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Job Changes, Hours Changes and the Path of Labour Supply Adjustment

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Abstract

This paper uses the first twelve waves of the British Household Panel Survey covering the period 1991-2002 to investigate single women's labour supply changes in response to three tax and benefit policy reforms that occurred in the 1990s. We find evidence of small labour supply effects for two of such reforms. A third reform in 1999 instead led to a significant increase in single mothers' hours of work. This increase was primarily driven by women who changed job, suggesting that labour supply adjustments within a job are harder than across jobs. The presence of hours inflexibility within jobs and labour supply adjustments through job mobility are strongly confirmed when we look at hours changes by stated labour supply preferences. Finally, we find little overall effect on wages.

JEL Classification: C23, H31, I38, J12, J13, J22
Keywords: Job mobility; Hours flexibility; Labour supply preferences; Hours-wage trade-off; Monopsony.

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I. Introduction

The use of the canonical model of labour supply for policy analysis is pervasive. A central tenet of this model is that workers have flexible choices over hours of work, selecting their desired utility-maximising level at any given wage. A number of studies have cast some doubt on this model by arguing that there is not free choice of hours within a job and limited choice across jobs, and providing evidence of job "packages" whereby wage and hours are tied together.[1] Most of the contributions in this literature however identify hours constraints relying on observed individual characteristics (e.g., number and age of children, or job mobility) or stated labour supply preferences (Ham, 1982; Moffitt, 1984; Lundberg, 1985; Altonji and Paxson, 1988; Stewart and Swaffield, 1997). This strategy is problematic because changes in labour supply preferences or other individual variables may not be exogenous to hours levels or changes.

Our strategy is to use a sequence of policy reforms that directly affected the labour supply incentives of specific groups of individuals while leaving the incentives faced by others unchanged. Our objective is to use these reforms to assess the degree of flexibility of hours changes within and across jobs. Specifically, we analyse the changes in hours worked by single women in response to (exogenous) tax and benefit policy changes that occurred in Britain in the 1990s. We use three different reforms to highlight likely actual movements along the labour supply curve, and combine these with information on stated preferences and job mobility to assess whether and how women adjust their labour supply in response to changes in the incentives to work a given number of hours.

Much of tax and benefit reforms in the United Kingdom, Canada, and the United States has been directed at increasing the labour market attachment of the lower skilled workers, in particular those facing high fixed costs of work such as childcare (Blundell, 2002). A significant part of the rise in employment among single mothers in the US over the late 1980s and 1990s has been attributed to the expansion of the Earned Income Tax Credit (Eissa and Liebman, 1996; Meyer and Rosenbaum, 2001). Similarly, it has been argued that much of the rise in the participation of single

[1] See Blundell and MaCurdy (1999) for an overview.

mothers in the UK has been due to increases in the generosity of the tax credit policies, namely Family Credit (FC) and Working Families' Tax Credit (WFTC).[2] The self-sufficiency experiment in Canada provided further experimental evidence on the effectiveness of financial incentives on the working decisions of low income single parents (Card and Robins, 1998). An interesting feature of the UK reforms has been the changing incentive structure towards part-time and full-time work engendered by these reforms. Not only has employment responded to these reforms but so has the distribution of weekly hours of work (Blundell et al., 2000; Brewer, 2001). However, the precise mechanism for these adjustments in labour supply has not been studied. Are adjustments to hours made by moving jobs or can workers adjust their hours of work over time with the same employer? This mechanism of adjustment is the focus of this paper.

For such an analysis panel data are essential, as it is necessary to know the initial employment position and hours worked of each specific individual before and after adjustment takes place. Since 1991 a high-quality panel data survey, the British Household Panel Survey (BHPS), has been collected annually for Britain, and that is the data source we use in our analysis covering the period 1991-2002. The BHPS also has the attraction of recording individuals' *stated* preferences toward hours of work, so that actual movements can be examined alongside changes in stated preferences.

Even if hours were completely fixed *within* jobs but mobility *between* jobs is costless, we would still expect workers to be located on their labour supply curve, i.e., at their most preferred level of hours given the market wage. But if there are individual costs to moving between jobs or firms collectively require a given number of hours due to facing fixed costs or technology-related coordination requirements,[3] then workers will face immobility (at least in the short run) on the hours they can work. This has implications for the interpretation of data on actual and preferred hours of work, rates of mobility between jobs, and for estimating models of labour supply. Various

[2] Blundell and Hoynes (2004) present a comprehensive review of the evidence.
[3] Card (1990) argues that constraints are the result of nonconvexities in the relationship between output and individual hours due to start-up costs or other aspects of the technology used.

strands of research have suggested models of hours choice where hours are fixed within jobs. One strand, which dates back to Barzel (1973) and Rosen (1976), grounds its analysis in models where jobs are packages of fixed hours-wage combinations (Ham, 1982; Moffitt, 1984; Lundberg, 1985; Biddle and Zarkin, 1989; Kahn and Lang, 1991; Altonji and Paxson, 1988, 1992; Dickens and Lundberg, 1993). Another more recent strand is developed within a monopsonistic environment, where employer preferences play a key role in determining hours of work in a given job (Manning, 2003).

In this study we are interested in checking if and how employed single mothers vary their hours in response to exogenous changes in the incentives to work a given level of hours, under the null hypothesis of complete flexibility in hours choice within the job. For this purpose, we use reforms to the tax and benefit system that changed the hours conditions for FC in 1992 and 1995 and the attractiveness of work through WFTC in 1999 to assess the 'canonical' model of complete hours flexibility. We also look at how changes in hourly wages relate to the introduction of the reforms both within and between jobs. Although this analysis can be biased by the usual endogeneity problems, it is likely to give us a more complete picture of the British labour market and an indication of the possible presence of imperfections.

Several studies have analysed the effect of the UK tax/benefit reforms, especially those of the WFTC programme, on employment, wages, and other outcomes (Bingley and Walker, 1997; Blundell et al., 2000 and 2004; Blundell and Hoynes, 2004; Gregg and Harkness, 2003; Brewer et al., 2005; Francesconi and Van der Klaauw, 2005). None of these studies, however, focuses on changes in worked hours. Stewart and Swaffield (2004) instead examine the working hours of low-wage employees in the UK, but analyse the impact of the introduction of the minimum wage in April 1999 rather than the impact induced by reforms that potentially changed the incentive to work a given number of hours per week. Their results indicate that the minimum wage had a negative effect on hours worked by low-wage women, although we do not know how single women with and without children have been differentially affected. In addition, neither these studies nor the earlier research on wage-hours packages analyse job changing behaviour as a mechanism to adjust hours of work or address the broader issue of labour supply adjustment.[4]

We find that the introduction of the WFTC reform in 1999 led to a substantial increase in single mothers' hours of work. This increase was primarily driven by women who changed job, suggesting that labour supply adjustments within a job are much less flexible than across jobs. There is a good deal of heterogeneity in the effects of the WFTC reform, with stronger evidence of hours inflexibility emerging among single mothers who were better educated, whose youngest child was aged 0-4, and who worked in larger firms, service industries, and the public sector. The presence of hours inflexibility within jobs is confirmed when we look at hours changes by stated labour supply preferences: unconstrained women showed the largest upward adjustments after the WFTC reform and overemployed women showed the largest downward adjustments after the 1992 FC reform (which reduced the minimum work requirement to receive FC from 24 to 16 hours a week) only if they changed job. Finally, we find relatively little effect on wages. However, there is evidence that some groups of women (especially single mothers who lived in London and the South East) operated under monopsonistic conditions, whereby changing job led to significantly lower wages after the introduction of WFTC.

Our research is likely to be relevant for many aspects of labour market policy, especially for the design of tax credit and benefit policies which specify a minimum number of hours of work per week as a precondition for entitlement to a given payment (e.g., the Working Tax Credit, and the current pilot for the Employment Retention and Advancement Scheme in the UK). From the result that hours are not flexible within jobs, we can infer that changes to the tax/benefit incentives to work a given number of minimum hours are likely to affect rates of job-to-job transitions for the affected groups of workers.

