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ABSTRACT

Was the Mid-2000s Drop in the British Job Change Rate Genuine or a Survey Design Effect?*

The year-on-year job change rate fell sharply, from 18% in 2005 to around 13% in 2006, according to British Household Panel Survey (BHPS) estimates. This fall coincides with the introduction of dependent interviewing to the BHPS, intended to reduce measurement error and improve consistency. Estimates from models of job change misclassification rates (Hausman et al., Journal of Econometrics, 1998) show that reduced measurement error cannot account for the fall in the job change rate. This suggests that the fall was genuine.

JEL Classification: J62, C25, C81

Keywords: job change, misclassification error, dependent interviewing, feed forward

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1. Introduction

This paper analyses whether a significant fall in the British year-on-year job change rate between 2005 and 2006 was genuine or due to a major survey design change. Having reliable estimates of job change rates is important because they are a fundamental aspect of labour market fluidity and because there is substantial policy interest in the relationship between job change and outcomes such as wage growth. Although countries such as the USA and Germany have linked employer-employee panels providing consistent series for job change rates, such series do not exist for the UK. Longitudinal surveys remain an important source of information.

The British Household Panel Survey (BHPS) is a premier information source for these topics. However, there was a change in data collection methods from 2006 onwards. Instead of asking about circumstances at each annual interview with no reference to past responses ('independent interviewing', INDI), a new 'dependent interviewing' (DI) approach was applied to multiple questionnaire domains including employment and nature of the job and employer. In each area, substantive information from the previous interview was fed-forward to respondents who were then asked if there was no change. Only if there was a change (or no usable past information), were there follow-up questions about change. Jäckle et al. (2007) provide details and explain that the 'main motivation … was to improve data quality, in particular the longitudinal consistency of responses' (2007: i).

Coincident with the BHPS's introduction of DI was a substantial reduction in the *observed* year-on-year rate at which individuals changed jobs, from 18% in 2005 to 13% in 2006, following a gentle decline from the end of the 1990s until 2005. See Figure 1. (Job change is defined in the next section.) By nature, DI is expected to reduce the prevalence of observed change, so to what extent was the 2005/2006 change a fall in the *true* job change rate or reflect a DI-induced change in the quality of observed measures?

<Figure 1 near here>

Suppose, plausibly, that INDI leads to over-estimates of change (the observed job change rate is greater than the true rate). In this scenario, one would expect DI introduction to decrease the probability of misclassifying true stayers as changers (a fall in the false positive rate, α_0), but it may also increase the probability of misclassifying true changers as stayers (a rise in the false negative rate, α_1) if e.g. respondents anticipate that this reduces their response burden. For DI introduction to account for the fall in the observed job change rate, we need a fall in α_0 that is greater than the rise in α_1.

I find this was not the case, suggesting that the observed fall in the job change rate was genuine. My conclusion is based on estimates from models of misclassification of binary responses fitted to BHPS data.

2. Models and data

I fit Hausman et al. (1998) misclassification models for job change occurrence separately for the pre- and post-DI introduction periods and compare the estimates of α_0 and α_1 from each model. The model relates responses on true job change between years t–1 and t to employee characteristics but there are also misclassification probabilities, α_0 and α_1, as above. The expected value of the observed binary response, y_{it}, for employee i in year t, is:

$$\Pr(y_{it} = 1 \mid X_{it-1}) = \alpha_0 + (1 - \alpha_0 - \alpha_1)F(\beta' X_{it-1}) \tag{1}$$

where X_{it-1} is a vector of characteristics and $F(.)$ is the cumulative normal distribution describing response probabilities if there is no misclassification. Misclassification probabilities depend on the true response but are independent of X_{it-1}. Parameters are identified by the non-linearity of $F(.)$ and a monotonicity condition, $\alpha_0 + \alpha_1 < 1$ (Hausman et al. 1998). With no misclassification, the model reduces to a standard probit regression. With misclassification, the marginal effects of characteristics on the observed response are smaller than those for the true response by a factor $(1 - \alpha_0 - \alpha_1)$. For DI introduction to account for a substantial fraction of the observed fall in job change rate, $(\widehat{\alpha_0} + \widehat{\alpha_1})$ should decrease between periods.

