Job Changes and Hours Changes: Understanding the Path of Labor Supply Adjustment

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We use British panel data to investigate single women's labor supply changes in response to three reforms that affected individuals' work incentives. We use these reforms to identify changes in labor supply. There is evidence of small hours of work effects for two of such reforms. A third reform in 1999 instead led to a significant increase in single mothers' hours of work. The mechanism by which the labor supply adjustments were made occurred largely through job changes

Financial support from the Department of Trade and Industry Labour Market Flexibility Small Grant Fund 2004 and from the Economic and Social Research Council Centre for the Microeconomic Analysis of Public Policy at the Institute for Fiscal Studies is gratefully acknowledged. We thank Alison Booth, Mark Bryan, Jan van Ours, Arthur van Soest, and seminar participants at the Universities of Essex, Milan, Swansea, and Tilburg, Institute for Fiscal Studies, Bank of England, Bank of Spain, U.K. Department *for* Education and Skills (London), the 2005 European Society for Population Economics (Paris), and the Econometric Society World Congress 2005 (London) for helpful suggestions. We are responsible for all remaining errors and shortcomings of the article. Contact the corresponding author, Marco Francesconi, at mfranc@essex.ac.uk.

[*Journal of Labor Economics,* 2008, vol. 26, no. 3]

0734-306X/2008/2603-0004$10.00

rather than hours changes with the same employer. This is little overall effect of the reforms on wages.

I. Introduction

The use of the canonical model of labor 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-maximizing 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 by relying on observed individual characteristics (e.g., number and age of children, or job mobility) or stated labor supply preferences (Ham 1982; Moffitt 1984; Lundberg 1985; Altonji and Paxson 1988; Stewart and Swaffield 1997; Euwals 2001). These two strategies are problematic because changes in labor 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 labor 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. The emphasis is more on the extent of within- and between-job flexibility—whether it is large or small and for which type of workers it is larger or smaller—rather than on the question of whether hours flexibility is complete or not. Specifically, we analyze transitions from positive hours of work to positive hours of work made 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 labor supply curve and combine these with information on stated preferences and job mobility to assess whether and how women adjust their labor supply in response to changes in the incentives to work a given number of hours.

Many of the tax and benefit reforms in the United Kingdom, Canada, and the United States have been directed at increasing the labor market attachment of the lower-skilled workers, in particular, those facing high fixed costs of work such as child care (Blundell 2002). A significant part of the rise in employment among single mothers in the United States 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 partici-

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

pation of single mothers in the United Kingdom 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 U.K. reforms has been the changing incentive structure toward 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 mechanism for these adjustments in labor supply has not been studied. Are adjustments to hours made by moving jobs, or do workers adjust their hours of work over time with the same employer? This mechanism of adjustment is the focus of this article.

For such an analysis panel data are essential, as it is necessary to know the 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 was costless, we would still expect workers to be located on their labor supply curve, that is, 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 because of 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 estimating models of labor supply. Various strands of research have suggested models of hours choice in which hours are fixed within jobs. One strand, which dates back to Barzel (1973) and Rosen (1976), grounds its analysis in models in which jobs are packages of fixed hours-wage combinations (Ham 1982; Moffitt 1984; Lundberg 1985; Altonji and Paxson 1988, 1992; Biddle and Zarkin 1989; Kahn and Lang 1991; Dickens and Lundberg

[2] Blundell and Hoynes (2004) and Brewer et al. (2006) 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.

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 examining if and how employed single mothers vary their hours in response to exogenous changes in the incentives to work a given level of hours. 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 hours flexibility. We also look at how changes in hourly wages both within and between jobs relate to the introduction of the reforms. Although this analysis can be biased by the usual endogeneity problems, it is likely to give us a more exhaustive picture of the British labor market and an indication of the possible presence of imperfections or technological rigidities.

Besides providing us with relatively "clean" experiments to test hours constraints, these three tax/benefit reforms (especially the WFTC program) have also been widely analyzed in previous studies (Bingley and Walker 1997; Blundell et al. 2000; Gregg and Harkness 2003; Blundell and Hoynes 2004; Brewer et al. 2006; Francesconi and van der Klaauw 2007). These studies have come up with broadly consensual evaluations of the reforms' effects on a number of outcomes, including employment and wages. None of these studies, however, focuses on changes in worked hours. Stewart and Swaffield (2004) examine the working hours of low-wage employees in the United Kingdom but analyze the impact of the introduction of the National 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 they do not show how single women with and without children have been differentially affected. In addition, neither these studies nor the earlier research on wage-hours packages analyzes job-changing behavior as a mechanism to adjust hours of work or address the broader issue of labor 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 adjustment

[4] There has been relatively little analysis of hours constraints in Britain. Two studies that have investigated the extent of constraints on desired hours are Stewart and Swaffield (1997) and Bryan (2007). Using data from the BHPS, they both find that a substantial proportion of male workers (Stewart and 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 labor supply adjustment.

primarily occurred through job changes rather than labor supply adjustments within a job. There is a good deal of heterogeneity in the effects of the WFTC reform, with evidence of even less adjustment within jobs emerging among single mothers whose youngest child was aged 0–4 and who worked in larger firms, service industries, and the public sector. The presence of some hours inflexibility within jobs is confirmed when we look at hours changes by stated labor supply preferences. Women who stated that they were unconstrained in their job showed the largest upward adjustments after the WFTC reform if they changed jobs. Similarly, and again in line with their stated preferences, 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 jobs. Finally, we find relatively little effect on wages. However, there is some weak evidence that certain groups of women (especially single mothers who lived in London and the South East) operated under monopsonistic conditions, whereby changing jobs led to significantly lower wages after the introduction of WFTC.

Our research is likely to be relevant for many aspects of labor market policy, especially for the design of tax credit and benefit policies that 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 demonstration project for the Employment Retention and Advancement Scheme in the United Kingdom). From the result that hours are not very flexible within jobs, we can infer that changes to the tax/benefit incentives to work a given number of minimum hours are likely to influence rates of job-to-job transitions for the affected groups of workers.

Section II 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

Programs to support low-income working families with children (hereafter called "in-work benefits," even though the more recent programs are officially designated tax credits) have a long history in the United Kingdom. A peculiar feature of the United Kingdom'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 been available only to families with children who usually work some minimum number of hours a week.[5]

Two in-work benefits were in operation during our sample period: Family Credit, which existed from April 1988 until September 1999, and the Working Families' Tax Credit, 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' labor supply is ambiguous: those working more than 16 hours had an incentive to cut hours to (no less than) 16, whereas those previously working fewer than 16 hours had an incentive to increase their labor supply to (at least) the new cutoff. In 1995, there was another reform to FC, 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 labor supply decisions of lone parents in obvious ways: there was an increased incentive for those working fewer 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 fewer than 30.

