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Working Time Mismatch and Subjective Well-Being

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Abstract

This study uses nationally representative panel survey data for Australia to identify the role played by mismatches between hours actually worked and working time preferences in contributing to reported levels of job and life satisfaction. Three main conclusions emerge. First, it is not the number of hours worked that matters for subjective well-being, but working time mismatch. Second, overemployment is a more serious problem than is underemployment. Third, while the magnitude of the impact of overemployment may seem small in absolute terms, relative to other variables, such as disability, the effect is quite large.

1. Introduction

Recent decades have witnessed noticeable changes in the distribution of working hours in many industrialised nations. In particular, while the average length of the work week has either continued to decline or remained largely unchanged, variation around the average has mostly increased. In the US, for example, Jacobs and Gerson (2004) reported on data from the Current Population Survey that reveal a marked decline between 1970 and 2000 in the proportion of employees reporting working a traditional 40-hour work week, and noticeable increases in the proportion of employees reporting either relatively short work weeks (less than 30 hours) or relatively long work weeks (50 hours or more). Similar trends have been reported in Australia (Wooden 2002; Wooden and Drago 2007), Canada (Sheridan *et al.* 2001), Japan (Japanese Ministry of Health, Labour and Welfare 2004), New Zealand (Callister 2005) and the UK (Green 2001). Even in continental Europe, the variability of weekly hours during the 1990s led the OECD (2004: 40) to conclude that the evidence "is suggestive of an overall trend towards greater diversification of weekly work schedules".

Such trends have been accompanied by rising levels of concern among researchers about the impacts that working time arrangements might be having on workers and their families. In particular, there is now a sizeable literature documenting the effects of regular exposure to extended work schedules on worker health and injury. Very differently, others have pointed to the potential for long working time regimes to adversely affect relationships within the home, and ultimately contribute to marital breakdown and adverse outcomes for children. At the other end of the working hours spectrum, part-time employment has also long been a subject of research interest, with many arguing that part-time jobs often provide insufficient hours to satisfy the needs and desires of workers.

Central to the latter argument, at least, is the premise that it is not the number of hours worked *per se* that matters, but whether those hours are in line with workers' preferences. In textbook models of labour supply, this distinction is irrelevant; individuals are assumed to freely choose the combination of work hours and non-work hours that maximises their personal utility subject to time and budget constraints. Actual hours worked are thus a direct reflection of preferences, and any mismatch should be only temporary. There is, however, a growing body of survey evidence that suggests that work time mismatches are common and found in many countries (e.g., Bell and Freeman 2001; Jacobs and Gerson 2004; Lee 2004; Reynolds 2004; Reynolds and Aletraris 2006; Stier and Lewin-Epstein 2003).

Despite this, relatively little research has taken working hours preferences into account when examining relationships between hours worked and outcome variables. This is the central feature of the study reported on here. Specifically, panel data from the first five waves of the Household, Income and Labour Dynamics in Australia (HILDA) Survey are used to test for relationships between the number of usual weekly hours of work and subjective measures of worker well-being. The HILDA Survey is relatively unusual in that it collects information on preferred hours of work conditional on any consequences for earnings. We are therefore able to identify workers who are underemployed (in the sense that preferred hours exceed hours usually worked) or overemployed (usual hours exceed preferred hours). Our hypothesis is that any adverse impacts of either short or long hours of work should be most prominent when those working hours are inconsistent with preferences.

A second distinguishing feature of the analysis is the use of panel data, and hence the ability to better control for unobserved worker heterogeneity. Most previous research into the effects of working hours arrangements has used cross-section data, and so findings may be sensitive to the availability and choice of control variables. In contrast, with panel data we are able to employ methods that effectively control for all worker characteristics that are time invariant, or, at least, do not vary much over the period under consideration. Panel data also provide the opportunity to identify the causal order of events. The role of timing is potentially important given the possibility that working time mismatch, and especially underemployment, could be the result (as well as a cause) of low levels of well-being.

Finally, the data are drawn from a large, nationally representative population sample. This stands in contrast to much of the earlier research, and especially studies investigating the consequences of long hours, which mostly employ data covering a small sub-group of the workforce, often taken from a single employer or occupation.

2. Previous research

2.1. Part-time work and underemployment

There are a number of distinct literatures that are relevant to the study of working time mismatch and its relationship to worker well-being. First, there is a long-standing literature concerned with the quality of part-time work. This body of research has been concerned predominantly with identifying whether part-time jobs should, on the basis of objective characteristics, be classified as good jobs or bad jobs. Until recently, most studies appeared to support the hypothesis that part-time jobs were generally of low quality, as reflected in relatively poor remuneration levels, the absence of fringe benefits, low levels of job security, and the lack of opportunities for career development (e.g., Blank 1990; Kalleberg *et al.* 2000; McGovern *et al.* 2005; Tilly 1996).

Recent research, however, suggests that much of the measured wage penalty for part-time employment disappears once differences in job and worker characteristics are controlled for. Hirsch (2005), for example, used micro-data from the US Current Population Survey and found that after controlling for measurable characteristics, and especially occupational skill requirements, much of the part-time wage differential disappeared. Indeed, for women, who account for the majority of part-time workers, the gap is almost entirely eliminated. Manning and Petrongolo (2004) reported very similar findings in an analysis of British Labour Force Survey data. Further, they found that the wage penalty becomes a premium once all time invariant worker and job characteristics are held constant. Using panel data from Australia, Booth and Wood (forthcoming) also report the same strong conclusion.