The BHPS began with a nationally representative sample of Great Britain's private household population in Autumn 1991 with adults re-interviewed annually each Autumn through to 2008 (the final survey year). I use data for working-age respondents aged 16–59 years with positive employment earnings at t–1 and t.

Prior to 2006, when the BHPS employed INDI, the job change measure is derived from responses to the following questions about the period between the current interview and the start of the reference year (1 September of the calendar year before the current interview year): 'What was the date you started working in your present position? If you have been promoted or changed grades, please give me the date of that change. Otherwise please give me the date when you started doing the job you are doing now for your present employer'.

From 2006 onwards, with DI introduction, employed respondents were fed forward information about their main job, occupation, and employer characteristics and asked to confirm if they described the current situation. Only if respondents said circumstances had changed was the question about the date of the present position asked. For all waves, a year-to-year job change is recorded if the respondent reports a new position and its start date is between the dates of interview in years *t*–1 and *t*.

The BHPS job change measure differs from those from the Panel Study of Income Dynamics and Current Population Survey used by Hausman et al. (1998), derived from questions about tenure in the year *t* job. The job change definition in the Living in Ireland Survey, a rotating panel using INDI data collection, is based on start dates of new jobs, as in the BHPS. Bergin (2015) reports Irish trends similar to those shown in Figure 1 for 1995–2001 (British levels are higher because they include promotions and grade changes). Neither study examines changes in job change misclassification probabilities over time, as I do.

I define the pre- and post-DI introduction periods as 2003–2005 and 2006–2008 respectively. (Lengthening the pre-DI introduction period does not change findings materially.) Following Hausman et al. (1998), I use a conventional set of covariates: sex, age, highest educational qualification, marital status, health status, region, whether there is a workplace trade union, and (log) earnings. Table 1 displays summary statistics. There is balance in the pre- and post-DI introduction covariates: means are similar in the two periods aside from small increases in average real earnings and the fraction with a degree or other post-A-level qualifications. (A-level exams qualify individuals for university entrance.)

<Table 1 near here>

I fit the misclassification models by maximum likelihood separately for pre- and post-DI periods, pooling the person-year data within each period, including year fixed effects, and clustering the standard errors by person to account for within-panel correlations. Following standard practice, I fit misclassification probabilities in the logit metric to constrain them to the (0,1) interval.

3. Findings

Table 2 shows the parameter estimates for the pre- and post-DI introduction models. Marital status, health problems, and region have no statistically significant association with job change in either period. In both periods, having higher qualifications or not having a workplace trade union are associated with more job change, and older workers are less likely to change jobs. Higher earners and women are more likely to change jobs – but these associations are statistically significant for the first period only.

<Table 2 near here>

At the bottom of Table 2 are estimates of the logits of the misclassification probabilities. Table 3 reports the misclassification probabilities transformed to their natural metric. For both periods, the estimates are small. For the pre-DI period, $\widehat{\alpha_1}$ is effectively zero (hence the lack of standard error estimate) and $\widehat{\alpha_0}$ is zero to 3 d.p. For the second period, $\widehat{\alpha_1}$ is again zero (to 3 d.p.) and $\widehat{\alpha_0}$ is 0.04 though its 95% confidence interval is wide. A specification constraining $\alpha_1 = 0$ provided the same estimates (not shown).

<Table 3 near here>

These estimates differ from those reported in previous research. For the USA in the 1980s, Hausman et al. (1998, Table 7, Partition *T* case) report $\widehat{\alpha_0} = 0.25$ and $\widehat{\alpha_1}$ near zero. For Ireland in the early 1990s, Bergin (2015, Table 10) reports that $\widehat{\alpha_0}$ is near zero and $\widehat{\alpha_1} = 0.51$. The heterogeneity in estimates across studies may arise from differences in data collection methods, country context, or time period. Although the BHPS is renowned as a high-quality household panel study, it is surprising that the BHPS-based estimates of misclassification probabilities are so small as I use the same model as the earlier studies and similar samples of working-age individuals and covariates.

