Job security perceptions and its effect on wage growth[1]

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Abstract

A concern that low job security is constraining wage growth has been expressed in many countries. In this paper, we use Australian household panel data to analyse the drivers of self-assessed job security and its relationship with wage growth. We construct measures of industry-level trade exposure and occupation-based automation risk to assess the conjecture that technological change and globalisation are leading to fears of job loss. We find that those in jobs with a higher trade exposure or automation risk or those working on a casual or fixed-term contract feel more insecure in their job. However, regardless of one's characteristics, there has been a broad-based fall in job security in recent years that cannot be explained by the model variables. Exploiting the panel dimension of the survey, we find that heightened job insecurity has been a statistically significant but small drag on wage growth. This result is robust to various model specifications.

JEL Codes: C23, F16, J28, J31, O33

Keywords: wages, job security, automation, panel data

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1. Introduction

Workers' perceptions of their own job security may influence how hard they bargain over their wages. Several factors may contribute to these subjective assessments of job security, including labour market conditions, the type of employment contract, union membership status, as well as the potential for their work to be automated or sent offshore. Shifts in these factors could make workers feel less secure in their job and so more likely to accept smaller wage rises. The concern that low job security is constraining wage growth has been expressed across a number of advanced economies recently (Haldane 2017; Lowe 2017), as well as in the United States two decades ago (Greenspan 1997). Yet there has been little research that empirically tests this relationship.

We study perceived (or self-assessed) job security, which is a worker's subjective expectation that they will retain their job. These beliefs – whether justified or not – can influence wage bargaining behaviour. For instance, if an employee believes that their job could be moved offshore, they may accept lower wage rises to encourage their employer to keep the job onshore. It does not matter whether or not the job is eventually moved, because it is this fear of job loss, or perceived contestability of the job, that would keep wage growth contained. This is in contrast to actual (or objective) job security, which is the likelihood that a worker does hold on to their job. Actual job insecurity is typically equated with jobs that have relatively high turnover or are temporary in nature.[3] Determinants of objective job security, such as casualisation and labour market conditions, can contribute to perceptions of job security. Personal circumstances, popular narratives, expectations of the future, and domestic or international events may also play a role.

This paper documents the decline in perceptions of job security in Australia in recent years, the cyclical and structural factors that have driven it, and shows that it explains a small part of recent weak wage growth. Like many advanced economies, Australia has experienced low wage growth in recent years that is difficult to explain in a conventional Phillips curve framework. As such, the Australian experience can shed light on wage dynamics more broadly. We use data from the Household, Income and Labour Dynamics in Australia (HILDA) Survey, which is a household panel survey that asks respondents about their feelings of job security and perceived likelihood of job loss. Along with questions on income, job and personal characteristics, this allows us to track whether workers who felt more secure in their job received a higher pay rise than their counterparts who were more anxious about their job security. We explore the existing literature on this topic in the next section. In Section 3 we describe the data and construction of the trade exposure and automation variables used. Next we explore the determinants of job security and its relationship to wage growth. We then perform a series of robustness checks, including accounting for selection and endogeneity effects. Section 6 concludes.

3 See de Ruyter and Burgess (2000) for an overview of the various dimensions of the job security definition.

2. Previous literature

Determinants of job security

The literature on the determinants of perceived job security in Australia is small. Some of the HILDA Statistical Reports have included a broad overview of the determinants of perceived job security, using the measure of the expected probability of losing one's job (Heady and Warren, 2008; Wilkins, 2015). Taken together they find age, tenure, education, hourly wage, occupation and industry influences job security perceptions, whereas gender and union membership do not. In addition, Borland (2002) found that perceived job-security moves pro-cyclically with the business cycle. Previous work using HILDA has found that job insecurity can be explained by regional unemployment, casual work, fixed term jobs (Milner et al. 2014) and job tenure (Green and Leeves 2013).

The international literature supports this, with job insecurity in the UK, for example, being partially explained by temporary contracts and short tenures (Green 2003). Other characteristics that are found to reduce perceived job security overseas include previous unemployment experience, regional unemployment, having a job that is manual, in the private sector, being male and being older (Böckerman 2004; de Bustillo and de Pedraza 2010; Campbell et al. 2007; Demoussis and Giannakopoulous 2007; Green, 2003; Hübler and Hübler 2006; Lurweg 2010; Näswall and de Witte 2003). All these studies examine this question with household cross-section or panel datasets. Despite globalisation and automation often being cited as drivers of job insecurity, there is little quantitative research to show this. Lurweg (2010) explores the relationship between globalisation and job security fears, finding that those in trade-exposed service industries in Germany had greater perceived and realised job insecurity. Vieitez et al. (2010) studied workers at a particular factory in Spain and showed that the types of technology used in their job influenced their perception of job security.

Job security and wages

The relationship between job security and wages can arise through either a bargaining framework model or a compensating wage differentials model. Under the bargaining model, job security influences the relative bargaining strength of employees (or unions) and employers. Greater perceived job security by employees raises their relative bargaining power, so they can demand higher wages; as such, wages and job security are complements and have a positive relationship. Bargaining models tend to be specified in terms of wage levels, but wage growth easily fits within this framework. The alternative framework sees monetary compensation (wages) and non-monetary features (such as job security) as substitutes within a total compensation package. For example, government employees may accept lower wages as a trade-off for greater job security, whereas a professional services firm may provide higher monetary compensation due to employment prospects fluctuating with the business cycle. However, this mechanism is more relevant for wage levels rather than wage growth.

Hübler and Hübler (2006) test both these theories and find support for the bargaining hypothesis. Using household-level panel data, they find that the level of

job insecurity negatively affects wage levels, and the change in job security negatively affects wage growth, in Germany and the United Kingdom, although the growth effect was not robust. Also using household level data, Campbell et al. (2007) found that job insecurity lowers the rate of wage growth for men, but not women, in the United Kingdom. For the United States, Aaronson and Sullivan (1998) use regional data to augment a wage Phillips curve with average regional perceived job security and find there is a negative, but not statistically significant, relationship between wage growth and fears of job loss. The relationship between perceived job security and wage growth has not been examined in detail for Australia.[4]

Before examining the role of self-assessed job security on wage growth, it is helpful to assess whether job security is useful in predicting unemployment itself. The literature shows that perceived job security can modestly predict job loss (Dickerson and Green 2012; McGuinness et al. 2014). However, workers tend to overestimate their probability of job loss. This overestimation lends credence to the idea that expectations of potential job loss can put downward pressure on wages, independent of actual labour market conditions.