The 1999 WFTC reform has a more complicated impact on labor 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' labor 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

[5] Hours rules are an important feature of the United Kingdom'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 labor supply effects, was the introduction of WFTC. We do not want to claim, however, that there has been a stable postreform period since October 1999. On this and other related and concurrent policy initiatives, see the discussion in the next subsection.

week) and a substitution effect toward 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; Brewer et al. 2006).[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 program in 1999) means that we can divide our sample into three periods: (*a*) autumn 1991 to March 1995, with the postreform 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 postreform period (labeled FC+) being defined over the years 1995–97; and (*c*) October 1999 to the end of the sample, with the postreform period (labeled WFTC) being between 1999 and 2002.[8] In our empirical analysis we take advantage of each of these separate reforms: not only did they have the potential to affect single mothers' hours of work, but they also could have done so in opposite directions. 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 and Identification Issues

To assess whether female labor 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' behavior through the differential tax and benefit treatment that they receive as compared to a control group, which is given by single women

[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 child care costs or high house rents (e.g., London and the South East of England). The availability of a more generous child care tax credit component under WFTC might reduce this problem (Francesconi and van der Klaauw 2007), 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 Sec. IV.B we will present and discuss estimation results obtained after stratifying the sample by child's age, housing tenure, and region of residence.

[8] In Sec. IV.A we shall return to the definition of the postreform 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).

without children.[9] 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. Therefore, identification relies on the assumption that variation in labor supply preferences of single parents be independent of the reforms conditional on the observed covariates and time effects.[10]

At the time of the introduction of the 1999 reform, however, there were other shocks that might have influenced single mothers' and childless women's labor supply differently. Three policy changes in particular could have interacted with the WFTC effects. First, there was an increase in basic child benefits under Income Support (the main welfare benefit, similar to Aid to Families with Dependent Children or Temporary Aid to Needy Families in the United States) between 1998 and 1999. In terms of labor 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. Second, the National Minimum Wage (NMW) was introduced in April 1999 (Dickens and Manning 2004; Stewart 2004). The NMW might have affected both the extensive margin of labor supply (inducing inactive women to get a job) and the intensive margin (increasing the incentives for working women to work more hours). But such incentives presumably had the same impact on single mothers' behavior that they did on single childless women's. The NMW-related shock, therefore, is not likely to have changed the employment outcomes of the treatment group differently than those of the control group.

Third, between July 1997 and October 1998, the British government launched a series of New Deal programs intended to help different groups of low-income people move from welfare into work using a combination

[9] The choice of single women without children as the control group in our analysis is somewhat arbitrary. Albeit not eligible to receive FC or WFTC because they do not have children, these women are different from single mothers along a number of observable characteristics (see Sec. III). Most of the existing studies on the effect of in-work benefits on lone mothers use the same control group that is used here, whether they look at the United Kingdom (Blundell et al. 2000; Gregg and Harkness 2003; Blundell and Hoynes 2004; Francesconi and van der Klaauw 2007) or the U.S. experience (Eissa and Liebman 1996; Meyer and Rosenbaum 2001). Blundell and MaCurdy (1999) lay out the identification conditions for such an analysis. Their credibility in the context of the analysis of tax reform is further discussed in Heckman's (1996) comment on Eissa (1996). In Sec. IV, however, we perform some sensitivity analysis in which the control group is restricted to single childless women with lower educational attainment.

[10] In general, conditioning can be accomplished nonparametrically by combining matching and difference in differences. We find that this makes very little difference to our estimates, which condition linearly on covariates.

of intensive job search assistance and small basic skills courses (Blundell et al. 2002; Van Reenen 2004). One of such initiatives, the New Deal for Lone Parents (NDLP), was aimed at all lone parents in receipt of Income Support with children under 16 and whose youngest child was over 5 years and 3 months (from April 2000 this lower age cutoff was dropped to 3).[11] Under NDLP, lone parents were assigned to a personal advisor, whom they were supposed to meet once every 2 weeks to receive advice on job vacancies, in-work benefits, child care arrangements, training, and job search techniques. One interesting aspect of NDLP, which was shared with some (but not all) other New Deal programs, was that involvement in the scheme and searching for work were entirely voluntary, and benefit entitlements did not depend on whether people decided to enter the scheme or not.[12] Single women without children were not involved in a similar initiative, unless they too were longer-term unemployed and had low income. Therefore, NDLP and any of the other New Deal schemes were likely to affect women—whether single mothers or not—only to the extent that they were unemployed. But since unemployed and out-of-the-labor-force women are excluded from our analysis (in fact, women must be employed for at least two consecutive years to be included in our sample; see Sec. III), the influence of NDLP on hours changes is likely to be limited. In any case, as single mothers, on average, have less education and are more likely to be unemployed, we performed sensitivity checks that will be discussed in Section IV.A by replicating our analysis using a more restricted control group, consisting of single childless women with low educational attainment.

Turning to the model specifications, 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 one if woman i is a lone mother at time $t-1$, and zero otherwise; and let Q_{it} be equal to one if

[11] Eligibility for and provisions of the various New Deal schemes have slightly changed over time. In relation to NDLP, since 2002 lone parents are eligible for NDLP not only if they are in receipt of Income Support (as they were in previous years) but also if they receive other benefits (such as Housing Benefit and Council Tax Benefit) and, importantly, WFTC (as well as maternity allowance and statutory maternity pay). Also eligible are lone parents working under 16 hours per week (and thus ineligible for WFTC) who are not claiming any benefits except child benefits. All these changes, however, were implemented outside our sample period.

