the LFS). The proportion of employed persons who are employees without paid leave entitlements is, in the most recent ABS figures (for August 2018), only slightly higher than the level in 2000 – 20.6% vs 20.0%. More importantly, the ABS data show that most of the growth in the casual employment share occurred in earlier decades, and especially in the 1980s and the first half of the 1990s.

***Figure 3 Trends in casual employment using a leave entitlements based definition (% of all employed)***




Notes: ABS estimates are determined by dividing the number of employees identified as having no paid leave entitlements in the August supplement to the LFS by the number of all employed persons as measured in the main LFS. The estimates for earlier years (ABS I) include owner managers of incorporated enterprises in the numerator. For later years (ABS II) this is not the case.

Sources: ABS

1984-1987 -- Dawkins and Norris (1990).

1988-1991 – ABS, *Weekly Earnings of Employees (Distribution), Australia* (ABS cat. no. 6310.0).

1992-2013 – ABS, *Employee Earnings, Benefits and Trade Union Membership, Australia* (ABS cat. no. 6310.0).

2014-2018: ABS*, Labour Force, Australia, Detailed, Quarterly* (ABS cat. no. 6291.0.55.003), Data Cube EQ04.

ABS, *Labour Force Australia* (cat. no. 6202.0), Time Series Spreadsheets Table 1.

HILDA Survey

HILDA Survey Release 17 (Department of Social Services / Melbourne Institute of Applied Economic and Social Research 2018).

### *What About Independent Workers?*

Another trend depicted in Figure 2, and one that again flies in the face of popular perceptions, is the decline in the self-employment share. Given the widespread concerns with both the level of outsourcing by firms (e.g., Bernhardt et al., 2016; Goldschmidt & Schmieder, 2017; Weil, 2014) and the emergence of the 'gig economy' (e.g., Friedman, 2014; Select Committee on the Future of Work and Workers, 2018; Senate Education and Employment References Committee, 2017; Stanford, 2017), one could be excused for believing that the number of self-employed workers must be growing relatively rapidly. The reality, however, is that self-employment growth is well behind rates of growth in other forms of employment. Further, this is not simply an artefact of a weakness in the HILDA Survey sample – it is a pattern that is also confirmed in LFS data. LFS data show that the number of self-employed has, since 2001, been growing by just 0.8% per annum, well below the average annual rate of growth in overall employment over the same period (2.0%).

As shown in Figure 4, however, this relative decline in self-employment is largely concentrated on those in full-time employment. Indeed, the proportion of employed persons who work part-time hours in their own business without any employees has been slowly but steadily rising – from about 3% of employment to around 4% – which may reflect growth in gig work. We suspect that the absence of employees and part-time hours are characteristics of many new gig jobs, especially those where digital platforms play a central role in connecting workers with customers (such as Uber drivers and food delivery couriers). However, it is just as clear that these sorts of jobs are both relatively uncommon[5] and not growing in sufficient numbers to offset the decline in more traditional forms of self-employment. Indeed, in some cases, the new gig workers may be simply displacing other self-employed workers (as might be happening in the taxi industry).

*Growth in Non-standard Employment and Real Wage Stagnation – Is there a Link?*

In summary, the share of non-standard employment in total employment at the end of our period (in 2017) was little different than the share at the start (in 2001). But clearly this share would have grown had it not been for the relative decline in rates of self-employment growth. Given our interest here in the association between non-standard employment and wages, it might be preferable to ignore the self-employed – after all, a meaningful (and comparable) measure of labour income for the self-employed cannot be readily constructed. Once we restrict attention to the population of employees, we find the share of non-standard work has risen from 44.9% of employees in 2001 to 48.3% in 2017. Growth in this share has not, however, been evenly distributed over time. Indeed, the non-standard employment share fell over the period 2001 to 2008. In the following 9 years, the share of non-standard employment in total employee employment rose by 5.5 percentage points. Casual employment and fixed-term contracts each account for 2.1 percentage points of this recent growth and permanent part-time employment 1.4 percentage points.

The coincidence of a rising non-standard employment share in the latter half of the period with a period of weaker aggregate wages growth is suggestive of a link between the two. Nevertheless, the aggregated data do not support a straightforward relationship. In Table 2 we divide our data period into three sub-periods: (i) the pre-GFC period, 2001-2008, during which the non-standard employee share fell; (ii) the period immediately following the GFC, 2008-2013; and (iv) the period since 2013, when nominal wages growth (at least according to the Wage Price Index [WPI]) has barely kept pace with the rate of price inflation. As can be seen, the period of weakest wages growth (2013 to 2017) has been associated with relatively modest increases in the non-standard employment share, whereas the period of strongest growth in non-standard employment (2008-2013) was a time when real wages growth was comparatively strong by recent historical standards.