3. Data

We use the HILDA Survey from 2001 to 2016. It is a representative household panel dataset that includes several measures of self-assessed job security, as well as labour income and a vast array of personal and job characteristics.[5] As such, we can use it to track workers and measure whether their levels of job security influenced their growth in wages over the following year. We also incorporate data from the Australian Bureau of Statistics (ABS) and Frey and Osbourne (2017) to measure risks of offshoring and automation, as well as data on state-level consumer price growth from the ABS.

The HILDA survey asks all employees about job security.[6] This leaves us with 124,617 responses over 16 years across 20,882 individual employees. For the analysis on determinants of job security we take no other exclusions to our sample. Nonetheless, with missing observations for some explanatory variables in the regression, we have 100,255 observations over 18,621 individuals. The summary statistics and graphs are shown with population-weighted data, while the regressions are unweighted. For the analysis on wage growth we omit those observations where the respondent reports zero wages or hours in a year. We also exclude those that changed jobs during the year.[7] A person may change jobs for many reasons and this will impact on their measured wage

4 Examining the relationship between job security and wage *levels*, in a compensating differentials analysis, Kelly, Evans and Dawkins (1998) found that job security matters most to those on low incomes.

5 The HILDA survey tracks households and individuals on an annual basis. It incorporated a top-up sample in 2011, to account for issues relating to attrition and recent migrant arrivals. The response rates are high, around 95 per cent for previous responders and 80 per cent for new individuals. For more information see Wilkins and Lass (2018) and references within with respect to sampling methodology.

6 As such, employers, own account workers and contributing family members are excluded.

7 A job change can be a change in employer, new occupation, industry etc. We have taken a self-assessed measure of job change in this paper, but the results are robust to several definitions available using the HILDA survey.

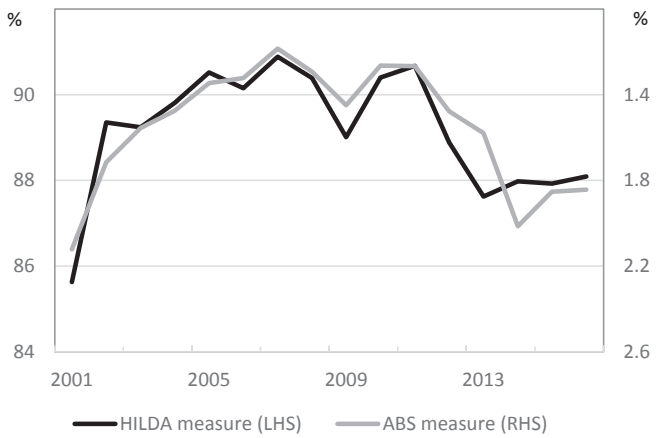
growth in ways that may or may not be related to job security concerns. In particular, we are interested in understanding how job security impacts an employee's ability to bargain for wage growth in their current job. This results in a reduction of 16 per cent of the observations. However, the robustness check in Section 5 demonstrates that it does not impact the overall conclusion. Overall, the sample for the wage regression falls to 65,864 observations, covering 13,707 employees.

Measures of subjective job security

The HILDA survey includes a range of indicators that can be used to gauge a worker's perception of their own job security. The key measure we focus on is a probability measure, which asks respondents the percentage chance of losing their job over the next 12 months. We invert this measure to frame it as the self-assessed probability of retaining one's job over the next year. That is, the probability of job retention is 100 less the stated probability of job loss. This measures perceptions of short-term job security, which should be more correlated with near-term wage outcomes, as opposed to concerns of future job loss through long-term changes in the structure of the economy.

The variable's movements closely match a measure of the probability of job retention from the Australian Bureau of Statistics (Figure 1). This gives some external validity that this measure is a sensible one to use. An alternative variable is one that measures subjective job security satisfaction. It is an ordinal measure, which asks how satisfied one is with their job security on a scale of 0–10. This ordinal job security satisfaction measure also moves closely with the probability measure and we examine its determinants in Section 5 for robustness.

Figure 1: Job security perceptions

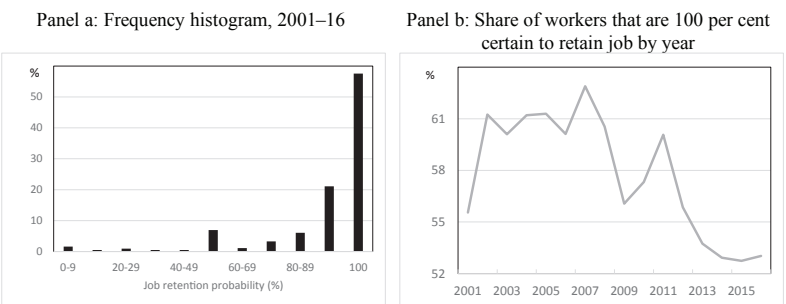


Notes: HILDA measure is the average job retention probability (100 less probability of job loss over the next 12 months). ABS measure is the annual average employment share of those expecting to leave their employer over the next 12 months due to downsizing/ closing or is a temporary worker.
Sources: ABS cat no. 6291.0.55.003, HILDA Survey Release 16.0.

While job security perceptions do not vary much, they appear to move in line with changes in overall labour market conditions. Job security is relatively low when there is ample spare capacity in the labour market or underutilisation in the economy. There was a brief spike in the unemployment rate and slowdown in growth in 2001. Low subjective job security at the time appears to reflect this. Aggregate wage growth also tends to be lower when workers felt less secure on average the year earlier. This suggests that perceived job security moves with the business cycle. As such, it is unclear whether job security can help explain movements in wage growth over and above general labour market conditions. These issues are explored in this paper.