[12] Compulsory Work Focused Interviews (CWFI) for lone parents claiming Income Support were introduced in April 2001. Under CWFI, people of working age seeking to claim Income Support are obliged to participate in a work-focused interview with an advisor at the start of their claim as a condition of receiving the benefit. Kirby and Riley (2004) find little evidence that CWFI increased labor market participation among inactive benefit-claiming lone mothers.

woman i changes a job between years $t-1$ and t, and zero otherwise. The four specifications are as follows:

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

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

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

and

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

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 independently and identically distributed 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, housing tenure, union coverage, and industry; the number and changes in the number of children by age group; and changes in health status and local unemployment rate.[13] The treatment effects for movers are captured by β_{FC} and β_{WFTC}, whereas b_{FC} and

[13] The levels of time-varying variables are all measured at $t-1$.

b_{WFTC} capture, respectively, the FC and WFTC treatment effects for workers who did not change jobs (stayers).[14]

The key differences across equations (1)–(4) involve the specification of time trends. In equation (1), time trends are not modeled, 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}), and 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 behavior after each reform (through α_{21} and α_{22}). If $\hat{b}_j = \hat{\beta}_j$ (with j = FC, WFTC), we cannot statistically reject the hypothesis of within-job flexibility in hours choice; whereas 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 refer to changes,[15] 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 12 waves of the BHPS collected over the period 1991–2002. Since autumn 1991, the BHPS has annually interviewed a representative random stratified sample of the population of Great Britain with about 5,500 households covering more than 10,000 individuals. All adults and children in the first wave are designated as original sample members. Ongoing representativeness of the nonimmigrant 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 houses or if their households split up), and there are interviews, at approximately 1-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

[14] It is worth noting at this point that were firms to adjust their overall shift lengths in response to these changes in desired labor supply, there could be important spillover effects on other workers in our control group. We have not been able to locate any evidence either way on changes in shift length (or, more generally, in labor demand) at the time of these reforms in Britain, but it would clearly be useful to document evidence on this.

[15] Women with 0 hours at the time of any of the 12 interviews are excluded from the analysis. For further discussion on this point, see Sec. III.

therefore remains broadly representative of the population of Britain as it changes over time.[16]

Our estimation sample includes employed unmarried, noncohabiting females (separated, divorced, widowed, or never married) who are at least 16 years old and were born after 1941 (thus aged at most 60 in 2002). Because equations (1)–(4) refer to changes in hours worked, we measure hours changes conditional on being in work in period $t - 1$ and remaining in work in period t. We exclude any female who was in school full-time or self-employed or out of the labor force in a given year.[17] 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% of the women are observed in the same marital state for all 12 years of the panel, about 60% 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 lone mothers and 7,774 childless women). Of the 1,280 single women in the 1999 wave of interviews, 25 lone mothers and 32 childless women (about 4.5% of the sample in that year) were interviewed just before the day on which the 1999 reform was implemented (October 5). To limit problems of interpretation, they were dropped from the estimating sample. Their inclusion, however, does not alter any of our main results.

Table 1 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.[18] Lone mothers are younger (30 vs. 38 years), less educated (56%

[16] Of the individuals interviewed in 1991, 88% were reinterviewed in wave 2 (1992). The wave-on-wave response rates from the third wave onward have been consistently above 95% (i.e., 95% of the previous wave respondents get interviewed). Detailed information on the BHPS is presented in Lynn et al. (2006) and 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.

[17] Excluding women who were long-term ill or were registered disabled, but satisfied the other sample selection restrictions, did not change any of the results presented below. In the present analysis, such women are included in the estimating sample.

[18] Restricting our analysis to women who work for at least two consecutive periods leads to a sample of women who are more educated and less poor than those who are observed out of the labor market more frequently. But crucially

Table 1
Summary Statistics

Variable	Unpartnered Women without Children		Lone Mothers	
Total weekly hours of work	34.74	(13.25)	25.61	(14.20)
Change in worked weekly hours*	.39	(11.40)	2.25	(12.31)
Absolute change in worked weekly hours*	6.00	(9.62)	6.91	(10.28)
Hourly pay	7.06	(6.01)	5.85	(5.19)
Monthly labor income conditional on working positive hours (in 2002 £'s)	1,110	(911)	694	(629)
Age (years)	38.1	(15.0)	30.00	(11.36)
Nonwhite	.043		.090	
Registered disabled	.049		.023	
Number of children by age group:†				
0–4			.231	(.510)
5–10			.588	(.755)
11–18			.798	(.771)
House owner	.578		.541	
In social housing	.229		.377	
A-level or higher educational qualification	.520		.438	
No qualification	.152		.144	
University degree or more	.144		.060	
Total work experience (years)	14.33	(11.47)	8.67	(7.88)
Employed in a firm with fewer than 50 workers	.660		.746	
Employed in service industries‡	.838		.820	
Employed in the public sector	.247		.171	
Union covered	.514		.530	
Changed job during previous year	.167		.179	
Local unemployment rate§	.065	(.032)	.063	(.031)
Number of person-wave observations	7,774		4,585	
Number of women	1,162		1,122	

NOTE.—The figures are means (standard deviations in parentheses) computed over all person-wave observations for which two consecutive years of data are available.
* The change is measured over two consecutive years.
† 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 (.448), 1.318 (.582), and 1.321 (.548), respectively.
‡ Service industries refer to banking, finance and insurance, distribution, hotels and catering, transport and communication, and other services (which include education and sanitary services).
§ Computed over 306 travel to work areas.

have qualifications short of A level vs. 48% among childless women, and only 6% of lone mothers have a university degree vs. 14.4%),[19] more likely to be nonwhite (9% vs. 4.3%) and in social housing (38% vs. 23%), less likely to be employed in the public sector (17% vs. 25%), and have fewer years of work experience. The two groups of women are instead

the differences between treatment and control groups in this larger sample (containing also women who are out of the labor market) are very similar to those found in the more restricted sample used in the article.

[19] For non-British readers, "A (advanced) level" corresponds to education beyond high school but short of a university degree.

relatively similar in terms of job-changing behavior, with 17% of childless women and 18% of single mothers moving across employers in any two given years.[20] Systematic differences emerge again in the case of labor 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 2.25 hours per week vs. less than 25 minutes).[21]

To gain a greater insight into how the reforms might have changed the distribution of hours worked among all single women, figure 1 plots histograms of total weekly hours of work for all women in the sample by survey year (with the vertical line in each panel indicating the eligibility hours cutoff).[22] Women can be found working any number of hours from 1 to 60 per week in any given year. In most years, we observe a great deal of variability, with bunching at about 20, 30, and 35–40 hours and, depending on the year, at some hours between 40 and 50 (see also Blundell, Duncan, and Meghir 1998; Blundell and Hoynes 2004). A striking feature is that, in every year up to 1996 (perhaps with the exclusion of 1991), there was only a small fraction of workers below the eligibility cutoff (accounting for about 15%–20% of the single women in the sample), whereas in the two years prior to the 1999 reform, there was clear evidence of bunching just below the 16-hour cutoff. From 1999 onward, the fraction of female workers with total weekly hours between 16 and 20 was around 12%, almost twice as large as the fraction of workers in the same hours range between 1992 and 1998. These features of the data provide some quantitative indication of the hours effect associated with WFTC. From 1998 onward, also discernible is a greater concentration of workers

[20] Our measure of job change does not include internal promotions or job changes within the same firm or establishment, but includes all moves from one firm to another (through either 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) produce results similar to those reported in this article. See also Sec. IV.C.