[4] There has been relatively little analysis of hours constraints in Britain. Two studies that have investigated the extent of constrains on desired hours are Steward and Swaffield (1997) and Bryan (2000). Using data from the British Household Panel Survey, they both find that a substantial proportion of male workers (Steward-Swaffield) and male and female workers (Bryan) are not putting in the hours they would like, with most of the dissatisfied workers wishing to work fewer hours per week. Both studies, however, abstract from the way in which job changes are related to hours changes, and, more broadly, from the issue of the path of labour supply adjustment.

The next section briefly explains the rules and structure of the FC/WFTC programs, and discusses our estimation approach and identification strategy. Section III introduces the data, and describes the variables used in the analysis. Section IV presents the empirical results, and Section V summarizes our main results.

II. 'In Work' Benefit Reforms in the United Kingdom

A. Institutional Background

Programmes to support low-income working families with children (hereafter called "in-work benefits", even though the more recent programmes are officially designated tax credits) have a long history in the United Kingdom. A peculiar feature of the UK's in-work benefits is that awards depend not just on the earned and unearned income and family characteristics, but also directly on (weekly) hours of work: since their inception, in-work benefits have only been available to families with children who usually work some the minimum number of hours a week.[5]

Two in-work benefits were in operation during our sample period: Family Credit (FC), which existed from April 1988 until September 1999, and the Working Families' Tax Credit (WFTC), which existed from October 1999 until March 2003.[6] In April 1992, the minimum work requirement in FC fell from 24 to 16 hours a week. This occurred between the first two waves of the BHPS. The impact of this reform on single parents' labour supply is ambiguous: those working more than 16 hours had an incentive to cut hours to (no less than) 16, while those previously working fewer than 16 hours had an incentive to increase their labour supply to (at least) the new cut-off. In 1995, there was another reform to Family Credit, in the form of an additional (small) credit for those adults working full time (i.e., 30 or more hours a week). This reform affected the labour supply decisions of lone parents in obvious ways: there was an increased incentive for those

[5] Hours rules are an important feature of the UK's welfare system more generally. Receipt of the basic safety-net welfare benefit (Income Support or income-related Jobseekers' Allowance) is conditional on both working less than a certain number of hours and having a sufficiently low income. For parents, the hours rules for welfare benefits and in-work benefits are aligned so that families can never be entitled to both.

[6] Since 1998, the transfer system affecting lone parents has undergone nearly continuous reform. However, the most important change, in terms of both government expenditure and potential labour supply effects, was the introduction of WFTC. We do not want to claim however that there has been a stable post-reform period since October 1999.

working less than 30 hours to increase their hours to 30, but an income effect meant that those already working at least 30 hours had an incentive to cut their hours worked to no less than 30.

The 1999 WFTC reform has a more complicated impact on labour supply. WFTC was more generous than FC in three ways: it had higher credits, particularly those for young children, families could earn more before the benefit began to be withdrawn, and it had a lower withdrawal/taper rate. Overall, the reform increased the attractiveness of working 16 or more hours a week compared to working fewer hours. But the last of the three aspects of the reform meant that the biggest income gains were experienced by families just at the end of the FC taper (i.e., families whose earnings had reduced their entitlement to FC just to zero), who tended to be working full time (Blundell et al., 2000). The expected impact of the WFTC reform on lone parents' labour supply, conditional on working 16 or more hours, is as follows: (i) people receiving the maximum FC award will face an income effect away from work, but not below 16 hours a week; (ii) people working more than 16 hours and not on maximum FC will face an income effect away from work (but not below 16 hours a week), and a substitution effect towards work; (iii) people working more than 16 hours and earning too much to be entitled to FC but not WFTC ("windfall beneficiaries") will face income and substitution effects away from work if they claim WFTC (see Blundell and Hoynes, 2004; Blundell et al., 2004; Brewer et al., 2005).[7]

The occurrence of such reforms (i.e., the 1992 fall in hours requirement for FC, the 1995 additional credit for working full time, and the introduction of the WFTC programme in 1999) means that we can divide our sample into three periods: (a) Autumn 1991 to (March) 1995, with the post-reform period (which in our analysis we label FC, i.e., under the in-work benefit regime of FC) covering the years 1992-94; (b) April 1995 to September 1999, with the post-reform period (labelled FC+) being defined over the years 1995-97; and (c) October 1999 to the end of the

[7] It is worthwhile noticing that, for all three reforms, work incentives were likely to be dampened for single mothers living in areas with high childcare costs or high house rents (e.g., London and the South East of England). The availability of a more generous childcare tax credit component under WFTC might reduce this problem (Francesconi and Van der Klaauw, 2004), although high and increasing rents had to be weighed within the trade-off between additional tax credit gains and lower Housing Benefit entitlements (Gregg and Harkness, 2003). In Section IV.B we will present and discuss estimation results obtained after stratifying the sample by child's age, housing tenure, and region of residence.

sample, with the post-reform period (labelled WFTC) being between 1999 and 2002.[8] In our empirical analysis we take advantage of each of these separate reforms, as they have potentially (and differently) affected single mothers' hours of work. However, although we use this three-group categorization, most of our analysis will only isolate the 1992 and 1999 reforms (as the additional credit under FC+ was small), and focus on the few years immediately following the introduction of each policy change.

B. Analytical Framework

To assess whether female labour supply adjustments operate through job changes in response to exogenous changes in the incentives to work a given number of hours, we estimate four different specifications of a simple model of hours changes. We perform this assessment using a difference-in-difference method (Ashenfelter, 1978; Heckman and Robb, 1985): that is, we identify the FC and WFTC effects on single mothers' behaviour through the differential tax and benefit treatment that they receive as compared to a control group, which is given by single women without children. The main identification condition underlying this approach is that, other than the introduction of the changes in in-work benefits, there are no contemporaneous shocks that affect the *relative* outcomes of the treatment and control groups.[9]

Let Δh_{it} denote the change in total (usual and overtime) weekly hours of work between year t-1 and year t; let d_{it-1} be a dummy variable that is equal to 1 if woman i is a lone mother at time t-1, and 0 otherwise; and let Q_{it} be equal to 1 if woman i changes a job between year t-1 and t, and zero otherwise. The four specifications are as follows:

[8] In Section IV.A we shall return to the definition of the post-reform periods. Brewer (2001) has a detailed time-line of reforms to in-work benefits between 1971 and 2000. This does not reflect the reforms in April 2003 which lie outside our sample, and which are described in Brewer (2003).

[9] In fact, there were other shocks that might have influenced single mothers' and childless women's labour supply differently. For example, there was an increase in basic child benefits under Income Support (the main welfare benefit, similar to AFDC or TANF in the United States) between 1998 and 1999, which may be problematic in this respect. In terms of labour supply, however, this increase implies a negative income effect that could lead to a downward bias in our effect estimates. Our estimates may then represent a lower bound of the true effect.

$$\Delta h_{it} = \alpha_0 + \alpha_1 d_{it-1} + \alpha_2 Q_{it} + \beta_{FC} d_{it-1} Q_{it} \mathrm{I}(1992 \le t \le 1994) + \beta_{WFTC} d_{it-1} Q_{it} \mathrm{I}(1999 \le t \le 2002) + \mathbf{X}'_{it}\gamma + \varepsilon_{it} \quad (1)$$

$$\Delta h_{it} = \alpha_0 + \alpha_1 d_{it-1} + \alpha_2 Q_{it} + (\alpha_{31} + \alpha_{32} d_{it-1})\delta(t) + \beta_{FC} d_{it-1} Q_{it} \mathrm{I}(1992 \le t \le 1994) + \beta_{WFTC} d_{it-1} Q_{it} \mathrm{I}(1999 \le t \le 2002) + \mathbf{X}'_{it}\gamma + \varepsilon_{it} \quad (2)$$

$$\Delta h_{it} = \alpha_0 + \alpha_1 d_{it-1} + \alpha_2 Q_{it} + (\alpha_3 + b_{FC} d_{it-1})\mathrm{I}(1992 \le t \le 1994) + (\alpha_4 + b_{WFTC} d_{it-1})\mathrm{I}(1999 \le t \le 2002) + \beta_{FC} d_{it-1} Q_{it} \mathrm{I}(1992 \le t \le 1994) + \beta_{WFTC} d_{it-1} Q_{it} \mathrm{I}(1999 \le t \le 2002) + \mathbf{X}'_{it}\gamma + \varepsilon_{it} \quad (3)$$

