The Socioeconomic Consequences of 'In-Work' Benefit Reform for British Lone Mothers*

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Abstract

In October 1999, the British government enacted the Working Families' Tax Credit which aimed at encouraging work among low-income families with children. This paper uses panel data collected between 1991 and 2001 to evaluate the effect of this reform on single mothers. We find that the reform led to a substantial increase in their employment rate of about 5 percentage points, which was driven by both a higher rate at which lone mothers remained in the labor force and a higher rate at which they entered it. Women's responses were highly heterogeneous, with effects double this size for mothers with one pre-school aged child, and virtually no effect for mothers with multiple older children. The employment increase was accompanied by significant increases in paid childcare utilization and our analysis in fact suggests that the generous childcare credit component of the reform played a key role in explaining the estimated employment and childcare usage responses. We also find that the increase in labor market participation was accompanied by reductions in single mothers' subsequent fertility and in the rate at which they married, behavioral responses which in turn are likely to influence the reform's overall impact on child poverty and welfare.

JEL classification: C23, H31, I38, J12, J13, J22

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I. Introduction

In October 1999, the United Kingdom enacted the Working Families' Tax Credit (WFTC), a generous tax credit program designed to improve the work incentives for low-income families with children. In addition to considerably higher tax credits for working parents, a striking feature of the program — which distinguishes it from its predecessor as well as its counterpart in the United States, the Earned Income Tax Credit (EITC) — is its provision of a generous childcare tax credit. Although recent research has documented positive effects of this reform on employment and earnings among low-income families, and especially single mothers (Blundell and Hoynes, 2004; Brewer *et al*., 2005), little effort has been devoted to determining which particular aspects of the reform contributed to these outcomes. In addition, our knowledge of the impact of the WFTC on other important socioeconomic behavior, such as childcare usage, marriage and fertility, is scant or nonexistent.

Our paper offers two substantive contributions. First, we provide a more comprehensive study of the impact of the WFTC reform on employment outcomes of single mothers by (a) documenting the nature of (and substantial heterogeneity in) the employment increases across individuals, and (b) linking these to specific components of the policy reform. More specifically, using longitudinal data drawn from the British Household Panel Survey 1991-2001, we assess whether the overall employment increase was due to an increase in the rate at which lone mothers entered eligible employment, a drop in the rate at which they left it, or both. Our study provides new insight into the role of the childcare tax credit component of the WFTC reform by analyzing whether and how employment responses varied across female-headed households with the number and ages of children, as well as the extent to which they covaried with changes in paid childcare usage.

Second, we analyze whether the employment increase was accompanied by a change in the rate at which single women married, and whether the WFTC reform affected single mothers' subsequent fertility decisions — outcomes which have not previously been analyzed in the British context. Examination of such responses is important both because it gives us a more complete picture of the consequences of the 1999 in-work benefit reform and because it allows us to check for the occurrence of unintended effects that may be crucial for the longer-term success of the reform itself.

Theoretically the reform was predicted to increase the probability of moving into eligible employment (that is, working 16 hours per week or more) among lone parents, since it increased the financial payoffs to working any given hours level above 16 hours per week.

But because of the interaction of WFTC with other benefits, income gains were expected to be most pronounced for women working 25 hours per week or more. Recent studies of the WFTC's effect on employment among lone mothers confirm these expectations, although they are discordant on the size of the effect. Using estimates of a static behavioral model of household labor supply which includes controls for childcare costs, Blundell *et al*. (2000) provide an ex-ante simulation of the impact of the reform. They predict that the introduction of WFTC would only lead 2.2 percent of single mothers to move from no work to either part-time or full-time employment. Based on post-reform data, the estimates in Blundell and Hoynes (2004) suggest that the employment impact among lone mothers was indeed positive and modest. However, more recent work by Brewer *et al*. (2005) based on estimates of a static structural model of joint labor supply and program participation reveals a substantially larger effect, with an estimated increase in lone mothers' employment of 5 percentage points. An employment response of similar size is estimated by Gregg and Harkness (2003), who use a difference-in-difference estimation technique combined with propensity score matching.

Based on a different data set and different estimation methodology, our estimate of an average employment increase of 5 percentage points matches these latter results. However, we find that this average conceals considerable heterogeneity in responses, which varied between a 10 percentage point increase for lone mothers with one pre-school aged child to essentially no effect for mothers of multiple older children. This employment growth was due to both an increase in the rate at which single mothers remained in the labor force and an increase in the rate at which they entered it. The reform also led to significant increases in formal childcare service utilization and childcare expenditures, especially among those who newly entered employment and those who decided to remain employed as a consequence of the reform. Our findings point to the importance of the generous childcare tax credit in WFTC's relative success in attracting single mothers to the labor market.

Equally important and new are our findings indicating that the reform had substantial impacts on other socio-economic outcomes, including substantial reductions in single mothers' subsequent fertility and in the rate at which they married. Given the importance of household income and family structure for later-life child outcomes (Duncan and Brooks-Gunn, 1997), a comprehensive assessment of one of the key objectives of the reform — that of reducing child poverty in single-parent households — should therefore incorporate an assessment of its indirect effects on child welfare through childcare usage, fertility, and marriage.

II. The WFTC Program

Up to April 2003, the main in-work support program in the UK has been the Working Families' Tax Credit, which replaced Family Credit (FC) on October 5th, 1999.[1] Along with other active labor market programs (such as the various welfare-to-work "New Deal" schemes) and the introduction of the National Minimum Wage, WFTC had a crucial part in the central government's antipoverty strategy. By the end of 2002, it reached almost 2.7 million children in 1.4 million families (a 70-percent increase from November 1998 when FC was still in place), and at a cost of around £6 billion a year (approximately 10 billion in current US dollars). To put these figures into perspective, in February 2003, a total of 2.5 million children were living in families claiming at least one of the other key means-tested welfare benefits, including Income Support and Housing Benefits, while the government spending on, say, Income Support alone — the primary cash transfer to low-income nonworking individuals (in many respects similar to AFDC or TANF in the United States) — was around £13 billion a year.

A family needs to meet three basic requirements in order to be eligible for WFTC. First, at least one adult in the family (or the lone parent in a single-parent family) must work 16 hours or more per week.[2] Second, the family must have at least one dependent child. A dependent child is a child, grandchild, stepchild or foster child of the family who is under the age of 16 or under 19 if in full-time non-university education. Third, family savings and capital must be below a given amount (which, in current prices, was set at £8,000 over our entire sample period) and net family income must be sufficiently low. In fact, families with incomes below a specified 'threshold' or 'applicable amount' (which increased from £62.25 per week in 1991 to £92.90 per week in 2001) receive maximum credit; when incomes are greater than the threshold, the maximum credit is reduced by a proportion (known as 'taper rate') of the difference between net family income and threshold. Besides family income and hours worked, the amount of the weekly credit to which a family is entitled also depends on

[1] A more detailed description of the program is in Blundell and Hoynes (2004) and Francesconi and Van der Klaauw (2004). It should be noted that, in April 2003 and thus after the end of our sample period, WFTC was itself replaced by two new tax credits: the Child Tax Credit (CTC) and the Working Tax Credit (WTC).

[2] At its introduction in 1988, Family Credit set the minimum hour cutoff at 24 hours per week, which was reduced to 16 in 1992. In addition, from July 1995, FC was modified to provide an extra £10 credit for those working 30 hours per week or more. This feature was retained by the WFTC reform (with the additional credit set at £11.15 in October 1999).

the number and ages of children and childcare costs, in the form of a basic child credit and a childcare credit.

There are four parameters through which the WFTC reform potentially increased the generosity of in-work support relative to FC. First, the WFTC system substantially increased the credit for younger children in the age group 0-10. The nominal increment of £5 per week represents a 34-percent increase between 1998 and 1999 (while the annual increase of the credit for children in the same age group between 1991 and 1998 was, on average, 6.6 percent only). Second, the income threshold grew by 14 percent from £79 to £90 per week, whereas its average annual growth in each of the previous eight years had been 3.4 percent. Third, the taper rate at which earnings above the threshold are taxed was lowered from 70 percent to 55 percent. Fourth, WFTC is more generous with eligible childcare costs.[3] From October 1994, FC allowed eligible childcare costs (up to a maximum amount, which was £60 per week just before WFTC was introduced) to be disregarded from the calculation of net family income.[4] Under WFTC instead, the disregard was replaced by a separate childcare credit worth 70 percent of childcare costs, subject to an overall limit of £100 per week for one child and £150 per week for two or more children. This meant that the maximum childcare support was £70 per week for a family with one child, and £105 per week for a family with two or more children.

To get an idea of the impact of the reform, consider a lone mother with one child aged 6, who works more than 16 hours per week (but less than 30), has net earnings of £150 per week and pays £60 per week for childcare. In 1999 under the WFTC regime, she would receive a credit of £81.15 per week. In 1998 under the FC regime, the same woman with the same characteristics would receive a credit of £56.80 per week (in 1999 prices), that is 43 percent less than in 1999. If the woman had net earnings of £200 per week, her credit in both years would be lower, but in 1999 she would receive 2.5 times the amount in 1998 (£53.65 versus £21.25 per week). For the empirical earnings and childcare expenditures distributions of mothers working 16 or more hours per week in 1998, and assuming no behavioral responses, the reform implied an average credit increase of about 20 percent (from an average

[3] To be ‘eligible’ (or ‘relevant’) childcare services must be provided by registered childminders, day nurseries and after-school clubs, or certain other special schools or establishments that are exempt from registration. Relevant childcare can be for any child in the family up to age 11 until May 1998, or up to age 12 from June 1998 to May 2000, or up to age 15 from June 2000 onward.

[4] The £60 amount was the disregard for families with one child. In 1998 a disregard of £100 was introduced for families with two or more children.

of £42 to almost £50 per week), with a quarter of these mothers seeing an increase of at least 50 percent.