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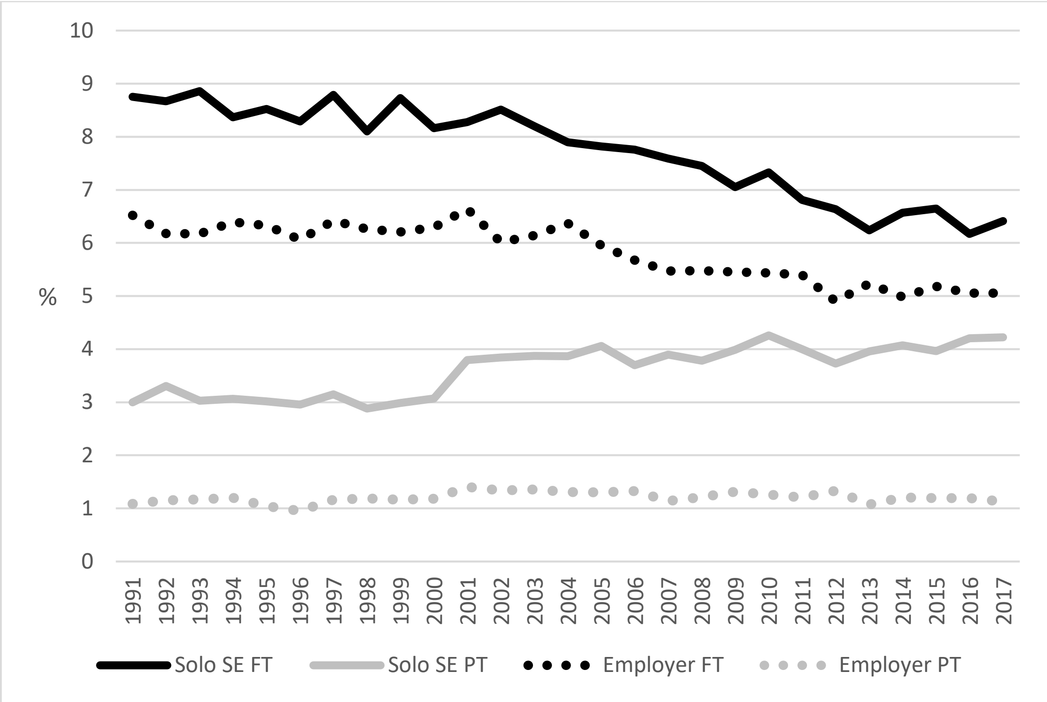
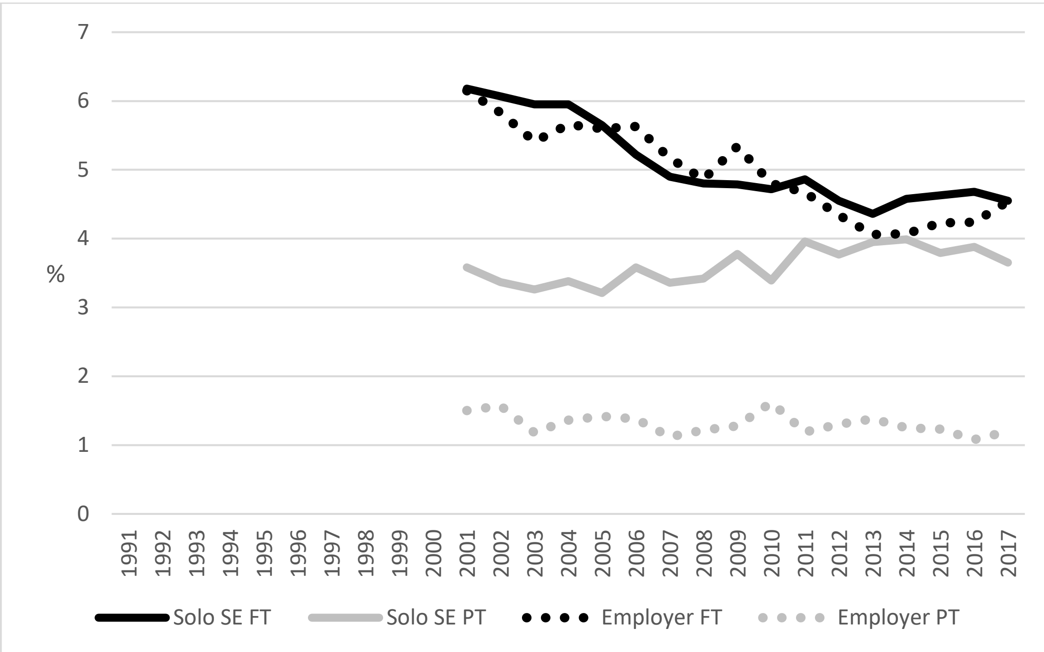
[5] The only study that we are aware of that has attempted to estimate the size of the gig economy in Australia is Minifie and Wilson (2016). Based on information obtained from the major digital platform operators, they concluded that less than half of one percent of Australian adults were, in late 2015, doing any work found or performed on an online platform on a regular basis (at least once a month), and in most cases this work involved very few hours each week.

***Figure 4 Trends in self-employment (% of all employed) by employer status and hours of work***

*A: ABS Labour Force Survey*




*B: HILDA Survey*




Note: The full-time / part-time distinction is based on hours worked in the main job in the HILDA Survey, but on hours worked in all jobs in the LFS.

Sources: ABS, *Labour Force, Australia, Detailed - Electronic Delivery* (cat. no. 6291.0.55.001), Time Series Spreadsheets Table 08.
HILDA Survey Release 17 (Department of Social Services / Melbourne Institute of Applied Economic and Social Research 2018).

*Table 2 The changing share of employees in non-standard jobs (2001-2017) by sub-period*

| *Period* | *Change in non-standard employee share* | | *Real wage growth (WPI)[(a)] (average annual % change)* |
|---|---|---|---|
| | *Total % point change* | *Annual average % point change* | |
| 2001-2008 | -2.1 | -0.3 | 0.7% – Modest real wage growth |
| 2008-2013 | +4.3 | +0.9 | 0.8% – Modest real wage growth |
| 2013-2017 | +1.2 | +0.3 | 0.2% – Low real wage growth |

*Note:* (a) The real wage series used here is the Wage Price Index for total hourly rates pay excluding bonuses (private and public sector) divided by the Consumer Price Index trimmed mean. We take numbers for each year from the September quarter to align with the median interview date in the HILDA Survey.

*Sources*: Non-standard employment: HILDA Survey Release 17 (Department of Social Services / Melbourne Institute of Applied Economic and Social Research 2018).
Wages: ABS, *Wage Price Index, Australia* (cat. no. 6345.0), Timeseries Spreadsheet Table 1.
Prices: ABS, *Consumer Price Index Australia* (cat. no. 6401.0), Timeseries Spreadsheet Table 8.

## Non-standard Employment and Wages

### *Method*

We now move on to examining more closely the links between employment contract type and wages. More specifically, we estimate:

(i) wage equations where the log of hourly wages ($\ln w_t$) is regressed against measures of employment contract type and other worker characteristics; and

(ii) wage growth equations where the outcome is the annual difference in log hourly wages ($\ln w_{t+1} - \ln w_t$) and where all explanatory variables are measured at time $t$.[6]

When estimating wage equations where wages are expressed as levels we report results from both pooled OLS models (but which allow for clustering of individuals over time) and fixed effects models (where estimates are identified by persons that change employment type over time). The importance of controlling for person-specific effects when attempting to estimate the association between employment contract type and hourly wages has previously been emphasised by Booth and Wood (2008). Indeed, based on an analysis of the first four waves of the HILDA Survey data set, they find that workers employed on part-time and / or casual contracts typically earn hourly wage premiums, but these premiums only became apparent once unobserved individual heterogeneity was taken into account.