[21] 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 table 1, 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 Sec. IV.C).

[22] For the sake of visual clarity, fig. 1 does not show the observations with more than 60 weekly hours. These, however, represent less than 1% of the subsamples in each survey period and are included in the regression analysis reported below.

at 30 hours, but we cannot detect any substantial change around the 30-hour cutoff immediately after the introduction of FC+.

We now turn to mean hours changes. Figure 2 plots the time trends for the year-on-year average changes in total hours worked over the sample period (with the dotted lines around the averages displaying the corresponding one-standard-deviation bands). Figure 2*a* shows the trends for all working women distinguishing lone mothers (straight line) from single childless women (dashed line), and figures 2*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 (fig. 2*a*). The mean hours changes for lone mothers instead are greater, and their time variability is higher too. The largest hours change is observed after the introduction of WFTC between 1998 and 1999, when lone mothers reported, on average, an increase of about 4.5 hours of work per week.[23] 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' labor supply by about 2.5 hours, but again this increase was not followed by further increases 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 its introduction in 1995, but it is followed by a steady increase even before the peak between 1998 and 1999.

Figure 2*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 figure 2*a*, 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 (fig. 2*c*).

[23] This contrasts with the estimates reported in Stewart and Swaffield (2004), which provide evidence of a female labor supply reduction of 1–2 hours per week as a result of the introduction of the NMW in April 1999. Their results are not robust across data sets and specifications and are obtained from data that stop in September 2000 at the latest (i.e., just before the second post-WFTC year in our sample). In addition, as pointed out in the introduction, the Stewart-Swaffield 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.

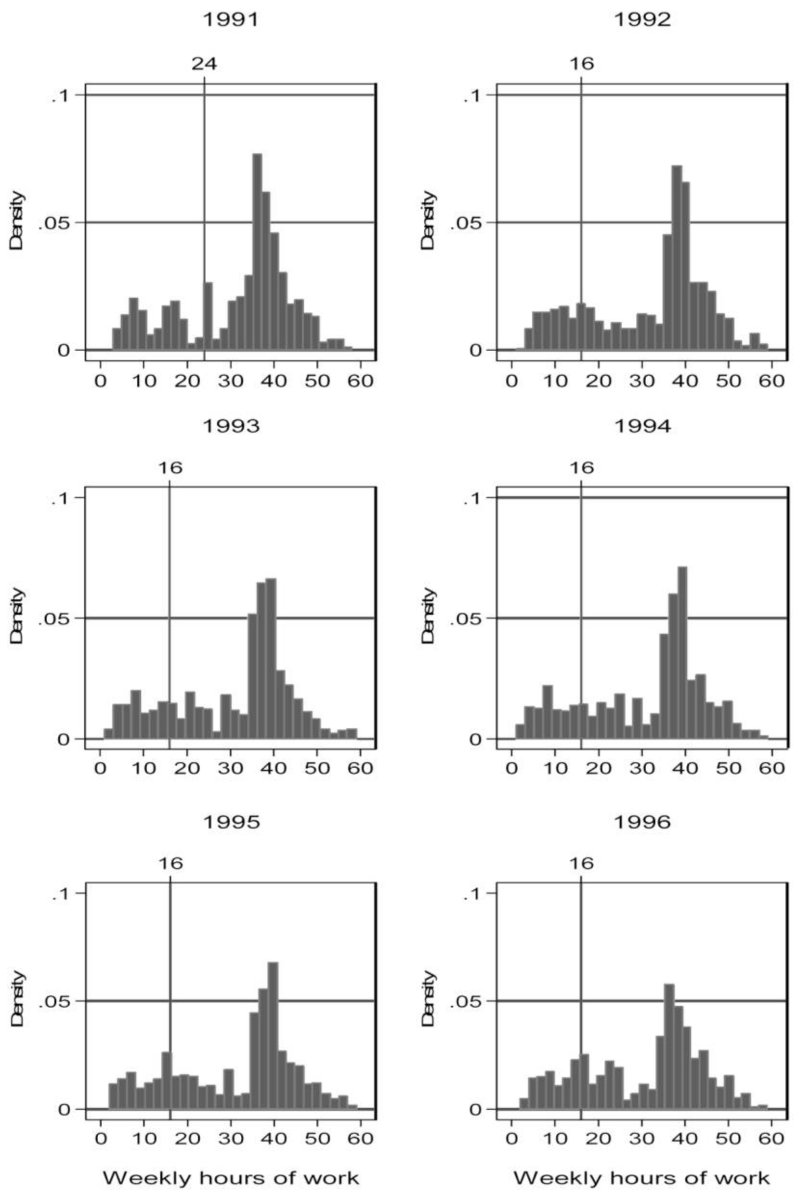

FIG. 1.—Female weekly hours of work by survey year. The vertical line indicates hours eligibility cutoff.