Unlike FC, WFTC was not administered by the Benefits Agency but by Inland Revenue. In line with the government's effort to reduce the potential stigma associated with claiming in-work benefits, this administrative feature and the receipt of the credit through the wage packet directly from the employer were intended to emphasize that WFTC was indeed a tax credit rather than a welfare benefit (H.M. Treasury, 1998).[5] Therefore, while in many respects WFTC is similar to EITC in the United States, it differs from it in that WFTC has no phase-in rate but instead a minimum hours requirement of 16 per week, it has a higher phase-out rate (taper rate), includes a generous childcare tax credit, and it is administered and paid out differently.

The WFTC reform was also accompanied, preceded and followed by the introduction of new programs and by changes in key parameters of other existing schemes, such as the National Minimum Wage and the various New Deal schemes. There are therefore a number of possible interactions between WFTC and other policy initiatives.[6] While disentangling the effect of each individual policy is beyond the scope of this paper, in our empirical analysis we will attempt to isolate, to the extent possible, the impact of WFTC. A detailed discussion of our methods is deferred until Section IV.

A simplified illustration of the main features of the reform is provided in Figure 1.[7] One of the parameters of the reform was an increase in the credit for those with a child under age 10, an increase which was also applied to those working less than 16 hours (or not working) and receiving Income Support. Therefore, to focus on the main work incentive effects of the program, in Figure 1 we control for this increase in basic benefits by also applying it to the old FC program. In absence of childcare subsidies, there was a gradual increase in benefits with higher hours of work levels. If the mother received Housing Benefit (a rent subsidy), the rate of increase was somewhat slower than shown in the figures, due to the fact that the tax credit was treated as income in other means-tested programs. However, the main features remained the same, with the greatest increases in benefits falling to those in

[5] Most families were paid through the pay packet. The most notable exception (which is not relevant in our study) was for couples in which the claimant was a nonworking partner: in these cases, it was paid to them directly.

[6] For a thorough description of such initiatives, see Dickens and Manning (2002), Blundell *et al*. (2002), and Card *et al*. (2004).

[7] The simplified versions of the benefit schedules and budget constraints apply to a mother with one child under age 11, earning an hourly wage rate of £5.00, and weekly childcare expenses which are linear in the number of working hours above 16, and equal to £28.00 (or £70.00) when working 16 (or 40) hours per week.

full-time employment, many of whom would not have been eligible for a tax credit before the reform. Depending on the amount of childcare expenditures, the childcare component of the tax credit could have represented a considerable increase in generosity of the in-work benefit program, beyond that associated with the reduced earnings tax rate and increased earnings disregard. This is illustrated in panel (a) which also shows the benefit schedule under WFTC in the case the childcare component had been computed as it was under FC (i.e., as an earnings disregard).

To assess the overall work incentives associated with the reform, panel (b) shows the mother's budget constraint in the case where she used paid childcare where again FC benefits and income at hours below 16 were calculated based on the higher basic child credit rate under WFTC. The reform unambiguously improved the financial incentive to take on eligible employment, and especially full-time employment. The effect on hours of work for those already in eligible employment was ambiguous, depending on the relative magnitude of income and substitution effects for this group. Similarly, the childcare tax credit, receipt of which was conditional on eligible employment, had an unambiguous positive effect on labor force participation and an ambiguous effect on hours for those in eligible work. As the credit represented a reduction in the net cost of childcare use, we expect an increase in paid childcare use and in the childcare credit take-up rate, among both those new and those already in eligible employment.

As credit receipt was conditional on having a dependent child, the reform was expected to create greater incentives for single childless women to enter single motherhood. However, it had an ambiguous effect on subsequent fertility choices of lone mothers. First, conditional on current working hours and childcare use, the increments in the child credit (under both IS and WFTC) and in the childcare credit (under WFTC) for existing children led to an income effect, which might have been positive or negative depending on whether children were seen as normal or inferior goods. Second, the same increments implied an increase in the added credit associated with an extra birth, which would have led to a rise in subsequent fertility. But, by increasing the gains from working and reducing the cost of childcare, these changes could have also influenced fertility choices through indirect income and substitution effects associated with changes in labor supply and childcare usage, which could have affected subsequent fertility decisions either positively or negatively. Third, the increase in the income threshold and the decrease in the taper rate similarly might have influenced fertility behavior through income and substitution effects associated with the increased incentives to work 16 or more hours.

Finally, because couples as well as lone parents may receive WFTC, the effect of the reform on marriage decisions is theoretically ambiguous. Eligibility for married couples required at least one of the parents to work sixteen or more hours, but eligibility for the childcare credit was conditional on eligible employment of both parents. As a result we expect that the reform may have been more beneficial to single mothers, but this is an empirical question which we investigate below.

III. Data

The data we use are from the first eleven waves of the British Household Panel Survey (BHPS) collected over the period 1991-2001.[8] Our sample includes unmarried non-cohabiting females who are at least 16 years old and were born after 1940 (thus aged at most 60 in 2001). We exclude any female who was long-term ill or disabled, or in school full time in a given year. The sample includes 3,333 women, of whom 1,507 are lone mothers and 1,826 remain childless during the observation period. 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 9 percent of the women are observed in the same marital state for all the 11 years of the panel, approximately 30 percent of them are observed for at least seven years in the same state. The resulting sample size, after pooling all 11 years for both groups of women, is 14,357 observations (5,283 on lone mothers and 9,074 on childless women). Of the 1,394 single women in the 1999 wave of interviews, 40 lone mothers and 61 childless women (about 7 percent of the sample in that year) were interviewed before the reform implementation date of October 5th. To limit problems of interpretation, they were dropped from the sample analyzed for that year. Their inclusion however does not alter any of our main results.

Table 1 presents summary statistics of the outcomes as well as background characteristics of the two groups of women which we will use as covariates in the analysis below. There are some noticeable differences in characteristics between the two groups. Those who have children tend on average to be younger, less educated, more likely to be nonwhite, and more likely to be in social housing. In addition there appear to be systematic

[8] The households from the European Community Household Panel subsample (followed since wave 7), those from the Scotland and Wales booster subsamples (added to the BHPS in wave 9) and those from the Northern Ireland booster subsample (which started in wave 11) are excluded from our analysis. Detailed information on the BHPS can be obtained at ⟨http://www.iser.essex.ac.uk/bhps/doc⟩.

differences in employment behavior of both groups of women. Compared to unmarried childless women, lone mothers are less likely to work 16 or more hours per week, and have a lower probability of staying in such labor market state and entering it in any given two successive years.[9] They are also less likely to work 30 or more hours per week, or work any positive number of hours, but they are equally likely to form marital or cohabiting unions. The other outcomes listed in Table 1 are analyzed only for single *mothers* (usage of and expenditures on childcare services, and having an additional child), with the exception of entry into lone motherhood which is analyzed for single childless women only.[10]

Figures 2 and 3 plot the time trends for all outcomes between 1991 and 2001. Figure 2 focuses on eligible employment and plots the labor market participation rates at 16 or more hours per week. Panel (a) shows the trends for the two groups of women, while panel (b) disaggregates the lone mothers' patterns into three groups stratified by the age of the youngest dependent child (ages 0-4, 5-10, and 11-18). The data reveal that single childless women had very stable participation patterns over the whole sample period. The participation rates of lone mothers too were stable with a small positive trend up to 1998, when they rose from about 40 to nearly 48 percent.[11] Figure 2(b) suggests that the strongest growth was experienced by women with children in the youngest age group (0-4 years), who increased their participation rate from approximately 30 percent during the 1991-1998 period, to 45 percent in the 1999-2001 period.[12]

The trends in average participation rates at 30 or more hours per week (full-time employment) in Figure 3(a) are similar to those shown in Figure 2(a). Usage of and expenditures on formal childcare services were stable up to 1998, and increased only in

[9] Worked hours are the sum of usual weekly hours of work and usual weekly hours of overtime work.

[10] Most of the figures in Table 1 conform to official statistics and to those reported in related studies (e.g., Blundell *et al*., 2000; Gregg and Harkness, 2003). Perhaps the most notable exceptions are childcare usage and expenditures. Using data from the Family Resources Survey (FRS) for the period 1994-1996, Blundell *et al*. (2000) report that nearly 18 percent of lone parents use formal childcare (rather than 11 percent as in Table 1), and the weekly childcare expenditure is about £57 (rather than £42). Besides differences in time period and data source, the FRS statistics in Blundell *et al*. (2000) refer to families where the youngest child is under 5, while the averages in Table 1 are computed over families where the youngest child is aged 12 or less. However, if we only look at families where the youngest child is under 5, the BHPS figures on average childcare expenditures are £54 per week, very close to the FRS figures.

[11] If the timing of WFTC's introduction were driven by a sudden fall in the employment rate of lone mothers in the years immediately preceding its introduction, then the evaluation can be affected by a "regression-to-the-mean" bias (Cook and Campbell, 1979). The figures presented here, however, do not reveal any such pre-program declines.

[12] Francesconi and Van der Klaauw (2004) explore the possibility that the slight employment rate increase in 1998, as shown in Figures 2 and 3(a), may represent a program announcement effect. Here we adopt a conservative approach by assuming that it is unrelated to the reform, which may cause a slight downward bias in our employment response estimates.

concomitance with the WFTC reform (Figures 3(b) and (c)). The 1999 reform appears also to be associated with a reduction in the entry rates into marital or cohabiting unions (as compared to women without children), a decline in the probability of an additional birth for lone mothers, as well as with a small decline in the entry rate into lone motherhood (Figure 3(d)).

