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Hours of Work and Gender Identity: Does Part-time Work Make the Family Happier?

By Alison L. Booth† and Jan C. van Ours††

†*Australian National University, University of Essex, IZA and CEPR* ††*Tilburg University, University of Melbourne, IZA and CEPR*

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Taking into account interdependence within the family, we investigate the relationship between part-time work and family wellbeing. We use panel data from the Household, Income and Labor Dynamics in Australia (HILDA) Survey. We find that part-time women are more satisfied with working hours than full-time women, and that women's life satisfaction is increased if their partners work full-time. Male partners' life satisfaction is unaffected by their partners' market hours but is increased if they themselves are working full-time. Our results are consistent with the gender identity hypothesis.

Introduction

In this paper we investigate the relationship between part-time work and working hours satisfaction, job satisfaction and life satisfaction. We account for interdependence within the family and use new panel data for partnered men and women.

Our paper makes three contributions to the existing literature. First, we investigate explicitly the degree to which part-time work may affect family subjective wellbeing. While there is a large and growing economics literature on the determinants of various components of satisfaction and happiness, few studies have explicitly investigated how part-time work status might affect individual life satisfaction, and none has looked at how part-time work affects family life satisfaction.[1] Second, following Ferrer-i-Carbonell and Frijters (2004), we use a fixed-effects ordered logit to estimate our models. This allows us to exploit more of the data than is usually done in the satisfaction literature, which typically uses a fixed-effects binary logit model. Finally, we utilize the time use module available in our data-set to illuminate our findings about partnered life satisfaction by investigating the relationship between the male shares of house work and market work. The distribution of working hours within a household may be driven by partners specializing in either market work or house work, as argued by for example Becker (1965).[2] However, social custom and conditioning—in particular gender identity as argued by Akerlof and Kranton (2000)—may influence the distribution of time spent on childcare and house work, and preferences for full-time and part-time jobs.

Although many women prefer to work part-time (OECD 2004), it is not clear *a priori* that part-time work contributes to the happiness of the family. To explore this empirically, we use the first four waves of the Household, Income and Labour Dynamics in Australia (HILDA) Survey to investigate the relationship between part-time work and various indicators of satisfaction. We use three indicators: satisfaction with working hours, overall job satisfaction, and life satisfaction. Since we are especially interested in the effects of part-time work on family life, we take into account that, for married or cohabiting couples, the distribution of paid work may not be unrelated to the distribution of home work. By studying the cross-partner effects of working part-time, we can determine whether or not part-time work makes families happier.

The setup of the paper is as follows. In Section I we give a brief overview of the literature on part-time work and job satisfaction, and relate this to the gender identity and specialization hypotheses. Section II describes the data. In Section III we examine the degree to which workers are satisfied with their current hours of work, before using a fixed-effects ordered logit model to estimate whether or not part-time work affects our three satisfaction indicators—hours, jobs and life. Our results show a gendered difference in the impact of part-time and full-time work on hours and life satisfaction. Since this suggests that Australian households are characterized by traditional gendered roles, in Section IV we exploit time-use data to estimate the relationship between the male shares of house work and market work. We find that men doing a small share of market work are also doing a small share of house work. This finding is consistent with the gender identity hypothesis of Akerlof and Kranton (2000), but inconsistent with a gender neutrality hypothesis. Section V concludes.

I. Background

While there is a large and growing economics literature on the determinants of various components of satisfaction and happiness, few studies have explicitly investigated the degree to which part-time work status might affect individual life satisfaction, as we noted in the Introduction. And none has looked at how part-time work affects family life satisfaction. In contrast, numerous studies have focused on unemployment status and individual happiness.[3] These have typically found that it is the experience of unemployment itself, rather than the loss of income through unemployment, that reduces life satisfaction. This finding has been rationalized by appealing to work as a source of social connection and self-esteem that is not found in unemployment. But these same arguments might also apply to individuals choosing to work part-time in the market sector rather than choosing home production or leisure. Moreover, a large—and in many countries growing—proportion of the workforce is in part-time work, and it would therefore seem important to know whether or not this work pattern is welfare-enhancing to the individuals and couples concerned.

Although many young people may choose to work part-time in order to finance educational investments or gain pocket money while at school, most part-time workers have family responsibilities. And family responsibilities involve partners in difficult choices, such as whether to buy in from the market sector goods and services that might alternatively be produced by one partner at home. Theories of household behaviour, such as that put forward by Becker (1965), predict that partnered households will be characterized by specialization of labour, whereby in the extreme case one partner engages in home work and the other in market sector work. Incomplete specialization, in which both partners perform part of the home work and and part of the market work, may arise because of nonlinear production functions or because cost functions associated with skills investment are characterized by economies of scope. Nonlinear production functions might arise if there is activity-specific fatigue or boredom, implying diminishing marginal productivity in each activity. Cost functions characterized by economies of scope occur if investment in market skills reduces the cost of investing in home skills. (See Rosen 1983 for a nice exposition of these.) Under incomplete specialization, there will be a monotonically declining relationship between the share of house work done by one partner and that same partner's share of market work.

Notice that Becker's theory is gender-neutral. If it is the male partner who does the lion's share of market work, his partner's share of house work should be larger; but if he does the minority share of market work, according to the specialization hypothesis he should do the majority share of house work.

Part-time jobs provide a means of combining domestic and market production, while maintaining workforce skills or experience capital for the future. Part-time work therefore facilitates incomplete specialization by either gender. The specialization hypothesis predicts gender differences in working hours because partners within a household specialize (completely or incompletely) in either market work or house work. However, the prediction is symmetric: if one partner specializes in market work the other will specialize in home production, and in principle there is no *a priori* reason why the partner specializing in market work should be female or male. We term this the *gender neutrality hypothesis*.

In contrast, the *gender identity hypothesis* of Akerlof and Kranton (2000) is based on the idea that gender matters. Here the distribution of household work and market work is determined by gender-specific 'utility'. According to this approach, since individuals operate within society's constraints, their happiness and the gender division of labour could be powerfully affected by social custom and conditioning. It is possible that, controlling for income, part-time jobs could make partnered women happier than either full-time work or no work, because such jobs allow them to gain esteem through working, while obtaining social and self-approbation from being with and caring for their families and their homes. Indeed, as argued by Akerlof and Kranton (2000), society's prescriptions about appropriate modes of behaviour for each gender might result in women and men experiencing a loss of identity should they deviate from the relevant code. If this is the case, men might be happier in full-time work and women in part-time work, since both are then adopting modes of behaviour dictated by social custom. Of course these prescriptions are endogenous to a society, as noted by thinkers such as John Stuart Mill (1869). Prescriptions might arise and continue because it is in the dominant group's interest to maintain them. They can be weakened or removed when this group loses power. For example, the female suffragette movement can be viewed as a movement aiming to remove the prescription that women were not capable of voting responsibly. Akerlof and Kranton (2000) also note that the women's movement has reduced the gender associations of particular tasks and made it more acceptable for women to work in the market sector. Whether or not this applies to men engaging in house work is a topic to which we return later in this paper. An empirical prediction of this gender identity model is that the average male share of house work will always be smaller than the female share, regardless of how the couple share their total hours of market work. (Later in the paper we will use the HILDA survey time-use module to discriminate empirically between the gender neutrality and gender identity hypotheses.)

How does the gender identity model affect life satisfaction? If women do feel a loss of identity by deviating from a particular prescription of responsibility for home production, we might expect part-time work to increase life satisfaction, *ceteris paribus*, since part-time work might be preferred simply because there is a finite amount of time in each day. If the responsibility for house work rests with the woman, then there are fewer hours available for market work and for this reason women might prefer part-time commitments. Although some partnered couples no longer follow the conventional gendered division of labour, the bulk of the evidence suggests that in the average household it is women who work part-time and assume the main domestic care role while the men work full-time. See also Section IV below.