$$\Delta h_{it} = \alpha_0 + \alpha_1 d_{it-1} + \alpha_2 Q_{it} + \alpha_{21} Q_{it} \mathrm{I}(1992 \le t \le 1994) + \alpha_{22} Q_{it} \mathrm{I}(1999 \le t \le 2002) + (\alpha_3 + b_{FC} d_{it-1})\mathrm{I}(1992 \le t \le 1994) + (\alpha_4 + b_{WFTC} d_{it-1})\mathrm{I}(1999 \le t \le 2002) + \beta_{FC} d_{it-1} Q_{it} \mathrm{I}(1992 \le t \le 1994) + \beta_{WFTC} d_{it-1} Q_{it} \mathrm{I}(1999 \le t \le 2002) + \mathbf{X}'_{it}\gamma + \varepsilon_{it}, \quad (4)$$

where I(w) is a function indicating that the event w occurs; $\delta(t)$ in equation (2) is a linear time trend; $\mathbf{X}_{it}$ is a vector of individual characteristics measured either at t-1 or between t-1 and t; and ε_{it} is an i.i.d. error term. The variables in $\mathbf{X}$, described in detail in the next section, are a cubic polynomial in total work experience, dummy variables for race, educational qualification, firm size, public sector, region of residence, and industry, the number and changes in the number of children by age group, and changes in health status, housing tenure, union coverage, and local unemployment rate.[10] The treatment effects for movers are captured by β_{FC} and β_{WFTC}, while b_{FC} and b_{WFTC} respectively capture the FC and WFTC treatment effects for workers who did not change job (stayers).

The key differences across equations (1)-(4) involve the specification of time trends. In equation (1), time trends are not modelled, except those operating through β_{FC} and β_{WFTC}. Equation (2) instead allows for group-specific linear time trends (captured by α_{31} and α_{32}), while in equation (3), we have a more flexible specification with group-specific discrete jumps for stayers after both the 1992 and 1999 reforms (b_{FC} and b_{WFTC}). Finally, equation (4) introduces even greater flexibility by allowing different trends in job changing behaviour after each reforms (through α_{21} and α_{22}). If $\hat{b}_j = \hat{\beta}_j$ (with $j = FC, WFTC$), we cannot reject the hypothesis of within-job flexibility

[10] The (levels of) time-varying variables are all measured at t-1.

in hours choice, while if $\hat{b}_j$ is statistically smaller than $\hat{\beta}_j$ there is evidence of hours constraints within jobs.

Estimation of (1)-(4) is performed using ordinary least squares (OLS). However, because our regressions are in changes,[11] all individual time-invariant permanent unobservables that enter additively in the determination of hours levels are eliminated from the estimation. In computing the standard errors we take advantage of the fact that we have multiple observations over time, and thus we allow for arbitrary serial correlation.

III. Data

The data we use come from the first twelve waves of the British Household Panel Survey (BHPS) collected over the period 1991-2002. Since Autumn 1991, the BHPS has annually interviewed a representative sample of about 5,500 households covering more than 10,000 individuals. All adults and children in the first wave are designated as original sample members. On-going representativeness of the non-immigrant population has been maintained by using a 'following rule' typical of household panel surveys: at the second and subsequent waves, all original sample members are followed (even if they moved house or if their households split up), and there are interviews, at approximately one-year intervals, with all adult members of all households containing either an original sample member, or an individual born to an original sample member whether or not they were members of the original sample. The sample therefore remains broadly representative of the population of Britain as it changes over time.[12]

Our estimation sample includes employed unmarried non-cohabiting females (separated, divorced, widowed and never married) who are at least 16 years old and were born after 1941 (thus

[11] Women with zero hours are excluded from our analysis. For further discussion on this point, see Section III.
[12] Of the individuals interviewed in 1991, 88 percent were re-interviewed in wave 2 (1992). The wave-on-wave response rates from the third wave onwards have been consistently above 95 percent. Detailed information on the BHPS can be obtained at ⟨http://www.iser.essex.ac.uk/ulsc/bhps/doc/⟩. The households from the European Community Household Panel subsample (followed since the seventh wave in 1997), those from the Scotland and Wales booster subsamples (added to the BHPS in the ninth wave) and those from the Northern Ireland booster subsample (which started in wave 11) are excluded from our analysis.

aged at most 60 in 2002). Because with equations (1)-(4) we estimate changes in hours worked, women are required to be in the labour market and report valid hours information for at least two consecutive years. We exclude any female who was long-term ill or disabled, and in school full time or self-employed or out of the labour force in a given year. The sample includes 2,284 women who have been observed working at least two consecutive times over the sample period and at some point were living alone, of whom 1,122 are lone mothers and the remaining 1,162 are childless. In line with the Inland Revenue's definition, a child must be aged 16 or less (or be under the age of 19 and in full-time education) to count as a dependent child for whom the single mother is responsible. Although only 16 percent of the women are observed in the same marital state for all the 12 years of the panel, about 60 percent of them are observed for at least 7 years in the same state. The resulting sample size, after pooling all available years for both groups of women, is 12,359 observations (4,585 on lone mothers and 7,774 on childless women). Of the 1,280 single women in the 1999 wave of interviews, 25 lone mothers and 32 childless women (about 4.5 percent of the sample in that year) were interviewed just before the day in which the 1999 reform was implemented (October 5th). To limit problems of interpretation, they were dropped from the estimating sample. Their inclusion however does not alter any of our main results.

Appendix Table A1 presents summary statistics of the outcomes and characteristics of the two groups of women, which we use as covariates in the empirical analysis below. There are some noticeable differences between the two groups. Lone mothers are younger (30 versus 38 years), less educated (56 percent have qualifications short of A level versus 48 percent among childless women, and only 6 percent of lone mothers have a university degree versus 14.4 percent), more likely to be nonwhite (9 versus 4.3 percent) and in social housing (38 versus 23 percent), less likely to be employed in the public sector (17 percent versus 25 percent), and have fewer years of work experience. The two groups of women are instead relatively similar in terms of job changing behaviour with 17 percent of childless women and 18 percent of single mothers moving across employers in any two given years.[13] Systematic differences emerge again in the case of labour market outcomes. Compared to unmarried women without children, lone mothers work about 9 fewer hours per week, earn £1.20 less per hour, and nearly £420 less per month, and report a larger change in worked hours from one year to the next (an increase of 2 hours and a quarter per week versus less than 25 minutes).[14]

Figure 1 plots the time trends for the year-on-year changes in hours worked over the sample period. Panel (a) shows the trends for all working women distinguishing lone mothers from single childless women, while panels (b) and (c) display the trends for female workers who moved between jobs and for workers who stayed with the same employer respectively. The data reveal that changes in hours worked among unmarried women without children are small and stable, ranging between 0 and 1 hour per week over the entire period (panel (a)). The hours changes for lone mothers instead are greater and their time variability is higher too. The largest change is observed after the introduction of the WFTC between 1998 and 1999, when lone mothers reported an increase of about 4.5 hours of work per week.[15] But, after 1999, lone mothers seem to have adjusted their hours changes downward. The 1992 reform, which reduced the hours requirement for FC eligibility from 24 to 16 per week, increased single mothers' labour supply by about 2.5 hours, but again this increase did not last in subsequent years. The additional FC for those working 30 or more hours does not appear to be associated with substantial changes in hours worked immediately after

[13] Our measure of job change does not include internal promotions or job changes within the same firm, but includes all moves to other firms (either through quits or layoffs). Alternative definitions of job change (e.g., dropping laid-off workers from the pool of movers, or dropping promoted workers from the group of the stayers) seem to produce similar results to those reported in this paper. See also Section IV.C.

[14] To account for potential differential attrition over the panel and individual/item nonresponse in each specific wave, we recomputed group-specific means using weighted data (with either cross-sectional or longitudinal enumerated individual weights). The results (not shown) are very similar to those obtained with unweighted data and presented in Appendix Table A1, suggesting that the problems induced by panel attrition and changing sample composition are likely to be relatively small in our data. We shall return to some of these issues while performing sensitivity analysis (see Section IV.C).