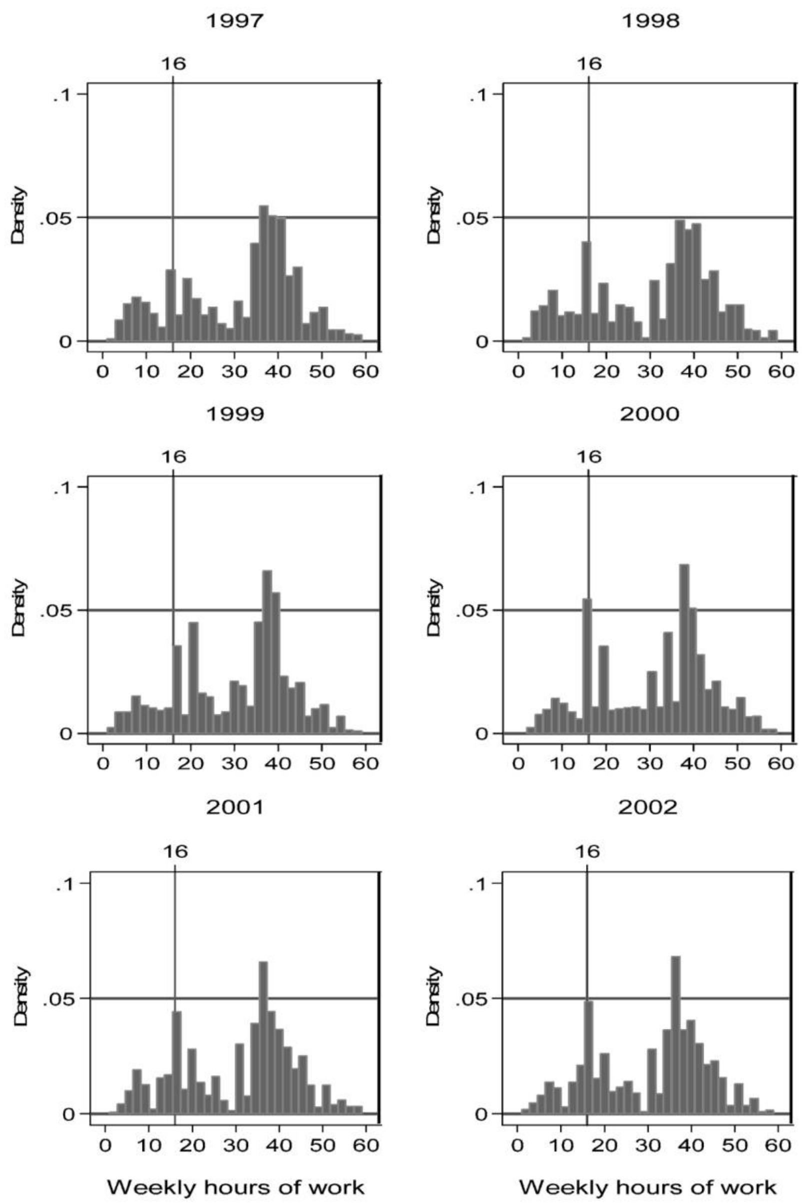

Fig. 1 (*Continued*)

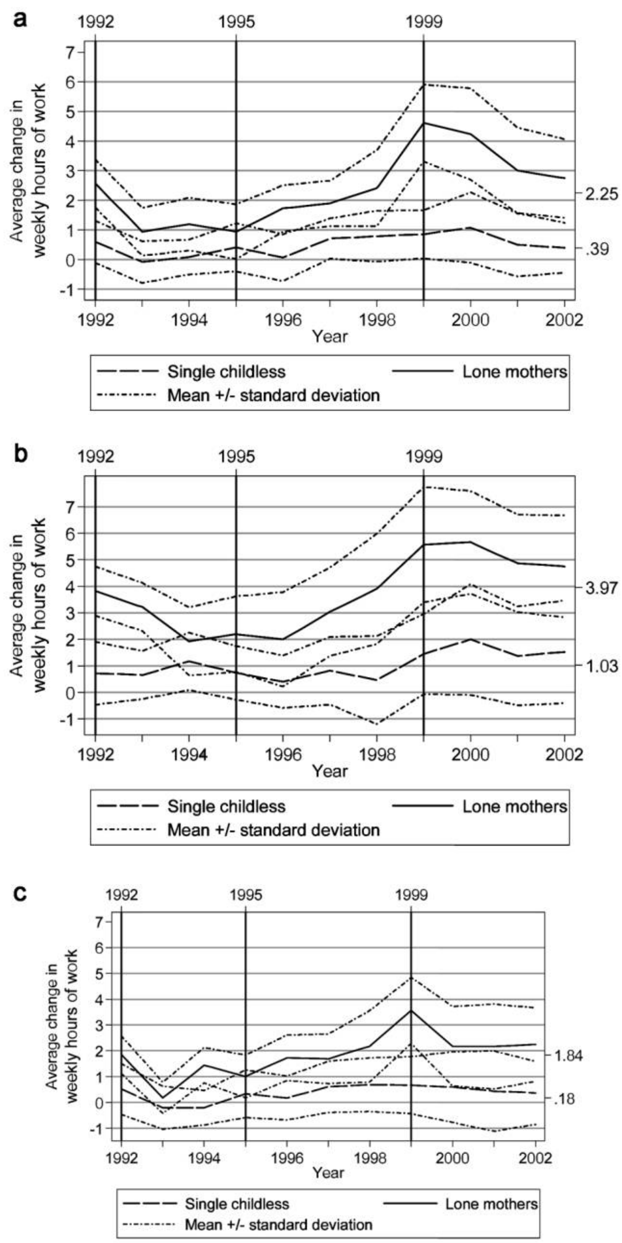

FIG. 2.—Average changes in total weekly hours of work: single childless women and lone mothers by job-changing status. *a*, All workers. *b*, Job movers. *c*, Stayers.

Table 2
The Impact of the In-Work Benefit Reforms and Job Changes on Hours Changes

	Without Controls (N = 12,359)				With Controls (N = 12,359)			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
α_1	1.58	1.27	1.54	1.36	.34	.44	.24	.19
	(7.94)	(2.68)	(3.51)	(2.98)	(1.08)	(.85)	(.59)	(.32)
α_2	.77	.74	.73	−.19	−.30	−.32	−.31	−.45
	(1.86)	(1.78)	(1.76)	(.28)	(.64)	(.70)	(.71)	(.82)
b_{FC}			−.25	.03			−.03	−.21
			(.40)	(.07)			(.01)	(.44)
b_{WFTC}			.16	.45			.20	.56
			(.33)	(.89)			(.42)	(.94)
β_{FC}	.11	.72	.95	.48	.21	.83	.89	.44
	(.08)	(.51)	(.67)	(.29)	(.15)	(.59)	(.62)	(.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)

Source.—British Household Panel Survey, 1991–2002.
Note.—Absolute values of t-statistics (obtained from standard errors that are adjusted to reflect multiple observations per person) are in parentheses. The labeling of cols. 1–4 corresponds to eqq. (1)–(4) described in the text. Controls include a cubic polynomial in total work experience; dummy variables for race, educational attainment, firm size, public sector, region of residence, housing tenure, union coverage, and industry; the number and changes in the number of children by age group; and changes in health status and local unemployment rate.