Workers tend to have a high rate of self-assessed job security; the probability of expected job retention averages 89 per cent across the whole sample period. The variable is also highly skewed (Figure 2a). Almost 60 per cent of workers said they had a 100 per cent chance of retaining their job in the next 12 months. For those who have some uncertainty around job loss, the probability they assign to job retention is still high, with less than 4 per cent of workers expecting more than a 50 per cent chance of job loss. This highlights that only a small share of the workforce has concerns about holding on to their job over the following year.

Figure 2: Subjective probability of job retention over the next year



Sources: HILDA Survey Release 16.0.

When we model the determinants of job security, we transform this probability measure into a dummy variable. We refer to it as 'job retention certainty' and it is equal to 1 if a worker is 100 per cent sure of retaining their job over the next year, and 0 otherwise.[8] We do this for a number of reasons. First, the job security dependent variable is non-normally distributed, being bounded by 0 and 1 and highly skewed. As such, making inferences on that measure would be problematic. Second, respondents may interpret the in-between probabilities differently (say, 90 per cent probability of job retention may feel high for one person, but low for another). The bunching at round numbers suggests there is difficulty in interpreting the probabilities. Finally, much of the variation in this measure over the past decade has been in the proportion

8 Sensitivity around this cut-off is examined in Section 5.

of workers that are completely certain that they will keep their job (Figure 2b). At its peak in 2007, 63 per cent of workers had absolute certainty that they would keep their job over the next year. This proportion troughed at 53 per cent in 2016. When we turn to the wage growth model, we use the raw probability variable, which goes from 0 to 1. This allows us to retain the full information from the variable. The non-normality of the distribution is no longer a concern as it is an independent variable, in this case.

Measuring wage growth

We use the HILDA data to calculate annual growth in average hourly earnings for each worker. We take weekly gross wages and salary divided by hours usually worked per week in one's main job. We omit observations where the respondent reports zero wages or hours in a year. These are log-differenced at the individual level and the highest and lowest 5 per cent of the hourly wage growth distribution for each year are winsorised, to mitigate the influence of outliers.[9] The 25th percentile, median, mean and 75th percentile of wage growth over the whole period is -6.9, 4.6, 5.7 and 18.2 per cent, respectively. Our measure tracks aggregate wage growth measured by the ABS fairly well, although it has been more stable of late (Figure 3).

Figure 3: Annual wage growth

Notes: HILDA measure is calculated as the mean of the log difference in wage per hour worked in one's main job.
Sources: ABS cat no. 6345.0, HILDA Survey Release 16.0.

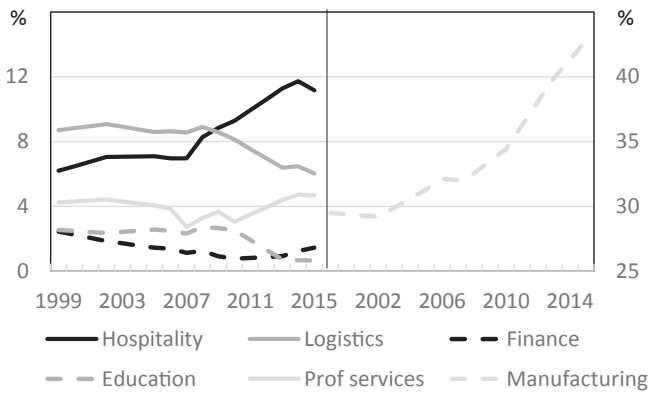
Measuring trade exposure and automation risk

Two factors that are commonly suggested to influence both actual job loss and perceptions of job security are globalisation and technological change. An industry's exposure to global trade can be measured using its import penetration ratio. These ratios are calculated as the import share of the total output supplied by an industry. It represents the foreign

9 That is, we take ln[W(i,t) – W(i,t-1)] for each year t and individual i.

share of the competition that firms in that industry face in the domestic market. Data are available from the ABS input-output tables for each industry subdivision across a range of financial years, with data for any missing years interpolated linearly. The import penetration ratio of the manufacturing industry is relatively high and has been increasing over the past decade as domestic demand for manufactured goods is increasingly met by imports (Figure 4). The hospitality and professional services industries have also faced rising import penetration, whereas trade exposure in the education industry is lower and has declined somewhat in recent years.

Figure 4: Import penetration ratios by supply industry



Notes: Selected industries; financial year; missing data are linearly interpolated.
Sources: ABS cat no. 5209.0.55.001.

Overall, the import penetration ratios of most industry subdivisions have been increasing in recent years. However, the effect of this on the aggregate labour market has been offset by stronger employment growth in industries that are less exposed to international trade. As a result of this compositional shift, the trade exposure faced by the representative worker has been little changed over the past decade, having declined over several years prior to that.

Another source of job insecurity is the risk that technology could be used to automate tasks that were previously completed by skilled workers. We capture this risk using the computerisation scores devised by Frey and Osborne (2017). These scores measure the likelihood that an occupation's tasks could be automated using current computing technology.[10] We map these scores to the Australian unit group occupations

10 A shortcoming of the Frey and Osborne (2017) approach is that the automation scores depend on the initial subjective decisions about which tasks have the potential to be automated using current technology (Borland and Coelli 2017). The method also does not allow for technology to complement labour and change the nature of an occupation. Therefore, these scores may not accurately measure the likelihood for each occupation to be automated. Nonetheless, the scores are well suited to measuring general perceptions of technological job loss, which is relevant to the topic of this article.

to capture the potential for each of these jobs to be automated. The scores are not available over time and so we cannot evaluate how job security is being affected by the pace of technological change. As expected, routine occupations are the most at risk of automation, and non-routine cognitive jobs are the least automatable (Table 1).[11] The automaton risk among non-routine manual occupations is varied, which includes service jobs, particularly in the health and hospitality industries. As is the case with industries highly exposed to global trade, occupations that have a greater potential to be automated have been decreasing as a share of total employment in recent years (Edmonds and Bradley 2015).

Table 1: Automation score by occupation type, weighted by population

Occupation type	*Median*	*25th percentile*	*75th percentile*
Routine cognitive	0.91	0.78	0.96
Routine manual	0.75	0.64	0.81
Non-routine manual	0.48	0.29	0.66
Non-routine cognitive	0.08	0.04	0.17

Sources: ABS Cat no. 1220.0, Frey and Osbourne (2017), HILDA Survey Release 16.0, author's calculations.