[15] This contrasts with the estimates reported in Stewart and Swaffield (2004), which provide evidence of female labour supply reduction of 1-2 hours per week as a result of the introduction of the National Minimum Wage in April 1999. These results are not robust across data sets and specifications, and are obtained from data that stop in September 2000 at the latest (that is, just before the second post-WFTC year in our sample). In addition, as pointed out in the Introduction, these estimates refer to all women, so we do not know how single women with and without children have been differentially affected by the minimum wage.

its introduction in 1995, but it is followed by a steady increase even before the peak between 1998 and 1999.

Panel (b) shows that the largest changes are experienced by women who moved between jobs, with lone mothers reporting an average change in hours of about 4 per week over the whole sample period and unmarried women without children of 1 per week. The time patterns for lone mothers are similar to those reported in the previous panel, although the peak in 1998-99 is followed by a further increase over the subsequent year. Lone mothers' increase in hours between 1991 and 1992 is also sizable, with an average close to 4 hours per week. Hours changes among those who stayed with the same employer instead are much smaller for both groups of women, especially for women without children (panel (c)).

IV. Results

A. Benchmark estimates

The estimates of the impact of job changing behaviour on hours changes are shown in Table 1. These are presented for the four specifications described in Section II.B, and separately for the cases in which the variables in $\mathbf{X}$ are excluded or included.

The regressions without controls indicate that changing job is associated with increases in women's labour supply by less than one hour per week (α_2), although this effect is only significant at the 10 percent level in the first three specifications, while single mothers experience significantly larger labour supply changes of about 1.5 hours per week (α_1). The treatment effects for stayers $(b_{FC}$ and $b_{WFTC})$ are small and never statistically significant, and so are the average treatment effects for job movers after the 1992 reduction in hours requirement under FC (β_{FC}). But the introduction of WFTC had a strong impact on job movers with a significant increase of their labour supply by 2.5-3 hours per week on average. Importantly, from specification (4) we can reject the hypothesis that $\hat{b}_{WFTC} = \hat{\beta}_{WFTC}$ at the 5 percent level (the *p*-value of the *t*-test of equality is 0.024), which provides evidence of hours inflexibility within jobs. Most of these results are robust to the inclusion of the control variables $\mathbf{X}$, with the only exception of α_1 which now becomes statistically insignificant.[16] With 16 and 20 percent of women changing job after the 1992 and 1999 reforms respectively, we can derive their overall effects on labour supply changes: FC had virtually no impact, while WFTC increased single mothers' weekly hours of work by about 1.2 hours irrespective of whether they changed job or not (specification (4)).

Because Figure 1 reveals that stayers also increased their worked hours immediately after the 1999 reform, we repeated the previous analysis excluding the last two years of the sample. Indeed, the WFTC effect for stayers is now larger and close to one extra hour per week, but its *p*-value is never below 0.11. In any case, even after this selection, all other results are confirmed, including the rejection of the hypothesis of complete flexibility in hours within jobs.[17] Thus, in response to the exogenous change in work incentives given by the WFTC programme, changing job seemed to have been the strongest mechanism of labour supply adjustment among single mothers after 1999.

B. Heterogeneous responses

It is possible that the labour supply responses to the policy reforms vary by observable characteristics of the women in the treatment and control groups. To allow for this, we look for heterogeneous responses by estimating models that distinguish women separately by individual attributes (such as education, number and age of their children, region of residence, and house tenure), work related attributes (such as firm size, industry, and sector), and stated labour supply preferences. The results from these regressions (based on specifications (3) and (4) only) are reported in Tables 2-8.

[16] To understand this lack of effect, we estimated simple variants of equations (1)-(4) with Q interacted with marital status (not shown). Regardless of whether we control for group-specific time trends, changing job is associated with increases for single mothers to their labour supply by about 1.2 hours per week and with reductions for single childless women by about 0.8 hours per week.

[17] We reach the same conclusion if we keep the entire sample as in Table 1 but redefine the post-reform period over either 1999-2000 or 1999-2001. Similarly, redefining the FC period over 1992-93 (rather than 1992-94) does not alter our baseline results.

The treatment effects for stayers do not differ between more educated women and less educated women. There are however asymmetric responses among movers. Changing job increased less educated single mothers' labour supply by 3-4 hours per week after the 1992 reform, and increased more educated single mothers' supply by 4-5 hours after the 1999 reform.[18] Equality tests of the estimated b and β coefficients from specification (4) can be rejected at the 5 percent level during the WFTC regime among the more educated, and at the 10 percent during the FC regime among the less educated.

Tables 3 and 4 demonstrate that the post-WFTC upward adjustment in single mothers' labour supply is primarily experienced by mothers of one child aged 0-4. Albeit smaller, the effect observed for mothers of children aged 5 or more is still sizeable and significant (Table 4). If we pool all women as we did in Table 1 and interact the variable on b_{FC} with the indicator of the youngest child being aged 0-4, this interaction term is negative and statistically significant $(b_{FC} = -1.35$ and $\text{s.e.} = 0.48,$ specification (4)), while the interaction with the indicator of the youngest child being older is never significant. This provides evidence that the 1992 reform induced some groups of women (in this case, single mothers of young children who did not change job) to reduce their labour supply over the 1992-94 period.

The UK in-work benefit system interacts with other welfare benefits (Blundell and Hoynes, 2004). One of these is Housing Benefit, which works as a rent subsidy. If a single mother receives Housing Benefit, she would benefit less from a given amount of tax credit because this is treated as income in other means-tested programmes. Rents in some parts of the country (in particular, London and the South East) are high and have rapidly increased over the 1990s, while owner-occupiers are not eligible to Housing Benefit. To capture part of the relationship between Housing

[18] If a large proportion of better-educated single mothers had not been eligible to WFTC, the effects reported in Table 2 should be attributed to shocks other than WFTC. However, using data from the Family Resources Survey, we find that tax credit eligibility has increased proportionally more for more educated lone mothers than for the less educated after the introduction of WFTC (albeit a greater fraction of the less educated are eligible). In particular, between 1995 and 1998, about 26 percent of better educated lone mothers who work 16 or more hours per week were eligible to FC. Between 2000 and 2002, 49 percent were eligible to WFTC (an increase in eligibility rate by 88 percent). For the less educated, the increase in eligibility rate was only 28 percent (from 65 to 83 percent).

Benefit and the tax credits of interest here, we stratified our sample by region of residence (London and the South East in one group and the rest of the country in the other) and by housing tenure (owner-occupier or not), both measured at *t*-1. For the sake of brevity, the results are not shown but are available from the authors. From this analysis it emerges that labour supply adjustments observed after the 1999 reform were greater for single mothers who lived outside the London/South-East region, and who were not owner-occupiers.[19]

Job specific characteristics provide other important sources of heterogeneity for the impact of job changes on hours changes after the 1999 reform. The strongest treatment effects are found for single mothers employed in relatively larger establishments (of the order of 4 additional hours per week, Table 5), in service industries (between 3 and 4 extra hours, Table 6),[20] and equally for those employed in the private sector or the public sector (between 2.5 and 4 additional hours per week, Table 7). Strong evidence of hours inflexibility emerges among lone mothers who work in larger firms, service industries, and the public sector.

Another important dimension along which we expect to see heterogeneous responses is given by stated labour supply preferences. At each interview, the BHPS asks respondents whether they would like to work fewer hours, or more hours, or continue to work the same number of hours "assuming that they would be paid the same amount per hour". We use this information to construct three labour supply preference variables for any given year of the sample period, labelled OVER (=1 if a worker would like to work fewer hours, and zero otherwise), UNDER (=1 if a worker would like to work more hours, and zero otherwise) and SAME (=1 if a worker would like to

[19] Stratifying the sample jointly by region and house tenure leads to small subsamples. But when we performed the analysis on the entire sample and included an interaction term between these two variables, the largest increases in worked hours occurred in association with changing job after the introduction of WFTC for single mothers who lived in rented accommodations outside the London/South-East region.