IV. Results

A. Benchmark Estimates

The estimates of the impact of job-changing behavior on hours changes are shown in table 2. These are presented for the four specifications described in Section II.B and separately for the cases in which the variables in **X** are excluded or included.[24]

The regressions without controls indicate that changing jobs is associated with increases in women's labor supply by less than 1 hour per week (α_2), although this effect is significant only at the 10% level in the first three specifications, whereas single mothers experience significantly larger 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 in their labor 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% level (the p-value of the t-test of equality is .024), which provides evidence

[24] In parentheses, this and the subsequent tables report the absolute value of t-ratios obtained from standard errors that are adjusted to reflect multiple observations per person (and are robust to arbitrary forms of serial correlation and misspecification). For the sake of brevity, the estimates on the control variables are not reported but are available from the authors.

of hours inflexibility within jobs. Most of these results are robust to the inclusion of the control variables **X**, with the only exception of α_1, which now becomes statistically insignificant.[25] With 16% and 20% of women changing jobs after the 1992 and 1999 reforms, respectively, we can derive their overall effects on hours changes conditional on working: FC had virtually no overall impact, whereas WFTC increased single mothers' weekly hours of work by about 1.1 hours (specification [4]).

Because figure 2 reveals that stayers also increased their worked hours immediately after the 1999 reform, the previous analysis was repeated after excluding the last two years of the sample. Indeed, the WFTC effect for stayers is now larger and close to 1 extra hour per week, but its *p*-value is never below .11. In any case, even after this selection, all other results are confirmed, including the rejection of the hypothesis of flexibility in hours within jobs.[26] Thus, in response to the exogenous change in work incentives given by the WFTC program, changing jobs seemed to have been the strongest mechanism of labor supply adjustment among single mothers after 1999.

We repeated the previous analysis with a different subsample, in which the control group of single childless women is limited to those with educational qualifications below A level. In the spirit of the discussion in Section II.B, this allows us to see whether our results are concentrated in specific subgroups of the population that might have been affected by other policy initiatives (such as the NDLP) that were introduced at approximately the same time as the WFTC reform. It also provides us with an important sensitivity check. Restricting our analysis to this different control group reduces only slightly the treatment effect for movers under WFTC (β_{WFTC}) to 3.28 (t-ratio $= 2.86$), changing neither the estimated effect for movers under FC nor the treatment effects for stayers in any significant way. The results illustrated so far, therefore, are robust to this change in the definition of the control group.

B. Heterogeneous Responses

It is possible that the labor 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 esti-

[25] To understand this lack of effect, we estimated variants of eqq. (1)–(4) with Q interacted with marital status (not shown). Regardless of whether we control for group-specific time trends, changing jobs is associated with increases of about 1.2 hours per week for single mothers and with reductions of 0.8 hour per week for single childless women.

[26] We reach the same conclusion if we keep the entire sample as in table 2 but redefine the postreform period as either 1999–2000 or 1999–2001. Similarly, redefining the FC period over 1992–93 (rather than 1992–94) does not alter our baseline results.

mating models that distinguish women separately by individual attributes (such as education and number and age of children), work-related attributes (such as firm size and industry), and stated labor supply preferences. The results from these regressions (based on specification [4] only) are reported in table 3.

The estimates in panel A of the table reveal that the increase in hours worked after the 1999 reform was predominantly observed among single mothers who changed jobs and used to work fewer than 16 hours per week. Women in this group experienced a labor supply increase in excess of 6 hours per week, whereas women who already worked 16 or more hours experienced a more modest growth of about 2 additional hours.[27] Both effects are significantly different from zero, and they are statistically different from each other at conventional levels (p-value = .006). These results suggest that an important part of the overall treatment effect of the 1999 reform was driven by greater entry into WFTC-eligible employment of already working single mothers. As the differences between $\hat{b}$ and $\hat{\beta}$ document, there is evidence of a greater degree of hours inflexibility within jobs after the WFTC reform for both groups of women and, for women who worked fewer than 16 hours per week, also after the 1992 FC reform (but this effect is significant only at the 6% level).

The treatment effects for stayers do not differ between more educated women and less educated women (panel B). There are, however, asymmetric responses among movers. Less educated single mothers increased their labor supply by 3–4 hours per week after the 1992 reform, whereas more educated single mothers' supply increased by 4–5 hours after the 1999 reform.[28] The null hypothesis that the estimated b and β coefficients are equal can be rejected at the 5% level during the WFTC regime among the more educated and at the 10% level during the FC regime among the less educated.

The next two panels demonstrate that the post-WFTC upward ad-

[27] These estimates are also accompanied by a significantly different impact of changing jobs on hours changes for the two groups of women (α_2). When moving from one job to another, women in the bottom part of the hours distribution faced an average increase of nearly 2 hours per week, whereas women in the top part of the distribution reduced their labor supply by about 1 hour per week.

[28] If a large proportion of better-educated single mothers had not been eligible for 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% of better-educated lone mothers who work 16 or more hours per week were eligible for FC. Between 2000 and 2002, 49% were eligible for WFTC (an increase in eligibility rate by 88%). For the less educated, the increase in eligibility rate was only 28% (from 65% to 83%).

Table 3
Heterogeneous Responses by Selected Observed Groups—Specification (4)

	α_1	α_2	b_{FC}	b_{WFTC}	β_{FC}	β_{WFTC}	Observations
A. Previous hours worked:							
Fewer than 16 per week	−.17	1.88	−.19	.26	2.84	6.39	3,117
	(.14)	(3.54)	(.31)	(.18)	(1.92)	(2.52)	
16 per week or more	.32	−1.24	−.23	.68	−.57	2.27	9,242
	(1.25)	(2.89)	(.34)	(1.47)	(1.55)	(2.51)	
B. Education:*							
Less educated	−.31	−.64	.51	.53	3.61	1.92	6,297
	(.61)	(.82)	(.85)	(.88)	(2.15)	(1.26)	
More educated	.56	−.25	−.45	.58	−2.65	4.89	6,062
	(.83)	(.39)	(.23)	(.77)	(1.52)	(2.88)	
C. Number of children:†							
One child	.60	−.97	.47	.39	−.47	4.15	6,427
	(1.07)	(1.49)	(.78)	(.64)	(.44)	(2.95)	
Two or more children	−.29	−.16	−.57	.53	1.98	1.37	5,932
	(.66)	(.27)	(.88)	(.84)	(.92)	(1.31)	
D. Age of youngest child:‡							
0–4 years	−.33	−.27	−.59	.30	−.51	3.82	5,438
	(.37)	(.36)	(.78)	(.19)	(.45)	(2.84)	
5 years or more	.62	−.76	.03	.64	.97	2.80	6,921
	(1.09)	(1.45)	(.82)	(1.05)	(.56)	(2.63)	
E. Firm size:							
Fewer than 50 employees	−.01	−.38	.47	.61	.37	2.18	8,553
	(.02)	(.61)	(.84)	(1.08)	(.25)	(1.84)	
50 employees or more	.74	−.93	−.38	−.16	.72	4.20	3,806
	(1.03)	(1.52)	(.38)	(.13)	(.41)	(2.64)	