These trends for those directly affected by the reform (single mothers) and those not directly affected (single women without children) strongly suggest that the changes in socioeconomic outcomes observed after 1998 were closely related to the in-work reform introduced around that time. The relative stability of trends in the different socio-economic outcomes for single women without children suggests that the observed changes for single mothers, which coincides with the 1999 reform, were not a result of improved wage opportunities for all women. However, it is also possible that the differential trends were due to changes over time in the composition of both groups. For example, single mothers' age, education, fertility or unobserved skill distributions might have changed compared to single childless women's. To investigate the causal link between the introduction of WFTC and these outcomes further, we will estimate a series of multivariate regression models that compare the outcomes of lone mothers to the outcomes of single women without children, controlling for demographic differences between the two groups as well as changes in these characteristics over time. A similar comparison group approach was used by Eissa and Liebman (1996) and Meyer and Rosenbaum (2000) to assess the impact of the EITC on the labor supply of single mothers, while Blundell and Hoynes (2004) compared employment outcomes of single women with and without children to evaluate the impact of the WFTC reform. The next section discusses our methodology for identifying the responses of British lone mothers to the WFTC reform more formally.

IV. Methods

Non-experimental program evaluations based on multiple pre- and post-treatment periods have been carried out in several different ways (see the comprehensive survey by Heckman *et al.* (1999) and the discussion of the interrupted time-series design by Cook and Campbell (1979)). To relate our approach to those previously adopted for evaluating the impacts of in-work benefit reforms in Britain and the United States, let d_i denote a dummy variable that is equal to 1 if individual i is a lone mother and 0 otherwise, and let s be the time period in which the reform occurs (i.e., s=1999). We model the outcome variable y_{it} as being determined by the following specification

$$(1) \quad y_{it} = \alpha_1 + \alpha_2 d_{it} + (\alpha_{31} + \alpha_{32} d_{it})t + [\alpha_{41} + \alpha_{42}(t-s)]\mathrm{I}(t \geq s) + \beta d_{it}\mathrm{I}(t \geq s) + \mathbf{X}'_{it}\gamma + \theta_i + \varepsilon_{it} ,$$

where the term I(w) is a function indicating that the event w occurs, $\mathbf{X}_{it}$ is a vector of individual characteristics, θ_i represents individual fixed effects, and ε_{it} is an i.i.d. error term, with $E(\varepsilon_{it} \mid \mathbf{X}_{it}, d_{it}, t) = 0$ where $E(\cdot)$ is the mathematical expectation operator. Equation (1) allows for different intercepts (when $\alpha_2 \neq 0$) and different linear trends (when $\alpha_{32} \neq 0$) for single women without children and with children. The parameters α_{41} and α_{42} measure possible shifts in the intercept and slope of the process generating y at the time of the reform. In our case, they capture the effects of all the other (non-WFTC) policy changes that occurred at s (e.g., the introduction of the minimum wage). While our control group of single women without children was ineligible for FC and WFTC benefits and therefore not directly affected by the in-work benefit reform, both groups were potentially affected by the other policy initiatives that took place in that year. By assuming that lone parents would have responded in the same way to these other reforms, we are able to net out the separate impact of WFTC, which is captured in the equation by β.

Our approach improves over the widely used "difference-in-difference" (DD) method (Eissa and Liebman, 1996; Blundell and Hoynes, 2004), applied to data from few repeated cross-sections, which assumes a common trend (imposing $\alpha_{32} = 0$).[13] As indicated by Figures 2-3 single childless women exhibit a number of pre-reform trends that differ from those of lone mothers, suggesting that it is important to allow for these in the analysis.[14] In addition, by using multiple observations before and after the occurrence of the policy change, unlike the DD approach (which imposes $\alpha_{42} = 0$), we allow for both a common jump and a change in the slope in 1999. In this respect our approach can be considered as a simple extension of a "difference-in-difference-in-difference" (DDD) estimation approach applied to panel data.[15]

[13] One concern with DD regressions is that standard errors may be misstated in the presence of serial correlation of outcomes for the same individual over time (Bertrand *et al*., 2004). Our fixed-effects specifications, however, directly account for time-invariant unobserved heterogeneity.

[14] We will also report results from alternative specifications with group-specific quadratic trends, in which case equation (1) includes the additional term $(\alpha_{51} + \alpha_{52} d_{it})t^2$.

[15] Under specification (1) without individual characteristics and fixed effects, a DDD estimator defined as

To control for potential differences in group-specific compositional changes over time, we have a set of standard individual characteristics in $\mathbf{X}_{it}$ (e.g., age, education, region of residence, number and ages of children). Because single childless women tend to be more concentrated at the bottom and top ends of the age distribution, it is important to account for age effects, which we do by including a quartic polynomial in age. In addition, as single mothers on average have less education we performed sensitivity checks that will be discussed later in the paper by replicating our entire analysis using a more restricted control group, consisting of single childless women with low educational attainment. Because we use panel data, we also account for compositional changes in unobserved characteristics with the inclusion of individual-specific fixed effects. Unlike studies based on cross-sectional data, this permits us to address the possibility that time changes in lone-mother status (i.e., inclusion in our sample) as well as changes in time-varying individual characteristics are endogenous to the policy reform as long as the fixed effect represents the source of this endogeneity.

By following the same individuals over time, we can examine whether the introduction of WFTC led to changes in the rate at which single women entered and left the labor force. That is, we can directly assess the impact of WFTC on year-to-year employment transitions. This can help us understand whether any given WFTC effect is associated with a change in the rate at which individuals entered the labor force and/or with a change in the rate at which people left it. We implement this analysis by estimating separate outcome equations (1) for each value of $y_{i,t-1}$.

It is important to point out the main identification condition underlying our approach. We explicitly assume that, other than the introduction of WFTC, there are no contemporaneous shocks that affect the *relative* outcomes of the treatment and control groups. The increase in basic child benefits under Income Support between 1998 and 1999 may be problematic in this respect. In terms of employment outcomes, however, this increase is modest and implies a negative income effect that could lead to a (small) downward bias in our effect estimates. Our estimates may then represent a lower bound on the true effect.

$$\{[\mathrm{E}(y_{i,s-1+k} \mid d_i = 1) - \mathrm{E}(y_{i,s-1} \mid d_i = 1)] - [\mathrm{E}(y_{i,s-1} \mid d_i = 1) - \mathrm{E}(y_{i,s-1-k} \mid d_t = 1)]\}$$
$$- \{[\mathrm{E}(y_{i,s-1+k} \mid d_i = 0) - \mathrm{E}(y_{i,s-1} \mid d_i = 0)] - [\mathrm{E}(y_{i,s-1} \mid d_i = 0) - \mathrm{E}(y_{i,s-1-k} \mid d_t = 0)]\}$$

will identify the treatment effect β (the time subscript k indicates the length of the time periods over which the differences are computed). See Francesconi and Van der Klaauw (2004) for a more extensive discussion of DD and DDD methods.

V. Impact on Employment

A. Benchmark Estimates

Table 2 shows the estimated impacts of the WFTC reform on three employment measures (eligible employment, full-time employment, and labor force participation).[16] For simplicity of interpretation, we only report least squares estimates based on linear probability models. Marginal effects estimates from Chamberlain fixed-effects logit models were very similar. In the table we present constant treatment effect estimates from model (1) with individual fixed effects θ_i and group-specific pre-program trends. Column (i) reports our baseline results obtained from the whole sample of single women with linear trends. The other three columns provide robustness checks. The estimates in column (ii) are obtained from the subsample in which the control group of single childless women is limited to those with educational qualifications below A level ('low education sample'). Columns (iii) shows results for the whole sample with quadratic trends, while column (iv) reports the estimates from the low education sample with quadratic trends.

Focussing on the first estimate in column (i), we find that the rate at which lone mothers worked 16 or more hours per week increased by a statistically significant 5.1 percentage points.[17] This estimate is remarkably close to those reported in Brewer *et al*. (2005) and Gregg and Harkness (2003), who applied different methods to different data from ours.[18] Our estimate implies an eligible-employment elasticity with respect to net income of about 1.10. This falls within (albeit at the high end of) the 0.69-1.16 range of comparable elasticities reported by Hotz and Scholz (2003) for the United States. Looking across columns, we notice that restricting the control group to childless women with qualifications lower than A level reduces only slightly the effect to 4.8 percentage points (column (ii)). Likewise, the introduction of quadratic pre-program trends leads to treatment effects of 5

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[16] Francesconi and Van der Klaauw (2004) present evidence on several additional outcomes, including monthly earnings, FC/WFTC receipt and award, and IS receipt. This analysis indicated that the WFTC reform led to significant increases in single mothers' average monthly earnings and receipt/award of the tax credit, and to a substantial reduction in IS participation.

[17] It is worthwhile noting that the estimates of α_{32} and α_{42} are around 0.0025 (s.e.=0.0011) and 0.017 (s.e.=0.007) respectively, implying that a DD estimator would lead to biased estimates. The same pattern of results emerges also from the other three specifications.

[18] Since there might have been delayed responses to the program, we also estimated variants to equation (1) that allow for a different treatment effect in each of the three years following the reform and available in our dataset. The estimates, not shown here for brevity, indicate that the largest employment impact emerged in 1999, that is, immediately after the introduction of the reform, when it led to an increase in the employment rate of 7 percentage points. The employment effect fell in subsequent years to 2 points in 2000 and 4 points in 2001. The finding of a program-introduction effect is consistent with results reported in Blundell and Hoynes (2004).

percentage points for the whole sample (column (iii)) and 4.6 percentage points for the restricted sample (column (iv)). In general, therefore, the baseline estimate in column (i) appears to be robust to alternative comparison group and trend specifications.

As discussed earlier, in absence of the childcare subsidy, the increase in net income induced by the introduction of WFTC can be expected to be small or modest below 25 hours of work per week. At higher hours levels, however, the reduction in the WFTC taper rate leads to greater returns to working. The childcare subsidy provides an additional incentive at any level of eligible employment to work and use paid childcare. The estimates in the second row of Table 2, in fact, suggest that the positive labor supply response of single mothers was predominantly driven by an increase in full-time employment (working 30 hours per week or more). The rate at which lone mothers worked full time increased by 4.7 percentage points over the post-reform period (column (i)). Different sample restrictions and trends modelling lead to similar results (columns (ii)-(iv)).