In summary, if women prefer part-time work because it satisfies their hours preferences, given their constraints, we should observe a positive correlation between part-time work and hours satisfaction. But although part-time work might increase hours satisfaction, it might not necessarily increase job satisfaction. (Part-timers may be doing more menial and less satisfying work than if they were working full-time.) So if part-time jobs are bad jobs, overall job satisfaction may be lower. The effect of part-time work on overall life satisfaction is unclear *a priori*. It is likely to provide flexible working hours while maintaining an individual's self-esteem and social connection. On the other hand, part-time jobs may be intrinsically unsatisfying and dead-end, and therefore may reduce life satisfaction through this avenue. Part-time jobs are often viewed as bad jobs with poor pay and promotion prospects. However, Rodgers (2004) and Booth and Wood (2008), also using the HILDA Survey data, show that there is a *ceteris paribus* part-time pay premium in Australia for women and men. Ultimately, it is an empirical issue as to which effect dominates.

To our knowledge, no studies have yet explored the nexus between the happiness of the partnered couple and their work status. Yet the observed patterns of higher female participation over the life cycle, and the combination of market and household production engaged in by couples, would suggest that the relationship between work status and happiness is an important issue to address.

While happiness research in the economics literature has been underway for over a decade, only relatively recently have panel data techniques been employed to control for unobserved individual heterogeneity. Cross-sectional equations facilitate the establishment of correlation rather than causation. This is because unobservables, e.g. an extrovert personality type, can be correlated both with the propensity to report happiness and with the explanatory variables of interest. Thus, the coefficients to the latter are possibly biased in cross-sectional work.[4] The use of panel data can overcome this problem, to the extent that personality traits are fixed over time, and can be differenced out.

To our knowledge, only a few studies apart from our own look at interdependence within the family. Van Praag and Ferrer-i-Carbonell (2004, chapter 6) investigate gender differences in happiness and explore covariances between satisfaction of the two partners in a household, using random effects from a cross-section of the BHPS. Winkelmann (2005) uses the GSOEP to examine interdependence across generations, using random effects estimation.[5]

In contrast to those two studies, we use fixed-effects *ordered logit* estimation on a panel of partnered men and women. This is in contrast to the bulk of the empirical literature on satisfaction analysis, in which the categorical satisfaction scale is typically reduced to a (0,1) scale, permitting fixed-effects estimation of a binomial logit model using Chamberlain's method. But unfortunately, that method comes at a large cost, since only those individuals moving across the cutoff point can be used in the estimation. Instead of adopting that procedure, we follow Ferrer-i-Carbonell and Frijters (2004) and use an ordered logit model. This introduces individual-specific fixed-effects and individual-specific thresholds, a simple reformulation that allows Chamberlain's method to be used, removing both individual-specific effects and thresholds from the likelihood specification. Moreover, the number of observations used in this approach is significantly greater relative to the binomial logit method. This is because *all* changes in satisfaction are exploited, and not just those across some arbitrary cutoff point.

II. Data

The empirical analysis is based on the first four waves of the HILDA Survey, a nationally representative random-sample survey of private households in Australia spanning the period 2001–04. The survey is a longitudinal study of representative households in Australia. (For details, see Appendix A; this appendix also gives an overview of the definitions of the variables used in the analysis.) We restrict our estimating sub-sample to married or cohabiting couples, because we are interested in the relationship between part-time work and family welfare. Since prime-age women in particular are confronted with choices concerning family life and paid work, we further restrict our analysis to couples in which the female partner was aged 25–50 in 2001, the first year of the HILDA survey. In addition, we dropped a few couples in which the male partner was over 60 in 2001, because such males are much less likely to participate in the labour market. We use an unbalanced panel, in which selected couples are present in at least two consecutive waves. These restrictions yield a sample of 2326 couples. For females in these couples, 29% have no job 37%, have a part-time job and 34% have a full-time job. For males in these couples, 9% have no job, 7% have a part-time job and 84% have a full-time job.

In our analysis we focus on three satisfaction variables: hours of work satisfaction, overall job satisfaction and life satisfaction. The hours and overall job satisfaction variables were obtained from the following question about the individual's main job:

> 'I now have some questions about how satisfied or dissatisfied you are with different aspects of your job . . .. If not currently employed, these questions refer to the most recent job you were working in.'

Respondents were then prompted for hours of work, and then for jobs. The precise question was:

> 'All things considered, how satisfied are you with your job?'

The responses could run from 0 to 10, with higher numbers denoting higher levels of satisfaction. The life satisfaction variable was obtained from the following question:

> 'All things considered, how satisfied are you with your life? Again, pick a number between 0 and 10 to indicate how satisfied you are.'

Our measure of part-time work is based on individuals' usual hours of work in their main job (including any paid or unpaid overtime for work done at the workplace or at home). Part-time status is assigned to individuals reporting working fewer than 35 hours per week. This is also the definition of part-time work used by Rodgers (2004) and Booth and Wood (2008), and is common in classifications of work status in Australia.

The distribution of each of the satisfaction variables across different groups is presented in Table 1. As shown, satisfaction ranges from 0 to 10, with most individuals being in the upper part of this scale. If we use the share of individuals in the two top grades as an indicator of satisfaction, it is clear that part-time working women and men are both more satisfied with their *hours of work* than full-time working individuals. The same holds for *overall job satisfaction*, although here the difference between part-timers and full-timers is smaller. If we use mean satisfaction as an indicator (see bottom of Table 1), full-time working men have a slightly higher overall job satisfaction than part-time working men.

TABLE 1
SATISFACTION INDICATORS BY GROUPS OF INDIVIDUALS, 2000–2004 (%)

	Hours of work satisfaction				Overall job satisfaction				Life satisfaction					
	Women		Men		Women		Men		Women			Men		
	PT	*FT*	*PT*	*FT*	*PT*	*FT*	*PT*	*FT*	0	*PT*	*FT*	0	*PT*	*FT*
0	0.6	1.0	1.0	0.6	0.3	0.2	0.7	0.3	0.3	0.0	0.0	0.3	0.2	0.1
1	0.8	1.1	2.3	1.2	0.7	0.3	1.0	0.6	0.1	0.1	0.0	0.4	0.0	0.1
2	1.6	2.5	3.7	2.8	0.8	0.8	1.9	1.1	0.5	0.1	0.3	1.3	0.5	0.2
3	2.6	4.7	4.2	4.3	1.3	1.7	1.9	1.6	1.0	0.2	0.4	2.7	0.4	0.3
4	3.5	5.0	5.1	5.1	1.9	1.7	3.7	2.0	1.3	0.8	0.9	2.4	1.6	1.2
5	8.0	11.6	11.4	10.9	5.6	6.5	8.6	6.4	4.9	2.9	3.5	8.5	4.5	3.0
6	5.9	10.8	7.3	11.3	6.6	7.9	7.7	8.0	4.9	4.3	6.2	10.3	7.3	5.3
7	11.1	18.5	13.8	19.0	14.3	19.0	19.0	20.7	15.8	18.3	20.2	19.4	21.1	20.4
8	20.3	20.4	19.9	22.7	26.7	28.6	25.3	30.3	29.5	35.0	34.4	25.0	33.4	38.2
9	19.0	13.4	14.2	12.4	23.3	21.8	16.1	19.0	24.0	26.1	22.8	15.2	19.7	21.6
10	26.6	11.0	17.1	9.7	18.5	11.5	14.1	10.0	17.7	12.2	11.3	14.5	11.3	9.6
>8	45.6	24.4	31.3	22.1	41.8	33.3	30.2	29.0	41.7	38.3	34.1	29.7	31.0	31.2
Mean	7.78	6.86	6.98	6.86	7.90	7.67	7.36	7.54	8.03	8.09	7.95	7.45	7.80	7.92
N	3031	2750	581	6855	3032	2750	582	6854	2384	3034	2750	730	583	6853

Concerning *life satisfaction*, we distinguish between part-timers, full-timers and non-working individuals. Table 1 shows that the highest mean life satisfaction for women is associated with part-time work, while for men it is associated with full-time work. However, the differences are not large.

Figure 1 gives the distribution of usual weekly hours worked in the main job for women and men, respectively (where observations are pooled across waves and more than 60 hours is the top category). For both men and women, there is a spike at 40 hours per week. However, female hours are also more dispersed across the lower part. of the distribution while men are more dispersed across the upper part. In addition, there are spikes at five hourly intervals, as is usual in reported hours per week.