F. Industry:§							
Services	.72	−.84	−.48	.33	−.82	3.74	9,262
	(1.12)	(.97)	(.28)	(.51)	(.88)	(3.15)	
Manufacturing	−.40	.23	.30	.86	1.57	2.95	3,097
	(.74)	(.15)	(.48)	(1.27)	(1.91)	(2.77)	
G. Sector:‖							
Private	−.08	−.26	.09	.55	1.08	3.09	9,659
	(.19)	(.27)	(.16)	(.91)	(.54)	(2.43)	
Public	.94	−.88	.43	−.18	−.52	4.10	2,700
	(1.14)	(1.52)	(.68)	(.15)	(.76)	(2.71)	
H. Labor supply preferences:#							
SAME = 1	.07	−1.21	.61	.84	.43	4.20	7,539
	(.17)	(1.58)	(.90)	(1.52)	(.74)	(2.97)	
OVER = 1	1.65	−3.92	−.53	−.35	−6.74	1.09	3,090
	(1.19)	(2.86)	(.86)	(.66)	(1.84)	(.31)	
UNDER = 1	.31	4.37	−.24	−.50	2.97	2.87	1,730
	(.42)	(2.73)	(.15)	(.27)	(1.44)	(1.64)	

Note.—Absolute values of t-statistics (obtained from standard errors that are adjusted to reflect multiple observations per person) are in parentheses. All regressions include the control variables used in table 2, except for education (panel B); number and change in the number of children by age group (panels C and D); firm size (panel E); industry (panel F); and sector (panel G). The variables defining each of the observed groups are measured at time $t-1$. For other definitions see the note to table 2.

* Less educated is defined as having less than A-level qualifications; more educated is defined as having A-level or higher qualifications.

† One child and two or more children pertain to lone mothers.

‡ Youngest child aged 0–4 and youngest child aged 5 or more refer to lone mothers.

§ Services include 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.

‖ Public sector includes civil service, central and local government, National Health Service, education, and nonprofit organizations.

\# OVER equals one 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 zero otherwise; UNDER equals one 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 zero otherwise; SAME equals one 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 zero otherwise.

justment in single mothers' labor 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 sizable and significant (panel D). If we pool all women as we did in table 2 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 ($b_{\mathrm{FC}} = -1.35$ and standard error $= 0.48$), whereas 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 workers (in this case, single mothers of young children who did not change jobs) to reduce their hours worked over the 1992–94 period.

The U.K. 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 programs. Rents in some parts of the country (in particular, London and the South East) are high and have rapidly increased over the 1990s, whereas owner-occupiers are not eligible for Housing Benefit. To capture part of the relationship between Housing 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 labor supply adjustments observed after the 1999 reform were greater for single mothers who lived outside the London/South East region (where house rents are lower and the interaction with Housing Benefit is likely to be more modest) and who were not owner-occupiers.[29]

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 (on the order of 4 additional hours per week, panel E), in service industries (about 3 extra hours, panel F),[30] and equally for those employed in the private sector or the public

[29] 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 jobs after the introduction of WFTC for single mothers who lived in rented accommodations outside the London/South East region.

[30] Single mothers who were employed in manufacturing industries also showed a significant increase of 3 hours of work per week if they changed jobs after the introduction of WFTC (panel F). For the same group of women there is also evidence (significant only at the 10% level) of positive labor supply adjustments

sector (between 3 and 4 additional hours per week, panel G). Strong evidence of hours inflexibility emerges among lone mothers who work in larger firms, service industries, and the public sector.[31]

Another important dimension along which we expect to see heterogeneous responses is given by stated labor 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 labor supply preference variables for any given year of the sample period, labeled OVER (one if a worker would like to work fewer hours, and zero otherwise), UNDER (one if a worker would like to work more hours, and zero otherwise), and SAME (one if a worker would like to continue to work the same number of hours, and zero otherwise).[32] 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 labor supply. The estimates on α_2 reported in panel H of table 3 confirm such expectations, with overemployed workers reducing their labor 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

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of about 1.6 hours per week if they changed jobs between 1992 and 1994 (i.e., during the FC regime). This effect involves only 25% of the whole sample, and this may be why it does not show up in the baseline estimates of table 2 for 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.

[31] To capture occupation-specific human capital, we examined some further interaction terms constructed using the Standard Occupation Classification (SOC) measure. In particular, we partitioned the sample into two occupation groups, i.e., "white-collar" occupations (i.e., managerial, professional, and technical occupations, corresponding to major groups 1–3 in the one-digit SOC system) and "blue-collar" occupations (i.e., clerical/secretarial occupations, crafts, personal service, sales, semiskilled, and unskilled, SOC groups 4–9). The results were similar to those for education and industry, with women in white-collar occupations showing a greater positive response after the 1999 reform (significant) and a negative response after the 1992 reform (not significant). All other estimates remain broadly similar to those already reported. These additional results are available from the authors on request.

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

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 jobs. But single mothers who wanted to continue working the same number of hours showed large upward labor supply adjustments of about 4 hours per week if they changed jobs after the WFTC reform.[33] Thus, initially "unconstrained" (i.e., neither over- nor underemployed) lone mothers did respond to the greater work incentives of the WFTC program, but only through a change of job.[34] 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 jobs, although none of such increases is statistically significant at conventional levels. After the 1992 reform, instead, we observe large (and significant at the 10% level) reductions of about 7 hours per week among overemployed single mothers who changed jobs. 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 labor 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% level, irrespective of the specification. Underemployed workers seem to be unable to adjust their labor supply upward if they did not change jobs. But those who moved did manage to increase their worked hours even after the 1992 reform by about 3 hours per week (although this increase is not statistically significant).[35]

We reestimated variants of equations (1)–(4) over the whole sample of women that included interaction terms between the variables on b_j and β_j (j = FC, WFTC) and stated labor supply preferences. The results from this analysis (not shown) confirm those previously discussed. In particular (from specification [4]), unconstrained single mothers who changed jobs after the 1999 reform increased labor supply by about 4 hours (t-value =

[33] The hypothesis that the estimated b and β coefficients are equal can be rejected at the 5% level (p-value = .027).