To analyze the extent to which the increase in full-time employment and in eligible employment was due to an overall increase in employment of single mothers, rather than an increase in hours among those working, we consider the WFTC impact on the labor force participation rate (third row of Table 2). Regardless of the specification, the estimates indicate that a large proportion of the increase in employment was due to the increase in the proportion of single mothers participating in the labor market. Indeed, using hours worked as dependent variable on the subsample of workers only (not shown), we find virtually no effect. Combined, these results suggest that the WFTC had a strong substitution effect on previously nonworking lone mothers and relatively similar income and substitution effects on already working women.

B. Employment Transitions

Depending on the tightness of the labor market, an increase in work incentives can differentially affect entry and exit rates. If new jobs are sparse and if most exits from the labor market by single mothers are voluntary, we might expect most of the employment response to come in the form of a lower exit rate out of eligible employment. More generally, we might also expect to see an increased entry rate into eligible employment. To investigate this, we next analyze WFTC's impact on employment transitions, by estimating its effect both on the probability of staying in eligible employment (i.e., conditioning on $y_{i,t-1} = 1$), and on the probability of starting a job with 16 or more hours of work per week (i.e., conditioning on $y_{i,t-1} = 0$). We define the former as the persistence probability and the latter as the entry probability. The corresponding effect estimates are reported in Table 3. They indicate that the introduction of the in-work benefit reform significantly increased lone mothers' persistence rates by 6.5 percentage points after 1998. Entry rates into WFTC-eligible jobs show similar patterns. They rose on average by almost 6 percentage points over the post-implementation period. These results are broadly confirmed by the estimates obtained with the restricted sample and the inclusion of quadratic trends.[19]

C. Estimates by Number of Children and Age of the Youngest Child

The WFTC reform could have generated different labor market responses depending on the number and ages of children, because it virtually eliminated existing differences in the child credit amount provided to children of different ages under Family Credit, and because of the new childcare credit. To analyze this we estimate separate treatment effects by the number of dependent children, and distinguishing by age of the youngest child in intervals 0-4, 5-10, and 11-18.[20] The results shown in the top panel of Table 4 indicate that the strongest increase in eligible employment emerges in the case of lone mothers with one pre-school aged child. Eissa and Liebman (1996) and Meyer and Rosenbaum (2001) in the case of the EITC, and Gregg and Harkness (2003) for the case of the WFTC, also reported larger employment effects for mothers with younger children (column (i)), but the differences they report are less dramatic than those found here. A lone mother with one child aged 0-4 increased her probability of being in eligible employment by 10 percentage points, and a lone mother with one child aged 5-10 experienced an increase of about 8 percentage points. But for a single

[19] The transition probability estimates line up well with the estimates shown in Tables 2. For example, given that 0.414 percent of lone mothers were in eligible employment in 1998, a rough calculation suggests that an additional 2.7 percent (0.414×0.065) were in employment in 1999 through the increase in persistence rates and a further 3.3 percent ((1-0.414)×0.057) were in employment in 1999 through the higher entry rates. These changes lead to a total effect of 6.0 percentage points, which is slightly higher than the 5.1 percentage point figure reported in Table 2.

[20] Specifically, our estimated equations take the form:

$$\begin{aligned} y_{it} = {} & \alpha_1 + \alpha_2 d_{it} + (\alpha_{31} + \textstyle\sum_j \alpha_{32}^{(j)} K_{it}^{(j)} d_{it})t + \left[\alpha_{41} + \alpha_{42}(t-s)\right]\mathrm{I}(t \geq s) + \textstyle\sum_m \beta^{(m)} N_{it}^{(m)} d_{it} \mathrm{I}(t \geq s) \\ & + \mathbf{X}'_{it}\gamma + \theta_i + \varepsilon_{it}, \end{aligned}$$

where $K_{it}^{(j)}$ is an indicator that equals one if the youngest child for mother *i* at time *t* is in age group *j*, *j* = 0-4, 5-10, 11-18, and zero otherwise; and $N_{it}^{(m)}$ is another indicator that equals one if lone mother *i* at time *t* has children in group *m*, where *m* is either of the following six mutually exclusive categories: one child aged 0-4, one child aged 5-10, one child aged 11-18, two or more children with the youngest aged 0-4, two or more children with the youngest aged 5-10, or two or more children with the youngest aged 11-18. The latter indicators are also included in $\mathbf{X}_{it}$.

mother with one child in the oldest group, that increase was of the order of only 3.5 percentage points.

Lone mothers with a greater number of children generally had a smaller labor supply response and, without exceptions, this was never statistically significantly different from zero. These results are robust to the other specifications (columns (ii)-(iv)), and the estimates for full-time employment and labor force participation (not presented here) show the exact same pattern.

The other two panels of Table 4 report the treatment effect estimates on persistence and entry rates for mothers with varying number of children and by age of the youngest child. In line with our previous results, the largest impact of the WFTC reform on employment persistence and entry rates emerged for women whose children were in the youngest group. From column (i) we see that the average persistence rate for a single mother with one child aged 0-4 increased by almost 13 percentage points relative to the rate of a corresponding single childless woman. This effect declined to 7 percentage points for a mother with one child aged 5-10, and dropped to 2.8 percentage points for a mother with one child aged 11-18. The entry rate estimates show a similar pattern by child's age, but their decline across age groups is less steep and their overall magnitudes are smaller compared to those of the corresponding estimates of the persistence rates, except for the oldest age group. For mothers with two or more children, the effect of the reform on the probabilities of remaining or entering eligible employment was always small and statistically insignificant regardless of the age of the youngest child. Again, all these results are robust across specifications (columns (ii)-(iv)).

D. Child Credit Component

The finding that the employment effects were stronger for lone mothers with young children provides important clues as to what WFTC parameters may have contributed to the observed responses. As mentioned earlier there are two components of the reform which could lead to differential treatment by child age. A first explanation, which relates to the increase in the child age-specific component of the tax credit, is analyzed here. The other is considered in the next section.

The credit for children aged 0-10 increased by about 25 percent relative to the credit for children aged 11 or more. However, the generosity of IS payments to workless lone mothers with children aged under 11 also grew by the same amount. The net result of these changes is a pure income effect on labor supply, which goes in the opposite direction: we would expect a smaller labor supply increase for mothers of younger children.[21]

To investigate further whether the childcare component of WFTC played any role, we compare differences in labor supply from 1998 to 1999 between lone mothers who had a child aged 10 in those two years and lone mothers who had a child aged 11 in the same two years. As the former group experienced a much larger increase in the child credit component relative to the latter, a comparison of the corresponding responses in employment behavior from 1998 to 1999 for these two groups provides an indication of the importance of the child credit schedule changes. We computed difference-in-difference estimates for both eligible employment and full-time employment (the results are not shown for brevity, but can be found in Francesconi and Van der Klaauw (2004)).[22] The Wald estimates, which are very small and statistically insignificant, indicate that the child credit increase for young children does not play any role in explaining the employment effects of the reform. Controlling for differences in individual characteristics and increasing the two comparison samples to mothers with children aged 6-10 and 11-14 respectively did not alter this result.

VI. Impact on Childcare Use and Role of the Childcare Credit

The second explanation is based on the increase in the tax credit provided to cover childcare costs. Although all lone-parent households may benefit from this provision, it is arguably lone mothers with children under school age who could benefit most from this incentive. WFTC provided much greater support for childcare than Family Credit did, in a number of ways. It added the childcare element towards the overall award, increased the maximum level of support for childcare costs, and offered support for a wider age range, for children aged up to 15 (or 16 if disabled) rather than up to 11. In conjunction with the large labor supply effects documented above, we therefore expect to observe a sizable impact of WFTC on both the use of eligible childcare services and childcare expenditures.

Although the BHPS distinguishes between formal (or paid) and informal childcare arrangements, it collects information on childcare only for working mothers who are responsible for children aged 12 or under. Our analysis therefore cannot consider childcare

[21] However, positive treatment effects associated with higher family income cannot be totally excluded. Such effects could occur if low-wage credit-constrained mothers of young children did not work because, for example, they could not afford available childcare services without the additional tax credit.

[22] This evaluation approach can be straightforwardly interpreted in a Regression-Discontinuity (RD) framework, where those with children just below the age cutoff (those aged 10) are compared to those just above the cutoff (those aged 11). See Hahn *et al*. (2001) and Van der Klaauw (2002) for discussions of the identification and estimation of treatment effects in case of an RD design.

arrangements for nonworking single mothers (who in any case were not WFTC eligible), and it cannot take account of childcare arrangements for children aged 13 or more (although this omission might have only minor consequences on our results since relevant childcare subsidies under WFTC started to cover 13-15 year olds only from June 2000 onward). The estimates in Table 5 confirm our expectation, showing that the introduction of WFTC led to an increase in the use of paid childcare services of about 3 percentage points regardless of the trend specification (first row in panel A).[23] This average effect represents a 35 percent increase relative to the pre-1999 levels, and lines up well with the statistics drawn from the Families and Children Study data reported in McKay (2003). The finding of a relatively large response in childcare use is also consistent with own price effect estimates reported for the US (Anderson and Levine, 2000). As will be discussed later, it represents a large proportional increase in the use of formal childcare services among those newly entering eligible employment. At the same time, the corresponding average childcare expenditures went up by about £17-19 per week (second row).

Differentiating by the number and ages of children we find a similar pattern to the reform's impact on eligible employment (panel B). Lone mothers with only one child aged 0-4 experienced the greatest increase in the probability of using paid childcare services of 4 percentage points, compared to an increase of 3 percentage points for those with one child aged 5-10. The effect further diminishes with the number of children in the household. These results are robust to the inclusion of quadratic trends.