Although partnered labour supply is not the focus of this paper, in Appendix B we briefly report our estimates of the main determinants of each partner's employment probability and hours of work. Cross-sectional and fixed-effects results show that, for a woman, having young children is associated with a significantly lower employment probability and a greater part-time employment probability, while having a partner in work significantly increases the employment and part-time work probabilities. For men, having a partner in work and being in good health are associated with a significantly higher employment probability. However, the fixed-effects estimates show that male hours of work are unaffected by these variables.

III. HOURS, JOB AND LIFE SATISFACTION

(a) Pooled cross-section satisfaction estimates

We start our analysis of the satisfaction indicators—hours of work satisfaction, overall job satisfaction and life satisfaction—using an ordered logit model estimated on pooled

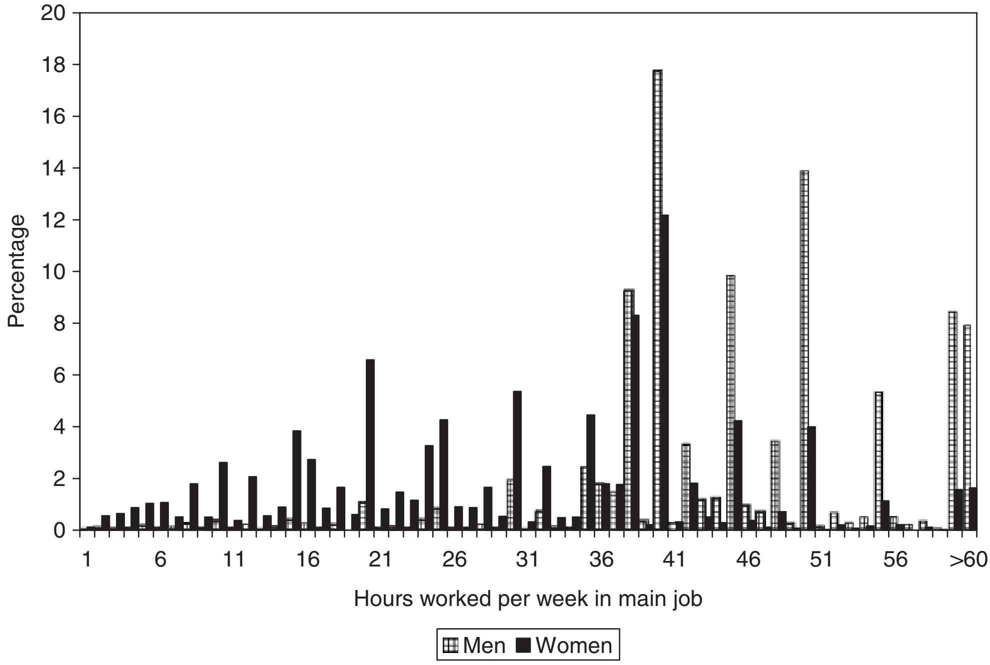

FIGURE 1. Weekly working hours, Australian women and men.

cross-sectional data. (In the following subsection we will discuss the panel estimates.) In the ordered logit model, j represents the response category ($j = 0, \ldots, 10$ for the satisfaction variables) and $\Pr(y_{it} = j) = \Lambda(\mu_j - \beta' x_{it}) - \Lambda(\mu_{j-1} - \beta' x_{it})$, with $\mu_0 = -\infty$, $\mu_1 = 0$, $\mu_{10} = \infty$. Here Λ is an indicator of the logistic cumulative distribution function, y indicates whether or not individual i is satisfied with working hours, t refers to the year, x is a vector of explanatory variables, and β is a vector of parameters to be estimated. Thus, the probability that the observed dependent variable y_{it} equals j is the probability that the latent variable y^*_{it} is between the boundaries $j-1$ and j. The μ_j are unknown parameters that are estimated jointly with β. (These are not reported in the interests of space but are available from the authors on request.)

The parameter estimates for hours satisfaction are shown in the first pair of columns in Table 2(*a*).[6] Women are more satisfied with their hours of work if they recently gave birth, if family income is higher, if they are in good health and if they work part-time. The health and working hours of their partner do not affect their own satisfaction about their working hours. Observe that the estimated coefficient to the part-time dummy variable is 0.87 (*t*-statistic 12.7). This is the largest coefficient of all, and is over twice as large as the next biggest coefficient of 0.41 (*t*-statistic 2.5) for the recent birth of a child. For men, hours satisfaction is significantly increasing in own health and working part-time.

The third and fourth columns of Table 2 show the parameter estimates for job satisfaction. For women these are by and large similar to the parameter estimates for hours satisfaction, which indicates that hours of work are an important attribute of job satisfaction. However, note that the estimated coefficient to part-time work is around a quarter of the estimated effect in the hours satisfaction regression. This is unsurprising, since hours satisfaction is but one facet of job satisfaction. Note also that female job satisfaction is increasing in the health of their partner. The estimates for men indicate

TABLE 2
PARAMETER ESTIMATES: HOURS SATISFACTION, JOB SATISFACTION AND LIFE SATISFACTION[a]

	Hours satisfaction		Job satisfaction		Life satisfaction		
	Women	Men	Women	Men	Women	Men	Sum
a. Pooled cross-section[b]							
Child born	0.41 (2.5)**	0.02 (0.2)	0.42 (2.8)**	– 0.16 (1.6)	0.17 (1.9)*	0.20 (2.1)**	0.22 (2.5)**
Family income	0.18 (2.9)**	0.01 (0.1)	0.14 (2.2)**	0.21 (3.4)**	0.27 (4.9)**	0.22 (4.0)**	0.25 (4.6)**
Health women	0.30 (3.4)**	0.03 (0.4)	0.29 (3.0)**	– 0.02 (0.3)	0.62 (7.3)**	0.02 (0.3)	0.42 (5.0)**
PT job women	0.87 (12.7)**	0.07 (1.0)	0.21 (3.1)**	– 0.02 (0.3)	– 0.04 (0.2)	0.02 (0.3)	0.03 (0.1)
FT job women	—	0.02 (0.3)	—	– 0.12 (1.5)	– 0.26 (1.3)	– 0.07 (0.9)	– 0.15 (0.7)
Health men	0.01 (0.2)	0.16 (2.1)**	0.21 (2.5)**	0.21 (2.8)**	0.08 (1.2)	0.37 (4.6)**	0.27 (3.5)**
PT job men	– 0.09 (0.5)	0.28 (2.5)**	– 0.27 (1.6)	– 0.04 (0.3)	0.09 (0.7)	0.17 (1.0)	0.19 (1.4)
FT job men	– 0.06 (0.4)	—	– 0.15 (1.0)	—	0.20 (1.8)*	0.24 (1.6)	0.30 (2.4)**
Observations	5757	7401	5758	7401	8131	8129	8127
b. Panel analysis[c]							
Child born	0.09 (0.5)	– 0.02 (0.2)	0.24 (1.2)	0.04 (0.3)	0.11 (0.9)	0.09 (0.8)	0.15 (1.3)
Family income	0.16 (1.7)*	0.05 (0.7)	0.18 (1.9)*	0.06 (0.7)	0.12 (1.6)	0.08 (1.0)	0.10 (1.3)
Health women	0.11 (0.8)	0.03 (0.3)	0.24 (1.8)*	0.11 (1.0)	– 0.02 (0.2)	– 0.06 (0.5)	0.01 (0.1)
PT job women	0.60 (6.2)**	– 0.09 (0.9)	0.10 (1.0)	– 0.05 (0.5)	– 0.04 (0.4)	– 0.06 (0.6)	– 0.05 (0.5)
FT job women	—	– 0.15 (1.2)	—	– 0.04 (0.3)	– 0.25 (2.2)**	– 0.06 (0.5)	– 0.19 (1.7)*
Health men	– 0.07 (0.6)	– 0.11 (1.2)	– 0.10 (0.8)	– 0.01 (0.2)	– 0.07 (0.8)	0.10 (1.0)	– 0.08 (0.9)
PT job men	0.09 (0.5)	0.03 (0.2)	– 0.20 (0.9)	– 0.13 (1.0)	– 0.00 (0.0)	0.17 (1.0)	0.24 (1.4)
FT job men	0.04 (0.2)	—	– 0.27 (1.4)	—	0.27 (1.8)*	0.37 (2.5)**	0.38 (2.6)**
Log likelihood	1910.8	2638.9	1815.0	2460.5	2600.5	2561.6	2953.3
Individuals	1510	1950	1433	1828	6848	6737	7658
Observations	5014	6795	4793	6404	1906	1875	2149

[a]Absolute t-values in parentheses (in the pooled cross-section estimates, corrected for clustering of observations across individuals).
**(*)Parameter estimate significant at the 5% (10%) level.
[b]*Pooled cross-section estimates*: ordered logit specification; the life satisfaction estimates also include other personal and family characteristics; in addition to this the other estimates also include other job characteristics (see Table A1 for an overview).
[c]*Panel analysis estimates*: fixed-effects ordered logit specification; all estimates included dummies for year of survey.

that men are more satisfied about their job if family income is higher and if they are in good health, but are unaffected by partner characteristics.