Very differently, many studies that employ subjective measures of job satisfaction have been unable to detect sizeable negative associations with part-time work (e.g., Bardasi and Francesconi 2004; Blanchflower and Oswald 1998; D'Addio *et al.* 2007; Kaiser 2002; Manning and Petrongolo 2004; Wooden and Warren 2004); indeed positive associations are often found. Such findings, perhaps, should not be surprising, given the widespread survey evidence that most part-time employees appear to prefer part-time hours.[1] Indeed, it is often argued that part-time work is especially attractive to women, and tends to be consistent with their own attitudes and preferences (e.g., Fortin 2005; Hakim 2000, 2002). Despite this, relatively few studies have given much consideration to the role that preferences might play in mediating the relationship between part-time work and subjective measures of job quality. Those studies that have (e.g., Dooley *et al.* 2000; Dooley and Prause 1998; Manning and Petrongolo 2004; Prause and Dooley 1997; Wilkins 2007) typically find that adverse outcomes, in the form of low job satisfaction, elevated levels of depression, or relatively poor self-esteem, are concentrated among underemployed part-time workers. Where part-time hours are consistent with worker preferences, the outcomes are generally no different than those for full-time workers.

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[1] Data compiled by the OECD, for example, reveal that involuntary part-time employment (defined as part-time workers who would prefer to work full-time hours) as a share of total of part-time employment averaged just over 15 per cent across OECD member nations in 2004 (OECD 2006: 42).

Nevertheless, it would be incorrect to think that research is unanimous in concluding that underemployment among part-time employees is consistently associated with negative outcomes. Friedland and Price (2003), in what we believe to be the only other study of hours-based underemployment that uses longitudinal data, found very little evidence to support the claim that underemployed workers are significantly worse off as opposed to otherwise comparable fully employed workers. Using data from the first two waves of the Americans' Changing Lives study, they estimated regression models using four different subjective well-being measures: life satisfaction; depression symptoms; positive self-concept (or self-esteem); and job satisfaction. A distinctive feature of their analysis was the inclusion of a lagged dependent variable in the list of controls in an effort to control for reverse causation. A weakly significant negative relationship was found with self-esteem; estimated relationships with both life satisfaction and depression were insignificant; while levels of job satisfaction were actually found to be significantly higher among the underemployed (presumably due to the absence of sufficient controls for individual characteristics).

2.2. Long hours and overemployment

While part-time employment is a pervasive feature of labour markets in many developed countries (and especially Australia, the UK, and northern Europe), in recent times it has been overshadowed by the growing proportions of employees reporting working hours well in excess of what once would have been considered the traditional norm or standard. Such trends have been accompanied by an increasing number of studies examining the impacts of extended work hours arrangements. These studies typically begin with the hypothesis that extended work schedules, by increasing fatigue levels, reducing the time available for recovery, and inducing unhealthy behaviour (e.g., smoking, poor diet, and lack of exercise), are harmful for worker health, and increase the risk of work-related injury. The reviews of this literature, which are numerous (e.g., Caruso 2006; Harrington 2001; Sparks *et al.* 1997;

Spurgeon *et al.* 1997; van der Hulst 2003), while typically highlighting the evidence demonstrating the link between long work hours and adverse health outcomes, are all forced to admit that the body of evidence is inconclusive, and that the magnitude of estimated associations is likely to depend critically both on sample selection and the extent to which confounding influences are controlled for.

Many of the stronger associations, for example, are found in populations working irregular hours or shifts, and thus it may be the timing of work hours, rather than their number, that is most critical to health outcomes (see Presser 2003). More generally, it is well recognised that long work hours are likely to co-vary with other job characteristics (Barnett 1998; Spurgeon *et al.* 1997; van der Hulst 2003), such as physical job demands, job control and autonomy, job complexity, and social support, as well as a range of other individual and personal characteristics. Identifying the impact of long work hours thus requires isolating the effects of these confounding influences, something that most studies employing cross-section designs have done with only limited success.

Most studies also employ quite small samples, typically covering workers at a single firm or employed in a specific occupation. Van der Hulst (2003), for example, reviewed 27 empirical studies, only four of which involved samples drawn from a range of employers or occupations and, of these, only two explicitly identified persons working long hours per week. It is thus difficult to know to what extent the results obtained from such studies can be generalised to the wider population, especially if the samples were selected because of the high incidence of long work hours. The evidence from the few existing studies employing nationally representative samples suggests that extended work schedules are associated with an elevated risk of workplace accidents or work-related injury, and moreover, that increased risk is not merely the result of spending more time exposed to risk (Dembe *et al.* 2005; Hänecke *et al.* 1998). Associations with other health outcomes are less consistent. Grosch *et* *al.* (2006), for example, used data from the 2002 General Social Survey in the US and found that the relative risk of adverse health outcomes was only noticeably higher among workers reporting very long hours – 70 or more per week – who account for a relatively small fraction of the population (just five percent in their sample). For persons working 49 to 69 hours a week, the relative risk was no higher than for those working standard hours (35 to 40 hours). Furthermore, the likelihood of being dissatisfied with a job was actually found to be highest for those working standard work weeks and lowest among those working the longest hours. That said, as in much of this literature, only a handful of controls were employed.

In a related but separate literature, it is argued that long hours of work both damage relationships within the home, especially marital relationships, and inhibit child development (e.g., Hochschild 1997; Cooper 1999; Relationships Forum Australia 2007). Such arguments flow from the assumptions that increased working time must come at the expense of time spent with family, and that time spent interacting with spouses and children is critically important to the quality of marital relationships and child behaviour and development, respectively. Despite this, the empirical research on the relationship between long working hours and family-related outcomes has failed to produce a consensus, leading Barnett to conclude that "the assumption that long work hours inevitably give rise to work / family conflict … is strongly challenged" (1998: 132). Recent evidence provides little reason to revise this conclusion. On the one hand, studies consistently report significant positive associations between working time and subjective measures of time-based conflict in the home (e.g., Grzywacz and Marks 2000; Major *et al.* 2002; Voydanoff 2004, 2005). On the other hand, research has been unable to find convincing evidence that such time pressures have impacted negatively the quality or stability of relationships (e.g., Crouter *et al.* 2001; Poortman 2005). Indeed, in Poortman's Dutch study, the probability of marital divorce is found to fall with the number of hours worked by the husband.