[20] Single mothers who were employed in manufacturing industries also showed a significant increase of 2-3 hours of work per week if they changed job after the introduction of WFTC (Table 6). For the same group of women there is also evidence (significant only at the 10 percent level) of positive labour supply adjustments of about 1.5 hours per week if they changed job between 1992 and 1994 (i.e., during the FC regime). This effect involves only 25 percent of the whole sample, and this may be why it does not show up in the whole sample. Manufacturing production is based on technologies that are traditionally less flexible than those used in services, such as batch methods and robotized assembly lines (Goldin and Katz, 1998), which may be reflected in a greater rigidity in (downward) adjustments in hours.

continue to work the same number of hours, and zero otherwise).[21] We expect that workers who are overemployed/underemployed at one point in time reduce/increase their worked hours over time, and those who want to continue working the same number of hours do not change their labour supply. The estimates on α_2 reported in Table 8 confirm such expectations, with overemployed workers reducing their labour supply by 3 hours per week on average, underemployed workers increasing it by about 4 hours, and the remaining group of workers showing no significant change. The 1992 and 1999 in-work benefits reforms did not affect hours worked by women who would have liked to keep working the same number of hours and did not change job. But single mothers who wanted to continue working the same number of hours showed large upward labour supply adjustments of about 3-4 hours per week if they changed job after the WFTC reform.[22] Thus, initially "unconstrained" (i.e., neither over- nor under-employed) lone mothers did respond to the greater work incentives of the WFTC programme but only through a change of job. This upholds our previous finding that there is evidence of hours inflexibility within jobs.

The 1999 reform also led to increases of 1-3 hours per week among both overemployed and underemployed workers who changed job, although none of such increases is statistically significant at conventional levels. After the 1992 reform, instead, we observe large (and significant at the 10 percent level) reductions of about 6-7 hours per week among overemployed single mothers who changed job. This lines up very well with the 8-hour fall in the minimum work requirement to receive FC (from 24 to 16 hours a week). Again, this labour supply adjustment occurs through movements across (rather than within) jobs, although equality tests of the estimated b and β coefficients can be rejected only at the 10 percent level, irrespective of the specification. Underemployed workers seem to be unable to adjust their labour supply upward if they did not

[21] Over the whole sample period, about 19 percent of lone mothers report being overemployed, 18 percent underemployed, and the remaining 62 percent report being satisfied with their hours of work. The corresponding proportions for single women without children are 28, 11, and 61 percent. Considering all women in the sample, the most mobile are the underemployed (with 27 percent of them changing job in any two consecutive years), while the job changing rates for the overemployed and the other group of workers are lower (19 and 15 percent respectively).

[22] The hypothesis that the estimated b and β coefficients are equal can be rejected at the 5 percent level in specification (4) (p-value = 0.027).

change job. But those who moved did manage to increase their worked hours even after the 1992 reform by about 2-3 hours per week (although this increase is not statistically significant).

We reestimated variants of equations (1)-(4) over the whole sample of women which included interaction terms between the variables on b_j and β_j $(j = FC, WFTC)$ and stated labour supply preferences. The results from this analysis (not shown) confirm those previously discussed. In particular (from specification (4)), unconstrained single mothers who changed job after the 1999 reform increased labour supply by about 4 hours (*t*-value=4.61), and overemployed single mothers who moved across jobs after the 1992 reform reduced their hours by about 7 a week (*t*-value=2.41). Further interactions with indicators of the age of the youngest child reveal that mothers of younger children (aged 0-4) who moved jobs experienced the greatest changes in labour supply. In particular, after the introduction of WFTC, unconstrained mothers whose youngest child was aged 0-4 and who changed job worked nearly 5.5 extra hours (*t*-value=3.27) as opposed to 3 among unconstrained mothers whose youngest child was aged 5-18. Similarly, after the FC reform, overemployed single mothers with younger children reduced their labour supply by 9 hours a week as compared to 5.6 among mothers of older children.

C. Sensitivity Analysis

We performed a number of sensitivity analyses to demonstrate the robustness of the results. For the sake of brevity we present the results from three exercises only.[23] First, we reestimated models (1)-(4) accounting for the 1995 FC reform that provided extra credit for full-time work. The estimates in Table 9 confirm our previous findings, and document that the 1995 reform was followed by no sizeable change in worked hours irrespective of whether women changed employer or stayed in the same job.

[23] We also reestimated the models eliminating laid-off workers from *Q*, or dropping promoted workers from the group of the stayers. Both these exercises produced results that were virtually identical to those shown in Table 1, and are thus not reported.

As mentioned in Section III, there may be concerns with changing sample composition over time, differential attrition, and missing data. Besides using weighted data, which provided similar results to those presented so far, we addressed these concerns by reestimating our models only on women who have been successfully interviewed for a given number of times (for example, six or more waves). If attrition or changing sample composition are important, the results from such selected subsamples are expected to differ from those discussed earlier. Table 10 reports the estimates found from two subsamples, one in which we include only women who have been observed for six or more years (i.e., at least half of the time between 1991 and 2002) and the other in which women have to be observed for at least 9 consecutive times. Regardless of the specification, the estimates from both subsamples support our previous results, suggesting that missing data problems are likely to have only minor consequences for our analysis.

Finally, we estimated the effects using propensity score matching (biweight kernel and local linear regression matching). Although, like standard OLS regressions, matching methods rely on a selection-on-observables assumption (Angrist and Krueger, 1999), they limit the potential bias due to differences in the support of $\mathbf{X}$ between single mothers and women without children and the bias due to the difference between the two groups of women in the distribution of $\mathbf{X}$ over its common support (Heckman et al., 1998). The estimates in Table 11 display patterns that are very similar to those illustrated above in this section.

D. Wage Estimates

The evidence so far indicates that British single mothers responded to the greater work incentives of the 1999 in-work benefits reform by substantially increasing their labour supply, whereas the two previous reforms to FC seemed to have induced only minor labour supply effects. The strong labour supply adjustment in conjunction with the introduction of WFTC was primarily achieved through a change of employer rather than changes in hours within the same job. This finding suggests that single mothers face some form of hours inflexibility within jobs. Against this background, we analyse wage responses. Of course, in-work benefits reforms were designed to change worked hours directly while wage determination is also under employers' control, and so our partial-equilibrium analysis is likely to provide biased estimates. Nonetheless, gauging wage responses is important because it gives us a more complete picture of the British labour market and some indication of the possible presence of labour market imperfections. We therefore estimated equations (1)-(4) with log hourly wages (expressed in 2002 prices) as dependent variable and the same set of explanatory variables used before. A number of checks, which were performed to test the robustness of such specifications, led to results that have the same qualitative implications as those reported here.

For both job movers and stayers and both the 1992 and 1999 reforms, we find no significant wage effect. There is also relatively little effect heterogeneity across different groups of women. Two important exceptions however are single mothers who lived in London and the South East and those who worked in small establishments. Among the former group of women, changing job after the introduction of WFTC implied not only a labour supply increase of almost 3 hours per week (*t*-ratio=3.11), but also a wage reduction of 2.7 percent (*t*-value=2.23). Among the latter, changing job after the 1999 reform led to 1.5 percent lower wages (*t*-value=1.51) and modest positive hours changes (see Table 5). Thus, despite the presence of hours inflexibility, the labour market generally operates quite competitively, although there is an indication of monopsony among some groups of single mothers.

V. Conclusions

By using three in-work benefits reforms during the 1990s in the UK, which either changed hours requirements to be eligible for the benefits or increased the attractiveness to work a given number of hours, we are able to assess the canonical labour supply model of complete hours flexibility, and analyse the path of labour supply adjustment among single women with and without children. We find that the 1992 and 1995 FC reforms had modest impacts on single mothers' labour supply, but the introduction of the WFTC reform in 1999 had large positive effects on their number of hours of work. This increase is largely driven by women who changed job, suggesting that labour supply adjustments within a job are much less flexible. This lines up well with the estimates we get when we look at hours changes by stated labour supply preferences: unconstrained women who changed job showed the largest labour supply increases after the 1999 reform, and overemployed women substantially reduced their labour supply after the 1992 reform (which did reduce the minimum work requirement to receive FC from 24 to 16 hours a week) only if they moved across jobs. There is evidence of considerable heterogeneity in the effects of the WFTC reform for different groups of women. The strongest evidence of hours inflexibility emerged among single mothers whose youngest child was aged 0-4. This was especially the case for those who worked in larger firms, service industries, and the public sector. Although there is little in the way of overall wage effects, we do find that after the introduction of WFTC hourly wages decreased significantly for single women who lived in London and the South East and moved jobs and, to a lesser extent, for movers who worked in small firms.