When estimating models of wages growth the case for controlling for person-specific effects is less clear. Such a specification would imply permanently different rates of wage increase across individuals and hence ever widening wage differentials, which does not

[6] This, for example, is the set up employed by both Mertens and McGinnity (2004) and Giesecke (2009) in their studies of the association between different forms of non-standard employment and wages growth in Germany (though Giesecke is slightly unusual in focusing on wages growth over a two-year period). Far more complex is the approach used by Ameudo-Dorantes and Serrano-Padial (2007) in their study of the impact of fixed-term employment on wages growth in Spain. They differentiate between job stayers and job movers and so apply a switching regression model. Nevertheless, the basic set up is one where the rate of growth in wages is regressed against contract type observed one year earlier.

accord with what is observed in reality. We thus only report results from pooled OLS models when examining annual wage growth.

*Data and Sample*

The data source is again the HILDA Survey, and more specifically Release 17 (Department of Social Services / Melbourne Institute of Applied Economic and Social Research 2018), which provides longitudinal data spanning the period 2001 to 2017.

The HILDA Survey commenced with a nationally representative sample of private Australian households. Members of households that responded at wave 1 are followed over time, with interviews sought on an annual basis with all original sample members aged 15 years or older, as well as any other adults who, in later waves, are residing with an original sample member.[7] In addition, a population refreshment sample was introduced in 2011.[8] In total, Release 17 contains 253,182 observations on 31,206 unique individuals.

Following the approach we have used in previous related research (Laß and Wooden forthcoming), the sample for this analysis is restricted to employees aged between 21 and 64 years (and hence avoiding the complications introduced by junior rates of pay). This exclusion does, however, markedly reduce the incidence of non-standard employment, especially casual employment, given such forms of employment are particularly pervasive among young workers.

Observations where the respondent did not report whether they were employed on a permanent, fixed-term or casual contract, where hourly wages could not be calculated due to missing information on usual weekly working hours, or where the derived hourly wage seemed extreme (below A$6 and above A$200 in 2017 prices), were also excluded. These exclusions resulted in the loss of just 266, 236 and 1565 observations respectively. The final working sample comprised 9312 men and 9504 women, contributing 57,615 and 58,062 observations respectively.[9]

For the analysis of wages growth the sample is smaller due to the effective loss of one wave of data given the need to observe working hours and earnings at two consecutive waves. The working sample for this analysis comprised 7103 men and 7296 women, contributing 44,884 and 44,242 observations respectively.

*Measuring Wages*

Our wage measure is real hourly earnings. It is constructed by dividing reported usual gross weekly wages and salaries in the main job by usual weekly hours of work in that job, and then deflating by the underlying rate of price inflation using the Consumer Price Index trimmed mean series.

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[7] All original sample members, children born to original sample members, and temporary sample members who have a child with an original sample member are added to the sample on a "permanent" basis. In addition, since wave 9 (2009) all temporary sample members who were born overseas and arrived in Australia after 2001 are converted to permanent sample members. From wave 16 this was changed to immigrants arriving after 2011.

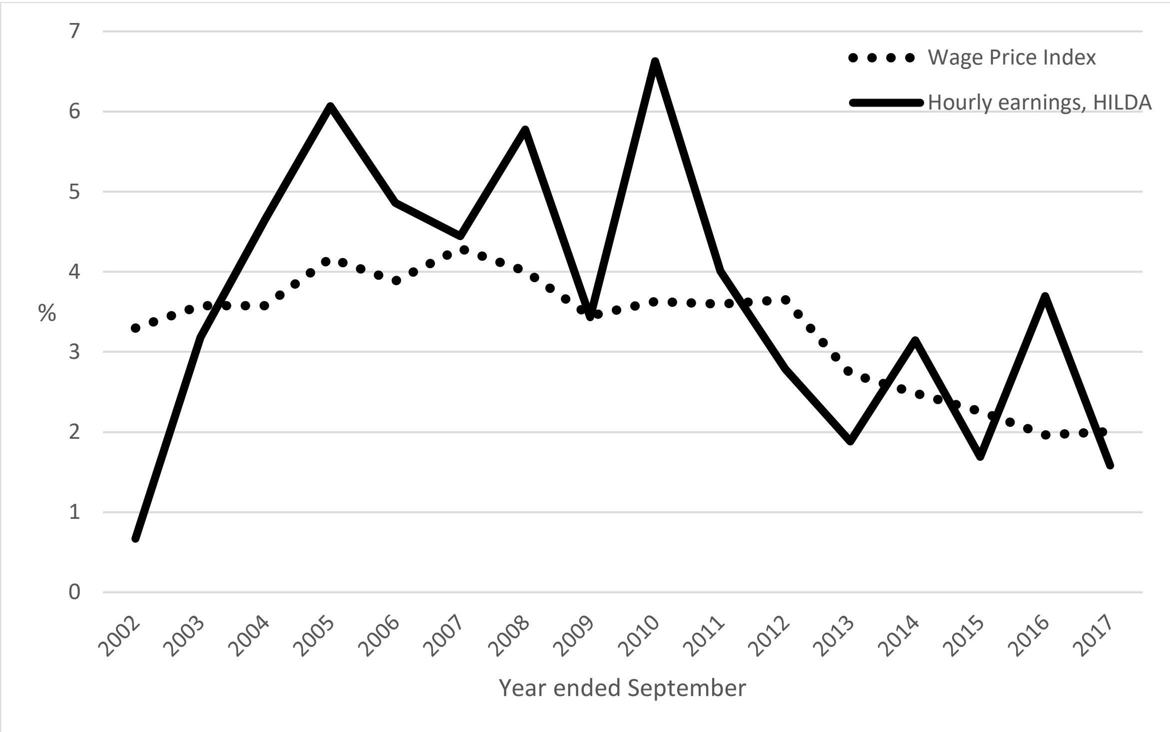
[8] For more details about the sample and survey design, see Watson and Wooden (2012).