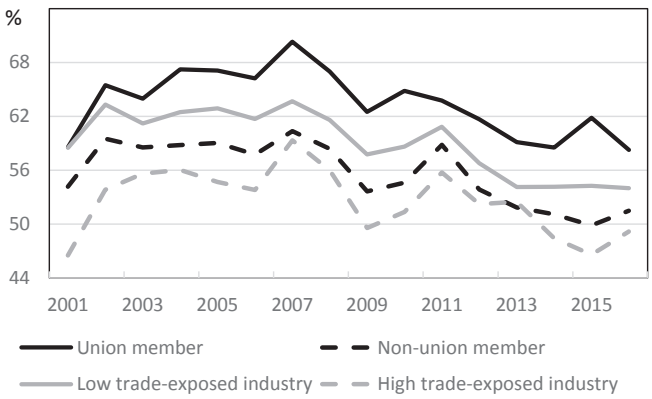
In the models, the trade exposure and automation variables are entered into the models linearly (quadratic terms were statistically insignificant).

Trends in job security

This section explores how economic conditions, structural factors and individual characteristics affect perceptions of job security in Australia. Perceived job security can be influenced by factors such as industry sector, occupation and work arrangements, as well as more general concerns about globalisation and automation. Figure 5 illustrates just two of these dimensions, related to the issues of offshoring and bargaining power, but we discuss the trends across additional factors too. Perceived job security is lower in industries that are more exposed to international competition.[12] The decline in perceptions of job security in recent years is a touch larger in more trade-exposed industries, but the fall has been broad based across all industry sectors. A fall in job security perceptions in business services has occurred despite strong employment growth in that sector, which suggests that labour market conditions are not the only driver of job security concerns.

11 The routine–non-routine and cognitive–manual occupation classification definitions for Australian data are described in Heath (2016).
12 An industry is classified as having high trade exposure if the import penetration ratio is above 5 per cent.

Figure 5: Job security by select job characteristics, share with job retention certainty, weighted by population



Sources: ABS cat no. 5209.0.55.001, HILDA Survey Release 16.0.

Union members and non-casual workers feel more secure in their job, relative to their non-union and casual counterparts. There is only a small difference in the job security of part-time and full-time workers. Those that are satisfied with their hours have higher job security than those wanting fewer or more hours, but concerns about job security have declined for both types of workers. Finally, those in regions with higher unemployment on average tend to have lower job security.

Despite some differences in the level of job security perceptions across various dimensions, a decline in these perceptions is apparent for almost all job characteristics, regardless of industry, occupation, contract type or union membership. The decline has also occurred across a range of personal characteristics, such as age, sex and educational attainment, as well as across most states. Overall, these patterns suggest that job and personal characteristics influence the *level* of perceived job security, but that the *decline* has been experienced across the board.

4. Job security trends, determinants and its relationship to wage growth

Determinants of job security

The descriptive analysis above suggests that a range of factors influence job security perceptions. To examine their relative importance, we estimate a random effects probit model for the job certainty dummy variable. We explore alternative panel data models and measures in Section 5. The model is estimated on an unbalanced panel of workers from 2001–2016. We include a range of job, firm and personal characteristics. The summary statistics for these independent variable and coefficient estimates are shown in Table 2.[13] We have not included interaction terms.

13 Summary statistics are shown for the wage regression sample. Concepts such as full-time/part-time and contract type used in the HILDA survey are based on definitions from the ABS (2018).

Table 2: Regression coefficient estimates – random effects models, 2002-16

	Variable Summary[a]		*Job Security*[b] *Probit model*		*Wage Growth*[c] *Linear model*	
	Mean	*S.D.*	*Coefficient estimate*	*S.E.*	*Coefficient estimate*	*S.E.*
Job security (t-1)[d]	0.92	0.17	--		1.62***	0.52
Economic Factors						
Unemployment (by region)	5.21	1.02	-0.03***	0.01	0.19	0.13
Trade exposure	0.05	0.11	-0.25***	0.07	-1.57**	0.71
Automation	0.48	0.35	-0.08***	0.03	-1.06***	0.35
State consumer price growth	2.50	0.85	--		0.61**	0.30
Industry (base: Goods production)[e]						
Business services	0.16		-0.09***	0.03	-0.38	0.30
Household services	0.38		0.16***	0.03	-1.03***	0.29
Goods distribution	0.18		0.08***	0.03	-1.12***	0.29
Other	0.11		0.08**	0.03	-0.40	0.34
Occupation (base: Non-routine cognitive)						
Routine manual	0.23		0.17***	0.03	0.19	0.31
Non-routine manual	0.15		0.09***	0.03	-0.45	0.28
Routine cognitive	0.25		0.10***	0.03	-0.40	0.31
Job Characteristics						
Lost a job in the past year	--		-0.28***	0.03	--	
Changed jobs in the past year	--		-0.09***	0.01	--	
Promoted in the past year	0.11		0.13***	0.02	3.61***	0.29
Full time	0.70		0.08***	0.02	-4.10***	0.23
Casual	0.17		-0.13***	0.02	-0.08	0.30
Temporary contract	0.08		-0.27***	0.02	0.54*	0.32
Ever work from home	0.19		0.06***	0.02	-0.85***	0.23
Number of years in a job	8.22	8.07	0.01***	0.00	-0.02*	0.01
Trade union member	0.32		0.01	0.02	0.07	0.18
Firm-type (base: private sector)						
NGO	0.09		0.10***	0.03	0.14	0.32
Public	0.30		0.19***	0.02	0.51**	0.23
Firm-size (base: small firms)						
Medium	0.32		-0.08***	0.02	0.08	0.20
Large	0.37		-0.12***	0.02	0.15	0.20
Hours preference (base: same)[f]						
Fewer	0.27		-0.09***	0.01	2.89***	0.20
More	0.12		-0.08***	0.02	-2.24***	0.30
Personal Characteristics						
Female	0.50		0.35***	0.02	-1.34***	0.17
Primary language is not English	0.08		-0.02	0.04	1.30***	0.31
Born outside of Australia	0.19		0.05	0.03	-0.58***	0.21
Outside a capital city	0.38		0.15***	0.02	-0.28	0.18
Age (base: over 55s)						
Under 25	0.14		-0.32***	0.03	6.39***	0.33
25 to 39	0.32		-0.23***	0.03	1.23***	0.23
40 to 54	0.38		-0.15***	0.02	-0.34*	0.20
Education (base: high school)						
Diploma or Certificate	0.33		-0.03	0.02	-0.04	0.17
University	0.30		-0.25***	0.03	-0.31	0.22

Continued / ….