So what remains of the canonical labour supply model? We have shown that labour supply adjustments in hours of work are made primarily by movements *between* jobs and there is little evidence of hours flexibility *within* jobs. Our analysis of stated preferences confirms this further showing that responses are highest among those who say they are unconstrained but also among those who are constrained but state that they would like to move in the direction suggested by the incentives. Thus, a labour supply model emerges in which labour supply adjustments are largely made by moving between workplaces. This could be achieved within an “adapted” canonical model in which establishments are organised around hours requirements and individuals move jobs to achieve hours flexibility. Of course, it could be also supported by theories that emphasise the importance of labour market frictions and imperfections, such as job search, wage-job packages, and/or dynamic monopsony. However, if there were such ‘imperfections’ we would expect these to be displayed in wage responses. The evidence is that such responses are not large and overall not statistically significant. Consequently, at least to a first approximation, an adapted canonical labour supply model with hours flexibility across jobs cannot be rejected. Nonetheless, our results by region, industry, and firm size suggest that employer preferences may not only reduce labour supply flexibility within firms but may also place constraints on hours mobility across firms.

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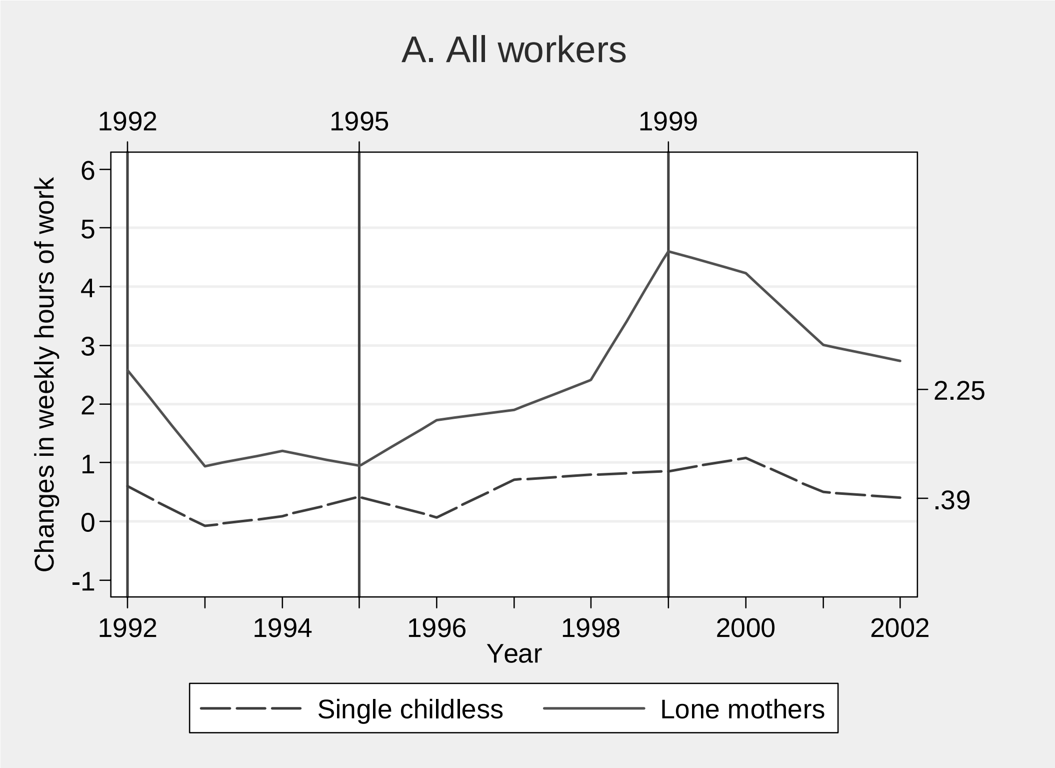
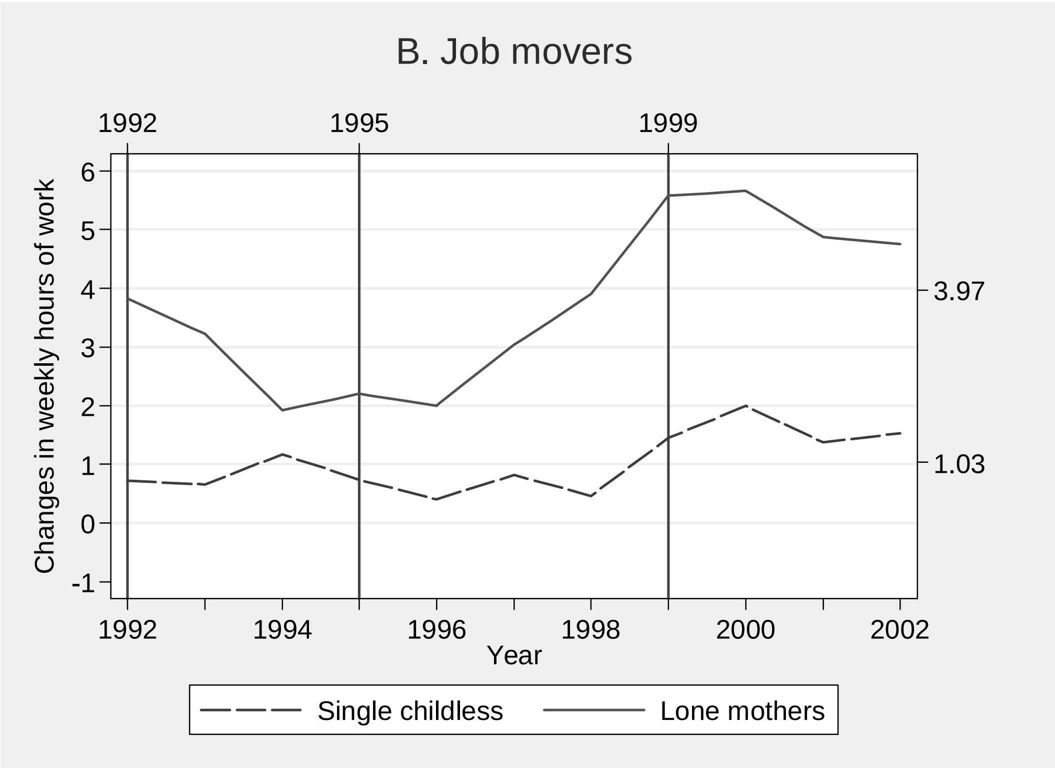
Figure 1
Changes in worked hours
Single childless women and lone mothers by job changing status





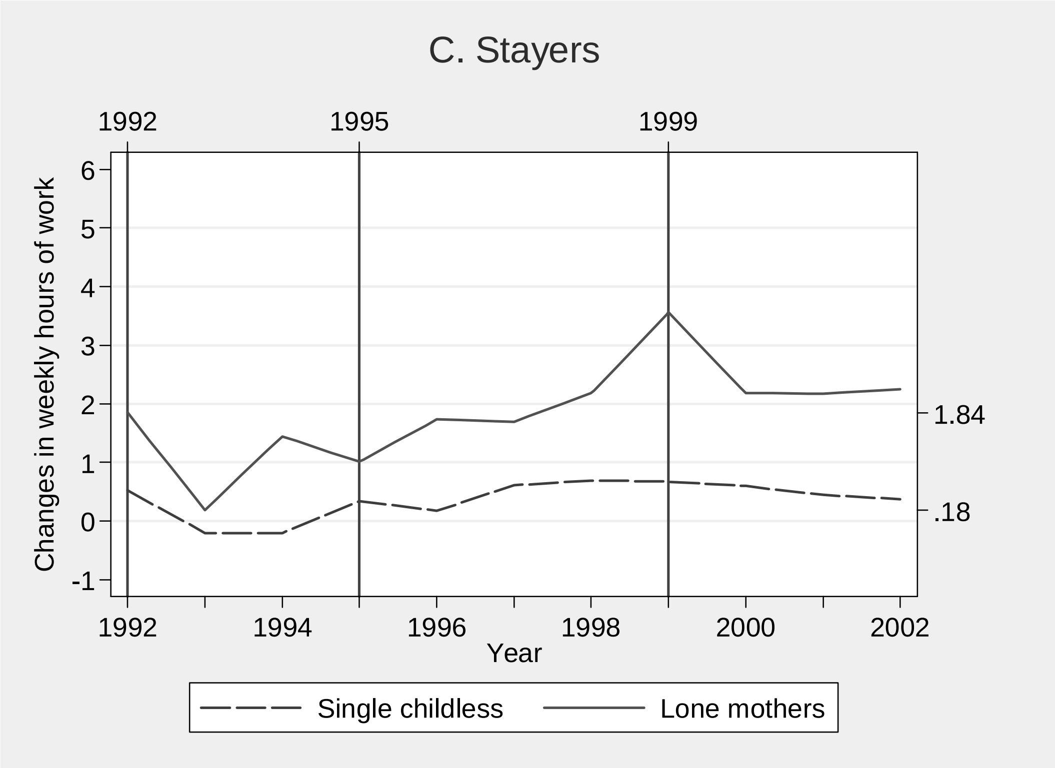

C. Stayers
1992
1995
1999
Changes in weekly hours of work
6
5
4
3
2
1
0
-1
1.84
.18
1992
1994
1996
1998
2000
2002
Year
Single childless
Lone mothers