[9] The incidence of non-standard employment in this sample is, as would be expected given the omission of young workers, a lot lower than in the wider workforce. The trend, however, is much the same with the non-standard employee share falling from 39.6% in 2001 to 37.5% by 2008 before rising to 43.1% in 2017.

Note that usual hours of work in the HILDA Survey covers all hours spent working on the (main) job, including both paid and unpaid overtime, and including hours worked both in the workplace and at home (and indeed at any location). Hours reported by workers need not (and often will not) align perfectly with the hours that an employer would record as the hours paid for.

While hourly earnings is not the same as what is measured by the ABS in its Wage Price Index (WPI), the annual rates of growth suggested by the two series follow similar paths, with both suggesting a slowdown in nominal wages growth in recent years. This can be seen in Figure 5. The nominal hourly earnings data from the HILDA Survey exhibit far more volatility than the WPI, which is to be expected given the greater noise that is inherent in survey data. Also, with the exception of the first observation, the rates of growth in hourly earnings from the HILDA Survey tend to lie above the rates of growth in the WPI. This is understandable given the WPI is designed to be unaffected by changes in the nature and quality of the work performed or by changes in the characteristics of job occupants. Clearly that is not the case with any survey-based measure of earnings.

***Figure 5 Annual % change in nominal wages: WPI vs HILDA Survey***




Note: Annual changes in the Wage Price Index are calculated for the year ended September using the series for total rates of pay excluding bonuses, private public and sector, and all industries (but note that enterprises primarily engaged in agriculture, forestry or fishing are not covered by the WPI).

Sources: ABS, *Wage Price Index, Australia* (cat. no. 6345.0), Time Series Spreadsheets Table 1.
HILDA Survey Release 17 (Department of Social Services / Melbourne Institute of Applied Economic and Social Research 2018).

*Employment Type*

The way we classify workers into different contractual employment types has been discussed at length earlier. We operationalise that here by including, at least in our main analyses, three dummy variables that distinguish between four different categories of employment: permanent full-time employment (the reference category); permanent part-time employment; fixed-term contract employment; and casual employment.

As previously noted, the sample is restricted to employees and hence the self-employed are excluded. This is dictated by the outcome measure, which can only be meaningfully measured for employees. Such an exclusion might be problematic for our analysis, especially if a significant proportion of the self-employed are working on contracts intended to disguise employment as independent contracting (what is often referred to as 'sham contracting'). It is difficult to identify such arrangements, but given the trends in self-employment reported earlier, it is hard to believe their incidence has been growing over time.[10]

*Control Variables*

Similar to Laß and Wooden (forthcoming), our preferred specification includes controls for a range of socio-demographic and work-related characteristics. We include age (in quadratic form), six dummies for the highest educational level, and a separate dummy for full-time students. The household context is considered by including a dummy for individuals living with a partner and another for those living with their own children below the age of 15. An indicator for the presence of a long-term health condition that impacts on work is also included. We account for regional differences in wage levels in two ways: first, by the inclusion of variables identifying residence in a major city, an inner-regional area, or a more remote area; and second, through the inclusion of indicators for the eight different Australian states and territories. We also control for time effects through the inclusion of year dummies. In the pooled cross-sectional models, we additionally control for indigenous origin and region of birth. Regarding employment-related variables, we include controls for length of tenure with the current employer (specified as a quadratic), whether employed in the public sector, whether has supervisory responsibilities, whether works a schedule other than a regular day schedule, whether had experienced any unemployment in the past year, and membership of a trade union or employee association.

Additionally, we also include a measure of the unemployment rate of the region in which the individual resides, where region is measured at the SA4 level – there are 107 SA4 regions (see ABS 2016). This required accessing more detailed data on geography from the restricted HILDA Survey data release and merging that into our analytical data file.

*Results: Hourly Wage Levels*

Table 3 summarises the mean hourly wages for men and women by employment type within the sample. The results reflect the well-known gender wage gap, with women on average receiving A$31 per hour and men A$36. Women also earn less than men in every single

[10] Results from one now very dated study (VandenHeuvel & Wooden 1995), which reported population estimates based on a sample survey conducted by the ABS as part of its (long discontinued) Population Monitor Survey, suggest that somewhere between 1.1 and 1.3 per cent of employed workers in May 1994 reported being both self-employed contractors and in a "dependent" relationship with the client. This estimate, however, represents at best an upper bound to the extent of sham contracting at that time. In this study the key feature defining 'dependence' was "the provision of services to one or mainly one organization" (p. 268), and of course this on its own does not imply that workers (i.e., the contractors) are being exploited.

employment type, although the difference is quite small among both permanent part-time and casual employees. When comparing wages across employment types, for both men and women, fixed-term contract workers receive the highest hourly wages (though among women the differences in the mean hourly wage of fixed-term contracts workers and permanent employees are not statically significant) and casual employees the lowest. Among men, but not women, we can also see a marked difference in the wages of permanent full-time and permanent part-time employees, with the latter earning on average around 13% less than the former.

***Table 3 Average pooled hourly wages (constant prices) by employment type and gender***

| | *Men* | *Women* |
|---|---|---|
| Hourly wages ($) | | |
| Permanent FT employees | 36.94 | 31.33 |
| Permanent PT employees | 32.24 | 31.94 |
| Fixed-term contract employees | 39.69 | 32.76 |
| Casual employees | 28.47 | 27.70 |
| All employees | 35.81 | 30.87 |
| Wage differences ($) | | |
| Permanent FT vs Fixed-term | -2.75 (3.75)** | -1.43 (3.31)** |
| Permanent FT vs Casual | 8.47 (16.80)** | 3.63 (11.04)** |
| Permanent FT vs Permanent PT | 4.70 (5.26)** | -0.61 (1.56) |
| Fixed-term vs Casual | 11.22 (13.53)** | 5.06 (10.25)** |
| Fixed-term vs Permanent PT | 7.45 (7.63)** | 0.82 (1.60) |
| Casual vs Permanent PT | -3.77 (4.35)** | 4.24 (10.19)** |

Notes: Cross-sectional responding person weights applied.
Wages are in constant (2017) dollars.
Figures in parentheses are absolute values of the t-test of significance of wage differences.
** and * denote statistical significance at the .01 and .05 levels, respectively.