	Variable Summary[a]		*Job Security*[b] *Probit model*		*Wage Growth*[c] *Linear model*	
	Mean	*S.D.*	*Coefficient estimate*	*S.E.*	*Coefficient estimate*	*S.E.*
State (base: NSW)						
Victoria	0.25		-0.02	0.02	-0.59***	0.19
Queensland	0.21		0.02	0.03	0.17	0.20
South Australia	0.09		0.15***	0.03	-0.32	0.27
Western Australia	0.09		0.35***	0.04	0.44	0.27
Tasmania	0.03		0.03	0.05	-0.62	0.38
Northern Territory	0.01		0.03	0.09	0.43	0.83
ACT	0.02		0.07	0.06	0.97**	0.49
Year (base: 2002)						
2003	0.06		-0.06**	0.03	0.69	0.63
2004	0.06		-0.04	0.03	1.14*	0.60
2005	0.06		-0.07**	0.03	1.07*	0.58
2006	0.06		-0.10***	0.03	1.68***	0.59
2007	0.06		-0.03	0.03	2.25***	0.64
2008	0.06		-0.13***	0.03	0.17	0.68
2009	0.06		-0.27***	0.03	1.55**	0.65
2010	0.06		-0.20***	0.03	0.53	0.54
2011	0.06		-0.20***	0.03	-0.14	0.54
2012	0.08		-0.36***	0.03	0.94	0.66
2013	0.08		-0.40***	0.03	-0.37	0.54
2014	0.08		-0.43***	0.03	-0.68	0.52
2015	0.08		-0.41***	0.03	-0.00	0.67
2016	0.09		-0.45***	0.03	-0.13	0.73
Constant	--		0.60***	0.03	4.30***	1.54
Pseudo log-likelihood	--		-56415.32		--	
R-squared	--		--		0.02	
Rho	--		0.48		0.02	
Individuals	13,707		18,621		13,707	
Observations	65,864		100,255		65,864	

***, ** and * show statistical significance at the 1, 5 and 10 per cent level, respectively. Standard errors are clustered by individual.

(a) Summary statistics shown for the wage regression sample. Standard deviations are shown for continuous variables only; remaining variables are dichotomous.

(b) Dependent variable equals 1 if 100 per cent certain to retain job over next year; 0 otherwise.

(c) Those that changed jobs over the year are omitted. The dependent variable, wage growth, is calculated as the log difference of weekly gross wages and salary divided by usual weekly hours worked in one's main job. It is winsorised at the top and bottom 5 per cent of the distribution for each year.

(d) Self-assessed probability of retaining job over the next year.

(e) Business services include: professional, scientific & technical; financial & insurance; administration & support; rental, hiring & real estate; and information, media & telecommunications. Household services comprise: accommodation & food; education & training; health care & social assistance; arts & recreation; and other services. Goods-production includes: mining; construction; manufacturing; electricity, gas, water & waste. Goods-distribution comprises: transport, postal & warehousing; wholesale trade; and retail trade. Public administration and agriculture is in the 'other' category.

(f) Lagged one year in the wage growth model.

Sources: ABS cat no. 1220.0, ABS cat no. 5209.0.55.001, Frey and Osborne (2017), HILDA Survey Release 16.0, author's calculations

We include time dummies for each survey year, allowing them to capture any non-observed determinants that vary over time and are common across individuals. This can include macroeconomic, technological, legislative and international factors. However, we attempt to isolate labour market slack by exploiting regional variation in the unemployment rate, where unemployment is included for the 13 ABS major statistical regions. The regions are larger than the ideal concept for this analysis of a local labour market. Nonetheless, there remains substantial variation in unemployment over time and by region (given the resources boom and financial crisis periods are included).

The results confirm that labour market conditions are an important influence on job security. Higher unemployment modestly contributes to lower job security, where a 1 percentage point rise in the unemployment rate is associated with a 1 percentage point fall in the probability of feeling certain about job retention (Table 4, Model I, shows marginal effects for select variables). Those who would prefer more hours have a lower expectation of job retention. This suggests underemployment may weigh on feelings of job security, although those wanting to work fewer hours also had low job security.

Trade exposure and automation have the expected significant negative relationship with perceived job security. In terms of marginal effects, an industry with no exposure to international competition has job retention certainty about 3 percentage points higher than one in which imports make up around one-third of an industry's total supply (such as some of the manufacturing industries). After controlling for trade exposure and other factors, those in the goods production and business services sectors still have relatively low perceptions of job security, with those in household services having relatively high security. The average likelihood of feeling certain about job retention falls 3 percentage points when moving from an occupation with no automation risk to one that has the highest likelihood of being automated. Once automation risk is controlled for, being in a routine or manual occupation contributes to higher job security, in contrast to what is suggested by the summary statistics.

Regarding other job and personal characteristics, those that are associated with higher job security include: working full time, having a permanent position, having longer tenure, being female, being older, having less formal education and living outside a capital city.[14] The effect of union membership on job security is not statistically significant. Thus, the difference in average job security perceptions between union and non-union members can be accounted for by their job and personal characteristics. However, there are a number of surprising results. For example, those in non-routine cognitive occupations have relatively low job security and those in regional areas have higher job security. This suggests there may be some issues with the job security measure on some dimensions. Nonetheless, for the most part the results are as expected and particularly so for our key variables of interest.