To examine the interaction between employment and childcare responses, we next analyze the WFTC impact on the rate at which nonworking lone mothers enter eligible employment while using paid childcare, as well as its impact on the rate at which single mothers who were previously in eligible employment increased their use of paid childcare while working. Estimates from both analyses are reported in panels C and D of Table 5. A significant fraction of lone mothers who entered eligible employment as a result of the reform did so by also choosing to use paid childcare. In fact, when compared to the estimates shown in Tables 3 and 4, the results in panel C of Table 5 suggest that, for mothers with one child aged 0-10, the work-childcare combination increased by about 3-5 percentage points, and accounted for approximately 60 percent of the increase in the labor market entry rate.

[23] Because the analysis in panels A and B is based on models estimated on the subsample of lone mothers only, there is no need to consider a low education comparison sample (former specifications (ii) and (iv)). The results shown in panels C and D are qualitatively similar to those found with a sample restricted to women with lower educational levels. These results therefore are not reported.

Similarly, the reform led to a sharp increase in the use of paid childcare amongst those who were already working 16 or more hours (panel D). The overall effect of 3.8 percentage points represents 55 percent of the estimated rise in the employment persistence rate reported in Table 3. Again, the largest effects are found among single mothers with one child aged either 0-4 or 5-10: the rate of their childcare utilization increased respectively by about 7 and 4 percentage points, regardless of how trends have been modelled.

Another way to illustrate the importance of the childcare subsidy is to relate it directly to the demand for paid childcare services by working mothers with pre-school children and school-aged children. As the demand by the former should be higher, we expect a greater response for this group. Of course, there may be a differential response to WFTC's work incentives by child age irrespective of the demand for childcare. However, in absence of any childcare needs we would expect mothers of four-year olds in 1999 to respond similarly to the WFTC reform as mothers of five-year olds. Any difference in response behavior between the two groups could then be legitimately attributed to differential childcare needs, given that children in the UK are legally required to start attending school after their fifth birthday. We performed a number of regressions on eligible employment and full-time employment distinguishing between specific child-age levels between age two and age eight. The results, not shown for brevity, document that the largest employment responses are found among mothers of three- and four-year olds, with no discernable effect for mothers of five-year olds and positive but smaller effects for mothers of children between the ages of six and eight. Especially noteworthy is the much smaller estimated increase in eligible employment around the age of school entry, which represents statistically significant drops from 15 and 18 percentage points for three and four year olds, to 2 and 4 percentage points for 5 and 6 year olds.

These estimates provide strong evidence suggesting that the childcare subsidy component of WFTC played a key role in producing the estimated large employment effects for single women with young children. The findings are new. Earlier studies on the WFTC reform did not explicitly examine it and generally assigned little importance to this component. As summarized by Blundell and Walker (2001), previous research either ignored childcare costs or has assumed a fixed relationship between childcare expenses and hours of work. For example, using simulations based on pre-reform data, Blundell *et al.* (2000) play down the importance of the childcare credit, arguing that while the reform was associated with a rapid increase in the take-up rate of the childcare credit (of over 150 percent), the overall take-up rate among WFTC recipients in 2000 was only 10 percent. Our new findings however suggest that the increase in the childcare credit take-up rate was highly concentrated among newly eligible lone mothers, and played a key role in explaining the employment effects, and in particular those for mothers with young children. Relying on a fixed pre-reform hours-childcare expenditure relationship in predicting employment responses is therefore likely to lead to invalid inferences, especially among mothers with children of pre-school age.

Finally, while our results point to the importance of the childcare credit, we cannot entirely exclude the possibility that the observed responses, including the larger employment responses among mothers of younger children, were in part due to a reduction in the taper rate and the way in which the tax credit was administered. The reduction in the taper rate is likely to have played an additional important role in the increase in the WFTC caseload and the overall positive employment effect for all lone mothers. Moreover, while these changes applied to all lone parents equally, independently of the age of their children, it could have had a larger impact on mothers with younger children, if a higher proportion of them were closer to their reservation value for eligible employment. However, the evidence presented earlier of a sharp drop in the employment response of single mothers with four- and five-year olds suggests that this cannot be the whole story, with the generous childcare credit providing a much more plausible explanation. Thus, it is likely that the interaction between the childcare subsidy and the other components of WFTC has played a crucial role. This line of argument is supported by the high childcare-price elasticities of labor force participation found for mothers of pre-school children both in Britain (Viitanen, 2005) and in the United States (Blau, 2001). Interestingly, an interpretation that stresses the importance of the childcare credit component of WFTC is further corroborated by evidence of increases in both the take-up of childcare subsidies and in the supply of registered childcare places during the post-1998 period reported in Gray and Bruegel (2003).

VII. Impacts on Marriage and Fertility Choices

As discussed in Section II, the structure of WFTC may have affected lone mothers' partnership decisions because the program's eligibility and benefit rules depend on a woman's living arrangement. In addition, through the increase in the basic child benefit amounts (both under IS and WFTC) as well as more generous childcare support, it may have affected subsequent fertility decisions of single mothers. There was also a cap on childcare support, and WFTC's positive impact on employment may in fact have increased the cost of having additional children. As compared to the huge body of research on the effect of welfare reforms on marriage and fertility in the United States (e.g., Eissa and Hoynes (2000), Schoeni and Blank (2000), Hotz and Scholz (2003), Bitler et al. (2004), and references therein), this literature is virtually nonexistent for Britain.

Table 6 presents effect estimates from linear probability models of transitions into partnership (marriage or cohabitation) for the sample of single childless women and lone mothers. We show overall effects and effects by age of the youngest child and number of children for the four usual sample/trend specifications. For each woman, the dependent variable takes value zero if the woman remains single, and value one if she married, and after that point her observations are censored.

The overall estimates from specification (i) indicate that the implementation of WFTC led to a sharp significant reduction in partnership rates.[24] Using single childless women as comparison group, the estimates imply that on average, with the WFTC reform in place, lone mothers were 2.4 percentage points less likely to form a union. This effect is large, representing a 28-percent change with respect to the average annual (re)-partnership rate of 8.5 percent during the sample period (see Table 1). Most of that effect was driven by mothers of pre-school children, who experienced a reduction in the chances of forming a partnership by about 2.8 percentage points if they had one child. If their child was older, their probability of marriage was instead not significantly reduced by the reform, although for mothers with two or more children with the youngest being aged 10 or less we found again strong negative and significant effects ranging between 1.5 and 2.5 percentage points. These results are robust to the more stringent selection on the educational attainment of single childless women (column (ii)), as well as to the inclusion of quadratic trends, although the overall effects in specifications (iii) and (iv), while of similar magnitude, lose their statistical significance.

To see if the effects in column (i) are stronger for younger individuals, we split the full sample in two groups of women (those aged 30 or less and those aged more than 30), and re-estimated transitions into partnership for the two groups separately using linear trends. A similar exercise was repeated for eligible employment. The results, not shown for convenience, reveal that the reduction in partnership rates affected all lone mothers regardless of their age. Likewise, the probability of being in eligible employment seems to have increased uniformly in the two groups of women, with the slightly greater effect for young mothers a difference of about 0.4 percentage points being not statistically significant at

[24] This result is qualitatively similar to much of the available evidence for the United States (e.g., Bitler *et al*., 2004).

conventional levels. These results suggest that the employment effects were not driven by younger women who could have been more likely to form a partnership or (as illustrated below) have an extra child.

Panel A of Table 7 reports estimates from models of transitions into birth for the subsample of lone mothers from the second year they were observed in the panel onwards. These estimates reveal that the partnership rate changes documented earlier were accompanied by a comparably large reduction in fertility rates among lone mothers. Overall, the post-reform risk of having an additional child decreased by 0.7 percentage points (which represents almost a 20-percent change over the average annual birth rate for lone mothers during the sample period), but this effect is not statistically significant regardless of how trends are modeled. Similar results emerge when we consider fertility responses of mothers with children in different age groups. We also found no significant difference in the estimated effects when distinguishing women by age, as was the case for the partnership rate.

To the extent that WFTC allocated more benefits to single women with children than the previous Family Credit system did, it could have provided greater incentives for entering lone motherhood. Thus, the social benefits documented so far in terms of better labor market outcomes for lone mothers could have been offset or reversed if the new program encouraged a greater proportion of women to become lone mothers. We explore this possibility by estimating linear probability models of transitions into lone motherhood for the subsample of single childless women only.[25] The results in panel B of Table 7 show little evidence of this unintended effect. In fact, the introduction of WFTC had the opposite impact, reducing the propensity of single childless women to form lone-mother households by almost 0.2 percentage points, which represents a further 15-percent decline over the average rate of entry into single motherhood for single childless women during the panel years (linear trend). This effect is however not statistically significant. Similar results are found with a quadratic trend specification, and with the low education sample (not shown). Interestingly, these results are in line with those presented for the United States by Moffitt (1994) and Hoynes (1997), although they analyzed welfare effects on female headship for married mothers, thus through separation or divorce.

[25] Leaving aside parental deaths, a single-mother household can be observed either after a marital dissolution among married mothers or after a fertility decision (becoming a mother) among single childless women. The focus given here is on the latter. We do not analyze the transition into single motherhood for married women with children because some of them were potentially eligible to FC/WFTC, and their behavioral responses then could have had complex interactions with other margins (e.g., employment, and partnership formation). This issue bears investigation in future work.

VIII. Conclusions

In October 1999, the Working Families' Tax Credit replaced Family Credit as the main package of in-work support for low-income families with children in Britain. This paper examines the impact of WFTC on lone mothers using for the first time data drawn from the British Household Panel Survey and collected between 1991 and 2001. Our study identifies the effect of the reform through comparisons of changes in behavior for lone mothers and single women without children. It contributes to the existing literature in two important ways. First it considers effects on a wide set of socioeconomic outcomes, some of which have never been examined before in the British context. Second, it uses specific aspects of the reform design and the panel nature of the data to understand how the estimated responses came about and identifies which parameters of the reform were more likely to explain such effects.