Finally, the last three columns of Table 2(*a*) present parameter estimates for life satisfaction—our indicator for the degree of happiness. The estimating sample now also includes women and men who do not have a job. Women are happier if a child is born. Furthermore, their happiness increases with family income and their own health. However, their life satisfaction is unaffected by part-time working (*PT*) although the coefficient is negative. Notice also that happiness in women is not affected by partner-health but is affected by the labour market position of their partner. When we experimented with removing family income as explanatory variable, this latter result becomes stronger, suggesting that it is the contribution of male full-time working to family income that explains a large part of females' life satisfaction. This appears at first blush to provide some corroboration of the income-pooling hypothesis (see e.g. Lundberg and Pollak 1996).

Men too are happier if a child is born, if they have a high family income and if they themselves are healthier. The health and labour market position of their spouse as well as their own labour market position is irrelevant for their life satisfaction.

The last column of Table 2 presents the estimates obtained when the LHS variable is the sum of both partners' happiness scores. These confirm the findings discussed above. Finally, we interacted the part-time variable with 'child born' and 'kids04' (estimates not reported in the table), but the interaction terms were never significantly different from zero. Moreover, the coefficient of the *PT* variable was hardly affected by the introduction of the interaction terms. This suggests that the effect of part-time working on the various satisfaction variables is independent of the family situation.

But do these cross-sectional estimates coincide with those obtained from fixed-effects estimation methods? Cross-sectional estimates are likely to be biased, as we argued at the start of this paper, and we therefore next report the results obtained from using fixed-effects estimation. As will be seen, we employ a less restrictive estimation method than that found in most of the panel data satisfaction literature.

(b) Panel satisfaction estimates

In the empirical literature on satisfaction analysis, a categorical scale is usually reduced to a (0, 1) scale—choosing an arbitrary common cutoff point—so that, instead of an ordered logit model, a binomial logit model may be used. This allows the introduction of fixed-effects and the estimation of the parameters using Chamberlain's method. However, this benefit comes at the cost of a large loss of observations, since only individuals who move across the cutoff point can be used in the analysis. This large loss of data may also mean that measurement errors become an important source of residual variation.

Instead of following that procedure, we use an ordered logit model, in which we introduce individual fixed-effects and individual specific thresholds: $\Pr(y_{it} = j) = \Lambda(\mu_{ij} - \alpha_i - \beta' x_{it}) - \Lambda(\mu_{i,j-1} - \alpha_i - \beta' x_{it})$. Ferrer-i-Carbonell and Frijters (2004) show that, by choosing for every individual a specific barrier k_i, the fixed-effects ordered logit specification can be reformulated as a fixed-effects binomial logit. So instead of a common cut-off point, individual specific cut-off points are chosen. This reformulation allows Chamberlain's method to be used and removes the individual specific effects α_i as well as the individual specific thresholds μ_{ij} from the likelihood specification.[7]

We start with the fixed-effects ordered logit estimates of hours of work satisfaction, reported in panel (*b*) of Table 2. The results are broadly consistent with those in panel (a). Working hours satisfaction is highest for women working part-time. It is also increasing in family income, although this is significant only at the 10% level. Compared with the cross-sectional estimate, the magnitude of the *PT* work dummy variable drops by over one-third to 0.60. However, this is still precisely estimated (the *t*-statistic is 6.2). Observe that male hours satisfaction is no longer affected by own-health or part-time status, in contrast to the cross-sectional estimates. Later we shall report the results of sensitivity analysis in which we further disaggregate the work status variables.

We next turn to the fixed-effects ordered logit estimates of overall job satisfaction, reported in the middle set of columns in the bottom panel of Table 2. It is striking that for neither partnered women or men does part-time work affect job satisfaction, in contrast to the cross-sectional estimates reported in the top panel of Table 2. For women, the fixed-effects estimates show that job satisfaction is increasing in family income and in own health and the magnitude of these estimates is quite similar to the cross-sectional results.

Finally, consider the estimates of life satisfaction reported in the last three columns of Table 2(*b*). Female life satisfaction increases if the partner gets a full-time job but declines if the woman herself moves into full-time work. Male life satisfaction is also increased if the man moves into full-time work. However, whether or not the spouse gets a job is irrelevant for the life satisfaction of Australian males.

How do the life satisfaction panel estimates compare with the cross-sectional ordered logits? We would expect that unobserved heterogeneity could be important, since, as we argued above, unobservables such as personality type may be correlated both with the propensity to report happiness and with the explanatory variables of interest.[8] Table 2(*a*) shows that the coefficient to *full-time work women* in the female cross-sectional life satisfaction equation in which family income is included is -0.26 with a *t*-statistic of 1.3. Table 2(*b*) reveals that the coefficient to the same variable in the fixed-effects ordered logit is -0.25 with a *t*-statistic of 2.2. The coefficient to *part-time work women* barely changes across estimation methods.

A comparison of the results for life satisfaction between panels (*a*) and (*b*) of Table 2 is striking in that, once the fixed-effects are introduced, the only variables that remain statistically significant are those for hours of work. The cross-sectional estimate of the coefficient to family income in the female life satisfaction equation was 0.27 (*t*-statistic 4.9), suggesting that women are happier if they have a high family income. But since this is not confirmed in the fixed-effects model (the coefficient drops to 0.12 with a *t*-statistic of 1.6), the implication is that intrinsically more satisfied women are found in households with higher family income. The fixed-effects estimate does not corroborate the income-pooling hypothesis.

In Section I we noted that, if women prefer part-time work because it satisfies their hours preferences given their constraints, we should observe a positive correlation between part-time work and hours satisfaction. This is indeed what we find. We also noted that, if part-time jobs were actually bad jobs, job satisfaction might be lower. But instead, we found no correlation between various hours of work patterns and job satisfaction for women and men. Finally, in Section I we suggested that the impact of part-time work on life satisfaction is unclear *a priori*. On the one hand, it provides a connection to the world of market work, allowing individuals to maintain human capital and some identity in that sphere. But on the other hand, the work might be dead-end and hence reduce life satisfaction. We found that for men, life satisfaction was unaffected by part-time work and increased by full-time work. For women, life satisfaction was increased by part-time work. Since we also controlled for family income, these findings are consistent with the hypothesis that, for women, part-time hours increase self-esteem or identity through work, while full-time hours do the same for men. We also found that men did not mind what their partners did with respect to market-sector work hours, but women's life satisfaction was increased if the men worked full-time. Such a gendered difference in responses is suggestive of households with traditional gender divides.

(c) Sensitivity analysis, fixed-effects estimates

Table 3 reports the estimates obtained when we disaggregated the two working-hours dummy variables (*PT job* and *FT job*) into six separate variables. These are 1–10 hours, 11–20 hours, 21–34 hours, 35–40 hours, 41–50 hours and > 50 hours. The results in the table confirm our broad findings. First, own-hours of work are statistically significant determinants of hours satisfaction and job satisfaction for both women and men.[9] Women's hours satisfaction and job satisfaction are reduced by full-time work, and especially for full-time jobs with longer hours. In contrast, male hours satisfaction is highest if they are working 35–40 hours and their job satisfaction is greatest in jobs involving 35–50 hours.