In summary, an increase in the casual employment share can be expected to be associated with an increase in the relative importance of relatively low-paying jobs, while the reverse is true of an increase in the share of fixed-term contract work. Permanent part-time jobs are also relatively low paid, though only among men.

But workers in casual and part-time jobs are also expected to have lower levels of human capital, as reflected in lower levels of experience, education and job-relevant skills and abilities. We thus estimate regression equations that attempt to control for these influences on worker wage outcomes. Estimates of the key parameters of interest are reported in Table 4. The results from simple OLS regressions on the pooled data set are consistent with the descriptive results discussed above. Among men, both casual employment and permanent part-time employment are associated with significant wage penalties relative to permanent full-time employment (of 3.9% and 6.4%, respectively) while fixed-term contract work is associated with a wage premium (of 5.4%). However, similar to what was found by Booth and Wood (2008), the results from the fixed effects regressions tell a very different story: Both casual employment and permanent part-time work are now associated with large and significant hourly pay premiums (of 6.0% and 9.3% respectively).[11] In other words, the relatively low hourly wages paid to men in casual or permanent part-time employment is not simply a function of their employment contract type, but instead a function of unobservable person-specific traits that are associated with lower productivity (e.g., lesser skills and abilities). Fixed-term contract work meanwhile continues to attract a pay premium, but in this case the magnitude of that premium declines in the presence of individual fixed effects, implying that such workers typically have unobserved traits that are pro-productivity.

***Table 4 Non-standard employment and real hourly wages [Outcome = ln real hourly wage]***

| *Employment type* | *Men* | | *Women* | |
|---|---|---|---|---|
| *(Reference category = Permanent full-time)* | *No fixed effects* | *Fixed effects* | *No fixed effects* | *Fixed effects* |
| Permanent part-time | -0.062** | 0.089** | 0.011 | 0.087** |
| Fixed-term contract | 0.053** | 0.026** | 0.007 | 0.045** |
| Casual | -0.038** | 0.058** | -0.029* | 0.089** |
| N | 56577 | 56587 | 56929 | 56941 |
| $R^2$ | 0.29 | | 0.27 | |
| $R^2$ (within) | | 0.15 | | 0.11 |
| $R^2$ (between) | | 0.15 | | 0.19 |
| Rho | | 0.70 | | 0.60 |
| Hausman test | | 1117.0** | | 935.9** |

Notes: ** and * denote statistical significance at the .01 and .05 levels, respectively.

Among women the situation is slightly different. In this case, it is only casual employment that attracts a significant negative coefficient in the pooled OLS regressions. However, in the presence of individual fixed effects the coefficients on all three non-standard employment types become positive and significant, and imply wage premiums ranging from 4.6% for fixed-term contract workers to 9.1% and 9.3% for permanent part-time and casual employees respectively.

We further checked whether the magnitudes of these differentials have been changing over time by re-estimating the fixed effects regressions for three different sub-periods: 2001-2008, 2009-2012, and 2013-2017. The results from these estimations are summarised in Table 5. There is some variation over time, with the coefficients on the indicators of casual and part-time employment largest in the immediate post-GFC period (when the non-standard employee share was rising fastest). Over the longer term, however, there are few differences; with the exception of men on fixed-term contracts, the magnitudes of the estimated wage differentials with permanent full-time employment are much the same in the last sub-period as they were in the first.

[11] The magnitude of the premium for permanent part-time work, however, is much smaller than that reported by Booth and Wood (2008), which ranged between 14% and 15%.

***Table 5 Non-standard employment and real hourly wages by time period [Fixed effects estimates; outcome = ln real hourly wage]***

| | *2001-2008* | *2009-2012* | *2013-2017* |
|---|---|---|---|
| *Men* | | | |
| Permanent part-time | 0.088** | 0.149** | 0.132** |
| Fixed-term contract | 0.034** | 0.030* | 0.012 |
| Casual | 0.069** | 0.092** | 0.063** |
| N | 23106 | 13321 | 20160 |
| $R^2$ (within) | 0.10 | 0.06 | 0.04 |
| $R^2$ (between) | 0.13 | 0.01 | 0.11 |
| Rho | 0.72 | 0.85 | 0.78 |
| *Women* | | | |
| Permanent part-time | 0.088** | 0.117** | 0.102** |
| Fixed-term contract | 0.046** | 0.041** | 0.051** |
| Casual | 0.102** | 0.142** | 0.094** |
| N | 23040 | 13384 | 20517 |
| $R^2$ (within) | 0.06 | 0.05 | 0.04 |
| $R^2$ (between) | 0.17 | 0.10 | 0.04 |
| Rho | 0.63 | 0.70 | 0.71 |

Note: ** and * denote statistical significance at the .01 and .05 levels, respectively.

We also subjected our results to a number of robustness checks. This involved re-estimating our models after:

(i) including additional controls (for industry, occupation and firm size);
(ii) including additional variables identifying persons who were no longer responding at *t+1* or were no longer an employee at *t+1* (as a crude means of identifying whether results are affected by selection);
(iii) omitting observations from wave 1 (given, as shown in Figutre 5, the relatively low rate of mean nominal wages growth between waves 1 and 2 of the HILDA Survey); and
(iv) omitting cases where reported usual hours were greater than 60 per week or fewer than 5 per week.

In each of these cases our main findings were qualitatively unaffected; that is, the magnitude of the estimated coefficients on the employment type dummies changed very little.

Finally, we draw attention to the fact that the estimated magnitude of the wage premium for casual employment – around 6% for men and 9% for women – is still well below the 25% premium currently mandated in awards (or the 20% premium that was widely accepted as the norm in the period prior to July 2010). Given the value of non-wage benefits that casual employees forego (such as paid leave and paid public holidays) is usually assumed to be substantially greater than 6% to 9%, our results support the claim that casual employment is generally not very well rewarded. However, no aggregate measure of wage inflation in use in Australia, including the WPI, takes account of non-wage benefits – they all measure growth in cash earnings. Growth in casual employment should thus be adding to wages pressure (as captured in the WPI) relative to growth in permanent full-time employment, while at the same time reducing overall labour costs.