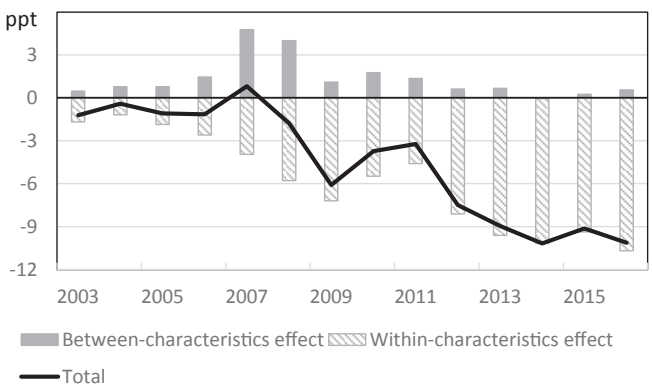
14 Some of these characteristics may be affected by selection bias. For instance, those groups with lower rates of labour market attachment – such as females and older workers – may leave the workforce if they cannot find secure employment, which would bias the results. Selection bias is examined in Section 5.

The time dummy variables are estimated to be increasingly negative in recent years. This suggests that there are aggregate factors weighing on job security that are not being captured by the other variables in the model.

Overall, the model estimates reveal the average propensity to feel secure for each personal, firm or job characteristic over time. However, these feelings of job security can and do change over time. Additionally, the changing composition of the workforce will affect aggregate job security. For instance, an increase in the share of part-time employment would be associated with more people feeling insecure, and so lower overall job security, all else being equal. A shift in the labour market to jobs that are less likely to be automated will increase the share of people feeling secure. As such, we decompose the change in job security over time into these within-characteristic and between-characteristic, or compositional, effects. By estimating the model for each year we can disentangle the roles of the within and between effects on aggregate job security. We conduct a pooled two-fold Oaxaca-Blinder decomposition for the probit model (Jann 2008). We run separate models for each year and compare the change in job security to the 2002 group.

The decline in job security has been broad based across the characteristics we examine in the model (Figure 6). This is highlighted by the persistently negative within-characteristics effect, and confirms the broad-based decline in job security shown by the trends and negative year fixed effects in the model. Compositional change had been working in the opposite direction, with employment moving towards characteristics that generally make workers feel more secure, providing a boost to aggregate job security. The drag from perceived job security within characteristics has been growing since around 2004, with the offset from compositional shifts diminishing over the post-crisis period.

Figure 6: Contributions to annual change in probability of job retention certainty, deviation from 2002



Sources: ABS cat no. 5209.0.55.001, HILDA Survey Release 16.0.

Job security and wage growth

Perceived job security is associated with higher wage growth the following year. Splitting job security into groups from low (0 per cent) to high (100 per cent secure), there is a rough positive relationship with wage growth in the following year (Table 3). A nonparametric K-sample test on the equality of medians shows the difference is statistically significant. However, the fall in wage growth in recent years has been experienced by both secure and less secure workers. And this simple correlation could be a result of a myriad of factors, including simply that higher insecurity is associated with higher unemployment, which may lead to lower wage growth.

Table 3: Median annual wage growth by lagged job security, 2002–16

Probability of Job Retention (%)	*Median Wage Growth (%)*
0–19	3.32
20–39	3.03
40–59	3.73
60–79	4.08
80–99	4.68
100	4.71
Equality of medians test	$\chi^2(5) = 19.5$; p-value=0.00

Notes: Those who changed jobs over the year are omitted.
Sources: HILDA Survey Release 16.0.

We examine this relationship by regressing annual hourly wage growth at time t on self-assessed probability of job retention at t-1. We use the full continuous probability measure of job security. The control variables included are the same as those used in the job security model above. The few differences are the: (1) inclusion of inflation – measured as growth in each capital city's consumer price index – as it should influence nominal wage growth; and (2) omission of those that changed jobs during the year. Therefore this model captures the effect of job security perceptions on wage growth over and above the many factors that can influence job security itself. Most controls in the model are included contemporaneously with the wage growth that was received in that year. However, we take the lag of job security, as our hypothesis is that one's perceptions will influence their wage negotiations in the *coming* year. Furthermore, we want to mitigate endogeneity, in that one's security perceptions are influenced by their annual pay outcomes.

We use random effects estimation, but report a suite of alternative models in Section 5. The Breusch-Pagan LM test recommends random effects over pooled OLS (χ^2= 1098.40, p-value = 0.00).

The results show that a higher perceived likelihood of job retention has a statistically significant positive relationship with wage growth over the following year (Table 1). A 1 percentage point decline in job security is associated with a 0.02 percentage point decline in wage growth. To put this into context, a 1 standard deviation fall in job security (a change of 17 percentage points) is associated with

a 0.28 percentage point decline in annual wage growth (i.e. 1.62 x 0.17). However, this finding explains little of the fall in average wage growth observed in recent years. From 2010 to 2014 (the most recent peak and trough), average self-assessed job retention probability declined by only around 3 percentage points. This explains around 0.05 percentage points of the 1 percentage point fall in average wage growth over the corresponding period.

Automation risk and trade exposure are both found to significantly contribute to lower wage growth. Despite the negative relationships, these measures have not been substantially weighing on aggregate wage growth. This is because, in recent years, employment growth has been shifting away both from industries with a higher trade exposure and occupations that are more easily automated.

Several variables that capture economic and labour market conditions were statistically significant. The coefficient on inflation is significant, and subdued inflation explains a sizeable amount of the fall in average wage growth in recent years. Those who would prefer more hours have lower wage growth in the subsequent year, which suggests an additional role of underemployment in suppressing wage outcomes. Surprisingly, unemployment has a weakly positive relationship with wage growth, though it does have the expected negative effect when the time dummy variables are excluded. Another surprising result is that average wage growth has been much higher for part-time workers than full-time workers (even in the raw data). With the wage level being higher for full-time workers, this suggests some convergence in part-time wages to full-time. This result could be explored further in future work.