Restricting the estimation sample to household heads, as Hausman et al. (1998) do, does not change the conclusions. I also refitted the models excluding the year fixed effects (estimates not shown), and derived similar results to those shown in Table 2. Although the pre-DI period model did not converge, the iteration log showed that this was because the maximizer was trying to set both misclassification rates to zero. For the post-DI period, $\widehat{\alpha_0}$ was 0.06 (95% CI 0.021, 0.139) and $\widehat{\alpha_1}$ was near zero again.

In sum, if anything, there is a slight increase in $(\widehat{\alpha_0} + \widehat{\alpha_1})$, from 0.0005 to 0.04, not a fall. Thus, there is no evidence that the BHPS's introduction of DI in 2006 reduced measurement error in a manner that explains the sharp fall in the job change rate between 2005 and 2006. (This is reminiscent of Krueger et al.'s (2017) conclusion that DI introduction in the CPS was not responsible for a rise in rotation group bias.)

Instead, arguably the fall in the rate is genuine – the true job change rate fell as well as the observed job change rate. There are no other British data series directly comparable with the BHPS's, but Quarterly Labour Force Survey estimates of job-job flow rates for workers aged 16–69 also show a fall between 2005Q4 and 2006Q4 (Office for National

Statistics 2020). Both series show a secular decline over the decade leading up to 2008, as does the US LEED-based job-job flow series (Hyatt and Spletzer 2013).

If misclassification of year on year job change is negligible, the standard probit model suffices. If I apply the pre-DI probit model's coefficients to the post-DI period sample, the predicted job change rate is 0.17, i.e. the same as the observed pre-DI period rate. However, application of the post-DI probit model's coefficients to the pre-DI sample leads to a predicted job change rate of 0.13 – the observed post-DI rate. Since the covariates in the two samples are balanced, this suggests that the fall in the job change rate is primarily accounted for by the changes in coefficients across periods (most obviously those for female and log(earnings): see Table 2.) Job change propensities were already falling before the Great Recession.

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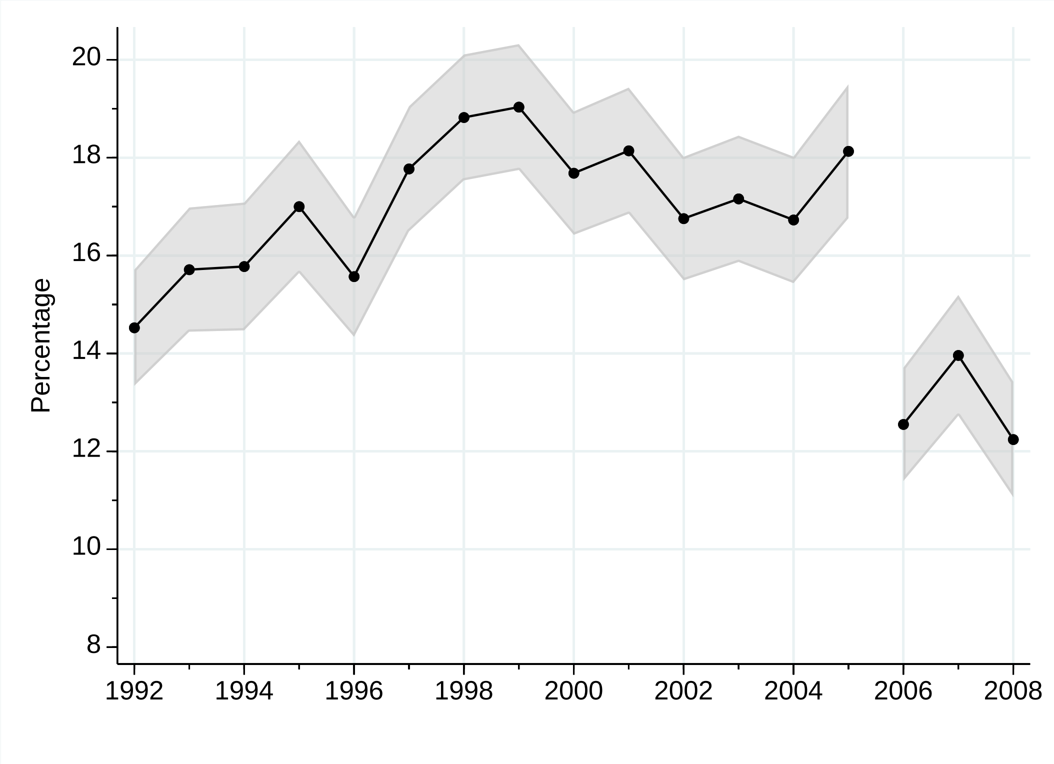
Figure 1. Prevalence of year-on-year job change (%), by year



Notes. The chart shows job change rates between survey year *t*–1 and *t*, for each *t* = 1992–2008, with pointwise 95% CIs. Dependent interviewing was introduced in 2006. Source: BHPS.