Finally, and as previously observed by Barnett (1998), rarely is any consideration given to whether the hypothesised adverse impacts of long hours might vary with the extent to which the hours worked are consistent with individual preferences. A notable exception is again Friedland and Price (2003), who analysed the impact of both overemployment and underemployment on their different subjective measures of well-being. Among the overemployed, the only evidence of any significant negative relationships was found for job satisfaction. For the other three outcome measures (life satisfaction, depression symptoms, and self-esteem), the overemployed fared better than the fully employed control group, and in the case of depression symptoms, significantly so.

3. Modelling subjective well-being

This study contributes to the literature by using nationally representative panel survey data to identify the role played by mismatches between hours actually worked and working time preferences in explaining variations across individuals in reported levels of job and life satisfaction.

Employed persons are categorised into discrete groups based on the number of hours usually worked each week, and whether those hours are below, match, or exceed stated preferences. These categories can then be correlated with our two subjective outcome measures to identify patterns of associations. However, given that the association between working time mismatch and subjective well-being is likely to be affected by a great many other intervening influences, the main focus of our analysis is on the estimation of multivariate regression models that attempt to hold constant these influences.

We begin with a simple model that pools data over the different survey years, as follows:

$$SWB_{it} = \alpha + X_{it}\beta + Z_{it}\gamma + \varepsilon_{it} \qquad (1)$$

$$i = 1, \ldots, N; \;\; t = 1, \ldots, T$$

where SWB_{it} is the level of well-being reported by individual i at time t, X_{it} is a measure of the interaction between hours usually worked and the presence and direction of mismatch between hours worked and preferred hours, Z_{it} is a vector of other individual-level time-varying covariates thought to influence well-being, and ε_{it} is an error term.

There are, however, many other potential influences on well-being that are individual-specific, and do not vary much over time. Furthermore, many of these characteristics of jobs and individuals will not be observable in most survey data collections (e.g., personality). We can, however, make use of the panel nature of the data to hold constant these influences. That is, we estimate the following fixed-effects model:

$$SWB_{it} = \mu_i + X_{it}\beta + Z_{it}\gamma + \varepsilon_{it} \quad (2)$$

where the μ_i terms are individual-specific constants.

Finally, following Friedland and Price (2003), we utilise the time dimension of the data, and include a one-period lagged value of the dependent variable:

$$SWB_{it} = \mu_i + X_{it}\beta + Z_{it}\gamma + SWB_{it-1}\delta + \varepsilon_{it} \quad (3)$$

We do this in an attempt to control for state dependence and the possibility that working time mismatch could be both the result and cause of low levels of subjective well-being. The latter effect is certainly plausible in the case of underemployment, with persons in depressed states being less attractive to employers, and thus more likely to have difficulties in securing jobs with attributes that satisfy their preferences. Note that in contrast to Friedland and Price (2003), our fixed effects specification helps to ensure that the estimated state dependence parameter will be net of likely strong correlations with unobservable personal characteristics (such as personality). Note further that, with the fixed effects specification, we have also dealt with the initial conditions problem that often affects regression estimates in the presence of a lagged dependent variable;[2] the initial condition is embodied in the individual fixed effects.

In this study, the regression models are all estimated with least squares methods, which requires the assumption of cardinality, and, strictly speaking, the outcome variables analysed here are not cardinal. A textbook approach would require that models where the dependent variable comprises a series of discrete values which are only ordinal be estimated using the ordered probit (or logit) framework. Maximum likelihood estimation of such non-linear models, however, gives rise to results that can be difficult to interpret. More importantly, when combined with fixed effects, the estimated coefficients will be inconsistent, especially when T is relatively small, as is the case here (see Lancaster 2000). The simpler least squares methods are thus very attractive. Further, as has been demonstrated using life satisfaction data from the German Socio-Economic Panel (GSOEP), the assumption of cardinality is relatively unimportant when analysing the type of dependent variables being used here, and instead what matters most "is how one takes account of time-invariant unobserved factors" (Ferrer-i-Carbonell and Frijters 2004: 655).

4. Data and variable construction

4.1. Sample

The data used in this study are drawn from Release 5.1 of the HILDA Survey data. Described in more detail in Wooden and Watson (2007), the HILDA Survey is a household panel survey with a focus on work, income, and family. Its design is closely modeled on the British Household Panel Survey (BHPS). Release 5.1 provides data covering the first five years (or waves) of data collection.

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[2] The introduction of a lagged variable into a conventional regression model typically gives rise to biased or inconsistent estimates, as a result of the 'initial conditions problem'; if outcomes this period depend on outcomes in the previous period, then it also follows that outcomes this period will be a function of outcomes at some much earlier unobserved, or 'initial', period.

The survey commenced in 2001 with a national probability sample of Australian households. Personal interviews were completed at 7,682 of the 11,693 households identified as in scope for wave 1, and while non-response is considerable, the characteristics of the sample appear to match the broader population quite well. The main weaknesses of the initial unweighted sample are a slight over-representation of females, and an under-representation of both immigrants from a non-English-speaking background and residents of Australia's largest city, Sydney (see Wooden *et al.* 2002).