Second, female life satisfaction is significantly lower in jobs that are full-time and in full-time jobs with overtime bringing the total to 41–50 hours. It is even lower in jobs of over 50 hours per week. (The coefficient to '41–50 hours' is -0.34 with a t-statistic of 2.2, while the coefficient to '> 50 hours' is -0.75, t-statistic 3.6.) On the other hand, while female life satisfaction is higher if their men are working full-time, women do not much mind if their husbands work overtime hours: the estimated coefficients to the three male full-time hours dummies in the female life satisfaction equation are of very similar magnitude. The impact of hours of work on life satisfaction is quite large. Using the parameter estimates of Table 3, we simulated the response categories for life satisfaction using an ordered logit model. Average life satisfaction goes down by 0.16 where women work full-time, while average life satisfaction goes up by 0.20 if their partner works full-time. Similarly, if men work full-time their satisfaction goes up by 0.26. Given that about half of all individuals have a life satisfaction of 8 or 9, we think these effects are substantial.

Third, male life satisfaction is significantly higher in jobs that are full-time with or without overtime hours involving up to a 50-hour week. This mirrors the result found for male job satisfaction, although the magnitude of the coefficients is larger here. Their partners' working hours do not affect male life satisfaction.

To summarize, our results yield a gendered difference in the impact of part-time or full-time work on hours and life satisfaction.[10] This remains even when account has been taken of unobserved heterogeneity using fixed-effects ordered logit estimation. This finding is suggestive of Australian households with traditional gender divides.

We next try to extract more information from the data by exploiting the time use module.

IV. What Explains These Findings?

What might explain these observed gender differences in partners' satisfaction with part-time work? We noted earlier that theories of household behaviour predict specialization

TABLE 3

PARAMETER ESTIMATES: PANEL ANALYSES, SENSITIVITY ANALYSIS[a]

	Hours satisfaction		Job satisfaction		Life satisfaction		
	Women (1)	Men (2)	Women (3)	Men (4)	Women (5)	Men (6)	Sum (7)
Child born	0.09 (0.5)	– 0.02 (0.2)	0.19 (1.0)	0.05 (0.4)	0.10 (0.8)	0.09 (0.8)	0.16 (1.4)
Family income	0.20 (2.1)**	0.09 (1.1)	0.20 (2.0)**	0.05 (0.7)	0.13 (1.6)	0.05 (1.1)	0.10 (1.4)
Health women	0.09 (0.7)	0.02 (0.2)	0.24 (1.8)*	0.11 (1.1)	– 0.02 (0.2)	– 0.05 (0.5)	0.01 (0.2)
Hours women							
1–10	—	– 0.07 (0.6)	—	– 0.03 (0.2)	– 0.01 (0.1)	– 0.09 (0.7)	– 0.07 (0.5)
11–20	0.34 (2.5)**	– 0.09 (0.8)	– 0.24 (1.7)*	– 0.10 (0.9)	0.03 (0.3)	0.00 (0.0)	0.07 (0.7)
21–34	0.16 (1.1)	– 0.09 (0.8)	– 0.50 (3.4)**	0.01 (0.1)	– 0.11 (0.9)	– 0.08 (0.6)	– 0.11 (1.0)
35–40	– 0.30 (2.0)**	– 0.15 (1.2)	– 0.48 (3.1)**	– 0.06 (0.5)	– 0.21 (1.7)*	– 0.07 (0.5)	– 0.18 (1.5)
41–50	– 0.97 (5.4)**	– 0.07 (0.4)	– 0.58 (3.2)**	0.16 (1.0)	– 0.34 (2.2)**	– 0.09 (0.6)	– 0.15 (1.0)
> 50	– 1.40 (5.9)**	0.06 (0.3)	– 0.87 (3.7)**	0.13 (0.6)	– 0.75 (3.6)**	0.24 (1.2)	– 0.39 (2.0)**
Health men	– 0.06 (0.5)	– 0.11 (1.1)	– 0.10 (0.8)	– 0.01 (0.24)	– 0.06 (0.6)	0.10 (1.0)	– 0.07 (0.8)
Hours men							
1–34	– 0.04 (0.2)	—	– 0.01 (0.1)	—	– 0.03 (0.2)	0.09 (0.6)	– 0.17 (1.2)
35–40	– 0.14 (0.9)	0.62 (4.8)**	– 0.02 (0.1)	0.28 (2.2)**	0.28 (2.1)**	0.37 (2.7)**	0.36 (2.8)**
41–50	– 0.04 (0.3)	– 0.03 (0.2)	– 0.02 (0.1)	0.24 (1.8)*	0.28 (2.0)**	0.31 (2.2)**	0.34 (2.6)**
> 50	– 0.31 (1.7)*	– 0.77 (5.4)**	– 0.05 (0.3)	0.12 (0.8)	0.25 (1.7)*	0.09 (0.5)	0.24 (1.5)
Log likelihood	1879.6	2561.1	1806.8	2456.1	2595.1	2557.6	2950.3
LR test[b]	62.4**	155.6**	16.4**	8.8	10.8	8.0	6.0
Individuals	1510	1950	1433	1828	6848	6737	7658
Observations	5014	6795	4793	6404	1906	1875	2149

[a]Fixed-effects ordered logit specification; all estimates include dummies for year of survey; absolute *t*-values in parentheses.
**(*)Parameter estimate significant at the 5% (10%) level.
[b]Likelihood ratio test statistic on whether the more detailed hours-per-week specification differs from the simple distinction between part-time and full-time jobs; $\chi^2_{0.05} = 15.5$, $\chi^2_{0.10} = 13.4$.

of labour in partnered households (see e.g. Becker 1965). In the extreme case one partner will engage only in home work and the other only in market-sector work. We also pointed out that under incomplete specialization (Rosen 1983) there will be a monotonically declining relationship between the share of house work done by one partner and that same partner's share of market work. Moreover, we noted that Becker's specialization theory is gender-neutral. If the male does the lion's share of market work, his partner's share of house work should be larger; conversely, if he does the minority share of market work, he should do the majority share of house work. In contrast, the gender identity hypothesis of Akerlof and Kranton (2000) is based on the idea that gender matters. Here the distribution of household work and market work is determined by gender-specific 'utility'.

Following the approach of Akerlof and Kranton, we next investigate the relationship between the male share of both partners' hours spent on house work (denoted by $hrwk_{it}$) and the male share of both partners' hours spent in market work (denoted by h_{it}). We use data on house work from the time-use module in the HILDA Survey. The time-use information was obtained from each partners' responses to the following question:

> 'How much time would you spend on each of the following activities in a typical week? (Please do not count any activity twice)'

There then followed a list of activities, including 'House work, such as preparing meals, washing dishes, cleaning house, washing clothes, ironing and sewing'. The information is given as hours per week, and the male share is calculated as the hours spent by men doing house work as a proportion of the combined hours of both partners.

We have 6214 observations for 2175 families. The number of cases is somewhat smaller than for the regression results presented above, because we had to drop those observations with missing information on time use. Women spend on average 20.3 hours per week in house work compared with 5.8 hours for men. Average hours of market work for women are 21.6 while for men they are 42.0. In our sample there are 874 observations in which women do the majority of market work, which accounts for approximately 14% of the sub-sample. However, there are only 260 observations households (4.2%) in which men do less than 20% of market work. We distinguish three types of family: those without children below age 14, those with the youngest child aged 0–4, and those with the youngest child aged 5–14.

Figure 2 presents men's share of house work in each decile of h for the three groups of families. As shown, there are not many differences between the groups provided the men's share of working hours is above 0.3. Below this there are differences between the groups, but note also that the number of observations is very small here; in fact, in the second decile there are only three observations for families with children of which the youngest child was aged 0–4.