We stress five main findings. First, lone mothers responded to the financial incentives of the reform by working substantially more. The introduction of WFTC is estimated to have led to an average increase of about 5 percentage points in the fraction of lone mothers who worked 16 or more hours per week, with almost all this increase being in full-time employment. Second, this large response was due to both the higher rate at which single mothers remained in the labor force and the higher rate at which they entered it. Third, the strongest effects emerged for mothers with one child under five, who increased their participation rates by about 10 percentage points. We instead find no effects for mothers with multiple older children. This is likely to reflect the smaller employment elasticities for (married) mothers with older children found in the literature (e.g., Blundell *et al.*, 1998; Blau and Kahn, 2005). The institutional cap imposed under WFTC on the childcare subsidy for mothers with two or more children might have also played a role but its extent could not be ascertained with our data. Fourth, there are important (and perhaps unintended) effects on lone mothers' behaviors other than on employment. In particular, there is evidence of a significant reduction in the rate at which single mothers formed cohabiting and marital unions, and some weaker evidence of a decline in single mothers' subsequent fertility. Fifth, among the policy parameters that had a part in explaining the estimated large employment responses, a great deal of evidence points to the role played by the generous childcare tax credit component of WFTC. The reform led to a 35 percent increase in the use of paid childcare. About 60 percent of the increased entry rate in eligible employment was attributable to lone mothers who also chose paid childcare arrangements, and the effect was concentrated among mothers with pre-school aged children. Similarly, among single mothers who continued to be in employment, more than 50 percent of their greater post-reform labor market attachment is observed in conjunction with paid childcare services.

This latter finding, indicating that a childcare credit can be an effective tool in attracting single mothers to the labor force, also informs the current childcare policy debate in other countries, including the United States. While several states have their own childcare subsidies and tax credits, the main federal childcare program is the Child and Dependent Care Tax Credit. Theoretically, the credit can be worth up to 35 percent of childcare costs, but in practice — given a maximum limit on eligible expenses and the fact that the rate phases down with income — most families receive a much lower 20 percent credit. The credit has been criticized for being regressive and poorly targeted, which is due in large part to the fact that the credit, unlike WFTC, is non-refundable (Burman *et al.*, 2005). As a result, low-income families who have no income tax liability to offset do not benefit from the credit, while higher-income households claim a disproportionate share of the subsidy. As current work requirements for welfare recipients in the United States appear to reflect society's preferences for low-income single parents to work outside the home, even if they have young children requiring care, it may be worthwhile to consider childcare subsidies that help make this feasible. Given its effectiveness, the British policy of a more generous, refundable credit which is paid on a monthly basis (instead of an end-of-the-year payment) seems worthy of consideration to help induce lone mothers, especially those too cash-constrained to pay for childcare, to enter the labor market.

The results that WFTC played a key role in the employment increases for lone mothers suggest therefore that in-work benefit policies can be successful in encouraging work among lone mothers. But a comprehensive evaluation must also take into account other results — such as the reduced marriage rates and increased childcare use among single women with children — that measure the effects of the reform not only on mothers' wellbeing but also on their children's. Whether and how WFTC and its successors, however, will alleviate child poverty or deprivation through better child outcomes (such as greater cognitive development and mental health, fewer truancy and early behavior problems, and higher educational attainment) remains to be seen.

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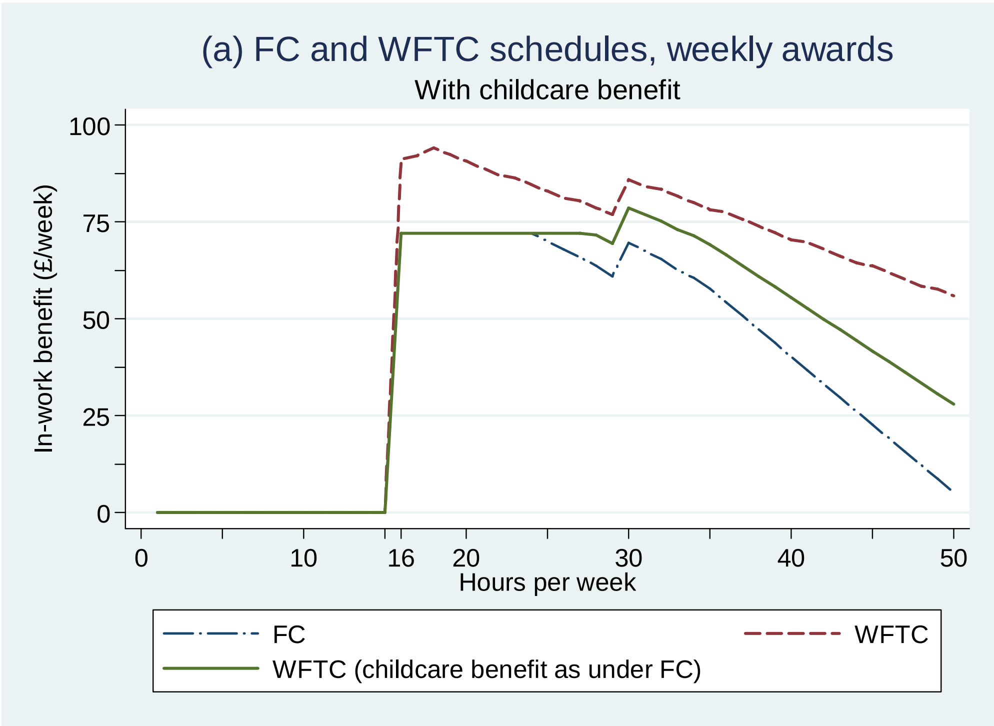
Figure 1. FC and WFTC schedules and budget constraints



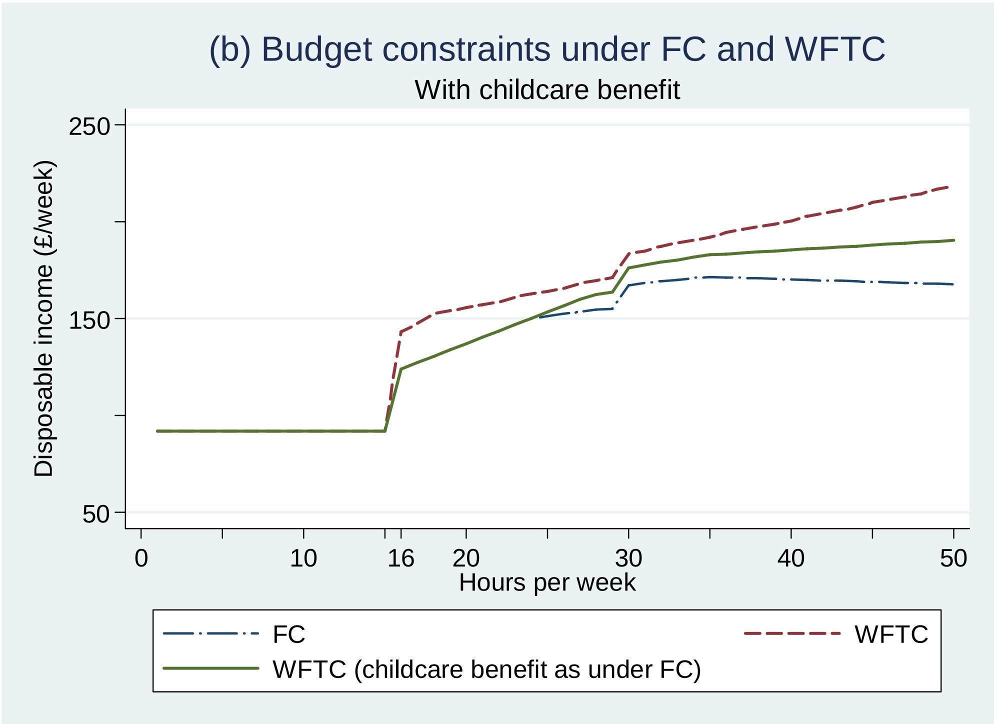
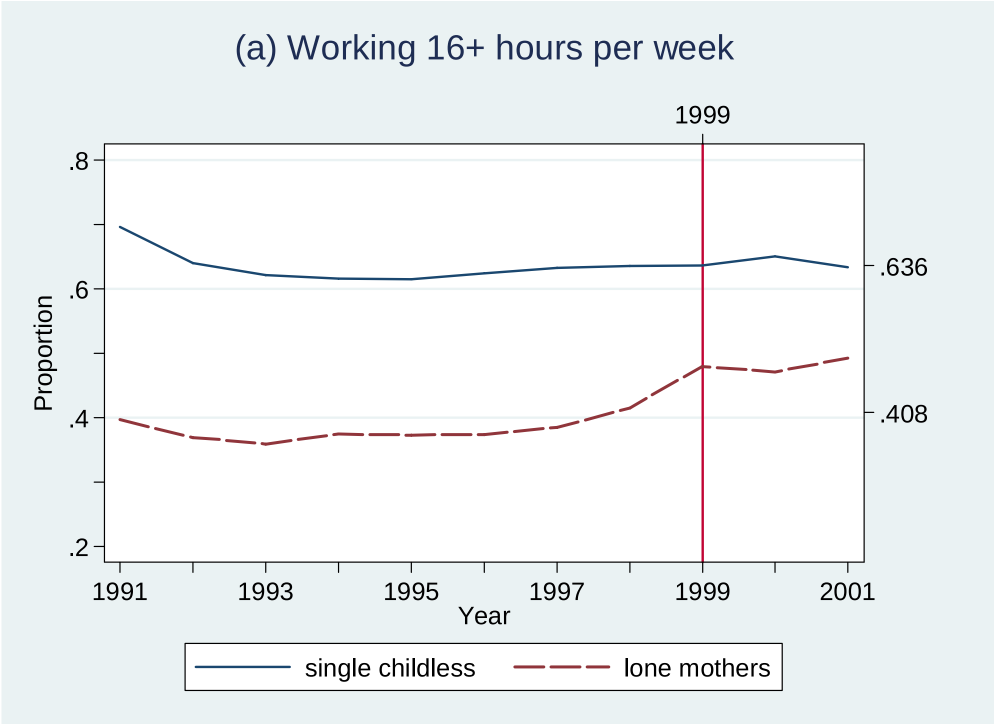