*Results: Annual Changes in Hourly Wages*

While the evidence presented shows that non-standard employment in Australia is associated with a wage premium once other characteristics and endowments are controlled for, it is still possible that non-standard employees have not done as well as other workers at securing pay rises in recent years. Descriptive statistics presented in Table 6 are mostly consistent with this hypothesis.

***Table 6 Median annual % change in real hourly wages by employment type and time period***

| | *2001-02 to-2016-17* | *2001-02 to 2007-08* | *2008-09 to 2012-13* | *2013-14 to 2016-17* |
|---|---|---|---|---|
| *Men* | | | | |
| Permanent full-time | +2.0 | +2.5 | +1.9 | +1.3 |
| Permanent part-time | -1.8 | -1.5 | -1.0 | -3.8 |
| Fixed-term | +2.7 | +2.6 | +3.3 | +1.5 |
| Casual | +1.6 | +1.1 | +0.3 | +2.6 |
| Total | +1.9 | +2.2 | +1.9 | +1.2 |
| *Women* | | | | |
| Permanent full-time | +2.8 | +3.0 | +2.7 | +2.8 |
| Permanent part-time | +0.1 | +0.6 | +0.2 | -1.0 |
| Fixed-term | +2.8 | +2.4 | +3.8 | +2.5 |
| Casual | -0.5 | +0.3 | -1.3 | -1.6 |
| Total | +1.7 | +2.0 | +1.7 | +1.4 |

Note: Paired longitudinal weights applied.

This table reports the median change in real hourly earnings.[12] Perhaps the first noteworthy feature of this table is that the median annual rate of real wages growth over this period (1.9% for men) and (1.7% for women) is considerably higher than the annual rates of change derived from cross-section data. Most workers who maintain employment from one year to the next have clearly experienced positive real wages growth. Even in the most recent years (2013 to 2017), while real wages growth has slowed, more than half of the employed adult workforce have experienced real wages growth in excess of 1.2%. By comparison, the WPI suggests wages have barely been keeping pace with inflation, growing by just 0.2% per annum over this period. The most obvious explanation for this difference lies in measurement. The WPI measures the wage attached to a specific job, and thus excludes the impact on wages from annual increments, promotions, reclassifications or changing employers, whereas the HILDA Survey is measuring the wages received by individuals and thus includes the effects of these influences.

[12] We report median rather than mean changes given the distribution of the percentages changes is highly asymmetric – by definition it has a lower bound of -100 and no upper bound, with all negative values obviously lying in the range -100 to 0.

But returning to the issue at the centre of this paper, Table 6 also shows that the median annual rate of growth in real hourly wages has been relatively low for permanent part-time employees and, among women, casual employees. Further, as wages growth has slowed, these growth differentials have mostly widened. This is most obvious with respect to permanent part-time employment, where median wage growth over the entire period has been negative (men) or close to zero (women), and where real wage declines are most marked in the most recent years.

Further confirmation for the hypothesis that real wages growth has been relatively slow among employees with non-standard working arrangements is provided by results from regression models. These are reported in Table 7. While the explanatory power of these models is extremely poor[13], two of our employment type variables attract relatively large and statistically significant negative coefficients. Among men the coefficients imply rates of real wage growth (when averaged over the entire sample period) that are about 7 and 3.5 percentage points lower among permanent part-time and casual employees, respectively, than among permanent full-time employees. Among women the comparable differences are around 4.5 and 6 percentage points lower. Table 7 also shows that these negative growth differentials are relatively stable over time.

In contrast, for both sexes, the difference in rate of wages growth between fixed-term contract workers and permanent full-time workers are small and statistically insignificant.

***Table 7 Non-standard employment and real hourly wages growth by sub-period [OLS estimates; outcome = annual change in ln real hourly wage]***

| | *2001-02 to 2016-17* | *2001-02 to 2007-08* | *2007-08 to 2012-13* | *2013-14 to 2016-17* |
|---|---|---|---|---|
| *Men* | | | | |
| Permanent part-time | -0.072** | -0.063** | -0.070** | -0.084** |
| Fixed-term contract | -0.008 | -0.005 | -0.005 | -0.013 |
| Casual | -0.035** | -0.033** | -0.036 | -0.039** |
| N | 44056 | 16599 | 13747 | 13718 |
| $R^2$ | 0.008 | 0.008 | 0.009 | 0.009 |
| *Women* | | | | |
| Permanent part-time | -0.045** | -0.037** | -0.050** | -0.049** |
| Fixed-term contract | -0.008 | -0.009 | -0.011 | -0.005 |
| Casual | -0.057** | -0.052** | -0.061** | -0.060** |
| N | 43349 | 16085 | 13573 | 13702 |
| $R^2$ | 0.008 | 0.007 | 0.009 | 0.010 |

Note: ** and * denote statistical significance at the .01 and .05 levels, respectively.

[13] The only other control variables included that achieve statistical significance are age (and its quadratic), partnership status, the presence of children (for women), public sector employment, and union membership (for men).

So we now have the puzzle that, other things held constant, workers in non-standard jobs appear to be relatively well paid (at last when earnings are measured solely in cash terms), yet both casual and permanent part-time employees have clearly been experiencing relatively low rates of real wages growth. One possible resolution to this puzzle lies in changing employment status – that is, that the largest declines in real wages are experienced by those that shift out of non-standard employment into permanent full-time employment. We would, for example, expect many casual employees to suffer an immediate reduction in hourly earnings upon conversion to permanent employment because of the loss of the casual pay loading. But even among permanent part-time employees (and fixed-term contract employees) there may be a willingness to trade off current wages for the benefits that 'standard' employment might bring, including most obviously access to steeper earnings trajectories in the future. Very differently, changes in hourly wages may simply reflect changes in working hours and, more importantly, the measured return to those hours.