Table 1. Covariate means, by pre- and post-DI period

Covariate	Pre-DI period (2003–2005)	Post-DI period (2006–2008)
Female	0.52	0.52
Highest educational qualification		
Less than A-level	0.30	0.25
A-level(s)	0.12	0.12
Degree or other post-A-level qualification	0.58	0.64
Age (years)	38.6	39.1
Marital status		
Married	0.75	0.76
Widowed, divorced, or separated	0.08	0.08
Single never-married	0.16	0.16
Has health problems	0.53	0.51
Region of residence		
London or rest of South-East England	0.08	0.07
Rest of England	0.78	0.79
Wales	0.05	0.06
Scotland	0.09	0.08
No trade union in workplace	0.47	0.47
Log(gross monthly earnings, 2011 prices)	7.43	7.49
No. person-years	9,545	9,149

Source: BHPS.

Table 2. Coefficient estimates, models of year-on-year job change allowing for misclassification, by pre- and post-DI introduction period

	Pre-DI (2003–2005)		Post-DI (2006–2008)	
Female	0.1109	***	0.0293	
	(0.0372)		(0.0575)	
A-level(s)	0.0668		0.1399	
	(0.0563)		(0.0984)	
Degree or other post-A-level qualification	0.1576	***	0.2078	*
	(0.0398)		(0.1192)	
Age (years)	–0.0195	***	–0.0308	**
	(0.0018)		(0.0137)	
Widowed, divorced, or separated	0.0607		0.1214	
	(0.0607)		(0.1076)	
Single never-married	0.0398		0.0032	
	(0.0470)		(0.0768)	
Has health problems	0.0051		0.0426	
	(0.0321)		(0.0557)	
Rest of England	0.0488		–0.0576	
	(0.0629)		(0.1260)	
Wales	–0.0033		–0.0977	
	(0.0988)		(0.1955)	
Scotland	–0.0442		–0.0219	
	(0.0822)		(0.1265)	
No trade union in workplace	0.0659	**	0.1414	**
	(0.0330)		(0.0668)	
log(gross monthly earnings, 2011 prices)	0.0880	***	–0.0086	
	(0.0274)		(0.0513)	
Year = 2004 (pre-DI) / 2007 (post-DI)	–0.0207		0.0805	
	(0.0360)		(0.0521)	
Year = 2005 (pre-DI) / 2008 (post-DI)	0.0341		–0.0107	
	(0.0367)		(0.0580)	
Constant	1.1102	***	–0.3618	
	(0.2361)		(0.5453)	
logit(α_0)	–7.6302	***	–3.1761	***
	(0.4342)		(1.0885)	
logit(α_1)	–99.3086		–9.1083	***
	(.)		(1.4403)	
Log pseudo-likelihood	–4283.4877		–3414.8807	
Mean of dependent variable	0.17		0.13	
No. persons	4,003		3,820	
No. person-years	9,535		9,149	

Notes. Table shows estimated coefficients and cluster-robust standard errors (clustering by person) in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The reference categories for the categorical variables are: has educational qualifications less than A-level standard, marital status is married or living as married, region is London and the South East of England, year is 2003 (pre-DI period) or 2006 (post-DI period). Source: BHPS.

Table 3. Estimates of misclassification probabilities

Probability	Pre-DI		Post-DI	
	Estimate	95% CI	Estimate	95% CI
α_0	0.0005	[0.0002, 0.0011]	0.0401	[0.0049, 0.2606]
α_1	0	–	0.0001	[0.00001, 0.0018]

Notes. α_0 is the false positive rate; α_1 is the false negative rate. Source: Table 2 estimates.