The members of these participating households form the basis of the panel pursued in the subsequent waves of interviews, which are conducted approximately one year apart. Interviews are conducted with all adults (defined as persons aged 15 years or older on the 30th June preceding the interview date) who are members of the original sample, as well as any other adults who, in later waves, are residing with an original sample member. Re-interview rates are reasonably high, rising from 87 percent in wave 2 to over 94 percent in wave 5.

The sample used here begins with the unbalanced panel, a dataset comprising a total of 64,905 observations from 17,375 people. For some of our analyses, however, the sample is restricted to persons in paid employment at the time of interview, which reduces the number of cases available for analysis to 40,673 (covering 12,367 individuals).

4.2. Measuring well-being

Two main outcome variables are used in this analysis. These variables measure satisfaction with the job and with life, respectively. Both are constructed from responses to a single item scored on a zero to 10 scale, with a score of zero labeled and described as "totally dissatisfied" and a score of 10 labeled and described as "totally satisfied". The life satisfaction question is asked of all respondents, while the question on job satisfaction is only asked of persons in paid employment at the time of interview.

While single-item scales are generally regarded as inferior to multi-item scales, they are now routinely included in large national and international surveys, and have formed the basis for a growing number of studies of both job satisfaction (e.g., Bardasi and Francesconi 2004; Blanchflower and Oswald 1998; Clark 1996; D'Addio *et al.* 2007; Heywood *et al.* 2002; Vieira *et al.* 2005; Wooden and Warren 2005) and life satisfaction (e.g., Bardasi and Francesconi 2004; Carroll 2007; Clark *et al.* 2001; Di Tella *et al.* 2003; Frijters *et al.* 2004, 2006). Indeed, the life satisfaction question included in the HILDA Survey is identical to one included every year in the GSOEP, while the wording of the question on job satisfaction is identical to a question included in the BHPS (though the BHPS employs a seven-point response scale rather than the 11-point scale used in both the HILDA Survey and the GSOEP).

4.3. Measuring working hours and working time mismatch

Actual working hours are based on self-reports of the numbers of hours usually worked each week, including any paid or unpaid overtime and including work undertaken in both the workplace and at home in all jobs. If respondents indicated that the number of hours varied from week to week, they were asked to provide the number of hours per week worked on average over a usual four-week period.

Respondents were then asked about their preferred hours. Specifically, they were asked the following:

> "If you could choose the number of hours you work each week, *and taking into account how that would affect your income*, would you prefer to work … fewer hours than you do now? about the same hours as you do now? or more hours than you do now?"

A further question yielded the precise number of preferred hours for respondents with preferences for fewer or more hours. For respondents claiming to prefer "about the same hours", preferred hours are assumed to equal usual hours.

The question on working time preferences is similar to one regularly included in the BHPS (see Böheim and Taylor 2004). In the BHPS, however, respondents are asked to condition preferences on the assumption that their hourly wage would remain the same as currently earned. This seems restrictive, given that many persons working quite long hours might expect minimal reductions in total current earnings following a reduction in hours. Further, it also assumes that workers are unable to work additional overtime hours for a wage premium. The HILDA Survey question relaxes these restrictions.

Descriptive data on the incidence of working time mismatches cross-classified by both the number of usual weekly hours of work and sex are presented in Table 1. The figures are population-weighted averages of the cross-section data collected over the first five survey waves, and cover the entire employed workforce. As can be seen, the incidence of mismatch is considerable. Almost 30 percent of employed men are overemployed and 15 percent are underemployed. For women the respective percentages are 25 and 18 percent.

4.4. Time-varying covariates

For the regression analyses, we opted for a parsimonious set of time-varying controls. Further, we restricted the list to variables that could sensibly be included in models of both life satisfaction and job satisfaction. The final list included controls for labour force status, age (specified as a quadratic), disability, marital (or more strictly, relationship) status, the number of dependent children aged less than 15 years, equivalised disposable household income, and whether any other adult was present at the time of interview (which is expected to upwardly bias responses on life satisfaction due to social desirability effects). We account for gender differences by estimating separate equations for men and women.

TABLE 1

Working Time Mismatch by Usual Weekly Hours Worked: Average, 2001 to 2005, HILDA Survey (% of employed workforce)

Weekly hours usually worked	*Type of working time match*			*% distribution*
	Over-employed	*Matched*	*Under-employed*	
Men				
<20	*	51.3	44.9	8.3
21-34	7.3	52.4	40.3	8.7
35-40	17.3	67.2	15.5	34.2
41-49	34.0	57.8	8.2	18.6
50+	54.3	42.5	3.1	30.1
Sub-total	29.5	55.4	15.0	100.0
Women				
<20	3.9	59.0	37.1	24.3
21-34	13.6	61.9	24.5	25.5
35-40	31.7	61.7	6.5	30.6
41-49	49.7	47.9	*	10.1
50+	64.3	34.9	*	9.5
Sub-total	25.3	57.1	17.6	100.0

Notes: All figures are weighted population estimates.
* denotes a cell where the sample size is too small to generate a reliable population estimate.

The construction of most variables is straightforward. There are, however, two exceptions – disability status and household income. A disability is defined as any long-term health condition or disability that has lasted, or is expected to last, six months or more. Further, and following other users of the HILDA Survey data (e.g., Headey and Wooden 2004; Shields *et al.* forthcoming), disabled persons are classified into three sub-groups according to the severity of the disability based on how much it limited the type of work that could be undertaken. Those who could not work at all were defined as severely disabled; those whose disability had no impact on their ability to work were classified as mildly disabled; while all other disabled persons were classified as having a moderate disability.