Table 1. The Impact of the In-Work Benefit Reforms and Job Changes on Hours Changes

	Without controls				With controls			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
α_1	1.58	1.27	1.54	1.36	0.34	0.44	0.24	0.19
	(7.94)	(2.68)	(3.51)	(2.98)	(1.08)	(0.85)	(0.59)	(0.32)
α_2	0.77	0.74	0.73	-0.19	-0.30	-0.32	-0.31	-0.45
	(1.86)	(1.78)	(1.76)	(0.28)	(0.64)	(0.70)	(0.71)	(0.82)
b_{FC}			-0.25	0.03			-0.03	-0.21
			(0.40)	(0.07)			(0.01)	(0.44)
b_{WFTC}			0.16	0.45			0.20	0.56
			(0.33)	(0.89)			(0.42)	(0.94)
β_{FC}	0.11	0.72	0.95	0.48	0.21	0.83	0.89	0.44
	(0.08)	(0.51)	(0.67)	(0.29)	(0.15)	(0.59)	(0.62)	(0.28)
β_{WFTC}	2.56	2.66	2.48	3.39	2.54	2.65	2.60	3.42
	(2.46)	(2.56)	(2.29)	(2.82)	(2.51)	(2.63)	(2.47)	(2.92)
Number of observations	12,359				12,359			

Source: British Household Panel Survey 1991-2002.
Notes: Absolute values of *t*-statistics (obtained from standard errors that allow for arbitrary serial correlation) are in parentheses. The labelling of columns (1)-(4) corresponds to equations (1)-(4) described in the text.

Table 2. Effects by education group

	Less educated		More educated	
	(3)	(4)	(3)	(4)
α_1	-0.22	-0.31	0.73	0.56
	(0.44)	(0.61)	(1.13)	(0.83)
α_2	-0.49	-0.64	-0.12	-0.25
	(0.78)	(0.82)	(0.21)	(0.39)
b_{FC}	0.54	0.51	-0.70	-0.45
	(0.86)	(0.85)	(0.82)	(0.23)
b_{WFTC}	0.85	0.53	0.17	0.58
	(0.96)	(0.88)	(0.04)	(0.77)
β_{FC}	2.79	3.61	-1.27	-2.65
	(2.28)	(2.15)	(0.86)	(1.52)
β_{WFTC}	1.62	1.92	3.95	4.89
	(1.16)	(1.26)	(2.55)	(2.88)
Number of observations	6,297		6,062	

Note: ‘Less educated’ is defined as having less than A level qualifications; ‘More educated’ is defined as having A-level or higher qualifications. Education is measured at time *t*-1. All regressions include the control variables used in Table 1 (except for education). For type of specification and other definitions, see Table 1.

Table 3. Effects by number of children

	One child (3)	One child (4)	Two or more children (3)	Two or more children (4)
α_1	0.62	0.60	-0.43	-0.29
	(1.08)	(1.07)	(0.71)	(0.66)
α_2	-0.59	-0.97	-0.08	-0.16
	(0.63)	(1.49)	(0.23)	(0.27)
b_{FC}	0.25	0.47	-0.44	-0.57
	(0.41)	(0.78)	(0.56)	(0.88)
b_{WFTC}	0.13	0.39	0.29	0.53
	(0.22)	(0.64)	(0.46)	(0.84)
β_{FC}	-0.41	-0.47	2.19	1.98
	(0.23)	(0.44)	(1.08)	(0.92)
β_{WFTC}	3.31	4.15	1.25	1.37
	(2.55)	(2.95)	(1.04)	(1.31)
Number of observations	6,427		5,932	

Note: The categories 'One child' and 'Two or more children' pertain to lone mothers. The number of children is measured at time *t*-1. All regressions include the control variables used in Table 1 (except for number and change in the number of children by age group). For type of specification and other definitions, see Table 1.

Table 4. Effects by age of youngest child

	Youngest child aged 0-4 (3)	Youngest child aged 0-4 (4)	Youngest child aged 5 or more (3)	Youngest child aged 5 or more (4)
α_1	-0.28	-0.33	0.59	0.62
	(0.30)	(0.37)	(1.03)	(1.09)
α_2	-0.11	-0.27	-0.55	-0.76
	(0.21)	(0.36)	(1.27)	(1.45)
b_{FC}	-0.75	-0.59	0.25	0.03
	(1.07)	(0.78)	(0.47)	(0.82)
b_{WFTC}	-0.11	0.30	0.57	0.64
	(0.20)	(0.19)	(1.10)	(1.05)
β_{FC}	-0.49	-0.51	1.21	0.97
	(0.31)	(0.45)	(0.79)	(0.56)
β_{WFTC}	3.06	3.82	2.63	2.80
	(2.70)	(2.84)	(2.35)	(2.63)
Number of observations	5,438		6,921	

Note: The categories 'Youngest child aged 0-4' and 'Youngest child aged 5 or more' refer to lone mothers. The age of the youngest child is measured at time *t*-1. All regressions include the control variables used in Table 1 (except for number and change in the number of children by age group). For type of specification and other definitions, see Table 1.

Table 5. Effects by firm size

	Fewer than 50 employees		50 or more employees	
	(3)	(4)	(3)	(4)
α_1	0.18	-0.01	0.46	0.74
	(0.38)	(0.02)	(0.86)	(1.03)
α_2	-0.10	-0.38	-0.77	-0.93
	(0.18)	(0.61)	(1.04)	(1.52)
b_{FC}	0.25	0.47	-0.72	-0.38
	(0.44)	(0.84)	(0.57)	(0.38)
b_{WFTC}	0.24	0.61	-0.05	-0.16
	(0.46)	(1.08)	(0.07)	(0.13)
β_{FC}	0.34	0.37	0.94	0.72
	(0.20)	(0.25)	(0.34)	(0.41)
β_{WFTC}	2.27	2.18	3.67	4.20
	(1.93)	(1.84)	(2.50)	(2.64)
Number of observations	8,553		3,806	

Note: All regressions include the control variables used in Table 1 (except for firm size). For type of specification and other definitions, see Table 1.

Table 6. Effects by industry

	Services		Manufacturing	
	(3)	(4)	(3)	(4)
α_1	0.82	0.72	-0.56	-0.40
	(1.61)	(1.12)	(1.01)	(0.74)
α_2	-0.36	-0.84	-0.20	0.23
	(0.72)	(0.97)	(0.28)	(0.15)
b_{FC}	-0.11	-0.48	0.32	0.30
	(0.01)	(0.28)	(0.49)	(0.48)
b_{WFTC}	-0.27	0.33	0.99	0.86
	(0.30)	(0.51)	(1.48)	(1.27)
β_{FC}	-0.35	-0.82	1.39	1.57
	(0.65)	(0.88)	(1.93)	(1.91)
β_{WFTC}	3.08	3.74	2.23	2.95
	(2.76)	(3.15)	(2.08)	(2.77)
Number of observations	9,262		3,097	

Note: The category 'Services' includes banking, finance and insurance, distribution, hotels and catering, transport and communication, and other services (which include education and sanitary services). 'Manufacturing' includes energy, extraction, metal goods, other manufacturing industries, construction, and primary industries. Industry is measured at time *t*-1. All regressions include the control variables used in Table 1 (except for industry dummies). For type of specification and other definitions, see Table 1.