Figure 2. Working 16 or More Hours per Week – Single Childless Women and Lone Mothers



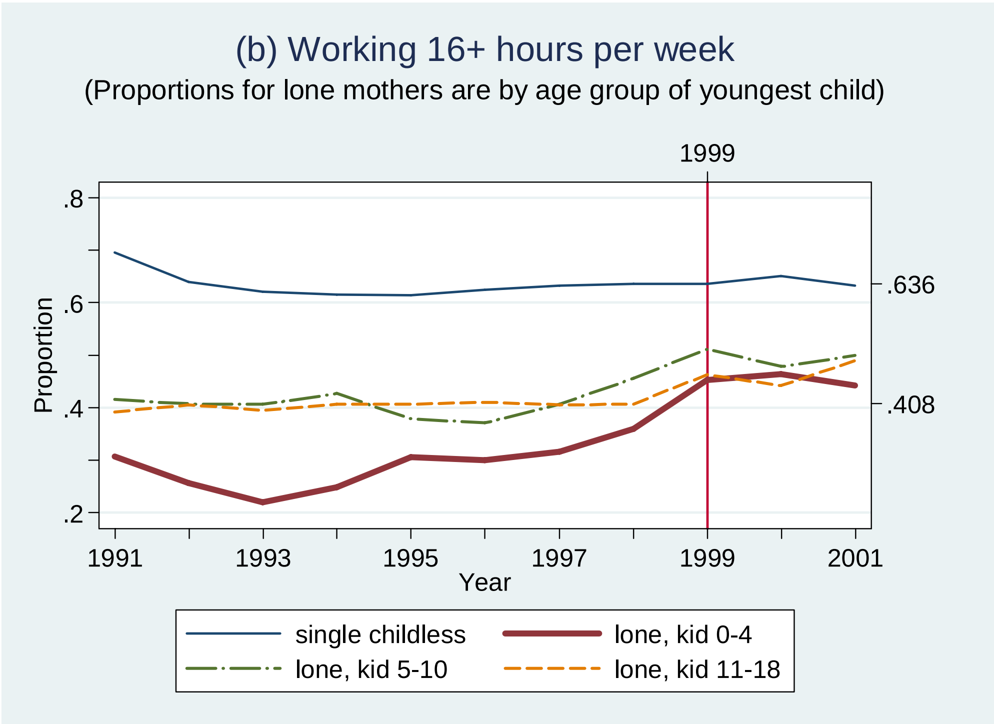


Figure 3. Other Outcomes for Lone Mothers and Single Childless Women

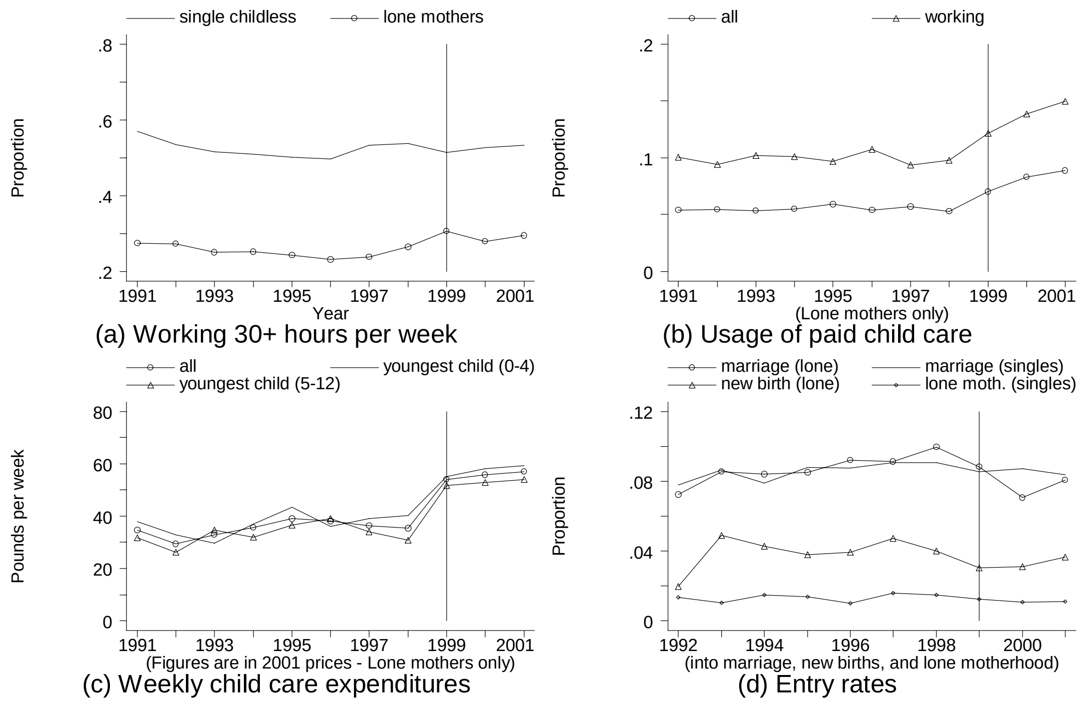


Table 1. Summary Statistics

Variable	Unmarried women without children	Lone mothers
Outcomes		
Working 16 or more hours per week	0.636	0.408
Working 16 or more hours per week by age of youngest dependent child:		
0-4		0.334
5-10		0.432
11-18		0.420
Transition probabilities of working 16+:		
Persistence probability	0.908	0.651
Entry probability	0.268	0.195
Working 30 or more hours per week	0.525	0.264
Labor force participation (working 1 or more hours per week)	0.726	0.595
Usage of paid childcare (all lone mothers)[a]		0.062
Usage of paid childcare (all working lone mothers)[a]		0.109
Weekly childcare expenditures (in 2001 pounds)[a]		42.70
		(33.34)
Entry into marriage/remarriage	0.079	0.085
Birth rates (lone mothers only)		0.037
Entry into lone motherhood (single childless only)	0.013	
Main explanatory variables		
Age	31.319	28.541
	(12.410)	(11.302)
Education:		
No qualification	0.168	0.177
Less than O level/GCSE	0.080	0.121
O level/GCSE (or equivalent)	0.209	0.343
A level (or equivalent)	0.192	0.133
Higher vocational qualification	0.186	0.161
University degree or more	0.143	0.045
Ethnic origin:		
White	0.957	0.916
Black	0.021	0.038
Indian	0.007	0.022
Pakistani/Bangladeshi	0.003	0.011
Chinese or other	0.012	0.013
Number of children by age group:[b]		
0-4		0.252
		(0.529)
5-10		0.589
		(0.754)
11-18		0.760
		(0.752)

Housing tenure:		
Owner	0.594	0.581
In social housing	0.203	0.346
In privately rented accommodation	0.202	0.073
Number of person-wave observations	9,074	5,283
Number of women	1,826	1,507

[a] Computed over single-mother households where the youngest child is aged 12 or less.
[b] Averages are computed over the entire subsample of lone mothers. If computed over the three specific subsamples of lone mothers in each child group, the averages (standard deviations) are: 1.178 (0.461), 1.314 (0.562), and 1.293 (0.523) respectively.
Notes: For convenience, the table does not report summary statistics on region (16 dummies). Standard deviations are in parentheses.

Table 2. Effects of Welfare Reform on Employment

	(i) Full sample with linear trends	(ii) Low education sample with linear trends	(iii) Full sample with quadratic trends	(iv) Low education sample with quadratic trends
Eligible employment[a]	**0.051**	**0.048**	**0.050**	**0.046**
	(0.018)	(0.018)	(0.022)	(0.021)
Full-time employment[b]	**0.047**	**0.042**	**0.046**	**0.042**
	(0.021)	(0.016)	(0.021)	(0.020)
Labor force participation[c]	**0.058**	**0.053**	**0.055**	**0.050**
	(0.022)	(0.020)	(0.022)	(0.023)
Number of women	3,333	2,322	3,333	2,322
Observations	14,357	9,427	14,357	9,427

[a] The dependent variable takes value one if a woman works 16 or more hours per week, and zero otherwise.

[b] The dependent variable takes value one if a woman works 30 or more hours per week, and zero otherwise.

[c] The dependent variable takes value one if a woman works (any positive number of hours), and zero otherwise.

Notes: Standard errors are shown in parentheses. Estimates are obtained from linear probability models with fixed effects on samples of single childless women and lone mothers. Estimates in bold are significant at the 5 percent level. The other variables used in estimation are: a quartic polynomial in age; number of children by age group (3 groups: ages 0-4, ages 5-10, and ages 11-18); dummy variables for: ethnic origin (4 dummies; white is the base category), highest educational qualification (5; no qualification), housing tenure (2; owner) region of residence (16; Greater London); and interactions between a woman's age and number of children by age group, age and the educational group dummies, and education and number of children by age group.

Table 3. Effects of Welfare Reform on Eligible Employment Transitions

	(i) Full sample with linear trends	(ii) Low education sample with linear trends	(iii) Full sample with quadratic trends	(iv) Low education sample with quadratic trends
Persistence Probability[a]	**0.065**	**0.060**	**0.065**	**0.058**
	(0.017)	(0.022)	(0.024)	(0.025)
Observations	6,123	4,020	6,123	4,020
Entry Probability[b]	**0.057**	**0.052**	**0.055**	**0.052**
	(0.021)	(0.023)	(0.024)	(0.025)
Observations	5,114	3,358	5,114	3,358

[a] Obtained from linear probability models of transitions in labor market states on the sample of single childless women and lone mothers. Conditional on $y_{i,t-1}$=1.
[b] Obtained from linear probability models of transitions in labor market states on the sample of single childless women and lone mothers. Conditional on $y_{i,t-1}$=0.
Notes: The term 'Observations' denotes the number of wave-on-wave state-specific transitions. Standard errors are shown in parentheses. Estimates in bold are significant at the 5 percent level. A list of all variables used in estimation is in the note to Table 2.