The results show quite unambiguously that there is incomplete specialization in market and house work. Households in which the male partner does the majority of market work do provide some evidence of specialization, since in those households the male share of house work is monotonically declining as their share of market work grows (see all points on the curve to the right of 0.5 on the horizontal axis). However, in households where the female does the majority of market work, the male share of house work remains proportionately low. Thus, the degree of specialization is partial and non-symmetric.[11]

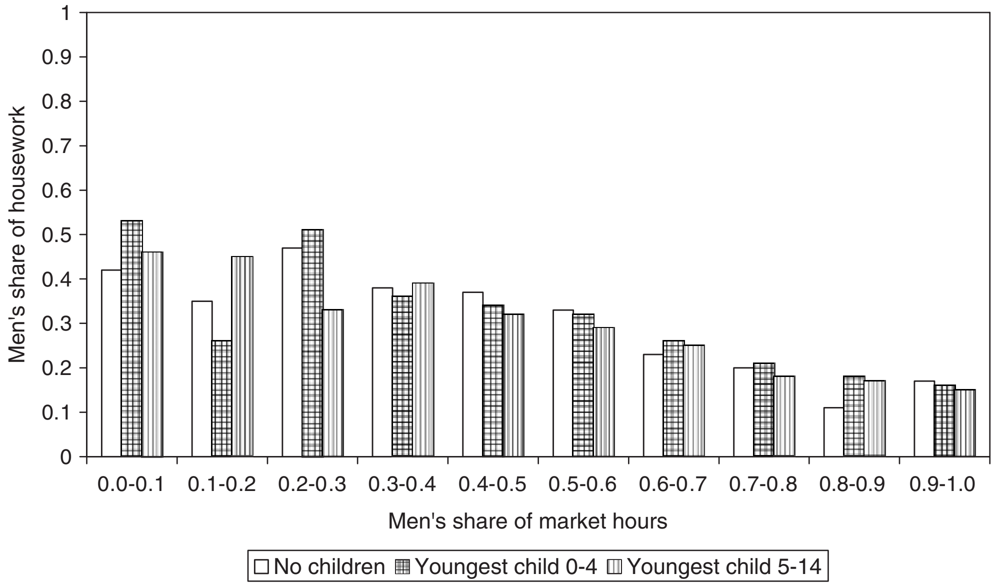

FIGURE 2. Men's share of house work hours versus their share of market work hours.

Note: Average values per decile in men's share of market hours.

A simple test for gender neutrality is whether or not the slope of a regression of $hrwk_{it}$ on h_{it} equals -1. This is clearly not the case. In a pooled regression, the slope is given by -0.21 (t-statistic 13.1). In a fixed-effects estimate the slope is -0.33 (22.6). This slope implies that a one-percentage-point decrease in men's share of market hours increases men's share of house work by only 0.33%.

Another simple test for gender neutrality is based on the symmetry of Figure 2. In case of incomplete specialization but gender neutrality, if the right-hand corner of Figure 3 is (1, a), with a being small but positive, the left-hand corner of Figure 2 should be (0, $1-a$). If men do all market work and still have a share of a in house work, symmetry requires that if women do all the market work they too should do a share of a in house work. The same holds for intermediate positions in Figure 2; gender-neutrality requires the average share of house work for men who have a share in market work higher than any arbitrary value is equal to 1 minus the the average share of house work for men who have a share in market work lower than 1 minus that arbitrary value. More concisely,

$$(1) \qquad (\overline{hrwk}|h > x) = ((1 - \overline{hrwk})|h < 1 - x)$$

for all values of $0.5 \leqslant x \leqslant 1$. This obviously is not the case. If we take $x = 0.5$, we find $(\overline{hrwk}h| > 0.5) = 0.217$, significantly different from $((1 - \overline{hrwk})|h < 0.5) = 0.629$. If we take $x = 0.25$, these numbers are 0.238 and 0.559. Clearly, the distribution of household work and market work is not gender-neutral.

But perhaps men who do a low share of market work do a larger share of other household-related activities, such as outdoor tasks or childcare. To investigate this, we use responses to the time-use questions about these other activities. The outdoor activities and childcare responses were obtained from the listing following the question:

> 'How much time would you spend on each of the following activities in a typical week? (Please do not count any activity twice)'

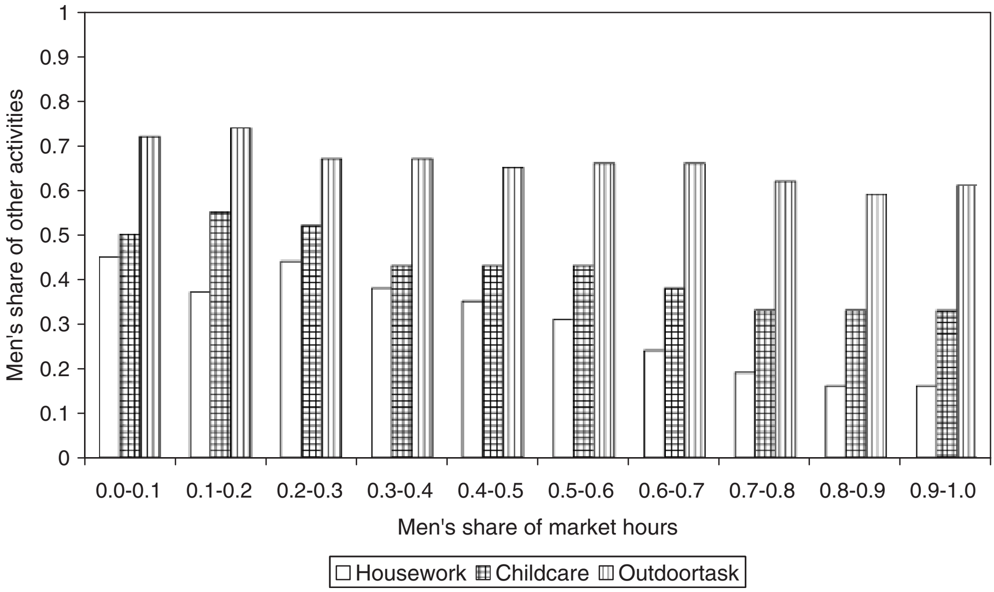

FIGURE 3. Men's share of house work hours, childcare hours and outdoor task hours versus their share of market work hours.

Note: Average values per decile in men's share of market hours.

The outdoor tasks question asked respondents to include time spent on 'home maintenance (repairs, improvements, painting etc.), car maintenance or repairs and gardening'. The childcare question asked respondents to include time spent on 'playing with your children, helping them with personal care, teaching, coaching or actively supervising them, or getting them to childcare, school and other activities'. On average, women spend 3.4 hours a week on outdoor activities while men spend 5.9 hours. Childcare activities absorb on average 17.3 hours of women's time and 8.3 hours of their partners' time. Figure 3 shows a breakdown of the male share of various household activities across the distribution of the male share of market hours. Thus, the figure gives average values per decile of men's share of market hours. The figure shows that outdoor tasks seem to be unrelated to market work. Men do a higher share of these tasks across the distribution. This suggests a gendered division of labour for outside work that is invariant to market hours shares. For childcare there is a trade-off between the male share of childcare and the male share of market work, but it is not large. Men who do 90%–100% of market work do about 30% of childcare, while men doing 0%–10% of market work do about 50% of childcare. This suggests incomplete specialization. The figure also confirms that the most striking finding is for house work.

In summary, we find a declining relationship between the share of house work done by men and their share of market work, and this is unaffected by the presence of dependent children. But there is certainly not complete specialization, for even the men doing all market work are also still doing some home work. Nor does there appear to be gender neutrality, since in households in which women spend more time making money they are also still doing more house work. This finding is inconsistent with the gender neutrality hypothesis, but is consistent with the gender identity hypothesis about time use within the household.[12]

V. Conclusions

This paper investigates the relationship between part-time work and three indicators of satisfaction: satisfaction with working hours, overall job satisfaction, and life satisfaction. The data used are from the first four waves of the Household, Income and Labour Dynamics in Australia Survey, spanning the period 2001–2003, and we focus on a sample of partnered men and women.

Our fixed-effects ordered logit results indicate that, conditional on observed characteristics, part-time women are more satisfied with their *hours of work* than full-time women. For men, hours of work satisfaction is greatest for those working 35–40 hours a week. However, for *job satisfaction* there is no such relationship. Indeed, for both men and women, job satisfaction seems to be independent of hours of work.

Finally, we found that partnered women's *life satisfaction* is reduced by working full-time, especially if their weekly hours are greater than 40. However, female life satisfaction is increased if their partners are working full-time, and they are particularly happy if their partners are working 35–50 hours per week. In contrast, male partners' life satisfaction is unaffected by their partners' market hours but is significantly increased if they themselves are working full-time, especially so if they are working 35–50 hours per week. Thus, it seems that full-time work for men in the region of 35–50 hours is the major contributor to both partners' life happiness, but that female part-time work has an asymmetric effect. Men do not mind how many hours per week their partners work, but women are happiest with part-time work.