5. Robustness checks

An alternative measure of job security

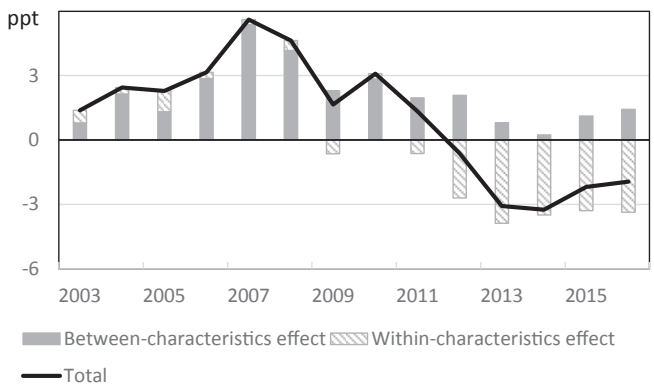
We introduced an alternative self-assessed job security measure in Section 3. It is an ordinal measure of how satisfied a worker is of their job security on a scale of 0–10. It is similarly skewed as the continuous job security measure used in the main analysis, with the vast majority of workers feeling secure. Almost a quarter of workers have the highest job security, 44 per cent chose the next two levels (i.e. 8 or 9 out of 10), close to a quarter chose the middle of the spectrum and only 8 per cent chose the bottom 5 categories. To model this variable, we divide it into three categories: high (8–10), medium (5–7) and low (0–4) job security satisfaction categories. The variable moves closely with the main job retention certainty measure and wage growth over time.

We estimated the determinants of job security satisfaction using a random effects ordered probit. The key results are similar to the main results above.[15] Unemployment, trade exposure and automation have a negative relationship with job security, although the latter is statistically insignificant. Those in routine manual roles had significantly lower perceived job security satisfaction, which is as expected but a different result to that of the job security certainty measure. Like the main model, the year dummies have substantial explanatory power. To explore the role of these unexplained aggregate economic factors further we decomposed the job security satisfaction into within- and between-characteristics factors. For ease of computation,

15 Model results are available on request.

we modelled a simple probit just for the high job security category. Compositional change appears to have had a much larger positive effect prior to 2010, with the model covariates explaining much of the rise in job security in that period (Figure 7). While compositional change continued to positively contribute to aggregate job security, the within-characteristics effect dominates in the latter part of the sample, putting substantial downward pressure on job security (similar to the job retention measure). This underlines the broad-based nature of the fall in job security in recent years.

Figure 7: Contributions to annual change in probability of high job security satisfaction, deviation from 2002



Sources: ABS cat no. 1220.0, ABS cat no. 5209.0.55.001, Frey and Osbourne (2017), HILDA Survey Release 16.0, author's calculations.

Alternative job security models

Our core results come from a random effects probit model, where variation in job security derives from differences between workers and differences over time for the same individual. Examining these various sources of variation allows us to better map the model to what is happening on aggregate in the economy. However, the effect of a particular factor on job security may be better isolated by examining within-individual changes, and so holding all fixed observed and unobserved characteristics constant.

As a fixed-effects probit model provides inconsistent estimates, we estimate a fixed effects linear probability model. For completeness, we also modelled a random effects linear probability model and a random effects model of the continuous probability variable (i.e. the data as presented in Graph 2, panel a). Finally, we re-estimate the job security random effects probit, where we define an individual as secure if they are 90 per cent certain of retaining their job (as opposed to 100 per cent). We refer to this as the 'expanded job certainty dummy'. Table 4 shows the results for key variables.

Table 4: Select determinants of job security – marginal effects

	Probit		*Linear probability model*		*Linear regression*
	(I)	*(II)*	*(III)*	*(IV)*	*(V)*
Dependent variable	Job certainty dummy	Expanded certainty dummy	Job certainty dummy	Job certainty dummy	Continuous probability
Unemployment (by region)	-0.01***	-0.01***	-0.01***	-0.01***	-0.00***
Trade exposure	-0.09***	-0.05***	-0.04*	-0.07***	-0.02**
Automation	-0.03***	-0.02***	-0.01	-0.02***	-0.01
Union member	0.00	0.01*	-0.01	0.00	0.00
Year FE	Yes	Yes	Yes	Yes	Yes
Individual FE	No	No	Yes	No	No
LL / R-Sq.	-56,415.32	-43,949.74	0.03	0.07	0.06
Rho	0.48	0.38	0.48	0.32	0.27
Individuals	18,621	18,621	18,621	18,621	18,621
Observations	100,255	100,255	100,255	100,255	100,255

Sources: ABS cat no. 1220.0, ABS cat no. 5209.0.55.001, Frey and Osbourne (2017), HILDA Survey Release 16.0, author's calculations.

The choice of cut-off for the probability of retaining one's job does not make a material difference to the results (Model II). The marginal effect of trade exposure and automation on job security is a little lower, but sign and statistical significance is the same. The coefficient on union membership becomes statistically significant at the 10 per cent level.

The signs on the coefficient estimates are largely the same for both linear probability models and the probit model (Models III and IV). However, the marginal effects in the fixed effects linear probability model are a bit smaller than in the random effects counterpart, with the former less likely to be statistically significant. Of the key variables of interest, unemployment and trade exposure remain statistically significant. Automation is not significant and, like for the job security satisfaction model, may reflect the limitation in the data in that automatability does not vary for a particular occupation over time. The regression using the continuous probability variable results in smaller marginal effects (Models V). This reflects the lower level of variability in the data.

Alternative wage growth models

Alternative models were estimated to test the robustness of the relationship between wage growth and job security. We examined a range of specifications and samples within the panel framework. We also address the issues of sample selection and

endogeneity. Table 5 presents the coefficient estimates of the lagged job security probability measure on wage growth for various models.

Table 5: Regression coefficients of lagged job security on wage growth[(a)]

	Coefficient	*St. Error*	*R-squared (overall / pseudo)*	*Observations[(b)]*
Main model				
Random effects	**1.62*** **	**(0.52)**	**0.02**	**65,864**
Alternate panel structure				
Pooled OLS	1.60***	(0.51)	0.02	65,864
Fixed effects	1.82***	(0.71)	0.02	65,864
Alternate model structure				
Heckman selection	1.55***	(0.51)	--	64,730
Instrumental variables	1.26	(5.64)	0.02	65,825
Median (quantile) regression	1.75***	(0.30)	0.01	65,864
Alternate specification or sample (random effects)				
Not promoted	1.68***	(0.55)	0.02	58,924
Includes job changers	2.16***	(0.45)	0.02	78,753
Post 2009 data	2.44***	(0.66)	0.02	35,221
No time dummies	1.76***	(0.52)	0.02	65,864
No controls	1.69***	(0.52)	0.00	68,765
Job security dummy variables				
Job retention certainty	0.32*	(0.17)	0.02	65,864
High job security satisfaction	0.58	(0.43)	0.02	66,582
Medium job security satisfaction	0.46	(0.46)	0.02	66,582

(a) Standard errors are clustered by individual. *, **, *** denotes significance at the 10, 5 and 1 per cent levels, respectively.
(b) The number of individuals is 13,707 for most of the models. The number of observations is 93,152 in the selection equation.
Sources: ABS cat no. 1220.0, ABS cat no. 5209.0.55.001, Frey and Osbourne (2017), HILDA Survey Release 16.0, author's calculations.