[34] Notice that "unconstrained" workers are defined to be those who would like to continue to work the same number of hours. This definition may not precisely reflect their entire preference ordering, since they may be constrained in other dimensions (e.g., job location and family responsibilities).

[35] Following Altonji and Paxson (1992) and Euwals (2001), we also checked whether the hours adjustments estimated in conjunction with the WFTC reform are in line with women's stated preferences. The results (which are not reported for convenience) show that this is the case, especially for underemployed lone mothers. Almost 80% of single mothers who wanted to work more did adjust their hours upward by changing jobs after the 1999 reform as opposed to only 30% among those who did not change jobs. The corresponding downward adjustments for women who wanted to work fewer hours were instead 55% and 18% for movers and stayers, respectively.

Table 4
Robustness Checks—Specification (4)

		Length of Time in the Panel†		Propensity Score Matching Models‡	
	Accounting for FC+* (1)	6 Years or More (2)	9 Years or More (3)	Biweight Kernel Matching (4)	Local Linear Regression Matching (5)
α_1	.30	.57	.66	.97	1.05
	(.69)	(1.20)	(1.13)	(.89)	(1.36)
α_2	−.56	−.81	−.60	−1.12	−.20
	(.88)	(1.56)	(.88)	(1.53)	(.51)
b_{FC}	.12	−.45	−.27	−.13	−.38
	(.11)	(.80)	(.40)	(.11)	(.75)
b_{FC+}[§]	−.22				
	(.44)				
b_{WFTC}	.24	.81	.75	−.49	.97
	(.48)	(1.54)	(1.02)	(.90)	(1.13)
β_{FC}	.81	.65	.17	.36	.3
	(.52)	(.50)	(.23)	(.48)	(.62)
β_{FC+}[§]	1.18				
	(1.32)				
β_{WFTC}	3.48	4.52	2.99	3.12	3.07
	(3.13)	(3.34)	(2.39)	(3.07)	(2.74)
Observations	12,359	8,314	5,153	12,359	12,359

NOTE.—Absolute values of *t*-statistics (obtained from standard errors that are adjusted to reflect multiple observations per person) are in parentheses. All regressions include the control variables used in table 2. For other definitions, see the note to table 2.

* These estimates are obtained from a regression accounting for the 1995 reform that provided extra credit for working 30 or more hours.

† 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.

‡ Absolute values of *t*-statistics (with standard errors obtained from 500 bootstrapped replications) are in parentheses. For the local linear regression matching regression, the estimates are obtained after imposing a tricube kernel.

§ Extra credit for full-time work.

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 hours conditional on working. In particular, after the introduction of WFTC, unconstrained mothers whose youngest child was aged 0–4 and who changed jobs 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 labor 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 table 4 presents the results only from three exercises using specification (4). The results obtained from the other specifications are qualitatively similar to those discussed here.[36]

First, we performed our analysis accounting for the 1995 FC reform that provided extra credit for full-time work. The estimates in column 1 confirm our previous findings and document that the 1995 reform was followed by no sizable change in worked hours irrespective of whether women changed employers or stayed in the same job.

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 results similar 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 (e.g., six or more waves). If attrition or changing sample composition is important, the results from such selected subsamples are expected to differ from those discussed earlier. Columns 2 and 3 of table 4 report the estimates found from two subsamples, one in which we include only women who have been observed for 6 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 nine consecutive times. In general, the estimates from both subsamples are relatively close to the corresponding figures reported in table 2. For example, β_{WFTC}, one of the key parameters in our study, is estimated to be 32% greater (col. 2) and 13% smaller (col. 3) than its counterpart in table 2. Despite such differences in magnitude, therefore, these estimates tend to 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 columns 4 and 5 of table 4 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 sub-

[36] We also reestimated the models eliminating laid-off workers from Q or dropping promoted workers from the group of stayers. Both these exercises produced results that were virtually identical to those shown in table 2 and are thus not reported.

stantially increasing their hours of paid work, whereas the two previous reforms to FC seemed to have induced only minor labor supply effects. The strong labor 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 analyze wage responses. Of course, in-work benefits reforms were directly designed to change the incentive to work specific hours leaving wages unaltered, whereas wage determination was affected more explicitly by the introduction of the NMW, and both hours and wages were (and still are) 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 labor market and some indication of the possible presence of labor market imperfections or rigidities in the matching technology. We therefore estimated equations (1)–(4) with log hourly wages (expressed in 2002 prices) as the 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 jobs after the introduction of WFTC implied not only a labor supply increase of almost 3 hours per week (t-ratio = 3.11) but also a wage reduction of 2.7% (t-value = 2.23). Among the latter, changing jobs after the 1999 reform led to 1.5% lower wages (t-value = 1.51) and modest positive hours changes (see table 3). Thus, despite the presence of hours inflexibility, the labor 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 United Kingdom, which either changed hours requirements to be eligible for the benefits or increased the attractiveness of working a given number of hours, we are able to assess the mechanism of labor supply adjustment among single women with children—the main target of these in-work benefit reforms. We find that the 1992 and 1995 FC reforms had modest impacts on single mothers' hours of work, 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 jobs, suggesting that the mechanism of labor supply adjustments is between rather than within jobs. This lines up well with the estimates we get when we look at hours changes by stated labor supply preferences: unconstrained women who changed jobs showed the largest hours increases after the 1999 reform, and overemployed women substantially reduced their hours worked 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 within jobs 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 labor supply model? We have shown that adjustments in hours of work are made primarily by movements between jobs, and there is little evidence of systematic labor supply–induced hours movements within jobs. Our analysis of stated preferences confirms this further, showing that responses are greater among those who say that they are unconstrained as well as among those who are constrained but state that they would like to move in the direction suggested by the incentives. Thus, a labor supply model emerges in which hours adjustments are largely made by moving between workplaces. This could be achieved within an "adapted" canonical model in which establishments are organized around hours requirements and individuals move jobs to achieve hours flexibility. Of course, it could also be supported by theories that emphasize the importance of labor market frictions and imperfections, such as job search, wage-job packages, and/or dynamic monopsony. However, if there were such "imperfections," we would expect them 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 labor supply model with hours flexibility across jobs cannot be rejected. Nonetheless, our results by region and firm size suggest that production technology or employer preferences not only may reduce labor supply flexibility within firms but also may place constraints on hours mobility across firms.

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