We thus re-estimate our wage growth models after replacing our three employment type dummies with a set of 15 dummies that identify all of the 16 possible combinations of observed employment type at both time *t* and time *t+1*. A summary of the key results is presented in Table 8. These results confirm that the groups that experience the lowest and highest rates of hourly wages growth are those where employment type changes over the one-year window. Perhaps surprisingly, however, it is not those exiting or entering casual employment who experience the largest wage changes. Instead, the largest changes in hourly wages are incurred by those moving between permanent full-time and permanent part-time employment. Among those that move from full-time to part-time employment, the rate of wages growth is almost 21 percentage points higher among men and 16 percentage points higher among women. Our hypothesis is that this largely reflects a shift from jobs where there is a relatively large number of hours being worked above what is specified in employment contracts and awards to jobs where the only hours worked are those specified in contracts and awards. Hourly wages thus rise because the reduction in weekly hours worked is relatively larger than the reduction in weekly pay. Conversely, but for similar reasons, permanent employees that transit in the opposite direction – from part-time work into full-time – experience large wage declines (19 percentage points lower among men and 16 percentage points lower among women).

Casual workers who transit into a permanent full-time job also typically experience relatively large drops in their real hourly wage, while movements in the other directions are associated with sizeable increases. As noted above, this is to be expected given the presence of the casual pay loading. And again, the surprise here is that the changes are not far greater given the magnitude of the casual pay loading. With a relative decline of 6 to 7% among men and 10% among women, shifts into permanent employment out of casual employment may well be welfare enhancing for the worker given both the value of the non-wage benefits and the access to greater promotion prospects and hence higher earnings in the long run.

In contrast, among workers who do not change employment type from one year to the next, relative annual changes in real hourly earnings are much smaller – no larger than 2.4 percentage points. Indeed, among fixed-term contract workers who remain in a fixed-term contract one year later, and male casual workers who remain in a casual job, there are no differences with the reference group (those remaining in permanent full-time employment).

***Table 8*** ***Transitions between employment type and real hourly wages growth [OLS estimates; outcome = annual change in ln real hourly wage]***

| *Employment type at t* | *Employment type at t+1* | *Men* | *Women* |
|---|---|---|---|
| Permanent FT | Permanent FT | [Ref. category] | [Ref. category] |
| Permanent FT | Permanent PT | 0.190** | 0.152** |
| Permanent FT | Fixed-term contract | 0.017* | 0.035** |
| Permanent FT | Casual | 0.052** | 0.118** |
| Permanent PT | Permanent FT | -0.174** | -0.147** |
| Permanent PT | Permanent PT | -0.024** | -0.011** |
| Permanent PT | Fixed-term contract | 0.030 | -0.001 |
| Permanent PT | Casual | -0.082* | 0.026 |
| Fixed-term contract | Permanent FT | -0.009 | -0.001 |
| Fixed-term contract | Permanent PT | 0.032 | 0.022 |
| Fixed-term contract | Fixed-term contract | 0.009 | 0.002 |
| Fixed-term contract | Casual | -0.042 | 0.062* |
| Casual | Permanent FT | -0.063** | -0.096** |
| Casual | Permanent PT | -0.049 | -0.033* |
| Casual | Fixed-term contract | -0.035 | -0.064** |
| Casual | Casual | -0.011 | -0.024** |
| N | | 43979 | 43275 |
| $R^2$ | | 0.015 | 0.022 |

Note: ** and * denote statistical significance at the .01 and .05 levels, respectively.

*Decomposition Analysis*

So, given the non-standard employee share has been rising, at least since 2008, and given that non-standard employment has been found to be associated with significantly lower rates of real hourly wages growth, it follows that the rise in the non-standard employee share should have contributed to low wages growth. But how large is this effect? To help us answer this question we decompose differences in the rate of real hourly wages growth at two periods in time using weighted and unweighted Oaxaca-Blinder decomposition methods. The two periods selected were 2009-10 and 2016-17. As reported in Table 9, in our data set annual real hourly wages growth in the latter period averaged (after weighting) about 0.7% for men and about 1.7% for women. This compares with about 4.3% and 3.4% in the earlier period. More importantly, and reflecting the low explanatory power of our regression models, very little of these differences over time is explained by changes in individual or job characteristics, including changes in the composition of employees by employment type. At most, changes in the non-standard employee share account for about 0.2 of a percentage point of the decline in the real wage growth rate among men, and less than 0.1 of a percentage point among women.

But all effects are, in statistical terms, insignificantly different from zero. In short, the groups where estimated wages growth is relatively low (or relatively high) are tiny. And the sizes of the groups that account for most of the change in employment status over these two periods (e.g., persons that remain in casual employment over each one-year interval) are not changing rapidly enough to be driving down wage growth to any large extent.[14]

## Conclusions

We have shown, using panel data from the HILDA Survey, that, if we ignore the self-employed, the share of non-standard employees in total employment is markedly higher today than at the start of the millennium and all of this increase occurred since the GFC. Further, we have also established that both casual and permanent part-time employment are associated with significantly lower rates of growth in real hourly wages. In some cases this reflects transitions into permanent full-time jobs that are associated with marked and immediate declines in the hourly wage (possibly because not all additional hours worked are directly remunerated). Nevertheless, decomposition analysis suggests that changes in workforce composition by employment type have had a very small impact on the overall rate of wages growth in recent years.