In the random effects wage growth model, the coefficient on job security was largely unchanged when we did not control for any other factors. This was also the case when we excluded those who were promoted over the year or when we excluded the year dummy variables. The relationship between job security and wage growth was stronger when we model only the post-crisis period or when we include those who changed jobs over the year. The relationship between job security and wage growth also appears robust to outliers, as the effect was similar under a median regression. However, we find a mostly insignificant effect of job security on wage growth when we include either the dummy variable of job retention certainty (i.e. the variable used in the job security determinants regression) or the job security satisfaction dummy variables. This is in line with the results of Dickerson and Green (2012), who find that continuous probability measures of subjective job loss explain labour market

outcomes better than ordinal measures.

Wages and job security are only observed for those in employment. Given a person with low perceived job security is more likely to be unemployed in the following period, their omission from the sample may lead to selection bias. In other words, the characteristics of the unemployed and those with low job security may be correlated. We estimate a Heckman selection model of wage growth to correct for the bias. To estimate the selection equation we need to include variables that will influence the probability of being employed, but be unrelated to annual wage growth. We used one's unemployment history (the number of years they have ever been unemployed), their health status (whether they have self-assessed poor or fair health and whether they have a chronic illness) and an interaction term for being female with a dependent child under five. We also included year dummies, age, education, regional unemployment rate, state, gender, having an English-speaking background, being an immigrant and living in a regional area. Despite the presence of selection effects (ρ=-0.07; χ^2=31.35; p-value =0.00), the estimate and statistical significance of job security was almost identical to the random effects model. This implies that selection bias is not a material concern in our analysis.[16]

The wage growth model may suffer from reverse causality, where the expectation of a large wage rise may make one feel more secure in their job. Alternatively, a common factor, such as the business cycle, may drive both variables (Hübler and Hübler 2006). Although job security enters the model as a lag, this does not necessarily adequately control for endogeneity (Bellemare et al. 2017). As such, we re-estimate a pooled model using instrumental variables. Our instrument is lagged subjective life satisfaction, based on a question that asks "how satisfied are you with your life (on a scale of 0–10)?". This captures how optimistic one feels about life and is expected to be correlated with one's perceptions of their job security, but unlikely to be related to wage growth in the following period.[17]

In the IV regression, we instrument lagged job security with two lagged dummy variables denoting high and medium life satisfaction (which equals 1 for those with life satisfaction of 9 or 10 or between 6 and 8, respectively). The first-stage regression F-statistic is large, so we reject the null of weak instruments (F=130.63; p-value = 0.00). We further test for instrument validity by performing the Wooldridge score test of overidentifying restrictions. We fail to reject the hypothesis of valid instruments (χ^2=0.41; p-value =0.52).

Given the instrument appears valid, we can test for endogeneity of job security with respect to wage growth. Using the Wu–Hausman test we fail to reject the null of exogeneity (F=0.01; p-value =0.91). As such, IV regression will lead to consistent but inefficient estimates, and our main random effects model is more appropriate. This is confirmed with the results of the model. The point estimate of job security on wage

16 We also estimated the job security probit model in this selection framework. The likelihood ratio test suggests selection effects are present (ρ=-0.16; χ^2=23.05; p-value =0.00). Nonetheless, most coefficient estimates of interest had similar sign and magnitude, although the effect of regional unemployment was smaller.

17 Ambrey and Fleming (2012) show personality, demographic factors and current circumstances affect life satisfaction.

growth in the IV regression is smaller to the main model result, it has an enormous standard error. As such, our main model estimates are not biased by endogeneity.

Overall, the wage growth models are robust to various specifications. In some models the results are larger and other smaller. But in each case, the overall conclusion that the role of job security on wage growth is statistically significant but small in economic terms.

6. Conclusion

This paper used household panel data to examine factors – including automation and globalisation risks – that influence workers' perceptions of their own job security. We then examined whether these job security perceptions have contributed to low wage growth in recent years. We find that perceptions of one's job security are lower for those in casual work, areas of high unemployment, jobs with a high risk of automation and industries more exposed to global trade. However, aggregate conditions not captured in our model appear to be the key driver of job security concerns. The recent decline in perceived job security has been broad based across industries, occupations, job structure and personal characteristics. As such, we cannot pinpoint any one factor that has led to the decline in perceptions of job security.

We find that greater perceived job security is associated with higher wage growth, over and above the other factors in the model, but the effect is small. A worker who is only 50 per cent sure that they will keep their job over the next year will have annual wage growth that is around 0.8 percentage points lower than a worker who is certain they will keep their job. As such, weaker job security perceptions have provided a small drag on wage growth in recent years. Furthermore, the relatively small magnitude of the fall in average job security means that the decline in aggregate wage growth is mostly explained by other factors, with the overall effect of job security being trivial. Subdued inflation contributed to more of the decline in wage growth, with most of the fall explained by aggregate factors that are not directly captured in our model.

Our strategy for identifying the effect of job security on wage growth was to use panel data to track whether those with low job security received lower wage rises the next year, controlling for a range of other factors. In theory this arises from lower job security leading to more constrained bargaining by the employee. A key limitation of our work is that wages are determined by the interaction between the firm and employees bargaining strengths, and yet we only observe the final outcome. Future work could look at surveying employees directly about their own bargaining behaviour to get at the issue directly. A more localised measure of unemployment may also improve the model, providing a more targeted and varied measure of regional unemployment.

An extension of our work would be to examine the role of self-assessed job security on hours worked. Part-time employment and underemployment have been increasing, allowing both workers and firms to adjust and bargain over hours. If workers are concerned about their job security they may not only accept lower wage rises, but also less hours, as a trade-off for stable employment.

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