Does this suggest that Australian families are characterized by complete specialization, with one partner engaged predominately in domestic work and the other in market-sector work? The answer is no. According to the specialization hypothesis, there will be a negative monotonic relationship between the share of house work done by one partner and that same partner's share of market work. This prediction is not supported by the data. In households where the female does the majority of market work, the male share of house work remains proportionately low. Thus, the degree of specialization is partial and non-symmetric. Men doing a small share of market work were also doing a small share of house work. This finding is consistent with the gender identity hypothesis, and it may suggest a reason why women are happier with part-time work.

APPENDIX A: THE HILDA DATA

The new Household, Income and Labor Dynamics in Australia (HILDA) Survey began in 2001. It is a nationally representative random-sample panel survey of private households in Australia. We use data from all four available waves, which span the period 2001–04.

All members of households providing at least one interview in the first wave form the basis of the panel. The sample has been gradually extended to include any new household members resulting from changes in the composition of the original households. The survey is a longitudinal study of representative households in Australia. For details, see http://www.melbourneinstitute.com/hilda. HILDA contains four survey instruments: the household form, a household questionnaire a person questionnaire, and a self-completion questionnaire. The information at the household level can be provided by any adult member of the household but preferably a person with knowledge of the household finances. The person-level questionnaires are for all persons aged 15 y and over in the household.

Table A1 provides an overview of the variables used in the analysis. There are five types of variable: personal characteristics, hours of work, job characteristics, family characteristics, and the time use information that was obtained from the self-completion questionnaire. We restrict our estimating sub-sample (for reasons given in the text) to married or cohabiting couples in which the

TABLE A1
DEFINITIONS OF VARIABLES AND MEANS

Variable	Definition	Women	Men
Personal characteristics			
Age	Respondent's age	39.3	41.8
Postgrad	Postgraduate degree (masters or doctorate)	0.03	0.05
Graddip	Graduate diploma or certificate	0.08	0.06
Bachelor	Bachelor degree	0.16	0.15
Advdip	Advanced diploma, diploma	0.10	0.10
Cert	Certificate	0.13	0.32
Year	Year 12 (base is year 11)	0.16	0.09
Born-oz	Australian born	0.76	0.75
Born-engsp	Born in English speaking country (not Oz)	0.10	0.12
Health	Dummy: in good health	0.85	0.81
Part-time job	Usual hours per week in main job <35	0.37	0.07
Full-time job	Usual hours per week in main job $\geqslant 35$	0.34	0.84
Hours of work[a]			
Hwork	Hours spent on house work in typical week	20.3	5.8
Chdcare	Hours spent on own childcare in typical week	17.3	8.3
Outdoor	Hours spent on outdoor activities per week	3.4	5.9
Home production	Total hours spent on home activities per week	41.0	19.9
Market work	Total hours spent in main job per week	21.6	42.0
Total hours of work		62.6	61.9
Job characteristics[b]			
Hours	Usual hours per week in main job	30.5	46.1
Casual	Casual contract	0.19	0.07
Contract	Fixed-term contract	0.08	0.06
Permanent	On going permanent employment	0.55	0.61
Siz20–99	Firm has 20–99 employees	0.27	0.24
Siz100–499	Firm has 100–499 employees	0.16	0.17
Siz500up	Firm has 500 or more employees	0.09	0.10
Industry dummies	One-digit industrial classification	–	–
Family characteristics			
Family income	Log(Total annual family gross income) – 1000 AUD	Log(84.9)	
Child born	Dummy: whether or not household had a new birth	0.05	
Child 0–4	Dummy: kids 0–4 years of age	0.27	
Child 5–14	Dummy: kids 5–14 years of age	0.50	
Urban	Living in major city	0.59	
Innreg	Inner regional	0.26	
Outreg	Outer regional (base is remote/very remote)	0.10	

[a]Hours of work refer to the main job.
[b]10.4% of the females and 7.7% of the males have more than one job.

female partner was aged 25–50 in 2001. We use an unbalanced panel, in which selected couples, are present in at least two consecutive waves. These restrictions yield a sample of 8170 observations of 2326 couples, of which 1601 were observed four times, 316 three times and 409 twice.

Partnered women and men are very much alike in terms of personal characteristics, as shown in Table A1, but there are substantial differences in their hours of work. While women on average spend about 20 hours per week on house work and 17 hours per week on childcare, men spend about 6 hours per week on house work and 8 hours on childcare. Of the women in our sample, 37% have a part-time job and 34% have a full-time job. Of the men, 7% have a part-time job and 84% a full-time job. These differences materialize in the usual hours per week in the main job, which is about 30 hours for women and 46 hours for men. In terms of job characteristics, the main difference between men and women concerns the share of workers with a casual contract, which is 19% for women and 7% for men.

APPENDIX B: PARTNERED LABOUR SUPPLY

To get some idea about the determinants of employment, we estimate discrete choice models using pooled cross-section data as well as exploiting the panel character of the data. To investigate the way in which the decisions of one partner affect the other, we also allow some individual characteristics to affect the partner's employment position. Thus, we ignore joint decision-making and assume that the decision of the partner is exogenous to the decision of the individual. The probability of having a job is analysed using a logit specification. We use the logit specification since it is a natural starting point for the introduction of fixed effects. In a bivariate probit model (not reported), we investigate to what extent there is correlation in the behaviour of partnered men and women conditional on their observed characteristics. We find that the estimated parameters are hardly affected by the introduction of possible correlation in the unobserved characteristics, whereas the correlation itself is positive and significantly different from zero. This indicates either joint decision-making or perhaps selective matching (individuals who are more likely to work match with similar individuals) that is orthogonal to observed characteristics. Thus, $\Pr(y_{it}=1)=\Lambda(\beta x_{it})$ and $\Pr(y_{it}=0)=\Lambda(-\beta x_{it})$, where Λ is an indicator of the logistic cumulative distribution function, y indicates whether or not an individual has a job, i refers to individual and t refers to the year of the survey. Furthermore, x is a vector of explanatory variables, and β is a vector of parameters.

The principal explanatory variables used in the analysis are: age, health, whether or not a household had a new birth in the period since the previous interview (or in the previous 12 months in the case of wave 1), whether the household has children in the age group 0–4 or 5–14, and the year of survey. Other variables included are education, country of birth and degree of urbanization. However, since these variables are time-invariant, they drop out of the panel analysis. In the interests of space, we do not report the estimated coefficients to these variables in the pooled cross-sectional models.

The first column of Table A2 reports the parameter estimates, where the upper panel gives the results for women and the lower panel, those for men. Age has a statistically significant effect for men only. Older men are less likely to have a job. For both women and men, being in good health has a positive effect on the probability of having a job. Having young children or teenage children has a negative effect only on the female probability of having a job. As shown, having a partner with a part-time or full-time job is positively related to an individual's own job probability. This association is consistent with studies showing the presence of work-rich and work-poor households (see e.g. Dickens and Ellwood 2003).

If we introduce fixed effects in a logit model, the specification becomes $\Pr(y_{it}=1)=\Lambda(\alpha_i+\beta x_{it})$, and $\Pr(y_{it}=0)=\Lambda(-\alpha_i-\beta x_{it})$, where the α_i represent individual fixed-effects. The parameters of this fixed-effects logit model are estimated using Chamberlain's conditional likelihood method. This means that the parameters are identified on the subset of observations where the dependent variable changes at least once over time.

As shown in the fourth column of Table A2, the number of observations reduces substantially if fixed effects are introduced. In total, 448 women and 243 found a job or lost a job at least once. However, by and large the results are not much different from the estimates based on pooled cross-sections. Note that in a fixed-effects setting we cannot identify the effects of age, since there is perfect correlation between age and calendar years. The results show, first, that the birth of a child increases the female probability of moving out of work. This is unsurprising—especially in view of the fact that Australia is one of the few OECD countries without statutory maternity leave provision (OECD 2001). Second, if a child moves from the age category 0–4 years to a higher age category the female probability of finding a job increases; however, for men these changes in family situation do not affect their labour market position. Third, an improvement in health significantly increases the probability of finding a job.