Table 4. Effect of Welfare Reform on Employment by Age of Youngest Child and Number of Children

	(i) Full sample with linear trends	(ii) Low education sample with linear trends	(iii) Full sample with quadratic trends	(iv) Low education sample with quadratic trends
A. Eligible employment				
One child aged 0-4	**0.103**	**0.088**	**0.092**	**0.082**
	(0.032)	(0.032)	(0.034)	(0.036)
One child aged 5-10	**0.083**	**0.073**	**0.077**	**0.069**
	(0.022)	(0.026)	(0.029)	(0.030)
One child aged 11-18	0.035	0.030	0.032	0.028
	(0.024)	(0.023)	(0.031)	(0.027)
Two children or more,	0.039	0.032	0.036	0.027
youngest 0-4	(0.041)	(0.043)	(0.043)	(0.048)
Two children or more,	0.004	0.012	-0.001	0.007
youngest 5-10	(0.033)	(0.048)	(0.026)	(0.036)
Two children or more,	-0.005	-0.0002	-0.003	-0.004
youngest 11-18	(0.027)	(0.012)	(0.035)	(0.021)
B. Persistence Probability				
One child age 0-4	**0.128**	**0.117**	**0.121**	**0.114**
	(0.044)	(0.047)	(0.050)	(0.051)
One child aged 5-10	**0.073**	**0.069**	**0.067**	**0.061**
	(0.028)	(0.032)	(0.031)	(0.030)
One child aged 11-18	0.028	0.022	0.027	0.020
	(0.040)	(0.036)	(0.038)	(0.034)
Two children or more,	0.022	0.017	0.021	0.019
youngest 0-4	(0.020)	(0.031)	(0.026)	(0.034)
Two children or more,	0.006	0.004	0.005	-0.002
youngest 5-10	(0.033)	(0.036)	(0.034)	(0.011)
Two children or more,	0.011	0.003	0.008	-0.004
youngest 11-18	(0.034)	(0.041)	(0.033)	(0.039)
C. Entry Probability				
One child age 0-4	**0.081**	**0.074**	**0.074**	**0.070**
	(0.033)	(0.031)	(0.032)	(0.032)
One child age 5-10	**0.049**	0.042	0.042	0.043
	(0.025)	(0.037)	(0.038)	(0.035)
One child age 11-18	0.041	0.029	0.036	0.030
	(0.027)	(0.026)	(0.030)	(0.028)
Two children or more,	0.021	0.015	0.017	0.011
youngest 0-4	(0.022)	(0.024)	(0.024)	(0.029)
Two children or more,	0.011	-0.001	0.006	-0.004
youngest 5-10	(0.023)	(0.015)	(0.021)	(0.014)
Two children or more,	0.0004	-0.008	-0.009	-0.011
youngest 11-18	(0.026)	(0.020)	(0.015)	(0.008)

Notes: Standard errors are shown in parentheses. Estimates in bold are significant at the 5 percent level. Definitions, number of observations, and list of variables used in estimation are in the notes to Tables 2 and 3.

Table 5. Effects of Welfare Reform on Childcare Use and Costs and Employment Transitions with Childcare Use, Overall and by Age of Youngest Child and Number of Children – Full Sample

	Linear trends		Quadratic trends	
A. Utilization and costs				
Paid childcare usage (N=5,283)[a]	**0.028**	(0.012)	**0.028**	(0.013)
Weekly childcare costs (N=351)[b]	**17.32**	(6.64)	**18.90**	(7.08)
B. Utilization by child's age (N=5,283)				
One child age 0-4	**0.040**	(0.014)	**0.037**	(0.015)
One child aged 5-10	**0.031**	(0.012)	0.030	(0.017)
One child aged 11-18	-0.007	(0.023)	-0.004	(0.020)
Two children or more, youngest 0-4	0.018	(0.022)	0.011	(0.025)
Two children or more, youngest 5-10	0.007	(0.026)	0.002	(0.031)
C. From nonworking to working 16+ and using paid childcare (N=5,114)[c]				
Overall	**0.035**	(0.015)	**0.033**	(0.015)
One child age 0-4	**0.050**	(0.021)	**0.046**	(0.021)
One child aged 5-10	**0.028**	(0.013)	0.024	(0.017)
One child aged 11-18	0.009	(0.020)	0.008	(0.023)
Two children or more , youngest 0-4	0.011	(0.024)	0.010	(0.029)
Two children or more, youngest 5-10	0.003	(0.027)	0.002	(0.034)
Two children or more, youngest 11-18	-0.004	(0.018)	-0.001	(0.030)
D. From working 16+ to working 16+ and using paid childcare (N=6,123)[d]				
Overall	**0.038**	(0.018)	**0.037**	(0.018)
One child age 0-4	**0.071**	(0.025)	**0.067**	(0.028)
One child aged 5-10	**0.040**	(0.017)	**0.041**	(0.019)
One child aged 11-18	0.007	(0.008)	0.007	(0.009)
Two children or more, youngest 0-4	**0.020**	(0.010)	0.019	(0.011)
Two children or more, youngest 5-10	0.014	(0.027)	0.013	(0.027)
Two children or more, youngest 11-18	0.004	(0.011)	0.004	(0.012)

[a] Obtained from linear probability models with individual fixed effects on the subsample of lone mothers. The dependent variable takes value one if the mother works, has at least one child aged 12 or less, and pays for childcare arrangements, and zero otherwise. N is number of observations.

[b] Obtained from linear regression models with individual fixed effects on the subsample of lone mothers who work, have at least one child aged 12 or less, and report positive expenditures on childcare arrangements. The weekly childcare expenditures are expressed in constant (2001) prices. N is number of observations.

[c] Obtained from linear probability models on the subsample of single childless women and lone mothers. The nonworking (origin) state includes women who work less than 16 hours per week. N is the number of wave-on-wave state-specific transitions.

[d] Obtained from linear probability models of transitions in labor market states on the sample of single childless women and lone mothers. N is the number of wave-on-wave state-specific transitions.

Notes: Standard errors are shown in parentheses. Estimates in bold are significant at the 5 percent level. All relevant information (explanatory variables, base categories, and definitions) is listed in the notes to Tables 2-4, except that the trend variables specific to single women without children are set to zero in panels A-C.

Table 6. Effects of Welfare Reform on Partnership Formation

	(i) Full sample with linear trends	(ii) Low education sample with linear trends	(iii) Full sample with quadratic trends	(iv) Low education sample with quadratic trends
Overall	**-0.024**	**-0.020**	-0.018	-0.019
	(0.008)	(0.008)	(0.013)	(0.016)
One child age 0-4	**-0.028**	**-0.025**	**-0.026**	**-0.023**
	(0.012)	(0.011)	(0.012)	(0.011)
One child aged 5-10	-0.021	-0.015	-0.017	-0.015
	(0.013)	(0.012)	(0.021)	(0.016)
One child aged 11-18	-0.010	-0.006	-0.008	-0.005
	(0.017)	(0.025)	(0.020)	(0.026)
Two children or more,	**-0.025**	**-0.022**	**-0.023**	**-0.020**
youngest 0-4	(0.011)	(0.009)	(0.011)	(0.008)
Two children or more,	**-0.015**	-0.014	-0.013	**-0.017**
youngest 5-10	(0.007)	(0.009)	(0.012)	(0.008)
Two children or more,	0.006	0.0007	0.0001	0.002
youngest 11-18	(0.019)	(0.013)	(0.026)	(0.018)

Notes: Standard errors are shown in parentheses. Estimates in bold are significant at the 5 percent level. Estimates are obtained from linear probability models of transitions into partnership (marriage or cohabitation) on the subsample of single childless women and lone mothers. For each woman, the dependent variable takes value zero if the woman is single, and value one in the period when she forms a union (after which her observations are censored). Multiple entries for the same woman are allowed. Explanatory variables are listed in the notes to Tables 2 and 4 above. The trends (either linear or quadratic) are allowed to differ by children's age group. N is equal to 15,634 in the regressions of columns (i) and (iii) and 10,265 in the regressions of columns (ii) and (iv).

Table 7. Effects of Welfare Reform on Fertility

	Linear trends	Quadratic trends
A. Birth rate (lone mothers only)[a]		
Overall	-0.007	-0.007
	(0.005)	(0.008)
One child aged 0-4	-0.006	-0.004
	(0.008)	(0.011)
One child aged 5-10	-0.010	-0.009
	(0.011)	(0.013)
One child aged 11-18	-0.009	-0.010
	(0.012)	(0.016)
Two children or more, youngest 0-4	-0.003	-0.001
	(0.017)	(0.017)
Two children or more, youngest 5-10	-0.002	-0.003
	(0.014)	(0.015)
B. Entry into lone motherhood (single childless only)[b]	-0.0016	-0.0010
	(0.0021)	(0.0027)

[a] Obtained from linear probability models of transitions into birth on the subsample of lone mothers from the second time they were observed in the panel onwards. Explanatory variables are listed in Tables 2 and 4, except that the trend variables specific to single women without children are set to zero. In the regressions that distinguish number of children and age of the youngest child, the base category is two or more children with the youngest aged 11-18. All regressions are performed on 4,782 observations.

[b] Obtained from linear probability models of transitions into lone motherhood for the subsample of single childless women. For each woman, the dependent variable takes value zero if the woman is single childless, and value one in the period when she has a child (after which her observations are censored). Explanatory variables are listed in Tables 2 and 4, except that the trend variables specific to single mothers are set to zero. Both regressions are performed on 6,410 observations.

Notes: Standard errors are shown in parentheses.