***Table 9 Results from decomposing the difference in mean log real hourly wage growth, 2009-10 vs 2016-17***

| | *Men* | | *Women* | |
|---|---|---|---|---|
| | *Spec I* | *Spec II* | *Spec I* | *Spec II* |
| *Weighted* | | | | |
| Mean log wage growth 2009-10 | 0.042 | 0.042 | 0.033 | 0.033 |
| Mean log wage growth 2016-17 | 0.007 | 0.007 | 0.017 | 0.017 |
| Difference | 0.035** | 0.036** | 0.016 | 0.016 |
| Explained component | 0.0010 | 0.0016 | -0.0010 | -0.0014 |
| Due to employment type | 0.0016 | 0.0021 | 0.0002 | 0.0007 |
| Unexplained component | 0.0338** | 0.0334** | 0.0173 | 0.0174 |
| Change in non-standard employment share (employees aged 21-64) | 0.043 | | 0.015 | |
| *Unweighted* | | | | |
| Mean log wage growth 2009-10 | 0.034 | 0.034 | 0.027 | 0.027 |
| Mean log wage growth 2016-17 | 0.021 | 0.021 | 0.018 | 0.018 |
| Difference | 0.014 | 0.014 | 0.018 | 0.018 |
| Explained component | 0.0014 | 0.0014 | 0.0009 | 0.0008 |
| Due to employment type | 0.0019** | 0.0015 | 0.0003 | 0.0016 |
| Unexplained component | 0.0124 | 0.0128 | 0.0116 | 0.0102 |
| Change in non-standard employment share (employees aged 21-64) | 0.025 | | 0.022 | |

Note: Specifications I and II only differ in the way employment type is represented. Specification includes three employment dummies measuring employment type at time t, while Specification II includes 15 dummies representing combinations of a person's employment types at time *t* and *t+1*.
** and * denote statistical significance at the .01 and .05 levels, respectively.

[14] Appendix A provides descriptive data on the incidence of change within our samples.

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## Appendix

***Table A1 Median annual % wage change by employment type, gender and whether employment type changed (figures in parentheses are the percentage of employed persons in each cell)***

| | *Employment type at t+1* | | | | |
|---|---|---|---|---|---|
| *Gender / Employment type at t* | *Permanent full-time* | *Permanent part-time* | *Fixed-term* | *Casual* | *Total* |
| Men | | | | | |
| Permanent full-time | +1.8 | +18.6 | +3.4 | +5.3 | +2.0 |
| | (68.0) | (0.9) | (3.8) | (2.1) | (74.8) |
| Permanent part-time | -10.7 | -0.2 | -6.5 | +7.4 | -1.8 |
| | (1.1) | (2.3) | (0.3) | (0.3) | (4.2) |
| Fixed-term | +2.1 | +3.0 | +4.3 | -2.0 | +2.7 |
| | (4.4) | (0.3) | (3.9) | (0.5) | (9.0) |
| Casual | -0.5 | +1.5 | +5.3 | +1.9 | +1.6 |
| | (3.0) | (0.7) | (0.9) | (7.5) | (12.1) |
| Women | | | | | |
| Permanent full-time | +2.1 | +14.8 | +4.9 | +11.1 | +2.8 |
| | (39.9) | (3.0) | (3.0) | (1.4) | (47.3) |
| Permanent part-time | -7.4 | +0.6 | +2.3 | +4.7 | +0.1 |
| | (3.1) | (17.5) | (1.4) | (1.8) | (23.8) |
| Fixed-term | +2.5 | +2.9 | +2.6 | +4.6 | +2.7 |
| | (3.4) | (1.5) | (4.7) | (0.7) | (10.2) |
| Casual | -5.0 | +1.8 | -3.9 | +0.4 | -0.5 |
| | (2.4) | (2.5) | (1.4) | (12.4) | (18.7) |

Notes: Sample = Employees aged 21 to 64 years.
Paired longitudinal weights applied.

***Table A2 Incidence of employment types, 2009-10 and 2016-17 (%)***

| *Employment type at t* | *Employment type at t+1* | *2009-10* | *2016-17* | *Change* |
|---|---|---|---|---|
| *Men* | | | | |
| Permanent FT | Permanent FT | 69.2 | 64.8 | -4.4 |
| Permanent FT | Permanent PT | 0.7 | 0.5 | -0.2 |
| Permanent FT | Fixed-term contract | 4.1 | 4.1 | -0.1 |
| Permanent FT | Casual | 1.9 | 2.2 | +0.3 |
| Permanent PT | Permanent FT | 0.9 | 0.9 | +0.0 |
| Permanent PT | Permanent PT | 2.3 | 3.3 | +1.0 |
| Permanent PT | Fixed-term contract | 0.3 | 0.5 | +0.1 |
| Permanent PT | Casual | 0.6 | 0.5 | -0.1 |
| Fixed-term contract | Permanent FT | 4.4 | 3.9 | -0.6 |
| Fixed-term contract | Permanent PT | 0.2 | 0.3 | +0.2 |
| Fixed-term contract | Fixed-term contract | 4.8 | 4.0 | -0.8 |
| Fixed-term contract | Casual | 0.9 | 0.6 | -0.3 |
| Casual | Permanent FT | 2.4 | 3.7 | +1.3 |
| Casual | Permanent PT | 0.4 | 0.6 | +0.2 |
| Casual | Fixed-term contract | 1.0 | 1.4 | +0.4 |
| Casual | Casual | 5.9 | 8.6 | +2.8 |
| *Women* | | | | |
| Permanent FT | Permanent FT | 40.0 | 38.4 | -1.6 |
| Permanent FT | Permanent PT | 3.2 | 3.3 | +0.1 |
| Permanent FT | Fixed-term contract | 3.3 | 3.2 | -0.1 |
| Permanent FT | Casual | 1.3 | 1.4 | +0.1 |
| Permanent PT | Permanent FT | 2.3 | 3.5 | +1.2 |
| Permanent PT | Permanent PT | 19.2 | 17.8 | -1.5 |
| Permanent PT | Fixed-term contract | 1.9 | 1.3 | -0.6 |
| Permanent PT | Casual | 1.8 | 1.4 | -0.4 |
| Fixed-term contract | Permanent FT | 3.1 | 3.4 | +0.3 |
| Fixed-term contract | Permanent PT | 1.4 | 1.7 | +0.2 |
| Fixed-term contract | Fixed-term contract | 4.5 | 6.3 | +1.8 |
| Fixed-term contract | Casual | 0.6 | 1.0 | +0.3 |
| Casual | Permanent FT | 3.0 | 3.4 | +0.4 |
| Casual | Permanent PT | 1.4 | 1.7 | +0.2 |
| Casual | Fixed-term contract | 1.0 | 1.6 | +0.6 |
| Casual | Casual | 11.3 | 12.1 | +0.7 |

Notes: Sample = Employees aged 21 to 64 years.
Paired longitudinal weights applied.