TABLE A2
PARAMETER ESTIMATES: EMPLOYMENT, PART-TIME WORK AND HOURS OF WORK[a]

	Pooled cross-section estimates			Panel estimates		
	Job (1)	Part-time (2)	Hours (3)	Job (4)	Part-time (5)	Hours (6)
Women						
Age	– 0.00 (0.5)	0.03 (4.9)**	– 0.12 (2.7)**	—	—	—
Health	0.84 (8.5)**	– 0.12 (1.1)	0.50 (0.6)	0.43 (1.7)*	0.44 (2.0)**	– 0.64 (1.2)
Child born	– 1.20 (9.8)**	0.22 (1.2)	– 2.15 (1.8)*	– 2.43 (7.1)**	– 0.02 (0.1)	– 1.82 (2.2)**
Child 0–4	– 1.26 (13.0)**	1.26 (10.6)**	– 8.71 (11.8)**	– 1.33 (4.2)**	1.91 (7.3)**	– 6.73 (10.9)**
Child 5–14	– 0.36 (4.4)**	0.91 (10.8)**	– 5.59 (10.0)**	– 0.08 (0.3)	0.42 (1.6)	– 1.65 (2.6)**
Partner full-time	1.38 (10.9)**	0.46 (2.6)**	– 0.97 (0.9)	0.93 (3.0)**	0.60 (1.7)*	– 0.06 (0.1)
Partner part-time	1.36 (8.0)**	0.54 (2.6)**	1.74 (1.4)	1.11 (2.8)**	0.90 (2.3)**	– 1.20 (1.3)
Observations	8170	5785	5639	1291	1770	5639
Individuals	—	—	—	448	519	1896
Men						
Age	– 0.03 (3.3)**	0.02 (2.0)**	0.30 (0.8)	—	—	—
Health	1.95 (6.7)**	– 0.73 (5.7)**	2.30 (3.9)**	0.78 (3.1)**	0.16 (0.6)	0.16 (0.4)
Child born	0.27 (1.2)	0.07 (0.3)	– 0.16 (0.2)	0.16 (0.5)	0.21 (0.6)	– 0.02 (0.5)
Child 0–4	0.24 (1.5)	– 0.12 (0.8)	1.00 (1.9)*	– 0.03 (0.1)	0.01 (0.0)	– 0.36 (0.8)
Child 5–14	0.36 (2.8)**	– 0.20 (1.6)	1.14 (2.5)**	– 0.14 (1.4)	0.37 (1.2)	– 0.18 (0.4)
Partner full-time	1.08 (6.9)**	– 0.11 (0.7)	0.91 (1.5)	0.53 (1.7)*	0.02 (0.1)	– 0.06 (0.1)
Partner part-time	1.56 (10.8)**	– 0.06 (0.4)	0.20 (0.4)	0.87 (3.3)**	0.06 (0.2)	0.05 (0.2)
Observations	8170	7439	7273	872	913	7273
Individuals	—	—	–	243	270	2206

[a]'Part-time' and 'Hours' concern choices conditional on having a job. 'Job' and 'Part-time' logit model specification, 'Hours' linear specification and the pooled cross-section estimates also contain other personal characteristics and family characteristics (see Table A1 for details). The panel estimates include individual fixed effects. All estimates contain dummy variables for survey years. Absolute t-values in parentheses (in the pooled cross-section estimates corrected for clustering of observations).
**(*)Parameter estimate significant at the 95% (90%) level.

The second column of Table A2 shows the pooled cross-sectional estimates of the determinants of the individual probability of having a part-time job conditional on being in work. For both men and women this probability increases with age. Females are less likely to have a part-time job if there are preschool children. We also find this result in the fixed-effects estimates reported in the fifth column. Furthermore, a woman is more likely to work part-time if her partner works, a result that we do not find if fixed effects are included. This suggests that the partner effect may be due to unobserved characteristics rather than being a causal effect. For males, apart from age, only health has an effect on the probability of their working part-time; but again, from the fixed-effects estimates, it seems as if this is not a causal effect.

Finally, columns (3) and (6) of Table A2 show the determinants of the hours of work decision from the pooled cross-sections and the fixed-effects estimation, respectively. In both, the presence of preschool children significantly reduces female but not male hours of work.

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NOTES

1. See e.g. Bardasi and Francesconi (2004). Some other studies examining individual life satisfaction include part-time work status as a control but do not comment on the estimated coefficients. Frijters *et al.* (2004a, b), using the GSOEP data find that life satisfaction is higher for full-time and part-time women—and for non-participating women—relative to the base of unemployed women. In job satisfaction studies, hours of work are frequently included as controls, and typically have a negative effect on job satisfaction (see *inter alia* Clark 1997; Clark and Oswald 1994), apart from overtime hours see Van Praag and Ferrer-i-Carbonell (2004, pp. 56–7).
2. See also Rosen (1983), who emphasizes how nonlinear production functions could lead to incomplete specialization. We return to this in Section IV.
3. For studies using panel data explicitly to investigate the relationship between happiness and unemployment, see Carroll (2007); Clark and Oswald (1994); Clark (2003); Clark *et al.* (2001); Gerlach and Stephan (1996); Winkelmann and Winkelmann (1998).
4. Studies using panel data techniques to examine the determinants of satisfaction include Carroll (2007); Clark (2003); Ferrer-i-Carbonell and Frijters (2004); Frijters *et al.* (2004a, b); and Hamermesh (2001).
5. Plug and Van Praag (1998) compare partners' responses to subjective wellbeing questions. While they find little difference, this is not the case with our variable of interest. In a companion paper (Booth and van Ours 2007), we use the BHPS to explore the determinants of partnered wellbeing in Britain. We find that life satisfaction of men and women is not affected by how many hours they work. Men have the highest hours-of-work satisfaction if they work full-time without overtime hours; women have the highest hours-of-work satisfaction if they work part-time, irrespective of whether the part-time jobs are small or large.
6. In an earlier version of the present paper, we also estimated the determinants of whether or not individuals would be happier working about the same hours as currently, or working more or fewer hours (for details, see Booth and van Ours 2005).
7. In our estimates we use $k_i = \Sigma_t y_{it}/n_i$, where n is the total number of observations of individual i. All observations for which $y_{it} > k_i$ is transformed into $z_{it} = 1$, and all observations for which $y_{it} \leqslant k_i$ is transformed into $z_{it} = 0$. Alternatively, we used $z_{it} = 1$ if $y_{it} \geqslant k_i$ and $z_{it} = 0$ if $y_{it} < k_i$. This hardly affected the parameter estimates.
8. To investigate the relevance of fixed effects formally, we performed Hausman tests for each satisfaction category. This compares the panel analysis estimates with those from the pooled cross-section. In all cases the chi-squared statistic indicates rejection of the null hypothesis that the difference in coefficients is not systematic. Hence the fixed-effects estimates are preferred.
9. We do not comment in the text on the other coefficient estimates, since they are very similar to those reported in Table 2(*b*), except that for women hours satisfaction and job satisfaction is increasing in family income. Female job satisfaction is also increased by own-health.
10. We also formally tested if the panel estimates of life satisfaction are significantly different for men and women. For the estimates presented in Table 3 (columns (5) and (6)) we find that the likelihood ratio (LR) test statistic for equal parameter estimates = 25.0, which is significant at the 10% level (the

critical $\chi^2_{0.10}$-value = 23.5). This result is to a large extent driven by the many insignificant parameter estimates. If we re-estimate the model with a limited number of explanatory variables (hours women: 35–40, 41–50, 50 + ; hours men: 35–40, 41–50, 50 + , year dummies) we find the LR test statistic = 22.4, which is significant at the 5% level (critical $\chi^2_{0.05}$-value = 15.5).

11. We also distinguished between those families in which the female partner had a high level of education (college and above) and those in which she had lower levels of education. Although in all cases the male share of house work was higher in those households with more highly educated females, the differences were small, and in no cases was the male share of house work as high as 50%.
12. This result was also found using US data by Akerlof and Kranton (2000).

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