As noted above, the income variable is equivalised disposable household income, and covers the financial year (the year ended 30 June) prior to interview. The equivalence scale used is that regularly used by the OECD, and assigns a weight of 1.0 for the first adult in the household, 0.5 for every other adult, and 0.3 for every child. The income tax payable, as well as some income transfers received (e.g., Family Tax Benefit and Maternity Allowance), were derived by the HILDA Survey team at the Melbourne Institute of Applied Economic and Social Research, thus enabling a measure of disposable income to be derived. Finally, since self-reported income data are known to be heavily affected by item non-response (and the HILDA Survey is no exception), the variable used here assigns imputed values for any missing cases. As with the derivation of taxes, the imputation of missing values was undertaken by the HILDA Survey team. Details of both the tax model and imputation procedures used in the construction of the income variables can be found in Goode and Watson (2006).

Descriptive statistics for all variables included in the regression analysis are provided in Table 2.

5. Results

5.1. Bivariate associations

Table 3 presents mean scores for our two outcome variables, cross-tabulated by hours usually worked per week and the type of working time match. Note that due to small cell sizes, we merge some of the less common mismatch categories (e.g., part-time workers who prefer fewer hours) into the matched group.

The table reveals that within each working hours category, average satisfaction levels are lower for those who report work hours mismatch. This gap in job satisfaction scores between matched and mismatched workers averages about 0.7 to 0.8 of a point. Quality employment, however, is only one among many elements that contributes to overall life satisfaction, and

TABLE 2

Variables: Descriptive Statistics (unweighted)

Variable	*Men*		*Women*	
	Mean	*Std. dev.*	*Mean*	*Std. dev.*
Dependent variables				
Life satisfaction	7.877	1.562	7.982	1.574
Job satisfaction	7.569	1.803	7.726	1.831
Explanatory variables				
Labour force status				
Not in labour force	0.262	0.440	0.413	0.492
Unemployed	0.043	0.202	0.033	0.179
<35 hours: underemployed	0.050	0.218	0.084	0.278
<35 hours: matched / overemployed	0.069	0.253	0.196	0.397
35-40 hours: underemployed	0.034	0.181	0.010	0.100
35-40 hours: matched [control group]	0.145	0.352	0.096	0.295
35-40 hours: overemployed	0.039	0.194	0.052	0.222
41-49 hours: matched / underemployed	0.086	0.280	0.029	0.167
41-49 hours: overemployed	0.045	0.208	0.029	0.167
50+ hours: matched / underemployed	0.104	0.305	0.021	0.142
50+ hours: overemployed	0.125	0.331	0.037	0.189
Age (years)	43.41	17.82	44.16	18.23
Age squared	2201.80	1685.97	2282.43	1772.92
Mild disability	0.080	0.271	0.068	0.252
Moderate disability	0.170	0.376	0.170	0.376
Severe disability	0.016	0.127	0.011	0.105
Partnered	0.579	0.494	0.555	0.497
Number of dependent children	0.469	0.928	0.572	1.004
Household income ($)	31867	22506	29772.4	21031.6
Others present during interview	0.408	0.491	0.349	0.477

Note: With the exception of job satisfaction, statistics apply to the unbalanced panel of all persons. For job satisfaction, the population is restricted to the unbalanced panel of employed persons.

hence we expect, and find, a smaller gap – about 0.4 of a point – when we move to life satisfaction. Interestingly, when we focus only on people who are working their preferred hours, we can see no evidence that either mean job satisfaction or mean life satisfaction scores vary much with the number of hours worked. The simple bivariate associations thus suggest that it is not the number of hours worked that matter for subjective well-being, but whether those worked are consistent with preferences.

TABLE 3

Mean Levels of Satisfaction by Working Time Mismatch and Sex: Average, 2001 to 2005, HILDA Survey

Hours usually worked per week / Type of working time match	*Job satisfaction*			*Life satisfaction*		
	Men	*Women*	*Persons*	*Men*	*Women*	*Persons*
Less than 35 hours						
Underemployed	7.0	7.5	7.3	7.6	7.8	7.7
Matched or overemployed	7.9	8.0	8.0	8.2	8.1	8.1
35-40 hours						
Underemployed	7.3	7.5	7.4	7.6	7.5	7.6
Matched	7.7	7.9	7.8	8.0	7.9	7.9
Overemployed	7.0	7.2	7.1	7.7	7.7	7.7
41-49 hours						
Matched or underemployed	7.8	8.1	7.9	8.0	8.0	8.0
Overemployed	7.1	7.2	7.1	7.6	7.7	7.6
50 hours or more						
Matched or underemployed	7.9	8.1	8.0	8.0	8.0	8.0
Prefers fewer hours	7.2	7.2	7.2	7.6	7.7	7.6

Note: All figures are weighted population estimates.

Note also that the pattern of differences is very similar for both men and women. The only obvious gender difference occurs among the part-time underemployed; women are not nearly as dissatisfied as men in this same situation. This result fits the claim that modern understandings of masculinity involve serving as the primary breadwinner in the household (e.g., Williams 2000).

5.2. Multivariate analysis

The results of the regression analyses are presented in Tables 4 and 5. In all equations, the overall explanatory power of the main covariates is extremely low. This finding likely reflects the fact that much of the variation in self-reported well-being is either inherently difficult to predict (e.g., because it varies with mood), or is influenced by other traits that are difficult to measure and hence observe (e.g., personality). The latter types of influence are held constant in the fixed effects specification, and in all of those models the variance accounted for by time invariant individual specific factors (measured by the estimated rho parameter) is much higher, ranging from 62 to 72 percent in the case of men, and from 55 to 68 percent in the case of women.