Table 7. Effects by sector

	Private sector		Public sector	
	(3)	(4)	(3)	(4)
α_1	0.03	-0.08	0.93	0.94
	(0.07)	(0.19)	(1.31)	(1.14)
α_2	0.10	-0.26	-0.79	-0.88
	(0.08)	(0.27)	(1.03)	(1.52)
b_{FC}	-0.14	0.09	0.51	0.43
	(0.10)	(0.16)	(0.76)	(0.68)
b_{WFTC}	0.28	0.55	-0.56	-0.18
	(0.53)	(0.91)	(0.46)	(0.15)
β_{FC}	1.18	1.08	-0.17	-0.52
	(0.66)	(0.54)	(0.46)	(0.76)
β_{WFTC}	2.49	3.09	2.82	4.10
	(2.06)	(2.43)	(2.18)	(2.71)
Number of observations	9,659		2,700	

Note: The category 'Public sector' includes civil service, central and local government, National Health Service, education, and non-profit organizations. Sector is measured at time *t*-1. All regressions include the control variables used in Table 1 (except for sector). For type of specification and other definitions, see Table 1.

Table 8. Effects by stated labour supply preferences

	SAME=1		OVER=1		UNDER=1	
	(3)	(4)	(3)	(4)	(3)	(4)
α_1	-0.45	0.07	1.90	1.65	0.35	0.31
	(0.52)	(0.17)	(1.56)	(1.19)	(0.25)	(0.42)
α_2	-0.42	-1.21	-2.72	-3.92	3.56	4.37
	(0.81)	(1.58)	(3.49)	(2.86)	(3.80)	(2.73)
b_{FC}	0.42	0.61	-0.83	-0.53	-0.15	-0.24
	(0.66)	(0.90)	(0.77)	(0.86)	(0.10)	(0.15)
b_{WFTC}	0.70	0.84	-0.24	-0.35	-0.36	-0.50
	(1.14)	(1.52)	(0.37)	(0.66)	(0.24)	(0.27)
β_{FC}	0.71	0.43	-7.03	-6.74	2.47	2.97
	(1.03)	(0.74)	(1.83)	(1.84)	(1.37)	(1.44)
β_{WFTC}	3.41	4.20	0.65	1.09	2.35	2.87
	(2.72)	(2.97)	(0.34)	(0.31)	(1.59)	(1.64)
Number of observations	7,539		3,090		1,730	

Note: OVER = 1 if the respondent indicated that she would like to work fewer hours "assuming that [she] would be paid the same amount per hour", and equals 0 otherwise; UNDER = 1 if the respondent indicated that she would like to work more hours "assuming that [she] would be paid the same amount per hour", and equals 0 otherwise; SAME = 1 if the respondent indicated that she would like to continue to work the same number of hours "assuming that [she] would be paid the same amount per hour", and equals 0 otherwise. Labor supply preferences are measured at time *t*-1. For type of specification and other definitions, see Table 1.

Table 9. Robustness check: Accounting for the 1995 FC reform that provided extra credit for full-time work (FC+)

	(1)	(2)	(3)	(4)
α_1	0.29	0.24	0.24	0.30
	(0.92)	(0.45)	(0.53)	(0.69)
α_2	-0.70	-0.72	-0.73	-0.56
	(1.64)	(1.70)	(1.70)	(0.88)
b_{FC}			0.25	0.12
			(0.90)	(0.11)
b_{FC+} (extra credit for FT work)			0.04	-0.22
			(0.10)	(0.44)
b_{WFTC}			0.21	0.24
			(0.41)	(0.48)
β_{FC}	0.65	1.05	1.02	0.81
	(0.47)	(0.83)	(0.91)	(0.52)
β_{FC+} (extra credit for FT work)	1.57	1.43	1.35	1.18
	(2.06)	(1.76)	(1.67)	(1.32)
β_{WFTC}	2.88	2.99	2.92	3.48
	(2.85)	(2.90)	(2.79)	(3.13)
Number of observations	12,359			

Table 10. Effects by length of time in the panel

	6 years or more		9 years or more	
	(3)	(4)	(3)	(4)
α_1	0.67	0.57	0.62	0.66
	(1.41)	(1.20)	(1.08)	(1.13)
α_2	-0.52	-0.81	-0.56	-0.60
	(1.30)	(1.56)	(1.16)	(0.88)
b_{FC}	-0.32	-0.45	-0.21	-0.27
	(0.57)	(0.80)	(0.31)	(0.40)
b_{WFTC}	0.62	0.81	0.68	0.75
	(1.25)	(1.54)	(1.10)	(1.02)
β_{FC}	0.76	0.65	0.39	0.17
	(0.68)	(0.50)	(0.61)	(0.23)
β_{WFTC}	3.97	4.52	2.89	2.99
	(3.26)	(3.34)	(2.46)	(2.39)
Number of observations	8,314		5,153	

Note: The categories '6 years or more' and '9 years or more' include only women who have been observed for at least 6 years and 9 years consecutively in the panel respectively. All regressions include the control variables used in Table 1. For type of specification and other definitions, see Table 1.

Table 11. Robustness checks: Propensity score matching models

	Biweight kernel matching		Local linear regression matching[a]	
	(3)	(4)	(3)	(4)
α_1	0.82	0.97	1.04	1.05
	(0.35)	(0.89)	(1.12)	(1.36)
α_2	-0.87	-1.12	-0.52	-0.20
	(0.71)	(1.53)	(0.33)	(0.51)
b_{FC}	-0.25	-0.13	0.20	-0.38
	(0.67)	(0.11)	(0.29)	(0.75)
b_{WFTC}	-0.31	-0.49	0.49	0.97
	(0.45)	(0.90)	(0.67)	(1.13)
β_{FC}	0.88	0.36	0.96	0.53
	(0.63)	(0.48)	(1.17)	(0.62)
β_{WFTC}	2.88	3.12	2.97	3.07
	(2.41)	(3.07)	(2.30)	(2.74)
Number of observations	12,359		12,359	

Notes: Absolute values of *t*-statistics (with standard errors obtained from 500 bootstrapped replications) are in parentheses.
[a] Estimates are obtained after imposing a tricube kernel.

Appendix Table A1. Summary statistics

Variable	Unmarried women without children	Lone mothers
Total weekly hours of work	34.74	25.61
	(13.25)	(14.20)
Change in worked weekly hours[a]	0.39	2.25
	(11.40)	(12.31)
Absolute change in worked weekly hours[a]	6.00	6.91
	(9.62)	(10.28)
Hourly pay	7.06	5.85
	(6.01)	(5.19)
Monthly labour income conditional on working positive hours (in 2002 pounds)	1,110	694
	(911)	(629)
Age (years)	38.1	30.00
	(15.0)	(11.36)
Nonwhite	0.043	0.090
Registered disabled	0.049	0.023
Number of children by age group:[b]		
0-4		0.231
		(0.510)
5-10		0.588
		(0.755)
11-18		0.798
		(0.771)
House owner	0.578	0.541
In social housing	0.229	0.377
A level or higher educational qualification	0.520	0.438
No qualification	0.152	0.144
University degree or more	0.144	0.060
Total work experience (years)	14.33	8.67
	(11.47)	(7.88)
Employed in a firm with fewer than 50 workers	0.660	0.746
Employed in service industries[c]	0.838	0.820
Employed in the public sector	0.247	0.171
Union covered	0.514	0.530
Changed job during previous year	0.167	0.179
Local unemployment rate[d]	0.065	0.063
	(0.032)	(0.031)
Number of person-wave observations	7,774	4,585
Number of women	1,162	1,122

Notes: The figures are means (standard deviations in parentheses) computed over all person-wave observations for which two consecutive years of data are available.
[a] The change is measured over two consecutive years.
[b] Averages are computed over the entire subsample of lone mothers. If computed over the three specific subsamples of lone mothers with children in each child group, the means (standard deviations) are 1.172 (0.448), 1.318 (0.582), and 1.321 (0.548), respectively.
[c] 'Service industries' refer to banking, finance and insurance, distribution, hotels and catering, transport and communication, and other services (which include education and sanitary services).
[d] Computed over 306 travel to work areas.