While the estimated coefficients on the control variables are not central to this study, one result deserves comment. Consistent with most other research studies employing large samples, our simple pooled data least squares regression suggests that household income is positively and significantly associated with life satisfaction. Once we hold constant all individual characteristics that are constant (or, at least, changed very little over the five-year data window), the size of this coefficient declines to zero for men, and becomes much smaller for women (though still retaining statistical significance). Such findings cast doubt on the claims advanced by Headey and Wooden (2004), among others, that money makes people happier. That said, the present study does not include measures of wealth or consumption, which, together with income, may provide a more complete picture of the relevant economic circumstances of individuals and of their families (see Headey, Muffels and Wooden forthcoming).[3]

Turning to the variables of most interest to this study, the first numeric column in Table 4 reports coefficients on the working hours variables. These are consistent with the mean scores for men reported in Table 3. It can be seen that among men whose usual working hours are matched to their preferences, there is very little difference in job satisfaction scores, and, if anything, those working the longest hours are most content. Job satisfaction, however, is lower among men who are either overemployed or underemployed and working part-time hours. Moreover, for workers in each of these groups, the size of the satisfaction penalty is

[3] In the HILDA Survey, detailed data on household wealth is only collected every four years starting with wave 2, while detailed data on household expenditure was only collected for the first time in wave 5.

TABLE 4

Regression Estimates: Men

	Job Satisfaction			*Life Satisfaction*		
	No fixed effects	*Fixed effects*		*No fixed effects*	*Fixed effects*	
Employment status:						
Not in labour force				-.115*	-.136**	-.102*
				(.047)	(.043)	(.051)
Unemployed				-.450**	-.306**	-.311**
				(.066)	(.051)	(.061)
<35 hours: underemployed	-.706**	-.446**	-.402**	-.240**	-.127**	-.071
	(.072)	(.066)	(.086)	(.055)	(.046)	(.054)
<35 hours: matched / overemployed	.043	.055	.060	.128**	.012	.018
	(.059)	(.065)	(.082)	(.045)	(.045)	(.051)
35-40 hours: underemployed	-.318**	-.148*	-.098	-.325**	-.082	.001
	(.080)	(.065)	(.081)	(.065)	(.049)	(.058)
35-40 hours: overemployed	-.693**	-.288**	-.307**	-.252**	-.104*	-.070
	(.066)	(.061)	(.070)	(.055)	(.046)	(.052)
41-49 hours: matched / underemployed	.074	.038	.048	-.023	-.016	.024
	(.047)	(.046)	(.055)	(.037)	(.035)	(.041)
41-49 hours: overemployed	-.677**	-.362**	-.280**	-.358**	-.163**	-.137**
	(.061)	(.058)	(.067)	(.047)	(.044)	(.050)
50+ hours: matched / underemployed	.177**	.091	.036	-.016	-.071	-.009
	(.048)	(.050)	(.059)	(.040)	(.038)	(.044)
50+ hours: overemployed	-.597**	-.401**	-.399**	-.378**	-.228**	-.226**
	(.051)	(.050)	(.059)	(.039)	(.038)	(.044)
Satisfaction at t-1			-.136**			-.177**
			(.010)			(.008)
Age	-.067**	-.033	-.020	-.070**	-.091**	-.102**
	(.009)	(.025)	(.038)	(.005)	(.013)	(.019)
Age squared (x 10^2)	.098**	.036	.018	.085**	.075**	.088**
	(.010)	(.030)	(.044)	(.005)	(.014)	(.019)
Mild disability	-.104*	-.054	-.055	-.177**	-.078**	-.045
	(.053)	(.049)	(.056)	(.037)	(.031)	(.035)
Moderate disability	-.299**	-.136*	-.131*	-.681**	-.291**	-.304**
	(.061)	(.056)	(.067)	(.039)	(.031)	(.036)
Severe disability				-1.472**	-.539**	-.412**
				(.130)	(.069)	(.076)
Partnered	.077	-.004	.096	.335**	.097**	.156**
	(.044)	(.056)	(.089)	(.035)	(.037)	(.059)
Number of dependent children	.072**	-.003	.004	-.008	.085**	.109**
	(.020)	(.032)	(.040)	(.017)	(.023)	(.028)
Household income ($100,000)	.368**	.123	.126	.342**	-.001	.021
	(.063)	(.078)	(.093)	(.051)	(.052)	(.060)
Others present during interview	.005	-.075**	-.032	.182**	.078**	.072**
	(.029)	(.027)	(.032)	(.022)	(.018)	(.020)
R-squared	.062	0.035	.067	.092	.004	.052
Rho (fraction of variance due to μ_i)		.626	.632		.645	.723
Model F	57.60**	13.01**	19.99**	69.65**	16.09**	36.04**
F test that all μ_i = 0		3.24**	2.22**		4.62**	2.86**
Observations	21005	21005	13880	30199	30199	21210
Individuals	6363	6363	4793	8356	8356	6735

Notes: Standard errors in parentheses; * and ** denote statistical significance at the five and one percent levels, respectively. In the pooled equations without fixed effects, the standard errors have been adjusted for the within-person clustering of observations. Constant terms have not been reported.

TABLE 5

Regression Estimates: Women

	Job Satisfaction			*Life Satisfaction*		
	No fixed effects	*Fixed effects*		*No fixed effects*	*Fixed effects*	
Employment status:						
Not in labour force				.101*	-.033	-.049
				(.041)	(.041)	(.048)
Unemployed				-.342**	-.201**	-.184**
				(.074)	(.054)	(.066)
<35 hours: underemployed	-.442**	-.337**	-.308**	-.131**	-.094*	-.092
	(.058)	(.060)	(.075)	(.046)	(.042)	(.048)
<35 hours: matched / overemployed	-.000	-.078	-.021	.169**	-.001	-.041
	(.047)	(.053)	(.067)	(.038)	(.038)	(.044)
35-40 hours: underemployed	-.282*	-.270*	-.047	-.253*	-.126	-.122
	(.112)	(.109)	(.136)	(.099)	(.080)	(.095)
35-40 hours: overemployed	-.723**	-.474**	-.425**	-.241**	-.164**	-.193**
	(.060)	(.058)	(.069)	(.046)	(.043)	(.050)
41-49 hours: matched / underemployed	.183**	.068	.198*	.052	-.030	-.058
	(.064)	(.069)	(.084)	(.056)	(.052)	(.060)
41-49 hours: overemployed	-.767**	-.566**	-.602**	-.275**	-.126*	-.188**
	(.075)	(.071)	(.085)	(.057)	(.053)	(.061)
50+ hours: matched / underemployed	.169*	.134	.189	.092	-.043	-.024
	(.080)	(.084)	(.101)	(.070)	(.063)	(.072)
50+ hours: overemployed	-.724**	-.600**	-.505**	-.276**	-.190**	-.186**
	(.071)	(.073)	(.090)	(.057)	(.054)	(.063)
Satisfaction at t-1			-.134**			-.158**
			(.011)			(.007)
Age	-.040**	-.031	-.024	-.051**	-.018	.012
	(.009)	(.028)	(.045)	(.004)	(.012)	(.018)
Age squared (x 10^2)	.066**	.025	.027	.065**	.006	-.009
	(.011)	(.034)	(.053)	(.005)	(.013)	(.018)
Mild disability	-.049	-.104	-.028	-.175**	-.044	-.053
	(.062)	(.061)	(.074)	(.037)	(.031)	(.035)
Moderate disability	-.411**	-.257**	-.144	-.812**	-.224**	-.195**
	(.071)	(.065)	(.081)	(.039)	(.029)	(.034)
Severe disability				-1.539**	-.476**	-.407**
				(.142)	(.076)	(.083)
Partnered	.224**	-.076	-.015	.329**	.113**	.017
	(.044)	(.062)	(.103)	(.030)	(.035)	(.054)
Number of dependent children	.035	.006	-.026	-.058**	.027	.045
	(.024)	(.042)	(.057)	(.015)	(.023)	(.029)
Household income ($m)	.063	.180	.181	.499**	.185**	.135*
	(.082)	(.095)	(.117)	(.056)	(.056)	(.064)
Others present during interview	.044	-.046	-.028	.162**	.080**	.078**
	(.032)	(.031)	(.039)	(.022)	(.018)	(.021)
R-squared	.051	.010	.035	.090	.001	.185
Rho (fraction of variance due to μ_i)		.558	.603		.618	.676
Model F	42.61**	15.55**	19.12**	73.09**	9.72**	27.63*
F test that all $\mu_i = 0$		2.91**	2.04**		4.47**	2.66**
Observations	18592	18592	11810	33589	33589	23928
Individuals	5940	5940	4300	8973	8973	7370

Notes: Standard errors in parentheses; * and ** denote statistical significance at the five and one percent levels, respectively. In the pooled equations without fixed effects, the standard errors have been adjusted for the within-person clustering of observations. Constant terms have not been reported.

similar – about 0.6 to 0.7 of a point. Once we control for fixed effects, however, the size of this penalty drops considerably, falling by almost half. In other words, much of the apparent penalty to working time mismatch was the result of unobserved heterogeneity. By way of contrast, the inclusion of a one-period lagged value for job satisfaction in the list of explanators appears to make little difference to the main results.

Note that we obtain a negative sign on the lagged term. Because people who report being highly satisfied in one period are much more likely to report being highly satisfied in the next, a large positive coefficient, as found by Friedland and Price (2003), might have been expected. The explanation for our finding lies in unobserved individual differences. We, too, find a large positive coefficient in the simple pooled data model, with the negative sign appearing only after controlling for fixed effects, suggesting that stable personality traits are responsible for the positive coefficient reported by Friedland and Price (2003). The negative effect indicates that net of fixed personal characteristics, a high satisfaction score one period is likely to be followed by a lower score next period. This is almost certainly a consequence of both the truncation of the satisfaction scales at 10 and the clustering of responses at the upper end. Individuals who report a 10 in one period can only report the same or a lower score in the next period. The scales, of course, are also bounded at the lower end, but this is of little consequence, given that so few respondents report a score of zero.

Considering the results for life satisfaction, we again see that the results, when using the pooled data model, yield conclusions that are very similar to those for the simple bivariate associations. As with job satisfaction, provided the number of hours worked is consistent with preferences, variations in hours worked make little difference to the overall life satisfaction scores of men. Among the mismatched, on the other hand, there is a satisfaction penalty, which is near 0.4 of a point for overemployed men working in excess of 40 hours per week, and around one-quarter of a point for workers in other mismatched categories.

Estimation of a fixed effects specification again makes a considerable difference, with these penalties again approximately halving in size. Finally, the introduction of a lagged dependent variable causes all of the hours variables, except for overemployment in the 41 to 49 hours and 50 or more hours per week categories, to lose statistical significance. In other words, in terms of men's life satisfaction, it is only both working more than 40 hours each week and working more hours than preferred, that is detrimental to self-reported indicators of satisfaction.

Are these effect sizes meaningful? The estimated magnitude of the largest variable in the fixed effects life satisfaction specification is only -.23, which seems small given that the dependent variable has an 11-point range. However, note that most respondents do not use the lower half of this scale. Further, it is well established that individual life satisfaction scores are relatively stable over time (e.g., Brickman and Campbell 1971; Headey and Wearing 1989), presumably reflecting the ability of people to adapt to, and cope with, most events and situations confronted in everyday living. Consider, for example, the estimated impact of a severe disability. Few would argue that the onset of an illness or disability that prevented the individual from engaging in any form of work activity is not a highly unfortunate and undesirable event. Yet in our final (and preferred) specification, such events are estimated only to reduce life satisfaction scores by just over 0.4 of a point. Working 50 hours per week when you would rather be working less thus an effect that is a little over half the size of the impact of debilitating injury and illness. Seen in this light, overemployment is indeed associated with meaningful reductions in life satisfaction.

Turning to the results for women, in Table 5, it can again be seen that among workers whose hours satisfy their preferences, the actual number of hours worked generally does not matter. Further, the measured penalty for overemployment on job satisfaction is much larger for women than for men. The impact on life satisfaction, however, is slightly smaller for women than for men. Similar to men, however, we discover that while underemployment has a negative impact on job satisfaction, it appears to have no significant adverse impact on life satisfaction.

One possible weakness with the approach to this point is that mismatch is assumed to have the same impact on subjective well-being irrespective of the absolute size of the gap between actual and preferred hours. We thus re-estimated the fixed effects regression models after replacing the hours variables with two variables, one for overemployed workers and another for underemployed workers, measuring the difference between usual hours worked and preferred hours. As an alternative specification, we also interacted these two variables with the number of hours usually worked per week. The estimated coefficients on these parameters are reported in Table 6.

These results suggest that underemployment is more important than was suggested by our earlier regression analyses. Nevertheless, it remains the case that, in terms of life satisfaction, overemployment exerts a larger negative impact. The size of these effects, however, is still quite small in absolute terms. If, for example, we compare a man working his preferred hours to one who is working 20 hours in excess of those preferred, the net projected loss in job satisfaction and life satisfaction would be .46 and .22, respectively. For a woman, the comparable projected effects are .60 and .14.[4]

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[4] Over the first 5 waves of data collection, an average of 25.8 percent of overemployed men and 19 per cent of overemployed women reported mismatches of 20 hours or greater.

TABLE 6

Summary of Fixed Effects Regression Coefficients Measuring the Extent of Hours Mismatch

	Job satisfaction		*Life satisfaction*	
	Men	*Women*	*Men*	*Women*
Absolute difference between usual and preferred hours				
Underemployment	-.024**	-.030**	-.005*	-.007**
	(.004)	(.004)	(.002)	(.002)
Overemployment	-.023**	-.035**	-.011**	-.011**
	(.002)	(.003)	(.001)	(.002)
Interactions between absolute difference and hours usually worked				
Underemployment * <35 hours	-.030**	-.033**	-.006*	-.007**
	(.004)	(.004)	(.003)	(.002)
Overemployment * <35 hours	-.031	-.065**	-.021	-.011
	(.018)	(.008)	(.011)	(.006)
Underemployment * 35-40 hours	-.017*	-.012	-.000	-.019
	(.007)	(.017)	(.005)	(.011)
Overemployment * 35-40 hours	-.024*	-.034*	-.003	-.015**
	(.006)	(.005)	(.004)	(.003)
Underemployment * 41-49 hours	-.013	.016	-.003	-.008
	(.012)	(.028)	(.009)	(.020)
Overemployment * 41-49 hours	-.030**	-.045**	-.011**	-.013**
	(.005)	(.005)	(.004)	(.004)
Underemployment * 50+ hours	-.011	.050	.003	-.021
	(.011)	(.065)	(.008)	(.047)
Overemployment * 50+ hours	-.023**	-.027**	-.012**	-.008**
	(.002)	(.003)	(.002)	(.002)

Notes: Standard errors in parentheses; * and ** denote statistical significance at the five and one percent levels, respectively. All models are estimated using fixed effects and include a lagged dependent variable and controls for labour force status, age, disability, relationship status, number of dependent children, household income, and presence of others at the interview. (Full results available, upon request, from the authors.)

6. Conclusions

The message yielded by our research is that neither self-reported job satisfaction nor self-reported life satisfaction scores vary much with the number of usual hours worked when the number of hours worked is consistent with preferences. This is exactly what conventional labour supply models would lead us to expect. Both relatively short hours and long hours of employment, however, are often associated with underemployment and overemployment, respectively, and workers in these situations are typically less satisfied than other workers, with the effects being larger for overemployment than for underemployment.

The absolute size of these effects, however, appears to be relatively small, especially once time-invariant, individual-specific characteristics are controlled for. Nevertheless, it also true that very few variables exert demonstrably large absolute effects on measures of subjective well-being, and that, relative to quite serious events, such as the onset of a severe illness or injury, the measured impact of overemployment should be viewed as important.

The results reported here fit neatly with Barnett's conclusion from previous studies that neither long nor short work hours *per se* are related to low levels of job or life satisfaction. Instead, the relationship is mediated by work hours preferences, and it is mismatch, rather than the absolute number of hours, that explains where long hours can be meaningfully viewed as representing undesirable overwork, and short hours as undesirable underwork. Only further research could confirm whether mismatch is also related to adverse health outcomes, but our findings suggest that such research might be worth pursuing.

Finally, we interpret our results as inconsistent with public policies imposing blanket restrictions on work hours for any and all employees, as is currently the case in France. Indeed, such restrictions would likely generate work hours mismatch among the substantial group (of mainly men) who prefer long hours, and might reduce job and life satisfaction as a result. Instead, our results are consistent with recent efforts by employers to provide flexible work arrangements, such as reduced hours options, and among public policy-makers to provide employees with the right to request flexible work arrangements subject to business needs, as currently exists in the UK. These efforts hold the promise of reducing the extent of work hours mismatch, thereby improving employee job and life satisfaction.

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